

This PDF is a selection from an out-of-print volume from the National Bureau of Economic Research

Volume Title: The Evolution of Retirement: An American Economic History, 1880-1990

Volume Author/Editor: Dora L. Costa

Volume Publisher: University of Chicago Press

Volume ISBN: 0-226-11608-5

Volume URL: <http://www.nber.org/books/cost98-1>

Publication Date: January 1998

Chapter Title: Income and Retirement

Chapter Author: Dora L. Costa

Chapter URL: <http://www.nber.org/chapters/c6109>

Chapter pages in book: (p. 32 - 59)

3 Income and Retirement

But it is pretty to see what money will do.

Samuel Pepys (1667)

Retirement requires income, whether in the form of state-provided retirement or disability benefits, private pensions, income from other family members, or assets. Researchers have investigated the role that each of these income sources plays in the retirement decision, largely using cross-sectional data for the years after the 1960s. But the applicability of cross-sectional estimates to periods outside the sample range is questionable. Seventy percent of the decline in the labor force participation rates of men age sixty-five or older occurred before 1960. Retirement rates were already high by 1960, and thus only large benefit increases could have enticed those remaining in the labor force to have withdrawn. To understand why retirement rates increased prior to 1960, we must examine earlier data.

An analysis of retirement requires information on retirement status, demographic characteristics (e.g., age), health, a proxy for the opportunity cost of not working (e.g., forgone income or occupation), and retirement income (e.g., pension amount). These are very strict data requirements. One of the few sources of information on the elderly of the past, the census, allows us to relate retirement status only to demographic characteristics, not to wealth. Fortunately, a data set that meets our requirements can be created from records generated by the Union army pension program.

This chapter will focus on the determinants of work levels in both 1900 and 1910 among white Union army veterans receiving Union army pensions. These men were the first cohort to reach age sixty-five in the twentieth century. They also represent a very broad cross section of the population. Eighty-one percent of all white, northern men born in 1843 served in the Union army during the Civil War. These men became eligible for an extremely generous pension, and the copious records generated by the Pension Bureau bureaucracy allow us to reconstruct their life histories. Additional information can be gathered by linking pension records to other sources, such as census manuscripts. The resulting

data set provides a unique picture of the life of the elderly at the turn of the century.

The data set created from Union army records provides information on retirement income in the form of Union army pensions. The receipt or level of Union army pensions, which replaced about 30 percent of the income of an unskilled laborer, did not depend on current income or past wages. Rather, their generosity was determined by the pensioner's health. Because the amount received also depended on whether the veteran could trace his disability to the war and not just on the seriousness of his infirmity, the effect of pensions on labor supply can be disentangled from that of health. Therefore, Union army pensions can be used to estimate a pure income effect on labor supply, thus revealing the effect of income growth on retirement and bearing on income effects arising from the Social Security program.

3.1 The Economics of Retirement

Many factors are likely to affect the retirement decision. As health deteriorates with age, work may become more arduous, and therefore men's desire to leave paid labor may increase. After initially rising with age, earnings generally decline with age, thus increasing the incentive to retire. Earnings today peak at age fifty to fifty-nine, whereas in the past earnings peaked at ages thirty to thirty-nine and declined by almost 30 percent by age sixty (Haber and Gratton 1994, 76). In addition, retirement income tends to be lower than income while working. By continuing to work, not only do men enjoy a higher income than they would if they retired, but they are also able to accumulate more savings or increase their entitlements to Social Security and private pension benefits. They can thus support higher retirement consumption at a later date—a date when there will be fewer remaining years of life over which support would have to be provided. A richer society is able to support more years of retirement because, when wages are low, men find it hard to accumulate enough savings to finance their retirement.

Not all retirement income is earned during the working years. Some men inherit wealth from their parents, and today most receive retirement benefits from the state. Until recently, retiring cohorts have received aggregate Social Security benefits that far exceed the present value of the contributions made by them and their employers (Boskin and Shoven 1987). The Civil War cohort was even more fortunate. Union army pensions were given regardless of labor force participation status and therefore both directly increased the income of retired veterans and allowed them to accumulate more wealth during their working years, thus enabling them to retire earlier.

Union army pensions represent a pure income effect on labor supply and therefore should have induced more men to retire at any given age. The question I pose is by how much Union army pensions reduced labor supply for veterans in their later years. This question can be answered with the help of a

simple model of the retirement decision. Although the model is so oversimplified that it cannot literally be true, when interpreted with care, it enables us to judge the effect of pensions on retirement rates.

At any date, a veteran can be thought of as making a choice between retirement and labor force participation. The well-being in each option will be determined by the income flows associated with each option, how enjoyable the veteran finds leisure and how unpleasant work, and the stigma costs of not working, among other factors. Well-being or utility when not working can be written as

$$U_w(Y + B + N, \bar{H}; \mathbf{Z})$$

and utility when not working as

$$U_l(B + N, 0; \mathbf{Z}),$$

where Y is labor market income, B is pension income, N is other non-labor market income, \mathbf{Z} is a vector of demographic and socioeconomic variables likely to affect utility (such as age and number of children), and \bar{H} is hours of market work. Then, assuming that the utility functions are linear in their arguments and that differences in tastes across individuals produce utility functions containing normally distributed random taste shifters, the individual can be thought of as evaluating the decision function

$$I^* = U_l(B + N, 0; \mathbf{Z}) - U_w(Y + B + N, \bar{H}; \mathbf{Z}).$$

Although the value of I^* is not observed, a discrete retirement indicator is observed, given by

$$I = \begin{cases} 0 & \text{if } I^* < 0, \\ 1 & \text{otherwise,} \end{cases}$$

where 1 represents retirement and 0 labor force participation. We do not know whether well-being when retired greatly or only narrowly outweighs well-being when not retired, nor are we even assuming that an individual can look forward to a comfortable retirement. However, if an individual is retired, then utility when retired must be greater than utility when not.

The decision function that the veteran evaluates can be rewritten as

$$\begin{aligned} I^* &= U_l(B + N, 0; \mathbf{Z}) - U_w(Y + B + N, \bar{H}; \mathbf{Z}) \\ &= -\mathbf{X}'\boldsymbol{\beta} + \boldsymbol{\varepsilon}, \end{aligned}$$

where \mathbf{X} is a vector containing Y , B , N , \bar{H} , and \mathbf{Z} , $\boldsymbol{\beta}$ is a parameter vector, and $\boldsymbol{\varepsilon}$ is a standard normal error term. Recall that pensions might have two different effects on the retirement decision. Pensions will directly affect income flows, as captured by the incorporation of the term B in the utility functions, and the

receipt of pensions in the past may affect the value of N , allowing veterans to retire earlier. Only the direct effect of pensions will be estimated. The estimated effect of pensions will therefore be a lower-bound estimate. Using the indicator function, I , the effect of characteristics such as pension amount included in X on the probability of retirement will be estimated by means of a probit,

$$\text{prob}(I = 1) = \text{prob}(\epsilon < X'\beta) = \Phi(X'\beta),$$

where $\Phi(\)$ is a standard normal cumulative distribution function. Knowing how the receipt and level of Union army pensions were determined, we can use this model to identify the effect of Union army pensions on retirement rates. The next section therefore provides a brief description of the Union army pension program and records. A more detailed description is given in appendix A at the end of the book.

3.2 Civil War Pensions and Union Army Records

By 1900 the scope of the pension program, run for the benefit of Union veterans and their dependent children and widows, was enormous. Benefits consumed almost 30 percent of the federal budget (Vinovskis 1990), and veterans lobbied vociferously for high tariffs to continue feeding the federal surplus (Glasson 1918a, 218–19). Even though Confederates were ineligible and immigration increased the population, a large percentage of the population was collecting benefits. Among all white males, 35 percent of those aged fifty-five to fifty-nine were on the pension rolls, 21 percent of those aged sixty-five to sixty-nine, 14 percent of those aged sixty-five to sixty-nine, and 9 percent of those aged seventy or older. The annual value of the average veteran pension was \$135, or 53 percent of the annual income of farm laborers, 36 percent of that of nonfarm laborers, 20 percent of that of carpenters, blacksmiths, or salesmen, and 12 percent of that of landlords or merchants.¹

The generosity of the Union army pension program arose from a number of causes. Like the elderly today, elderly Union veterans wielded considerable political might. Union pensions were a prominent election issue throughout the latter half of the nineteenth century and the beginning of the twentieth. While Union veterans constituted a relatively small group, they were, however, extremely well organized, again like the elderly today. The veterans' organization sent lobbyists to Congress and communicated with its members through local chapters and through newspapers. This organization was able to form an effective political alliance with manufacturing interests to maintain high tariffs and to redistribute the resulting revenue to its members. Because veterans had defended the Union, and because the federal treasury was relatively flush, veterans' pensions had the backing of many Americans.

