

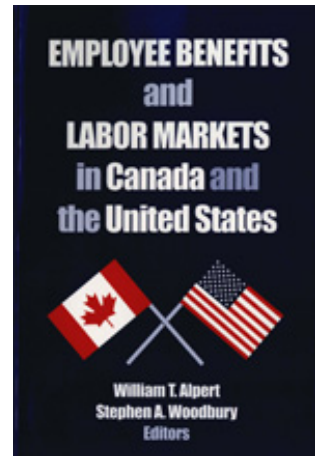


Upjohn Institute Press

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Chapter 13 (pp. 475-495) in:
Employee Benefits and Labor Markets in Canada and the United States
William T. Alpert, and Stephen A. Woodbury, eds.
Kalamazoo, MI: W.E. Upjohn Institute for Employment Research, 2000
DOI: 10.17848/9780880995511.ch13

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13 Did the Decline in Marginal Tax Rates during the 1980s Reduce Pension Coverage?

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After years of constancy or increase, private pension coverage rates declined during the 1980s. Because private pensions are an important source of retirement income, the decline in their coverage raises concern over the adequacy of future retirement income.¹ Between 1979 and 1988, the percentage of full-time male private sector employees participating in a pension plan fell from 55 to 51 percent, where it remained in 1993 (Beller and Lawrence 1992; U.S. Department of Labor 1994).

The coverage decline was particularly large for young males. Coverage for full-time male private sector employees aged 25 to 29 declined by nearly a quarter, from 53 percent in 1979 to 41 percent by 1993 (U.S. Department of Labor 1994).

Because the decline in pension coverage rates has been particularly great for young males, researchers have looked for determinants of coverage that changed more for that group than for older males. Bloom and Freeman (1992) and Even and Macpherson (1994) used this approach to argue that the decline in coverage for young males is explained primarily by disproportionately large declines in their unionization and contemporaneous real income.² The fact that marginal tax rates declined most for high-income workers while coverage declined most for younger low-income workers led these researchers to ignore the potentially important effect of contemporaneous declines in marginal tax rates.³

The tax code encourages both pension coverage and generosity by exempting pension savings from the double taxation associated with other savings vehicles (Turner 1981; Woodbury and Bettinger 1991; Woodbury and Hamermesh 1992; Gentry and Peress 1995).⁴ Workers' earnings are taxed, for example, before they contribute to savings accounts. The returns on savings are again taxed when they are realized (Munnell 1982). In contrast, pension contributions made by firms on behalf of workers are not taxed. Pension benefits are only taxed when they are disbursed, thereby avoiding double taxation.

Preferential tax treatment causes the tax advantage of pensions to increase with marginal income tax rates. Workers with high marginal tax rates tend to seek jobs with pensions, suggesting that pension coverage was reduced by declines in marginal tax rates during the 1980s and early 1990s. Woodbury and Bettinger (1991) found that decreases in marginal tax rates did reduce pension coverage for a sample pooled by gender and age. Woodbury and Huang (1991) and Feldstein (1997) found that the large cuts in marginal income tax rates encouraged high-income workers to take less compensation as fringe benefits and more as income.⁵

Because the tax expenditure for pensions is the largest tax expenditure for individuals in the federal budget, it is important to understand the effects of that expenditure on pension coverage.⁶ The tax expenditure for pensions could lead to increased national savings through increased pension coverage, but of itself reduces government revenue, reducing savings.

We examine whether the decline in marginal tax rates during the 1980s caused a decline in pension coverage rates. We empirically test the assertion that tax changes cannot explain the disproportionate decline in coverage for young males because "the 1980s fall in coverage was smallest among high-income (older) workers, for whom marginal tax rates declined the most" (Bloom and Freeman 1992, p. 543). We explore causal links between declines in tax rates and observed declines in pension coverage using cross-sectional data over a 15-year period, from the 1979, 1988, and 1993 Current Population Surveys.⁷

Our estimates suggest that, on average, a 1 percentage point increase in the marginal tax rate leads to a 0.4 percentage point increase in private pension coverage. Declining tax rates explain almost 20 percent of the total decline in coverage for young males

between 1979 and 1988. Our model predicts that changes in exogenous variables lowered coverage rates 7.3 percentage points for young males between 1979 and 1988. Declining tax rates account for 1.4 percentage points of the total predicted decline in coverage. Declining unionization accounts for only 0.9 percentage points of the predicted decline, whereas declining earnings account for 5.7 percentage points of the predicted decline.

