Pensions, Health Insurance, and Tax Incentives

Leonard E. Burman
Richard W. Johnson
Deborah I. Kobes

Discussion Paper No. 14

January 2004

Leonard E. Burman is a senior fellow at the Urban Institute and codirector of the Tax Policy Center. Richard W. Johnson is a senior research associate at the Urban Institute. Deborah I. Kobes is a research assistant at the Urban Institute and the Tax Policy Center.

The authors gratefully acknowledge helpful comments from Joe Antos, Peter Orszag, Sheila Zedlewski, participants at the PWBA research conference in Washington, DC, and seminar participants at the University of Virginia and Fedesarrollo (Bogotá, Colombia). This research was funded by the U.S. Department of Labor, Employee Benefits Security Administration (contract J-9-P-7-0044). The opinions expressed are those of the authors.

The Urban-Brookings Tax Policy Center

The Tax Policy Center (TPC) aims to clarify and analyze the nation's tax policy choices by providing timely and accessible facts, analyses, and commentary to policymakers, journalists, citizens and researchers. TPC's nationally recognized experts in tax, budget and social policy carry out an integrated program of research and communication on four overarching issues: fair, simple and efficient taxation; long-term implications of tax policy choices; social policy in the tax code; and state tax issues.

Support for the Center comes from the Annie E. Casey Foundation, the Brodie Price Philanthropic Fund, the Charles Stewart Mott Foundation, the Ford Foundation, the George Gund Foundation, the Lumina Foundation, the Nathan Cummings Foundation, the Open Society Institute, the Sandler Family Supporting Foundation, and others.

Publisher: The Urban Institute, 2100 M Street, N.W., Washington, DC 20037

Contents

Abstract

In 2001, Congress significantly expanded the scope of Individual Retirement Accounts (IRA). This paper uses variation in IRA eligibility rules in the 1980s and 1990s to determine whether more widespread access to IRAs undermines traditional employer-sponsored pensions, especially 401(k)-type plans. This paper develops a model of the value of fringe benefits with a special focus on the valuation of pensions. The theoretical model illustrates how IRAs can reduce demand for employer-sponsored pensions. The empirical results, based on a model of compensation and job turnover, indicate that workers highly value pensions and health insurance. Workers with benefits are less likely to separate from their employers than workers without benefits. In addition, workers who face high tax prices for health insurance appear to value coverage less than workers for whom taxes make employer-provided health benefits inexpensive. However, based on the historical evidence, higher IRA limits appear to have negligible effect on employees' valuation of employer pensions. The findings suggest that restrictions enacted in 1986 on Individual Retirement Accounts (IRAs) did not significantly increase workers' demand for employer-sponsored pensions. Similarly, there is no evidence that the increase in IRA limits enacted in 2001 will undermine employer-sponsored pension plans. These conclusions, however, must be viewed with caution because the legislated increases are significantly larger than the accounts available historically.

Pensions, Health Insurance, and Tax Incentives

The Economic Growth and Taxpayer Relief Act of 2001 (EGTRRA) made sweeping changes to the rules governing pensions and Individual Retirement Accounts (IRAs). As these provisions were being debated, critics pointed out that the higher IRA limits—which will reach as much as $6,000 for some taxpayers, up from $2,000 under previous law—could crowd out private pensions, especially 401(k) plans, because the comparative tax advantage for employer-sponsored plans would diminish and in some cases even be eliminated (Lubick 1999). Although this critique has merit on theoretical grounds, no evidence exists about how IRAs have affected employees' demand for pension plans. This study attempts to shed some light on the empirical magnitude of the crowd-out effect.

This paper develops a model of the value of fringe benefits with a special focus on the valuation of pensions. The underlying approach is inspired by two models: the standard model of job search—in which people look for work until the wage offer exceeds a certain reservation wage—and so-called disequilibrium models of home valuation—in which people remain in their home until the value of living somewhere else exceeds the value of remaining in their current home plus the transaction costs of moving. The model developed here assumes that the value of a current job depends on a set of job attributes, including the fringe benefits offered, in much the same way that the value of a home depends on the size of the home, location, number of bedrooms, bathrooms, and other amenities. There are search costs or transaction costs to changing jobs just as there are costs to moving from one home to another. In the context of this model of job valuation, the search model is reframed in terms of reservation compensation, rather than a reservation wage, where the value of compensation is a function of job attributes and the attributes of the individual.

A key attribute of the individual is the tax treatment of fringe benefits and substitutes purchased outside of work. The tax treatment varies across individuals because people's tax rates vary with income under a progressive income tax. In addition, the tax price of substitutes for fringe benefits purchased outside of work, such as an IRA, depends on the individual's eligibility, which is a function of income and whether the individual or spouse is covered by a pension. Thus, eligibility varies across individuals. But it also varies over time because changes in tax laws have changed the rules for eligibility. The empirical model exploits both cross-sectional and time-series variation in prices to measure the value of defined contribution (DC) pensions.

The model parameters are estimated using maximum likelihood methods based on data from the Survey of Income and Program Participation (SIPP) for 1984, 1990, 1992, and 1996. The pooled cross sections span a major change in the tax law—the Tax Reform Act of 1986 (TRA86), which restricted eligibility for IRAs. The estimates suggest that the widespread availability of IRAs before TRA86 took effect did not significantly crowd out employer pensions. This result suggests that the increase in the IRA limits in 2001 and more recent proposals to expand tax-free savings account limits to $7,500 will not undermine the system of employer pensions. But the usual caveat about the perils of predicting outside the range of the sample applies. That is, $7,500 retirement savings accounts may crowd out some employer-sponsored plans even though $2,000 IRAs had only negligible effects.

This paper begins by reviewing the evidence on how employees value fringe benefits. Section II develops a model of job choice, which values fringe benefits indirectly as a device to retain workers, and explicitly accounts for the tax treatment of pensions and IRAs. Section III describes the data from the SIPP and the Internal Revenue Service (IRS), and the measures we compute to implement our empirical model. Section IV presents the empirical results. Section V concludes the paper.

I. Background

Employee benefits, particularly health insurance and pension plans, are important components of worker compensation. In 2001, benefits accounted for 27 percent of the total compensation received by civilian employees (U.S. Bureau of Labor Statistics 2001). About 68 percent of nonelderly adults received health insurance through the workplace in 2000 (Holahan and Pohl 2002), and about 57 percent of full-time wage and salary workers participated in pension plans in 1999 (U.S. Pension and Welfare Benefits Administration 1999).

For a number of reasons, employers can generally provide health insurance and retirement savings at lower costs than workers would face on their own. Employers have greater purchasing power than individual workers and access to better information. In addition, employers can pool risks across many individuals, and benefits are tax-free to the workers.

In providing fringe benefits, the employer is acting as the employee's agent. In this role, an employer arranges for the package of benefits that employees would choose to purchase if they had the same knowledge and purchasing power as the employer. However, most employers are not perfect agents, because their own interests are not always aligned with those of their workers. The firm will take into consideration the heterogeneous nature of worker preferences and balance these against the costs of tailoring benefits to each worker. Tax rules and state and federal insurance and pension laws further limit the firm's ability to tailor benefits.

Moreover, the firm must consider the changing future demand and supply of labor. When labor supply is abundant, employers may wish to structure pensions, health insurance, and other benefits to induce older, higher paid workers to retire relatively early. When labor supply tightens, the firm may choose to structure pension and health benefits to be neutral with respect to retirement age, or even to encourage older workers to remain with the firm. The firm-agent must structure pension benefits that appeal to workers with diverse preferences and yet reflect current and expected future labor market conditions.

Assessing the performance of the employer as an agent for the employee depends in part on how workers value the benefits they receive in the workplace. Despite the importance of benefits in the compensation package, there is much we do not know about how workers value them. The standard approach to measuring the willingness of workers to pay for benefits is to estimate a hedonic wage equation that models the equilibrium locus of wage and job characteristics (Antos and Rosen 1975; Rosen 1974). The locus of combinations of wages and benefits represents the different values that employers and employees place on benefits. For example, some workers, such as those who are older, more risk-averse, or those with high marginal tax rates or low rates of time preference, may prefer relatively more pension benefits in their compensation packages. Others, such as younger, less risk-averse workers or those in low tax brackets, may be less willing to forgo current cash wages in exchange for future pension benefits.

Hedonic wage equations estimate wages as a function of worker characteristics and job attributes (including benefits). In principle, the coefficients of the job characteristic variables on the right-hand side of the equation may be interpreted as implicit (or hedonic) prices of employee benefits—that is, what an employee is willing to sacrifice in wages for the particular benefit. In practice, these estimates are likely to be biased downwards, because unobservable factors that increase wages, such as unobserved productivity, are likely to be positively correlated with the receipt and generosity of employee benefits (Biddle and Zarkin 1988; Epple 1987). Probably because of these biases, many studies of compensating wage differentials have found that job attributes have relatively small effects on wages (Brown 1980; Schiller and Weiss 1980).[1] Another limitation with the hedonic wage approach is that it requires strong assumptions to extract the underlying structural parameters useful for policy analysis (Bartik 1987; Epple 1987). These models do not explain how changes in policy parameters might change the way workers value fringe benefits.

An alternative, potentially less biased approach to assessing how workers value fringe benefits is to compare the effects of wages and benefits on employee turnover. Quit probabilities decline as the value of compensation increases, all else equal. By measuring the relative effect of cash wages and benefits on quit rates, models of job change can reveal how much workers value the components of their compensation package.

Some previous work has used information on employee turnover to measure how individuals value working conditions and employee benefits. Gronberg and Reed (1994) matched employment data on young men from the National Longitudinal Survey of Youth (NLSY) to information about job attributes from the Dictionary of Occupational Titles (DOT). They found that poor working conditions, defined as exposure to extremes of heat, cold, wetness, vibrations, or hazards, significantly increased quit rates. They estimated that workers were willing to forgo \$1.23 per hour in earnings, or more than one-fourth of the average hourly wage, to avoid working under adverse conditions. However, when they used the same data to estimate hedonic wage equations, they found that workers were willing to forgo only \$0.33 per hour to avoid poor working conditions, suggesting that hedonic wage equations substantially understate the marginal willingness to pay for job attributes. Bartel (1982) estimated probit equations of job separations in the late 1960s for full-time workers in the National Longitudinal Surveys of Young Men and Mature Men. Using data from the U.S. Bureau of Economic Analysis, she matched the total value of fringe benefits to survey respondents by three-digit industry codes. Working conditions were imputed from the DOT. She found that the value of fringe benefits significantly reduced quit rates for mature men (age 45 to 59) across all occupations and for

[1] A number of studies have attempted to correct for these potential biases. For example, Duncan and Holmlund (1983) concluded that they can substantially reduce the bias from unobserved productivity differences by estimating models of wage changes, instead of wage levels. They found that dangerous and stressful work conditions were associated with relatively high wages when they differenced the wages, but not when they modeled wage levels. Montgomery, Shaw, and Benedict (1992) found significant trade-offs between wages and pensions, but only when they modeled the natural log of wages as a function of the ratio of annual accruals in pension wealth to wages.

young men (in their late teens and early 20s) in blue-collar occupations only. The impact of fringe benefits on turnover was substantially larger for older men than younger men.

Another strand of the literature emphasizes the importance of employer-sponsored health and pension plans to turnover rates. For example, a number of studies have found that employer-sponsored health insurance can discourage workers from quitting, "locking" them into their jobs, because some workers fear that insurance from potential new employers might exclude coverage for pre-existing conditions (Gruber and Madrian 1994; Madrian 1994; Stroupe, Kinney, and Kniesner 2000). Other research has shown that workers with employer-sponsored health insurance who lack access to retiree health benefits are less likely to retire than those with access to retiree benefits, because the costs of replacing employer-sponsored coverage with private nongroup insurance before reaching the Medicare eligibility age can be prohibitive (Blau and Gilleskie 2001; Johnson, Davidoff, and Perese 2003; Rogowski and Karoly 2000).[2]

Pension plans have also been shown to have important effects on employee turnover. One study, for example, found that workers are three times more likely to separate from jobs that do not provide pension coverage than from covered jobs (Gustman and Steinmeier 1993). Some economists argue that defined benefit (DB) plans are designed to bond workers to their employers by imposing capital losses on covered workers who separate before retirement. Payments from DB plans are often calculated as a percentage of workers' salary earned near the end of their careers. As a result, pension wealth falls when workers in DB plans leave their employers before they are eligible to receive retirement benefits, even if they are already vested (as long as nominal wages increase over time). Allen, Clark, and McDermed (1993) and Ippolito (1991) find that these capital losses explain much of the difference in job tenure and turnover rates between workers with and without pensions. Gustman and Steinmeier (1993), however, note that defined contribution (DC) plans, which are basically tax-deferred savings accounts into which employers and employees can make deposits, are not inherently backloaded and do not penalize separating workers. Nevertheless, they find that workers in DC plans are no more likely to quit than those in DB plans, suggesting that employee turnover rates are not driven by large potential losses in retirement benefits. Gustman and Steinmeier instead conclude that mobility rate differentials arise from the compensation premium earned by workers with pension coverage. According to this view, workers without pensions are more likely to quit because they stand to gain more from changing jobs.

II. Model of Job Choice

To model the valuation of health insurance and pensions, we develop a hybrid of the standard model of job search (Sattinger 1991; Stigler 1962) and the so-called disequilibrium model of

[2] Federal legislation was enacted during the 1980s and 1990s to improve portability of health benefits across jobs. The Consolidated Omnibus Reconciliation Act of 1985 (COBRA) guarantees individuals access to health insurance coverage from their former employers for up to 18 months after they separate, but employers can require former workers to pay 102 percent of the full premium costs. The Health Insurance Portability and Accountability Act of 1996 (HIPAA) limits the use of preexisting condition exclusions and forbids group health plans from denying coverage based on the health status of workers or their families. Although these laws are designed to reduce job lock, health insurance concerns still appear to reduce employee mobility (EBRI 1998).

housing demand (Dynarski 1985; Fair and Jaffee 1972). In the simplest model of job search, a worker searches for employment until the wage offer, W, reaches or exceeds a reservation wage, W^*.[3] In the disequilibrium model of housing demand, a household tries to match up the quantity of housing units, H—an artificial index of the value to the family of an array of housing attributes—with its desired housing consumption, H^*, which is a function of price, income, and demographic characteristics. When someone moves into a home, housing consumption equals desired consumption, so that $H = H^*$. Over time, H^* can diverge from H because of changes in such factors as income, marital status, family size, and age. If the divergence grows large enough to offset the transaction costs of moving, the household will move to a new home and once again regain equilibrium.

Applying this model to job search, the value of a job offer is not simply the wage, but includes a number of costs and benefits of the employment offer (Brown 1980). Thus, W should be replaced in the search equation with C, the subjective value of a particular compensation offer to a particular household. C depends on all the tangible and intangible benefits and costs of a particular job, and the valuation of a particular job depends on each individual's personal circumstances. Thus, two people might value the same job offer differently, just as two families might value the same home differently.

Over time, C can fall short of C^* for several reasons. The worker's valuation of the job could change because of personal or family circumstances or income. For example, workers may value pensions more as they age or their incomes increase. They may value health insurance if they marry and plan to have children. Disequilibrium could also occur if the employer changes the package of cash and noncash compensation. Or the worker's reservation compensation, C^*, could change because the value of nonmarket labor increases—for example, because of childrearing—or because the worker perceives that he or she could get a higher (or lower) wage offer elsewhere. Workers will leave their current employment when $C^* - C \geq k$, where k is the transaction cost associated with job change.

