Employer Matching and 401(k) Saving: Evidence from the Health and Retirement Study

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Abstract

Employer matching of employee 401(k) contributions can provide a powerful incentive to save for retirement and is a key component in pension-plan design in the United States. Using detailed administrative contribution, earnings, and pension-plan data from the Health and Retirement Study, this analysis formulates a life-cycle-consistent two-limit censored regression model of 401(k) saving and estimates the effect of matching on 401(k) saving accounting for non-linearities in the household budget set induced by matching. Parametric and semi-parametric estimates indicate that an increase in the match rate by 25 cents per dollar of employee contribution raises 401(k) saving by $500-$800 (in 1991 dollars), and the estimated elasticity of contributions with respect to matching ranges from 0.06-0.17 overall, with two-thirds of this effect on the extensive margin and one-third on the intensive margin.

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I. Introduction

As 401(k)s have come to dominate the private pension landscape in the United States, researchers and policy makers have given increased attention to the impact of plan characteristics on retirement-saving decisions.[1] One important characteristic is whether and to what extent the employer matches employee contributions. A typical match might be 50 cents for each dollar of contribution, up to a maximum percentage of pay, say, 6 percent. Although much of the discussion by the popular press and policy makers presumes employer matching raises saving, there is actually strikingly little consensus among researchers. Some studies have found that increases in the match rate raise 401(k) saving (Papke and Poterba, 1995; Clark and Schieber, 1998; VanDerhei and Copeland, 2001; and Choi, Laibson, Madrian, and Metrick, 2002). Others have found that it is not the match rate *per se* that matters, but whether the firm offers a match at all (Even and Macpherson, 1996; Bassett, Fleming, and Rodrigues, 1998; Papke, 1995; Kusko, Poterba, and Wilcox, 1998). That is, providing a match raises 401(k) saving, but an increase in the level of the match rate (conditional on providing a match) does not. Finally, still other studies (Employee Benefit Research Institute, 1994; Andrews, 1992; Munnell, Sunden, and Taylor, 1998; and GAO, 1997) have suggested that, conditional on being eligible for a match, an increase in the match rate *lowers* 401(k) contributions, which, when interpreted in the context of a simple two-period model of saving, suggests that the income effect dominates the substitution effect from the higher rate of return matching

[1] This includes work on automatic enrollment (Madrian and Shea, 2001; Choi, Laibson, Madrian, and Metrick, 2002, 2004), investment in company stock (Poterba, 2003; Brown, Liang, and Weisbenner, 2004; Mitchell and Utkus, 2002), portfolio choice and trading in 401(k) plans (Benartzi and Thaler, 2001; Agnew, Balduzzi, and Sunden, 2003).

provides.[2] Overall, this ambiguity has emerged as an important empirical puzzle in the literature on saving behavior (Bernheim, 2003).

A central shortcoming in this literature has been the failure to exploit the fact that employer matching, based either on multiple-match-rate schedules or caps on the generosity of the match, results in a non-linear household budget set.[3] As has been long recognized in the study of taxation on labor supply, reduced-form estimates of behavioral elasticities are biased and inconsistent unless the non-linearity is accounted for explicitly (Hausman, 1985; Moffitt, 1990; Blundell and MaCurdy, 1999). Indeed, the presence of budget-set kinks may reconcile some of the findings of previous studies: for example, the provision of a match may raise 401(k) saving if the substitution effect dominates, but variation in match rates may not matter if employees are bunched at kinks.

Unlike previous studies, this paper includes a detailed theoretical framework that models the budget set defined by employer matching and federal tax treatment as twice continuously differentiable and then uses the first-order conditions from the consumer's optimization to derive a life-cycle-consistent two-limit censored regression model of 401(k) saving. As an alternative to the maximum-likelihood piecewise-linear-budget-set estimation summarized in Hausman (1985)—and the recent, related non-parametric extensions by Blomquist and Newey (2002)—and the maximum likelihood differentiable-budget-constraint methodology of MaCurdy, Green, and Paarsch (1990), this paper employs instrumental-variable techniques that linearize the budget set at the observed outcome to calculate the price and virtual-income terms and then instruments to

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[2] Throughout the paper, "401(k) saving" and "401(k) contributions" are used synonymously as per period flows. In a multi-period model, this would suggest the income effect dominates the substitution and wealth effects (Summers, 1981).

[3] The two exceptions are the recent reduced-form studies by Choi, Laibson, Madrian, and Metrick (2002) and VanDerhei and Copeland (2001) that we discuss below.

correct for endogeneity, where the instrumental-variable Tobit estimator of Newey (1986, 1987b) and an instrumental-variable symmetrically censored least squares (SCLS) estimator based on Powell (1986) and Newey (1986) are used. To calculate budget-set slopes and virtual income in a neighborhood around kink points, kernel regression is used to smooth the budget set non-parametrically.

Empirically, the paper makes three additional contributions. First, to circumvent difficulties with measurement error in 401(k) contributions and matching incentives that have plagued previous studies, administrative data from three sources are used: contributions from W-2 earnings records provided by the Social Security Administration (SSA) and Internal Revenue Service (IRS); detailed matching formulas from pension Summary Plan Descriptions (SPD) provided by the employers of Health and Retirement Study (HRS) respondents; and, a combination of Social-Security-covered-earnings histories for 1951-1991 and W-2 earnings for 1980-1991, pension SPDs, and pension-benefit calculators to construct public and private pension entitlements and accruals. The sample consists of 1,042 individuals in 1991 eligible for 401(k) plans in the HRS. Second, the analysis includes a calculation of the dollar amount of unused employer matching contributions due to workers' failure to contribute at least until the point at which the employer match is exhausted. Most of this occurs because of non-participation. For non-participants, the unclaimed employer match represented 3.7 percent of pay. However, even *participants* left "money on the table" equal to 1 percent of pay in unclaimed employer match. Based on measures of liquidity constraints used by others in the literature, reduced-form evidence is not inconsistent with the presence of liquidity constraints as a potential explanation for this phenomenon. Finally, unlike previous pension studies that have used the employer-provided SPDs in the HRS, which are available only for a non-random sub-sample of HRS respondents, the estimation uses methods laid out in Vella (1992) and Das, Newey, and Vella (2003) for the Tobit and SCLS estimators, respectively, to correct for potential sample selection using a set of plausible exclusion restrictions derived from Internal Revenue Service (IRS) Form 5500 administrative pension-plan filings. The exclusions have substantial predictive power for determining who is in the analysis sample, and there is statistically significant evidence of selection. However, the economic impact of the selection on the estimates is small.

The parametric and semi-parametric estimates from the life-cycle-consistent two-limit censored regression specifications indicate that the estimated marginal effect of an increase in the employer match rate by 25 cents per dollar of employee contribution raises 401(k) contributions by $500-$800 dollars (in constant calendar year 1991 dollars). Comparing the Tobit and SCLS estimates using the Hausman-type test in Newey (1987a), the validity of the Tobit model cannot be rejected. When the Tobit estimates are expressed in terms of elasticities, the results suggest that the impact of the match rate on 401(k) saving is quite inelastic: the estimated elasticity of 401(k) saving with respect to the match rate ranges from 0.06-0.17 overall, with two-thirds of this effect on the extensive (participation) margin and the remaining one-third on the intensive (contributions conditional on participation) margin.

The paper is organized as follows. Section II lays out the theoretical model that directly motivates the empirical analysis. Section III lays out the econometric framework and construction of the key variables. Sections IV and V describe the data. Section VI describes the empirical analysis of the relationship between matching and measures of liquidity constraints. Section VII discusses the estimation and identification. Section VIII presents the estimation results. There is a brief conclusion.

II. Theoretical Framework

Previous studies have had two important shortcomings. First, they have not couched their analyses in formal models of intertemporal choice, even though saving involves the substitution of resources across time. This means that previous estimates cannot be interpreted as estimates of life-cycle-consistent determinants of 401(k) saving necessarily, because the empirical specifications may not have been consistent with underlying utility maximization. So, while the existing literature has provided quite informative descriptive analyses, it has said little about how 401(k) saving may respond to prospective changes in employer matching or what the optimal match rate should be to achieve a saving target. Second, with the exception of Choi, Laibson, Madrian, and Metrick (2002) and VanDerhei and Copeland (2001), previous studies have failed to exploit the fact that multiple-match-rate schedules and caps on matching induce kinks in the budget set.[4]

To illustrate the kink, Figure 1 shows budget sets with and without matching in a simple two-period model of consumption (without borrowing) typically used in

[4] Both of these studies examined the impact of match rates and caps on 401(k) participation and contributions, but from rather different perspectives than in this paper. Specifically, Choi, Laibson, Madrian, and Metrick (2002) examined large samples of employee administrative data from two firms, one that introduced a match of 25 percent up to 4 percent of pay and another that held the match rate constant but increased the match cap. Their reduced-form estimates were not inconsistent with a relatively moderate substitution effect and a very small income effect. Copeland and VanDerhei (2001) estimated the impact of match rates and caps using a reduced-form sequential-response regression model for 137 matching formulas and a very large sample of participants, in which a separate equation was estimated for each match rate the participant faced. They, too, found evidence that matching raised participation and contributions. Probably the biggest difference between these studies and the current paper is that the analysis below is based on a life-cycle-consistent empirical specification, constructs the household's budget set (rather than the employee), and incorporates virtual income into the estimation. A companion paper, Engelhardt and Kumar (2003), provided a further discussion of the previous literature.

undergraduate textbooks. Let C_1 and C_2 be consumption in period 1 when working and period 2 when retired, respectively, and y_1^l be first-period gross labor earnings. For simplicity, assume federal marginal income tax rates, θ_1^F, in the first period, and θ_2^F, in the second period, and no labor earnings in the second period. Let W^{SS} be the present value of traditional Social Security wealth, r be the gross return, and assume that Social Security is not taxed when retired. Assume that the 401(k) is the only form of voluntary saving, and contributions, Q^{VOL}, measured in dollars from right to left on the horizontal axis, are matched at a fixed rate, m^{VOL}, up to a cap, Q^{match}, based on the amount of the voluntary contributions. The upper limit on voluntary contributions is $Q^{\max}$. The budget constraint without matching is $\overline{abcd}$. With matching, it is $\overline{abkhi}$, and from 0 to Q^{match} (reading right to left, beginning at point a), the slope is $-(1-\theta_2^F)(1+m^V)(1+r)/(1-\theta_1^F)$, from Q^{match} to $Q^{\max}$, the slope is $-(1-\theta_2^F)(1+r)/(1-\theta_1^F)$, and beyond $Q^{\max}$, the slope is zero. At the match cap, Q^{match}, there is a kink point, k, at which the match is exhausted.

The introduction of the employer match with a cap has differential effects depending upon the location on the budget set. For low levels of contributions, the budget set rotates outward, and there are opposing income and substitution effects, but for higher levels of contributions, the budget set shifts out in parallel, and there is only an income effect, for which contributions fall with the imposition of the match. Furthermore, once the match cap is in place, changes in the match rate, m^V, may not induce changes in contributions if individuals are bunched at the kink point, k, and standard income and substitution effects are not well defined. Instead, income and substitution effects on each budget segment are defined by the slopes given above and the virtual incomes shown in the figure, y^{v_1} and y^{v_2}, respectively.

Motivation

By specifying a detailed theoretical framework, this paper represents a stark departure from the previous literature. There were eight criteria that guided the model detailed below. First, it should incorporate non-linear budget sets explicitly. In particular, the budget sets individuals actually face are substantially more complicated than the ones depicted in Figure 1, because they may have multiple kinks due to variable-rate matches, and there may be multiple kinks because there are multiple marginal tax rates, and contributions are tax-deductible, so that making a contribution may change the marginal tax rate. Surprisingly, none of the previous studies even have accounted for the effect of taxation on 401(k) saving. Second and third, respectively, the model should allow for goods other than consumption, such as leisure, that enter utility—because a change in the match rate may induce intratemporal substitution across goods—and for uncertain lifetimes and bequests, facets that are important to individuals of the ages found in the HRS. Fourth, the model should allow for wealth accumulation in many forms, because 401(k) plan participants can save through means that differ from 401(k)s in terms of tax treatment and liquidity, especially those that have been identified by previous studies as potential channels for 401(k)-saving offsets: taxable and IRA wealth (Poterba, Venti, and Wise, 1996; Hubbard and Skinner, 1996; Engen, Gale, and Scholz, 1996); Social Security, defined-benefit and non-401(k) defined-contribution pensions

(Gale, 1998; Engelhardt, 2001); and mortgage debt (Engen and Gale, 1998).[5] Fifth, when formulating the household budget constraint, the model should specify in detail the tax treatment of 401(k) and IRA contributions, respectively, IRA withdrawals, and the interrelationship between employer matching, 401(k) plan characteristics, and the price of leisure. This is a critical part of the analysis because the non-linear-budget-set estimation requires that non-linearities be well specified, and it is something that has been omitted from previous reduced-form studies. Sixth, it should allow for an uncertain rate of return and liquidity constraints, two central features of models of consumption in the literature. Seventh, in addition to the optimal asset allocation decision across the different forms of wealth, the model should follow a series of papers, including Shoven (1999), Shoven and Sialm (1998, 2003), Poterba, Sialm, and Shoven, (2001), Dammon, Spatt, and Zhang (2004), and Gomes, Michaelides, and Polkovnichenko (2004), among others, and allow for an optimal asset *location* decision, whereby the consumer must decide which assets to hold in taxable and tax-deferred forms. Finally, it should be tractable enough to be used directly to formulate an econometric model of 401(k) saving that is consistent with intertemporal optimization.

Model Set-Up

Formally, intratemporal direct utility, $U(C,l;\mathbf{z})$, is derived from leisure, l, with an associated price, p^l, consumption of a composite good, C, with an associated price, p^c, and a vector of demographics, $\mathbf{z}$. Intratemporal utility is weakly separable, and

[5] Technically, 401(k)s are defined-contribution plans under federal law, but a distinction will be made throughout this analysis between 401(k)s and all other DC plans, where the latter category will be referred to as "non-401(k)" DC plans and includes plans such as money-purchase, ESOP, target benefit, and profit-sharing plans. The exact type of plan is indicated in the pension plan data used in the analysis.

intertemporal utility is additively separable. The consumer faces a per period probability of survival of ρ, with period T being the known maximum length of life. With probability $1-\rho$, the consumer dies and receives the terminal payoff $\Phi(W^T)$, the utility of bequests, which is a function of total wealth, W^T. The lifetime is composed of two parts: from period N to T, the consumer is retired and no hours of labor are supplied to the market, so leisure equals the time endowment, L^l; from period 0 to $\mathrm{N}-1$, the consumer works, and the timing of retirement in period N is endogenous. Total wealth is accumulated in seven forms when working: wealth from IRAs, W^{IRA}; wealth from 401(k)s, W^{401k}; wealth from non-401(k) defined-contribution pension plans, W^{DC}; wealth from defined-benefit pension plans, W^{DB}; Social Security wealth, W^{SS}; housing equity, W^H; and taxable wealth, W^{TA}. Each period when working, the consumer chooses consumption, leisure, voluntary 401(k) contributions, Q^{VOL}, IRA contributions, Q^{IRAC}, IRA withdrawals, Q^{IRAW}, and the housing loan-to-value ratio, ϕ, (discussed below). Each period when retired, the consumer chooses consumption, IRA contributions, IRA withdrawals, the housing loan-to-value ratio, and receives eligible pension and Social Security benefits. Because 401(k) contributions can be made only while employed, the remaining description of the model focuses on the period when working, in the interest of exposition.[6]

The literature on optimal asset allocation and location to taxable and tax-deferred accounts has argued that, because capital gains are taxed more lightly than interest income in the United States, the consumer has the incentive to hold relatively heavily

[6] Details of the complete model are available upon request.

taxed bonds in tax-deferred accounts and relatively lightly taxed stocks in taxable accounts.[7] The model allows the consumer to choose the optimal allocation of IRA, DC, 401(k), and taxable wealth between stocks and bonds, in addition to the choice variables described above. Specifically, the model follows this literature and collapses IRA, DC, and 401(k) wealth into a single group, called retirement-account wealth, $W^{RA} \equiv W^{IRA} + W^{DC} + W^{401k}$, when formulating the intertemporal budget constraint. This is done for three reasons. First, IRAs, DC plans, and 401(k)s are all tax-deferred forms of saving, in that contributions are not taxed when made, they allow for inside build-up at the pre-tax rate of return, and qualified withdrawals are taxed as ordinary income at the time of withdrawal, so, in principle, there is little, if any, difference between vehicles.[8] Second, federal law allows an employee to roll a DC or 401(k) account balance over to an IRA upon job severance, so, in practice, assets that were once accumulated in an employer-provided pension may appear on the household balance sheet as IRA wealth (Engelhardt, 2002 and 2003a).[9] Third, this assumption helps keep the model tractable. Therefore, total wealth is defined as

$$W^{T} \equiv W^{RA} + W^{TA} + W^{DB} + W^{SS} + W^{H} . \tag{1}$$

Let s^h be the beginning-of-period share of total wealth in asset type h, $h = RA, TA, DB, SS, H$.

Defined-benefit pension and Social Security wealth evolve, respectively, as

[7] The obvious exception to this result is that tax-exempt municipal bonds should be held in the taxable account.

[8] There is a clear distinction in the model between deductible and non-deductible IRA contributions. Related to this, note that Roth IRAs were not available in 1991, the calendar year of the empirical analysis below.

