Do pensions increase the labor supply of older men?

By: Christopher J. Ruhm

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Abstract:

This paper investigates the relationship between pension coverage and the retirement behavior of older men. Pensions are associated with higher work involvement for males in their late fifties and early sixties but with lower rates of job holding for those aged 65-69. Age of entry into pension employment is shown to be positively correlated with future labor supply. When combined with evidence on age-specific changes in labor force participation rates, this pattern casts doubt on the hypothesis that broadened pension coverage explains a substantial portion of the trend towards earlier male retirement.

Keywords: Labor supply; Retirement; Pensions JEL classification: J14; J26

Article:

1. INTRODUCTION

The labor force participation rates of older males have dramatically declined in the United States since the end of the Second World War, falling from 90% to 67% for 55 to 64 year olds and from 47% to 16% for those over 65 between 1948 and 1993 (US Department of Labor, 1989, 1994). In explaining this reduction, economists have recently focused attention on the role of private pensions.[1] Pensions often contain complicated incentives that encourage workers to remain with the pension provider at some ages but to leave those jobs subsequently. These incentives are typically strongest for defined benefit (DB) plans, where pension payments depend on some combination of seniority and earnings, and often include vesting periods, early retirement options, and social security offsets.[2]

The actuarial value of pensions is frequently maximized at fairly young ages.[3] The changes in pension incentives, occurring with age and length of service, can be extremely large and there is strong evidence that workers respond to these incentives.[4] Furthermore, the number of full-time

1. *Peracchi and Welch (1994) provide an in-depth discussion of the labor supply trends of older men and women. Earlier research on retirement emphasized mandatory retirement provisions, health limitations, changes in wealth and earnings, and social security system incentives. For literature reviews, see Lazear (1986), Quinn et al. (1990), or Ruhm (1994).*

2. *Sixty-eight per cent of pension-covered workers were enrolled in primary DB plans in 1987, a reduction from 87% in 1975 (Beller and Lawrence, 1992). Some of these persons received supplemental coverage from defined contribution (DC) plans.*

3. *Kotlikoff and Smith (1983) found that half of civilian employees could retire at 62 and 15% by age 55, with no actuarial reduction in pension benefits. Similarly, in Mitchell and Fields' (1984) survey of 10 pension plans, lifetime benefits were maximized at 62 or earlier in 9 of the plans and by age 60 in 5. Mitchell (1992) shows that the early retirement incentives became still more prevalent during the decade of the 1980s.*

4. *For instance, in their case study of one large firm, Kotlikoff and Wise (1989) calculate the accrual to pension wealth for an additional year of service at $72,527 for 54 year old managers with 25 years seniority but negative $14,936 for 65 year olds with 30 years on the job. They estimate that workers were 30 percentage points more likely to depart the firm prior to age 60 (44% vs. 14%) than if the company had offered an age-neutral pension plan.*

private wage and salary workers covered by pension plans more than doubled between 1950 and 1975, rising from 26% to 52% (Turner and Beller, 1989). Although coverage rates have been static in recent years, the number of retirees with pensions continues to increase. Many economists have concluded that this combination of pension incentives and expanding coverage rates is responsible for a significant share of the secular decline in the labor force participation of older American men.

As suggested by its title, this paper questions this conclusion and argues that the effects of pensions are far more complicated than is commonly understood. Although pensions hasten the retirements of some workers, they delay the labor force withdrawal of others and so the net impact on retirement ages is not clear.

Pensions have ambiguous impacts for at least three reasons. First, departures from covered jobs may be followed by subsequent full- or part-time work, rather than immediate retirement, thus moderating the effects of the pension incentives.[5] Second, the incentives to leave the firm prior to normal retirement ages are likely to be limited to persons entering the pension plan at relatively young ages, since vesting provisions and length-of-service requirements often discourage early departures for persons beginning pension-covered employment later in life.[6] Third, even for persons entering covered employment at relatively young ages, defined benefit plans are likely to discourage mobility prior to the age at which actuarial benefits are maximized (Allen et al., 1993; Gustman and Steinmeier, 1990), reducing the probability of early, as well as late, retirement.[7]

The empirical analysis below confirms that pensions have more ambiguous effects on retirement behavior than has generally been realized. In particular, pensions are associated with increased labor supply for men in their late fifties and early sixties but with reductions in employment and work hours after age 64. Delayed entry into pension jobs correlates with higher labor supply among mature males and the pattern of age-specific changes in labor force participation rates casts doubt on the hypothesis that broadened pension coverage accounts for a major portion of the trend towards earlier male retirement.

2. PENSION COVERAGE AND RETIREMENT AGES

This section examines the relationship expected between pension coverage and retirement ages.[8] Following Mitchell and Fields (1984), V is the discounted value of lifetime compensation, L the number of periods of retirement, and T and R the ages of death and retirement, respectively. Specifying Y and P as the discounted values of lifetime earnings and pensions benefits, workers choose retirement ages to maximize the utility function $U(V, L)$, subject to the time constraint $L = T — R$ and the budget constraint $V = Y + P$.