Government policies can alter the trade-off between fringe benefits and wages. The most obvious factor is a change in marginal income tax rates. Since wages are taxable, a reduction in marginal tax rates, all else equal, should increase the demand for cash compensation relative to untaxed fringe benefits. Also, the existence of nontaxed or lightly taxed substitutes for fringe benefits can change their demand. For example, the availability of tax-free IRAs has varied over time due to statutory changes—most notably in 1986, when new legislation restricted IRA eligibility to moderate-income workers and those who did not participate in employer pensions—and the erosion of the contribution limit by inflation. All else equal, the reduction in access to IRAs and their lower real contribution limits should have increased the value of DC pension plans. Similarly, proposals to allow tax credits or deductions for individually purchased health insurance would, if enacted, tend to reduce the attractiveness of employer-sponsored health insurance (Blumberg and Nichols 2000; Burman and Gruber 2001).

[3] In more complicated models, W^* may be updated to reflect information gleaned from the job search.

A. Valuation of Pension Tax Benefits

A key research question is how the taxation of savings outside of pensions affects employees' valuation of employer pensions. A simple model illustrates how taxes might affect pensions. Suppose, for simplicity, that employees may save an unlimited amount in their employer pension tax free, but may save only a limited amount tax free in an IRA. As a first-order approximation, this is not a bad assumption. Only 4 percent of workers contributed the maximum ($9,500) to their 401(k) plan in 1996 (Richardson and Joulfaian 2001).[4] As a final simplification, assume that all saving is done for retirement.

In a two-period model, in which period 1 corresponds to working and period 2 corresponds to retirement, the value of pensions depends on the marginal tax rate, the effective tax rate on non-IRA savings, and the amount that may be contributed to IRAs. Figure 1 depicts a simple model of how an IRA affects saving and consumption.[5] For an individual who saves more than the IRA limit, L, labeled a "high saver" in the figure, the tax deduction for IRA contributions is equivalent to a lump-sum addition to income. On the margin, the subsidy does not affect the price, because marginal saving is done outside IRAs. Equilibrium consumption and saving occur where the individual's highest indifference curve is tangent to the budget constraint.

A low saver, whose saves less than L, would face a marginal incentive to save from the IRA tax deduction. Equilibrium consumption (and implicitly saving) for a low saver is shown on the right side of figure 1.

For estimation purposes, the utility function is represented as an indirect utility function. The utility level for the high saver, at (C_1^*, C_2^*), may be written as $V_h\{R_A, Y + Lrt_s/R_A\}$, where R_A is the gross after-tax return to retirement savings done outside of IRAs or pensions, defined as $1 + r(1 - t_s)$, r is the pre-tax rate of return on savings, t_s is the effective tax rate on non-IRA savings, Y is after-tax household income (before considering IRA tax benefits), and L is the IRA limit. Put differently, R_A is the price of current consumption in terms of forgone future consumption. By similar logic, an individual who saves less than L would achieve a utility level of $V_l(1 + r, Y)$. Because all saving can be done in the nontaxable form, the gross return of that saving is simply $1 + r$.

Individuals with access to employer pensions are assumed to be unconstrained by contribution limits. They can achieve utility level $V(1 + r, Y)$, even if they are high savers. Thus, the value of a pension to a high saver is

$$V_p = V_h(1 + r, Y) - V_h(R_A, Y + Lrt_s/R_A) \qquad \text{if } S > L, \tag{1}$$

---

[4] Some individuals may have contributed less than $9,500, but still been constrained by limits imposed by their employers or required by law, but it is unlikely that this group is very large.

[5] See, e.g., Burman, Cordes, and Ozanne (1991). Some authors argue that the simple life-cycle model is not consistent with empirical evidence on IRA participation and savings (e.g., Poterba, Venti, and Wise 1996). Other authors, however, strongly disagree (Engen, Gale, and Scholz 1996). Resolving that debate is beyond the scope of this paper.

where S is saving. V_p will be positive as long as $t_s > 0$ (see figure 2), that is, if the tax benefits of savings outside pensions and IRAs are less than the tax benefits of pension saving. The value of pensions to low savers is 0, however, since they could garner the same tax benefits through IRAs.[6]

Some straightforward comparative statistics arise from this simple model. The partial derivatives of the utility function with respect to gross return and income, V_1 and V_2, are both positive under standard assumptions about consumer preferences. Thus, an increase in the IRA limit reduces the value of employer pensions, for two reasons. First, a higher IRA limit raises the value of IRAs to high savers (shifts out the IRA budget constraint in figure 2) without changing the value of pensions, so V_p declines directly. Second, when L increases, more people are able to do all their saving in IRAs, i.e., the cut-off for "high savers" moves to the left in figure 2. Those who are "low savers" get no additional tax benefit from pensions.

An increase in the gross return on non-IRA savings, R_A, which could arise from a reduction in the tax rate on such savings, also reduces the value of pensions. It increases the slope of the left-hand segment of the IRA budget constraint and reduces the difference between the pension and IRA utility levels.

To parameterize the effect of taxes on pension valuation, consider the following linear approximation to the indirect utility function:

$$V(R, Y) = a_0 + a_1 R + a_2 Y + e. \tag{2}$$

Substituting this approximation into equation (1) yields the following expression:

$$\begin{aligned} V_p &= \begin{cases} a_1\left[(1+r) - R_A\right] - \dfrac{a_2 rLt_s}{R_A} & \text{if } S > L \\ 0 & \text{otherwise} \end{cases} \\ &= \begin{cases} a_1 rt_s - a_2 \dfrac{rLt_s}{R_A} & \text{if } S > L \\ 0 & \text{otherwise} \end{cases} \end{aligned} \tag{3}$$

For people who save more than L, the IRA limit, the value of pensions depends on two terms. The first is the difference in after-tax return on savings between taxable and nontaxable accounts, rt_s. The second is the value of the IRA tax benefit, which has a pure income effect. For people who would save less, pensions have no value (at least in this simple model), because they can earn the same return on IRAs as on pensions.

[6] There may be nontax benefits from pension saving. For example, people may value the convenience of regular payroll deductions. Also, to the extent that workers do not pay for employer matching contributions in the form of lower wages (or do not perceive that they do), there are additional benefits to employer pensions.

A key insight is that the price variable for pensions depends on whether the taxpayer would save more than the IRA limit in the absence of a pension. Of course, saving behavior in the absence of a pension is not observable for taxpayers who have pensions. We can, however, estimate the probability that an individual would save more than the IRA limit, *L*, based on variation in IRA limits across individuals and over time, as discussed below. In addition to income, other control variables include income, gender, education, components of household wealth, and other demographic variables. That estimated probability, $\mathbf{p}$, of saving more than the IRA limit is used in a structural model of employee benefit valuation developed in the next section.

B. Empirical Specification

Suppose there are *J* fringe benefits. The value of fringe benefit *j* to employee *i* is F_{ij}, which is determined by an equation of the form:

$$F_{ij} = Z_{ij}\mathbf{a}_j + X_{ij}\mathbf{b}_j + ZX_{ij}g + e_{ij}\ . \tag{4}$$

The explanatory variables are functions of two vectors, Z_{ij}, a vector of fringe benefit attributes, and X_{ij}, a vector of individual-specific attributes. The specification also allows for interaction effects between individual and benefit attributes, denoted ZX_{ij}. To limit the number of independent variables, most of the possible interactions are constrained to be zero. Finally, the unobservable disturbance term is assumed to be independent and identically distributed.

The focus of our analysis is the two most important fringe benefits: pensions and health insurance. For pensions, the *Z* vector includes the type of pension (DC or DB), vesting requirements,[7] and whether the employer contributes (for those with DC plans). The individual-specific variables, *X*, include spouse's pension coverage and the valuation of tax benefits for pensions relative to alternative nonemployer retirement savings, developed above.[8]

The *X* vector also includes a set of demographic variables, including age, after-tax income (other than from own employment), marital status, education, industry, and occupation.

For health insurance, the *Z* vector includes participation in employer-sponsored insurance (ESI) and the tax price of ESI. The tax price depends on whether the employee has to pay a share of the premium out of after-tax income, i.e., whether the employer pays for only a portion of the

---

[7] The vesting index could be defined as $\sum_{t=1}^{T} p_t v_t$, where p_t is the probability a worker terminates employment after *t* years and v_t is the percentage vested after *t* years. The probabilities could be measured based on industry code and firm size, or at a higher level of aggregation. Unfortunately, data were not available to calculate this index.

[8] Other variables that could be included in this type of analysis, were they available, are the match rate by employers, the number of investment options available to the worker, and penalties on early withdrawals (Burman, Coe, and Gale 1999).

insurance.[9] The X vector includes health status, whether a spouse is covered by health insurance, the tax price of private nongroup health insurance, and the demographic variables.

The total value of compensation is

$$(5) \qquad C_i = W_i + \sum_{j=1}^{J} F_{ij} = W_i + \sum_{j=1}^{J} \left[Z_{ij}\boldsymbol{\alpha}_j + X_{ij}\boldsymbol{\beta}_j + Z_i X_{ij}\boldsymbol{\gamma} + \boldsymbol{\varepsilon}_{ij} \right]$$

Equation (5) can be rewritten in somewhat simpler terms as

$$(6) \qquad C_i = W_i + \mathbf{Z}_i\boldsymbol{\alpha} + \mathbf{X}_i\boldsymbol{\beta} + \mathbf{Z}_i\mathbf{X}_i\boldsymbol{\gamma} + \boldsymbol{\varepsilon}_i ,$$

where the bold-faced vectors represent the concatenation of the corresponding vectors in equation (5) across all the fringe benefits, with duplicate variables eliminated, and corresponding changes to the parameter vectors.

In addition to the fringe benefit variables, Z also includes other working condition variables not related to particular benefits that might affect how an employee would value a job offer, such as firm size, location, union status, industry, and occupation.

The reservation compensation for individual i is

$$(7) \qquad C_i^* = \tilde{X}_i\boldsymbol{\theta} + \boldsymbol{\eta}_i$$

The information set for this equation, $\tilde{X}_i$, is an augmented version of X including variables that affect only the reservation compensation. For example, the location of the worker's residence can affect C^* because search costs are lower in areas with many employers (e.g., urban and suburban areas). In addition, the variables specific to the valuation of benefits will generally be excluded from equation (7). The disturbance term, η, is assumed to be a vector of independent random normal error terms.

For a new job, the compensation package exceeds the reservation compensation; that is, $C - C^* \geq 0$. Substituting from equations (6) and (7) yields the following expression:

$$(8) \qquad C_i - C_i^* = W_i + \mathbf{Z}_i\boldsymbol{\alpha} + \tilde{X}_i(\boldsymbol{\beta} - \boldsymbol{\theta}) + \mathbf{Z}_i\mathbf{X}_i\boldsymbol{\gamma} + \mu_i \geq 0 ,$$

where $\mu_i \equiv \boldsymbol{\varepsilon}_i - \boldsymbol{\eta}_i$.

[9] The ideal dataset (but not SIPP) would allow the Z vector to include the actuarial value of the employer's contribution to employee health insurance, and the number of plan choices. If the employer pays some of the premiums, the tax-price of ESI depends on whether employees can use flexible spending accounts to pay for their premiums out of pre-tax dollars.

A worker will remain in an old job for another year if $C - C^* \geq -k_i$, where k_i represents the transaction costs of search, lost firm-specific human capital, and possibly lost pension benefits from leaving a job. The costs of switching are determined by the following equation

(9) $$k_i = S_i f + u_i$$

where vector S_i includes tenure on the present job, a measure of lost benefits from a defined-benefit pension plan, and proxies for firm-specific human capital (firm size, industry, and occupation).

Substituting (9) into (8) yields the condition for remaining in a job:

(10) $$W_i + Z_i a + \tilde{X}'_i(b - q) + Z_i X_i g + S_i f + m_i + u_i \geq 0\,,$$

where the combined disturbance term, $m_i + u_i$, will be independent and identically distributed under the assumptions of the model. Note that some elements of S_i are also in X_i, which means that the corresponding coefficients of f will not be separately identifiable from (10).

Assuming that $m_i + u_i$ is normally distributed, equation (10) takes on the form of a probit maximum likelihood model. However, the pension variables of equation (3) depend on whether the individual would be at the IRA limit or not, as discussed above. Although we do not know with certainty whether the individual would be at the IRA limit, we can estimate the probability, p_i. Equation (10) can then be estimated by a simple extension of the probit model.

To develop the estimator, we can rewrite equation (10) in simpler form. The individual will remain in the job if:

(11) $$Q_i d_i + x_i \geq 0,$$

where $Q_i = \{W_i \mid Z_i \mid \tilde{X}'_i \mid Z_i X_i \mid S_i\}$, d_i is the corresponding vector of parameters from equation (10), and $x_i = m_i + u_i$. Two of the terms of Z_i—the terms on the right-hand side of equation (3)—depend on whether the individual would save more than the IRA limit. Define Q_{i1} as the value of the right-hand side variables when the IRA limit would be binding, and Q_{i0} as the value when it is not binding (i.e., with the two terms set equal to zero).

Define y_i as 1 when individual i changes jobs and 0 when he or she stays. Then the likelihood that individual i changes jobs equals:

(12)
$$L(y_i, Q_i \mid d_i) = y_i\left[p_i \Phi(-Q_{i1}d_i) + (1-p_i)\Phi(-Q_{i0}d_i)\right] + (1-y_i)\left[p_i \Phi(Q_{i1}d_i) + (1-p_i)\Phi(Q_{i0}d_i)\right].$$

As in the probit model, the parameters of (12) may only be measured up to a scalar constant, since multiplying all the coefficients and the standard error of x_i by a constant would not change the value of the likelihood function. Thus, we normalize the equation by setting the variance of x_i to 1.

We estimate the model in two steps. In the first stage, we estimate p_i by tobit maximum likelihood, as described in appendix 2. In the second stage, we substitute the predicted value for p_i into the likelihood function in (12) and estimate the parameters using maximum likelihood methods.

C. Identification of Key Parameters

Assuming that all the right-hand-side variables are exogenous, that the errors are asymptotically uncorrelated with the right-hand side variables, and that p_i is consistently measured, the parameters in equation (12) can be consistently estimated by simple maximum likelihood methods. Many elements of β and θ may not be identified, but the effects of the demographic variables are of only secondary interest in this research. The marginal effect of firm and fringe-benefit specific attributes in Z may be estimated consistently. Thus, for example, we can estimate how changing the tax price of pensions or health insurance would change the value of employer pensions.

A key advantage of estimating the model based on job switchers is that wages and fringe benefits may plausibly be treated as exogenous variables for analysis purposes. A classic problem in standard hedonic models of fringe benefits is that both wages and benefits are determined in part by the same unobservable individual characteristics, which cannot be adequately controlled for in a regression of benefits against wages. But it is less problematic to include wages and benefits on the right-hand side of equation (12), because the worker may be assumed to take both as given in making decisions about staying or leaving. The only problem would occur if the unobservable determinants of wages and fringes were also explanatory variables for the job change decision. In that case, their exclusion could bias the parameter estimates. But even then, if workers value both wages and fringe benefits, it is likely that the biases in coefficients on both elements would be roughly proportional. Since we are primarily interested in the valuation of fringe benefits *relative to* wages, potential biases are likely to cancel out.

Another potentially more important problem is that job changers may have consistently different preferences in wages and fringe benefits than those who stay with the same employer. For example, Ippolito (1997) has found evidence that pension benefits are used to attract and retain employees with low discount rates, those who tend to save and may have other desirable work characteristics. People with low discount rates may also value health insurance more than those with high discount rates. If the employee's unobservable discount rate belongs on the right-hand side of the job change equation, its omission may bias coefficient estimates for both pension and health variables downward (away from zero).[10] That bias will be mitigated to the extent that other variables, such as job tenure, serve as adequate proxies for the discount rate.

A final concern is that tax rates themselves are endogenous to decisions about wages and fringe benefits. Individuals who face high marginal tax rates will prefer to receive more compensation in the form of fringe benefits, which in turn lower their tax rates. Thus, the tax rate, wages, and compensation are correlated with each other. In their model of health insurance offers by employers and take-up by employees, Gruber and Lettau (forthcoming) address this problem by constructing a tax rate instrument that is independent of actual compensation. Endogeneity is a less likely problem in our model, because benefits are not on the left-hand side of the equation. However, it is possible that people may value fringe benefits for unobservable reasons, i.e., that the marginal value of fringe benefit j in equation (4), a_j, has a stochastic component. If the stochastic component is correlated with the tax rate, because people who especially value fringe benefits also tend to have lower marginal tax rates, all else equal, then this correlation could bias parameter estimates. Lacking a good alternative, however, we will follow standard practice and treat the coefficients as fixed parameters.