[9] There is very clear evidence of these rollovers in the HRS data used below. In particular, there are individuals who have IRA wealth that is substantially higher than what they would have accumulated had they made limit contributions since 1981, when IRAs became legally widely available, at plausible rates of return. Theses individuals also self-reported having rolled pension assets over to an IRA.

$$W_{t+1}^{DB} = (1+\alpha_t^{DB}) s_t^{DB} W_t^T , \tag{2}$$

and

$$W_{t+1}^{SS} = (1+\alpha_t^{SS}) s_t^{SS} W_t^T . \tag{3}$$

For defined benefit plans, the accrual, α^{DB}, is determined by a plan formula that is usually a complicated function of age, earnings, and years of service. The Social Security accrual, α^{SS}, is determined by benefit formulas in the federal law. The model assumes that DB and Social Security wealth are illiquid until retirement and cannot be used as collateral.[10]

Following Gomes, Michaelides, and Polkovnichenko (2004), retirement-account and taxable wealth can be invested either in bonds, with riskless pre-tax return r^B, or in stocks, with a risky pre-tax return

$$\tilde{r}_t^S = \frac{1+\tilde{g}_t + d}{1+\pi} - 1, \tag{4}$$

where π is the constant inflation rate, d is the constant nominal dividend yield, and $\tilde{g}$ is the stochastic nominal capital gain earned from the beginning of period t to the beginning of period $t+1$.[11] Throughout the analysis, a return with a tilde will indicate an *ex ante* uncertain return, whereas one without will indicate either an *ex post* realization or

[10] Federal law prevents the use of Social Security or pension assets as collateral for loans, so that the illiquidity of these assets is a standard assumption in the literature.

[11] With respect to the empirical analysis described below, the HRS pension data tell whether employer stock was an investment option both for the employee voluntary contributions and the employer match, but only for plans that allowed individual-directed investment. Unfortunately, for plans that did not allow for individual-directed investment, the HRS did not electronically code this information off of the SPDs, so that it is not known whether employee and employer contributions were required to be invested in employer stock. Therefore, it is not possible to model and estimate the impact of employer stock on contribution behavior here specifically. Poterba (2003), Brown, Liang, and Weisbenner (2004), and Mitchell and Utkus (2002) have studied various aspects of investment in company stock.

a certain return. Let R^{RA} be the weighted-average return on wealth in retirement accounts,

$$R_t^{RA} \equiv (1-\vartheta_t^{RA})(1+r_t^B)+\vartheta_t^{RA}(1+\tilde{r}_t^S), \tag{5}$$

where ϑ^{RA} is the share of retirement-account wealth invested in stocks. Then retirement-account wealth evolves as

$$W_{t+1}^{RA} = R_t^{RA}[s_t^{RA}W_t^T + F_t + Q_t^{VOL} + M_t^{VOL} + Q_t^{MAN} + M_t^{MAN} + Q_t^{IRAC} - Q_t^{IRAW}]. \tag{6}$$

Some employers mandate 401(k) contributions, shown as Q^{MAN} in (6), and the employer may match those mandatory contributions (Cunningham and Engelhardt, 2002). In (6), M^{MAN} is the employer's matching contribution in dollars on the employee's mandatory contribution, and M^{VOL} is the employer's matching contribution in dollars on the employee's *voluntary* contribution, Q^{VOL}. The matching functions are

$$M_t^{VOL} = M^{VOL}(Q_t^{VOL}, y_t^l, \mathbf{m}_t^{VOL}) \tag{7}$$

and

$$M_t^{MAN} = M^{MAN}(Q_t^{MAN}, y_t^l, \mathbf{m}_t^{MAN}), \tag{8}$$

respectively, and are twice continuously differentiable; y^l is labor earnings; $\mathbf{m}^{VOL}$ and $\mathbf{m}^{MAN}$ are vectors of plan-specified match rates for voluntary and required contributions, respectively. In (6), F is the employer's non-matching contribution for the plan.[12] The

[12] These would be contributions the employer makes on a periodic basis, as specified in the SPD, but are not related to any voluntary or mandatory contributions by the employee: for example, those defined as a percentage of pay in a money purchase plan or as a function of some measure of firm performance in a profit-sharing plan.

model assumes that DC and 401(k) wealth are illiquid until retirement and cannot be used as collateral, but allows IRA withdrawals when working (discussed below).[13]

Housing equity is defined as $W^H \equiv P - D$. Housing value, P, evolves as

$$P_{t+1} = (1 + r_t^P) P_t , \tag{9}$$

where r^P is the return on housing value. Mortgage debt is $D = \phi P$, where ϕ is the loan-to-value (LTV) ratio, and D evolves as

$$D_{t+1} = (1 + r_t^D) D_t , \tag{10}$$

where r^D is the real mortgage-interest rate,

$$r_t^D \equiv \frac{1 + n_t^D}{1 + \pi} - 1 , \tag{11}$$

and n^D is the nominal mortgage-interest rate. Equations (9) and (10) can be used to write housing equity in period $t+1$ in terms of the LTV:

$$W_{t+1}^H = (1 + r_t^P) \frac{1}{1 - \phi_t} s_t^H W_t^T - (1 + r_t^D) \frac{\phi_t}{1 - \phi_t} s_t^H W_t^T . \tag{12}$$

In the model, the real mortgage-interest-payment portion on the right-hand side of (12), defined as $B \equiv r^D (\phi / 1 - \phi) s^H W^T$, is required to be paid out of cash on hand and appears in (14) below, so that the evolution equation for housing equity reduces to

$$W_{t+1}^H = \frac{(1 + r_t^P) - \phi_t}{1 - \phi_t} s_t^H W_t^T . \tag{13}$$

[13] There are 401(k) plans that allow hardship withdrawals and/or borrowing against plan balances when working that provide some liquidity to 401(k) saving. However, even though the SPD for each plan indicates whether, the extent to which, and the terms under which, hardship withdrawals and borrowing can occur, these features were not electronically coded by the HRS. Therefore, liquidity through hardship withdrawals and borrowing could not be incorporated as choice variables directly into the intertemporal budget constraint. Instead, the existence of these liquidity channels, for plans that offer them, are controlled for directly as explanatory variables in the empirical analysis.

Finally, the equation of evolution for taxable wealth is

$$W_{t+1}^{TA} = R_t^{TA}[s_t^{TA}W_t^T + w_t(L^l - l_t) + y_t^o - p_t^c C_t - Q_t^{VOL} - Q_t^{IRAC} + Q_t^{IRAW} - B_t - T_t], \qquad (14)$$

where R^{TA} is the return on taxable wealth,

$$R_t^{TA} \equiv (1 - \vartheta_t^{TA})(1 + r_t^B) + \vartheta_t^{TA}(1 + \tilde{r}_t^S), \qquad (15)$$

ϑ^{TA} is the share of taxable wealth invested in stocks, y^o is other income, w is the gross hourly wage rate, $w(L^l - l) \equiv y^l$ is labor earnings, and B is the real mortgage-interest payment.[14] In (14), T is the sum of income and payroll tax liability. It is a twice continuously differentiable function,

$$T_t = f^T(\boldsymbol{\theta}_t^F, \boldsymbol{\theta}_t^P, \boldsymbol{\theta}_t^E; I_t^F, \min(y_t^l, \mathbf{L}_t^P), D_t^{Age59-1/2} \cdot Q_t^{IRAW}), \qquad (16)$$

of a vector of statutory marginal income tax rates, $\boldsymbol{\theta}^F$, Social Security and Medicare payroll tax rates, $\boldsymbol{\theta}^P$, and the penalty-tax rate, $\boldsymbol{\theta}^E$, on early, non-qualified IRA withdrawals, which are defined by $D^{Age59-1/2} \cdot Q^{IRAW}$, where $D^{Age59-1/2}$ is a dummy variable that is one if the individual is under age 59½ and is zero otherwise.[15] Federal taxable income is

$$I_t^F = \{(y_t^l - Q_t^{VOL} - Q_t^{MAN}) + \lambda_t Q_t^{IRAW} + K_t - \zeta_t Q_t^{IRAC}\} - G_t. \qquad (17)$$

[14] Because the mean age in the HRS sample below is 55, and Carroll (1992) and Gourinchas and Parker (2002), among others, have estimated that most lifetime income uncertainty has been resolved by this age, at which point households have transitioned from buffer-stock to life-cycle savers, income uncertainty likely was not central to the households under study and is not included in the model, but would be an important consideration in comparing the empirical results below to those based on a younger sample. However, note that, as stressed by MaCurdy and Blundell (1999), the two-stage budgeting empirical approach adopted below explicitly can handle multiple forms of uncertainty, including, for example, wage uncertainty, and, therefore, in principle, this assumption does not compromise the basic framework underlying the empirical specification.

[15] For simplicity in exposition, we have suppressed notation for state income taxes. However, we include them in the empirical analysis below. Engelhardt (2002) discussed lump-sum distributions in detail.

The term in parentheses within the braces in (17) is income for federal tax purposes reported on Form W-2, the term in braces is adjusted gross income (AGI), and the term G represents the sum of personal exemptions and deductions, and includes deductible nominal mortgage interest for homeowners.[16] The term K is taxable capital income and is the sum of taxable interest income from bonds and taxable dividend and realized capital-gain income from stocks.[17] The factor λ is the fraction of IRA withdrawals that is federally taxable and depends on age and whether the withdrawal was qualified or not.[18] The factor ζ is the fraction of IRA contributions that is federally tax-deductible. It depends on adjusted gross income (AGI), and itself is a function of 401(k) contributions and IRA withdrawals, because 401(k) contributions are excluded from AGI and the taxable portion of withdrawals enters AGI.[19] $\mathbf{L}^P$ is a vector of covered-earnings caps for payroll taxes.

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[16] In the model, the housing service flow enters the composite commodity, C, and its rental cost for renters and the implicit rental cost for homeowners is subsumed into the price index, p^c.

[17] Dammon, Spatt, and Zhang (2001, 2004) have provided detailed models of consumption and investment, with and without tax-deferred assets, which specified detailed tax treatment of realized and unrealized capital gains.

[18] Engelhardt (2002) outlined and examined empirically in the HRS the federal tax treatment of non-qualified pre-retirement withdrawals from 401(k) and IRA plans. Federal law allows penalty-free withdrawals from IRAs at age 59½, at which point IRA assets become liquid; prior to 59½, tax-qualified withdrawals are allowed for a very limited number of reasons, and non-qualified withdrawals are assessed a penalty tax. The HRS surveyed individuals who were 51-61 in 1992, so some in the sample analyzed below are eligible for qualified withdrawals.

[19] Prior to the Tax Reform Act of 1986 (TRA86), IRA contributions were fully tax-deductible up to the limit of $2,000 for single individuals and $2,250 for married couples. TRA86 limited the deductibility of contributions. For single individuals, contributions remained fully deductible if adjusted gross income was less than $25,000, were linearly phased out for incomes between $25,000 and $35,000, and not deductible for incomes above $35,000. For married couples, contributions remained fully deductible if adjusted gross income was less than $40,000, were linearly phased out for incomes between $40,000 and $50,000, and not deductible for incomes above $50,000. Therefore, ζ varies according to a non-linear interaction of income, and marital status.

Equations (2), (3), (6), (13), and (14) sum to yield the intertemporal budget constraint that determines W_{t+1}^{T}. In addition, there are five other constraints on behavior. First, following Deaton (1991), there is a liquidity constraint,

$$W_t^{TA} \geq 0\,, \tag{18}$$

which means that total per period full expenditure (also referred to as "full income" in the two-stage budgeting literature), y, defined as

$$y_t \;\equiv p_t^c C_t + w_t l_t + B_t\,, \tag{19}$$

must be less than or equal to total net cash on hand, X, defined as beginning-of-period liquid taxable wealth and other income on hand, plus the market value of the leisure endowment, less the tax liability, plus any IRA wealth made liquid through a withdrawal, less any tax-deferred saving:

$$X_t \equiv W_t^{TA} + y_t^o + w_t L^l - T_t - Q_t^{VOL} - Q_t^{IRAC} + Q_t^{IRAW}\,. \tag{20}$$

Let μ_t be the associated Kuhn-Tucker multiplier. This formalizes the assumption that 401(k), defined-contribution, defined-benefit pension, and Social Security assets are illiquid prior to retirement. Second, although IRA withdrawals help to loosen the liquidity constraint by increasing total net cash on hand in (20), there are the following minimum and maximum constraints on withdrawals, with multipliers in square brackets:

$$Q_t^{IRAW} \geq D_t^{Age70-1/2} \cdot \frac{W_t^{IRA}}{h_t}\,, \qquad [\,\varphi_t^0\,] \tag{21}$$

$$Q_t^{IRAW} \leq W_t^{IRA}\,. \qquad [\,\varphi_t^L\,] \tag{22}$$

In (21), $D^{Age70-1/2}$ is a dummy variable that is one if the individual is age 70½ or older and is zero otherwise, and h is the individual's life expectancy. Thus, if under age 70½, the withdrawal must be greater than or equal to zero, and, if age 70½ or older, the withdrawal must satisfy the minimum-distribution requirements under federal law that are a function of life expectancy. Equation (22) states that the withdrawal cannot exceed the beginning-of-period IRA wealth.[20] Third, the minimum- and maximum-contribution constraints on 401(k)s and IRAs with multipliers in square brackets, respectively, are

$$Q_t^{VOL} \geq 0, \qquad [\eta_t^0] \quad (23)$$

$$Q_t^{VOL} \leq L_t^{VOL}, \qquad [\eta_t^L] \quad (24)$$

$$Q_t^{IRA} \geq 0, \qquad [\upsilon_t^0] \quad (25)$$

and

$$Q_t^{IRA} \leq L_t^{IRA}. \qquad [\upsilon_t^L] \quad (26)$$

The terms L_t^{VOL} and L_t^{IRA} are the upper limits on 401(k) and IRA contributions, respectively. L_t^{IRA} is governed by federal law and depends on marital status and pension coverage.[21] L_t^{VOL} is governed by the employer's plan, but may not exceed the federal statutory maximum.[22] Fourth, the loan-to-value ratio must lie in the unit interval for homeowners, but there may be a minimum home-equity constraint—such as a downpayment constraint (Engelhardt, 1996 and 2003b; Campbell and Cocco, 2003)—that further constrains the LTV to be below some exogenous threshold, $\bar{\phi}$. Formally, these constraints are

$$\phi_t \geq 0, \qquad [\varsigma_t^0] \quad (27)$$

[20] Note that the penalty tax on non-qualified withdrawals (e.g., those prior to age 59½) is accounted for in the taxes-paid function in (16).
[21] These limits apply to the sum of deductible and non-deductible IRA contributions.
[22] Leisure is not modeled as bounded by zero and the leisure endowment because 401(k) matching contributions are available only to those who work.

$$\phi_t \leq \bar{\phi}_t, \qquad [\varsigma_t^L] \quad (28)$$

where $\bar{\phi} \leq 1$. Finally, there are constraints on the shares of retirement-account and taxable assets allocated to stocks:

$$\vartheta_t^{RA} \geq 0, \qquad [\varpi_t^0] \quad (29)$$

$$\vartheta_t^{RA} \leq 1, \qquad [\varpi_t^L] \quad (30)$$

$$\vartheta_t^{TA} \geq 0, \qquad [\chi_t^L] \quad (31)$$

and

$$\vartheta_t^{TA} \leq 1. \qquad [\chi_t^L] \quad (32)$$

To summarize, the only forms of "active" saving when working are through contributions to 401(k), IRA, or taxable assets; adjustments can be made to the mortgage-debt position as well. However, the primary technology for smoothing resources across periods when working is through taxable-asset saving, because 401(k) saving is illiquid; IRA contributions are not necessarily illiquid because of the availability of withdrawals, but IRA withdrawals may incur a tax penalty; traditional pensions and Social Security are illiquid; and the extent of mortgage borrowing is limited. This means that the consumer's optimization does not imply automatically that all active saving be allocated first to the tax-preferred asset with the highest net return, because, in the face of uncertainty, the consumer must balance the desire for a high return with the need for liquidity.

First-Order Conditions

As explained in the data section below, consumption and hours are not fully observed in the HRS, so that, from the perspective of the empirical analysis, it is desirable to work with the indirect, rather than the direct, utility function. Specifically, let $V(\mathbf{p}, y; \mathbf{z})$ be the intratemporal indirect utility function. It takes as arguments the vector of prices of leisure and consumption, $\mathbf{p}$, full income, y, given in (19), and the vector of demographics, $\mathbf{z}$. Following Browning, Deaton, and Irish (1985), let $V_t^*(W_t^T)$ be the sum of current and future expected utility based on total wealth in period t. The individual makes all decisions at the beginning of the period, based on the information set, Ω_t, after which, r^s is realized. E is the expectations operator conditional on the information set. Then for any time t, $t < \mathrm{N}$, the dynamic optimization problem can be written as

$$\begin{aligned} V_t^*(W_t^T) = \max_{\{\mathbf{c}\}} \Big\{ & V(\mathbf{p}_t, y_t; \mathbf{z}) + E_t[\rho_t V_{t+1}^*(W_{t+1}^T) + (1-\rho_t)\Phi(W_{t+1}^T)] + \mu_t (X_t - y_t) \\ & + \varsigma_t^L (\bar{\phi}_t - \phi_t) + \varsigma_t^0 (\phi_t - 0) + \eta_t^L (L_t^{VOL} - Q_t^{VOL}) + \eta_t^0 (Q_t^{VOL} - 0) \\ & + \upsilon_t^L (L_t^{IRAC} - Q_t^{IRAC}) + \upsilon_t^0 (Q_t^{IRAC} - 0) + \varphi_t^L (W_t^{IRA} - Q_t^{IRAW}) \\ & + \varphi_t^0 (Q_t^{IRAW} - D_t^{Age70-1/2} \cdot (W_t^{IRA} / h_t)) \\ & + \varpi_t^L (1 - \vartheta_t^{RA}) + \varpi_t^0 (\vartheta_t^{RA} - 0) + \chi_t^L (1 - \vartheta_t^{TA}) + \chi_t^0 (\vartheta_t^{TA} - 0) \Big\}, \end{aligned} \tag{33}$$

where $\mathbf{c}$ contains the choice variables: $y_t, Q_t^{VOL}, Q_t^{IRAC}, Q_t^{IRAW}, \phi_t, \vartheta_t^{RA}, \vartheta_t^{TA}$; and there is a two-stage budgeting interpretation.[23] In the first-stage, the individual chooses full income, dis-saving through IRA withdrawals, the mortgage-debt position, and the portfolio allocations to stock of retirement-account and taxable wealth, and must allocate total "active" saving to three asset categories─401(k), IRA, and taxable wealth─to maximize the expected present discounted value of lifetime indirect utility. Note that the choice variable for taxable-asset saving is made redundant by the intertemporal budget

[23] The necessary condition for two-stage budgeting is that utility be weakly separable (Gorman, 1959). The model assumes strongly intertemporally and weakly intratemporally separable preferences, so that a two-stage budgeting interpretation is valid. The choices of 401(k) contributions and portfolio shares do not apply when not working, where with regard to the latter it is assumed for simplicity that pension wealth is annuitized at retirement.

constraint and does not appear explicitly in $\mathbf{c}$. In the second stage, optimal full income, y^*, in each period is allocated statically between the goods that enter direct utility: consumption and leisure.

