THE EFFECT OF TAXES ON PORTFOLIO CHOICE: EVIDENCE FROM PANEL DATA SPANNING THE TAX REFORM ACT OF 1986

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Abstract:

This paper estimates the effect of marginal tax rates on demands for various financial assets, using the 1983-1989 Survey of Consumer Finances (SCF) panel. In the cross-section, marginal tax rates appear to have a strong and statistically significant influence on portfolio allocations, even after controlling for income and wealth. For example, people with high tax rates tend to hold larger shares of their assets in equities, which are taxed relatively lightly due to the treatment of capital gains. Cross-sectional estimates could be biased, however, if the marginal tax rate is correlated with unobserved influences on portfolio choice, such as financial sophistication, preferences for risk or liquidity, or degree to which portfolio choices are constrained by Social Security. These unobserved variables, like the marginal tax rate, are likely to be related to income and wealth in a positive and non-linear fashion, so that the marginal tax rate may serve as a proxy for them. This paper utilizes the SCF panel to estimate fixed-effects and correlated random-effects (Chamberlain, 1984) models, which effectively control for this unobserved heterogeneity. Identification is provided by the Tax Reform Act of 1986, which changed marginal tax rates by very different amounts for different people. The empirical results confirm that, without controlling for unobserved heterogeneity, marginal tax rates have a strong influence on portfolio choices. After controlling for unobserved heterogeneity, however, the magnitude of the effect is much smaller and often statistically insignificant. These results are found to be robust to efforts to distinguish between unobserved heterogeneity and state dependence (i.e. slow adjustment of portfolios).

Introduction

The U.S. personal income tax has a major impact on the incentives facing a household when constructing a financial portfolio. Different types of assets are subject to widely varying effective tax rates. For example, corporate equities are treated relatively favorably by the individual tax code because much of the return comes in the form of capital gains. In addition to the low rates or exclusions often applied to capital gains in the U.S. tax code, taxes on capital gains are deferred until the asset is sold, thus reducing the present value of the tax burden, and are eliminated if the asset is held until death. Interest on state and local government bonds is exempt from federal income taxation. At the same time, interest from savings accounts, corporate bonds, and most federal government bonds is fully taxable. Income taxation may also influence portfolio choices by affecting the relative risk of different assets. By taxing away part of any positive return, and providing partial tax deductions for losses, it may affect the variances and covariances of returns on various assets.

The degree to which portfolio choices are affected by income taxation is an interesting and important question for a number of reasons. First, it is an interesting economic exercise to investigate whether people respond to incentives in the way we expect. Second, if people alter the risk and liquidity characteristics of their portfolios in order to avoid taxes, there is some deadweight loss. A greater responsiveness of asset demands to taxation would suggest a larger deadweight loss. Evidence on this question would be relevant to evaluating proposals to raise or lower marginal tax rates, make capital income taxation more uniform, or eliminate taxation of capital income altogether, as in the flat tax and other consumption tax proposals. Third, the government needs information on the responsiveness of portfolios to tax changes in order to estimate projected revenues. Government agencies already incorporate explicit assumptions about portfolio behavior in their estimates, and evidence on this question could help improve the estimates.[1]

This paper is the first to utilize the 1983-1989 Survey of Consumer Finances (SCF) panel to estimate the effect of marginal tax rates on the share of total financial assets invested in each type of asset. The SCF panel includes detailed information on asset holdings and personal characteristics, and importantly, spans the Tax Reform Act of 1986 (TRA86). This act changed marginal tax rates dramatically and in different ways for different people, thus providing a very useful source of identification. In cross-sectional data, the correlation between portfolio allocations and marginal tax rates appears to be strong and in the expected direction, even after controlling for income and wealth. For example, people with high tax rates tend to hold larger shares of their assets in equities, which are taxed relatively lightly. Cross-sectional estimates could be biased, however, if the marginal tax rate is correlated with unobserved influences on portfolio choice, such as the degree of financial sophistication, preferences for risk, or degree to which portfolio choices are constrained by social security. These unobserved variables, like the marginal tax rate, are likely to be related to income in a positive and non-linear fashion, so the marginal tax rate may be correlated with them even after controlling for income.