Congress established the basic system of pension laws, known as the General Law pension system, in 1862 to provide pensions to both regular recruits

and volunteers who were disabled as a direct result of military service. The dollar amount received depended on the degree of disability, where disability was determined by the applicant's capacity to perform manual labor. Under later reinterpretations the total disability standard soon meant incapacity to perform even lighter types of manual labor. In fact, men judged disabled continued to labor in physically demanding, manual occupations. Inability to perform manual labor remained the standard in this and all subsequent laws, regardless of the wealth of the individual or his ability to earn a living by other than manual means. Withdrawal from the labor force was not a necessary prerequisite for the receipt of a pension. If the claimant had lesser disabilities, then he received an amount proportionate to his disabilities. Application was made through a pension attorney, and the degree of disability was determined by a board of three local doctors employed by the Pension Bureau and following guidelines established by the bureau.

An act of 27 June 1890 instituted a universal disability and old-age pension program for Union veterans. According to the veterans' lobby, the new law would "place upon the rolls all survivors of the war whose conditions of health are not practically perfect" (quoted in Glasson 1918a, 233). In fact, within a year of the act's passage, the number of pensioners on the rolls more than doubled. Any disability now entitled a veteran to a pension. However, an applicant who could trace his disability to wartime service received substantially more for the same disability than his counterpart who could not. By 1900 men who could not claim a disability of service origin received from \$6.00 to \$12.00 per month or from 19 to 38 percent of the monthly income of a laborer, while men who could claim a war-related disability generally received a pension ranging from \$6.00 to \$35.00 per month or up to 109 percent of the monthly income of a laborer. Although few men were eligible, pensions for war-related disabilities could be as much as \$100 per month, close to one-third the yearly income of a laborer. By 1900, 58 percent of veterans who were Union army pensioners were collecting a pension for disabilities unrelated to wartime service (U.S. Bureau of Pensions 1900).

Table 3.1 illustrates differences in pension amounts according to whether a veteran could trace his disability to the war and thus fell under the 1862 law rather than the 1890 law. Controlling for health, men who fell under the 1862 rather than the 1892 law received much larger pensions. Fifty-six percent of the very disabled were receiving pensions of over \$12.00 and 44 percent pensions of \$12.00 or less. That individuals of the same health status received different pension amounts will prove to be very important to my subsequent estimation strategy. The difference in pension amount will allow me to identify the effect of pensions and of health on the retirement decision.

The Pension Bureau instructed the examining surgeons in 1890 to grant a minimum pension to all men at least sixty-five years of age, unless they were unusually vigorous. At age seventy-five men became eligible for an even larger pension. In 1904, Executive Order 78 officially authorized the Pension Bureau

Table 3.1 Monthly Pension Means and Pension Rates by Percentile, by Health Status and Law, 1900 (\$)

	Mean	Percentile				
		10	25	50	75	90
All veterans	12.9	6	8	12	14	24
General Law	17.6	8	12	14.5	24	30
1890 law	9.4	6	8	10	12	12
Health:						
Good	9.8	6	8	8	12	12
Fair	11.4	8	8	12	12	16
Poor	17.5	10	12	15	24	30
General Law and:						
Health good	14.3	8	8	12	15	18
Health fair	14.1	10	12	12	16	17
Health poor	20.1	12	15	17	24	30
1890 law and:						
Health good	8.6	6	6	8	12	12
Health fair	9.6	6	8	10	12	12
Health poor	10.9	8	10	12	12	12

Source: Calculated from the data used in the estimation. The health variable used is based on the ratings of the examining surgeons.

to grant pensions on the basis of age. The Service and Age Pension Act of 6 February 1907 marks official congressional recognition of age as sufficient cause to qualify for a pension. Veterans aged sixty-two to sixty-nine now received \$12.00 per month, those aged seventy to seventy-four \$15.00 per month, and those older than seventy-four \$20.00 per month. Because most eligible veterans were already on the rolls, this act mainly induced pensioners to switch from the 1890 law to the 1907 law; it did not increase the total number of pensioners. Whereas in 1900 slightly more than half of all veterans were collecting a pension under the 1890 law, by 1910 64 percent of all veterans were collecting under the 1907 law, 22 percent under the General Law, and only 14 percent under the 1890 law (U.S. Bureau of Pensions 1910). Pension amount now depended primarily on age and whether a veteran could trace his disabilities to the war (see table 3.2).

The samples of veterans that I use are random samples of the veteran population and contain disproportionate numbers of rural, native-born men.² Nonetheless, as discussed in greater detail in appendix A, we can draw inferences from these samples for the entire population. The men in these samples do not differ in observable characteristics (e.g., home ownership, marital status, literacy, and age) from the male population in northern states. Controlling for rural residence, they resemble the northern population of men in occupation and foreign birth. The samples are also representative of the northern population in terms of mortality and wealth.

Table 3.2 Monthly Pension Means and Pension Rates by Percentile, by Age, Health Status, and Law, 1910 (\$)

	Mean	Percentile				
		10	25	50	75	90
All veterans	16.5	12	12	15	20	24
General Law	22.3	17	17	24	24	30
1890 law	11.7	12	12	12	12	12
1907 law	14.5	12	12	12	15	20
Good health	16.6	12	12	15	20	24
Poor health	18.0	12	12	15	20	30
Age < 70	15.0	12	12	12	17	24
Age ≥ 70	18.8	15	15	17	20	30
Age < 70 and:						
General Law	21.3	14	17	17	24	30
1890 law	11.7	12	12	12	12	12
1907 law	12.2	12	12	12	12	12
Age ≥ 70 and:						
General Law	23.9	17	17	24	30	30
1890 law	11.5	10	12	12	12	12
1907 law	16.9	15	15	15	20	20

Source: Calculated from the data later used in estimation. The health variable used was based on the ratings of the examining surgeons.

The reconstruction of the life histories of men who fought for the Union army in the American Civil War represents a unique data source on a past population. Because such a large percentage of the male population fought in the Civil War, we can generalize from this sample to the population as a whole. Because the peculiar rules of the Union army pension program led to men who were equally disabled receiving very different pension amounts, we can identify the effect of pensions on the retirement behavior of veterans. Because neither demographic nor occupational characteristics nor the lawyer through whom the pensioner applied predicts either the ratings of the examining surgeons or pension amount, we can be sure that our results are not tainted by past fraud. Furthermore, whether a man could trace his disability to the war can be used to identify the relation between retirement and pension amount free from the confounding effects of potential endogeneity between pension amount and retirement status. This is because the ability to claim war-related disabilities or receive a pension under the General Law, and therefore to receive a larger pension, is arguably unrelated to unobservable determinants of retirements. Whether a veteran received a pension under the General Law depended on the often incorrect medical theories of the time.

3.3 Pensions and Retirement

Compared with the general population, Union army veterans retired at a greater rate at all ages. This is evident in figure 3.1, which compares retirement

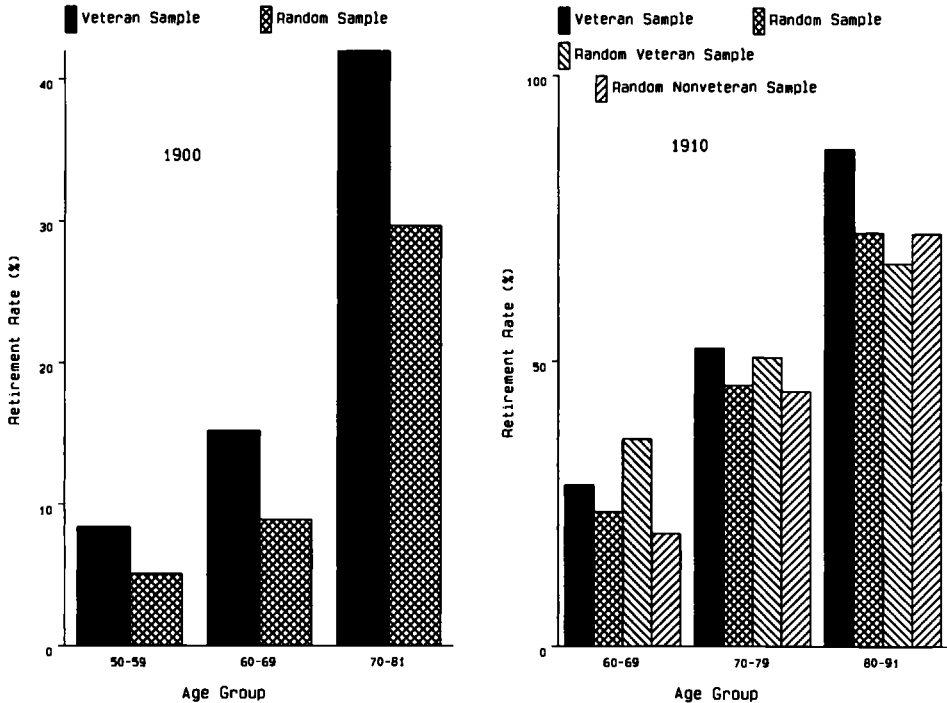


Fig. 3.1 Comparison of retirement rates among veterans and a random sample of white men, 1900 and 1910