The first-order conditions when working for 401(k) contributions, IRA contributions, and full income can be expressed as

$$\eta_t^L - \eta_t^0 = E_t \left\{ [R_t^{RA}(1 + M_{Q_t^{401k}}^V) - R_t^{TA}(1 - T_{It}(1 - \zeta_{y^I{}_t} Q_t^{IRA}))] \cdot [\rho_t V_{W_{t+1}^T}^* + (1-\rho_t)\Phi_{W_{t+1}^T}] \right\} - \mu_t (1 - T_{It}(1 - \zeta_{y^I{}_t} Q_t^{IRA})), \tag{34}$$

$$\upsilon_t^L - \upsilon_t^0 = E_t \left\{ [R_t^{RA} - R_t^{TA}(1 - T_{It}\zeta_t)] \cdot [\rho_t V_{W_{t+1}^T}^* + (1-\rho_t)\Phi_{W_{t+1}^T}] \right\} - \mu_t (1 - T_{It}\zeta_t), \tag{35}$$

and

$$V_y(\mathbf{p}_t, y_t; \mathbf{z}) = E_t \left\{ R_t^{TA} [\rho_t V_{W_{t+1}^T}^* + (1-\rho_t)\Phi_{W_{t+1}^T}] \right\} + \mu_t, \tag{36}$$

respectively. Note that subscripts indicate a partial derivative (other than t, which denotes time): for example, T_I is simply the marginal tax rate; $M_{Q^{401k}}^V$ is the marginal employer match rate for an additional dollar of 401(k) contribution; ζ_{y^I} is the change in the fraction of an IRA contribution that is deductible for an additional dollar of AGI; and V_y is the marginal utility of full income.[24]

Even though the typical employer match yields a net return far exceeding that on other assets, so that it would appear obvious that the individual always would want to make the maximum possible contribution to the 401(k), equation (34) indicates the role of liquidity in the 401(k) contribution decision. In particular, there are two ways for this

[24] The first-order conditions for the other choice variables are available upon request and are not shown here simply in the interest of exposition.

equation to be satisfied when contributions are less than the plan maximum even when the match and tax rates are positive, $M^V_{Q^{401k}} > 0$ and $T_I > 0$, respectively, and these occur when the liquidity constraint binds, $\mu > 0$: the corner solution of no 401(k) saving, for which the Kuhn-Tucker multipliers on the contribution constraints in (23) and (24) are $\eta^0 > 0$ and $\eta^L = 0$, respectively; and an interior solution, for which $\eta^0 = 0$ and $\eta^L = 0$ (i.e., the contribution is positive, but not at the plan maximum), and the multiplier on the liquidity constraint is large enough that the second term just equals the first term on the right-hand side of (34). A particularly important example of this latter case is when the contribution is less than the match cap and the employee leaves "money on the table" by not contributing up until the point the match is exhausted. Therefore, even in the presence of an employer match, binding liquidity constraints can explain why 401(k) participation can be less than 100 percent, contributions can be less than the plan maximum, and employees rationally can leave money on the table. This implication of the model is examined in Section VI below, in which reduced-form specifications are estimated to see whether variables used by others in the literature to measure liquidity constraints can explain who fails to contribute at least until the match is exhausted.

III. Econometric Specification

There are two major obstacles to the use of (34) as a direct basis for estimating the impact of employer matching on 401(k) saving. First, who is liquidity constrained, and, therefore, μ, is not perfectly observed. Second, the sum of future expected utility, V^*, is not observed.

To overcome these obstacles, define the following tax and match prices: $p_t^{IRA} \equiv 1 - T_{It}\zeta_t$, $p_t^m \equiv 1 + M^V_{Q_t^{401k}}$, and $p_t^{401k} \equiv 1 - T_{It}(1 - \zeta_{y^l_t} Q_t^{IRA})$.[25] Then distribute terms in (34) and divide both sides by p_t^{401k} to yield

$$\begin{aligned}\frac{\eta_t^L - \eta_t^0}{p_t^{401k}} = \frac{p_t^m}{p_t^{401k}} \cdot E_t\left\{R_t^{RA}[\rho_t V^*_{W^T_{t+1}} + (1-\rho_t)\Phi_{W^T_{t+1}}]\right\} \\ - E_t\left\{R_t^{TA}[\rho_t V^*_{W^T_{t+1}} + (1-\rho_t)\Phi_{W^T_{t+1}}]\right\} - \mu_t.\end{aligned} \tag{37}$$

The last two terms on the right-hand side of (37) are equal to $-V_y(\mathbf{p}_t, y_t; \mathbf{z})$ according to (36), so that (37) reduces to

$$\frac{\eta_t^L - \eta_t^0}{p_t^{401k}} = \frac{p_t^m}{p_t^{401k}} \cdot E_t\left\{R_t^{RA}[\rho_t V^*_{W^T_{t+1}} + (1-\rho_t)\Phi_{W^T_{t+1}}]\right\} - V_y(\mathbf{p}_t, y_t; \mathbf{z}), \tag{38}$$

which eliminates μ. Similarly, (35) can be re-expressed as

$$\frac{\upsilon_t^L - \upsilon_t^0}{p_t^{IRA}} = \frac{1}{p_t^{IRA}} \cdot E_t\left\{R_t^{RA}[\rho_t V^*_{W^T_{t+1}} + (1-\rho_t)\Phi_{W^T_{t+1}}]\right\} - V_y(\mathbf{p}_t, y_t; \mathbf{z}). \tag{39}$$

Solve (39) for the expectations term, then substitute into (38), multiply through by p_t^{401k}, and combine terms to yield

$$\eta_t^L - \eta_t^0 = \Delta p_t \cdot V_y(\mathbf{p}_t, y_t; \mathbf{z}) + p_t^m(\upsilon_t^L - \upsilon_t^0), \tag{40}$$

where $\Delta p \equiv p^m p^{IRA} - p^{401k}$, which eliminates V^*.

Equation (40) serves as a basis for the functional form for the estimation and lends itself naturally to a limited-dependent-variable econometric model of the determinants of 401(k) contributions. In particular, for the first term on the right-hand

[25] These three prices are not for goods that enter the intratemporal direct utility function and, therefore, are not in the price vector **p** that is an argument in the indirect utility function.

side, the marginal indirect utility of full income, V_y, is measured in utils, not dollars. However, at the optimum,

$$y^* \equiv e(\mathbf{p}, V; \mathbf{z}), \tag{41}$$

where the expenditure function, $e(\mathbf{p}, V; \mathbf{z})$, is expressed in dollars, not utils. Differentiate both sides of (41) with respect to y and rearrange to yield

$$V_y = \frac{1}{e_V}. \tag{42}$$

To substitute out for V_y, assume the expenditure function is

$$\ln e(\mathbf{p}, V; \mathbf{z}) = \ln[a(\mathbf{p})] + \Psi(\mathbf{z})^{-1} V \ln[b(\mathbf{p})], \tag{43}$$

associated with the PIGLOG indirect utility function (Muellbauer, 1976),

$$V(\mathbf{p}, y; \mathbf{z}) = \Psi(\mathbf{z}) \cdot \frac{\ln(y) - \ln[a(\mathbf{p})]}{\ln[b(\mathbf{p})]}, \tag{44}$$

that has been used extensively in the literature on consumption, where, following Blundell, Browning, and Meghir (1994), Ψ is a utility scaling factor that is a function of the exogenous demographic characteristics. In (43), a is homogeneous of degree one, and b is homogeneous of degree zero and modeled as a Cobb-Douglas price aggregator

$$b(\mathbf{p}) = \prod_k p_k^{\beta_k}, \tag{45}$$

across the k goods that enter the direct utility function, where

$$\sum_k \beta_k = 0. \tag{46}$$

Because there are only two goods that enter direct utility, leisure $(k = 1)$ and consumption, respectively, (46) implies $\beta_2 = -\beta_1$, so that (45) can be re-written as

$$b(\mathbf{p}) = \omega^{\beta_1}, \tag{47}$$

where $\omega \equiv p^l / p^c$ is the real relative price of leisure. Differentiate both sides of (43) with respect to V and substitute in (41) and (47) to yield

$$e_V = \Psi(\mathbf{z})^{-1} \beta_1 y \ln(\omega) \tag{48}$$

at the optimum, and model Ψ as

$$\Psi_i(\mathbf{z}) = \psi_0 + \sum_m \psi_m z_{i,m} \,, \tag{49}$$

where $\mathbf{z}$ is an $m \times 1$ vector. The second term on the right-hand side of (40) is zero when IRA saving is at an interior solution, positive when constrained by the upper IRA limit, and negative when at the lower IRA limit (of zero). Therefore, let

$$\kappa \equiv p^m (D^{L^{IRA}} - D^0) \,, \tag{50}$$

where $D^{L^{IRA}}$ is a dummy variable that is one if IRA contributions, which are measured in the HRS, are at the upper limit and zero otherwise, and D^0 is a dummy variable that is one if IRA contributions are zero and zero otherwise. Let Q^{VOL*} denote desired 401(k) contributions. Then equation (40), combined with equations (41), (43), (48), (49), and (50), suggests that desired 401(k) contributions can be modeled as

$$Q_{ijt}^{VOL*} = \delta_0 \frac{\Delta p_{ijt}}{y_{ijt} \ln(\omega_{ijt})} + \sum_m \delta_{1m} \frac{\Delta p_{ijt}}{y_{ijt} \ln(\omega_{ijt})} z_{i,m} + \delta_2 \kappa_{ijt} + \alpha \mathbf{x}_{ijt} \,, \tag{51}$$

where i and j index individuals and 401(k) plans, respectively, and the regression index is linear in parameters.

In the empirical analysis, $\mathbf{z}$ includes the worker's education (in years), age, and dummy variables for whether the worker was married, white, and female, respectively. These demographic characteristics enter parsimoniously, in a manner which allows the impact of employer matching to be heterogeneous across demographic groups.[26] The last term on the right-hand side of (51) includes $\mathbf{x}$, a vector that contains a constant and exogenous employer and employment characteristics. These are additional factors, explained in section VIII below, which fall outside of the scope of the theoretical framework, but may affect contributions. In the baseline specifications, $\mathbf{x}$ is limited to a constant, so that (51) follows (40) and the theoretical framework directly; additional specifications allow the employer and employment characteristics to enter $\mathbf{x}$.

The upper and lower limits on contributions in (23) and (24) lead to a two-limit censored regression model. Specifically, when there is an interior solution ($\eta^L - \eta^0 = 0$) in (40), actual 401(k) contributions are defined as the sum of desired contributions and optimization error, ε:

$$Q_{ijt}^{VOL} = Q_{ijt}^{VOL*} + \varepsilon_{ijt}. \tag{52}$$

At the upper contribution limit ($\eta^L - \eta^0 > 0$), $Q_{ijt}^{VOL*} + \varepsilon_{ijt} \geq L_{ijt}^{VOL}$, but $Q_{ijt}^{VOL} = L_{ijt}^{VOL}$; at the lower contribution limit ($\eta^L - \eta^0 < 0$), $Q_{ijt}^{VOL*} + \varepsilon_{ijt} \leq 0$, but $Q_{ijt}^{VOL} = 0$. Appendix B provides a detailed discussion of the nature of the optimization error.[27]

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[26] It is useful to note in passing that, unlike *ad hoc* reduced-form specifications, in which researchers often include measures of the present value of future entitlements to public and private pension wealth as regressors, those wealth measures play no direct role in the specification here, because their effects are transmitted through the allocation of full expenditure across periods, and, hence, the level of the marginal utility of full income.

[27] An alternative estimation strategy would be to combine (41), (43), (48), (49), and (50) and estimate (40) directly:

$$(\eta^L - \eta^0)_{ijt} = \delta_0 \frac{\Delta p_{ijt}}{y_{ijt} \ln(\omega_{ijt})} + \sum_m \delta_{1m} \frac{\Delta p_{ijt}}{y_{ijt} \ln(\omega_{ijt})} z_{i,m} + \delta_2 \kappa_{ijt} + \alpha \mathbf{x}_{ijt},$$

which could be done, for example, by maximum likelihood. Indeed, in a companion paper, Engelhardt and Kumar (2004b), this was done, and the structural estimates were used to simulate the responsiveness of voluntary participation in prospective Social Security personal accounts with respect to government matching contributions. The advantage of this approach is that it is structural and directly estimates the parameters governing the first-order condition. The disadvantages are that it only provides estimates the effect of matching on the extensive (participation) margin, but says nothing about the intensive margin, and

IV. Data

Previous research primarily has used nationally representative, individual-level survey data, such as the Current Population Studies (CPS) and Surveys of Consumer Finances (SCF), which are plagued by two important sources of measurement error. First, even though the researcher must know the entire match schedule for a plan to account for the individual's full opportunity set, as well as whether the match is discretionary or through profit-sharing, the typical survey respondent has great difficulty in accurately conveying even relatively simple pension provisions to interviewers, no less detailed matching schedules.[28] Second, self-reported contribution data also suffer from substantial reporting error.[29] In addition, as the theoretical framework showed, the data required to model saving are quite extensive: contributions, components of household (including spousal) income, assets, debts, demographics, marginal tax rates, spousal pension coverage, and expected entitlements from Social Security and traditional

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does not use any of the variation in contributions that are observed to be interior solutions to identify the parameters. The specification in the current paper, shown in (51)-(52), though technically not the primitive, structural relationship, reflects the first-order condition embodied in (40) and utilizes the additional information in the variation in contributions from interior solutions, and, importantly, allows for the calculation of separate responses on the extensive and intensive margins for the IV-Tobit specifications shown below in Table 6. For comparative purposes, the maximum-likelihood estimated extensive-margin (participation) elasticity from the structural estimation in Engelhardt and Kumar (2004b) for the specification shown in column 7 of Table 6 was 0.10 with a standard error of 0.05. This estimated elasticity is very similar in magnitude to the estimated extensive-margin elasticity from (51)-(52) shown in column 7, Panel E of Table 6 below of 0.07 with a standard error of 0.03, and there is substantial overlap in the 95-percent confidence intervals around these two estimated elasticities. This illustrates that the estimation of (51)-(52) and the structural model are really getting at the same underlying behavioral response. Engelhardt and Kumar (2004b) estimated the structural relationship because there the focus was solely on participation; the current paper estimates (51)-(52) because the focus is on 401(k) saving, not just participation, so that the intensive-margin response needs to be estimated.

[28] See Mitchell (1988), Starr-McCluer and Sunden (1999), Johnson, Sambamoorthi, and Crystal (2000), Gustman and Steinmeier (1999), Rohwedder (2003a, 2003b), and Engelhardt (2001) for evidence on measurement error in pension data.

[29] For example, some plans mandate employee contributions as a condition of eligibility (Cunningham and Engelhardt, 2002). Surveys like the CPS and SCF do not distinguish between mandated and voluntary contributions, so that voluntary contributions are measured with error. Overall, there is likely substantial measurement error in survey data.

pensions, which require lifetime and job earnings histories, respectively. Previous studies have not had all of these data.[30]

In this paper, these problems are overcome by using remarkably detailed data from the first wave of the HRS, a nationally representative random sample of 51-61 year olds and their spouses (regardless of age). The first wave asked detailed questions about wealth (including IRA and taxable assets), demographics, and spousal characteristics in 1992. The survey also asked detailed questions about household income, tax information, and IRA contributions, but, as is true in many household surveys, these questions were for the *previous* calendar year, 1991. So, for the purposes of the empirical analysis, periods t and $t+1$ refer to 1991 and 1992, respectively.

Questions on employment were asked for the job (if any) held at the time of the interview, as well as previous jobs. A unique feature of these data is that the HRS used the job rosters from the household interviews and collected Summary Plan Descriptions (SPDs), which are legal descriptions of pensions written in plain English, from employers of HRS respondents for all current and previous jobs in which the respondent was covered by a pension. These descriptions help to sidestep the problems with measurement error outlined above, and, instead, measure the exact incentives to contribute by using the employer matching formulas given in the SPD.[31] Specifically,

[30] Studies that instead have used detailed pension plan descriptions and employer personnel records have circumvented these reporting-error issues and yielded very useful insights, but at the cost of relatively little or no knowledge about other factors that affect the worker's 401(k) saving decision, such as total household income, wealth, and spousal characteristics. (Choi, Laibson, Madrian, and Metrick, 2002, 2003a, 2003b, 2003c; Clark and Schieber, 1998; Kusko, Poterba, and Wilcox, 1998; VanDerhei and Copeland, 2001). In addition, these studies have focused on a limited set of firms that are not necessarily nationally representative. Other studies have used firm-level data from Form 5500 (Papke, 1995; General Accounting Office, 1997), which have no information on individual employees. Joulfaian and Richardson (2001) used W-2 data on contributions, but lacked detailed data on wealth, family characteristics, and pension plans.