To the extent that the unobserved variables are constant across time, panel data can help solve this problem. This paper uses fixed-effects and correlated random-effects (Chamberlain, 1984) approaches to control for time-invariant unobserved heterogeneity and obtain consistent estimates. Treating the data as a pooled cross-section, without controlling for unobserved heterogeneity, yields estimates suggesting that marginal tax rates have a very strong influence on portfolio choices. After controlling for unobserved heterogeneity, however, the magnitude of the effect is much smaller and often statistically insignificant. These results are found to be robust to efforts to distinguish between unobserved heterogeneity and state dependence (i.e., slow adjustment of portfolios).

Portfolio theory

Numerous papers in the Capital Asset Pricing Model (CAPM) literature have derived optimal asset shares as a function of the expected returns, variances, and covariances of asset returns, using a variety of approaches. Examples include Merton (1969, 1971, 1973), Samuelson (1969), Friend and Blume (1975), and Friedman and Roley (1979).

To illustrate the implications of CAPM theory for empirical work on the determinants of portfolio allocation, consider the following simple model, adapted to a choice between two risky assets in the presence of differential taxation.[2] An individual chooses s_j, the share of wealth (W_t) at time t held in asset type j, in order to maximize the expected utility of future wealth. Asset returns (r_j) are assumed to be normally distributed, and following a Wiener process. Each asset's return has an expected value r_j^e and variance s_j^2, and the covariance between the returns to assets j and k is s_{jk}. There is an income tax rate τ, and the tax code's varying treatment of different types of assets is approximated by assuming that a fraction f_j of asset j's return is subject to taxation. A derivation is provided in the Appendix. The optimal share of assets to invest in asset j=1 is given by the following equation:

$$
\begin{aligned}
s_1 = {} & \frac{1}{c_t} \frac{\left((1-f_1 t)r_1^e - (1-f_2 t)r_2^e\right)}{\left((1-f_1 t)^2 s_1^2 + (1-f_2 t)^2 s_2^2 - 2(1-f_1 t)(1-f_2 t)s_{12}\right)} \\
& + \frac{(1-f_2 t)^2 s_2^2 - (1-f_1 t)(1-f_2 t)s_{12}}{\left((1-f_1 t)^2 s_1^2 + (1-f_2 t)^2 s_2^2 - 2(1-f_1 t)(1-f_2 t)s_{12}\right)},
\end{aligned}
\qquad (1)
$$

where c_t is the coefficient of relative risk aversion, $W_t[-U''(W_t)/U'(W_t)]$.

In the event that the assets do not have identical risk characteristics, individuals face a tradeoff between the advantages of investing in relatively tax-preferred assets, and the desire to maintain a portfolio with balanced risk characteristics. This tradeoff is incorporated in equation (1). In the special case where utility functions display constant relative risk aversion, everyone faces the same tax rates, and everyone has the same beliefs about expected returns and risk characteristics, equation (1) implies that everyone holds identical portfolios containing all of the available assets. Variation in tax rates (τ) across individuals would provide one reason why we observe significant variation in portfolio allocations across people.

In general equilibrium, we would expect some tax clientele effects to develop (Miller, 1997, Auerbach and King, 1983). For example, demand for tax exempt bonds would drive their rate of return down to the point where just enough people with high tax rates invest in them to exhaust the supply. Households with marginal tax rates below the implicit tax rate on tax-exempt bonds (i.e., the percentage difference between the returns to exempt and non-exempt bonds) would receive a higher return from taxable bonds and would specialize in them. Similar tax clienteles might develop for corporate equities, which are taxed relatively lightly, versus assets that pay taxable interest. However, assets with preferred tax attributes may not be perfect substitutes for less tax-preferred assets due to varying risk characteristics, so it is unlikely that strict tax clienteles would develop. Households would weight their portfolios more heavily towards the assets for which they are the clienteles, but they would still hold the full range of assets for diversification reasons. One implication of this theory is that when marginal tax rates change, the way asset allocations respond may depend more on how one's tax rate changes *relative* to other peoples' rates, rather than on the magnitude one's rate change considered in isolation.