We test the robustness of our results by reestimating the coverage equation for a sample of female private sector workers. In contrast to the males, who experienced declining coverage rates in the 1980s and early 1990s, females experienced slightly rising coverage rates during this period. Our results from the female sample corroborate our earlier conclusions that on average a 1 percentage point increase in the marginal tax rate leads roughly to a 0.4 percentage point increase in pension coverage.

In the next section, we discuss the empirical specification. We then discuss the data and variables, with special attention to the problem of endogeneity of tax rates, and we estimate our model for males and then for females. In the final section, we offer concluding comments.

EMPIRICAL SPECIFICATION

Observed compensation packages consisting of wages and fringe benefits result from decisions made by firms and workers, subject to market and regulatory constraints. Woodbury (1983) and Woodbury and Huang (1991) estimated a demand equation for pensions as a share of total compensation. They modeled the determinants of pension provision and other nonwage compensation by assuming that employers offer a menu of compensation packages, given their costs. Utility-maximizing workers then choose their preferred compensation packages from the menu of available alternatives. As suggested by Deaton and Muellbauer (1980), the authors cited above specified a flexible form expenditure function, from which they derive a system of demand equations for wages and pension benefits as a share of total compensation. Since share data are more readily available at the firm level, they variously used the establishment and the two-digit industry as the unit

of observation. Explanatory variables are firm or industry average characteristics.

Other studies, such as Woodbury and Bettinger (1991), Bloom and Freeman (1992), and Even and Macpherson (1994), focused on accounting for changes over time in observed coverage rates. These studies used household data and the individual as the unit of observation. Household data sets, however, do not contain information about the amount that the firm contributes to an individual's retirement pension. The data available are discrete and measure whether an employer offers a pension and, if so, whether the employee participates in the plan. These authors estimated a discrete model of the probability that a worker with given economic and demographic characteristics, employed at a firm with given attributes, is covered by a pension. The coverage equation is interpreted as the probability of a pension coverage outcome and is not interpreted as a behavioral equation. The estimated coefficients in the coverage equation, appropriately transformed, measure the effect of a change in an exogenous variable on the probability that a worker/firm match leads to coverage for the worker.

Like Woodbury and Bettinger, Bloom and Freeman, and Even and Macpherson, we use household data to estimate a pension coverage equation. However, Bloom and Freeman and Even and Macpherson do not include a tax variable and maintain the hypothesis that changing tax rates have no effect on coverage. Woodbury and Bettinger, on the other hand, include a tax variable, but pool by age and gender. None of these authors have tested whether declines in tax rates contributed to the decline in coverage for young males. They also have not tested the hypothesis that declining tax rates put downward pressure on coverage rates for women, during a time period where observed coverage rates for women were rising.

To formalize the model, let Z represent a vector of worker and firm characteristics that affect the probability of coverage.⁹ The equilibrium outcome is represented by an indicator variable, P , which takes a value of 1 if the worker is covered by a pension plan offered by the worker's firm. Let e represent a random variable interpreted as unobserved heterogeneity in the rates at which firms and workers are willing to substitute pension benefits for wages. We can write the probability that worker i employed at firm j has pension coverage as:

$$\text{Prob}(P=1) = F(Zg + e > 0) = F(-Zg) \quad (1)$$

where F is a cumulative distribution function and g is the vector of parameters to be estimated. We assume that e has a normal distribution and estimate a probit model.

Rather than report the estimated probit coefficients, \hat{g} , we report the marginal effects of the continuous variables and the delta effects of the dichotomous variables.⁹ The delta effect is the discrete analog of the marginal effect. We report t -statistics for the marginal and delta effects themselves.¹⁰

DATA, VARIABLES, AND ENDOGENEITY

We use data from the 1979 and 1988 May Current Population Surveys (CPS) and the 1993 April CPS, which include a special survey of workers concerning pension plan coverage and other employer attributes. These data have been matched to the March CPS of the same year, which provides income and other economic and demographic data. The sample is limited to full-time employed, private, wage and salary workers aged 21 to 55 who did not work in agriculture or the railroad industry and had valid responses to questions relevant for this study.

The dependent variable is worker self-reported pension plan participation, which includes participation in both defined-benefit and defined-contribution plans. This is the best definition of pension coverage for our purposes because it represents the worker's intention to use a plan for retirement. Some workers who are in defined-contribution plans, which are like savings accounts, may intend to use those plans for preretirement consumption rather than for retirement and respond that they are not covered by a pension plan.