[31] Also note that in the employment section of the survey, there were questions on pensions that were asked to the respondent. This information was not used to construct our dataset because of well-documented

the job in which the respondent was employed in 1991 was identified and then SPDs associated with that job that had dates of adoption after 1991 were excluded. In addition, the date of last amendment and dates for changes in plan provisions indicated in the text of the SPD were used to exclude plans that were in existence in 1991 but whose features changed between 1991 and the time the SPD was collected.

The HRS also asked in the first wave the respondents' permission to link their survey responses to administrative earnings data from SSA and IRS. These administrative data include Social Security covered-earnings histories from 1951-1991 and W-2 earnings records for jobs held from 1980-1991, and were made available for use under a restricted-access confidential data agreement. They are the basis for two critical measures in the analysis dataset. First, the W-2 data provide administrative data on earnings and 401(k) contributions for 1991. Unlike the contributions data used in previous studies, these data are not subject to measurement error, as they are the employer's official report to the government on annual earnings and elective deferrals. Second, when combined, the W-2 earnings histories, Social Security covered-earnings histories and self-reported earnings histories, allowed for the construction of complete earnings histories from 1951-1991 for each member of our sample. When used with Social Security and pension-benefit calculators, which are described in Appendix A, these data allowed for the calculation of the public and private pension wealth, accruals and changes in accruals, for 1991 and 1992, respectively.

Overall, when all of the sources are combined, the data are a comprehensive description of the household's financial situation and exact pension incentives in 1991

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measurement error in these data, e.g., Johnson, Sambamoorthi, and Crystal (2000), Gustman and Steinmeier (1999), Rohwedder (2003a, 2003b), and Engelhardt (2001).

and 1992 to estimate the parameters in the empirical specification in (51)-(52) and a significantly richer data source than previous studies. Specifically, the sample consists of 1,042 HRS individuals eligible to contribute to a 401(k) in 1991.[32]

V. Descriptive Statistics

Many plans limit the amount of the match. These caps are usually expressed as a percent of pay in the SPD, but also can be a percent of contributions, and even a fixed-dollar amount. Table 1 shows the distribution of matching caps in the analysis sample, expressed as a percent of annual pay. About 19 percent of these plans had caps on employer matching that were less than four percent of pay. The median cap was 6 percent of pay, but 15 percent of plans had higher caps. Plans also vary according to the

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[32] It should be emphasized that, for a number of reasons, the HRS data cannot be assembled in a way to perform either joint estimation of all of the parameters that determine all of the choice variables (such as consumption, leisure, etc.) or to perform longitudinal analysis of 401(k) saving. First, the 401(k) contributions are measured from the W-2 data, which end in 1991, but there are no income data (from all sources) prior to 1991. This precludes longitudinal analysis using years prior to 1991. Second, there is no effort currently underway to extend the W-2 data for the original HRS cohort (b. 1931-41) beyond 1991, because this is not covered under the original Memorandum of Understanding with SSA and IRS and would require additional permission from respondents. There is an effort underway to get W-2 data for the War Babies cohort (b. 1942-46), who entered the HRS/AHEAD in 1998, that would cover earnings through 1997, but those individuals do not appear in the analysis dataset, so this is not useful in this context. Third, the respondent-reported contributions to pensions in the employment sections of the survey in the first and subsequent waves have substantial measurement error, many missing values, and do not distinguish between mandatory and voluntary contributions; hence, they are not appropriate for longitudinal analysis. Fourth, even if there were good contributions data after 1991 (which there are not), SPDs were collected only in 1993. They were collected again in 1998 for those still working, but are not available for the intervening years, so that it is not possible to track exact changes in pension incentives over time in the HRS at the level of detail needed for the analysis in this paper. Fifth, questions on hours per week and weeks worked were asked for the job held at the time of the first interview. But because these interviews were held after 1991, the responses to these questions cannot be used to determine labor supply and leisure in 1991 (as they refer to 1992), which means that joint estimation of the labor supply equation is not possible. Fifth, only a limited set of questions on consumption expenditures were asked in the first interview (primarily food and housing), and they were asked about behavior in 1992, not 1991, so that this information cannot be used to jointly estimate the consumption equation from the theoretical framework. Therefore, even though the HRS has very rich data, the structure and timing of the survey prevent anything beyond cross-sectional estimation for 1991 for this analysis. One implication of this data limitation is that it will not be possible to estimate any intertemporal preference parameters, and nothing can be said about the elasticity of intertemporal substitution (EIS) from this analysis; in particular, the estimation will only provide estimates from (51)-(52) of *some*, but not all, of the intratemporal preference parameters in (44).

match rate. Table 2 shows the distribution of "first-dollar" match rates in the analysis sample. Columns 1 and 2 indicate that these match rates were clustered at 25, 50, and 100 percent, where the median match rate was 50 percent. However, 27 percent of the plans offered matches of 100 percent, and three plans offered match rates of 200 percent.

Descriptive statistics for selected variables used in the empirical analysis are shown in Table 3. Column 1 shows sample means for the full sample, with the standard deviation in parentheses, and the median in square brackets. Overall, the sample consists of mostly white, married individuals in their mid-50s, with some college education and relatively few children at home. Only 56.4 percent of the sample actively participated (defined as having made a positive contribution) in 1991. The sample mean 401(k) contribution in calendar year 1991 was $1,377, but among contributors, the average contribution was $2,446 (shown in column 4). A comparison of contributions between those without and with employer matching in columns 2 and 3, respectively, indicates that individuals with matching contributed just over $400 more on average than those without matching (i.e., $1,640-$1,232=$408). The difference in the median contributions between these two groups was $800.

A comparison of columns 2 and 3 in Table 3 also indicates that plans with employer matching differ along other dimensions that may make saving attractive. For example, if there is an employer match, the individual is much more likely to be able to borrow against the plan balance, direct the investment of plan balances, less likely to have another traditional pension plan, more likely to have the plan annual contribution limit lower than the federal limit, and more likely to be allowed to make after-tax contributions to the plan.

VI. Explaining Unused Employer Matching Contributions

Because the typical employer match yields a return far exceeding that on alternative investments, 401(k) participation would be predicted to be 100 percent if all individuals were fully informed, financially rational, with access to perfect capital markets and no transactions costs. In addition, at a minimum, all participants would be predicted to maximize total compensation and contribute up to the point at which the employer match was exhausted and then engage in a set of borrowing and lending arrangements to achieve the desired level of consumption and leisure.[33] Yet in Table 3, 401(k) participation is 56.4 percent among all sample individuals, and only 54 percent among those offered a match. This suggests that individuals left "money on the table" by not capturing the total potential employer match.

Column 1 of Table 4 shows the total potential employer match in the sample for individuals eligible for a match. The mean potential match was $1,249, or 3.8 percent of annual pay. The average employer match that went unused because contributions were not made up to the level of the match cap was $550, or 1.9 percent of pay. Naturally, non-participants accounted for most of this, with the unused match equal to 3.7 percent of pay (column 3). Even more striking, though, is that among *participants* the average unused employer match represented 1 percent of pay (column 2).

As described in the theoretical framework, one possible explanation for this is that individuals were liquidity constrained. To explore this, *ad hoc* reduced-form models were estimated measuring whether and to what extent the 401(k) contribution was less

[33] A similar argument would imply that all individuals would be predicted to contribute up to the plan limit, and engage in a set of borrowing and lending arrangements to achieve the desired consumption and leisure, and thus exploit the tax arbitrage from deferral to minimize the lifetime tax liability.

than the cap on the employer match as a function of a set of explanatory variables that measure (to varying degrees) the ability to borrow, which have been used by others in the previous literature on liquidity constraints, but none of which measure such constraints definitively: demographics (white, age, married, and years of education); dummy variables for whether the household has no capital income; has access to borrowing against home equity through a home equity line of credit, conditional on being a home owner; has experienced financial distress in the past due to unanticipated medical expenses and unemployment, respectively; and has access to informal private support from friends and family if under financial distress.[34]

The results are shown in Table 5 for the sub-sample of individuals offered a match. Sub-sample means are shown in column 1. Column 2 shows estimates from a probit model in which the dependent variable is one if the individual contributed below the employer match cap (including a zero contribution) and zero otherwise. Individuals who were more educated or with a home equity line of credit were statistically significantly less likely to have contributed below that match cap, whereas those individuals having had financial distress from unanticipated medical expenses or no capital income were statistically significantly more likely to have contributed below that match cap. Columns 3 and 4 show estimates for Tobit models of the dollar amount and

[34] It is important to emphasize to the reader that this section of the paper includes the results of *ad hoc* reduced-form models of the impact of liquidity constraints on "money on the table" only to provide some evidence in support of (or at least not inconsistent with) the mechanism in the theoretical framework that would explain individuals' failure to fully exploit the employer matching and tax-deferral in 401(k)s—namely, liquidity constraints. In particular, measures of liquidity constraints used by others in the literature are used here simply for comparative purposes; indeed, there is debate as to how well these variables actually measure constraints (e.g., Jappelli, Pischke, and Souleles, 1998). As is clearly stated in section III, direct measures of liquidity constraints (such as the variables used in Table 5) are not used in the econometric specification in (51)-(52) because who is liquidity constrained is not perfectly observed in the HRS data. Indeed, this is reflected in (38), where the multiplier on the liquidity constraint was substituted out. One other note: the results in Table 5 did not differ if a dummy for negative or zero financial wealth was used in place of the dummy for no capital income.

the percentage of pay of the unused employer match, respectively, and the results are qualitatively similar. These results are not inconsistent with the theoretical result that constrained borrowing is a plausible explanation for why workers fail to capture the total potential employer match.[35]

VII. Estimation and Identification

There are a number of approaches to the estimation of non-linear-budget-set models, including the maximum-likelihood estimation pioneered by Burtless and Hausman (1978) and summarized in Hausman (1985), in which each segment and kink point on the individual's budget set has its own contribution to the likelihood function—and the recent, related non-parametric extensions by Blomquist and Newey (2002)—the maximum likelihood differentiable-budget-constraint methodology of MaCurdy, Green, and Paarsch (1990), and instrumental-variable techniques that linearize the budget set at the observed outcome to calculate the price and virtual income terms and then instrument to correct for endogeneity, which also has a long history, but a recent example of which is

[35] Again, it should be emphasized that these measures of constraints, though used in the literature, are not definitive. There are at least two other explanations for the presence of unused employer matching contributions. First, employees may have imperfect information about the details of the plan. A series of studies, including Bayer, Bernheim, and Scholz (1996), Bernheim and Garrett (2003), Bernheim, Garrett, and Maki (2001), Duflo and Saez (2002, 2003), among others, suggest that financial education raises participation in retirement saving programs. Second, employees may procrastinate, perhaps due to indirect costs of learning about the 401(k) plan and why employer matching is such a good deal, or have present-biased preferences. A provocative set of studies by Madrian and Shea (2001), Choi, Laibson, Madrian, and Metrick (2002, 2003a, 2003b, 2003c, 2004), Choi, Laibson, and Madrian (2004), O'Donoghue and Rabin (1999a, 1999b), Diamond and Koszegi (2003), Benartzi and Thaler (2004), Laibson (1997), and Laibson, Repetto and Tobacman (1998), among others, have addressed various empirical and theoretical aspects on this. While we are not necessarily unsympathetic to these alternative explanations, they play little role in our analysis simply because the HRS data we use, though incredibly rich, are not rich enough to model these additional aspects of behavior. We do note, however, that automatic enrollment (Madrian and Shea, 2001) does not play a role in the plans we analyze. Specifically, we asked Bob Peticolas of the HRS to review the SPDs in the HRS and none of them had default enrollment like that analyzed in Madrian and Shea (2001). We also control in our specifications below for whether respondent's and spouse's employers offer retirement seminars and whether the respondent and spouse discussed retirement with their respective co-workers.

Ziliak and Kniesner (1999). In the empirical analysis below, the parameters in (51) are estimated along the lines of the latter approach.

In particular, for all observed 401(k)-contribution outcomes in the dataset, the tax and match prices, p^{IRA}, p^{m}, and p^{401k}, the net wage, ω, and full income, y, must be calculated in order to construct $\Delta p / y \ln(\omega)$ and κ in (51). Because budget-set slopes are not defined at kink points, a variant of the method of MaCurdy, Green, and Paarsch (1990) was used to calculate p^{IRA}, p^{m}, and p^{401k} for each individual in the sample. Specifically, the matching formulas in the SPDs, tax-rate information from NBER's TAXSIM calculator, and detailed household financial and demographic characteristics were used to lay out the budget set in detail, then the kinks in the budget set were smoothed non-parametrically using kernel regression of the implicit (negative) tax rate from the employer matching and tax subsidy to contributions on AGI over the federal legally allowable range of 401(k) contributions of 0 to \$9500 using a second-order Gaussian kernel, $K(z) = (1/\sqrt{2\pi})e^{-z^2/2}$, with bandwidth chosen by Silverman's rule of thumb, $h = 0.9m/n^{1/5}$, where $m = \min(\sqrt{\mathrm{var}_x}, iqr_x/1.349)$ and iqr_x is the inter-quartile range.[36] This regression was done on an individual-by-individual basis, so that the smoothing is individual-budget-set specific, and the estimates allow for budget-set slopes to be measured for those individuals located at kink points.

To measure full income, y, note that, from (14), it can be expressed as

$$y_{it} \equiv \Delta A_{it}^{TA} + w_{it} L^{l} + y_{it}^{o} - Q_{it}^{VOL} - Q_{it}^{IRAC} + Q_{it}^{IRAW} - T_{it} \,, \tag{53}$$

[36] See Appendix A for details. MaCurdy, Green, and Paarsch (1990) used a cubic polynomial. The kinks were smoothed using cubic polynomial and fractional polynomial regression as alternatives to kernel regression, and, overall, the results from the three methods were qualitatively similar, but the kernel regression gave a better fit to the budget sets. In the rare cases of non-convexity in the budget set, the convex hull of the budget set was used, following the labor-supply literature.

and includes the market value of the leisure endowment.[37] Under two-stage budgeting, the capital income and net (dis-)saving terms embodied in ΔA are sufficient statistics for the past and the expectations of future variables (Blundell and Macurdy, 1999). Because of the non-linear structure of matching and marginal tax rates, the tax and match prices, p^{IRA}, p^{m}, and p^{401k}, change depending upon the budget-set segment (either because the marginal match rate or tax rate changes), and, hence, the taxes-paid measure, T, which, in turn, incorporates the dollar amount of the implicit tax liability from the employer-matching and tax subsidies, will change depending upon the budget-set segment as well. Therefore, full income is actually measured as "virtual" full income, y^{v}, according to the respective budget segment, where T^{v} denotes the associated implicit tax liability, which is calculated by numerically integrating the estimated kernel-smoothed implicit tax function described above.

The real relative price of an additional hour of leisure, ω, is,

$$\begin{aligned}\omega_{ijt} = (w_{it} / p_{it}^{c})[(1 - T_{Iit}(1 - Q_{y^{l}ijt}^{MAN} - \zeta_{y^{l}it} Q_{it}^{IRAC})) \\ + (\alpha_{y^{l}ijt}^{DB} W_{ijt}^{DB} + \alpha_{y^{l}it}^{SS} W_{it}^{SS} + F_{y^{l}ijt} + M_{y^{l}ijt}^{VOL} + Q_{y^{l}ijt}^{MAN} + M_{y^{l}ijt}^{MAN})]\end{aligned} \tag{54}$$

where the subscript y^{l} indicates the partial derivative with respect to labor-market earnings. In particular, the measure in (54), which appears as part of the first-order condition with respect to leisure, is substantially more complicated than, say, a simple measure of the after-tax wage, because of the greater detail in the tax treatment and the

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[37] The sum of consumption, real mortgage interest, and the market value of leisure to measure full income directly was not used for two reasons. First, as discussed in the data section, the timing of the consumption and hours' information in the first wave of the HRS did not coincide with the timing of the income and contribution data. Second, the HRS only gathered information on a limited number of categories of consumption expenditures (mainly food and housing). In (53), $\Delta A_t \equiv W_{t+1}^{TA} / R_t^{TA} - W_t^{TA}$.

presence of public and private pensions in the theoretical framework above.[38] In particular, the additional factor $1-Q_{y^l}^{MAN}-\zeta_{y^l}Q^{IRA}$ multiplying the marginal tax rate in (54) occurs because additional hours raise earnings and AGI, which may change the employee's mandatory pension contribution (which is deductible) and the deductibility of IRA contributions, respectively. The treatment of public and traditional private pensions in (2)-(3) adds $(w\ /\ p^c)[(\alpha_{y^l}^{DB}W^{DB}+\alpha_{y^l}^{SS}W^{SS}+F_{y^l}+M_{y^l}^{VOL}+Q_{y^l}^{MAN}+M_{y^l}^{MAN})]$ to (54), which represents the effect of additional earnings (through greater hours) on DB and non-401(k) DC pension accruals, Social Security accruals, 401(k) employer matching of voluntary contributions, mandatory employee contributions, and employer matching of mandatory employee contributions. For example, if the employer contributes five percent of earnings to a pension plan as part of a non-401(k) DC plan (such as a money purchase plan), then the opportunity cost of leisure will depend not just on the net wage, but also the lost employer contribution, F_{y^l}.