While equation (1) is useful as a starting point, it should be noted that it leaves out several important considerations. For example, in order to reduce opportunities for sophisticated tax-avoidance schemes, the income tax does not allow full deductions for losses, so the tax rate would not enter symmetrically into the variance and covariance terms as it does above.[3] In the period covered by this study, net capital losses in excess of $3,000 could not be deducted in the year they occurred, but rather had to be carried forward to later years, reducing the present value of the tax savings. Moreover, taxes only affect the variances and covariances of the assets if the government is able to absorb some of the risk faced by the taxpayer. But low returns or losses caused by non-diversifiable economy-wide risk would reduce government revenues, and the government may transfer the risk back to taxpayers through higher tax rates or reduced services.[4] In addition, transaction costs, such as the tax on capital gains realizations, may prevent portfolios from adjusting quickly to their new optimal allocations in response to a change in tax rates or other relevant influences.[5]

Another consideration is that a large proportion of households have no holdings at all of many types of assets, which is not a prediction of a standard CAPM model like the one above. One explanation would be that individuals face actual or perceived costs of becoming informed about certain types of investments, particularly corporate equities. Only those households who clear the information cost hurdle for a given asset would invest in it. Certain assets may also involve monitoring or holding costs, which have similar implications. Haliassos and Bertaut (1995) and Bertaut (1998) explore this theory, and present some evidence in support of it based on cross-sectional relationships between equity ownership and variables likely to be correlated with information costs (e.g. education) in the SCF data. A second explanation commonly cited in the literature is that because of tax considerations, risk preferences, or beliefs, some households would optimally hold negative positions in some assets, but are unable to do so because of constraints on short sales. A third possibility is that social security and/or defined-benefit pension plans already constrain many households to do more long-term saving than they otherwise would. Any financial assets held by these households would be largely for near-term consumption purposes, and thus would only be held in fairly liquid form.

In addition to the large variation across people in which assets are owned, we also observe a great deal of variation in asset shares across the population. While this may be explained in part by tax considerations, variation in non-tax factors, such as preference for risk or liquidity, level of financial sophistication, beliefs about returns and variances, age, or degree to which decisions are constrained by social security and defined benefit plans, may also be important. Income and wealth are likely to be correlated with and serve as proxies for many of the aforementioned influences, which are often unobserved.

Wealth may also affect optimal portfolio shares directly in equation (1) if the utility function does not display constant relative risk aversion.

Some general issues in estimation

Equation (1) does not lend itself easily to structural estimation, particularly because the expected rates of return and variance-covariance structures of the assets are unknown. A more realistic version of equation (1) that incorporated all of the important additional considerations discussed above would render structural estimation intractable. For this reason, all previous empirical studies of the effect of taxes on portfolio choice have estimated simple reduced-form equations, typically regressing the portfolio share (sometimes in log form) on a tax rate measure and a set of control variables, with some adjustment for the limited nature of the dependent variable. This paper follows in the tradition of reduced-form estimation, and focuses on how unobserved heterogeneity influences the reduced form estimates.

A problem all empirical work on this subject must face is that the dependent variable is limited. When the dependent variable is defined as the share of total assets held in each particular form, a disproportionately large number of people are observed to settle at the limits, especially at zero, but also at one.[6] In general, a linear regression will produce biased estimates of the slope parameter under these circumstances. Unbiased estimates can be obtained via a tobit model if the errors are normally distributed, and the effect of each explanatory variable on the probability of being a non-limit observation is strictly proportional to its effect on asset shares. The Heckman (1979) two-step model can be used to relax this second restriction, allowing explanatory variables to have different effects on the decision whether or not to own an asset and the decision regarding how much of the asset to hold.[7] The ownership decision can be modeled as a separate probit equation. To correct for selection bias, the error term from the probit is then included in the share equation, estimated on the sample of non-limit observations.[8] The greater flexibility of this approach comes at a cost, however. In the absence of any variables that obviously affect the discrete ownership decision, but not how much is owned, the selection correction term in the level equation is just a non-linear function of the other right hand side variables, which can introduce severe multicollinearity.[9] Given that the problems with the Heckman two-step approach are compounded by my relatively small sample size, my estimation strategy will focus on the tobit framework.