5. *Continued work after the end of pension employment deserves greater attention in light of recent research (e.g. Gustman and Steinmeir, 1984; Honig and Reimers, 1987; Quinn et al., 1990; Ruhm, 1990, 1991; Blau, 1994) indicating the prevalence of postcareer employment and partial retirement.*

6. *The importance of age of entry into the pension plan is discussed by Kotlikoff and Wise (1989).*

7. *Stock and Wise (1990) simulate the effects, in one large firm, of switching from a DB to a DC plan. If the employer contributes to the latter plan at a rate such that the annuity value for a 60 year old worker with 30 years of service is the same as for the DB plan, annual retirement probabilities are calculated to rise prior to age 59, fall between 59 and 64 and increase at ages 65 and 66. Because of the high propensities to leave the firm at relatively young ages, however, cumulative retirement probabilities are raised at all ages studied. Simulations for a second firm by Lumsdaine et al. (1994) similarly indicate that switching from a DB to a DC plan would increase retirements through to age 58 and then lower them subsequently.*

8. *The results below do not depend upon a specific definition of retirement.*

Retirement therefore occurs at

$$R = \underset{t}{\operatorname{argmax}}\ U[V(t), L(t)]\ . \tag{1}$$

First-order conditions imply that $U_L = U_v \cdot V_t$, at R, where subscripts indicate partial derivatives and $V_t = \gamma(t) + P_t$ for $\gamma(t)$ and P_t the earnings and marginal pension accrual at t.[9]
It is useful to define the function:

$$D(_{t,\gamma}) = U_v \cdot [\gamma(t) + P_t] — U_L\ , \tag{2}$$

where γ, the age of entry into the pension job, is set to T (the age of death) for persons never holding pension employment. Individuals work at t if $D(\cdot)$ is positive and retire when $D(\cdot) = 0$, provided that $D_t < 0$. Optimal retirement ages for each γ, denoted by R_γ, occur where $D(t,\gamma) = 0$.

The effect of pensions on retirement ages can be examined by comparing persons obtaining covered positions in the first period of work with those employed in non-covered jobs their entire working lives. The latter group are paid their marginal product, $\gamma_0(t)$, and receive no pension. If total compensation is the same in pension-covered positions as in non-covered employment, $y(t) + P_t = \gamma_0(t)$ for all t, which implies that $D(\cdot)$ will be independent of γ and pensions will have no effect on retirement behavior. It is only when pay in pension jobs differs from that in corresponding non-covered positions that labor supply may be affected.

For instance, assume that the compensation of covered workers, at time t, exceeds that of non-covered individuals by a positive differential $\pi(t)$. Persons never holding pension employment retire at age Ω where

$$D(\Omega, T) = U_v \cdot [\gamma_0(\Omega)] — U_L = 0. \tag{3}$$

Persons entering the pension sector at time 0 retire at R, such that

$$D(R_0, 0) = U_v \cdot [\gamma_0(R_0) + \pi(R_0)] - U_L = 0\ . \tag{4}$$

In this situation, pensions may either hasten or delay retirement. Covered workers receive higher lifetime compensation, creating an income effect that lowers U_v (if $U_{vv} < 0$) and speeds withdrawal from the labor force. Compensation is also greater by $\pi(\Omega)$ at Ω, however, generating an offsetting substitution effect. The ambiguous impact of pensions occurs even if the actuarial value of pension wealth is declining over time. Although covered workers exit the labor force earlier than if pensions were actuarially neutral *and earnings remained unchanged*, with competitive labor markets, the wage and pension profiles are inextricably linked.

There are a number of reasons to expect higher compensation in pension-covered employment. Compensation premia and pension coverage may go together in unionized employment. Pay differentials could also exist with competitive labor and product markets. For instance, pensions might be provided, as a means of preventing shirking, in jobs where output is sensitive to effort. The intuition is that effort will only be maximized if sufficiently large costs are imposed for detected shirking. As retirement approaches, the penalty must take the form of reduced pension benefits,

9. *This implies that retirement occurs when the marginal utility of earnings is exactly offset by the disutility of lost leisure.*

since the number of periods over which earnings can be decreased is small. Pension jobs will also be rationed because new job entrants will not have built up sufficient deferred payments and so will have incentives to shirk if they can immediately re-enter a new pension job.[10]

If jobs offering pensions pay more than those that do not, persons will repeatedly attempt to obtain covered employment and there will be heterogeneity in the age of entry into the pension sector.[11] In this situation, late entrants are likely to retire after their counterparts starting in pension employment at younger ages. This occurs because the substitution effect is equally strong for the two groups, whereas early entrants have higher lifetime incomes and a larger (negative) income effect.[12]

To summarize, pension coverage may be associated with either early or late retirement, depending on the strength of the offsetting income and substitution effects. Among workers ever employed in covered jobs, the ages of entry into pension employment and of retirement should be positively correlated. A key condition required to generate this result is that total compensation is higher in covered than non-covered positions.

3. DATA

Data for this study were obtained from the Social Security Administration Retirement History Longitudinal Survey (RHS). The RHS contains a representative national sample (for the United States) of males aged 58-63 in 1969, the initial survey year, with follow-up interviews at two-year intervals through 1979. The sample analyzed includes non-self-employed men whose longest job was in the private sector.[13] Data were pooled across RHS survey years to increase the sample sizes of respondents of given ages.[14]

Fairly detailed information is available on labor force histories during the survey period (1969-1979), as is less specific data for the presurvey years. Importantly, it is possible to ascertain whether the individual was covered by a pension plan in up to three jobs: the job (if any) held at the time of the 1969 interview, the employment held immediately previously, and the longest job, if different from the survey date and previous employment.