Unfortunately, the explanatory variables in (51) have components based on choice variables. Therefore, the instrument set, $\mathbf{Z}$, includes the vector of demographics, $\mathbf{z}$, and three additional variables, Z_{it-1}^{FC}, $p^{mz}\cdot p^{IRAz}$, and p^{401kz}: the first is a dummy variable if the household was in poor financial condition in 1990, and the second and the third are based on "first-dollar" match and marginal tax rates for a synthetic taxpayer in 1989. Appendix A gives a detailed description of the construction of the instruments, but the methods are highlighted here briefly.

[38] The data are cross-sectional for 1991. The time subscript is maintained to distinguish the timing of the lagged variables used to construct the instruments.

There were three important considerations in constructing the instruments. First, the instruments were drawn from the information set Ω_t. Because t is 1991, all information from 1989 and 1990 is in the information set and orthogonal to decisions made in 1991 under rational expectations. Second, because the observed marginal match and tax rates depend upon 401(k) and IRA contributions, $p^{mz} \cdot p^{IRAz}$ and p^{401kz} are based on first-dollar measures: the employer match on the first dollar contributed and the marginal tax rate at which the first dollar contributed is deductible (which equals the tax rate on the last dollar of earnings). Third, to minimize dependence on individual-specific income and family size that might be correlated with saving behavior, the first-dollar rates were calculated for a synthetic individual of each marital status assumed to have no capital income, no children, under age 65, and taking the standard deduction—where marital status is assumed exogenous—with synthetic annual labor earnings constructed as follows: individuals were divided into cells based on exogenous demographic characteristics, and the cell mean gross hourly wage rate, $\overline{w}_{\bullet t-2}$, was multiplied by 2,000 annual hours, $\overline{H}$. Let the subscript $\bullet$ denote a synthetic measure and the superscript 0 denote a first-dollar measure, then

$$p^{mz}_{\bullet j} \cdot p^{IRAz}_{\bullet t-2} \equiv (1 + M^{V0}_{Q^{401k}{}_{\bullet j}}) \cdot (1 - T^{0}_{I \bullet t-2} \zeta^{0}_{\bullet t-2}) \tag{55}$$

and

$$p^{401kz}_{\bullet t-2} \equiv 1 - T^{0}_{I \bullet t-2} (1 - \zeta^{0}_{y^l{}_{\bullet t-2}} \overline{Q}^{IRA}), \tag{56}$$

where $\overline{Q}^{IRA}$ was set to one dollar for all individuals.

There are two primary sources of variation in the instruments. First, p^{mz} varies by plan, j. That is, it is assumed that the variation in matching schedules across plans is exogenous. Second, p^{IRAz} and p^{401kz} vary across synthetic individuals because the tax function is non-linear in income and marital status. It is important to note that the tax function for Δp in the endogenous variable $\Delta p / y^v \ln(\omega)$, is based on the tax system in 1991, but the tax function for the instruments is *different* because it is based on the tax system in 1989 (indicated by the subscript $t-2$ in (55)-(56) above). Figure 2 plots the federal marginal tax rate by real AGI (in 1991 dollars) for a single individual under 65 in 1989 and 1991. For individuals with AGI below $50,000, the functions are essentially the same, but differ for those above this level. Specifically, above this income level in 1989, the marginal tax rate increased from 28 to 33 percent due to the phase-out of the personal exemption. However, the Budget Act of 1990 raised the top marginal tax rate to 31 percent and changed the phase-out of the personal exemption. Therefore, the non-linearity in the instruments' tax function differs from that for the endogenous regressor due to the tax-law change, which is taken as exogenous to the individual.[39] About 15 percent of the sample is affected by this differential non-linearity in the instruments.

A final issue is that the sample is likely non-random because it is based on individuals for whom the HRS was able to obtain an employer-provided SPD for the 401(k) plan. Although previous pension studies using the HRS employer-provided SPDs have not corrected for selection because of the lack of plausible exclusion restrictions, two exclusion restrictions based on IRS Form 5500 data were used to estimate the model

[39] In principle, one could use state-level variation in state marginal income tax rates in the instrument. Unfortunately, the Memorandum of Understanding between the Social Security Administration and the University of Michigan concerning the use of restricted-access HRS data prevents the merging of any information based on state of residence to the Social-Security-covered-earnings and W-2 earnings files used in this analysis, so that it is not possible to construct the instruments in this manner. However, a weighted average state tax liability for the individual's Census division of residence was added to the total tax liability, T^v, calculated assuming hypothetical residence in each state in the division, weighted by that state's share of the division adult population in the individual's income group.

using a number of methods to correct for selection. The first exclusion is the incidence of pension-plan outsourcing by Census region, employment-size category, one-digit SIC code, and union status (union plan vs. non-union plan) cell in 1992, where outsourcing means the plan was administered by an entity other than the employer.[40] The intuition is that the HRS is less likely to obtain an SPD from the employer if (on average in its cell) plan administration is outsourced, because more than one contact is needed (first the employer, then the plan administrator) to receive the SPD.[41] The second exclusion is the incidence of pension-plan consolidation due to mergers and acquisitions by cell from 1988-1992. The intuition is that the HRS is less likely either to obtain an SPD from the employer or match it to the employee if (on average in its cell) there has been a lot of plan consolidation, because plan names and detail are often changed upon consolidation. Finally, two other variables based on HRS data were used as exclusions in the selection equations: dummies for whether the individual left the job because the business closed or was laid off, respectively. These help to measure whether the employer possibly was in financial difficulty at severance, which, if that resulted in a business failure, would have made it more difficult for the HRS to have obtained an SPD. The construction of the exclusions and the selection equation are discussed in detail in Appendix A.

VIII. Estimation Results

[40] There is a restricted-access HRS dataset that provides industry and occupation information at a finer level of detail than the one-digit level. Unfortunately, the Memorandum of Understanding between the Social Security Administration and the University of Michigan concerning the use of restricted-access HRS data prevents the merging of any information based on the more detailed industry and occupation data to the Social-Security-covered-earnings and W-2 earnings files used in this analysis, so that it is not possible to construct these exclusions more finely .

[41] It may well be that plans that are outsourced are better administered and therefore more likely to return the pension provider survey and SPD. However, this is likely more than offset because the SPD request is significantly less likely to get fulfilled with multiple entities to contact.

To begin, the parameters in (51)-(52) are estimated in the baseline specification by Tobit maximum likelihood, where $\varepsilon \sim N(0, \sigma^2)$, using the estimator of Newey (1986, 1987b). In the baseline specifications, the vector $\mathbf{x}$ in (51) is limited to a constant. If an increase in the employer match raises contributions, then the null hypothesis that $\delta_0 = \delta_{11} = \delta_{12} = \delta_{13} = \delta_{14} = \delta_{15} = \delta_2 = 0$ should be rejected, and the estimated elasticity of contributions to matching should be positive. Appendix B provides a detailed discussion of the relationship of the structure of the econometric model here to those in the non-linear-budget-set literature on the impact of taxes on labor supply and why the critique of MaCurdy, Green, and Paarsch (1990) is not a concern for the estimation in this paper. The interested reader is referred there for details. Appendix C provides estimates from *ad hoc* reduced-form models similar to those used in the previous literature for those readers who wish to compare the results in Table 6 with previous methods.

Panel A of Table 6 presents parameter estimates with standard errors in parentheses.[42] Panel D lists the set of additional explanatory variables in the models. Panel E shows the *p*-values for the tests of the null hypothesis that each of the key economic variables in (51)—the employer match rate, virtual full income, the net wage, and the marginal tax rate—have no effect on contributions.

Column 1 shows the Tobit parameter estimates for the baseline specification without instrumenting. The marginal effect for a one unit change in the match rate—defined as $\partial E(Q^{401k} | \mathbf{X}) / \partial M^V_{Q^{401k}}$, where $\mathbf{X}$ is the vector of all the explanatory variables in (51), which accounts for the fact that a change in the match may induce some

[42] The standard errors are calculated using the variance-covariance estimator of Newey (1986) with an analytic correction to account for the presence of the estimated selection-correction term in the model.

individuals to begin contributing to the 401(k)—represents an increase of one dollar of match per dollar of employee contribution and is shown in Panel B. The estimated marginal effect of an increase in the match rate by one dollar is a *reduction* in contributions of \$533 (measured in constant, calendar year 1991 dollars), which is statistically different than zero. In addition, the estimated marginal effect of virtual full income is positive, the *wrong* sign if consumption and leisure are normal goods, but not statistically significantly different than zero. Column 2 gives the instrumental-variable Tobit estimates. Now, the sign of the estimated marginal effect flips: an increase in the match rate of one dollar increases contributions by an estimated \$2,167 and is statistically different than zero. In addition, the income effect is negative, the correct sign, and indicates that an increase of one unit in virtual full income, which represents an increase of \$100,000, results in an estimated reduction in contributions of \$2,750 (or a marginal propensity to save through the 401(k) of 0.0275), which is statistically different than zero.

To get a better sense of the economic magnitude of these marginal effects, Panel E shows the estimated elasticity of contributions with respect to the employer match rate, $\eta_{QM} = [\partial E(Q^{401k} | \mathbf{X}) / \partial M^V_{Q^{401k}}] \cdot [M^V_{Q^{401k}} / E(Q^{401k} | \mathbf{X})]$, evaluated at the sample means. The estimated elasticity is 0.06, with a standard error of 0.03, and is statistically significantly different from zero. If the employer match were raised from twenty-five to fifty cents (i.e., doubled), contributions would rise by six percent.[43] Thus, contributions appear to be quite inelastic with respect to the employer match. The estimated income elasticity is -0.25, with a standard error of 0.06, and, in comparative terms, represents a fairly sizable income effect. The estimated net wage elasticity is -0.10 and statistically

[43] The estimated elasticities are very similar when based on individual characteristics.

different than zero; the estimated elasticity with respect to the marginal tax rate is 0.05 and statistically significant. Panel E also shows the decomposition of the total elasticities into elasticities along the extensive (participation) and intensive (contributions conditional on participation) margins, respectively, based on the method of McDonald and Moffitt (1980). Two-thirds of the total match rate elasticity occurs along the extensive margin and the other one-third on the intensive margin, which suggests that participation is relatively more responsive to variation in the employer match than contributions.

Panel C shows the parameter estimates on the exclusion restrictions in the selection model.[44] The *p*-value for the test of the null hypothesis that the exclusions jointly do not explain who has a matched SPD, though not shown in the table, is less than 0.01, which indicates the exclusions have predictive power for who is in the sample. In particular, greater plan outsourcing, consolidation, and business closure significantly decrease the likelihood of having a matched SPD. The parameter estimate on the selection term in the contribution equation is negative, and, based on the associated standard error, the null hypothesis of no selection bias can be rejected at the 3 percent level of significance. This is evidence of selection: high savers are less likely to have an employer-provided SPD in the HRS, consistent with the reduced-form analysis of Gustman and Steinmeier (1999). To gauge whether this selection is economically important, column 3 shows the IV Tobit results without selection correction. The estimated contribution elasticity with respect to the match rate in Panel E is now 0.08,

[44] The estimation uses methods laid out in Vella (1992) and Das, Newey, and Vella (2003) for the Tobit and SCLS estimators, respectively, to correct for potential sample selection. Parameter estimates for the full selection model are available upon request. The first-stage regressions for the endogenous variables in (51) were selection corrected as well.

with a standard error of 0.04, which is 25 percent higher than the selection-corrected elasticity of 0.06 in column, but the 95-percent confidence intervals around the selection- and non-selection-corrected estimates overlap substantially, so that it appears that the impact of selection is minimal in this application.[45]

Column 4 presents estimates using an instrumental-variable version of the Symmetrically Censored Least Squares (SCLS) estimator of Powell (1986) and Newey (1986) extended to allow for both upper and lower censoring and individual-specific upper limits, as dictated by these data. The primary advantage of this semi-parametric estimator is that it is robust to heteroskedasticity and any departures to normality that were assumed for the Tobit error term. The estimated marginal effect of an increase in the match rate by one dollar in column 4 is $963, but with a standard error of $873, is not precisely estimated. The estimated marginal income effect indicates an increase of $100,000 results in an estimated reduction in contributions of $2,020, which is statistically different than zero and is close in magnitude to the Tobit-estimated effect of $2,750 from column 2.[46]

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[45] This result is very robust to the set of exclusion restrictions employed.

[46] One drawback of the SCLS estimator is that the total, extensive-, and intensive-margin match rate elasticities shown in panel F cannot be calculated from the estimates. In addition to SCLS, the specifications were estimated using an instrumental-variable version of Powell's CLAD estimator extended to allow for two limits, with the upper limit varying across individuals, as is dictated by these data, accounting for selection using the method of Blundell and Powell (2004) for semiparametric estimators. The CLAD point estimates produced marginal effects that were very similar to the SCLS and Tobit estimated marginal effects shown in Table 6. However, when bootstrapping the standard errors, the estimator frequently failed to converge. In particular, as is well-known for LAD estimators, the estimator can fail to converge when the specification becomes more and more saturated with control variables and the cells become thin, as in the richer specifications in subsequent columns in Table 6. This is a particularly severe problem for CLAD estimators that rely on iterative trimming of the data when the analysis dataset is not large, because the estimation dataset becomes smaller with each iteration, even though the number of control variables remains the same. In this application, the estimation algorithm trimmed the original 1,042 observations down to around 400 by the last iteration. For this reason, reliable bootstrapped standard errors were not able to be obtained for CLAD estimates and they are not shown in Table 6. Currently, we are experimenting with alternative estimators proposed by Blundell and Powell (2004), Honore and Powell (2002), and Honore and Hu (2004).

Robustness Checks and Extensions

There are two practical concerns in the estimation that fall outside of the scope of the theoretical framework. First, firms may offer matching as a way to try to avoid failing federal pension non-discrimination rules because they have low-saving employees (McGill, *et al*., 1996). This would tend to bias downward the estimated match elasticity. Second, firms that match may adopt other plan features to stimulate employee saving (e.g., allow for borrowing against plan balances, self-directed investment, offer after-tax saving options, offer retirement seminars, etc.) or offer different fringe benefit packages that might affect saving behavior than firms that do not match. This would tend to bias upward the estimated match elasticity.

The reduced-form relationship between employer match rates and these factors using the HRS data was examined in a companion paper, Engelhardt and Kumar (2004a). As described there, the non-discrimination rules are set up so that employers with a greater proportion of workers with earnings large enough to be deemed "highly-compensated employees" under federal law face greater pressure to meet non-discrimination rules if they offer a 401(k). In particular, a variable that measured the share of workers with earnings above the federal threshold for the definition of a "highly-compensated" employee under federal non-discrimination regulations in the respondent's Census-region-by-employment-size-category-by-one-digit-SIC-code-by-union-status cell in 1989 was constructed from the March CPS. This measure was then weighted by the difference in combined federal and state marginal tax rates on earnings for the median highly- and non-highly-compensated workers in the cell to reflect the value a highly-compensated worker would put on a dollar of tax-deferred salary through a 401(k) relative to that for a non-highly-compensated worker. This tax-difference-weighted share was used as a measure of the non-discrimination "pressure" faced by the typical employer in the respondent's cell in a reduced-form model of the determinants of match rates in the HRS data.

The estimation results in Engelhardt and Kumar (2004a) showed that the measure of pressure and other plan characteristics were highly significant. For example, the greater the pressure (tax-difference-weighted share) the more likely the respondent's plan offered a match and the higher the match rate. Also, plans that allowed borrowing, self-directed investment, had other traditional features, had limits less than the federal limit, and after-tax saving options had significantly higher first-dollar employer match rates, as was suggested in the comparison of unconditional means in Table 3 of the current paper.[47]

With this in mind, two groups of additional explanatory variables were included in the vector $\mathbf{x}$ in (51) for the contribution specification in Table 6: 1) *fringe benefits offered*: dummy variables for whether the firm offered long-term disability and group term life insurance, respectively, as well as the number of health insurance plans, number of retiree health insurance plans, weeks paid vacation, and days of sick pay; 2) *other plan characteristics*: dummy variables for whether the 401(k) allowed borrowing, hardship withdrawals, self-directed investment, had an after-tax saving option, a 401(k)

[47] Another potential concern is that high-saving individuals, such as those with long horizons, might sort to firms that offer employer matching contributions (Ippolito, 1997). This would tend to bias upward the estimated elasticity of voluntary contributions to matching. However, the estimation results in Engelhardt and Kumar (2004a) showed no correlation of the employer match rate with measures of the demographics and horizon and offered no support for endogenous sorting.

contribution limit less than the federal limit, respectively, whether the firm offered other traditional pensions, and the measure of non-discrimination "pressure" described above.

Columns 5 and 6 in Table 6 show the Tobit and SCLS estimation results for this specification, respectively. In particular, in column 5, Panel B, the estimated Tobit marginal effect of a one-dollar increase in the match rate is an increase in contributions of \$1,639, whereas the comparable SCLS marginal effect, shown in column 6, Panel B, is \$1,326, both of which are somewhat lower in magnitude than the associated estimated marginal effects in columns 2 and 4 (although there is substantial overlap in the 95-percent confidence intervals). However, the estimated match rate elasticity, shown in column 5, Panel E, is 0.07, compared to 0.06 in column 2, so that the addition of the fringe benefit and other plan characteristics does not have an important impact on the results. The estimated marginal effects for income, net wage, and the marginal tax rate in columns 5 and 6, respectively, are very similar to those in the baseline specifications in columns 2 and 4, respectively.