It is worth noting one more potentially important implication of equation (1) for empirical work involving panel data. The coefficient on the tax rate in reduced-form estimation is likely to be a function of rates of return, variances and covariances of all assets, as well as the degree to which this asset and each of the other assets are taxable (f_j). This suggests that the tax coefficient is likely to change over time, especially if a change in tax law changes the f_j's. The responsiveness of a particular asset's demand to the marginal tax rate is not some immutable characteristic, but rather depends on the particular structure of the tax law at any given time. TRA86 changed many features of the tax code besides marginal rates; for instance, it increased the share of realized capital gains included in adjusted gross income from 40 percent to 100 percent. This point will be taken into account in the econometric work.

Empirical Literature

Several studies have analyzed the impact of taxation on household asset allocation using cross-sectional data, with mixed results. Feldstein (1976) examines the influence of taxes on portfolio choice using data on 1,799 households from the 1962 Survey of Financial Characteristics of Consumers. He runs an OLS regression of asset shares on a set of dummy variables for income ranges (roughly corresponding to tax brackets), dummies for net worth ranges, age, sex, and the ratio of estimated human capital to non-human net worth. Individuals in upper income classes are found to hold significantly larger shares of their assets in tax-preferred forms. Feldstein interprets this as evidence of "a very powerful effect" of taxation on asset demands. However, his approach cannot distinguish whether the relationship between income range and asset demand is due to the influence of marginal tax rates, the direct influence of income, or the effects of some omitted variable, such as preference for risk or financial sophistication, which may be correlated with income.

Hubbard (1985) analyzes the effects of taxes on asset demands using a cross-sectional data set collected in 1979 and 1980 by the President's Commission on Pension Policy, including 3,084 households. Using a Heckman two-step approach, he regresses the log-odds ratio of asset shares, $\ln(s_j/(1-s_j))$, on the marginal tax rate, an estimate of the log of permanent income, estimates of private wealth, social security wealth, and pension wealth all divided by permanent income, age, and other variables.[10] Higher marginal tax rates are found to have a statistically significant influence on the probability of owning several types of tax-preferred assets. The effect of the marginal tax rate on asset shares is generally inconclusive because of large standard errors, but there is a significant positive effect on equities as a share of net worth.

King and Leape (1998) examine 6,010 households from the 1978 Survey of Consumer Financial Decisions, taking an approach similar to Hubbard's. Again using a Heckman two-step approach, they regress log asset shares on marginal tax rate, age, age squared, marital status, occupation, education, employment, and response to a survey question about risk preference.[11] Additional flexibility is introduced by allowing the intercepts of each share equation to vary depending on the particular combination of assets owned.[12] The decision whether to own a particular combination of assets is endogenous, so predicted probabilities based on probit estimates are used as instruments. The same set of explanatory variables are used for the share equations, selection corrections, and asset combination equations, so multicollinearity is a serious problem. Not surprisingly, very few statistically significant results are found for the level equations, although once again the marginal tax rate is found to be a significant determinant of which assets are owned.

Poterba and Samwick (1999) investigate the impact of taxes on the allocation of financial asset portfolios using data on 15,451 households from the 1983, 1989, 1992, and 1995 cross-sections of the SCF. They estimate tobit and probit models explaining asset shares and probability of ownership, controlling for the marginal tax rate, as well as dummy variables for year, income category, net worth category, occupation, industry, education, age, sex, and marital status. They find that the marginal tax rate generally has a statistically significant impact in the expected direction on both asset demands and probability of asset ownership. Households with high tax rates are found to hold larger shares of their financial assets in relatively tax-preferred forms such as equity mutual funds, retirement accounts, and tax exempt bonds, and to hold smaller shares of their assets in relatively heavily-taxed forms such as interest bearing accounts. Results for probit equations are similar. Samwick (2000) takes a similar approach to the same data and finds broadly similar results, with a notable exception being a significant impact of the marginal tax rate on equity holdings.

Both Samwick (2000) and Scholz (1994) provide descriptive statistics based on multiple years of SCF cross-sectional data.[13] They find that the although the patterns of asset ownership in the cross section are consistent with a tax explanation, changes in asset allocations across time at different points in the income or wealth distribution do not match up very well with the way tax incentives changed for these groups.