Men working at the 1969 survey were defined as having pensions if covered in any of the three jobs for which information is available and to lack coverage if none of the three positions provided a pension. The pension status of men not employed in 1969 was determined according to whether or not they were covered in their most recent or longest job. Coverage is calculated to begin at the start of the first job identified as offering a pension.

10. *See Ruhm (1993) for a theoretical model of partial pension coverage with these characteristics.*

11. *Several types of evidence suggest that pension jobs offer pay premiums. Gustman and Steinmeier (1990) indicate that turnover rates are lower in jobs providing pensions because total compensation is substantially higher in these positions. Similarly, Ellwood (1985) shows that pension coverage is disproportionately obtained by highly paid and highly educated workers and is relatively rarely supplied to disadvantaged labor market groups. He also provides evidence of the dispersion of entry ages into pension employment.*

12. *Since $D(R_0,0) = 0$, $D(R_0, \gamma) > 0$, for $\gamma > 0$ (and $U_{vv} < 0$), which implies that retirement ages rise with the age of entry into the pension sector.*

13. *Women were excluded because information on unmarried females is extremely limited in the RHS.*

14. *For example, data on 63 year olds was obtained from the 1969, 1971, and 1973 interviews. Cohort differences were tested for in preliminary analysis but are not accounted for in the estimates presented below.*

The RHS contains a number of shortcomings. First, the data are quite old and so may not be fully representative of the cohort now approaching retirement age. Second, details on the type of pension coverage are sketchy. In particular, it is not possible to determine whether respondents were covered by DB or DC plans. This limitation is made less severe, however, by the predominance of defined benefit pensions during the survey period. Third, some workers may have obtained coverage in jobs for which the RHS obtains no information or have been employed with firms providing pension plans for which they ultimately did not qualify for benefits.

Despite these deficiencies, the RHS is a useful data set to analyze because it contains a larger representative sample of retirement age men and better retrospective information on presurvey employment than most other longitudinal sources.[15] Furthermore, the RHS has been widely employed to study retirement behavior and so the results obtained are directly comparable with much of the previous research on pensions.

Economists have variously defined retirement as occurring when individuals classify themselves as retired, receive public or private pension benefits, experience a discontinuous reduction in earnings or hours worked, or exit the labor force. The last criteria is implicitly used when considering the relationship between pension coverage and labor force participation rates and ignores any impact of pensions on positive work hours. To avoid depending on a single definition, four alternative measures of labor supply are considered below. The first three are dummy variables indicating whether the individual does not work, works full-time (more than 35 hours per week), or is employed in a 'career' job (defined as a position that has lasted at least 10 years). The fourth is a continuous indicator of weekly hours of work.

4. EMPIRICAL ANALYSIS OF PENIONS AND LABOR SUPPLY

Labor supply declines monotonically with age for RHS respondents. Particularly sharp reductions are observed at 62 and 65, reflecting a combination of Social Security system provisions, mandatory retirement requirements, and pension plan incentives (see Table 1).[16] Sample non- employment probabilities are somewhat higher than corresponding national non-participation rates.[17] This is partially explained by the exclusion of unemployed workers from the latter measure.

Figs. 1-4 summarize information on the relationship between pension status and the alternative measures of labor supply for 58-68 year old men. The patterns are fairly consistent and quite striking. Between the ages of 58 and 61, pension-covered males are substantially more involved in the labor force than their counterparts. Differences of 7-14 and 9-17 percentage points are observed in the probabilities of working and of holding full-time employment, respectively. Men with pensions are also 22-27 percentage points more likely to remain in career positions and average 4-5 more hours per week on the job.

By contrast, the employment probabilities and average work hours of 62-64 year old men are virtually unrelated to pension coverage. Small differentials are still observed, however, when considering the probability of working full-time (3-6 percentage points) and particularly of continuing in the career job (6-13 percentage points).

15. *Although the new Health and Retirement Survey will provide more current information in the near future, its sample members are presently too young for the type of analysis undertaken below.*

16. *The legal minimum mandatory retirement age was raised from 65 to 70 by the 1978 amendments to the Age Discrimination in Employment Act and has now been eliminated for most workers.*

17. *For example, 17% and 18% of 58 and 59 year old RHS respondents were not working at the survey date in 1969. The corresponding national non-participation rates were 11% and 13%, in that year.*

Table 1
Labor supply as a function of age

Age at survey date	Probability of working 0 hours	≥35 hours	In career job	Average hours worked	N
58	0.170	0.797	0.575	36.3	1003
59	0.191	0.782	0.544	34.9	934
60	0.209	0.750	0.511	33.7	1817
61	0.256	0.698	0.477	31.8	1747
62	0.357	0.597	0.410	27.1	2541
63	0.401	0.536	0.346	25.0	2371
64	0.505	0.422	0.288	19.8	2364
65	0.684	0.222	0.140	11.6	2180
66	0.737	0.153	0.098	8.7	2203
67	0.755	0.131	0.082	7.8	2004
68	0.781	0.099	0.055	6.3	2021
	(1)	(2)	(3)	(4)	(5)

Note: Career jobs are defined as positions that have lasted 10 or more years.