In columns 7 and 8, a third set of explanatory variables was added to $\mathbf{x}$ in (51): 3) *additional employment characteristics*: dummy variables for both the worker and spouse for whether the firm offered a retirement seminar, discussed retirement with co-workers, whether responsible for the pay and promotion of others, the number of supervisees, spousal pension coverage, as well as controls for firm size, Census division, and union status. In addition, these additional employment characteristics were interacted with the fringe benefit and plan characteristics described above to allow a more flexible functional form for $\alpha\mathbf{x}$ in (51). The estimated marginal effects for the match rate, income, net wage, and marginal tax rate for the Tobit and SCLS estimators in columns 7 and 8, respectively, are very similar to those in the baseline specifications in columns 2 and 4, respectively. The estimated elasticity of contributions to the match rate rises to 0.12 in column 7, Panel E.

Finally, to allow for a significantly more flexible functional form for $\alpha\mathbf{x}$ in (51), in columns 9 and 10, occupation dummies were added; the fringe benefit, plan, and other employment characteristics were interacted with occupation; and, the other plan characteristics were interacted with the fringe benefit variables. Hence, the specifications in columns 9 and 10 are essentially fully interactive in the elements of $\mathbf{x}$. The estimated elasticity of contributions to the match rate rises to 0.17 (column 9, Panel E), with roughly two-thirds of this effect on the extensive margin and one-third on the intensive margin. Because the Tobit elasticities are non-linear in the match rate, Figure 3 shows the relationship between the total match elasticity and the level of the match. The elasticity is increasing in the match rate. At match rates of 10, 25 (roughly the sample mean for all plans), and 50 percent (the sample median for plans with matches), the elasticity is 0.06, 0.13, and 0.23, respectively.

Newey (1987b) proposed a specification test for the distributional assumption of normality and homoscedasticity in the presence of endogenous variables. The test is based on the difference of an estimator that is dependent on these assumptions and another estimator that is consistent if these distributional assumptions are not valid. For the richest model, shown in the last two columns of Table 6, the Hausman statistic was 96.6 with a *p*-value of 0.99 for the test of the null hypothesis of no systematic differences between the instrumental variable two-limit Tobit estimator and the SCLS estimator.

As noted in section III, the demographic characteristics enter (51) parsimoniously, in a manner that allows the impact of employer matching to be heterogeneous across demographic groups. Across columns in Table 6, the two demographic groups for which the employer match consistently appears to have statistically significant differential effects on contributions are females and the relatively highly educated. In particular, females (conditional on marital status) are less responsive to the match, and the responsiveness to the match rises with education.

Table 7 shows the estimated marginal effect and elasticities for the employer match rate by sex and education group for the richest specification, shown in column 9 of Table 6. In columns 1 and 2 of Table 7, an increase in the match rate of one dollar increases contributions of males and females by an estimated \$2,568 and \$1,602, respectively. However, the estimated elasticities are very similar for the sexes. In columns 3-7, the marginal effect rises sharply by education group; however, again, when measured in terms of the elasticities, there is little difference across education group, primarily because the expected contribution, $E(Q^{401k}|\mathbf{X})$, that enters the denominator for the elasticity formula, also rises sharply with education, causing the elasticities to be relatively flat across groups.

IX. Summary and Implications

Previous studies have produced a puzzling array of estimates of the impact of employer matching on 401(k) saving. This probably stemmed from the use of less than ideal data and, more importantly, the failure to incorporate into estimation match-induced kinks in the budget set. Overall in this analysis, based on the life-cycle consistent specification derived, the estimated elasticity of 401(k) saving with respect to the match rate ranged from 0.06-0.17, and, hence, 401(k) saving was quite inelastic with respect to employer matching.

There are two potential implications of these findings. First, a number of commonly-advocated reforms to Social Security call for the introduction of voluntary private accounts, to which individuals could choose to contribute additional funds toward Social Security. Under some proposals, the federal government would match those contributions as an incentive. In designing such a system, it would be instrumental for policy makers to know how individual contributions would respond to the government match. Clearly, much could be learned in this context from the experience of employer matching for 401(k)s. This is examined in more detail in a companion paper (Engelhardt and Kumar, 2004b). Second, a number of prominent companies have reduced or eliminated matching contributions recently due to declining profits. Although it remains to be seen if this is a long-term trend, understanding the impact of matching is critical to understanding the impact of these changes on retirement income security for a workforce increasingly dependent on 401(k) plans for retirement. The fact that the estimated response of contributions to the employer match was quite inelastic suggests that overall 401(k) activity at these firms might not be greatly affected by these changes in matching.

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Appendix A

This appendix describes the construction of and gives background on the analysis dataset. Additional detail can be provided upon request.

The sample consists of 1,042 individuals from wave 1 of the HRS who were employed in 1991, eligible for a 401(k), whose employer provided a SPD for the plan, and who had linked administrative W-2 and Social Security earnings data. The restricted-access employer-provided SPDs are distributed as the *HRS Wave 1 Pension Plan Detail Data Set* (Gustman, Mitchell, Samwick, and Steinmeier, 1999). This dataset contains plan type, eligibility rules, benefit formulae, employer contribution and matching formulae, early and normal retirement dates, and other information described in the SPD, but not any information for individual employees. The W-2 data are distributed as the *HRS Wages and Self-Employment Income in Covered and Non-Covered Jobs* dataset (Mitchell, Olson, and Steinmeier, 1996). The dataset is a cross-section for 1991 because even though there are earnings and deferral data prior to 1991, there are no data on other income and wealth needed to construct full income prior to 1991 in the HRS. Some of the individuals in the sample worked in 1991 but were retired at the time of the first interview in 1992. Exclusion of these individuals had no impact on the estimation results.

There are four types of employer matching: fixed-rate, discretionary, profit-sharing, and variable-rate matching. Engelhardt and Kumar (2003) discussed these in detail. Because the extent of matching is not always known in advance to employees making deferral decisions in profit-sharing and discretionary plans, these plans were not included in our sample. The SPDs were used to construct the complete schedule of employer matching contributions for each individual in our sample and applied all relevant restrictions on plan eligibility in the SPD, including those due to tenure, hours, earnings, age, and vesting of the employer matching contributions.

Because workers' budget sets can have multiple kinks and, therefore, multiple points of non-differentiability, from changes in match and marginal tax rates, a smooth, differentiable budget set around all kink points was constructed, following the methodology of MaCurdy, Green, and Paarsch (1990). Specifically, kernel regression of the implicit subsidy from employer matching and tax deductibility on the set of potential contributions from 0 to \$9500 (the federal maximum contribution in 1991) by \$50 increments was used to smooth the budget set, using the Gaussian kernel, $K(z) = (1/\sqrt{2\pi})e^{-z^2/2}$, with bandwidth chosen by Silverman's rule of thumb, $h = 0.9m/n^{1/5}$, where $m = \min(\sqrt{\mathrm{var}_x}, iqr_x/1.349)$ and iqr_x is the inter-quartile range. A smooth marginal implicit subsidy function was constructed from the kernel-regression estimates. The respondent-reported income in wave 1 of the HRS referred to behavior in calendar year 1991; hence, ΔA was formed by using taxable wealth in 1992 taken from wave 1, taxable wealth in 1991, which was capitalized from 1991 capital income, a technique is commonly used in the literature, and R^{TA} constructed from a weighted-average gross return based on returns in Ibbotson (2003). Tax rules from 1991 were used to construct an IRA phase-out calculator to determine ζ and ζ_{y^l}, and household income, tax, and demographic data and NBER's TAXSIM calculator (Feenberg and Coutts, 1993) were used to construct marginal tax rates and taxes paid for each household. The estimated kernel-regression function was numerically integrated to obtain the dollar amount of implicit subsidy for every level of potential 401(k) contribution, which, with full income, was used to construct virtual full income along all portions of the budget sets.

For the private and public pension components in ω in (54), individuals were divided into cells based on exogenous demographic characteristics and the Social Security covered earnings from 1951-1991 and W-2 earnings records from 1980-1991 were used to calculate earnings histories for a synthetic-cell individual. These synthetic earnings histories were input as follows: 1) into the University of Michigan's *Pension Estimation Program* to calculate defined benefit pension wealth, W^{DB}, accrual, α^{DB}, and change in accrual for additional earnings, $\alpha^{DB}_{y^l}$, for individuals with DB plans; 2) into the *HRS DC/401(k) Calculator* (Engelhardt, 2004) developed to calculate for individuals with defined contribution plans their DC pension wealth, W^{DC}, non-matching contributions and the effect of additional earnings thereon, F and F_{y^l}, respectively; the impact of additional earnings on employer match on voluntary contributions, $M^{VOL}_{y^l}$; required 401(k) contributions and the impact of additional earnings thereon, Q^{MAN} and $Q^{MAN}_{y^l}$, respectively; and, 3) into the Social Security benefit calculator developed by Coile and Gruber (2000) to calculate Social Security wealth, W^{SS}, accrual, α^{SS}, and change in accrual for additional earnings, $\alpha^{SS}_{y^l}$. The effect of additional earnings on the employer match to voluntary contributions, $M^{VOL}_{y^l}$, was calculated assuming a 401(k) contribution of 50 dollars for all individuals (regardless of actual contribution level).

Finally, the sample is likely non-random because it is based on individuals for whom the HRS was able to obtain 1) an employer-provided SPD for the 401(k) plan, and 2) permission from the individual to match SSA covered earnings and IRS W-2 earnings histories. To understand the exclusion restrictions that were developed, it is useful to note the manner in which the HRS obtained the SPDs and administrative earnings data. The HRS asked all respondents who reported being in a (current or past) pension-covered job to provide the name and address of the employer. To maintain respondent confidentiality, the HRS attempted to contact the employer, not about the respondent's pension(s), but more generally as part of a survey of pension providers in which the HRS requested copies of SPDs for the universe of pensions the employer provided (to all employees). The HRS then "matched" from this universe the appropriate pension(s) to the respondent based on the respondent's characteristics, e.g., union status, method of pay (hourly, salaried, commission, piece rate), occupation, tenure, etc. The "match" rates were well below 100 percent: 65 percent of those currently working in pension-covered jobs, 66 percent for the last job for those not working, and 35 percent for jobs held five years or longer prior to the current (last) job for those working (not working).

There are a number of important reasons for the failure to match an SPD to the respondent. First, the respondent may not have given correct employer name and address. Second, the HRS may have failed to receive the SPD because the employer may have refused to comply with the pension provider survey, the employer could not be located at the address given, or the employer went out of business or merged with another company and no longer existed under the name given by the respondent. Third, the employer may have submitted an SPD, but the HRS was unable to match the SPD to the respondent based on the plan detail and the respondent's characteristics. This is less likely for union and public sector workers, who are easy to identify and whose plans are easy to obtain, and more likely for workers whose employers had undergone mergers and acquisitions with subsequent plan modifications.

The exclusion restrictions were constructed as follows. First, Form 5500 data for 1988-1992 from the Department of Labor, Employee Benefit Security Administration, on the universe of pension plans with 100 or more participants and a 5 percent random sample of plans with less than 100 participants were obtained. Second, plans were divided into cells defined by Census region, employment size category, one-digit SIC code, year, and union status (union plan vs. non-union plan). The first exclusion is the incidence of pension plan outsourcing by cell in 1992, where outsourcing means the plan was administered by an entity other than the employer (weighted using sampling weights provided by DOL). The intuition here is that the HRS was less likely to have obtained an SPD from the employer if (on average in its cell) plan administration was outsourced, because more than one contact was needed (first the employer, then the plan administrator) to have received the SPD. (It may well have been that plans that were outsourced were better administered and, therefore, employers that outsourced were more likely to have returned the pension provider survey. However, this was likely more than offset because the SPD request was significantly less likely to have been fulfilled when multiple entities needed to be contacted.) The second exclusion was the incidence of pension plan consolidation due mergers and acquisitions by cell from 1988-1992. The intuition here is that the HRS was less likely either to have obtained an SPD from the employer or to have matched it to the employee if (on average in its cell) there had been a lot of plan consolidation, because plan names and detail were often changed upon consolidation. Two other variables were used as exclusions for pensions on past jobs in our selection equations: dummies for whether the individual left the job because the business closed or was laid off, respectively. These helped to measure whether the employer possibly was in financial difficulty at severance, which, if that resulted in a business failure, would have made it more difficult for the HRS to have obtained an SPD.

Appendix B

This appendix discusses in detail the nature of the error term in the econometric model presented in (51)-(52) in the text and why the critique of MaCurdy, Green and Paarsch (1990) does not apply to the estimation in this paper.

From equation (51) in the text, the latent-variable econometric model was

$$Q_{ijt}^{VOL*} = \delta_0 \frac{\Delta p_{ijt}}{y_{ijt} \ln(\omega_{ijt})} + \sum_m \delta_{1m} \frac{\Delta p_{ijt}}{y_{ijt} \ln(\omega_{ijt})} z_{i,m} + \delta_2 \kappa_{ijt} + \alpha \mathbf{x}_{ijt}, \tag{B.1}$$

where i and j index individuals and 401(k) plans, respectively. Although most non-linear-budget-set models, such as those applied to taxes and labor supply, assume a two-error structure, in which one source of error is preference heterogeneity and the other is optimization error, the error term in (B.1) is optimization error, ε; hence, (B.1) is a single-error model.

The primary reason why, in a strict sense, preference heterogeneity does not appear in the form of a second error term in the model is because the functional form in (B.1) was based on the assumed expenditure function,

$$\ln e(\mathbf{p}, V; \mathbf{z}) = \ln[a(\mathbf{p})] + \Psi(\mathbf{z})^{-1} V \ln[b(\mathbf{p})], \tag{B.2}$$

associated with the class of PIGLOG utility functions (Muellbauer, 1976) that have been used extensively in the literature on consumption, and the typical manner in which heterogeneity has been introduced in studies with the PIGLOG has been through the $\ln[a(\mathbf{p})]$ term. From (B.2), the indirect utility function is

$$V(\mathbf{p}, y; \mathbf{z}) = \Psi(\mathbf{z}) \cdot \frac{\ln(y) - \ln[a(\mathbf{p})]}{\ln[b(\mathbf{p})]}. \tag{B.3}$$

However, equation (40) in the text, shown here without the time subscripts,

$$\eta^L - \eta^0 = \Delta p \cdot V_y + p^m (\upsilon^L - \upsilon^0), \tag{B.4}$$

was derived directly from the first-order conditions, and the basis for (B.1), required V_y, and, from (B.3), the $\ln[a(\mathbf{p})]$ term drops out upon differentiating the indirect utility function with respect to y, so that the main channel through which preference heterogeneity typically has appeared in other contexts has been eliminated here. Indeed, what is left is

$$V_y = \Psi(\mathbf{z}) \frac{1}{y \ln[b(\mathbf{p})]}, \tag{B.5}$$

But b is homogeneous of degree zero and modeled as a Cobb-Douglas price aggregator

$$b(\mathbf{p}) = \prod_k p_k^{\beta_k}, \tag{B.6}$$

across the k goods that enter the direct utility function, where

$$\sum_k \beta_k = 0, \tag{B.7}$$

and so there is no natural role for preference heterogeneity to enter $\ln[b(\mathbf{p})]$.

This highlights a somewhat subtle assumption underlying the identification of the model: the instrumental variables are valid as long as preference heterogeneity does not enter V_y. To see this more clearly, write (B.1) more generally, and in accordance with (B.4) as

$$Q_{ijt}^{VOL*} = \delta_0 \Delta p_{ijt} V_{yijt} + \sum_m \delta_{1m} \Delta p_{ijt} V_{yijt} z_{i,m} + \delta_2 \kappa_{ijt} + \alpha \mathbf{x}_{ijt}, \tag{B.8}$$

If y and preference heterogeneity were not independent in the indirect utility function, so that the heterogeneity entered V_y, then because Δp pre-multiplies V_y in (B.8), the heterogeneity would enter the error term in (B.8), but also be pre-multiplied by Δp, and the instrumental variables, $p^{mz} \cdot p^{IRAz}$ and p^{401kz}, would not be uncorrelated with the error term and would no longer be valid.

That said, it is very important to understand that the independence of y and preference heterogeneity in the indirect utility function is a feature shared by many specifications of preferences commonly used in modeling consumption and labor supply. For example, the indirect utility function,

$$V(w, y) = e^{\beta w}\left(y + \frac{\alpha}{\beta} w - \frac{\alpha}{\beta^2} + \frac{\gamma}{\beta} \right) \tag{B.9}$$

that generates the linear labor-supply equation

$$H = \alpha w + \beta y + \gamma \tag{B.10}$$

that has been used extensively to model the impact of taxes on labor supply with non-linear budget sets, has the property that preference heterogeneity, embodied in γ in

(B.11), does not appear in V_y, yet it does appear in the hours equation (because it appears in V_w, and, therefore, through Roy's Identity, in hours). Other preference specifications that share this property include the PIGLOG (and related functions like AIDS), the linear expenditure system, quadratic indirect utility, and the preferences that yield the quadratic, semi-log, and share-linear-in-logarithms labor-supply functions (Stern, 1986). Random-utility models that assume additive error terms (e.g., $V + v$) also share this property.

There are a number of approaches to the estimation of non-linear budget set models, including the maximum-likelihood estimation pioneered by Burtless and Hausman (1978) and summarized in Hausman (1985), in which each segment and kink point on the individual's budget set has its own contribution to the likelihood function, and instrumental-variable techniques that linearize the budget set at the observed outcome to calculate the price and virtual income terms and then instrument to correct for endogeneity. Most applications of the former approach to the impact of taxes on labor supply specified a two-error model—one for heterogeneity and one for optimization error—and the early work utilized piece-wise linear budget constraints.