Only two papers have examined the relationship between taxes and portfolio choice using panel data, and both were based on the 1983-86 version of the SCF panel. Ioannides (1992) runs a cross-sectional Heckman two-step using 1,622 households from the 1986 SCF,[14] with $ln[s_j/(1-s_j)]$ as the dependent variable. A broad array of explanatory variables is included.[15] The marginal tax rate is included only in the level equation. The panel aspect of the data is utilized by including the 1983 value of $ln[s_j/(1-s_j)]$ as an explanatory variable in the 1986 equation, in an effort to capture the dynamic adjustment process of the portfolio. Lagged values of the other right-hand-side variables are used as instruments for $ln[s_j/(1-s_j)]$. No allowance is made for unobserved heterogeneity. The effects of the marginal tax rate in the level equation are generally inconclusive because of large standard errors. Ioannides does find evidence that the lagged value of $ln[s_j/(1-s_j)]$ generally has a significant positive effect on the current value of $ln[s_j/(1-s_j)]$. This could be due either to slow adjustment of portfolios, or correlation of the lagged explanatory variables (which serve as instruments for the lagged dependent variable) with unobserved time-invariant influences.

Guell (1991) is the only paper to examine the impact of unobserved heterogeneity on the relationship between tax rates and portfolio choice, utilizing data on 2,376 households from the 1983-86 SCF panel. He estimates a “corelated random-effects” tobit model developed by Chamberlain (1980, 1984) and Jakubson (1983, 1988), which is described further below. Unlike the classic random-effects specification, but like fixed-effects, this approach is consistent in the presence of correlation between the regressors and time-invariant omitted variables. The dependent variables are asset shares, and the explanatory variables are the marginal tax rate, marginal tax rate interacted with age, year dummies, age, education, total assets, age of oldest child, and income. Guell finds some evidence that the marginal tax rate has a statistically significant positive effect on ownership of stocks and mutual funds, expressed as a share of total assets. Results are otherwise inconclusive due to large standard errors.[16]

The present paper is the first to estimate the effect of tax rates on portfolio choice using the 1983-89 SCF panel. The 1983-89 panel has much more detailed and reliable information on portfolios than the 1983-86 panel. Moreover, the 1983-89 period provides stronger identification of the tax effects because it spans a major change in tax law. In a fixed-effects or correlated random-effects framework which includes time dummies, identification requires that tax rates change over time in different ways for different people. The only change in tax rates occurring between the 1983 and 1986 surveys was a 10 percent (not percentage *point*) across-the-board reduction between 1983 and 1984.[17] This was the last year of a gradual three-year phase-in of rate reductions enacted in the Economic Recovery Act of 1981. Because marginal tax rates were reduced by similar amounts for everyone, the effect of this change would largely be absorbed into the time dummies.[18] The 1983-89 panel, by contrast, provides excellent identification of the tax effects because of the Tax Reform Act of 1986, which will be discussed further below. An additional advantage of the newer data is that information from at least three different time periods are necessary to distinguish between unobserved heterogeneity, as emphasized by Guell, and state dependence, as emphasized by

Ioannides. At the end of the paper, I address this issue, combining the 1983-89 panel with information from 1986.

Data

The Survey of Consumer Finances is a unique data set that provides very detailed information on respondents' asset holdings, income, other financial variables, and personal characteristics. It heavily over-samples people with high incomes, which is important for our purposes given that financial wealth is so heavily concentrated at the top of the income distribution. A SCF survey has been conducted every three years since 1983. The 1986 and 1989 surveys re-interviewed a sub-sample of households present in 1983, to construct a panel. No panel element has been included since then. The panel experienced significant attrition over time; the nature of the attrition is described in Kennickell and Woodburn (1997). In the 1983 cross-sectional survey, 4,103 observations were included. For cost reasons, only a subset of the 1983 respondents were chosen for re-interview in 1989, with the selection based mainly on geography. Of the subset that was re-contacted, the population-weighted response rate was 67 percent for members of the random component of the sample and 81 percent for members of the high-income sample. This yielded 1,479 total respondents in 1989. According to Kennickell and Woodburn, the main difference between respondents and non-respondents in the random component of the sample was a somewhat higher average 1983 income and wealth level among respondents. Income and wealth may be correlated with more stable households, and thus higher likelihood of response. In the high-income sample, respondents and non-respondents exhibited very similar characteristics. For attrition to cause bias in my estimates, there would have to be a correlation between changes in portfolio shares and changes in other characteristics that influence the likelihood of responding to the survey, are correlated with my explanatory variables, and are not controlled for in my regressions. Although there are no obvious candidates, this remains an open possibility. The 1986 survey collected only limited information on portfolio allocations and other important variables. For this reason, my empirical work mainly focuses on the data from 1983 and 1989, although I do present some estimates based on all three years of data at the end of the paper.