Note: To ensure comparability with the veteran sample, the random samples were reweighted to have the same age distribution as the veteran sample and were restricted to men who were either born in a Union state or who, if foreign-born, immigrated prior to the Civil War. The random samples contain both veterans and nonveterans. Because one of the questions asked in the 1910 census was veteran status, by type of veteran, retirement rates for both veterans and nonveterans in the random sample are given in 1900. The random samples were drawn from Ruggles and Sobek (1995). All samples were limited to the noninstitutionalized.

rates by age group among men in the Union army sample with the general population in both 1900 and 1910. Thus, among men aged sixty to sixty-nine in 1900, retirement rates among veterans were 15 percent, whereas they were only 9 percent in the general population. Among men of the same age in 1910, they were 28 percent among veterans but only 22 percent in the general population. In 1910, when veteran status of the general population is known, retirement rates of veterans in the general population can be compared to those of nonveterans. With the exception of ages eighty to ninety-one, an age group of which veterans composed a relatively small fraction in 1910, the retirement rates of veterans are sharply higher than those of nonveterans. This difference between veteran and nonveteran retirement rates is underestimated because underenumeration of veterans in the 1910 census implies that the retirement rates of nonveterans are overestimated.

Figure 3.1 suggests that retirement rates were higher for veterans because

of Union army pensions, but other reasons could explain this finding. Morbidity rates may have been higher for veterans, and poor health, not the pension, may have been the driving factor. After all, 31 percent of veterans claimed to have had an injury or gunshot wound while in service. Even those who had not been injured may still have suffered long-term effects from the infectious diseases endemic in the army. Over three-quarters of men paid a visit to the camp hospital or saw a camp surgeon during their service. The effect of pensions therefore needs to be distinguished from that of health by comparing retirement rates of veterans with similar health status but different pension levels within the Union army sample.

I use the ratings of the examining surgeons to construct health variables. The surgeons rated each specific disability, and I use the surgeons' ratings to classify each veteran as being in good, fair, or poor health. Other health proxies could have been used, but the results remain unchanged regardless of whether the surgeons' ratings, the Body Mass Index (see sec. 4.1.1), or the presence of a chronic disease is used.

Figure 3.2 shows that men younger than seventy were more likely to be retired either if they were receiving higher pensions or if they were in poorer

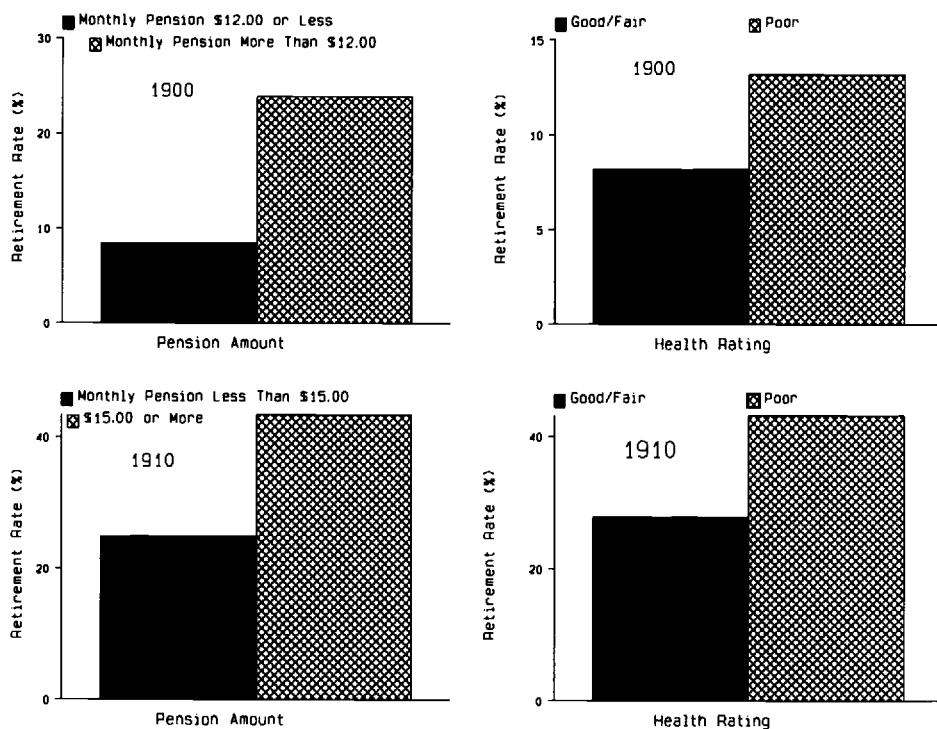


Fig. 3.2 Retirement rates among veterans under seventy years of age by pension amount, 1900 and 1910

health. In 1900, 8 percent of veterans receiving a monthly pension of \$12.00 or less were retired, as opposed to 24 percent of those receiving a pension of more than \$12.00. In 1910, 43 percent of veterans receiving a monthly pension of \$15.00 or more were retired, as opposed to only 25 percent of those receiving a pension of less than \$15.00. Retirement rates were about 40 percent higher among those in poor health than among those in good or fair health.

Figure 3.2 does not provide conclusive proof that veterans were responding to increased income. The receipt of a large pension and disability were correlated because of the rules of the pension program. Therefore, figures 3.3 and 3.4 show retirement rates by health category and by pension amount among

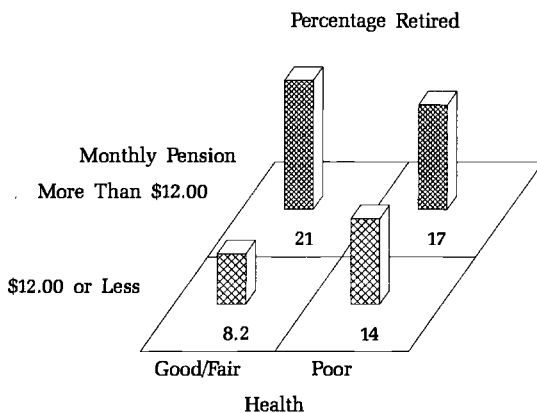


Fig. 3.3 Retirement rates by disability status among veterans under seventy years of age, 1900

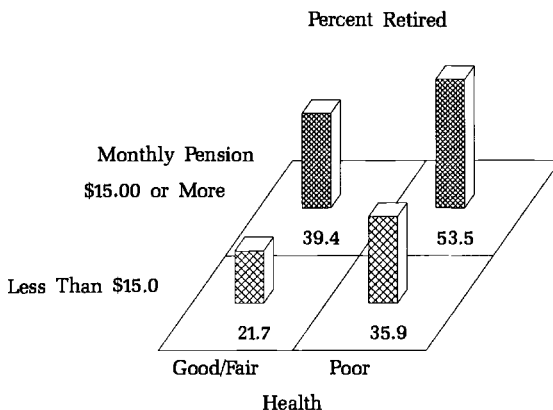


Fig. 3.4 Retirement rates by disability status among veterans under seventy years of age, 1910

men younger than seventy years of age. Retirement rates were higher at higher pension amounts among those in good or fair health and for those in poor health, and the difference in retirement rates by health status is statistically significant in 1910. But, because men with war-related disabilities, and thus those eligible for a large pension, were not a random sample, this evidence is still insufficient proof of a relation between income and retirement.

Men with war-related disabilities had different service, occupation, and health histories. Compared with men whose disabilities were not service related, men who claimed a disability of service-related origin were more likely to have been discharged for a disability, to have been prisoners of war, and to have been volunteers. They were also more likely to have been farmers and less likely to have been professionals and proprietors. Interestingly, there is no significant difference in the percentage claiming injury or gunshot wounds, but men without service-related disabilities were more likely to claim rheumatism, gastrointestinal disorders other than diarrhea, respiratory disorders, hernias, and conditions that could not be classified. Men who could trace their disabilities to the war entered the rolls earlier, were judged by the surgeons to be in worse health, and were less likely to live out their expected life span. Even though men who could trace their disabilities to the war had different characteristics than those who could not, an income effect from pensions can still be identified provided that I can control for these characteristics.