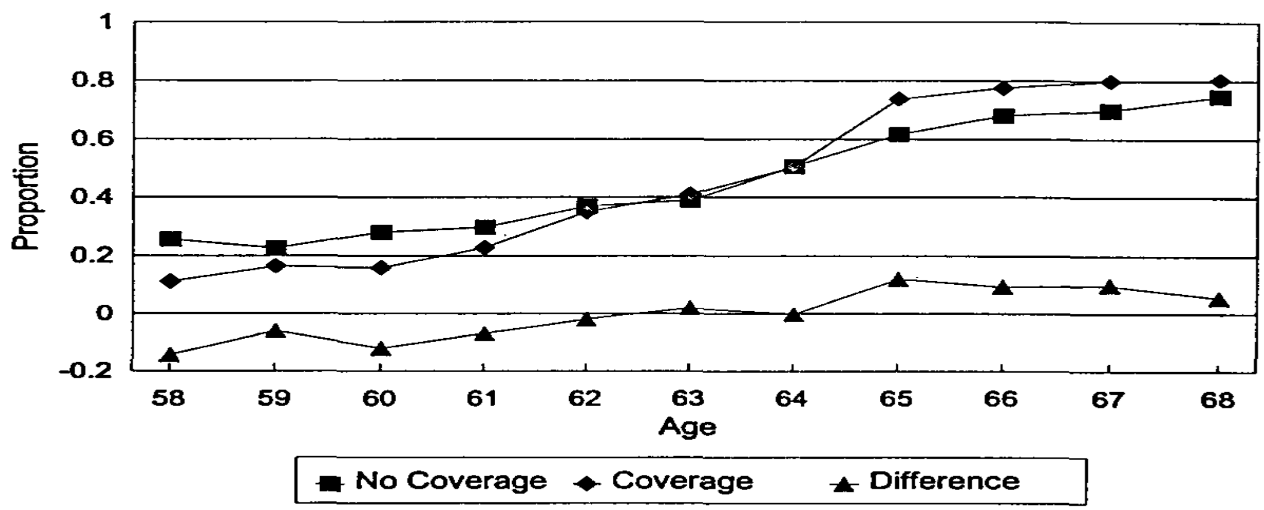

Fig. 1. Proportion of men not working, as a function of age and pension status.

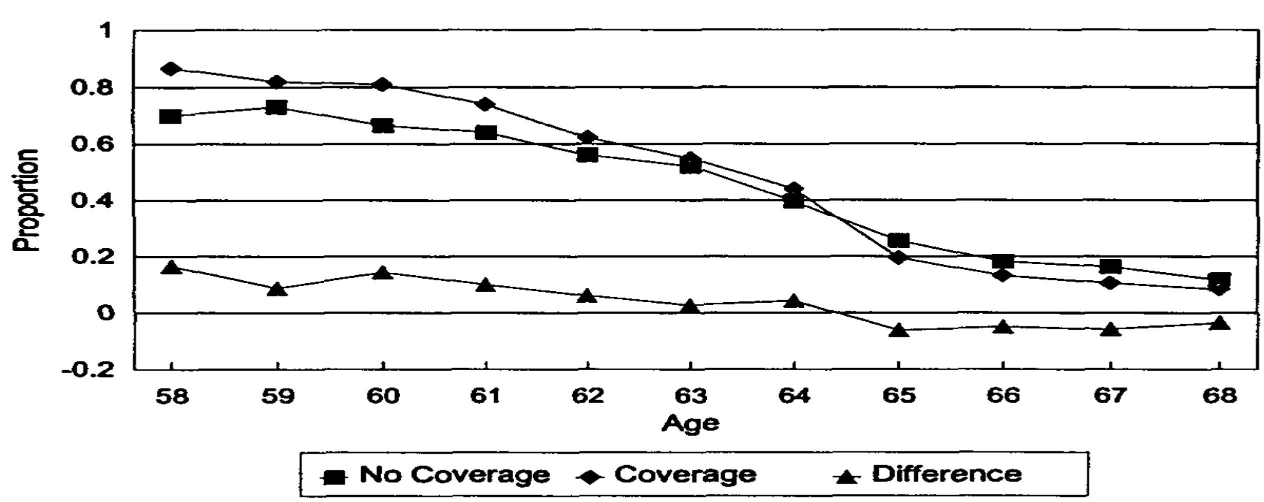

Fig. 2. Proportion working full-time, as a function of age and pension status.

Beyond age 64, pensions are associated with significantly reduced labor supply. For example, whereas 74% of covered 65 year olds did not work, 20% were employed more than 35 hours per week, 12% held career jobs, and weekly employment averaged 10 hours, the corresponding figures for 65 year old men without pensions were 61%, 26%, 17%, and 14 hours per week, respectively. Pensions continue to be correlated with reduced labor supply through age 68, with the differences in non-employment, full-time work, career job-holding, and hours ranging from 6-12, 1-6, and 2-5 percentage points and 2-4 hours per week, respectively, depending on the exact age considered.