III. Data and Methods

The dependent variable in (12) is the dichotomous variable indicating job change. The sample is limited to full-time workers between the ages of 25 and 55. The independent variables include indicators of pension and health benefits (the Z variables in equations 4–10); the tax price of health insurance and the value of savings outside of pensions; information about spouses' benefits; health conditions that would affect valuation of ESI; and an extensive set of economic and demographic control variables.

Because our aim is to measure the effects of tax policy changes on employees' valuation of benefits, the sample pools cross sections from before and after passage of TRA86—a dataset collected from 1984 to 1986, and its follow-ups collected in 1990 to 1992, 1992 to 1994, and 1996 to 2000. The data are described in detail below.

A. SIPP Data

We analyze the interaction of compensation and job choice using data from SIPP, which follows a national sample of individuals over several years, re-interviewing them at four-month intervals (known as SIPP waves). The survey samples a new panel of respondents every few years. Our analysis uses the 1984, 1990, 1992, and 1996 panels because respondents in these panels were followed for at least one year after they provided information about their pension coverage. Thus, we can use these panels to examine how fringe benefits affect turnover rates. Most of our data are taken from the wave with pension information (wave 4, which was collected a year after

[10] Suppose people with high discount rates are more mobile, all else equal. Demand for health insurance and pensions will likely be negatively correlated with the discount rate. Health insurance and pensions will be negatively correlated with mobility if people value fringe benefits. Thus, excluding a measure of the discount rate implies a negative bias (away from zero) on the fringe benefit coefficients.

the initial interview); other waves provide employment, health insurance, health status, and tax information.

Our sample consists of wage and salary workers age 25 to 55 who worked at least 35 hours per week. SIPP interviewed each member of a household separately, generating self-reported information about income, pension, and health insurance. One respondent, however, provides tax information for the entire household.

The dependent variable is whether the individual changed jobs. Job changers are respondents who changed jobs within a year of reporting their pension benefits. The independent variables are pension and health coverage details, the tax prices of alternative retirement savings and health coverage, and spouse coverage. The model also includes demographic control variables.

We base our measure of job change on employer identifiers. Respondents were asked about their current employers each month, and each new employer is identified by a unique job ID. We identify job changers as respondents whose employer ID changed between wave 4 and wave 7 (which are one year apart).

A drawback of this approach is that job changes may be especially sensitive to measurement error in the employer ID variable. Miscoding (or misreporting) of the employer ID in either wave 4 or wave 7 would lead us to misidentify the respondent as a job changer. Even if these errors are random, Freeman (1984) shows that they can bias parameter estimates toward zero, possibly by substantial amounts. If reporting error is correlated with some independent variables in the model, such as age or education, the bias could also go in the other direction.

A related issue is that we cannot distinguish involuntary from voluntary job changes in every year of the SIPP data (although the reason for a job change is reported in some years). What we want to measure with our models is voluntary quits. If involuntary job changes are random, or at least uncorrelated with the compensation variables, the key parameter estimates are likely to be biased toward zero, because of the presence of random measurement error in our turnover variable. If, on the other hand, poor job performance reflects a low valuation of the employee compensation package (and other unobserved job attributes) then "involuntary" job changers may be similar to those who quit.[11] However, some separations, such as those related to plant closings, are not related to any employee decisions. Including them with the voluntary separations would tend to dilute the measured effect of employee benefits on transitions. We examine the sensitivity of our results to this measurement issue in appendix 1, using the subset of data that includes information on the reason for job change.

The pension information available in SIPP indicates whether the respondent was offered coverage, the types of plans offered (DB, DC, 401(k) or 403(b)-type plan, and profit-sharing plans), and, for those who are covered, whether the employer matches contributions, the number of years that the respondent participated in DB plans, and whether the spouse has pension coverage. We identified DC plan participants as those who said they participated in 401(k),

[11] There is indirect evidence that quitters resemble those who are fired. McEvoy and Cascio (1987) find that poor performers are more likely to quit their jobs than strong performers.

403(b), or profit-share plans, or that their pensions were based on the amounts they contributed. DB participants are those people who said their pensions were based on years of service and pay. Additionally, we considered respondents who said they were in multiple plans to have both DC and DB plans (although some of them may actually have had more than one kind of DC plan). Employer contribution to a pension plan is included as a dummy variable. The job tenure variable proxies for the number of years in a DB plan (among other things).

We calculated total income tax liability and marginal tax rates based on a measure of taxable income and the tax schedules in effect for each year of the sample. Tax information reported on the SIPP varies over time. The 1984 SIPP does not include data on adjusted gross income (AGI), taxable income, or tax liability, although it does identify those who itemize and file jointly with their spouses. We calculate AGI for that panel based on income variables reported on the SIPP. We set total income in 1985 equal to individual income for single filers.[12] Total income for joint filers included spousal income.[13] We deducted largely tax-free forms of income—Social Security, Supplemental Security Income (SSI), and railroad retirement—from total income to estimate AGI.[14] We impute itemized deductions as a percentage of AGI for those who reported that they itemized, as a function of filing status, homeownership, and AGI, based on an OLS regression on data from a random sample of 1985 tax returns.[15] Non-itemizers were assigned the standard deduction for their filing status. A two-earner deduction was calculated for dual-earner married couples equal to 10 percent of wages of the lower-earning spouse up to a maximum of $30,000.[16] Taxable income equaled AGI minus deductions. We computed tax payments based on total taxable income and the applicable 1985 tax tables.

Respondents in the 1990, 1992, and 1996 panels were asked to report tax liability and taxable income, but nonresponse rates were high, and taxable income and liability were only reported within broad categories that did not correspond to marginal tax brackets. For consistency across years, we chose to estimate marginal tax rates and tax liabilities using the same procedure as in 1984. The imputation for itemized deductions is based on data from 1990 tax returns.[17]

---

[12] Respondents in the 1984 SIPP were asked tax questions in 1985.

[13] We do not use the SIPP family income variable, because the family unit is not clearly defined in the SIPP, can change during the course of the year, and does not necessarily coincide with filing units.

[14] This is a simplification for Social Security recipients because those with higher incomes may owe tax on a portion of their benefits. However, because we restrict our sample to full-time workers between age 25 and 55, only a tiny fraction receive Social Security benefits (through a retired or disabled spouse).

[15] Tax data come from the University of Michigan Tax Panel, a public-use panel of tax returns compiled by the IRS for tax years 1979 to 1990.

[16] The two-earner deduction was repealed in 1986, so this adjustment was not made in later years.

[17] Although significant tax law revisions were enacted in 1990 and 1993, there were no significant changes to the definition of AGI between 1990 and 1996. The 1990 law did include a provision that phases out 3 percent of itemized deductions for some high-income taxpayers. That provision increased marginal tax rates by at most 1.2 percentage points for taxpayers with very high incomes. We did not account for that provision or numerous other complexities in the tax code in our simple tax calculator.

The measure of the tax price of health insurance is based on the computed marginal tax rate (t). For workers offered health insurance whose employers contribute to premiums, the tax price is $(1 - t - t_p)/(1 + t_p)$, where t_p is the Social Security and Medicare payroll tax rate (7.65 percent for all but those with high incomes).[18]

The measurement of the pension tax variables followed the two-period specification in equation (3), but with an adjustment to account for differences in age (and implicitly on time until retirement). The effective tax rate on savings outside of pensions or IRAs, t_s, depends in a complicated way on the array of fully taxed, lightly taxed, and untaxed investments available to individuals. To simplify, we examined two alternate assumptions—that the alternative investment is fully taxed, as a bond or interest-bearing account would be, and that the alternative is taxed at the long-term capital gains tax rate.

The expected growth in asset value is assumed to be the interest rate yield on 30-year Treasury bonds. The rate, r, was 10.79 percent in 1985, 8.61 percent in 1990, 7.67 percent in 1992, and 6.61 percent in 1997.[19]

The lump-sum equivalent value of a limit contribution to an IRA (the second term in equation (3)) is slightly more complicated in a multiperiod context than in the two-period model. It varies with length of time until retirement and thus with age.

For simplicity, we assume that all funds are withdrawn in a single lump-sum when the individual turns 65 and that tax rates remain constant over the lifetime.[20] Thus, a taxpayer who contributes x to an IRA will withdraw $x(1 + r)^N(1 - t)$ in after-tax retirement cash in year N (defined as 65 minus current age), where t is the tax rate on ordinary income.

Alternatively, a taxpayer can make the same after-tax contribution, $x(1 - \tau)$, to a taxable account. The contribution is lower because unlike contributions to IRAs, those to taxable accounts do not qualify for up-front tax deductions. That investment would yield $x(1 - t)(1 + r(1 - t_s))^N$ in retirement cash after N years.

Assuming discounting at the after-tax interest rate, the present value of the additional retirement income due to the IRA tax exclusion is

$$PV_{IRA} = x(1-t)\left[\frac{(1+r)^N}{[1+r(1-t_s)]^N} - 1\right]$$

The present value increases with the tax rate on alternative investments, t_s, and with the length of the holding period (N). That is, all else equal, it is greater for younger contributors.

---

[18] We initially imputed the employer share of premiums but the imputation was poor so we dropped it.

[19] Tax questions in the 1996 SIPP panel referred to 1997 payments, the time period covered by the pension questions.

[20] In fact, tax rates tend to fall after retirement, which raises the value of tax-free retirement saving. See Burman, Gale, and Weiner (2001).

The IRA limit variable, L, varies over time because of the 1986 change in tax law. Prior to 1987, the IRA limit was the lesser of earnings or $2,000 per individual ($4,000 per couple). The limit was $2,200 for a working individual with a nonworking spouse. After 1986, people who had pension coverage or whose spouses had pension coverage were only eligible to contribute to a deductible IRA if their incomes were below certain thresholds—$25,000 for singles and $40,000 for couples. Each individual's limit was the lesser of earnings or $2,000 per individual. The limit for a couple with one worker earning at least $2,000 and a nonworking spouse was $2,250. In 1997, the rules were adjusted for couples so that their limit was the lesser of combined earnings or $4,000, rather than the sum of two individual limits. The relevant question in valuing nonemployer retirement savings is whether employees could take advantage of IRAs if they left the current employer. Thus, for most single individuals in most years, L = $2,000, and for most working couples L = $4,000. The variation arises after 1986 when $L = 0$ for married people with incomes greater than $40,000 whose spouses are covered by a pension.

The other factor in equation (12) is the probability that a taxpayer would want to save more than the IRA limit. We estimated the probability of maximum contributions to IRAs in two stages. First, we estimated a right-censored regression model for those who made any contributions to a deductible IRA in 1990, 1992, and 1996, as a function of income and demographic variables. We excluded 1984 because the amount of contributions to IRAs was not reported on the 1984 SIPP panel. Second, assuming that pension participants would want to contribute to IRAs if they could not contribute to pensions, we estimated the probability that their contribution would exceed the IRA limit based on the applicable limit and the parameters estimated on the IRA sample. These results were applied to all years, including 1984. We estimated the probability for IRA participants using the same estimated equation. The probability of exceeding the IRA limit was set to zero for those in households that did not participate in pension plans or contribute to IRAs. (See appendix 2.)

For the empirical specification, we allowed for the possibility that workers treat employer contributions to pensions or 401(k)s differently than their own contributions. In theory, employees pay for employer contributions through lower wages, but this offset is not exact and employees may not fully understand the trade-off. Employees may view IRAs as very good substitutes for DC plans with no employer contributions, but an inferior alternative if employers contribute (to either DC or DB pension plans). Thus, we interact the two pension valuation variables with an indicator of whether the employer pays. This variable is only available for those who take up a pension plan. There is no measure for whether an employer would contribute to a plan in which an employee does not participate.

The other key element of compensation is the individual's earnings. In addition, after-tax household income, including spousal earnings and income from capital, are also included.[21] All income items are adjusted for inflation to be in constant 1992 dollars, based on the Consumer Price Index.

[21] Earnings are the annualized equivalent of monthly earnings in the wave in which the pension questions were asked. Nonlabor income is the sum of the respective components of income over the course of the year. This variable is the sum of interest income from money market funds and government bonds, dividends, rentals, mortgages, and royalties, and spouse's earnings. After-tax income is the sum of earnings plus nonlabor income minus tax liability, calculated as described above.

The job tenure variable was taken from the wave that includes employment history questions. Those who changed jobs since the job tenure question was asked had been at their job less than a year, and were therefore assigned a job tenure value of zero. Responses to the job tenure question indicate that a few individuals began their current job before age 16. We assigned those respondents missing job tenure values.

Sample statistics for each year and the pooled sample are displayed in table 1.

IV. Results

Table 2 reports the parameter estimates from the maximum likelihood estimation. The coefficients indicate the marginal effects of each variable on the criterion function that determines job changing (equation 11), and may be interpreted in the same way as probit estimates. A positive coefficient indicates a higher probability of leaving the job, equivalent to a lower valuation for the compensation package. As in the probit maximum likelihood, the variance of the error term, x, is constrained to equal unity to permit identification of parameters up to a scale factor.

The three columns show variants on results for the full specification. The first and third columns assume that t_s equals the tax rate on capital gains; the second assumes that it is the tax rate on ordinary income. The first two columns assume that employees value pensions and 401(k) plans only if they participate in the employer's plan. The specifications include a dummy variable for the offer of a pension, even if the employee does not participate. The third column assumes that employees value pensions if offered, even if they do not participate, and the specification includes a dummy variable for take-up.

The results are similar in the three specifications. The value of the log-likelihood function is essentially identical in each case. The estimates suggest that IRAs are not a statistically significant factor in employees' valuation of pensions, and the measured effect is quantitatively quite small. As explained earlier, we expect that the coefficients on the first two terms—rt_s—would be negative, and on the second two terms—PV_{IRA}—positive. In addition, if employees do not treat employer contributions as their own money, the coefficients should be larger in absolute value when the employer does not contribute, i.e., the case in which IRAs are the best substitutes for employer pensions.

Although the coefficients have the expected signs in the case when the employer does not contribute toward the pension plan in the first two columns, they are not statistically significant. Moreover, the measured income effect (PV_{IRA}) has the wrong sign when employers contribute, and is significantly different from zero in the second column. In the third column, based on employer-plan offers rather than participation, the relative price variable is statistically insignificant and has the wrong sign. The income effect term has the expected sign and is statistically significant when the employer does not contribute. Collectively, the IRA variables are only statistically significant at the 5 percent level in the second column, but that is not actually evidence in support of the model, since the only statistically significant term has the wrong sign.[22]

The implication of these findings is that, at the levels available in the 1980s and 1990s, IRAs did not crowd out participation in employer pension plans. As discussed below, this result appears to be robust with respect to alternate specifications.

Interestingly, after controlling for the value of IRAs, the offer of a pension itself has no explanatory power, but employer plan contributions (toward a DB or DC plan) significantly reduce the likelihood of quitting. If take-up is added as an explanatory variable in the alternative specification (column 3), it is statistically significant and with the expected sign. The incremental effect of employer contributions, conditional on take-up, is small and statistically insignificant in that case.

As found in other research, health insurance plays a significant role in employee turnover decisions. Workers with employer-sponsored insurance coverage on the current job are much less likely to separate from their employers than those without ESI coverage.

Conditional on participation, workers appear insensitive to the tax price of health insurance. Although the coefficient for the tax price variable has the expected sign—positive, meaning that ESI is less valuable to employees when it is more costly—it is always statistically insignificant.