Recall that the empirical specification that will be estimated is a probit,

$$\text{prob}(I = 1) = \text{prob}(\epsilon < X'\beta) = \Phi(X'\beta),$$

where I is equal to one if the individual is retired and 0 otherwise; X is a vector containing both labor and nonlabor income, pension income, and demographic and socioeconomic variables; β is a parameter vector; and $\Phi(\)$ is a standard normal cumulative distribution function. Although direct evidence on earnings and wealth is not available, several proxies are. I use past occupation as a proxy for the opportunity cost of not working, where occupation is classified as farmer, professional or proprietor, artisan, and semiskilled or unskilled laborer, including farm laborer. For men in 1910 this is either occupation in 1900 or, for those retired in 1900, occupation at an earlier date as given in the pension records. For men in 1900 this is either occupation in 1900, if working, or past occupation as given in the pension records. Men whose occupation could not be discerned from the pension records were assigned to an occupation class on the basis of their occupation at enlistment and their probability of switching occupation category given their individual characteristics. These men could just as well have been assigned to random occupations; the results on pension amount remain unchanged. Past occupation may be a poor proxy for opportunity cost if the ill are no longer able to work in their usual occupation, and for these men the opportunity cost of retirement is underestimated. Therefore, the effect of pensions may be overestimated.

Other indicators of lower earnings are illiteracy and foreign birth. Marital status may also be measuring earnings if employers favored married men or if married men were more skilled. In 1900, married males in manufacturing earned 17 percent more than unmarried males, controlling for the observable characteristics of workers and their jobs (Goldin 1990, 102). Among home owners in cities, letting rooms to boarders increased family income but may be symptomatic of economic difficulties (Modell and Hareven 1973). The hire of a servant is an indicator of affluence, but servants may have played an important role in family market enterprises, particularly among farmers or small business owners. Home ownership meant that the person had wealth because a substantial down payment, generally equal to half the value of the purchased property, was required (Haines and Goodman 1992).³ Because long-term unemployment was often a prelude to retirement, higher unemployment in the veteran's current state, available in 1900 but not in 1910, may have induced more retirement.⁴ Additional control variables are region of residence, extent of urbanization, and whether the veteran was discharged for disability from the army, a measure of early health status.

The probit results for 1900 are presented in table 3.3. It should not come as a surprise that the older were more likely to retire. Note that a single linear term in age is included because tests revealed that in 1900 the probability of retirement did not increase sharply at a specific age.⁵ Those living in states with high unemployment rates and those in poor health were significantly more likely to retire. Those who owned no property were also significantly more likely to retire, but, because property ownership is known only for household heads, this variable is also an indicator of living arrangements. Having servants, boarders, and four or more dependents in the household reduced the probability of retirement, but not significantly so, as did being married, foreign born, or illiterate. Nonfarmers were less likely to retire than farmers, and, in the case of professionals and proprietors, the difference in retirement rates was statistically significant. I discuss the effect of occupation on retirement, particularly of farm occupation, in more detail in chapter 5. Interestingly, those who had been discharged from the service for disability were significantly less likely to be retired in 1900 even though they were in worse health than men who had not been discharged for disability. These men changed to a less physically demanding occupation after enlistment, which may have enabled them to remain in the labor force longer.

The effect on the probability of retirement of a unit change in one of the independent variables is given by the partial derivative of the probability function P with respect to that independent variable. Thus, a \$10.00 increase in monthly pension income raises the retirement probability by 0.09.⁶ Coefficients on interactions of pension amount with poor health and dummies for older ages and an above-average unemployment rate are small and insignificant.

As previously noted, those with higher pensions may have been less healthy,

Table 3.3 Probit of Determinants of Probability of Retirement, with Retirement Status as the Dependent Variable, 1900 (526 observations, pseudo $R^2 = .22$)

Variable	Mean	Est.	S.E.	$\partial P/\partial x$
Dummy = 1 if retired	.17			
Intercept		-12.14‡	2.24	
Monthly pension	12.94	.05‡	.01	.0090
Age	61.28	.05‡	.01	.0106
Dummy = 1 if does not own home	.34	.35†	.17	.0695
Discharged disability	.25	-1.63‡	.19	-.1229
Health good	.22			
Health fair	.35	.39*	.23	.0765
Health poor	.25	.37*	.25	.0717
Health status unknown	.18	.46*	.26	.0905
Farmer	.46			
Professional or proprietor	.18	-.48†	.24	-.0935
Artisan	.14	-.09	.23	-.0168
Laborer	.22	-.02	.21	-.0046
Servant in house	.02	-.96	.67	-.1891
Boarder in house	.05	-.26	.41	-.0515
4 or more dependents	.14	-.46	.29	-.0895
Married	.85	-.25	.20	-.0486
Foreign-born	.10	-.13	.25	-.0249
Illiterate	.06	-.02	.31	-.0031
Lives in East	.21			
Lives in Midwest	.73	.42*	.24	.0828
Lives in other region	.06	-.28	.47	-.0540
Urban county	.37	.41†	.17	.0799
Mean duration of unemployment for manufacturing workers by state	3.62	1.86‡	.63	.3644

Note: The omitted dummies are good health, farmer, and eastern residence. The symbols *, †, and ‡ indicate that the coefficient is significantly different from zero at at least the 10 percent, 5 percent, and 1 percent levels, respectively. $\partial P/\partial x = \beta(1/n) \sum \phi(x'\beta)$, where ϕ is the standard normal density, and $\partial P/\partial x$ is in probability units.

but their poorer health may be unobservable. Furthermore, although pensions were awarded regardless of participation status, nonparticipation may have been viewed by employees of the Pension Bureau as evidence of an inability to perform manual labor. It is therefore unclear whether we are measuring the effect of pensions on retirement rates or the effect of retirement on pensions. Pension status is potentially endogenous, and all coefficients may be biased. Fortunately, unbiased estimates of the coefficients can be obtained through the use of a proxy that is highly correlated with pension amount but uncorrelated with the error term appearing in the regression equation. Such a proxy is known as an *instrumental variable*.

The instrumental variable that I use is whether the veteran received a pension under the General Law, that is, whether he could trace his disability to his

wartime service. Recall that the ability to establish that a disability was related to wartime service depended on the recruit's record of military service and prevailing medical theory. The ability to establish whether a disability could be traced to wartime service predicts pension amount and is arguably not related to unobserved retirement determinants conditional on measured health status. Although the war disabled entered the pension rolls eight years earlier than those not war disabled, the war disabled were not necessarily disabled earlier in life and therefore did not necessarily receive less job training and therefore have lower opportunity costs of not working. Whether a recruit could trace his disability to the war does not predict whether his occupation in 1900 was of lower socioeconomic status than his occupation at enlistment. Furthermore, the fraction of men who were property owners does not vary by ability to establish whether a disability was war related. A dummy variable indicating whether a recruit could trace his disability to the war, that is, whether he fell under the General Law, is therefore used as an instrumental variable.

Assuming that whether a recruit could trace his disability to the war is a legitimate instrument, consistent estimates of pension amount on retirement can be obtained easily. In the first stage, pension amount is regressed on the exogenous variables, that is, all variables except for pension amount, and a dummy equal to one if the recruit could trace his disability to the war. In the second stage, a probit is estimated in which a retirement dummy is regressed on pension amount, the exogenous variables, and the residuals from the first stage. This method, known as two-stage conditional maximum likelihood, was developed by Rivers and Vuong (1988).⁷ A convenient feature of this estimation procedure is that it becomes possible to test statistically whether pension amount is determined by retirement status.⁸ If pension amount is not determined by retirement status, then the coefficient on the residuals will be equal to zero when uncorrected standard errors are used. In fact, the hypothesis that the coefficient on the residuals is not equal to zero can be rejected only at the 85 percent level of significance, suggesting that endogeneity is not a problem.

Table 3.4 compares probit estimates with those from a two-stage conditional maximum likelihood procedure among men for whom information on whether the disability can be traced to the war is available.⁹ The first-stage estimates are also presented. The two columns that should be compared are those giving the derivatives and marked $\partial P/\partial x$. These columns show that the change in the coefficient on pension amount is small, with the estimated mean effect of a dollar increase in monthly pension amount on retirement probability rising from 0.0092 when a probit is estimated to 0.0101 when two-stage conditional maximum likelihood estimation is used. As in the simple probit, coefficients on interactions of pension amount with other variables are small and insignificant.