Most of our explanatory variables are standard in equations estimating pension coverage. They include age, race, firm size, education, marital status, years with employer, and union status (Table 1). We also use nine industry and four occupation dichotomous variables.

In addition, we use two variables not universally included in pension coverage equations—the predicted combined state and federal

marginal income tax rate and predicted yearly earnings at age 55. To avoid bias due to the endogeneity of earnings and thus marginal income tax rates, we use the predicted value of both variables.¹¹

We calculate the predicted marginal income tax rate, reflecting both state and federal income taxes, using current predicted family income (rather than actual family income), marital status, and number of children. These predicted tax rates are not subject to endogeneity bias arising from idiosyncratic variations in labor supply and earnings. To calculate marginal tax rates, we use the income tax codes for each of the 50 states for each of the three years of analysis. Marginal tax rate variability across states provides exogenous variation in tax rates.

Table 1 Variable Definitions

Variable	Definition
Covered	Equals 1 if covered by a pension on the current job
Tax	State plus federal marginal income tax rate based on current predicted family earnings
Age	Age in years
Pearn55	Predicted yearly earnings at age 55 in 1993 dollars, assuming real earnings from the cross-sectional age/earnings profile grow 1 percent annually
African American	Equals 1 if African American
Married	Equals 1 if married with spouse present
Newhire	Equals 1 if worked for current employer for no more than one year
Union	Equals 1 if covered by a collective bargaining agreement
Mult1000	Equals 1 if employer operates at more than one location and employs 1,000 or more workers

Since the family income question in the CPS is retrospective, we use tax rates for the year prior to each CPS.¹²

The average predicted marginal income tax rate, state plus federal, for male workers aged 21–35 in our regression sample fell from 30.8 percent in 1979 to 26.0 percent in 1988 and rose slightly to 26.1 percent in 1993. The average for workers aged 36–55 fell from 35.3 percent in 1979 to 30.4 percent in 1988 and to 29.8 percent in 1993 (Table 2).

In addition to controlling for marginal tax rates, it is important to control for wealth or lifetime income. The argument for including such a measure is based on the normality of consumption during retirement. Individuals who have greater earnings or wealth over their lifetime wish to consume more during retirement and thus have a higher demand for pension coverage. Because income and marginal income tax rates are positively correlated, if income is not adequately controlled for, a finding of a significantly positive effect of marginal income tax rates on coverage could merely indicate that higher income workers have a higher demand for coverage.

Some authors, particularly Bloom and Freeman (1992) and Even and Macpherson (1994), include current earnings as a variable explaining pension coverage. However, current earnings are endogenous and so the coefficient estimate on this variable is biased. The direction of the bias cannot be determined *a priori*. The estimated coefficient is likely to be upward biased if unobserved heterogeneity in ability is positively correlated with both earnings and pension coverage. However, compensating differentials for pension coverage would cause the coefficient on earnings to be downward biased since unobservables that are positively correlated with pension coverage may be negatively correlated with earnings.

Instead of using current earnings, we use the instrumental variables approach suggested by Dorsey (1982) and Woodbury and Bettinger (1991). First, for each data set we estimate an earnings equation, with a standard human capital formulation. Included in the explanatory variables is potential experience, measured as age minus years of education minus 6. Using current job characteristics, we predict earnings for each individual at age 55. To do so, we assume that, in addition to age/earnings growth due to greater work experience as indicated by cross-sectional age earnings profiles, there is a 1 percent growth rate in real earnings over the life cycle.¹³

Table 2 Variable Means for Male Workers
(standard deviations of continuous variables in parentheses)

Variable	Ages 21–35			Ages 36–55		
	1979	1988	1993	1979	1988	1993
Covered	0.569	0.481	0.453	0.694	0.674	0.657
Tax	30.8 (8.0)	26.0 (9.3)	26.1 (8.7)	35.3 (7.9)	30.4 (9.0)	29.8 (7.5)
Pearn55 ^a	31.52 (9.75)	29.08 (10.85)	26.80 (10.86)	36.73 (11.08)	36.85 (13.46)	33.90 (13.48)
Union	0.285	0.169	0.128	0.343	0.250	0.196
Mult1000	0.423	0.403	0.397	0.469	0.485	0.469
African American	0.055	0.058	0.056	0.054	0.050	0.057
Married	0.705	0.611	0.584	0.880	0.831	0.785
Newhire	0.206	0.188	0.193	0.074	0.094	0.095

SOURCE: 1979 and 1998 March and May Current Population Surveys and 1993 March and April Current Population Surveys. 1978 tax rates are from Commerce Clearinghouse (1979). 1987 and 1992 rates are from Advisory Commission (1988 and 1993).