Earnings, of course, are the largest component of compensation. All else equal, higher earnings are strongly negatively correlated with turnover. Long-tenured workers and those who work in union jobs are much less likely to quit than others. Workers over age 34 are also significantly less likely to quit than young workers, but older workers—those age 45 to 55—appear somewhat more likely to quit than workers age 35 to 44. This difference may reflect a slightly higher propensity among older workers to retire early or leave on disability.

Urban residents are more likely to quit than suburban and rural workers, probably because there are more alternative job opportunities in cities.

Women and workers with children under two are less likely to separate than other workers, but the number of older children and marital status are not statistically significant factors. So-called secondary earners—defined here as the lower-earning spouse—are neither more nor less apt to leave a job than primary workers. There do not appear to be significant differences by race. But less educated workers are statistically less likely to leave their job than those with at least some college, probably because the alternative job prospects for those with less education are not as good.

[22] Although not shown in the table, we also estimated the model using probit maximum likelihood and constraining the coefficients to be the same regardless of whether the employer contributed toward the pension plan. Using probit rather than the slightly more complicated maximum likelihood technique is equivalent to constraining p_i to 1 for all i. The log likelihood in this case was -22,259 compared with -22,258 in the unconstrained specification in column (1). The two models are statistically indistinguishable. The coefficients on the IRA variables were again small and statistically insignificant, and both had the wrong sign.

Workers in fair or poor health are significantly more likely to leave their jobs than those in good or excellent health. The small percentage of the sample who do not know or refuse to report their health status (3 percent or less) are much more likely to leave their jobs than those who answer the health status question. Some other missing indicators also are statistically significant, suggesting that the willingness to answer certain survey questions is not entirely random. We test the sensitivity of results to observations with missing variables below.

Those with more nonlabor income are much less likely to leave their jobs than others. This is a bit surprising as leisure is a normal good, so its demand should increase with income. It is likely that after-tax income proxies for unobservable individual and job characteristics that tie individuals to jobs. For example, people with low discount rates may choose to save more (so they have more income from capital) and marry higher-income people; since workers with low discount rates are also less apt to quit, income would have a spurious negative correlation with job mobility.

Firm size does not seem to matter, but administrative and clerical workers are significantly less likely to leave their jobs than workers in other occupations, and agriculture, mining, and construction workers are more mobile than those in other industries. Otherwise, there are no significant occupation and industry effects.

Finally, there are significant differences by year. Workers were significantly more mobile in 1984 and 1996 than in 1990 or 1992, after controlling for all of the other right-hand side variables. This could reflect business cycle factors—the economy was relatively weak in 1990 and 1992, making alternative job opportunities less promising—or it could reflect systematic differences in the sample design in different years.

We can use the estimated results to simulate how the increase in IRA limits already enacted and proposed by the president would affect the probability of job separation. By 2008, the IRA limits for workers under age 50 will increase to $5,000 under legislation enacted in 2001. Based on the point estimates reported in column 1 of table 2, this would increase the likelihood of quitting by 0.004 percentage points per year—a tiny and statistically insignificant amount.

The effects would be marginally greater if the president's proposals for expanded tax-free saving were enacted. In his 2003 budget, the president proposed $7,500 per person of tax-free "lifetime savings accounts" that could be used for any purpose (Burman, Gale, and Orszag 2003). Contributions to these accounts would be made from after-tax income, but all withdrawals would be tax-free. This makes the effective contribution limit significantly higher than the $7,500 for taxable individuals. In addition, IRAs would be replaced with similar "back-loaded" accounts, called retirement savings accounts, which would also have a $7,500 limit. The consequence is that an individual with earnings of at least $7,500 would face a combined contribution limit for both accounts of $15,000, and there would be no income limit. Simulating the proposed rules using our model, the quit rate would increase by about 0.1 percentage points, or about ½ of one percent of annual quits. Since we did not take account of the higher effective limit, this could be an underestimate.

However, it should also be noted that these simulations are far outside the range of historical experience for the model. Even though the historical response was statistically insignificant, it is possible that with overall annual contribution limits many times what had been available for IRAs, some employees and their employers would decide that sponsoring pension plans is no longer worth the cost.

A. Sensitivity Tests

We tested the sensitivity of our estimates to various specifications of the model and sample. Adding or deleting control variables does not alter any of the key conclusions. The first column of table 3 shows the results from the first model, but using a parsimonious list of control variables. The key coefficient estimates are statistically indistinguishable from those in column 1 of table 2. The only difference is that the variable for the number of children becomes larger and statistically significant when the separate measure of children under age 2 is excluded from the model.

The second column shows the results when more detailed pension and health insurance variables and more interaction terms are added to the model. Although the IRA variables take on the expected sign under the full specification, they are still statistically insignificant (p-value = 0.46). The estimates for the other pension variables are not much changed by adding additional detail, but the story for health insurance is much more interesting. Employer contributions to a DB or DC pension plan reduce quit rates, but it does not seem to matter much whether it is to a DB or DC plan. The coefficient estimates are not statistically different from each other, and only the coefficient on DC coverage is statistically significant at the 5 percent level.

Spouse's pension coverage does not seem to affect an employee's valuation of a job. We tested for age differences in valuation of DB and DC pension plans. The largest effect is for workers 45 to 55 years of age with DB plans, who are more likely to quit than younger workers or those with DC plans.[23] Older workers with DC plans are also significantly more likely to quit than younger workers and those without pension coverage, but the marginal effect is much smaller than for older workers covered by DB plans. The large DC effect likely reflects the availability of early retirement starting at age 50 in some DB plans for long-tenured employees. More generally, older workers with pension coverage may be more likely to quit because they are wealthier and better able to finance early retirement or a transition to a second career.

The health insurance coefficients change markedly when additional detail is added. First, the coefficient on ESI participation, which was already large in the base case, more than doubles. The tax price of health insurance becomes statistically significant and of the expected sign—the higher the price of health insurance, the more likely someone will leave a firm in which he or she is covered. Workers covered by health insurance are also more likely to separate from their employers when they are second earners.

---

[23] Recall that the sample is limited to full-time workers between the ages of 25 and 55. The age effects are measured relative to workers between 25 and 34 years of age.

Health status does not seem to affect workers' valuation of health insurance, but that could be because health problems force some workers to leave their jobs. There do not appear to be significant age-related differences in quit probabilities among those covered by health insurance.

A seemingly odd result is that, conditional on participation, workers are more likely to leave their jobs if their employers contribute toward health insurance than if the employers do not contribute. Employees who participate in their employers' health insurance plans without an employer match (the comparison group here) generally receive no tax benefits and have to pay the full premium. They are likely to place exceptionally high value on health insurance coverage, and might have health conditions that would preclude their coverage in the nongroup market if they moved to a job that did not offer health insurance. Thus, the story here may be that those workers are especially eager to retain their coverage, as indicated by the fact that they pay the most for it.[24]

Finally, the augmented specification suggests that income effects reverse sign with age. Workers under age 35 are significantly less likely to quit if they have higher nonlabor income, as noted above. However, this effect diminishes (and becomes statistically insignificant) for workers between 35 and 44, and reverses sign for workers 45 and over.[25]

Tables 4 and 5 decompose the sample by gender and marital status. There are significant differences between men and women in their valuation of health insurance. Both men and women are significantly less likely to separate from their employers when they have ESI coverage, but the tax price is more important (and only statistically significant) for men.

Men and women appear to value pensions similarly, and the effect of IRAs remains statistically insignificant. Employer matching is a significant factor for both men and women. There appear to be some small differences in pension valuation among older men and women. Older men (age 45 to 55) appear more likely to leave a firm where they are covered by either a DB or DC pension than women in the same age category, although older women covered by DB plans are more likely to quit than younger women and those without a DB plan. DC plans make older men more apt to quit, but not older women.

Both men and women appear to become much less likely to quit as they get older. Men with more children are significantly less likely to quit than men with fewer or no children and than women (with or without children). When the sample is split by gender, the effect of children under 2 becomes statistically insignificant for both sexes.

Married and single workers also appear to value benefits somewhat differently, as shown in table 5. Health insurance affects married workers' turnover rates much less than those of

[24] Note that the employer contribution variables are not available in 1984.

[25] The effect for workers in each age group equals the coefficient on the log of after-tax income plus the coefficient on the income term interacted with that age group. Thus, the net income effect for workers 35 to 44 is -0.036, and for those age 45 to 55, 0.057.

singles. The tax price of health insurance also makes a big difference to singles, but is not statistically significant for married households.

The effects of pensions on job change are similar for married and single workers, but many coefficients are statistically insignificant when the sample is split by marital status. The IRA variables have the correct sign, but remain statistically insignificant for married people. The coefficients where the employer contributes are statistically significant for singles, but they have the wrong sign. This anomalous result may occur because the tax price is only weakly identified for singles. The main source of independent variation—the applicable limit for IRA contributions—primarily varies across individuals due to the pension status of a spouse, which is not applicable for singles.

There are also some differences between singles and couples with respect to the demographic variables.

Table 6 repeats the estimates for the basic model (in table 2), but with the observations with missing values excluded (column 1) and an alternative measure of income tax (column 2). Excluding missings has virtually no effect on the results—all the statistically significant variables remain so and their coefficient estimates are nearly identical to those shown in table 2. To compute the estimates in the second column, tax rates were inferred from the tax liability reported on the SIPP starting in 1990. Respondents in the 1990, 1992, and 1996 panels were asked to report tax liability and taxable income, but nonresponse rates were high. For nonrespondents, we estimated marginal tax rates and tax liabilities using the same procedure as in 1984 (and as we used for the estimates reported in tables 2–5). For people who reported tax information, we calculated marginal income tax rates based on the tax liability variable in the SIPP tax module. Unfortunately, actual tax liability is not reported; the values are grouped in brackets that do not correspond to marginal tax brackets. SIPP tax liability categories generally span two marginal tax brackets. For tax liability brackets that span two marginal income tax rates, we used the IRS tax panel to impute the probability that the tax liability fell in a particular marginal income tax bracket, given that the individual's tax liability was within specified bounds. We then created two records for the respondent, one for each marginal income tax rate, and weighted them based on that probability.[26] Because the Michigan tax panel does not extend beyond 1990, the 1992 and 1996 SIPP estimates were also based on the 1990 tax panel.[27]

The results are not much different using the SIPP derived tax rates. A noticeable difference is that the relative price term for IRAs (rt_s) becomes statistically significant in the case where the employer does not contribute. However, overall the IRA variables remain statistically

[26] For example, suppose an individual has tax liability that spans the 15 and 28 percent tax brackets. Suppose, in addition, that SOI data indicate that 30 percent of people in that tax liability category are in the 15 percent tax bracket and the other 70 percent are in the 28 percent bracket. We would create two records for this respondent—one with an assigned marginal tax rate of 15 percent and a weight of 0.30, and a second with a marginal tax rate of 28 percent and a weight of 0.70.

[27] To verify that our two measures were comparable, we compared the tax rates created by our imputation process and those determined by tax liability. For 1996, 76 percent of individuals were assigned matching tax rates. All but 6 percent of individuals had a tax rate based on the imputation process that was within one of the range predicted by the tax liability bracket.

insignificant (p-value 0.14). In addition, dummy variables indicating that the tax rates are calculated rather than derived from SIPP are highly statistically significant in two of the three years, suggesting that the differences between the two methods of calculating tax rates are not random. Nonetheless, the similarity of the estimates to those in table 2 suggests that the results are robust with respect to the measure of marginal tax rate.

V. Conclusions

This paper is the first attempt to measure how IRAs affect the demand for employer-sponsored pension plans. The theoretical model shows that in a standard life cycle framework, higher IRA limits and broader eligibility would reduce the attractiveness of pensions for workers who are eligible for both. However, the empirical estimates based on that theoretical model suggest that IRAs have had a negligible effect on pension demand in the 1980s and 1990s.

More generally, the results indicate that workers value benefits quite highly. Those with pensions and/or health insurance are much less likely to leave their job than those without. Employer contributions to pension plans, either in the form of DB pensions or matching contributions to DC plans, significantly reduce quitting. Similarly, workers with ESI are much less likely to quit than those without. There is also some evidence that worker demand for health insurance is sensitive to its tax price, although this result is not robust to all specifications of the model.

These findings have several potential policy implications. The first is that the restrictions on IRAs enacted in 1986 had little or no effect on workers' demand for pensions. The second is that the increase in the IRA limits enacted in 2001 along with the decline in marginal tax rates can be expected to have little effect on workers' demand for employer-sponsored pensions. That is, the concerns that IRA expansion could undermine the employer pension system may turn out to be overblown.

Several major caveats are in order, however. Our findings may be distorted by the limitations in the SIPP data. Our data on IRA contributions applied to a very small sample of households, and many observations included missing values that had to be discarded. Our measures of both the dependent variable—job change—and the key independent variable—marginal tax rates—are subject to error. We attempted to test for the possible effect of those errors on our results, and found them relatively robust, but any errors would tend to bias our results against finding an effect of IRAs on pensions. It is also possible that households do not optimize in the way suggested by the simple two-period model. Many might not even be aware of the rules that apply to IRAs, especially after 1986 when financial institutions sharply curtailed their advertising.

On the other hand, it is possible that the much higher IRA contribution limits scheduled to take effect later this decade would have a much bigger effect on employees' psyche than the modest IRA limits in effect in our sample period. It is certainly possible that the very large tax-free savings accounts allowed under President Bush's budget proposals could have much larger effects than would be predicted by our model. In particular, employers may be tempted to shirk their role as employees' agents when they can satisfy all own needs for tax-free savings through individual accounts, even if employees still value their pensions. This type of response could not be measured in our model.

What is certain is that the new law should provide excellent data for reexamining these questions in a few years.

References

Allen, Steven G., Robert L. Clark, and Ann A. McDermed. 1993. "Pensions, Bonding, and Lifetime Jobs." *Journal of Human Resources* 28(3): 463–81.

Anderson, Patricia, M., Bruce D. Meyer, John Pencavel, and Mark J. Roberts. 1994. "The Extent and Consequences of Job Turnover." *Brookings Papers on Economic Activity, Microeconomics* 1994: 177–248.

Antos, Joseph R., and Sherwin Rosen. 1975. "Discrimination in the Market for Public Sector Teachers." *Journal of Econometrics* 3(1): 123–50.

Bartel, Ann P. 1982. "Wages, Nonwage Job Characteristics, and Labor Mobility." *Industrial and Labor Relations Review* 35(4): 578–89.

Bartik, Timothy J. 1987. "The Estimation of Demand Parameters in Hedonic Price Models." *Journal of Political Economy* 95(1): 81–88.

Biddle, Jeff E., and Gary A. Zarkin. 1988. "Worker Preferences and Market Competition for Job Risk." *Review of Economics and Statistics* 70(4): 660–67.

Blau, David M., and Donna B. Gilleskie. 2001. "Retiree Health Insurance and the Labor Force Behavior of Older Men in the 1990s." *Review of Economics and Statistics* 83(1): 64–80.

Blumberg, Linda J., and Len M. Nichols. 2000. "Decisions to Buy Private Health Insurance: Employers, Employees, the Self-Employed, and Non-Working Adults in the Urban Institute's Health Insurance Reform Simulation Model." Final report to the Pension and Welfare Benefits Administration, U.S. Department of Labor (Contract No.: J-9-P-7-0044). Washington, DC: The Urban Institute.

Brown, Charles. 1980. "Equalizing Differences in the Labor Market." *Quarterly Journal of Economics* 94(1): 113–34.

Burman, Leonard E., and Amelia Gruber. 2001. "First, Do No Harm: Designing Tax Incentives for Health Insurance." *National Tax Journal* 54(3): 473–93.

Burman, Leonard E., Norma B. Coe, and William G. Gale. 1999. "What Happens When You Show Them the Money? Lump-Sum Distributions, Retirement Income Security, and Public Policy." Final report to the Pension and Welfare Benefits Administration, U.S. Department of Labor (Contract No. J-9-P-7-0044). Washington, DC: The Urban Institute.