Endogeneity is not the only source of potential bias. Another possible source of bias is that from sample selection. If pensions affected survivorship, then the men surviving to 1900 will be a selected sample, and the coefficient on

Table 3.4 **Comparison of Probit and Two-Stage Conditional Maximum Likelihood Estimates of Determinants of Probability of Retirement, 1900 (485 observations)**

Variable	Probit: $\partial P/\partial x$	Two-Stage Conditional Maximum Likelihood				
		First Stage: Adj $R^2 = .33$		Second Stage: Pseudo $R^2 = .22$		$\partial P/\partial x$
		Est.	S.E.	Est.	S.E.	
Intercept		22.36‡	8.49	-12.48‡	2.57	
Monthly pension	.0092			.05†	.03	.0101
Age	.0116	.06	.05	.06‡	.01	.0115
Dummy = 1 if does						
not own home	.0651	-1.00	-1.00	.34*	.18	.0664
Discharged						
disability	-.1342	.91	.72	-.69‡	.21	-.1358
Health good						
Health fair	.0651	.15	.84	.34	.24	.0658
Health poor	.0745	3.42‡	.94	.33	.29	.0692
Health status						
unknown	.0921	2.12‡	.96	.45*	.27	.0892
Farmer						
Professional or						
proprietor	-.0802	-1.06	.85	-.40	.25	-.0787
Artisan	-.0255	-.64	.95	-.12	.24	-.0246
Laborer	.0190	-1.27	.83	.11	.22	.0246
Servant in house	-.2207	9.23‡	2.22	-1.15	.70	-.2269
Boarder in house	-.0617	-.06	1.34	-.31	.41	-.0621
4 or more						
dependents	-.0892	-1.22	.88	-.45	.29	-.0890
Married	-.0185	-.84	.90	-.10	.22	-.0189
Foreign born	-.0381	.03	1.00	-.19	.26	-.0376
Illiterate	-.0301	-.08	1.33	-.15	.32	-.0296
Lives in East						
Lives in Midwest	.0921	1.43*	.85	.49*	.25	.0961
Lives in other region	-.0447	3.14†	1.48	-.24	.48	-.0467
Urban county	.0893	-.62	.67	.46‡	.18	.0905
If disability						
traceable to war		6.83‡	.66			
Mean duration of						
unemployment in						
manufacturing by						
state	.3563	-4.73†	2.26	1.83‡	.65	.3615
Residuals first stage				-.01	.03	-.0010

Source: Costa (1995b).

Note: The first stage is a regression of pension amount on the exogenous variables and whether the disability was traceable to the war, that is, whether the veteran fell under the General Law. The second stage is a probit with the exogenous variables, pension amount, and the first-stage residuals as explanatory variables. The standard errors have been corrected. The symbols *, †, and ‡ indicate that the coefficient is significantly different from zero at at least the 10 percent, 5 percent, and 1 percent levels, respectively. $\partial P/\partial x = \beta(1/n) \sum \phi(x'\beta)$, where ϕ is the standard normal density, and $\partial P/\partial x$ is in probability units.

pensions will be biased.¹⁰ I tested whether pension amount affects life expectancy using a proportional hazards model where the dependent variable was the number of years lived after 1900. Controlling for health, pension amount did not affect the probability of mortality. Neither did date of entry, suggesting that duration of the receipt of a pension was not an important predictor of mortality.

Now that we can be confident that our results are not tainted by bias, we can estimate how responsive retirement rates are to changes in pension income. One way of measuring responsiveness is to calculate elasticities of labor force nonparticipation with respect to pension income from mean derivatives and retirement probabilities. In the probit specification in table 3.3 above, and at the pension mean of \$12.90, the elasticity is 0.73 (= 0.0090 [12.9/0.1589]), indicating that a 1 percent increase in pension amount increases retirement by 0.73 percent. Evaluated half a standard deviation below and above the pension mean, the elasticities are 0.53 (= 0.0076 [9.0/0.1289]) and 0.88 (= 0.0104 [16.8/0.1990]), respectively. Hence, the larger the pension income, the larger the elasticity. Using the two-stage conditional maximum likelihood method used in table 3.4, and evaluating at the pension mean, the elasticity of labor force nonparticipation with respect to pension income rises slightly to 0.80 (= .0101 [12.9/0.1625]).

The results obtained for 1900 should be compared with those for 1910, when veterans were ten years older. Probit results for 1910 are given in table 3.5. As in 1900, the older, those in poor health, and those who owned no property are significantly more likely to be retired. Having either a servant or two or more dependents in the household becomes a significant predictor of retirement. In contrast to the 1900 results, being foreign born raised the probability of retirement in 1910, a finding consistent with the pattern seen in the general population. Although the coefficient on whether a veteran was discharged for disability was no longer significant, those so discharged were still less likely to be retired. Professionals, proprietors, and artisans were less likely to be retired than farmers, and the difference between farmer and artisan retirement rates was statistically significant. Compared with farmers, laborers were more likely to be retired, but the difference in retirement rates is not statistically significant.

A \$10.00 increase in monthly pension income raises the retirement probability by 0.112. Once again, coefficients on interactions of pension amount with poor health and age dummies are small and insignificant, but the coefficients on the interactions of pension amount with the age dummies suggest that the effect of pensions is lower among older men. Two-stage conditional maximum likelihood estimation, using the 1862 law as an instrument, yielded coefficients almost identical to those obtained from the probit estimates. The estimated elasticity for 1910, evaluated at the pension mean of 16.94, is 0.47 (= 0.0112 [16.94/0.3989]), somewhat lower than the elasticity estimated for 1900.

Although an interaction between pension amount and age in the regression

Table 3.5 Probit of Determinants of Probability of Retirement, with Retirement Status as the Dependent Variable, 1910 (923 observations, pseudo $R^2 = 0.16$)

Variable	Mean	Est.	S.E.	$\partial P/\partial x$
Dummy = 1 if retired	.40			
Intercept		-6.42‡	.71	
Monthly pension	16.94	.03‡	.01	.0112
Age	69.19	.08‡	.01	.0246
Dummy = 1 if does not own home	.28	.34‡	.11	.1101
Discharged disability	.18	-.14	.12	-.0458
Health good or fair	.53			
Health poor	.34	.22†	.11	.0703
Health status unknown	.13	-.17	.16	-.0552
Farmer	.49			
Professional or proprietor	.19	-.11	.13	-.0360
Artisan	.14	-.39‡	.14	-.1249
Laborer	.17	.16	.13	.0527
Servant in house	.05	-.87‡	.25	-.2796
Boarder in house	.05	-.16	.21	-.0530
2 or more dependents	.21	-.30‡	.12	-.0976
Married	.78	.12	.12	.0385
Foreign born	.08	.34†	.17	.1114
Illiterate	.05	.14	.22	.0441
Lives in Midwest	.86	.21	.14	.0680
Urban county	.18	-.04	.13	-.0141

Note: The omitted dummies are good or fair health and farmer. The symbols *, †, and ‡ indicate that the coefficient is significantly different from zero at at least the 10 percent, 5 percent, and 1 percent levels, respectively. $\partial P/\partial x = \beta(1/n) \sum \phi(x'\beta)$, where ϕ is the standard normal density, and $\partial P/\partial x$ is in probability units.

equations was insignificant, there is some suggestion that the effect of pensions on retirement varied by age. When the 1910 sample is divided into those aged seventy or less and those older than seventy, the respective elasticities, evaluated at the pension means, are 0.62 (= 0.0123 [15.3/0.3033]) and 0.28 (= 0.0083 [20.2/0.5930]). The elasticity is lower in the older sample both because the increase in the probability of retirement for a dollar increase in monthly pension amount is smaller than in a younger sample and because the probability of retirement is much higher. At older ages, men's participation decision is less sensitive to changes in income.

Figure 3.5 simulates the effect on retirement rates of eliminating Union army pensions, showing that retirement rates in the Union sample and in the general population, which contains veterans collecting Union army pensions, would have fallen—which they did in fact do.¹¹ The resulting narrowing of differentials in retirement rates between the general population and the Union army sample suggests that much of the difference in retirement rates between veterans and the general population is due to pensions.