In a highly-cited paper, MaCurdy, Green, and Paarsch (1990) found that estimating the labor-supply function by maximum likelihood with heterogeneity error imposed restrictions on the estimated labor-supply function. Specifically, they found that for the statistical model to make sense and for the likelihood function to be well defined, the Slutsky restrictions must be imposed at the *interior* kink points, and that this explained why many studies using piecewise-linear budget sets, such as those summarized in Hausman (1985), found large compensated effects on labor supply. They argued that the reason why the restrictions are imposed is that the likelihood function in a piecewise-linear-budget-constraint setting contains $\Pr(\theta_i = 1)$, where θ_i indicates whether the individual locates on interior kink i. Letting the heterogeneity error term for segment i be denoted by v_i, $\Pr(\theta_i = 1)$ is negative unless the heterogeneity error component satisfies restrictions such that $v_i^U - v_i^L \geq 0$, where v_i^U, v_i^L are the upper and lower bounds of the heterogeneity error associated with a kink. These restrictions, in turn, placed restrictions on preferences that can be expressed mathematically as the imposition of the Slutsky restrictions at the interior kinks. MaCurdy, Green, and Paarsch (1990) went on to show that if the budget set were modeled as twice-continuously differentiable, such that kink points were smoothed, rather than as piecewise-linear, and the labor-supply parameters were estimated by maximum likelihood, that the differentiable-budget-set methodology still imposed restrictions on preferences at the smoothed kink points, but those restrictions were much weaker than those associated with the piecewise-linear method.

There are three important points about the critique of MaCurdy, Green, and Paarsch (1990). First, it only arises because of heterogeneity error; the presence of optimization error plays no role in the critique, and this was something that MaCurdy, Green, and Paarsch (1990) were very clear about in their exposition. Second, it applies to maximum-likelihood estimation; the use of alternative estimators circumvents this problem. An example of an alternative estimation approach is that described above: an instrumental-variable technique that linearizes the budget set at the observed outcome to calculate the price and virtual income terms and then instruments to correct for endogeneity. Third, it only applies to *interior* kink points. For example, probit maximum-likelihood estimation of the participation decision in the labor-supply context would not be subject to this critique. This also was something that MaCurdy, Green, and Paarsch (1990) were very clear about in their exposition.

The remainder of this appendix explains why the critique of MaCurdy, Green, and Paarsch (1990) does not apply in the current paper. To begin, and to be perfectly clear to the reader, note that the approach employed in estimating the parameters of the 401(k) contribution decision is not the Hausman-type maximum-likelihood estimation, in which each segment and kink point on the individual's budget set has its own contribution to the likelihood function. Instead, the approach taken was to use instrumental-variable estimation that linearized the budget set at the observed outcome to calculate the price and virtual income terms and then instrument to correct for endogeneity. So, it would appear that this critique would not be an issue because of the use of IV. However, because the dependent variable is limited in range, the parameters were estimated in the baseline specifications using the Tobit instrumental-variable estimator of Newey (1986), which, obviously, is a maximum-likelihood estimator. So, the use of IV *per se* cannot be used as a justification for why the critique does not apply in this case.

In particular, the 401(k)-contribution likelihood function is mathematically equivalent to the labor-supply likelihood function MaCurdy, Green, and Paarsch (1990) analyzed, but without the heterogeneity error. And therein lies the reason why their critique does not apply to the current analysis: there is no heterogeneity error in the 401(k) specification, only optimization error. The Slutsky restrictions implied by the critique derive solely from the presence of heterogeneity error, and MaCurdy, Green, and Paarsch (1990) are quite clear about this, so that their critique does not apply here.

Furthermore, there are three empirical robustness checks that indicate that the critique is not a concern. First, in a dual-error labor-supply model with heterogeneity and optimization error, the heterogeneity error captures the variation in the utility function across individuals, whereas the optimization error represents the deviation of actual hours from the desired hours. Bunching precisely at the kink points is an indication of the relative strength of the variation due to optimization error and heterogeneity: substantial bunching at kink points indicates that the heterogeneity error variance is large; relatively little bunching is evidence in favor of large optimization error and inconsequential heterogeneity error (Moffitt, 1986 and 1990; Triest, 1990). There is little evidence of bunching precisely at the kink points in the 401(k) data; instead, twenty-five percent of the sample observations are located within a fifty-dollar interval on either side of interior kink points. The difference between the true-kink contribution and the actual contribution is optimization error. One example of optimization error that is apparent in the data is the frequency with which contribution amounts that are chosen are "round numbers." For example, for an employee who earns, say, $40,550 annually in a typical 401(k) plan with a match of 50 cents per dollar up to 6 percent of pay, the kink point occurs at a contribution of $2,433. However, it is much more common in the data to see a contribution of \$2,400, \$2,450, or even \$2,500, i.e., the nearest round number to the kink in \$50 or \$100 increments. Of course, such rounding would not generate necessarily normally-distributed errors, but, again, it should be emphasized that is just one potential source of optimization error, and the sum totals from all individual and plan sources of optimization error are assumed to be normally-distributed in the Tobit specifications. Overall, the fact that few individuals locate exactly at the kink suggests that the heterogeneity error is insignificant and is evidence in favor of optimization error. Second, the critique only applies at interior kinks and, therefore, probit maximum-likelihood estimation of the 401(k) participation decision is not subject to this critique. In particular, the probit estimates of the parameters in

$$D_{ijt}^{VOL*} = \delta_0 \frac{\Delta p_{ijt}}{y_{ijt} \ln(\omega_{ijt})} + \sum_m \delta_{1m} \frac{\Delta p_{ijt}}{y_{ijt} \ln(\omega_{ijt})} z_{i,m} + \delta_2 \kappa_{ijt} + \alpha \mathbf{x}_{ijt}, \tag{B.13}$$

where D^{VOL*} indicates whether a contribution was made, will not be affected by the critique and can be compared to the extensive margin effects from the McDonald-Moffitt decomposition of the structural Tobit parameter estimates as a robustness check. For the specification shown in column 9 of Table 6, the probit estimated match-rate elasticity is 0.08, which is very close to the extensive margin elasticity of 0.10 from the McDonald-Moffitt decomposition. The null hypothesis that the two elasticities are the same cannot be rejected, so that there is no support for the critique. Finally, the critique only applies to maximum-likelihood estimation, so that the IV-SCLS estimates can be used as a robustness check, as they are not subject to the critique. As shown in Table 6, the parameter estimates from the IV-Tobit and IV-SCLS specifications were very similar, and the implied estimated elasticities of the latent variable to the match rate were similar as well. So, there is no empirical support for the critique.

Appendix C

To compare the non-linear budget set approach with that from the previous literature, this appendix gives the estimation results from a series of *ad hoc* reduced-form specifications similar in spirit to those in the literature, using the same estimators as used in the literature, but with the HRS data. Selected parameter estimates are shown in Appendix Table C-1 below. Standard errors are in parentheses.

Table C-1. Selected Parameter Estimates from Specifications Similar to Previous Literature, Standard Errors in Parentheses

	(1)	(2)	(3)	(4)
	Estimator and Dependent Variable			
	Probit	OLS	OLS	One-Limit Tobit
Explanatory Variables	Dummy if Contributed	Contribution	Contribution	Contribution
Dummy if Plan Offers a Match	0.284 (0.175)	595.4 (209.7)	549.4 (206.9)	865.7 (331.7)
First-Dollar Match Rate	0.220 (0.228)	-255.2 (271.2)	-170.3 (266.3)	-42.6 (427.2)
Dummy if Female	0.316 (0.112)	389.1 (121.2)	304.7 (135.6)	639.3 (224.9)
Dummy if White	0.262 (0.107)	302.1 (133.2)	275.6 (131.2)	617.1 (223.2)
Education (Years)	0.043 (0.018)	105.8 (22.4)	89.4 (22.5)	147.6 (38.2)
Earnings Entered as	Quartic	Linear	Quartic	Quartic
Match Rate Elasticity	0.037 (0.043)	-0.042 (0.045)	-0.028 (0.044)	-0.005 (0.003)
R^2	---	0.31	0.34	---

Note: This table shows selected parameter estimates for *ad hoc* reduced-form models similar to those in the previous literature, estimated with the sample of 1042 individuals described in the text. Standard errors in parentheses. In columns 1, 3, and 4, earnings were entered as a quartic function and in column 2 linearly. Additional explanatory variables in the specifications included a quartic in age, married, number of children, spouse's education, a quartic in spouse's age, and a constant. For the Probit equation in column (1) and the Tobit equation in column (4) bootstrapped standard errors reported for the match rate elasticity.

Column 1 shows a probit specification for the decision to contribute to the 401(k) plan in 1991. In columns 2-4, the dependent variable is the dollar amount of contributions. Contributions are modeled as a function of earnings, demographics, a dummy for whether the firm matches contributions and the marginal match rate in column 2. Column 3 expands the specification to include quartic functions in age and earnings. Column 4 presents one-limit Tobit estimates. Like previous studies, all specifications in the table indicate that the presence of a match raises contributions. However, conditional on offering a match, the point estimates suggest that increases in the match rate may increase or decrease contributions, but none of these effects are statistically significant. The estimated elasticity of contributions with respect to the match rate (conditional on having been offered a match) is shown in the last row of the table for each specification, with standard errors in parentheses.

Table 1. Cap on Matching Contributions, as a Percentage of Pay, for Plans that Offer Employer Matching in the Analysis Sample

Cap on Employer Matching Contributions as a Percentage of Pay	(1) Number of Plans	(2) Percent of Plans	(3) Number of Individuals	(4) Percent of Individuals
Less than 2%	7	3.3	10	2.7
2	11	5.3	12	3.2
2.5	1	0.5	1	0.3
3	19	9.1	24	6.5
3.75	1	0.5	4	1.0
4	23	11.0	40	10.8
5	17	8.1	53	14.2
5.5	1	0.5	1	0.3
5.7	1	0.5	1	0.3
6	56	26.8	109	29.3
Greater than 6%	32	15.3	57	15.3
No Cap	41	19.6	60	16.1
Total	209	100.0	372	100.0

Note: Authors' calculations from the HRS restricted-access pension plan data for the 209 plans associated with the 372 of the 1,042 HRS individuals in the analysis sample in plans with matching provisions.

Table 2. Distribution of First-Dollar Match Rates as a Percentage of Contributions

(1) First-Dollar Match Rate (%)	(2) Number of Plans	(3) Percent of Plans	(4) Number of Individuals	(5) Percent of Individuals
0 to 24	9	4.3	11	3.0
25	23	15.3	43	11.6
26 to 49	5	2.4	9	2.4
50	90	43.1	143	38.4
51 to 99	22	8.1	34	12.4
100	57	27.2	116	31.2
200	3	1.4	4	1.1
Total	209	100.0	372	100.0

Note: Authors' calculations from the HRS restricted-access pension plan data for the 209 plans associated with the 372 of the 1,042 HRS individuals in the analysis sample in plans with matching provisions.

Table 3. Sample Means of Selected Variables in the Empirical Analysis Sample,
Standard Deviations in Parentheses, Medians in Square Brackets

	(1)	(2)	(3)	(4)	(5)
Variable	Full Sample	Subsample without Matches	Subsample with Matches	Subsample with Positive Contributions	Subsample with Zero Contributions
401(k) Contributions (in 1991 dollars)	1377	1232	1640	2446	0
	(1920)	(1895)	(1938)	(1982)	(0)
	[500]	[100]	[900]	[1892]	[0]
Match Rate (in percent)	23	0	65	28	17
	(37)	(0)	(32)	(38)	(33)
	[0]	[0]	[50]	[0]	[0]
After-Tax Wage (in 1991 dollars per hour)	10.04	10.09	9.96	10.91	8.91
	(5.55)	(5.56)	(5.54)	(5.96)	(4.75)
	[8.92]	[9.12]	[8.51]	[9.66]	[8.23]
Age (years)	54.9	54.9	54.8	54.7	55.1
	(5.2)	(5.1)	(5.4)	(5.0)	(5.5)
	[55.0]	[55.0]	[55.0]	[55.0]	[55.0]
Education (years)	13.3	13.5	13.0	13.8	12.7
	(2.7)	(2.7)	(2.6)	(2.5)	(2.7)
	[13.0]	[13.0]	[12.0]	[14.0]	[12.0]
Percent Female	47	47	47	48	45
Percent White	82	81	85	86	78
Number of Dependents	0.70	0.68	0.75	0.71	0.70
	(0.93)	(0.93)	(0.94)	(0.95)	(0.91)
	[0.0]	[0.0]	[0.0]	[0.0]	[0.0]
Percent Married	80	79	82	81	79
Spouse's Education (Years)	10.6	10.6	10.6	11.0	10.1
	(5.5)	(5.7)	(5.2)	(5.5)	(5.5)
	[12.0]	[12.0]	[12.0]	[12.0]	[12.0]
Percent with Plans that Allow Borrowing	36	19	68	42	29
Percent with Plans that Allow Hardship Withdrawals	4	4	5	6	2

Variable	(1) Full Sample	(2) Subsample without Matches	(3) Subsample with Matches	(4) Subsample with Positive Contributions	(5) Subsample with Zero Contributions
Percent with Plans that Allow Self-Directed Investment	63	46	92	66	58
Percent with Other Pensions at the Firm	47	53	34	45	48
Percent with Plan Limit less than Federal Limit	80	73	92	76	85
Percent with Plan that Allows After-Tax Saving	23	9	47	26	18
Percent that had Employer-Sponsored Retirement Seminar	23	23	23	25	20
Percent with a Spouse who has a Pension	39	39	38	42	35
Percent in a Union	34	39	27	28	43
Number of Observations	1042	670	372	588	454

Note: Authors' calculations based on the sample of 1042 HRS individuals working in 1991 with matched employer-provided pension plan data and W-2 data, excluding those in plans with discretionary and profit-sharing-based employer matching provisions, as described in the text.

Table 4. Mean Potential and Unused Employer Matching Contributions for the Sub-sample of Individuals Eligible for Employer Matching Contributions, Standard Deviations in Parentheses, Medians in Square Brackets

Variable	(1) Overall	(2) Sub-sample with Positive Contributions	(3) Sub-sample with Zero Contributions
Potential Employer Matching Contributions in 1991 dollars	1249 (1409) [939]	1362 (1153) [1068]	1021 (1804) [714]
Potential Employer Matching Contributions as a Percentage of Pay	3.8 (4.1) [3.0]	3.9 (2.8) [3.5]	3.8 (5.8) [3.0]
Unused Employer Matching Contributions in 1991 dollars	550 (1243) [205]	319 (741) [0]	1013 (1798) [710]
Unused Employer Matching Contributions as a Percentage of Pay	1.9 (4.0) [1.0]	1.0 (2.0) [0]	3.7 (5.9) [3.0]

Note: Authors' calculations based on the sub-sample of 372 HRS individuals working in 1991 eligible for employer matching contributions.

Table 5. Selected Parameter Estimates of the Effect of Liquidity Constraints on the Extent that Contributions are Less than the Cap, Standard Errors in Parentheses

	(1)	(2)	(3)	(4)
		Estimator and Dependent Variable		
		Probit	Tobit	Tobit
Explanatory Variables	Sample Mean	Dummy if Contribution Below the Cap	Dollar Amount that Contribution is Below the Cap	Amount Contribution is Below the Cap as a Percent of Pay
Dummy if Has a Home Equity Line of Credit	0.19	-0.524 (0.198)	-649.4 (171.3)	-0.014 (0.004)
Dummy if Financial Distress Due to Unexpected Medical Expenses	0.11	0.725 (0.291)	81.1 (202.4)	0.004 (0.004)
Dummy if Financial Distress Due to Unemployment	0.06	-0.569 (0.366)	-733.8 (360.8)	-0.017 (0.014)
Dummy if No Capital Income	0.23	0.552 (0.180)	220.1 (136.0)	0.009 (0.003)
Education (years)	12.97	-0.077 (0.036)	-41.4 (26.5)	-0.002 (0.001)

Note: This table shows selected parameter estimates from *ad hoc* reduced-form selection-corrected specifications of the impact of selected variables that proxy for the ability to borrow on whether and to what extent employer matching contributions go unused on the sub-sample of 372 individuals in 401(k) plans that offer employer matching. Standard errors are in parentheses. The fraction of individuals with contributions below the cap is 0.54. The average dollar amount that the contribution is below the cap is $953, which, for average pay of $33,377, represented 3 percent of pay. The specifications also control for age, dummy variables if married, white, home owner, and can rely on private income transfers under financial distress, as well as a constant. The selection equations for the specifications used the same exclusion restrictions as in Table 7 and explained in the text.