Burman, Leonard E., Joseph Cordes, and Larry Ozanne. 1991. "IRAs and National Savings." *National Tax Journal* 43(3): 259–83.

Burman, Leonard E., William Gale, and Peter Orszag. 2003. "The Administration's Savings Proposals: A Preliminary Analysis." *Tax Notes* 98(10): 1423–46.

Burman, Leonard E., William Gale, and David Weiner. 2001. "The Taxation of Retirement Saving: Choosing Between Front-Loaded and Back-Loaded Options." *National Tax Journal* 54(3): 689–702.

Davis, Steven J., and John Haltiwanger. 1999. "Gross Job Flows." In *Handbook of Labor Economics,* Volume 3, edited by David Card and Orley Ashenfelter. Amsterdam: Elsevier Science.

Duncan, Greg J., and Bertil Holmlund. 1983. "Was Adam Smith Right After All? Another Test of the Theory of Compensating Wage Differentials." *Journal of Labor Economics* 1(4): 366–79.

Dynarski, Mark. 1985. "Housing Demand and Disequilibrium." *Journal of Urban Economics* 17(1): 42–57.

Employee Benefit Research Institute (EBRI). 1998. "Health Insurance Portability and Job Lock: Findings from the 1998 Health Confidence Survey." EBRI Notes No. 8. Washington, DC: Employee Benefit Research Institute.

Engen, Eric M., William G. Gale, and John Karl Scholz. 1996. "The Illusory Effects of Savings Incentives on Saving," *Journal of Economic Perspectives* 10(4): 113–38.

Epple, Dennis. 1987. "Hedonic Prices and Implicit Markets: Estimating Demand and Supply Functions for Differentiated Products." *Journal of Political Economy* 95(1): 59–80.

Fair, Ray C., and Dwight M. Jaffee. 1972. "Methods of Estimation for Markets in Disequilibrium." *Econometrica* 40(3): 497–514.

Farber, Henry. 2003. "Job Loss in the United States: 1981-2001." Princeton University Industrial Relations Section Working Paper 471. Princeton: Princeton University.

Freeman, Richard. 1984. "Longitudinal Analyses of the Effects of Trade Unions." *Journal of Labor Economics* 2(1): 1–26.

Gronberg, Timothy J., and W. Robert Reed. 1994. "Estimating Workers' Marginal Willingness to Pay for Job Attributes Using Duration Data." *Journal of Human Resources* 29(3): 911–30.

Gruber, Jonathan, and Michael Lettau. Forthcoming. "How Elastic is the Firm's Demand for Health Insurance?" *Journal of Public Economics*.

Gruber, Jonathan, and Brigitte C. Madrian. 1994. "Health Insurance and Job Mobility: The Effects of Public Policy on Job-Lock." *Industrial and Labor Relations Review* 48(1): 86–102.

Gustman, Alan L., and Thomas L. Steinmeier. 1993. "Pension Portability and Labor Mobility: Evidence from the Survey of Income and Program Participation." *Journal of Public Economics* 50(3): 299–323.

Holahan, John, and Mary Beth Pohl. 2002. "Changes in Insurance Coverage: 1994–2000 and Beyond." *Health Affairs* Web exclusive, 3 April. W162-71. Available at http://content.healthaffairs.org/cgi/reprint/hlthaff.w2.162v1

Ippolito, Richard A. 1991. "Encouraging Long-Term Tenure: Wage Tilt or Pensions?" *Industrial and Labor Relations Review* 44(3): 520–35.

———. 1997. *Pension Plans and Employee Performance*. Chicago: The University of Chicago Press.

Johnson, Richard W., Amy J. Davidoff, and Kevin Perese. 2003. "Health Insurance Costs and Early Retirement Decisions." *Industrial and Labor Relations Review* 56(4): 716–29.

Lubick, Donald C. 1999. Testimony before the House Committee on Ways and Means Subcommittee on Oversight. U.S. Department of the Treasury, March 23. Available at http://www.ustreas.gov/press/releases/rr3036.htm.

Madrian, Brigitte C. 1994. "Employment-Based Health Insurance and Job Mobility: Is There Evidence of Job-Lock?" *Quarterly Journal of Economics* 109: 27–51.

McEvoy, Glenn M., and Wayne F. Cascio. 1987. "Do Good or Poor Performers Leave? A Meta-Analysis of the Relationship between Performance and Turnover." *Academy of Management Journal* 30: 744–62.

Montgomery, Edward, Kathryn Shaw, and Mary Ellen Benedict. 1992. "Pensions and Wages: An Hedonic Price Theory Approach." *International Economic Review* 33(1): 111–28.

Poterba, James, Steven Venti, and David Wise. 1996. "How Retirement Savings Programs Increase Saving." *Journal of Economic Perspectives* 10(4): 91–112.

Richardson, David, and David Joulfaian. 2001. "Who Takes Advantage of Tax-Deferred Saving Programs? Evidence from Federal Income Tax Data." Washington, DC: Office of Tax Analysis, U.S. Treasury Department.

Rogowski, Jeannette, and Lynn Karoly. 2000. "Health Insurance and Retirement Behavior: Evidence from the Heath and Retirement Survey." *Journal of Health Economics* 19(4): 529–39.

Rosen, Sherwin. 1974. "Hedonic Prices and Implicit Markets: Product Differentiation in Pure Competition." *Journal of Political Economy* 82(1): 34–55.

Sattinger, Michael. 1991. "Consistent Wage Offers and Reservation Wage Distributions." *Quarterly Journal of Economics* 106(1): 277–88.

Schiller, Bradley R., and Randall D. Weiss. 1980. "Pensions and Wages: A Test for Equalizing Differences." *Review of Economics and Statistics* 62(4): 529–38.

Stigler, George. 1962. “Information in the Labor Market.” *Journal of Political Economy* 70(5, Part 2): 94–105.

Stroupe, Kevin T., Eleanor D. Kinney, and Thomas J.J. Kniesner. 2001. “Chronic Illness and Health Insurance-Related Job Lock.” *Journal of Policy Analysis and Management* 20(3): 525–44.

U.S. Bureau of Labor Statistics. 2001. “Employer Costs from Employee Compensation–March 2001.” Bureau of Labor Statistics News. USDL 01-194. Washington, DC: U.S. Department of Labor. Accessed at www.bls.gov/ncs/ect/.

U.S. Pension and Welfare Benefits Administration. 1999. *Coverage Status of Workers under Employer Provided Pension Plans: Findings from the Contingent Work Supplement to the February, 1999 Current Population Survey.* Available at http://www.dol.gov/ebsa/programs/opr/CWS-Survey/hilites.html.

Table 1. Weighted Sample Means for Employment Characteristics, Income, and Demographic Variables, by Year

Variable	1984	1990	1992	1996	All years
Leave job	0.248	0.220	0.212	0.220	0.223
Pension variables					
Pension offered by employer	0.647	0.638	0.671	0.702	0.668
Participate in pension plan	0.581	0.543	0.590	0.572	0.572
Participate in DB plan	0.508	0.441	0.448	0.379	0.437
Participate in DC plan	0.150	0.276	0.330	0.411	0.308
Employer contributes to pension plan	0.570	0.526	0.574	0.549	0.554
rt_s	0.006	0.009	0.008	0.008	0.008
PV_{IRA}	472.9	407.9	375.7	332.4	388.6
Prob. of exceeding IRA contribution limit	0.136	0.317	0.346	0.307	0.290
Spouse covered by pension	0.289	0.263	0.267	0.259	0.267
Contribute to an IRA	0.169	0.034	0.033	0.030	0.057
Health insurance variables					
Participate in ESI	0.775	0.744	0.744	0.734	0.747
Employer pays some of ESI	N/A	0.445	0.494	0.497	0.394
Employer pays all of ESI	N/A	0.287	0.238	0.228	0.205
Spouse covered by ESI	0.347	0.272	0.263	0.223	0.269
Job variables					
Annual earnings (thousands of 1992 dollars)	10.077	10.124	10.097	10.127	10.109
Job tenure	7.625	7.152	7.986	7.543	7.550
Union member	0.238	0.202	0.182	0.172	0.194
Demographics					
Female	0.396	0.429	0.429	0.439	0.426
Age: 25–34	0.432	0.393	0.367	0.321	0.372
Age: 35–44	0.331	0.360	0.377	0.374	0.364
Age: 45–55	0.237	0.247	0.256	0.305	0.264
Urban resident	0.728	0.755	0.779	0.819	0.776
Marriage variables					
Married	0.712	0.659	0.666	0.657	0.670
Lower-earning spouse	0.171	0.180	0.175	0.198	0.182
Children					
Number age 18 or younger	0.999	0.898	0.930	0.940	0.937
Number age 2 or younger	0.120	0.121	0.128	0.101	0.117
Race					
Hispanic	0.052	0.074	0.085	0.095	0.079
Black, non-Hispanic	0.102	0.102	0.105	0.112	0.106
White, non-Hispanic	0.816	0.791	0.773	0.753	0.779
Non-white, non-black, non-Hispanic	0.030	0.033	0.037	0.041	0.036

Variable	1984	1990	1992	1996	All years
Education					
Less than a high school degree	0.144	0.115	0.104	0.099	0.112
High school degree	0.359	0.344	0.336	0.287	0.328
Some college	0.241	0.261	0.273	0.314	0.276
At least four years of college	0.256	0.280	0.287	0.300	0.284
Health status					
Excellent or very good	0.707	0.715	0.706	0.741	0.719
Good	0.234	0.236	0.243	0.215	0.232
Fair or poor	0.059	0.049	0.051	0.043	0.050
Missing	0.012	0.012	0.030	0.000	0.014
Income variables					
Log of after-tax income	10.099	10.113	10.129	10.304	10.170
Firm size					
Fewer than 25 people	0.220	0.259	0.179	0.156	0.199
25–99 people	0.151	0.116	0.101	0.115	0.116
At least 100 people	0.629	0.625	0.721	0.729	0.685
Missing	0.159	0.000	0.000	0.000	0.028
Occupation					
Managers and technicians	0.315	0.337	0.358	0.363	0.346
Sales	0.088	0.086	0.087	0.094	0.089
Administrative and clerical	0.165	0.164	0.168	0.156	0.163
Service, mechanics, and construction	0.197	0.201	0.183	0.190	0.192
Craft workers and machine operators	0.150	0.123	0.122	0.120	0.127
Transportation	0.039	0.043	0.039	0.034	0.038
Laborers	0.047	0.047	0.044	0.044	0.045
Missing	0.010	0.006	0.005	0.003	0.005
Industry					
Agriculture, mining, construction	0.074	0.072	0.074	0.075	0.074
Manufacturing	0.265	0.221	0.230	0.208	0.228
Transportation, public utilities, communication	0.087	0.095	0.085	0.081	0.087
Wholesale and resale trade plus Finance, Insurance, and Real Estate (FIRE)	0.211	0.212	0.226	0.231	0.221
Services and professional	0.287	0.322	0.318	0.331	0.317
Public administration	0.076	0.078	0.067	0.074	0.073
Missing					
Number of observations	7,680	11,096	11,212	16,597	46,585

N/A = not available

Table 2. Marginal Effects on the Probability of Job Change

Independent variables	Full take-up ($t_s = t_g$)	Full take-up ($t_s = t$)	Full offer ($t_s = t_g$)
Pension variables			
rt_s * employer contributes toward pension	-0.021	-0.699	0.210
	(1.872)	(1.789)	(1.874)
rt_s * employer does not contribute	-11.006	-12.446	3.399
	(6.845)	(6.642)	(5.448)
PV_{IRA} * employer contributes toward pension	-4.17E-05	-6.46E-05*	-3.72E-05
	(3.80E-05)	(2.90E-05)	(3.79E-05)
PV_{IRA} * employer does not contribute	1.17E-04	1.18E-04	2.56E-04*
	(1.50E-04)	(1.25E-04)	(1.19E-04)
Pension offered by employer	0.037	0.038	…
	(.023)	(.023)	
Employee has pension	…	…	-0.195**
			0.068
Employer contributes to pension plan	-0.222***	-0.210***	-0.006
	(.026)	(.026)	(.069)
IRA contributions unknown	-0.170*	-0.171*	-0.172*
	(.085)	(.085)	(.085)
Health insurance variables			
Participate in ESI	-0.387***	-0.359***	-0.398***
	(.084)	(.085)	(.085)
Spouse covered by ESI	-0.027	-0.027	-0.027
	(.020)	(.020)	(.020)
ESI coverage * tax price of health insurance	0.154	0.114	0.174
	(.115)	(.116)	(.116)
Job variables			
Log of annual earnings	-0.055***	-0.054***	-0.056***
	(.015)	(.015)	(.015)
Job tenure	-0.033***	-0.033***	-0.033***
	(.001)	(.001)	(.001)
Job tenure unknown	-0.832***	-0.832***	-0.828***
	(.109)	(.109)	(.108)
Union member	-0.094***	-0.094***	-0.092***
	(.020)	(.020)	(.020)

Independent variables	Full take-up ($t_s = t_g$)	Full take-up ($t_s = t$)	Full offer ($t_s = t_g$)
Demographics			
Female	-0.113***	-0.116***	-0.111***
	(.017)	(.017)	(.017)
Age: 35–44 years old	-0.109***	-0.110***	-0.108***
	(.017)	(.017)	(.017)
Age: 45–55 years old	-0.076***	-0.079***	-0.075***
	(.020)	(.020)	(.020)
Urban resident	0.139***	0.139***	0.139***
	(.017)	(.017)	(.017)
Number of children age 18 or younger	-0.009	-0.009	-0.009
	(.007)	(.007)	(.007)
Number of children age 2 or younger	-0.065**	-0.065**	-0.065**
	(.021)	(.021)	(.021)
Marriage variables			
Married	-0.025	-0.031	-0.022
	(.020)	(.020)	(.020)
Lower-earning spouse	0.004	0.005	0.003
	(.022)	(.022)	(.022)
Race			
Hispanic	-0.028	-0.027	-0.030
	(.025)	(.025)	(.025)
Black, non-Hispanic	0.020	0.017	0.022
	(.023)	(.023)	(.023)
Non-white, non-black, non-Hispanic	0.026	0.026	0.023
	(.035)	(.035)	(.035)
Education			
Less than a high school degree	-0.073*	-0.073*	-0.075*
	(.029)	(.029)	(.029)
High school degree	-0.060**	-0.060**	-0.060**
	(.022)	(.022)	(.022)
Some college	-0.020	-0.020	-0.019
	(.020)	(.020)	(.020)
Health status			
Good	-0.004	-0.005	-0.004
	(.017)	(.017)	(.017)
Fair or poor	0.191***	0.190***	0.192***
	(.031)	(.031)	(.031)
Missing	1.158***	1.159***	1.159***
	(.060)	(.060)	(.060)
Income			
Log of after-tax income	-0.137***	-0.135***	-0.138***
	(.015)	(.015)	(.015)

(continued)

Table 2 (continued)

Independent variables	Full take-up ($t_s = t_g$)	Full take-up ($t_s = t$)	Full offer ($t_s = t_g$)
Firm size			
Fewer than 25 employees	0.015	0.014	0.007
	(.019)	(.019)	(.019)
25–99 employees	0.022	0.023	0.018
	(.022)	(.022)	(.022)
Missing	-0.110*	-0.103*	-0.108*
	(.049)	(.049)	(.049)
Occupation			
Sales	0.024	0.024	0.023
	(.028)	(.028)	(.028)
Administrative and clerical	-0.062**	-0.063**	-0.061**
	(.023)	(.023)	(.023)
Service, mechanics, and construction	-0.033	-0.034	-0.033
	(.024)	(.024)	(.024)
Craft workers and machine operators	-0.019	-0.021	-0.017
	(.029)	(.029)	(.029)
Transportation	-0.019	-0.021	-0.018
	(.041)	(.041)	(.041)
Laborers	-0.049	-0.047	-0.048
	(.039)	(.039)	(.039)
Occupation unknown	-0.065	-0.065	-0.064
	(.101)	(.101)	(.101)
Industry			
Agriculture, mining, construction	0.188***	0.189***	0.186***
	(.029)	(.029)	(.029)
Manufacturing	0.025	0.027	0.023
	(.023)	(.023)	(.023)
Transportation, public utilities, communication	0.020	0.022	0.019
	(.030)	(.030)	(.030)
Wholesale and retail trade, FIRE	0.028	0.029	0.027
	(.021)	(.021)	(.021)
Public administration	-0.058	-0.061	-0.055
	(.031)	(.031)	(.031)
Industry unknown	0.101	0.103	0.100
	(.086)	(.087)	(.086)