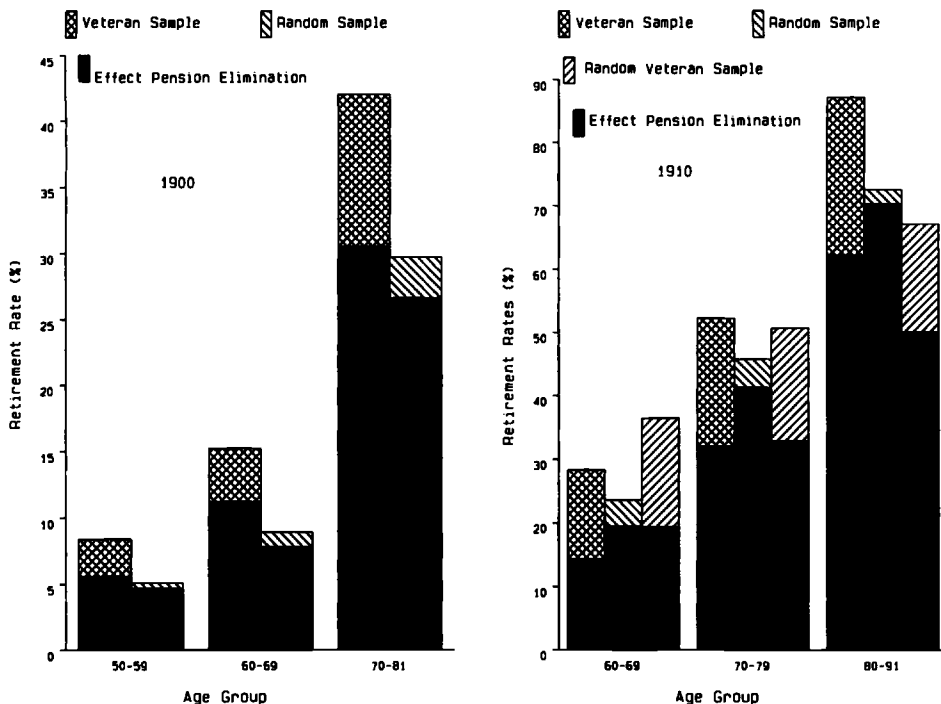


Fig. 3.5 Estimated Effect of Pension Elimination on Retirement Rates in 1900 and 1910

Note: Retirement rates assuming a pension elimination were calculated using the coefficients in tables 3.3 and 3.5. The random samples were drawn from Ruggles and Sobek (1995) and were limited to men who either were born in a Union state or, if foreign born, immigrated prior to the Civil War. The random samples contain both veterans and nonveterans and were reweighted to have the same age distribution as the veteran sample. Estimates of the fraction collecting Union army pensions were used to calculate retirement rates under a pension elimination. For details, see Costa (1995b).

3.4 Confederate and Union Veterans

Variation in pension amount by whether a recruit was able to trace his disability to the war has enabled me to separate the effect of pensions from that of health. Another source of variation in the Union army pension program was disparate treatment by type of veteran. Confederates were ineligible. In 1910 Union pensioners were collecting an average pension of \$171.90 per year, and about 90 percent of all Union veterans were collecting a pension. Although some Confederate states provided pensions, the average pension amount was just \$47.24 per year, and fewer than 30 percent of all Confederate veterans were collecting a pension.¹² Because the Union army pension was extremely generous while Confederate pensions were honorariums, then, if pensions matter, the difference in retirement rates between Union veterans and nonveter-

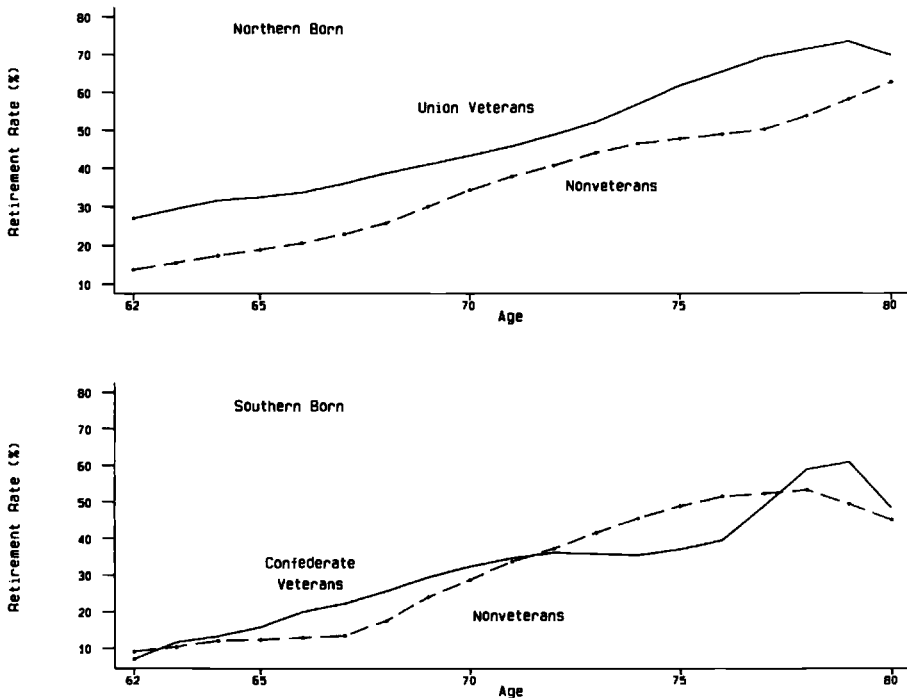


Fig. 3.6 Retirement rates by veteran status among northern-born and southern-born men aged sixty-two to seventy-nine in 1910

Note: Estimated from Ruggles and Sobek (1995). The trend in retirement rates was smoothed using Cleveland's lowest running line smoother with a band width of 0.4.

ans should be large and that between Confederate veterans and nonveterans small. This pattern is indeed observed in figure 3.6, where to control for differences in regional characteristics the sample was divided into those born in a Union state, that is, those at risk to serve in the Union army, and those born in a Confederate state, that is, those at risk to serve in the Confederate army. Because disability rates were probably higher among Confederate veterans than among Union veterans, relative differences in retirement rates between veterans and nonveterans in the two samples cannot be explained by health status.¹³

An alternative way to compare retirement rates among veterans and nonveterans within northern- and southern-born samples is to estimate two probits in which retirement status is the dependent variable. The advantage is that individual characteristics such as marital status, illiteracy, property ownership, region of residence, extent of urbanization, and whether servants or boarders are in the household can be controlled for. Table 3.6, which gives the probits, bears

Table 3.6

Probit Predicting Probability Retirement for Northern-Born and Southern-Born Aged 62–79 in 1910, with Retirement Status as the Dependent Variable (from Public-Use Census Sample)

Variable	Northern Born (4,517 Observations, Pseudo $R^2 = .10$)				Southern Born (1,224 Observations, Pseudo $R^2 = .17$)			
	Mean	Parameter Est.	S.E.	$\partial P/\partial x$	Mean	Parameter Est.	S.E.	$\partial P/\partial x$
Dummy = 1 if retired	.31				.24			
Intercept		-5.37‡	.30			-7.07		
Dummy = 1 if:								
Union veteran	.23	.36‡	.05	.1160				
Confederate veteran					.38	.03	.09	.0071
Married	.71	-.10*	.05	-.0305	.77	-.29‡	.11	-.0767
Illiterate	.04	-.11	.10	-.0345	.12	-.15	.13	-.0380
Has servant	.07	-.15*	.08	-.0484	.04	-.22	.22	-.0577
Takes in boarder	.13	-.11*	.06	-.0361	.11	-.22	.14	-.0568
Owns no property	.40	.14‡	.05	.0461	.39	.28‡	.10	.0723
Lives in South	.03	-.10	.12	-.0308	.78	-.03	.11	-.0085
Lives in urban county	.31	-.07	.05	-.0218	.07	-.05	.16	-.0127
Age	68.22	.07‡	.00	.0233	67.84	.09‡	.01	.0227
Number of dependents	1.24	-.14‡	.02	-.0454	1.63	-.20‡	.04	-.0527

Source: The sample consists of white, noninstitutionalized, native-born men aged 62–79 drawn from the integrated 1910 Census (Ruggles and Sobek 1995).

Note: The symbols *, †, and ‡ indicate that the coefficient is significantly different from zero at at least the 10 percent, 5 percent, and 1 percent levels, respectively. $\partial P/\partial x = \beta(1/n) \sum \phi(x'\beta)$, where ϕ is the standard normal density, and $\partial P/\partial x$ is in probability units.

out the results of figure 3.6. In the southern-born sample Confederate veteran status is not a significant predictor of retirement. In the northern-born sample, Union army veteran status is, suggesting that Union pensions led to higher retirement rates among veterans. There were a few Confederate veterans in the northern-born sample and a few Union veterans in the southern-born sample (classified as nonveterans in table 3.6).¹⁴ Although no strong conclusions can be drawn from such a small fraction of men in either category, when dummies are included for these men, the coefficient on Confederate status in the northern-born sample is insignificant and that on Union veteran in the southern-born sample is significant, suggesting once again that the receipt of a pension was an important determinant of retirement status.