NOTE: The sample includes all full-time, male, private, wage and salary workers aged 21 to 55. The sample is further restricted to those who have valid responses to questions relevant to this study. The sample size is 5,496 in 1979, 6,241 in 1988, and 6,157 in 1993.

^a Predicted earnings at age 55, in units of 10,000 1993 dollars. The variable used in the regressions was transformed by taking logarithms.

The predicted earnings measure is affected not only by changes in the worker's current earnings but also by changes in the entire age/earnings profile for workers of that gender for the given year. It is also affected by changes in the rate of return to experience, unionization, industry of employment, and firm size.

This measure of predicted earnings is then included as an explanatory variable in the coverage equation. In addition to circumventing the endogeneity problem associated with current earnings, the instrumental variable approach measures (although imperfectly) the earnings power of all individuals at the same age. The imprecision is greater for young workers, for whom we project for more years.

Predicted earnings at age 55 for young male workers fell by \$2,400 (1993 dollars) between 1979 and 1988 and fell another \$2,300 between 1988 and 1993. The fall during the 1980s occurred because young males were moving to lower paid occupations and industries in greater numbers than older males. In addition, young males experienced relatively large declines in unionization and in employment in large firms.¹⁴

COVERAGE ESTIMATES FOR MALE WORKERS

In this section, we present evidence from our data on the effect of the decline in income tax rates on the pension coverage of males. We examine effects separately for young and older workers for each of the three years of data.

We tested whether we could pool our data by age group, gender, or year, and the equality of coefficients across groups was always rejected. This result is in itself interesting because pension antidiscrimination rules limit firms' ability to target specific groups of workers. Whether a firm provides a pension to a worker should depend on the collective characteristics of the workers in the firm rather than the individual characteristics of the worker. Sorting in the labor market may account for the differing coefficients across age and gender groups.

Table 2 contains variable means for the three sample years while Tables 3 (young males) and 4 (older males) contain the estimates of the marginal effects for the continuous variables and the delta effects for the discrete variables.

A decline in marginal tax rates can result in reduced pension coverage rates through several paths. Some firms may decide to terminate plans that have diminished value to their workers. New firms that otherwise would have offered a pension plan may decide not to do so. Workers may change jobs, leaving firms offering a pension plan and moving to firms without one. Workers in firms where pension participation is optional may choose not to participate. Finally, workers entering the job market who would have otherwise sought firms offer-

ing a pension plan may instead seek employment with nonpension firms.

The extent to which workers change their pension coverage status in reaction to a change in marginal tax rates depends on the extent of job change within the economy, which depends on the phase of the business cycle. It also depends on the length of time workers have had to adjust to tax rate changes and the length of time workers expect those new tax rates to be in effect. Workers in firms offering only a 401(k) plan can adjust their pension status more quickly than other workers because they can simply decide not to participate in the plan. We do not attempt to distinguish by which path a change in marginal income tax rates influences pension coverage. Because of dynamic aspect of these factors, however, we expect the estimated tax coefficients to vary over time.

The estimated coefficient on marginal tax rates is positive and significant in all six of the male samples (Tables 3 and 4).¹⁵ Thus, these results suggest that the decline in marginal tax rates during the 1980s reduced pension coverage for both young and older males.

The coefficient on predicted earnings at age 55 is positive and significant for all samples.¹⁶ Given the positive correlation between marginal income tax rates and employee income, the finding of significant positive effects for predicted earnings as well as taxes is important because it suggests that we have isolated separate income and tax effects.¹⁷

One way to quantify the predicted effect of changes in tax rates on pension coverage is to multiply the estimated marginal effect of taxes by the observed change in taxes. This approach is equivalent to taking the difference in the predicted probabilities of coverage with mean tax in the base year and mean tax in the comparison year, evaluated at the means of the other variables in the base year. The difference in the predicted probabilities gives the change in the estimated probability attributable to the tax change for an "average" individual.

This approach has the weakness that the sum over changes in all variables does not equal the change in coverage predicted by the model. Even and Macpherson (1990) developed a technique without this defect for calculating the predicted effect of changes in one variable. With their technique, the sum over changes in all variables is constrained to equal the total predicted change.