Independent variables	Full take-up ($t_s = t_g$)	Full take-up ($t_s = t$)	Full offer ($t_s = t_g$)
Panel year			
1984	0.184***	0.192***	0.185***
	(.024)	(.024)	(.024)
1990	0.030	0.030	0.029
	(.020)	(.020)	(.020)
1996	0.078***	0.077***	0.081***
	(.019)	(.019)	(.019)
Intercept	1.676***	1.653***	1.698***
	(.166)	(.166)	(.166)
Fraction of sample that changed jobs	0.223	0.223	0.223
Number of observations	46,585	46,585	46,585
Number of parameters	50	50	50
Log likelihood	-22,258	-22,255	-22,255
Significance level for IRA variables ~ c^2**(4)**	0.4285	0.0474	0.0816

… = excluded variable

* statistically significant at the 5% level
** statistically significant at the 1% level
*** statistically significant at the 0.1% level

Table 3. Marginal Effects on the Probability of Job Change, Alternative Control Variables

Independent variables	Limited take-up ($t_s = t_g$)	Augmented take-up ($t_s = t_g$)
Pension variables		
rt_s * employer contributes toward pension	1.200	-0.136
	(1.842)	(2.057)
rt_s * employer does not contribute	-11.115	-8.936
	(6.830)	(6.985)
PV_{IRA} * employer contributes toward pension	-1.83E-05	5.07E-05
	(3.69E-05)	(4.06E-05)
PV_{IRA} * employer does not contribute	1.38E-04	1.93E-04
	(1.49E-04)	(1.48E-04)
Pension offered by employer	0.020	0.033
	(.022)	(.023)
Employer contributes to pension plan	-0.253 ***	-0.206 ***
	(.026)	(.038)
IRA contributions unknown	-0.163	-0.177 *
	(.085)	(.085)
Covered by DB plan	…	-0.057
		(.034)
Participate in DC plan	…	-0.070 *
		(.033)
Spouse has pension plan	…	-0.013
		(.023)
Employee has DB plan * 35–44 years old	…	0.001
		(.035)
Employee has DB plan * 45–55 years old	…	0.142 ***
		(.039)
Employee has DC plan * 35–44 years old	…	0.050
		(.038)
Employee has DC plan * 45–55 years old	…	0.083 *
		(.041)
Health insurance variables		
Participate in ESI	-0.376 ***	-0.871 ***
	(.084)	(.138)
Spouse covered by ESI	-0.027	-0.007
	(.019)	(.021)
ESI coverage * tax price of health insurance	0.119	0.490 ***
	(.115)	(.136)
ESI coverage * excellent or very good health	…	0.007
		(.035)
ESI coverage * fair or poor health	…	0.007
		(.068)
ESI coverage * lower-earning spouse	…	0.101 **
		(.036)

Independent variables	Limited take-up ($t_s = t_g$)	Augmented take-up ($t_s = t_g$)
Participate in ESI * 35–44 years old	…	0.011
		(.037)
Participate in ESI * 45–55 years old	…	0.073
		(.042)
Employer pays all of ESI	…	0.215 ***
		(.055)
Employer pays some of ESI	…	0.220 ***
		(.054)
Employer contributions to ESI unknown	…	0.179 **
		(.066)
Job variables		
Log of annual earnings	-0.033 *	-0.137 ***
	(.014)	(.021)
Job tenure	-0.035 ***	-0.034 ***
	(.001)	(.001)
Job tenure unknown	-0.879 ***	-0.841 ***
	(.109)	(.109)
Union member	…	-0.093 ***
		(.021)
Demographics		
Female	-0.120 ***	-0.103 ***
	(.015)	(.017)
Age: 35–44 years old	-0.102 ***	-0.974 ***
	(.016)	(.260)
Age: 45–55 years old	-0.074 ***	-1.991 ***
	(.019)	(.283)
Urban resident	…	0.139 ***
		(.018)
Number of children age 18 or younger	-0.018 **	-0.013
	(.007)	(.007)
Number of children age 2 or younger	…	-0.058 **
		(.021)
Marriage variables		
Married	-0.034	-0.021
	(.019)	(.020)
Lower-earning spouse	…	-0.051
		(.031)
Race		
Hispanic	-0.021	-0.032
	(.025)	(.026)
Black, non-Hispanic	0.015	0.028
	(.023)	(.024)
Non-white, non-black, non-Hispanic	0.038	0.027
	(.035)	(.035)

(continued)

Table 3 (continued)

Independent variables	Limited take-up ($t_s = t_g$)	Augmented take-up ($t_s = t_g$)
Education		
Less than a high school degree	…	-0.074 *
		(.029)
High school degree	…	-0.065 **
		(.022)
Some college	…	-0.025
		(.020)
Health status		
Good	-0.013	0.003
	(.017)	(.029)
Fair or poor	0.176 ***	0.201 ***
	(.031)	(.053)
Missing	1.139 ***	1.164 ***
	(.060)	(.065)
Income		
Log of after tax income	-0.125 ***	-0.122 ***
	(.015)	(.015)
Income * 35–44 years old	…	0.086 **
		(.027)
Income * 45–55 years old	…	0.179 ***
		(.029)
Firm size		
Fewer than 25 people	…	0.017
		(.020)
25–99 people	…	0.021
		(.022)
Missing	…	-0.094
		(.049)
Occupation		
Sales	…	0.025
		(.028)
Administrative and clerical	…	-0.064 **
		(.023)
Service, mechanics, and construction	…	-0.030
		(.024)
Craft workers and machine operators	…	-0.017
		(.029)
Transportation	…	-0.015
		(.041)
Laborers	…	-0.050
		(.039)
Occupation unknown	…	-0.061
		(.101)

Independent variables	Limited take-up ($t_s = t_g$)	Augmented take-up ($t_s = t_g$)
Industry		
Agriculture, mining, construction	…	0.186 ***
		(.029)
Manufacturing	…	0.027
		(.023)
Transportation, public utilities, communication	…	0.021
		(.030)
Wholesale and retail trade/FIRE	…	0.031
		(.021)
Public administration	…	-0.052
		(.031)
Industry unknown	…	0.099
		(.087)
Panel year		
1984	0.158 ***	0.215 ***
	(.023)	(.037)
1990	0.022	0.026
	(.020)	(.020)
1996	0.084 ***	0.081 ***
	(.018)	(.019)
Intercept	1.457 ***	0.215 ***
	(.149)	(.055)
Fraction of sample that changed jobs	0.223	0.223
Number of observations	46,585	46,585
Number of parameters	28	65
Log likelihood	-22,347	-22,195
Significance level for IRA variables ~ $\chi^2(4)$	0.5068	0.4569

… = excluded variable

* statistically significant at the 5% level
** statistically significant at the 1% level
*** statistically significant at the 0.1% level

Table 4. Marginal Effects on the Probability of Job Change by Gender, Assuming Non-IRA Savings Are Taxed at Capital Gains Tax Rates

Independent variables	Female	Male
Pension variables		
rt_s * employer contributes toward pension	2.472	-2.193
	(3.172)	(2.768)
rt_s * employer does not contribute	0.756	-14.089
	(11.043)	(9.086)
PV_{IRA} * employer contributes toward pension	1.47E-04	3.39E-05
	(8.50E-05)	(5.00E-05)
PV_{IRA} * employer does not contribute	4.62E-05	2.49E-04
	(2.96E-04)	(1.75E-04)
Pension offered by employer	4.07E-04	0.054
	(.033)	(.033)
Employer contributes to pension plan	-0.202***	-0.201***
	(.055)	(.053)
IRA contributions unknown	-0.077	-0.227*
	(.130)	(.114)
Covered by DB plan	-0.060	-0.060
	(.052)	(.046)
Participates in DC plan	-0.080	-0.068
	(.050)	(.044)
Spouse has pension plan	0.019	-0.041
	(.034)	(.031)
Employee has DB plan * 35–44 years old	0.002	0.004
	(.054)	(.047)
Employee has DB plan * 45–55 years old	0.118*	0.169**
	(.058)	(.053)
Employee has DC plan * 35–44 years old	0.051	0.054
	(.058)	(.051)
Employee has DC plan * 45–55 years old	0.023	0.136*
	(.063)	(.055)
Health insurance variables		
Participate in ESI	-0.785***	-0.900***
	(.209)	(.191)
Spouse covered by ESI	-0.047	0.033
	(.033)	(.029)
ESI coverage * tax price of health insurance	0.292	0.593**
	(.206)	(.188)
ESI coverage * excellent or very good health	0.023	-0.006
	(.050)	(.048)
ESI coverage * fair or poor health	0.138	-0.119
	(.097)	(.098)
ESI coverage * lower-earning spouse	0.125**	0.030
	(.047)	(.068)

Independent variables	Female	Male
Participate in ESI * 35–44 years old	0.005	0.009
	(.053)	(.051)
Participate in ESI * 45–55 years old	0.033	0.092
	(.060)	(.060)
Employer pays all of ESI	0.288**	0.163*
	(.085)	(.073)
Employer pays some of ESI	0.282**	0.177*
	(.083)	(.072)
Employer contributions to ESI unknown	0.253*	0.109
	(.099)	(.093)
Job variables		
Log of annual earnings	-0.152***	-0.108***
	(.034)	(.029)
Job tenure	-0.038***	-0.032***
	(.002)	(.002)
Job tenure unknown	-0.972***	-0.780***
	(.200)	(.130)
Union member	-0.167***	-0.053*
	(.033)	(.027)
Demographics		
Age: 35–44 years old	-1.295**	-0.777*
	(.403)	(.360)
Age: 45–55 years old	-1.799***	-1.712***
	(.446)	(.385)
Number of children age 18 or younger	0.013	-0.034***
	(.010)	(.009)
Number of children age 2 or younger	-0.060	-0.049
	(.035)	(.026)
Marriage variables		
Married	-0.036	0.013
	(.031)	(.029)
Lower-earning spouse	-0.059	0.022
	(.041)	(.055)
Race		
Hispanic	0.007	-0.059
	(.039)	(.034)
Black, non-Hispanic	0.030	0.021
	(.033)	(.035)
Non-white, non-black, non-Hispanic	0.015	0.028
	(.053)	(.048)

(continued)

Table 4 (continued)

Independent variables	Female	Male
Health status		
Good	0.025	-0.019
	(.040)	(.041)
Fair or poor	0.127	0.277 ***
	(.073)	(.077)
Missing	1.210 ***	1.129 ***
	(.100)	(.087)
Income		
Log of after-tax income	-0.145 ***	-0.136 ***
	(.021)	(.024)
Income * 35–44 years old	0.120 **	0.067
	(.042)	(.036)
Income * 45–55 years old	0.162 ***	0.151 ***
	(.046)	(.039)
Fraction of sample that changed jobs	0.222	0.220
Number of observations	20,606	25,979
Number of parameters	65	65
Log likelihood	-9,861	-12,266
Significance level for IRA variables ~ c^{2}**(4)**	0.3553	0.4911

Note: Same right-hand side variables as in augmented take-up model in table 3. The specifications also control for education, race, firm size, occupation, industry, urban residence, and year.

* statistically significant at the 5% level
** statistically significant at the 1% level
*** statistically significant at the 0.1% level

Table 5. Marginal Effects on the Probability of Job Change by Marital Status, Assuming Non-IRA Savings Are Taxed at Capital Gains Tax Rates

Independent variables	Married	Single
Pension variables		
rt_s * employer contributes toward pension	-0.333	21.175*
	(2.287)	(8.327)
rt_s * employer does not contribute	-5.013	929.234
	(7.184)	(681.077)
PV_{IRA} * employer contributes toward pension	3.07E-05	-2.79E-04*
	(5.23E-05)	(1.41E-04)
PV_{IRA} * employer does not contribute	3.14E-04	-5.36E-02
	(1.81E-04)	(4.09E-02)
Pension offered by employer	0.054	0.025
	(.030)	(.037)
Employer contributes to pension plan	-0.231 ***	-0.256 ***
	(.046)	(.067)
IRA contributions unknown	-0.201 *	-0.110
	(.099)	(.167)
Covered by DB plan	-0.047	-0.010
	(.042)	(.059)
Participates in DC plan	-0.059	0.019
	(.041)	(.056)
Spouse has pension plan	-0.020	…
	(.023)	
Employee has DB plan * 35–44 years old	-0.013	-0.008
	(.043)	(.064)
Employee has DB plan * 45–55 years old	0.131 **	0.087
	(.047)	(.077)
Employee has DC plan * 35–44 years old	0.032	0.020
	(.047)	(.070)
Employee has DC plan * 45–55 years old	0.086	-0.055
	(.050)	(.081)
Health insurance variables		
Participate in ESI	-0.675 ***	-1.352 ***
	(.171)	(.243)
Spouse covered by ESI	0.011	…
	(.022)	
ESI coverage * taxprice of health insurance	0.281	1.061 ***
	(.166)	(.244)
ESI coverage * excellent or very good health	0.038	-0.088
	(.042)	(.061)
ESI coverage * fair or poor health	0.035	-0.027
	(.086)	(.113)
ESI coverage * lower-earning spouse	0.064	…
	(.039)	

(continued)

Table 5 (continued)

Independent variables	Married	Single
Participate in ESI * 35–44 years old	0.029	-0.048
	(.045)	(.068)
Participate in ESI * 45–55 years old	0.088	0.038
	(.050)	(.082)
Employer pays all of ESI	0.179 *	0.316 ***
	(.070)	(.091)
Employer pays some of ESI	0.178 *	0.329 ***
	(.069)	(.089)
Employer contributions to ESI unknown	0.162 *	0.205
	(.082)	(.117)
Job variables		
Log of annual earnings	-0.161 ***	-0.061
	(.028)	(.034)
Job tenure	-0.033 ***	-0.038 ***
	(.001)	(.002)
Job tenure unknown	-0.876 ***	-0.835 ***
	(.141)	(.173)
Union member	-0.095 ***	-0.089 *
	(.025)	(.037)
Demographics		
Female	-0.116 ***	-0.074 **
	(.023)	(.075)
Age: 35–44 years old	-0.919 **	-0.981 *
	(.332)	(.445)
Age: 45–55 years old	-2.241 ***	-1.075 *
	(.354)	(.507)
Number of children age 18 or younger	-0.015 *	-0.009
	(.008)	(.016)
Number of children age 2 or younger	-0.047 *	-0.051
	(.022)	(.075)
Marriage variables		
Lower-earning spouse	-0.021	…
	(.033)	
Race		
Hispanic	-0.028	-0.045
	(.031)	(.045)
Black, non-Hispanic	0.033	0.020
	(.032)	(.035)
Non-white, non-black, non-Hispanic	0.048	-0.010
	(.042)	(.067)

Independent variables	Married	Single
Health status		
Good	0.024	-0.068
	(.034)	(.052)
Fair or poor	0.219 **	0.132
	(.066)	(.090)
Missing	1.238 ***	1.014 ***
	(.087)	(.101)
Income		
Log of after tax income	-0.124***	-0.153 ***
	(.019)	(.027)
Income * 35–44 years old	0.082*	0.087
	(.034)	(.046)
Income * 45–55 years old	0.206***	0.085
	(.036)	(.053)
Fraction of sample that changed jobs	0.203	0.258
Number of observations	31,551	15,034
Number of parameters	65	63
Log likelihood	-14,494	-7,661
Significance level for IRA variables ~ χ^2(4)	0.506	0.0676

Note: Same right-hand side variables as in augmented take-up model in table 3. The specifications also control for education, race, firm size, occupation, industry, urban residence, and year.