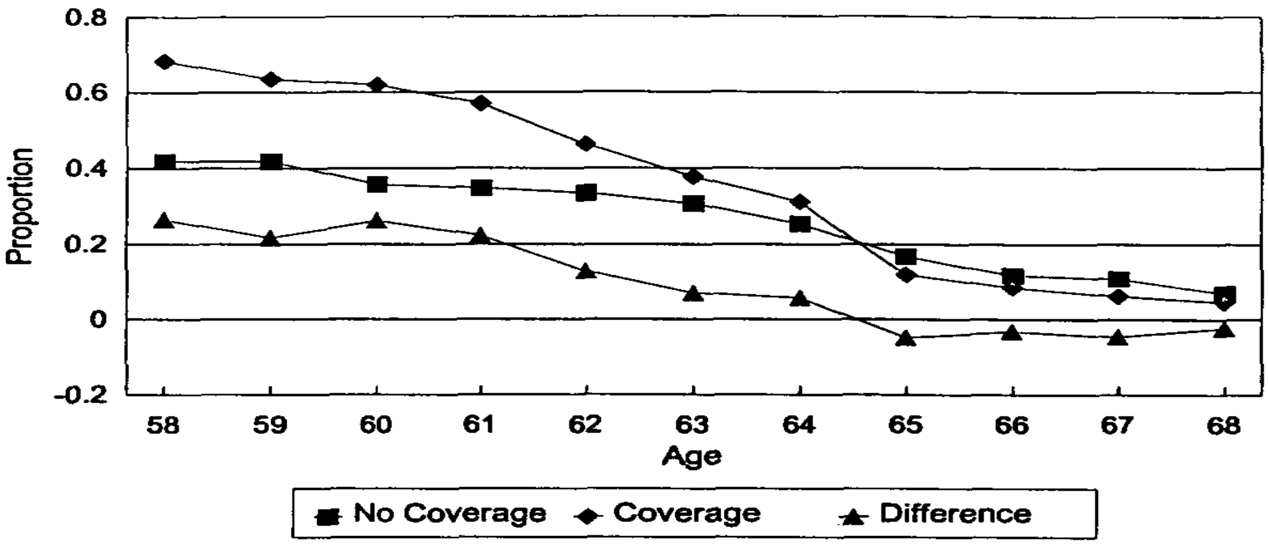

Fig. 3. Proportion working in career job, as a function of age and pension status.

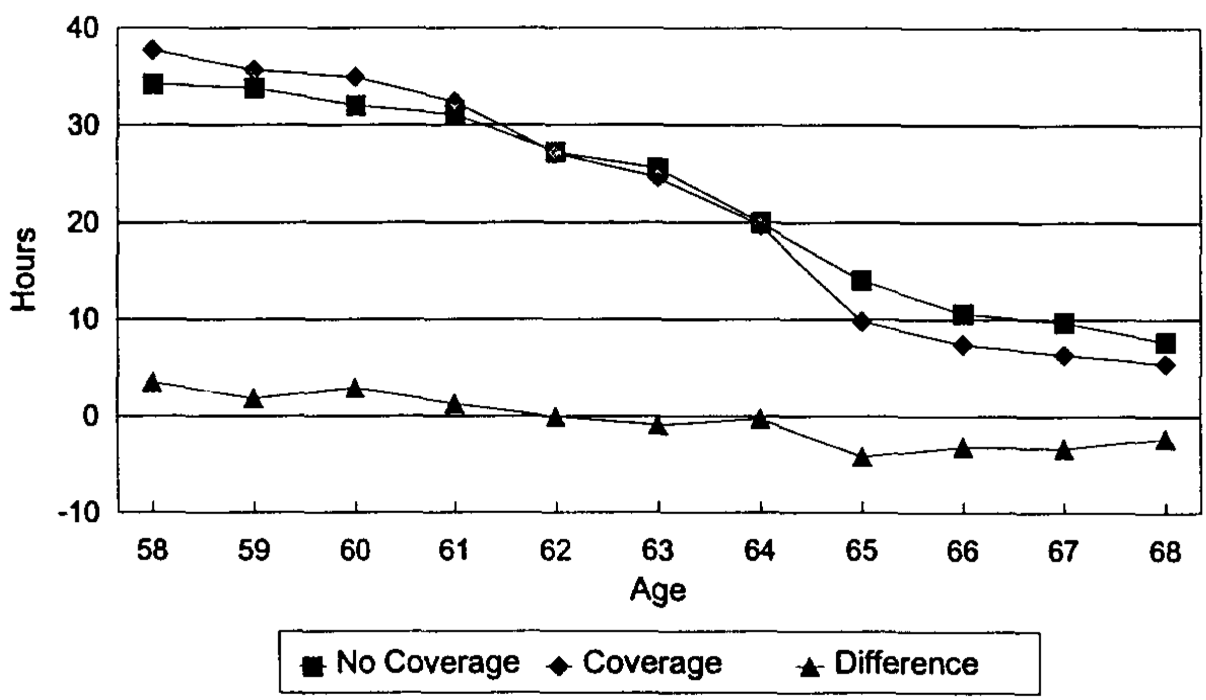

Fig. 4. Average work hours, as a function of age and pension status.

To examine whether these patterns persist after controlling for observable characteristics, probit models of the form

$$Pr(Z_j = 1) = \Phi(X_j\alpha + ßP_j), \quad (5)$$

were (separately) estimated for subsamples of respondents aged 58-68. The subscript refers to the *j*th respondent, *Z* is a dichotomous outcome (i.e. non-employment, full-time work, or career job-holding), $\Phi(\cdot)$ is the cumulative normal distribution function, *X* a vector of covariates, including a quadratic for years of education and dummy variables indicating whether the respondent was non-white, married, a household head, or residing in an SMSA, and *P* is the pension coverage dummy variable.