The results presented in table 3.6 can also be used to estimate whether the difference in participation rates between the northern- and the southern-born sample is largely due to differences in observable characteristics or in participation behavior. When retirement rates from the northern-born sample are compared with those from the southern-born sample, then the difference in retirement rates should consist of two components. The first will be the component due to observable differences, such as region of residence or fraction of veterans in the population. The second component should be due to differences in participation behavior. Union army pensions will lead to differences in the participation functions. So might other variables. For example, the northern born who lived in the South may have differed in unobservable retirement determinants. More formally, let R^n be the probability of retirement among the northern born, R^s the probability among the southern born, and X^n and X^s the vectors of northern- and southern-born characteristics, respectively. Then

$$R^n - R^s = [R^n(X^n) - R^n(X^s)] + [R^n(X^s) - R^s(X^s)],$$

where the first term is predicted using the northern-born participation equation for both samples, and the second term is the residual component due to differences in participation behavior between northern and southern born using the southern-born sample. The actual difference in retirement rates is 0.0722. Using the northern-born participation equation for both samples yields a value of 0.0445 for the second term, suggesting that differences in participation behavior and thus pensions account for about 62 percent of the difference in retirement rates between the southern and the northern born. When men aged eighty to ninety-one are included in the sample, regressions on the northern- and southern-born samples suggest that differences in participation behavior account for about 39 percent of the difference in retirement rates between the southern and the northern born. Although the coefficient on an interaction term between age and Union army veteran was statistically significant only at the 20 percent level, it was negative, suggesting that being a Union army veteran had a smaller effect on the participation decision of older relative to younger veterans.

3.5 Implications

I have demonstrated that Union army pensions exerted a large effect on male labor force participation rates. The elasticity of labor force nonparticipation with respect to pension income was 0.73 in 1900 and 0.47 in 1910. I argued that these estimates were unbiased. They can therefore be used to calculate the effect of a secular increase in income on the secular decline in male labor force participation rates, under the assumption that the elasticity of labor force nonparticipation with respect to pension income remained constant after 1900. The retirement rate for men sixty-five years of age or older rose from 35 percent in 1900 to 83 percent in 1990, and per capita fixed reproducible tangible wealth rose by 415 percent.¹⁵ Therefore, using the 1910 pension elasticity of nonparticipation, an increase in income, holding wages constant, could explain 60 percent of the decline in labor force participation rates.

Mounting evidence, however, suggests that the elasticity of labor force nonparticipation with respect to income was lower in the period after 1940 than in my estimates. According to my results, the elasticity of labor force nonparticipation was 0.73 in 1900, when the median age of veterans was fifty-six, and 0.47 in 1910, when the median age was sixty-six. But others find far lower estimates for the period after 1940.

Friedberg's (1996) estimates of the effect of Social Security Old Age and Assistance in 1940 and in 1950, given to those age sixty-five and older, imply an elasticity of labor force nonparticipation with respect to benefits of around 0.25–0.42. The majority of studies using data from the late 1960s on find a similar or smaller effect on labor force participation rates of either assets or Social Security retirement and disability payments (Bound 1989; Bound and Waidmann 1992; Hausman and Wise 1985; Haveman and Wolfe 1984a, 1984b; Krueger and Pischke 1992). Among men in their early sixties Hausman and Wise (1985) estimate an elasticity of 0.23 and Krueger and Pischke (1992) one of 0. Elasticities of labor force nonparticipation with respect to assets in these studies are close to 0. Bound (1989) finds an elasticity of labor force nonparticipation with respect to Social Security disability of 0.16 among men aged forty-five to sixty-four. Studies finding a more sizable effect of Social Security payments on labor force nonparticipation are those of Hurd and Boskin (1984), Leonard (1979), and Parsons (1980). For example, Parsons (1980) calculates an elasticity with respect to Social Security disability of 0.63. The results of selected studies are summarized in table 3.7.

Statistical problems may bias some of the estimates presented in table 3.7 upward. Leonard (1979) and Parsons (1980) compare the labor force participation rates of those whose potential disability benefits would replace a relatively large fraction of their predisability earnings to those whose potential benefits would not. Since replacement rates for disability benefits are decreasing functions of past earnings, it is difficult to determine whether generous replacement

Table 3.7 Elasticities of Labor Force Nonparticipation in Selected Studies

Study	Age of Sample	Year Studied	With Respect to	Elasticity
Costa (this chapter)	Median age is 56	1900	Union army pensions	.73
Parsons (1980)	48–62	1969	Social Security Disability (SSDI)	.63
Leonard (1979)	45–54	1972	SSDI beneficiary status	.35
Haveman and Wolfe (1984a, 1984b)	45–62	1978	SSDI	.21–0.06
Bound (1989)	45–64	1972 and 1978	SSDI	.16
Costa (this chapter)	Median age is 66	1910	Union army pensions	.47
Friedberg (1996)	66–80	1940 and 1950	Old Age Assistance	.25–0.42
Hurd and Boskin (1984)	60–64		Old Age and Survivors Insurance (OASI)	.71
Hausman and Wise (1985)	58–63 in 1969	1969–79	OASI	.23
			Assets	≈0
Krueger and Pischke (1992)	60–68	1976–88	OASI	≈0

Source: The elasticities given for Friedberg (1996), Hurd and Boskin (1984), and Hausman and Wise (1985) were estimated using the information provided by the authors.

rates or low earnings induce the individual to leave the labor force. Haveman and Wolfe (1984a, 1984b) try to avoid the problem of the endogeneity of the replacement rate through the use of an instrumental variables procedure in which they first predict disability benefits as a function of exogenous information and then incorporate these predicted values into the final estimation equation. They estimate an elasticity of 0.06–0.21. Bound (1989) avoids the endogeneity problem by using those who applied for disability benefits but were rejected as a control group for beneficiaries. He also estimates a low elasticity (0.16).

Some of the estimates of the effect of Social Security retirement benefits on retirement may also be biased upward. A potential problem with most studies is that the source of variation across individuals, differing levels of Social Security benefits, arises because of past earnings history. Past earnings are likely to be correlated with present labor supply and thus bias upward estimates of the effect of Social Security.¹⁶ A few studies use other sources of variation. Friedberg (1996) uses state variation in benefits to identify the influence of Old Age Assistance on labor supply. Krueger and Pischke (1992) examine an unexpected legislative change that substantially reduced benefits to individuals born after 1916, leading to a worker who retired at age sixty-five after a career of earning the average wage to receive Social Security benefits that were 13 percent lower than he would have received had he been born in 1916. They

concluded that Social Security wealth had a negative, but insignificant, effect on the probability of retirement. The income elasticity of retirement therefore appears to have fallen from 0.47 in 1910, to 0.25–0.42 in 1940 and 1950, to 0 in recent years.

What therefore needs to be explained is why elasticities estimated from the Union army sample are so much higher than elasticities with respect to transfer income estimated from recent data. One possibility is that there has been a change in the income elasticity of retirement. Another is that elasticities of nonparticipation with respect to assets or Social Security benefits may not be comparable to those calculated with respect to Union army pensions. Assets are not necessarily exogenous, and Social Security payments will have both an income and a substitution effect. With the exception of the unique circumstances examined by Krueger and Pischke (1992), it is plausible to assume that the future stream of Social Security payments was predicted with greater accuracy by men in their working years than was the future stream of Union army payments. Another difference exists because Union army pensions represented a larger fraction of earnings than do Social Security disability payments today, the former constituting 36 percent of the annual earnings of nonfarm laborers in 1900, the latter 36 percent. Furthermore, Union army pensions were the only available retirement program, whereas Social Security disability payments represent 75 percent of all transfer payments (estimated from Center for Human Resource Research 1985).

If noncomparability of elasticities calculated with respect to Union army pension income and those calculated with respect to Social Security payments arises from Union army pensions being at the time the only available transfer program, then the retirement elasticity with respect to Union army pensions can be adjusted to account for this. Thus, if total transfer income were \$12.90 per month in 1900, which is what the average Union army pension paid, a program equivalent in scale and scope to Social Security Disability would have paid \$9.70 per month in 1900 (or 0.75 [\$12.90]). Retirement income includes not only transfer income but also private pensions. Disability payments represent 67 percent of the sum of transfer income and private pensions today, translating into \$8.60 of \$12.90 per month in 1900. Using the 1900 regression produces an elasticity of 0.56 ($= 0.0078 [9.7/0.1343]$) for the equivalent of Social Security disability and one of 0.51 ($= 0.0075 [8.6/0.1259]$) for that of transfer income plus private pensions.¹⁷ Similar calculations using the 1910 regression estimates yield an elasticity of 0.40 ($= 0.0112 [12.4/0.3473]$), also substantial.¹⁸ Furthermore, the average Union army pension in 1900 and 1910 was about as generous as the average Social Security retirement benefit.