… = excluded variable

* statistically significant at the 5% level
** statistically significant at the 1% level
*** statistically significant at the 0.1% level

Table 6. Marginal Effects on the Probability of Job Change Excluding Missings and Using a Tax Rate Derived from SIPP Responses

Independent variables	Excluding missings ($t_s = t_g$)	SIPP response tax rate ($t_s = t_g$)
Pension variables		
rt_s * employer contributes toward pension	-0.455	-0.141
	(1.922)	(1.777)
rt_s * employer does not contribute	-11.705	-15.160 *
	(6.991)	(6.347)
PV_{IRA} * employer contributes toward pension	-5.19E-05	-3.81E-05
	(3.98E-05)	(3.49E-05)
PV_{IRA} * employer does not contribute	9.68E-05	1.70E-04
	(1.56E-04)	(1.49E-04)
Pension offered by employer	0.034	0.038
	(.024)	(.023)
Employer contributes to pension plan	-0.225 ***	-0.222 ***
	(.027)	(.026)
IRA contributions unknown	…	-0.183 *
		(.081)
Health insurance variables		
Participate in ESI	-0.340 ***	-0.394 ***
	(.087)	(.082)
Spouse covered by ESI	-0.024	-0.023
	(.020)	(.020)
ESI coverage * tax price of health insurance	0.093	0.169
	(.119)	(.113)
Job variables		
Log of annual earnings	-0.054 ***	-0.052 **
	(.015)	(.015)
Job tenure	-0.033 ***	-0.033 ***
	(.001)	(.001)
Job tenure unknown	…	-0.825 ***
		(.114)
Union member	-0.097 ***	-0.095 ***
	(.021)	(.020)

Independent variables	Excluding missings ($t_s = t_g$)	SIPP response tax rate ($t_s = t_g$)
Demographics		
Female	-0.114***	-0.111***
	(.018)	(.017)
Age: 35–44 years old	-0.112***	-0.109***
	(.017)	(.017)
Age: 45–55 years old	-0.078***	-0.074***
	(.020)	(.019)
Urban resident	0.136***	0.137***
	(.018)	(.017)
Number of children age 18 or younger	-0.010	-0.009
	(.007)	(.007)
Number of children age 2 or younger	-0.064**	-0.064**
	(.022)	(.020)
Marriage variables		
Married	-0.029	-0.025
	(.020)	(.019)
Lower-earning spouse	0.006	0.010
	(.023)	(.022)
Race		
Hispanic	-0.034	-0.035
	(.026)	(.026)
Black, non-Hispanic	0.023	0.007
	(.024)	(.023)
Non-white, non-black, non-Hispanic	0.021	0.021
	(.036)	(.035)
Education		
Less than a high school degree	-0.073*	-0.084**
	(.030)	(.029)
High school degree	-0.062**	-0.069**
	(.022)	(.022)
Some college	-0.020	-0.026
	(.021)	(.020)
Health status		
Good	-0.002	-0.006
	(.017)	(.017)
Fair or poor	0.190***	0.192***
	(.032)	(.031)
Missing	…	1.142***
		(.064)
Income		
Log of after-tax income	-0.134***	-0.131***
	(.015)	(.015)

(continued)

Table 6 (continued)

Independent variables	**Excluding missings ($t_s = t_g$)**	**SIPP response tax rate ($t_s = t_g$)**
Firm size		
Fewer than 25 people	0.008	0.010
	(.020)	(.019)
25–99 people	0.025	0.020
	(.022)	(.021)
Missing	…	-0.116 *
		(.048)
Occupation		
Sales	0.027	0.020
	(.029)	(.028)
Administrative and clerical	-0.058 *	-0.062 **
	(.024)	(.023)
Service, mechanics, and construction	-0.029	-0.033
	(.024)	(.023)
Craft workers and machine operators	-0.030	-0.018
	(.030)	(.029)
Transportation	-0.017	-0.021
	(.042)	(.040)
Laborers	-0.058	-0.050
	(.041)	(.040)
Occupation unknown	…	-0.068
		(.099)
Industry		
Agriculture, mining, construction	0.190 ***	0.190 ***
	(.030)	(.029)
Manufacturing	0.039	0.025
	(.024)	(.023)
Transportation, public utilities, communication	0.030	0.022
	(.030)	(.029)
Wholesale and retail trade, FIRE	0.033	0.025
	(.021)	(.021)
Public administration	-0.068 *	-0.051
	(.033)	(.031)
Industry unknown	…	0.116
		(.086)

Independent variables	Excluding missings ($t_s = t_g$)	SIPP response tax rate ($t_s = t_g$)
Panel year		
1984	0.207***	0.220***
	(.024)	(.030)
1990	0.059**	-0.039
	(.021)	(.035)
1996	0.101***	0.046
	(.019)	(.030)
Tax rate from tax calculator		
1990	…	0.146***
		(.031)
1992	…	0.056
		(.029)
1996	…	0.099***
		(.024)
Intercept	1.628***	1.549***
	(.171)	(.167)
Fraction of sample that changed jobs	0.218	0.215
Number of observations	43,853	51,635
Number of parameters	50	53
Log likelihood	-21,003	-22,332
Significance level for IRA variables ~ χ^2(4)	0.3115	0.1359

… = excluded variable

* statistically significant at the 5% level
** statistically significant at the 1% level
*** statistically significant at the 0.1% level

Figure 1. IRA Budget Constraint and Utility

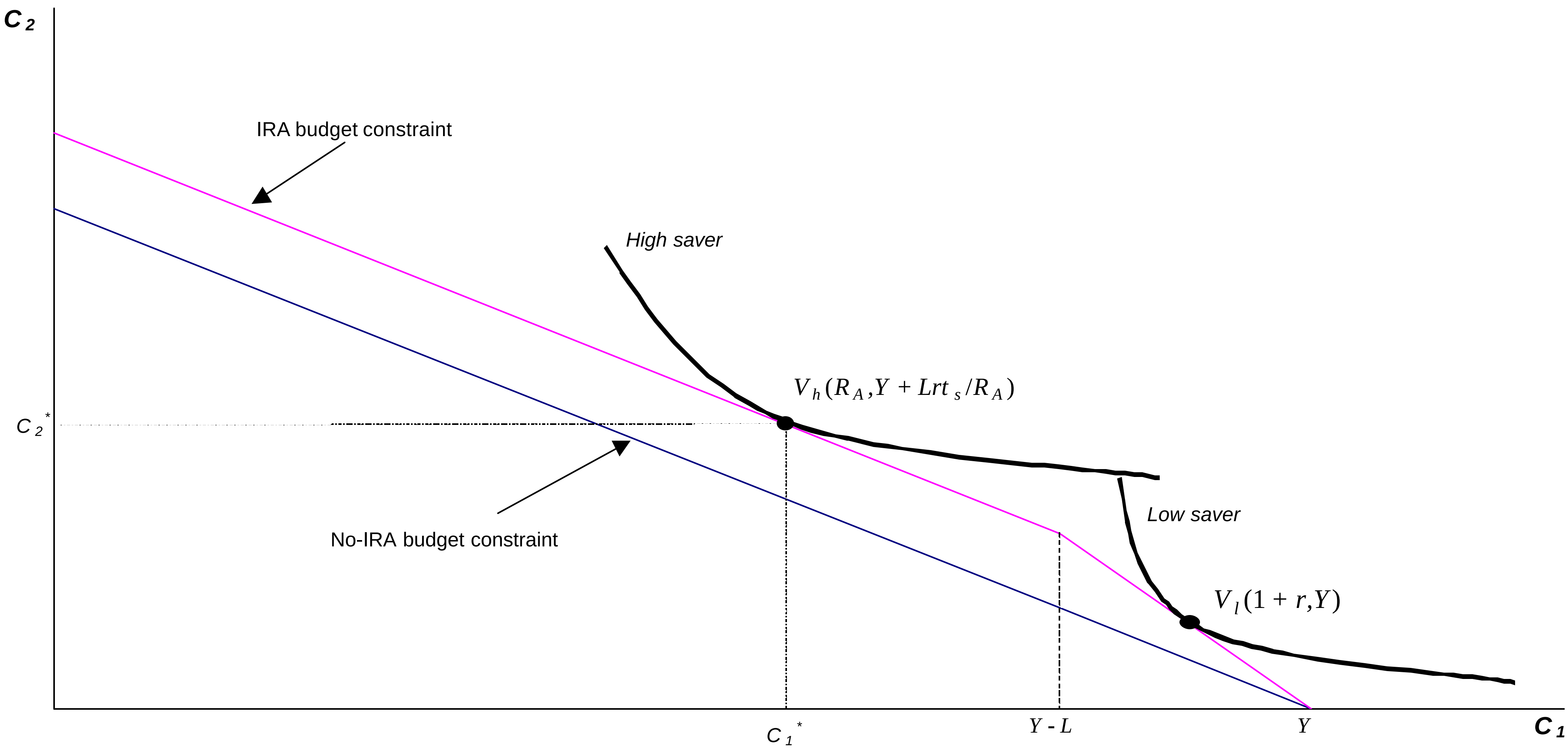
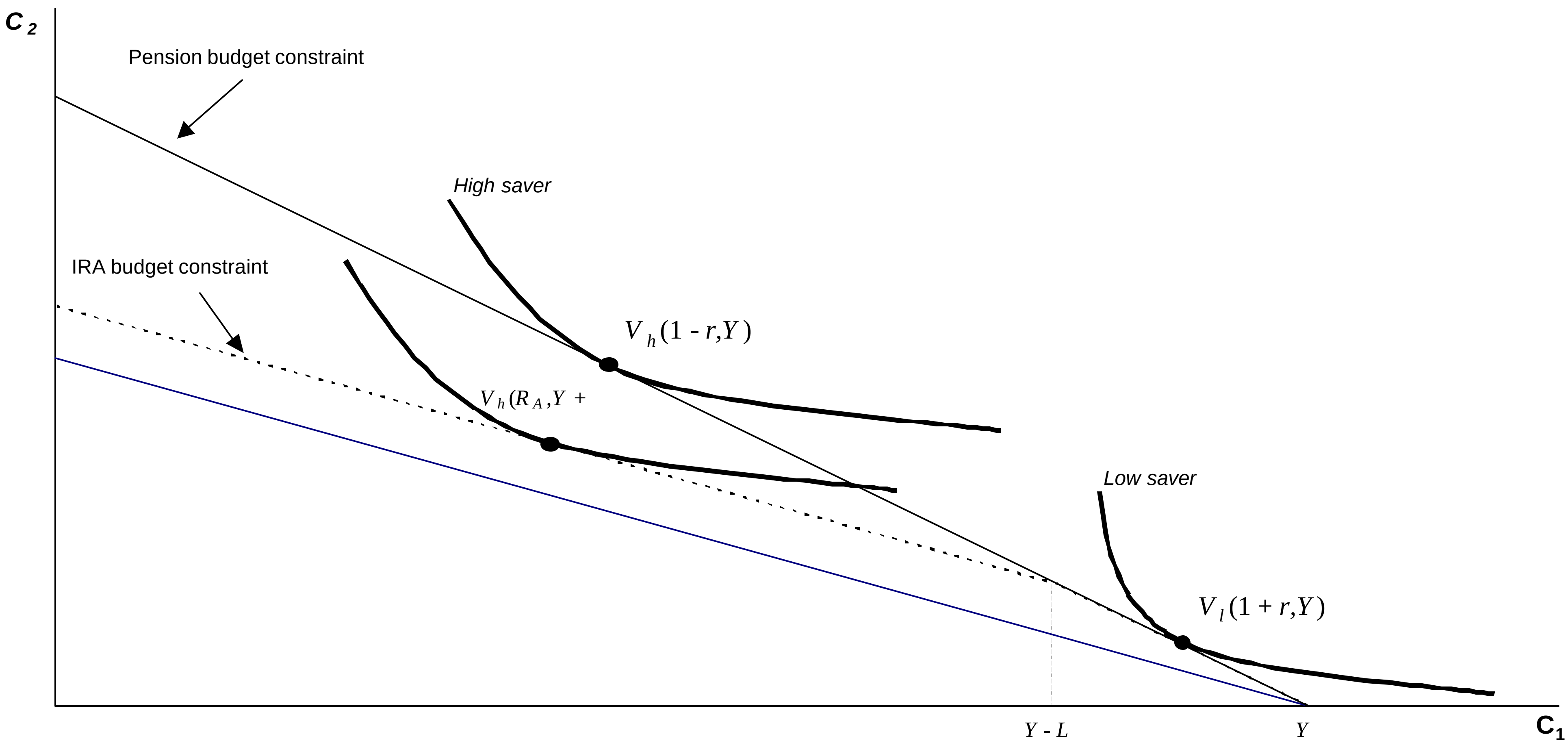


Figure 2. Valuation of Pension Tax Benefits



Appendix 1. Sensitivity of Results to Job Change Measure

As discussed in the text, a potentially serious problem arises if job changing is measured with error. An additional complication is that our analysis combines involuntary job changers with voluntary movers. This appendix examines data on reasons for job change that are available for a subset of our sample to test the validity of our estimation results.

The SIPP data include direct questions about whether respondents changed jobs between waves and, if so, why, for 1990 and later panels. Unfortunately, those data do not exist in 1984, the only year in our sample before IRA eligibility rules changed in 1986. In addition, it is unclear how reliable the reason-for-change variables are.

Table A1 compares the answers to the direct question on job change asked in the 1990, 1992, and 1996 SIPP panels with the indirect measure of job change that we used in estimation for the same three cross sections. The pair of questions asked in these panels address employer change specifically, and not only job change within a single firm: "Last time we recorded that you worked for [Employer name]. Do you still work for [Employer name]?" and "What is the main reason you stopped working for [Employer name]?"[28] Based on the "reason for job change" variable, it appears that about one-third of separations were involuntary, which could reduce the power of our results. Perhaps more importantly, the two measures are not very consistent. More than 1,000 respondents who reported working for the same employer in waves 4 and 7 apparently reported that they had voluntarily or involuntarily changed jobs between the waves. And almost 8,000 respondents reported no reason for a job change, which SIPP codes as no change in employers, even though they separately told the interviewer that their current employer is different from the employer in a previous wave.

The alternate measure of job change (based on the "reason to change jobs") appears to be less reliable than the measure we used in estimation. Based on the alternate measure, only 10.3 percent of respondents changed jobs in a year, but other surveys suggest that about 20 to 24 percent of workers leave their jobs each year.[29] Based on our job change variable, 22.7 percent of workers changed jobs in the three SIPP samples, and the percentage is fairly stable—ranging from 22.2 to 23.9 percent—across samples. However, the estimate of involuntary separation is roughly consistent with Farber (2003), who finds that about 10 percent of workers lose their jobs over a three-year period, or about 3 percent a year.

We used the data on reason for job change to see whether involuntary job changers behave differently from voluntary changers, and also to test the extent of bias created by failure to exclude involuntary movers. Table A2 compares estimates for the key parameters using the 1990–96 SIPP samples under our original measure of job changers, and using the alternative definition based on the reported reason for job change.

[28] These questions are from the 1996 panel, wave 2, Labor Force Section Part 1, but are consistent with the 1990 and 1992 panels.

[29] Anderson and Meyer (1994) report a turnover rate of 23 percent. Davis and Haltiwanger (1999) report a rate of about 20 percent.

For reference purposes, the first column reproduces the results reported in table 3 for specification 1 using the full sample, including 1984. The second column shows the parameter estimates for the same data and specification, but excluding 1984. The results are similar, although the estimate for the first IRA variable becomes statistically significant and takes on the sign predicted by the theoretical model. Collectively, the estimates for the IRA variables become marginally significant. The estimated probability that the IRA variables are collectively zero falls from 0.51 to 0.07.