The probit results were supplemented with an hours-of-work equation, estimated as a tobit model:

$$H_j^* = X_j\alpha + ßP_j, + \in_j, \quad (6)$$

for *H* indicating weekly work hours and *H** a latent (observed) variable with *H* equal to *H** (0) if *H** is greater than (less than or equal to) 0.

After estimating (5) and (6), the average effect of pensions on the various measures of labor supply was calculated. The first three columns of Table 2 present the predicted impact on non-employment, full-time work, and career job-holding probabilities, while column 4 displays the corresponding differential in weakly work hours.[18] For example, the first row of the table indicates that 58 year old men with pensions were 11, 12, and 23 percentage points more likely than their counterparts to work, hold full-time jobs, and to continue in career positions, respectively, and that they worked an extra 2.3 hours per week.

The results in Table 2 are similar to those obtained without controls for individuals characteristics. Specifically, pensions correlate with increased labor supply prior to age 62, with lower work propensities after 65, and with small differences for 62-64 year olds. Thus, pensions coverage was associated with a 7 percentage point reduction in average non-employment probabilities between ages 58 and 61, a 4 point increase for 62-64 year olds, and an 11 percentage point rise for those aged 65-68. The corresponding changes in full-time work propensities averaged 9, 0, and —7 percentage points, for the three age groups, while the mean differences in career job-holding and weekly work hours were 21, 5, and —5 points and 1.5, —2.3, and —4.1 hours, respectively.

These findings show that estimated pension effects vary with age and are sensitive to the measure of labor supply used. They further suggest that it may be problematic to use only a single criteria (e.g. participation rates) when evaluating the impact of pensions. For example, averaged over the 58-68 year age range, pensions are predicted to decrease employment probabilities and average hours worked but also to increase slightly the frequency of full-time employment and to raise considerably the likelihood of remaining in a career job.

18. *For the probit models, these are calculated as the expected value of 4,(x/91 +13)— 0(X16), averaged over all respondents, where cr and / indicate probit coefficients. For the tobit models, the differences are estimated by 0(.)fi , where is the expected probability of censoring at zero hours, evaluated at the regressor means.*

Table 2
Predicted effects of pension coverage on labor supply

Age at survey date	Probability of working			
	0 hours	≥35 hours	In career job	Hours worked
58	−0.107	0.120	0.232	2.27
	(4.37)	(4.61)	(7.01)	(1.66)
59	−0.034	0.059	0.194	0.79
	(1.28)	(2.10)	(5.72)	(0.55)
60	−0.094	0.110	0.232	2.44
	(4.71)	(5.18)	(9.36)	(2.27)
61	−0.047	0.067	0.187	0.45
	(2.12)	(2.90)	(7.49)	(0.38)
62	0.021	0.012	0.093	−2.06
	(1.15)	(0.61)	(4.51)	(2.02)
63	0.058	−0.014	0.037	−3.22
	(2.81)	(0.67)	(1.81)	(3.06)
64	0.030	0.013	0.027	−1.74
	(1.40)	(0.62)	(1.40)	(1.70)
65	0.141	−0.085	−0.059	−5.67
	(6.79)	(4.60)	(3.75)	(6.57)
66	0.114	−0.064	−0.045	−4.10
	(5.66)	(3.90)	(3.29)	(5.54)
67	0.116	−0.077	−0.058	−4.17
	(5.66)	(4.80)	(4.40)	(5.69)
68	0.065	−0.046	−0.028	−2.37
	(3.32)	(3.20)	(2.59)	(3.69)
	(1)	(2)	(3)	(4)

Notes: Coefficients show the average effect of pension coverage on the dependent variable. Columns (1)–(3) are obtained by estimating probit models of the form $\Pr(Z_j = 1) = \Phi(X_j\alpha + \beta P_j)$, with predicted pension effects calculated as the difference in the average values of $\Phi(X_j\hat{\alpha} + \hat{\beta})$ and $\Phi(X_j\hat{\alpha})$, where $\hat{\alpha}$ and $\hat{\beta}$ are probit coefficients. Estimates in column (4) are from a corresponding Tobit model. The dependent variable is hours worked and predicted effects are calculated as $\hat{\beta}\Phi(X'\hat{\alpha} + P'\hat{\beta})$, where X' and P' are sample means and $\Phi(\cdot)$ shows the estimated probability that an observation is censored. Estimates are separately obtained for each age group and control for marital status, years of education, race (white vs. non-white), household status (head vs. non-head), whether the individual resides in an SMSA, and pension coverage. Absolute value of t statistics are shown in parentheses.

5. DELAYED ENTRY INTO PENSION EMPLOYMENT

We next examine whether the labor supply of mature men is positively correlated with the age of entry into the pension sector. As discussed, this is expected if compensation is higher in covered than non-covered employment, since the income effect is weaker for individuals starting pension jobs at later ages, while the substitution effect remains substantial.[19] By contrast, labor supply may be independent of entry ages if total compensation is the same in the two types of jobs.[20]

To test whether later entry into pension employment is associated with elevated labor supply in late middle age, the probit and tobit models specified by Eqs. (5) and (6) were re-estimated, with the addition of a covariate indicating the commencement age of the covered job. Results of these regressions, summarized in Table 3, support the hypothesized relationship, at least for ages 62 and older. A one-year delay in entering the pension job reduces the probability of non-employment by 0.1 to 0.7 percentage points, depending on the exact age considered (see column 1). The corresponding increase in work hours ranges from —0.06 to 0.26 hours per week and the probability of working full-time rises by as much as 0.6 percentage points (columns 2 and 3).[21]

Postponed entry into pension employment has particularly pronounced effects for 62-66 year olds, with a smaller impact continuing through age 68. For example, a one standard deviation (12.5 year) delay in starting the pension job raises predicted employment rates by 5.2 percentage points at age 62, 7.8 percentage points at 65, and 3.0 points at 68. Corresponding increases in the probabilities of working full-time are 4.6, 6.5, and 2.4 percentage points, respectively. Persons entering covered jobs after their early (middle) forties are more likely to be working or employed full-time at age 62 (64) than their counterparts never obtaining pensions.