Table 6. Instrumental-Variable Parameter Estimates of 401(k) Contributions for Selected Variables, Standard Errors in Parentheses

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Estimator									
Explanatory Variable	Tobit with Selection	IV Tobit with Selection	IV Tobit without Selection	IV SCLS with Selection	IV Tobit with Selection	IV SCLS with Selection	IV Tobit with Selection	IV SCLS with Selection	IV Tobit with Selection	IV SCLS with Selection
A. Parameter Estimates										
$\Delta p / y \ln(\omega)$	-10776	21227	19628	5374	6590	-1588	1242	-9322	3859	-10828
	(3485)	(19592)	(18904)	(16856)	(13830)	(12261)	(14147)	(14025)	(14123)	(17952)
$Age \times \Delta p / y \ln(\omega)$	51	-284	-268	-91	-108	23	-54	72	-2	155
	(52)	(208)	(203)	(178)	(146)	(128)	(151)	(149)	(149)	(189)
$Female \times \Delta p / y \ln(\omega)$	156	-7420	-7205	-6068	-3746	-3972	-3336	-2464	-3197	-2686
	(541)	(2431)	(2321)	(2109)	(1872)	(1680)	(1934)	(1960)	(1787)	(2317)
$White \times \Delta p / y \ln(\omega)$	2179	-1056	-1161	259	-618	750	-607	838	-789	52
	(847)	(2735)	(2655)	(2517)	(2104)	(2074)	(2062)	(2244)	(2043)	(2757)
$Married \times \Delta p / y \ln(\omega)$	-118	-4504	-4263	-4280	-1704	-1919	-776	-173	107	1056
	(667)	(2841)	(2727)	(2449)	(1985)	(1793)	(1945)	(1943)	(1854)	(2357)
$Education \times \Delta p / y \ln(\omega)$	471	1495	1492	1405	1120	1025	1198	1277	681	1235
	(101)	(413)	(399)	(358)	(292)	(263)	(285)	(292)	(313)	(438)
κ	745	4420	4147	4169	2726	2589	1903	1679	1588	1792
	(115)	(1950)	(1885)	(1691)	(1438)	(1286)	(1413)	(1386)	(1486)	(1852)
B. Latent Marginal Effects										
Match Rate	-533	2167	2176	963	1639	1326	2084	2074	2113	3221
	(125)	(1039)	(1023)	(873)	(901)	(775)	(960)	(919)	(955)	(1269)
Full Income	13	-2750	-2656	-2020	-1857	-1641	-1795	-1708	-1697	-2204
	(56)	(669)	(633)	(586)	(550)	(504)	(551)	(565)	(567)	(797)
Net Wage	0.44	-90	-86	-66	-60	-53	-58	-55	-55	-72
	(1.82)	(22)	(21)	(19)	(18)	(16)	(18)	(18)	(18)	(26)
Marginal Tax Rate	-11.61	2384	2303	1752	1610	1423	1557	1482	1471	1911
	(179)	(772)	(754)	(658)	(538)	(477)	(562)	(566)	(533)	(743)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
C. Parameter Estimates for Selection Term in Contribution Equation and Exclusion Restrictions in the Selection Equation										
Selection Term in Contribution Tobit	-6608 (3588)	-7227 (3363)		-7224 (3268)	-6573 (3188)	-5670 (3355)	-4768 (2719)	-6639 (3588)	-4683 (2729)	-14078 (5373)
p-Value on Selection Term	0.06	0.03		0.04	0.04	0.09	0.08	0.07	0.09	0.01
Selection-Equation Exclusions:										
Plan Administration Outsourcing	-0.08 (0.03)	-0.08 (0.03)		-0.08 (0.03)	-0.09 (0.03)	-0.09 (0.03)	-0.12 (0.03)	-0.12 (0.03)	-0.13 (0.03)	-0.13 (0.03)
Plan Consolidation	-0.12 (0.04)	-0.12 (0.04)		-0.12 (0.04)	-0.15 (0.04)	-0.15 (0.04)	-0.19 (0.04)	-0.19 (0.04)	-0.19 (0.04)	-0.19 (0.04)
Left Job Due to Business Closure	-0.068 (0.017)	-0.068 (0.017)		-0.068 (0.017)	-0.05 (0.02)	-0.05 (0.02)	-0.05 (0.02)	-0.05 (0.02)	-0.05 (0.02)	-0.05 (0.02)
Left Job Because Laid Off	-0.02 (0.24)	-0.02 (0.24)		-0.02 (0.24)	-0.01 (0.02)	-0.01 (0.02)	-0.01 (0.02)	-0.01 (0.02)	-0.001 (0.02)	-0.001 (0.02)
D. Additional Controls										
Fringe Benefit, and Plan Characteristics?	No	No	No	No	Yes	Yes	Yes	Yes	Yes	Yes
Interaction of firm size with Fringe Benefits and Plan characteristics	No	No	No	No	No	No	Yes	Yes	Yes	Yes
Occupation and Interactions of Occupation with Demographics, Fringe Benefits, Plan Characteristics, and Other Employment Characteristics?	No	No	No	No	No	No	No	No	Yes	Yes

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
E. Estimated Elasticity with Respect to										
					Total					
Match Rate	-0.033	0.06	0.08		0.07		0.12		0.17	
	(0.015)	(0.03)	(0.04)		(0.04)		(0.06)		(0.09)	
Full Income	0.002	-0.25	-0.29		-0.26		-0.32		-0.43	
	(0.010)	(0.06)	(0.07)		(0.08)		(0.11)		(0.18)	
Net Wage	0.001	-0.10	-0.11		-0.10		-0.13		-0.17	
	(0.004)	(0.03)	(0.03)		(0.04)		(0.04)		(0.07)	
Marginal Tax Rate	-0.0004	0.05	0.05		0.05		0.06		0.08	
	(0.007)	(0.02)	(0.02)		(0.02)		(0.02)		(0.03)	
					Extensive					
Match Rate	-0.005	0.04	0.05		0.04		0.07		0.10	
	(0.008)	(0.02)	(0.02)		(0.03)		(0.03)		(0.06)	
Full Income	0.0003	-0.14	-0.18		-0.15		-0.18		-0.26	
	(0.002)	(0.04)	(0.04)		(0.06)		(0.07)		(0.14)	
Net Wage	0.0001	-0.06	-0.07		-0.06		-0.07		-0.10	
	(0.0006)	(0.02)	(0.02)		(0.02)		(0.03)		(0.05)	
Marginal Tax Rate	-0.00007	0.03	0.03		0.03		0.03		0.05	
	(0.001)	(0.01)	(0.01)		(0.01)		(0.01)		(0.03)	
					Intensive					
Match Rate	-0.029	0.03	0.03		0.03		0.05		0.07	
	(0.009)	(0.01)	(0.01)		(0.02)		(0.02)		(0.03)	
Full Income	0.002	-0.10	-0.11		-0.12		-0.15		-0.07	
	(0.009)	(0.03)	(0.03)		(0.03)		(0.05)		(0.03)	
Net Wage	0.0009	-0.04	-0.04		-0.05		-0.06		-0.07	
	(0.003)	(0.01)	(0.01)		(0.01)		(0.02)		(0.02)	
Marginal Tax Rate	-0.0004	0.02	0.02		0.02		0.03		0.03	
	(0.006)	(0.01)	(0.01)		(0.01)		(0.01)		(0.01)	

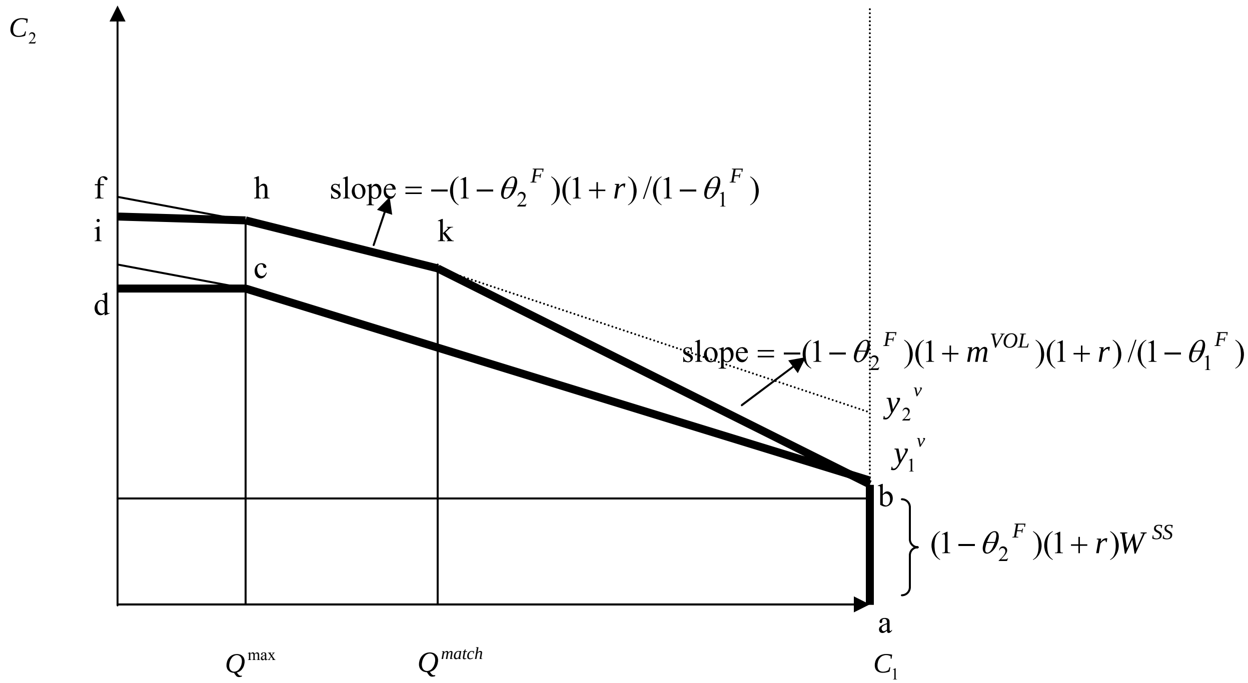
Note: Columns 1-10 of Panel A of this table present parameter estimates estimating (51) in the text. For all columns, the sample consists of 1,042 individuals, and the upper contributions limits are individual varying, as described in the text. All the columns except column (1) assume prices, net wage, and virtual income are endogenous. The IV SCLS estimator of Newey (1986) described in the text is used in Columns 4, 6, 8, and 10. The remaining columns use the IV Tobit estimator of Newey (1986). Columns 2 and 4-10 present selection-corrected estimates using the exclusion restrictions discussed in the text and the appendix. Marginal effects evaluated at the mean of the regressors are shown in panel B. The parameter estimates for the selection term in the structural contribution equation are shown in Panel C. Panel C also presents parameter estimates for the exclusion restrictions from the selection equations. The estimates in column 3 are not selection corrected. Panel D presentd the additional controls included. Panel E presents estimates of elasticities of 401(k) contributions based on the structural parameter estimates in panel A, evaluated at the sample means. The elasticities on the extensive and intensive margins were calculated using the McDonald Moffitt (1980) decomposition, respectively.

Table 7. Estimated Marginal Effects and Elasticities by Demographic Groups

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Measure	Male	Female	High School Dropout	High School Diploma	Some College	College Degree	Graduate School
Latent Marginal Effect from the Match Rate	2568	1602	1279	1905	2282	2729	2935
	(1034)	(944)	(942)	(937)	(976)	(1058)	(1107)
Total Match Elasticity	0.20	0.17	0.07	0.09	0.10	0.10	0.10
	(0.10)	(0.10)	(0.06)	(0.05)	(0.05)	(0.05)	(0.05)
Extensive Margin (Participation) Elasticity	0.11	0.12	0.03	0.02	0.02	0.01	0.01
	(0.07)	(0.08)	(0.03)	(0.02)	(0.02)	(0.01)	(0.01)
Intensive Margin Elasticity	0.08	0.05	0.05	0.06	0.08	0.09	0.09
	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.04)	(0.04)

Note: This table shows estimated marginal effects and elasticities of contributions with respect to the employer match rate by sex and education group based on the parameter estimates from the richest IV-Tobit specification, shown in column 9 of Table 6, evaluated at the sample means.

Figure 1



C_2
f
h
slope $= -(1-\theta_2{}^F)(1+r)/(1-\theta_1{}^F)$
i
k
c
d
slope $= -(1-\theta_2{}^F)(1+m^{VOL})(1+r)/(1-\theta_1{}^F)$
$y_2{}^v$
$y_1{}^v$
b
$(1-\theta_2{}^F)(1+r)W^{SS}$
a
$Q^{\max}$
Q^{match}
C_1

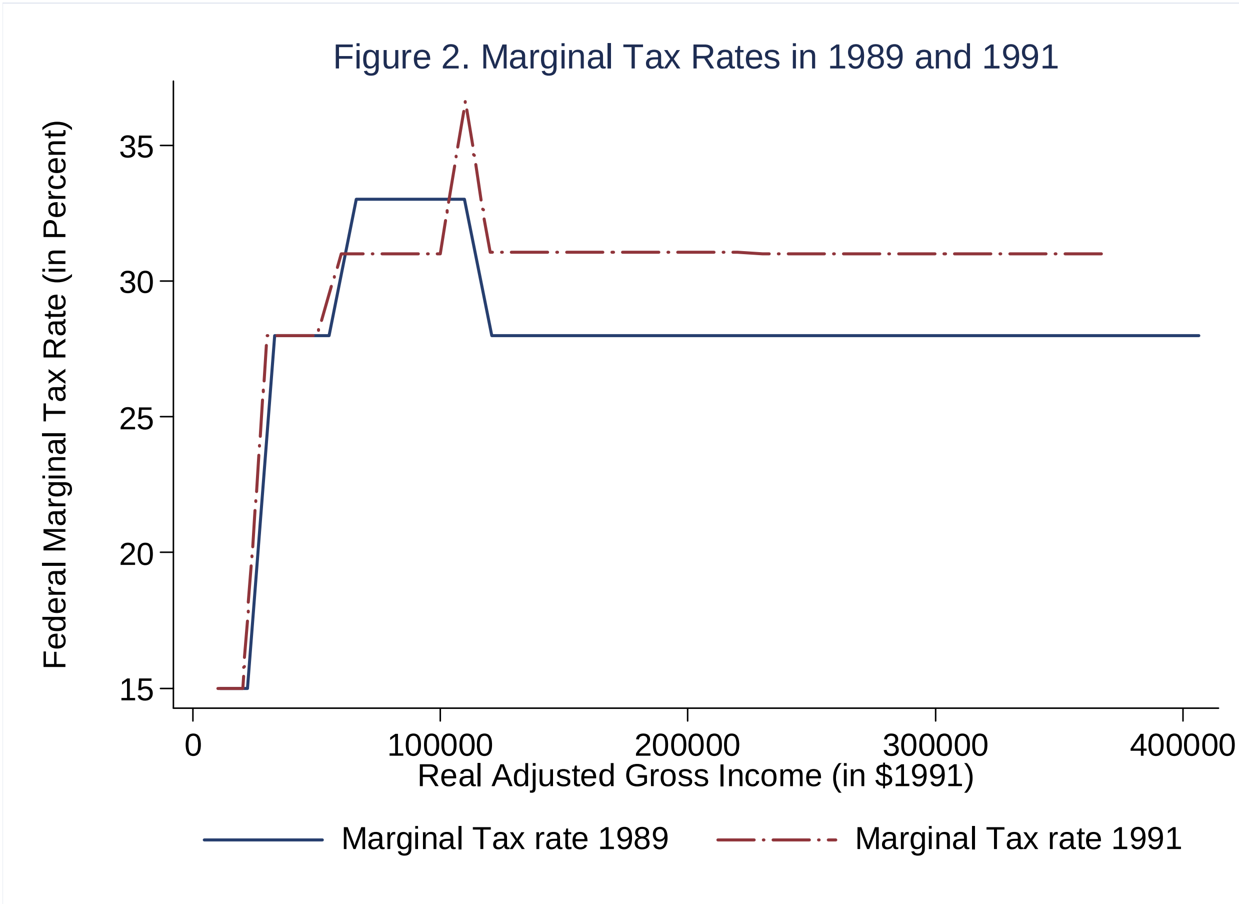


Figure 2. Marginal Tax Rates in 1989 and 1991
Federal Marginal Tax Rate (in Percent)
15
20
25
30
35
0
100000
200000
300000
400000
Real Adjusted Gross Income (in $1991)
Marginal Tax rate 1989
Marginal Tax rate 1991

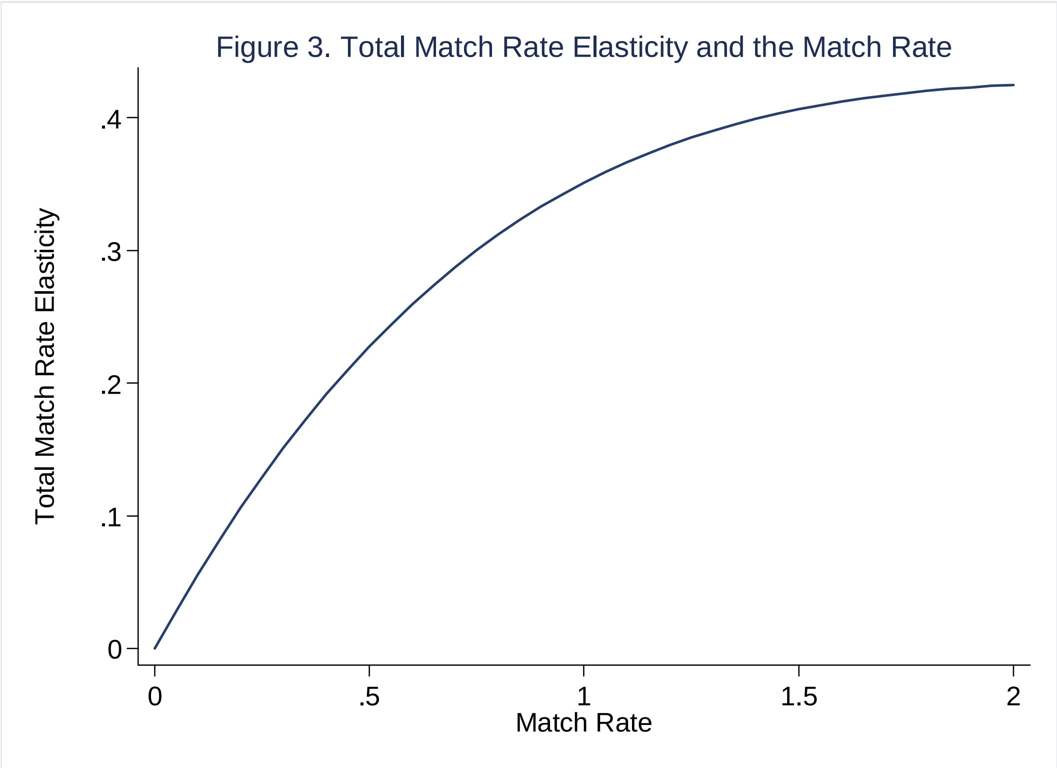


Figure 3. Total Match Rate Elasticity and the Match Rate
Total Match Rate Elasticity
Match Rate
0
.1
.2
.3
.4
0
.5
1
1.5
2