If those identified as involuntary movers are excluded from the sample, the parameter estimates are not significantly different. Collectively, the significance level for the IRA variables changes from 0.071 to 0.075. Thus, it appears that including in the sample respondents whom SIPP identifies as involuntary movers does not markedly alter the parameter estimates or conclusions about statistical significance.

The last two columns show parameter estimates based on the alternative measure of job changers—those who identify themselves as voluntary or involuntary movers. The results are consistent with the conclusion that this variable is measured with substantial error. The IRA variables become much less statistically significant. The *p*-level increases to 0.56 if movers include both voluntary and involuntary movers. If involuntary movers are eliminated from the sample, the *p*-level increases further to 0.81, although the point estimates are not markedly different.

Overall, these results support the definition of job change as the best available option using the SIPP data.

Table A1. Comparison of Different Job Change Variables in SIPP, 1990–96

Value of job change measure used in estimation	Reason for job change		Did not change	All	Percent
	Voluntary	Involuntary			
Change jobs	2,641	1,609	*7,881*	12,131	22.7
No change	*799*	*440*	39,992	41,231	77.3
All	3,440	2,049	47,873	53,362	100.0
Percent	6.4	3.8	89.7	100.0	

Note: Inconsistent responses in italics.

Table A2. Sensitivity of Main Results to Definition of Mover, Based on Table 3, Specification 1

		Sample excluding 1984			
				Alt. move definitions	
Independent variables	Full sample	All movers	Exclude involuntary	All movers	Exclude involuntary
rt_s * employer contributes toward pension	1.20	-4.50	-4.65	0.69	0.76
	(1.84)	(1.94)	(2.02)	(2.46)	(2.86)
rt_s * employer does not contribute	-11.11	-12.01	-11.87	-13.19	-12.44
	(6.83)	(6.82)	(7.05)	(9.14)	(10.71)
PV_{IRA} * employer contributes toward pension	-1.83E-05	-1.40E-05	-2.02E-05	3.51E-05	1.28E-05
	(3.69E-05)	(3.88E-05)	(4.05E-05)	(4.43E-05)	(5.12E-05)
PV_{IRA} * employer does not contribute	1.38E-04	1.52E-04	1.30E-04	1.47E-04	1.24E-04
	(1.49E-04)	(1.50E-04)	(1.56E-04)	(1.85E-04)	(2.15E-04)
Pension offered by employer	0.02	0.02	0.02	0.06	0.05
	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)
Employer contributes to pension plan	-0.25	-0.20	-0.19	-0.23	-0.20
	(0.03)	(0.03)	(0.03)	(0.03)	(0.04)
Participate in ESI	-0.38	-0.20	-0.12	-0.36	-0.22
	(0.08)	(0.09)	(0.10)	(0.11)	(0.13)
Log of annual earnings	-0.03	-0.07	-0.04	-0.13	-0.08
	(1.40E-02)	(1.51E-02)	(1.61E-02)	(1.75E-02)	(2.05E-02)
Job tenure	-0.03	-0.03	-0.03	-0.03	-0.03
	(1.18E-03)	(1.26E-03)	(1.31E-03)	(1.65E-03)	(1.92E-03)
Log likelihood	-22,347	-18,574	-16,782	-12,069	-8,528
Number of observations	46,585	38,905	37,291	38,905	37,291
Fraction of sample moving	.2211	.2161	.1919	.1062	.0675
Significance level for IRA variables ~ $\chi^2(4)$	0.5068	0.0710	0.0747	0.5627	0.8107

Appendix 2. Estimation of Desired IRA Contributions

In the theoretical model, the tax value of pension benefits depends on whether an individual would want to save more than the deductible IRA limit. We estimate the likelihood of contributing the maximum amount to an IRA for those who contribute at all, and then apply the results to everyone who saves for retirement either through an IRA or a pension plan. The likelihood of being at the IRA limit is a function of demographic variables and total liquid assets. Liquid assets include savings and checking accounts, stocks, U.S. savings bonds, and other interest-earning financial investments. The estimation is split according to gender. Because the size of IRA contributions is not reported for the 1984 SIPP, the sample is limited to 1990, 1992, and 1996 observations.

We estimate a censored regression model where the dependent variable, log of IRA contributions, is censored from above. (Recall that the sample is restricted to people who contribute to an IRA.) Estimation is by maximum likelihood, assuming that the error term is normally distributed. The upper limit is variable, because an individual's IRA limit depends upon contribution rules for the sample year, an individual's income, and whether the spouse is covered by a pension plan.

Based on parameter estimates from this model, the probability of contributing the limit (p) may be calculated as $p_i = P(Y_i > L_i) = P(X_i ß + e_i > L_i) = P(e_i > L_i - X_i ß)$, where $P(\cdot)$ signifies probability, Y_i is the dependent variable (contribution level) for individual i, L_i is the IRA limit, X_i is the vector of independent variables, $ß$ is the regression coefficient vector, and e is the equation error term (assumed to be distributed normally with mean 0 and variance s ^{2}). This term may be consistently estimated as $p_i = 1 - F((L_i - X_i b)/s)$, where b is the estimated parameter vector and s is the estimated standard error from the censored regression, and $F(\cdot)$ is the standard normal distribution function.

The estimated parameters from the censored regressions are shown in tables A3 and A4. Table A4 shows the distribution of predicted probabilities of limit contribution by gender, marital status, and year based on the estimated parameters and the values of the independent variables in each year.

Table A3. Females, Two-Limit Tobit Model with Variable Upper Limit

Independent variables	Coefficient	Standard error	*t*	*P* > |*t*|	95% confidence interval	
Income						
Liquid assets	0.0016	0.0005	3.00	0.003	0.0005	0.0026
Annual earnings (in $10,000s)	-170.144	70.561	-2.41	0.016	-308.952	-31.34
Annual earnings squared	13.3934	4.4394	3.02	0.003	4.6602	22.1266
Non-wage income (in $10,000s)	122.987	43.216	2.85	0.005	37.97	208.002
Non-wage income squared	-0.0121	0.0142	-0.85	0.397	-0.0400	0.0159
Demographics						
Age	-232.528	84.415	-2.75	0.006	-398.59	-66.466
Age squared	2.8970	1.0464	2.77	0.006	0.8386	4.9554
Married	-380.51	178.69	-2.13	0.034	-732.03	-28.98
Race						
Hispanic	-404.73	345.44	-1.17	0.242	-1,084.3	274.82
Black	-1451.3	461.35	-3.15	0.002	-2,358.8	-543.68
Non-white, non-black, non-Hispanic	362.83	308.86	1.17	0.241	-2,44.76	970.4
Education						
Less than high school degree	232.96	539.67	0.43	0.666	-828.68	1,294.59
Some college	192.95	185.70	1.04	0.300	-172.36	558.26
At least 4 years college	84.52	195.28	0.43	0.665	-299.64	468.68
Health status						
Unknown	511.1	1,103.133	0.46	0.643	-1,659.032	2,681.138
Very good	496.65	164.35	3.02	0.003	173.35	819.96
Good	120.2082	197.81	0.61	0.544	-268.92	509.33
Fair	-2.56	415.94	-0.01	0.995	-820.8	815.68
Poor	-593.92	817.42	-0.73	0.468	-2,202.0	1,014.11

(continued)

Table A3 (continued)

Independent variables	Coefficient	Standard error	*t*	*P* > |*t*|	95% confidence interval	
Firm size						
Fewer than 25 people	191.680	155.03	1.24	0.217	-113.29	496.65
25–99 people	96.202	207.84	0.46	0.644	-312.66	505.07
Occupation						
Managers and technicians	324.54	254.31	1.28	0.203	-175.7	824.826
Sales	-24.46	176.86	-0.14	0.890	-372.4	323.45
Administrative and clerical	-67.31	273.03	-0.25	0.805	-604.4	469.7912
Service, mechanics, and construction	-130.58	394.27	-0.33	0.741	-906.2	645.03
Craft workers and machine operators	-1,440.64	1,166.67	-1.23	0.218	-3,735.7	854.44
Laborers	7620.2	…	…	…	…	…
Industry						
Manufacturing	1,064.8	1,128.2	0.94	0.346	-1,154.60	3,284.3
Transportation, public utilities, communication	1,158.1	1,133.5	1.02	0.308	-1,071.82	3,387.9
Wholesale and resale trade, FIRE	691.6	1,111.3	0.62	0.534	-1,494.644	2,877.8
Services and professional	887.5	1,102.3	0.81	0.421	-1,281.00	3,056.0
Intercept	5,546.4	1,960.7	2.83	0.005	1,689.4	9,403.4
Standard error	1,081.668	57.63726				
Number of observations	361					
Number of parameters	31					
Log likelihood	-1,729					
Pseudo R^2	0.0316					

Note: Some demographic variables were excluded because of collinearity.

Table A4. Males, Two-Limit Tobit Model with Variable Upper Limit

Independent variables	Coefficient	Standard error	*t*	*P* > |*t*|	95% confidence interval	
Income						
Liquid assets	0.0004	0.0033	0.13	0.895	-0.0061	0.0069
Annual earnings (in $10,000s)	276.1409	123.8209	2.23	0.026	32.5264	519.7555
Annual earnings squared	-2.07	4.936	-0.42	0.675	-11.7831	7.64
Non-wage income (in $10,000s)	551.5940	173.3091	3.18	0.002	210.6126	892.5755
Non-wage income squared	-0.42	0.2	-2.70	0.007	-0.7	-0.1
Demographics						
Age	-158.892	211.926	-0.75	0.454	-575.85	258.067
Age squared	1.7972	2.6004	0.69	0.490	-3.3190	6.9135
Married	-2,345.25	436.92	-5.37	0.000	-3,204.89	-1,485.61
Race						
Hispanic	1,032.91	984.35	1.05	0.295	-903.77	2,969.60
Black	-2,505.9	1,924.62	-1.30	0.194	-6,292.6	1,280.72
Non-white, non-black, non-Hispanic	-453.00	799.78	-0.57	0.572	-2,026.555	1,120.6
Education						
Less than high school degree	-845.84	835.97	-1.01	0.312	-2,490.59	798.92
Some college	-410.23	491.20	-0.84	0.404	-1,376.66	556.19
At least 4 years college	-605.96	508.58	-1.19	0.234	-1,606.58	394.66
Health status						
Very good	630.75	367.93	1.71	0.087	-93.13	1,354.6420
Good	-391.90	513.86	-0.76	0.446	-1402.90	619.1037
Fair	0.00	0.00	0.00	0.000	0.00	0.00

(continued)

Table A4 (continued)

Independent variables	Coefficient	Standard error	*t*	*P* > |*t*|	95% confidence interval	
Firm size						
Fewer than 25 people	-334.06	379.86	-0.88	0.380	-1,081.43	413.32
25–99 people	-115.36	490.16	-0.24	0.814	-1,079.74	849.03
Occupation						
Managers and technicians	338.56	534.23	0.63	0.527	-712.5	1,389.643
Sales	-1,063.27	798.53	-1.33	0.184	-2,634.4	507.83
Administrative and clerical	-845.9	574.05	-1.47	0.142	-1,975.3	283.5827
Service, mechanics, and construction	-1,608.9	672.57	-2.39	0.017	-2,932.2	-285.63
Craft workers and machine operators	-1,007.2	923.66	-1.09	0.276	-2,824.5	810.03
Laborers	-361.7	1,811.39	-0.20	0.842	-3,925.6	3,202.14
Industry						
Agriculture, mining, construction	1,418.64	2,107.17	0.67	0.501	-2,727.2	5,564.44
Manufacturing	742.5178	2,089.75	0.36	0.723	-3,369.01	4,854.04
Transportation, public utilities, communication	388.25	2,114.8	0.18	0.854	-3,772.6	4,549.14
Wholesale and resale trade, FIRE	203.528	2,087.6	0.10	0.922	-3,903.8	4,310.86
Services and professional	-65.945	2,098.1	-0.03	0.975	-4,193.92	4,062.03
Public administration	-345.44	2,229.96	-0.15	0.877	-4,732.8	4,041.95
Intercept	6,743.34	4,748.4	1.42	0.157	-2,599.0	1,6085.7
Standard error	2,502.684	73.031				
Number of observations	348					
Number of parameters	31					
Log likelihood	-1,904					
Pseudo R^2	0.0168					

Note: Some demographic variables were excluded because of collinearity.

Table A5. Estimated Probability at IRA Limit, by Gender, Marital Status, and Survey Year

Probability at limit	Gender		Marital status	
	Male	Female	Single	Married
0	42.3	45.1	54.6	38.2
0–0.2	16.4	13.0	13.4	15.6
0.2–0.4	14.5	20.0	15.7	17.5
0.4–0.6	10.6	12.2	9.4	12.3
0.6–0.8	7.3	2.3	4.4	5.4
0.8–1.0	8.9	7.5	2.5	11.0

Probability at limit	Year				All
	1984	1990	1992	1996	
0	39.3	46.8	41.1	44.5	43.5
0–0.2	18.2	11.5	12.8	17.0	14.9
0.2–0.4	20.5	12.9	14.5	19.6	17.0
0.4–0.6	13.1	10.3	11.0	11.5	11.3
0.6–0.8	6.2	5.3	5.3	4.4	5.1
0.8–1.0	2.8	13.4	15.4	2.9	8.3

Other Publications by the Tax Policy Center

Discussion Paper Series

Number	Title	Authors	Date
13	Searching for a Just Tax System	Rudolph G. Penner	December 2003
12	Tax Incentives for Health Insurance	Leonard E. Burman, Cori E. Uccello, Laura L. Wheaton, and Deborah Kobes	May 2003
11	State Fiscal Constraints and Higher Education Spending	Thomas J. Kane, Peter R. Orszag, and David L. Gunter	May 2003
10	Budget Blues: The Fiscal Outlook and Options for Reform	Alan J. Auerbach, William G. Gale, Peter R. Orszag, and Samara R. Potter	May 2003
9	Private Pensions: Issues and Options	William G. Gale and Peter Orszag	April 2003
8	The Economic Effects of Long-Term Fiscal Discipline	William G. Gale and Peter Orszag	April 2003
7	Charitable Bequests and Taxes on Inheritances and Estates: Aggregate Evidence from across States and Time	Jon Bakija, William G. Gale, and Joel Slemrod	April 2003
6	The Enron Debacle: Lessons for Tax Policy	Jane G. Gravelle	February 2003
5	The Individual AMT: Problems and Potential Solutions	Leonard E. Burman, William G. Gale, Jeffrey Rohaly, and Benjamin H. Harris	September 2002
4	Providing Federal Assistance for Low-Income Families through the Tax System: A Primer	Frank Sammartino, Eric Toder, and Elaine Maag	July 2002
3	An Economic Evaluation of the Economic Growth and Tax Relief Reconciliation Act of 2001	William G. Gale and Samara R. Potter	March 2002
2	How Marriage Penalties Change under the 2001 Tax Bill	Adam Carasso and C. Eugene Steuerle	May 2002
1	The Effect of the 2001 Tax Cut on Low- and Middle-Income Families and Children	Leonard E. Burman, Elaine Maag, and Jeff Rohaly	April 2002

Issues and Options Series

Number	Title	Author	Date
7	Promoting 401(k) Security	J. Mark Iwry	September 2003
6	Effects of Estate Tax Reform on Charitable Giving	Jon M. Bakija and William G. Gale	July 2003

Number	Title	Author	Date
5	The AMT: Out of Control	Leonard E. Burman, William G. Gale, Jeffrey Rohaly, and Benjamin H. Harris	September 2002
4	Saying “I Do” after the 2001 Tax Cuts	Adam Carasso and C. Eugene Steuerle	August 2002
3	EGTRRA: Which Provisions Spell the Most Relief?	Leonard E. Burman, Elaine Maag, and Jeff Rohaly	June 2002
2	The Estate Tax Is Down, But Not Out	Leonard E. Burman and William G. Gale	December 2001
1	Designing Tax Cuts to Benefit Low-Income Families	Frank Sammartino	June 2001