Table 3 Marginal And Delta Effects for Young (age 21–35) Male Pension Coverage Probit, 1979, 1988, and 1993 (t-statistics in parentheses)

Variable	1979	1988	1993
Tax	0.0040 (2.34)	0.0036 (2.03)	0.0043 (2.32)
Pearn55	0.467 (3.97)	1.071 (371.99)	1.064 (222.71)
Union	0.303 (17.29)	0.156 (15.54)	0.166 (25.81)
Mult100	0.278 (24.48)	0.180 (35.82)	0.123 (25.83)
African American	0.055 (3.19)	-0.180 (0.64)	-0.013 (1.35)
Married	0.147 (33.48)	0.010 (31.60)	0.129 (21.26)
Newhire	-0.192 (36.10)	-0.205 (18.11)	0.029 (3.16)
Intercept	-0.863	-1.744	-1.877
Log likelihood	-1492.9	-1534.3	-1524.1
N	3008	2931	2949

SOURCE: 1979 and 1988 March and May Current Population Surveys and 1993 March and April Current Population Surveys. The 1978 tax rates are from Commerce Clearinghouse (1979), 1987 and 1992 rates are from Advisory Commission (1988 and 1993).

NOTE: The sample includes all full-time, male, private, wage and salary workers aged 21 to 55. The sample is further restricted to those who have valid responses to questions relevant to this study. The probit equation also contains four controls for education, four controls for occupation, and nine industry controls. All controls were coded as dichotomous variables.

Table 4 Marginal and Delta Effects for Older (age 36–55) Male Pension Coverage Profit, 1979, 1988, and 1993
(*t*-statistics in parentheses)

Variable	1979	1988	1993
Tax	0.0033 (1.97)	0.0066 (4.18)	0.0048 (2.86)
Pearn55	0.370 (5.28)	0.627 (18.76)	0.662 (24.01)
Union	0.224 (11.74)	0.195 (9.64)	0.154 (7.81)
Mult100	0.237 (14.67)	0.190 (13.95)	0.224 (17.14)
African American	-0.023 (0.78)	-0.033 (1.26)	0.148 (5.02)
Married	-0.021 (0.85)	0.077 (5.01)	0.035 (2.90)
Newhire	-0.150 (10.74)	-0.126 (8.80)	-0.012 (0.44)
Intercept	-0.628	-1.137	-1.312
Log likelihood	-1102.9	-1248.4	-1452.9
<i>N</i>	2516	2886	3213

SOURCE: 1979 and 1988 March and May Current Population Surveys and 1993 March and April Current Population Surveys. The 1978 tax rates are from Commerce Clearinghouse (1979), 1987 and 1992 rates are from Advisory Commission (1988 and 1993).

NOTE: The sample includes all full-time, male, private, wage and salary workers aged 21 to 55. The sample is further restricted to those who have valid responses to questions relevant to this study. The probit equation also contains four controls for education, four controls for occupation, and nine industry controls. All controls were coded as dichotomous variables.

The predicted change in coverage between 1979 and 1988 is calculated:

$$EXP = \sum_{i=1}^{N88} \Phi(Z_{i88}\hat{g}_{88}) - \sum_{i=1}^{N79} \Phi(Z_{i79}\hat{g}_{79}), \quad (2)$$

where $N88$ is the number of observations in 1988, $N79$ is the number of observations in 1979, and Φ is the standard normal cumulative distribution function. EXP is the average predicted coverage rate in 1988 minus the average predicted coverage rate in 1979. Using 1988 as the base year, the portion of the predicted change attributable to changes in variable Z_k is

$$EXP_k = EXP * (\bar{Z}_{k88} - \bar{Z}_{k79})\hat{g}_{k88} / [(\bar{Z}_{88} - \bar{Z}_{79})\hat{g}_{88}] \quad (3)$$

where \bar{Z}_{79} and \bar{Z}_{88} are the vectors of variable means in 1979 and 1988; and \bar{Z}_{k79} and \bar{Z}_{k88} are the means of variable k in 1979 and 1988. A similar formula applies for base year 1988 in comparison to 1993.

Calculations of the effects of the changes in selected variables are presented in Table 5. These calculations indicate that the changes in marginal income tax rates, in the earnings measure, and in the percentage of the workforce covered by a union help explain the decline in pension coverage.¹⁸ Our results regarding the effects of declining unionization and earnings are comparable to those found by Bloom and Freeman (1992) and by Even and Macpherson (1994). However, we find that between 1979 and 1988, the effect of declining taxes was twice as large for young workers as for older workers. Between 1988 and 1993, for both young and older workers, the estimated effect is so small as to be economically insignificant.