6. AGE-SPECIFIC CHANGES IN LABOR FORCE PARTICIPATION

If broadening pension coverage plays a key role in explaining the secular decline in the labor force participation rates of older men, the decreases would probably be most pronounced for men aged 65 and above. This would be expected since pensions are associated with the greatest reductions in labor supply for this group. Smaller declines would then be anticipated for younger workers for whom pensions have more modest negative impacts or are correlated with increased work involvement. Surprisingly, scant attention has been paid to the pattern of age-specific changes in participation rates.[22]

The first three columns of Table 4 indicate levels and changes in participation rates for men between the ages of 58 and 68. This information was obtained from unpublished US Department of Labor data and is presented as three-year averages for periods centered around the dates listed in the table

19. *If the pension plan has service requirements for 'early' or 'normal' retirement, the substitution effect may be stronger for late than for early entrants, which will reinforce the positive relationship between entry ages into the pension sector and labor supply of mature adults.*

20. *Compensation could be equal even when the average earnings of covered workers exceed those of persons without pensions if pensions are provided as a compensating differential in jobs with particularly unpleasant working conditions, or if more productive workers choose to take a larger proportion of their pay in the form of fringe benefits.*

21. *Since career jobs are defined a positions lasting at least 10 years, persons recently beginning pension employment will automatically be precluded from holding career jobs. When the sample is restricted to persons starting survey date employment more than 10 years previously, a positive relationship is observed between pension entry ages and probabilities of working in career jobs for most age groups.*

22. The best previous analysis is by Quinn et al. (1990).

Table 3
Predicted effects of delaying start of pension-covered job for one year

Age at survey date	Probability of working: 0 hours	Probability of working: ≥35 hours	Hours worked
58	−0.0015	0.0025	0.102
	(1.29)	(2.00)	(1.36)
59	0.0013	−9.3E-4	−0.057
	(0.96)	(0.68)	(0.74)
60	−0.0010	0.0016	0.094
	(1.14)	(1.61)	(1.69)
61	−0.0012	8.6E-4	0.057
	(1.07)	(0.74)	(0.90)
62	−0.0039	0.0035	0.207
	(3.84)	(3.36)	(3.94)
63	−0.0053	0.0049	0.246
	(4.91)	(4.52)	(4.56)
64	−0.0045	0.0042	0.220
	(4.06)	(3.77)	(4.17)
65	−0.0066	0.0057	0.260
	(6.69)	(6.45)	(6.75)
66	−0.0044	0.0041	0.173
	(4.68)	(5.34)	(5.17)
67	−0.0026	9.0E-4	0.082
	(2.84)	(1.27)	(5.59)
68	−0.0024	0.0020	0.084
	(2.51)	(3.09)	(2.74)
	(1)	(2)	(3)

Notes: See notes to Table 2. For the probit estimates, predicted effects are calculated as the average change in the outcome variable, for persons with pension coverage, from delaying the start of coverage from the sample mean age it begins until one year later. For the Tobit model, they are estimated as $\hat{\gamma}\Phi(\cdot)$, for $\hat{\gamma}$ the coefficient on the age of first pension coverage, and $\Phi(\cdot)$ the predicted probability that an observation is censored, with the regressors evaluated at their sample means.

Table 4
Changes in labor force participation rates of older men

Age	LFPR in 1969	Change in LFPR (1969 to specified year)				Predicted effect of pensions on LFPR
		Absolute change		Relative change		
		1979	1988	1979	1988	
58	0.886	−0.080	−0.104	0.050	0.074	0.107
59	0.868	−0.082	−0.120	0.049	0.058	0.034
60	0.846	−0.106	−0.148	0.025	0.030	0.094
61	0.815	−0.111	−0.160	0.020	0.018	0.047
62	0.754	−0.161	−0.235	−0.030	−0.057	−0.021
63	0.699	−0.172	−0.249	−0.041	−0.071	−0.058
64	0.665	−0.177	−0.271	−0.046	−0.093	−0.030
65	0.515	−0.154	−0.201	−0.023	−0.023	−0.141
66	0.456	−0.148	−0.169	−0.018	0.009	−0.114
67	0.406	−0.134	−0.149	−0.003	0.029	−0.116
68	0.377	−0.113	−0.151	0.018	0.027	−0.065
	(1)	(2)	(3)	(4)	(5)	(6)

Notes: Age-specific labor force participation rates were obtained from unpublished Department of Labor data and calculated as three-year averages, centered on the year listed. The predicted pension effect is derived from the probit results summarized in column 1 of Table 2.

(e.g. column 1 presents average rates for the years 1968-1970). The use of three-year averages minimizes the (sometimes considerable) effects of year-to-year variations in labor force participation.