The elasticities given in table 3.7 therefore suggest that the responsiveness of retirement to income has fallen since 1900. Workers may now be less responsive to changes in transfer income because they are no longer close to subsistence levels; instead, they reach retirement age with enough to satisfy their consumption needs. Each additional dollar of income will therefore have

less of an effect on their decision. Alternatively, workers' choices may now be constrained by a retirement ethos. Once a sizable fraction of older men are retired, unresponsiveness to pension payments may be the outcome of a "bandwagon" effect or of a desire to conform to societal expectations. By establishing age sixty-five and later age sixty-two as an "official" retirement age, Social Security may have led individuals to want to retire at that age and therefore reduced the effect of income on the work decision. The men remaining in the labor force may be those who are greatly attached to work and who can be induced to leave the labor force only by a very large sum of money.

Workers may also be less responsive to changes in transfer income because leisure has become relatively more attractive and less expensive. In chapter 7 I discuss how in the 1920s and 1930s new technologies such as the car, the phonograph, the radio, and movies increased the variety of recreational activities and lowered their price. These new goods diffused rapidly throughout the population. At the same time public recreational facilities, such as parks, golf courses, and swimming pools, put recreational sports within the reach of more and more individuals. In 1934 the pension advocate Isaac Rubinow noted that wintering in Florida and summering on Michigan lakes was how many of the well-to-do spent their lives after age sixty-five (Rubinow 1934, 243). Even during the 1930s large numbers of the elderly migrated to the Pacific Coast and South Atlantic regions. Their numbers increased after the Second World War. By 1940 private insurance companies offering retirement income plans advertised those plans as offering "freedom from money worries. You can have all the joys of recreation or travel when the time comes at which every man wants them most."¹⁹ Graebner (1980, 215–41) argues that there was a national effort to glorify retirement in the 1950s and describes how retirement was aggressively marketed as a consumable commodity by corporations, labor unions, and insurance companies that were pension plan trustees. Companies established retirement preparation programs, and journals aimed at retired employees were increasingly filled with idyllic depictions of the retired life. What is not known is the extent to which increased retirement induced the marketing and the extent to which the marketing induced retirement.

The marketing of retirement in the 1950s, however, was accompanied by the continued growth of leisure industries. Now mass tourism and mass entertainment, such as films, television, golf, and spectator sports, provide activities for the elderly at a low price. As the desirability of leisure increased, the elasticity of labor force nonparticipation with respect to pension income may have decreased. I take this point up again in chapter 7. Once retirement was seen as a period of personal fulfillment rather than a period before death when men were too ill to work, more men may have found retirement desirable, even if their retirement activities were limited.

A fall in the income elasticity of retirement implies that, while secular increases in income explain a larger share of the rise in retirement rates at the beginning of the century, they explain much less at the century's end. The 1910

estimate of the elasticity of nonparticipation implies that 90 percent of the decline in labor force participation rates between 1900 and 1930 could be attributed to secularly rising incomes. Friedberg's (1996) estimates suggest that at least half the decline between 1930 and 1950 can be accounted for by secularly rising incomes. In contrast, rising incomes may explain none of the 1950–80 decline.²⁰ The findings also suggest that, as leisure continues to grow more attractive, changes in transfer policies alone may not be enough to induce large increases in labor force participation rates among the elderly.

3.6 Summary

This chapter investigated the effect of Union army pensions on veterans' retirement rates, finding that the elasticity of nonparticipation with respect to Union army pension income was 0.73 in 1900 and 0.47 in 1910, when veterans were older and their participation decision was less sensitive to changes in income. The findings suggest that secularly rising income explains a substantial part of increased retirement rates, particularly before 1940. Rising incomes cannot, however, account for most of the recent increase in retirement rates. Comparisons with elasticities of nonparticipation with respect to Social Security income imply that the income elasticity of retirement has decreased with time, either because of changing societal expectations or because of increasingly attractive leisure-time opportunities. Not only can most men now afford to retire, but, when they do retire, they can look forward to a variety of low-cost leisure activities.

Notes

1. Imputations for annual incomes are given in Preston and Haines (1991, 212–20, table A.1). The pension represented an even greater proportion of the earnings of older men because of the sharp decline in the age-earnings profile.

2. The pension records, which include the successive reports of the examining surgeons of the Pension Bureau, are currently being linked to the 1850, 1860, 1900, and 1920 censuses and to army service records to reconstruct the life histories of a random sample of men who served in the Union army. The collection of these data is still ongoing; therefore, nonrandom subsamples of the data are used in the statistical analysis.

3. However, because property was one of the primary modes of savings, men who had retired might already have liquidated their property.

4. The statewide unemployment numbers are from Keyssar (1986, 340–41, table A.13).

5. The coefficients on a spline in age are insignificant, and the use of a spline leaves the regression results unchanged. Similarly, the use of a quadratic term in age does not affect the results.

6. The values of $\partial P/\partial x$ were calculated as $\beta(1/n) \sum \phi(x'\beta)$, where $\phi(\)$ is a standard normal density function, β is the probit coefficient, and n is the number of observations.

7. Using symbols,

$$I_i^* = \alpha P_i + Z_i' \beta + u_i,$$

$$P_i = \Pi' Z_i + V_i,$$

where I_i^* is not observed, only the dummy variable, $I_i = 1$ if retired and 0 otherwise, is. P_i is pension amount; Z_i is the vector of exogenous variables (all variables except for pension amount) and is a subset of Z_i , which also contains the instrumental variable indicating whether a recruit fell under the General Law. The instrumental variable is not included in the retirement equation. The normally distributed error terms are represented by u_i and V_i . Rivers and Vuong (1988) present formulas for the standard errors. When the coefficient on the residuals in the second stage is equal to zero, the standard errors are the usual probit standard errors. There was little difference between the corrected and the uncorrected second-stage standard errors.

8. A Hausman test for exogeneity of pension amount is used.

9. Men for whom such information is unavailable do not differ in observable characteristics from men for whom this information is available.

10. The bias could go either way.

11. Among the white male population in 1900, 15 percent of those aged fifty to fifty-nine, 18 percent of those sixty to sixty-nine, and 9 percent of those seventy to eighty-one were collecting a Union army pension. Among the white male population either born in a Union state or, if foreign born, who immigrated prior to the Civil War, 28 percent of those aged fifty to fifty-nine, 33 percent of those sixty to sixty-nine, and 22 percent of those seventy to eighty-one were collecting a Union Army pension (for sources, see Costa 1995b). Retirement rates among nonveterans in 1900 are calculated from

$$R_g = R_v(\text{fraction veterans}) + R_n(\text{fraction nonveterans}),$$

where R_g is the retirement rate of the general population, R_v that of veterans, and R_n that of nonveterans. Assuming that a pension elimination affects only veteran retirement rates, new retirement rates for the veteran general population can be estimated.

12. Glasson (1918a, 1918b) gives the number of Civil War veterans on the pension rolls in 1910. Because of underenumeration of veterans in the 1910 census, the total number of Union army veterans is estimated from a life table and is from Series Y 957-970 in U.S. Bureau of the Census (1975, 1145). Assuming that underenumeration of veterans did not vary among Union and Confederate veterans, the number of Confederate veterans can be calculated from the 1910 public-use sample.

13. Although disability levels for Confederate veterans are unavailable, young men in the South were almost three times as likely to die during the Civil War as were young northern men (Vinovskis 1990).

14. If these men are classified as either Confederate veterans in the northern-born sample or as Union army veterans in the southern-born sample, then the coefficients in table 3.6 do not change.

15. The increase in per capita fixed reproducible tangible wealth was calculated from U.S. Bureau of Economic Analysis (1986, 322-70). Retirement incomes have kept pace with per capita wealth between 1950 and 1990, increasing by 90 percent, but between 1900 and 1990 rose by more than per capita wealth.

16. An additional problem arises with Hurd and Boskin's (1984) study. Their high elasticity may be an artifact of the way they selected the population at risk to retire (see Diamond and Hausman 1984).

17. Of course, savings, wages of family members, and income from part-time work are now greater than in 1900. Therefore, a monthly transfer is likely to have a smaller effect at high than at low levels of retirement.

18. Hurd and Shoven (1982) estimated that Social Security retirement benefits account for 73 percent of nonwage income, excluding Medicare and Medicaid payments. Including Medicare and Medicaid payments, Social Security payments account for 37 percent of nonwage income, and the elasticity becomes 0.24 ($= 0.0103 [6.7/0.286]$).

19. Insurance company advertisement from a 1940 issue of *Newsweek*, reproduced in Haber and Gratton (1994, 64).

20. Secular increases in income were proxied by increases in per capita fixed reproducible wealth.