A calculation indicates that pension coverage rate for males aged 21 to 36 was 1 percentage point lower in 1993 than it would have been had the marginal tax rates in 1979 been in effect.¹⁹ Dividing the estimated effect due to the change in tax rates by the change in tax rates, we find that a 1 percentage point increase in marginal tax rates on average leads to a 0.4 percentage point increase in pension coverage rates for this group.

As a test of robustness, we reestimate the model for males making three changes in the regression (Table 6). First, we pool the data. Sec-

Table 5 Predicted Changes in Male Pension Coverage Attributed to Changes in Observed Characteristics

Variable	Age 21–35		Age 36–55	
	1979–1988	1988–1993	1979–1988	1988–1993
Total predicted change	-0.073	-0.041	-0.011	-0.026
Change explained by				
Tax	-0.014	-0.3-e5	-0.007	-0.6-e3
Union	-0.009	-0.002	-0.004	-0.002
Pearn55	-0.057	-0.038	-0.004	-0.019

NOTE: These predicted changes are calculated using the 1988 estimates of the probability of coverage. Qualitatively similar predicted changes are found using the 1979 estimates of the probability of coverage. The percentage change attributable to changes in representation in manufacturing predicted a decline of 0.002 for young male workers and a rise of 0.002 for old male workers. Since the predicted change in coverage is sufficiently close to zero, we do not report the percentage of total predicted change attributable to underlying variables.

ond, we test for the effect of lagged taxes. Third, we use the log of current salary rather than our permanent earnings variable. Our measure of lagged taxes is the tax rate that would have applied in the second and third years of our data had the tax laws applying to the first or second years of the data prevailed. The interpretation of the lagged variable is complicated. It can indicate the effect of a lag in adjustment to taxes. It can also indicate for a particular time period that workers view lagged taxes to be more representative of the long run tax regime than they view current taxes.

COVERAGE ESTIMATES FOR FEMALE WORKERS

We test the robustness of our estimated tax effects for males by reestimating the model using data on females from the 1979, 1988, and 1993 CPS. The additional estimates of the marginal effect of taxes on pension coverage are an independent measure to assess the plausibility of our estimated tax effects for males. These comparisons across gender are particularly useful because changes in average tax rates were

Table 6 Estimated Marginal Tax Effects from Probit Regressions for Females, 1979, 1988, and 1993 (*t*-statistics in parentheses)

Sample	1979	1988	1993
Young females	0.0046 (2.28)	0.0048 (2.63)	0.004 (2.09)
Older females	0.0073 (2.30)	0.0046 (2.11)	0.0040 (1.90)

SOURCE: 1979 and 1988 March and May Current Population Surveys and 1993 March and April Current Population Surveys. The 1978 tax rates are from Commerce Clearinghouse (1979), 1987 and 1992 rates are from Advisory Commission (1988 and 1993)

NOTE: The sample includes all full-time, female, private, wage and salary workers aged 21 to 55. The sample is further restricted to those who have valid responses to questions relevant to this study. The probit equation also contains four controls for education, four controls for occupation, and nine industry controls. All controls were coded as dichotomous variables.

similar for both men and women, while changes in coverage rates were not. While coverage for men declined, women generally experienced rising coverage rates during the 1980s.²⁰

We estimate female pension coverage equations using the same specification used for males. To economize on space, we summarize results concerning tax effects (Table 6). Coverage rates for females in our sample rose by 0.5 percentage points between 1979 and 1988 and by 2 percentage points between 1988 and 1993.²¹ The mean tax rates in each year are virtually identical to those of men and display a similar trend.²²

Predicted earnings at age 55 for females rose over the period. The increase in predicted earnings reflects the rising wages and narrowing of the gender gap in wages that women have experienced from the late 1970s. In addition, the percentage of the workforce that is unionized fell by 50 percent between 1979 and 1993. The overall percentage decline in unionization is comparable to the percentage decline for men, although in absolute levels men are twice as likely to be unionized.

The percentage of women who were new hires, defined to have less than one year of tenure with the employer, fell from 19 to 18 percent between 1979 and 1988 and then fell to 14 percent in 1993. Since

eligibility for coverage usually requires some minimum level of tenure, the 4 percent decline in new hires is potentially important in explaining women's rising coverage rates between 1988 and 1993.