The second column shows changes in average participation rates occurring between 1969 and 1979, the period covered by the RHS survey. Column 3 displays corresponding changes taking place between 1969 and 1988. Thus, the labor force participation rate of 58 year old males averaged 89% during the three-year period centered around 1969 and fell by 8 and 10 percentage points (to 81% and 78%), respectively, by the periods centered on 1979 and 1988 (see row 1). The participation rate of 58-68 year old men fell by an average of 13 percentage points between 1969 and 1979 and by 18 points over the 1969-1988 interval.[23]

'Relative changes' in participation rates, shown in columns 4 and 5 of the table, were next calculated by taking the difference between the change for the particular age group and the average reduction for all 58-68 year olds. For instance, the 10 percentage point decrease in the labor force participation rates of 58 year olds, between 1969 and 1988, represents a *growth* of 7 points (-0.104--0.178 = 0.074) relative to the average for all 58-68 year olds. The sixth column displays the predicted effect

23. *These are unweighted averages, which do not account for age-group size differences.*

of pension coverage on the probability that RHS respondents worked positive hours at specific ages. These estimates were obtained by reversing the sign of the non-employment effects displayed in column 1 of Table 2.

The table shows that 62-64 year olds experienced the largest declines in participation (columns 4 and 5). At these ages, however, pensions were correlated with (at most) only modest reductions in labor supply (column 6). Conversely, the pension disincentives were concentrated among 65-68 year olds, for whom the declines in aggregate participation rates were much smaller.

These results suggest that broadened pension coverage is unlikely to explain much of the trend towards earlier male retirement but do not rule out other roles for pensions. Since coverage rates grew rapidly between 1950 and 1975, recent cohorts will have been enrolled in pension plans for longer periods of time than their previous counterparts at equivalent ages. This could lead to a reduction in the age of peak pension wealth and disincentives for continued work. Second, even holding entry ages into the pension sector constant, the characteristics of DB plans appear to have changed over time in ways that encourage earlier retirement.[24]

7. CONCLUSIONS

Previous research suggests that private pensions reduce the labor supply of mature adults and, when combined with higher rates of coverage, may account for an important share of the secular decline in the labor force participation rates of older men. This paper questions these conclusions, arguing that pensions are likely to reduce labor supply at some ages but to increase it at others, resulting in a more complicated and ambiguous aggregate impact than is frequently realized.

Empirically, pension coverage is associated with higher labor supply for RHS respondents in their fifties to early sixties, with negligible or small effects for men aged 62-64, and with large negative impacts among 65-69 year olds. Disparate results are obtained using alternative measures of labor supply, suggesting shortcomings with relying on any single criterion. The RHS data also provide strong evidence that the labor supply of mature males is positively correlated with the commencement age of pension employment. This last result is anticipated if pay premiums are provided in pension jobs and it may be inconsistent with theories that require equality of compensation across the covered and non-covered sectors.

The evidence that pension effects vary with age may explain why earlier analysis that constrains the impact to be constant (e.g. Hanoch and Honig, 1983; Blau, 1994) frequently uncovers a negative relationship between pensions and work propensities.[25] The age variation also raises the possibility of important interactions between private pensions and other forms of retirement wealth, such as Social Security. Recent research (e.g. Samwick, 1993; Lumsdaine et al., 1994; Ruhm, 1995) suggests that these interactions may be important but provides little consensus on the direction or timing of the effects.

The age pattern of changes in labor force participation rates also casts doubt on the hypothesis that broader pension coverage explains a substantial portion of the trend reduction in male retirement

24. *The percentage of the male workforce eligible for full pension benefits prior to 65 rose from 4% in 1970 to 15% in 1985 (Ippolito, 1990). Mitchell and Luzadis (1988) and Luzadis and Mitchell (1991) provide additional evidence that the age of maximum pension wealth has been falling over time.*

25. *Gordon and Blinder's (1980) analysis of age-specific pension effects, using the RHS, obtains results consistent with those reported in this paper.*

ages. Whereas the age-specific decreases in participation are largest for 62-64 year olds, the negative impact of pensions is most pronounced among men between the ages of 65 and 68. These divergent patterns do not rule out a role for pensions, and in particular do not account for secular changes in the age profile of pension accruals; however, a closer correspondence would be expected if increases in private pension coverage are of primary importance.

Previous research on pensions has typically focused on the microeconomic incentives implicit in pension plans. Plan incentives are only one component of total compensation, however, and the latter is determined in the larger labor market. Thus, pensions could contain incentives that discourage continued employment of persons enrolled in the plans, while having no aggregate effect on, or even increasing, average retirement ages.

The age of entry into pension employment has a substantial impact on subsequent labor supply. This highlights the importance of better understanding the process by which workers obtain jobs offering pensions. Subsequent empirical work also needs to investigate the dynamics of entry into, and exit out of, pension employment, the role of worker heterogeneity in determining pension coverage, and the relationship between the broad labor supply effects focused upon in this paper and the results obtained from detailed analyses of individual plans.

Future research should also attempt to ascertain the applicability of this and previous studies for the cohort of workers who are currently approaching retirement. Significant differences could result from the greater use of defined contribution pension plans, an economic environment that is increasingly characterized by stagnant earnings and high rates of labor displacement, and the rising labor force involvement of women and other groups receiving limited attention in prior work.

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