The estimated coefficient on marginal tax rates is positive and significant in five of the six female samples.²³ It is similar in magnitude to the estimated coefficients in the male regressions.

Table 7 presents the total predicted change in coverage and the predicted change in coverage attributable to changes over time in marginal tax rates. As we did for men, we use the 1988 coefficients as the base from which to extrapolate. The change in coverage for women predicted by our model is much smaller than it is for men, consistent with their smaller change in coverage.

Table 7 Predicted Changes in Female Pension Coverage Attributed to Changes in Observed Characteristics

	Age 21–35		Age 36–55	
	1979–1988	1988–1993	1979–1988	1988–1993
Total predicted change	0.024	-0.035	0.013	0.030
Change explained by				
Tax	–	-0.002	-0.001	-0.0–e3
Union	–	0.005	-0.004	-0.001
Pearn55	–	-0.045	0.018	0.026

NOTE: These predicted changes are calculated using the 1988 estimates of the probability of coverage. Qualitatively similar predicted changes are found using the 1979 estimates of the probability of coverage. Although the model predicts a 2.4 percentage point decline in coverage for young women between 1979 and 1988, almost all the change came about by changes in the estimated coefficients in the two years and not from changes in variable means. The decomposition described by Eq. 5 is meaningful only if the denominator is not close to zero. When this occurs, the predicted change attributed to any one variable becomes implausibly large because of division by a number close to zero. Therefore, we do not report predicted changes attributable to individual variables for young women between 1979 and 1988.

CONCLUSIONS

Private pension coverage rates for males declined during the 1980s, especially for young males. Previous studies of pension coverage for young males have ignored the decline in marginal income tax rates. Using data from the 1979, 1988, and 1993 CPS Pension Supplements, we find that the probability of coverage for an individual, in both our young and older samples, increases with increases in the marginal income tax rate. Declining marginal income tax rates are found to be nearly as important as the decline in unionism in explaining trends in coverage for young males. While our estimates vary, a rough summary indicates that a one percentage point increase in marginal tax rates causes a 0.4 percentage point increase in pension coverage rates. We find comparable tax effects for women.

Our results indicate that workers and firms react to changes in marginal income tax rates when making decisions concerning pension plans. Higher pension tax expenditures associated with higher marginal income tax rates “pay for” increased pension coverage.

Our results have implications for a number of issues not directly addressed in the paper. The decline in generosity of pension plans that many analysts believed occurred during the 1980s may have been due in part to the fall in marginal income tax rates. To the extent that pension saving is new saving, rather than replacing saving that would have occurred in another form, the decrease in marginal tax rates may have caused a decrease in savings. Finally, our results suggest that the reduction in tax rates partially paid for itself because the lower tax rates were associated with reduced tax expenditures on pensions.

Notes

We have received helpful comments from Daniel Beller, Tasneem Chipty, Stephen Cosslett, William Even, William Gale, Richard Hinz, David McCarthy, Donald Parsons, Alan Viard, participants at the Center for Pension and Retirement Research conference at Miami University of Ohio, and participants at seminars at the University of Vienna, the University of Southern California, the American Economic Association meetings, the Tax Economists Forum, and the Congressional Budget Office. Tzu-Kuang Hsu and Zooyob Anne provided excellent research assistance. Patricia Reagan thanks the Center for Labor Research at the Ohio State University for financial support.

The opinions expressed in this paper are the responsibility of the authors and do not represent the position of the U.S. Department of Labor.

1. See Doescher (1994) for an extensive survey of studies on pension coverage.
2. We follow traditional usage and define coverage to indicate that a worker is a participant in an employer-provided pension plan. When discussing 401(k) plans, we draw the distinction between being offered a plan and choosing to participate in it.
3. The primary legislative change in tax rates during the 1980s occurred with the passage of the Economic Recovery Tax Act of 1981, which cut the top federal marginal income tax rate from 70 percent to 50 percent and reduced marginal income tax rates in all other brackets by 23 percentage points over three years. The Tax Reform Act of 1986 reduced the top rate on wealthiest households to 38 percent, effective 1988. It provided for a transitional top rate of 38.5 percent, effective 1987. The highest rate in 1988 was 33 percent, which applied that year for single (unmarried head of household) [married couple] households with taxable income between \$44,315 and \$100,480 (\$61,650 and \$156,550) [\$71,900 and \$192,930]. The brackets increased in subsequent years.
4. The basic tax rules concerning pensions were established in the Internal Revenue Acts of 1922, 1926, and 1928. Employer contributions to private pension plans are not treated as income to workers. The investment earnings on those contributions accrue tax free. Benefits are taxed under the federal and state personal income taxes when received.
5. If managers of firms decide on whether to offer pension plans based, in part, on the desirability of pension benefits to themselves, the decline in tax rates at upper income levels will also affect the probability that lower income workers have pension coverage.
6. In 1979, 26 percent of all male workers with tax rates below 10 percent were covered. Coverage rates rose to 75 percent for the 40–49 percent tax bracket. A similar profile emerges from the 1988 and 1993 data.
7. We follow the convention of referring to the data by the year of the survey that it is from. The income data, and the income tax rate data derived from it, are for the year preceding the survey.
8. Since the unit of observation is a match between worker i and firm j , we should subscript the vector and subsequent stochastic terms by ij . However, for ease of notation we suppress these subscripts.
9. The probit estimates are available from the authors on request.
10. The practice of reporting marginal effects of continuous regressors is standard to the literature (see Even and Macpherson 1994). The authors, however, use the same formula to calculate the marginal effects of discrete regressors. We instead report delta effects for discrete variables. We also report t -statistics based on standard errors of the marginal and delta effects. These differ from the t -statistics on the coefficient estimates of the probit equation. We believe that our approach represents a technical improvement over previous work. The variance-covariance matrix for the marginal and delta effects is calculated by pre- and post-multiply-

ing the variance-covariance matrix of the probit estimates by the matrix of the derivatives of the vector of marginal and delta effects with respect to the elements in the vector g . The code is available on request.

11. Gustman, Mitchell, and Steinmeier (1994) criticized previous studies for not addressing the issue of the endogeneity of marginal tax rates. In principle, the worker's expected marginal income tax rates for all future years affect the demand for pensions. We do not pursue that approach empirically because of colinearity.
12. State tax data for 1987 and 1992 are contained in reports of the Advisory Commission on Intergovernmental Relations (1988, 1993). Data for 1978 were supplied by Commerce Clearing House (1979). The marginal income tax rate is calculated as follows. First, we take the family income data from the CPS, which is categorical, and replace it with the mean family income in each category. Since the data on family income is top coded, we use IRS Statistics of Income tables to obtain average family income conditional on income exceeding the maximum reported by the CPS. We then use information on marital status and number of children in the family from the CPS, coupled with information about allowed exemptions and deductions from the federal and state income tax codes to obtain a measure of taxable income. Taxable income was calculated separately for state and federal tax purposes. The combined federal and state tax rates take into account the deductibility of state income taxes in computing federal income tax rates.
13. We experimented with a 2 percent growth rate and found our results to be robust to the assumption of 1 percent growth.
14. Unionism is another variable that previous studies have found to have an important effect on pension coverage. Between 1979 and 1988, the percentage of workers covered by a union contract dropped 11.7 and 9.3 percentage points for young and old workers, respectively. Between 1988 and 1993, these rates dropped an additional 4.1 and 5.4 percentage points. Although older workers are more likely to be covered by union contracts in all years, the magnitude of the decline in unionization was large for both groups.
15. We found statistically significant positive effects for all samples when we entered marginal tax rates calculated from actual family earnings rather than predicted family earnings.
16. The reported t -statistics are for the marginal and delta effects. The marginal and delta effects and their t -statistics are calculated by a nonlinear transformation of the probit estimates. Because of the nonlinearity of the transformation, it does not preserve the t -values in the probit estimates. The transformation increases the t -value for variables with already large t -values, explaining the very large reported t -values for some of the earnings coefficients.
17. In addition, we find the standard results that pension coverage increases with earnings, education, firm size, union status, and a marital status dichotomous variable (1=married). When we entered age in regressions not shown, it is insignificant, age having been controlled for already in the choice of samples.

18. For the calculations for both 1979 to 1988 and 1988 to 1993, we use 1988 as the base year. These calculations are entirely based on statistically significant estimated coefficients.
19. We calculate this by multiplying the predicted tax effect on coverage rates by the number of male full-time private sector wage and salary workers not covered by a pension plan (U.S. Department of Labor 1994).
20. Between 1988 and 1993, young women experienced a slight decline
21. As with males, the coverage rates within the sample exceed the population coverage rates due to restrictions on valid responses to questions used in the regression analysis.
22. This is not surprising since marginal tax rates are based on family income.
23. The predicted earnings at age 55 is positive and significant, which suggests that the tax effect has been isolated.

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