My 1983-89 panel sample excludes households with a head younger than 25 in 1983, as well as year 1989 observations that represent spouses or partners who split off from the main household in the intervening years. In order to reduce outlier problems, my sample also excludes 408 households who had total financial assets of less than $2,000 (in 1989 dollars) in either year; these households account for less than 3 percent of the aggregate value of financial assets in the population-weighted sample. This leaves a balanced panel of 984 households, or 1,968 observations counting both years.

To limit the scope of investigation, I concentrate the analysis on financial assets. Because of their relative liquidity, financial assets are more likely than other assets to adjust to tax changes within the span of a few years. The dependent variables are the share of total financial assets invested in each of six asset classes. Definitions of asset categories are generally consistent with those in Poterba and Samwick (1999), with some minor modifications.[19] Financial assets are listed below roughly in order from most-heavily taxed to most-lightly taxed. Total financial assets are defined as the sum of these six categories.

Dependent variables

Interest bearing accounts include checking accounts, savings accounts, money markets, and CDs. Interest on these accounts is in general taxable.

Taxable bonds include federal government bonds, corporate bonds, and foreign bonds that are held directly (not in retirement accounts or mutual funds). Interest on these bonds is generally fully taxable.[20]

Taxable mutual funds include all mutual funds except those consisting primarily of tax-exempt bonds. The SCF does not provide sufficient information to distinguish taxable bond funds from equity funds in 1983. The tax treatment of these funds is a mixture of the treatment of taxable bonds and equities, depending on what is in them. Equity mutual funds are in general taxed more heavily than directly held equities, because they generate more frequent capital gains realizations (Dickson and Shoven, 1995).

Corporate equities include stocks held directly, in investment clubs, in brokerage accounts, or publicly traded stock in one's own company. Stocks held in retirement accounts or mutual funds are counted elsewhere. Equities are relatively tax favored because much of the return comes in the form of capital gains, which benefit from deferral of tax until realization, and exclusion from tax if held until death. In addition, only 40 percent of long-term capital gains realizations were counted as adjusted income in 1983. As a result of TRA86, 100 percent of realizations were included in AGI and taxed at the ordinary income tax rate by 1989. Dividends are fully taxable.

Tax-free bonds include bonds issued by state and local governments and held directly or in mutual funds or money market accounts. Interest on these bonds is exempt from federal income tax.

Retirement accounts include Individual Retirement Accounts (IRAs) and Keogh accounts. Defined contribution pension plans such as 401(K)s, thrift plans, profit sharing plans, stock option plans, and employee stock ownership plans are also included if they are from a job where a household member is currently employed. Most contributions to these plans are deductible from income, all returns are exempt from tax, and withdrawals are taxable, which replicates consumption tax treatment. Therefore, except to the extent that tax rates differ between the time contributions are made and the time withdrawals are made, the tax rate on the returns to assets held in these accounts is zero. These plans are often subject to penalties for early withdrawal before retirement, and are subject to limitations on tax-deductible contributions. No information on the types of assets held in retirement accounts is available in the SCF for 1983, so the accounts are treated as a single unified asset class in estimation.

Explanatory variables

The *marginal tax rate* is calculated by applying the tax schedule for the appropriate year and filing status to an estimate of taxable income, which is adjusted gross income (AGI)) less deductions and personal exemptions. The SCF contains information on many important determinants of taxable income, including AGI, marital status, family size, and age.[21] It also provides sufficient information to calculate interest paid deductions, which account for about 40 percent of itemized deductions in IRS aggregate statistics during this period.[22] Other itemized deductions are imputed based on information from IRS data.[23] In an effort to protect the privacy of respondents, the SCF omits state of residence for most observations.[24] So unfortunately, state tax rates cannot be calculated.

One relevant problem at this juncture is that the marginal tax rate is likely to be endogenously related to the dependent variable. A change in the share of assets held in taxable forms will affect taxable income, which in turn can affect the marginal tax rate by pushing the taxpayer into a different tax bracket. Thus, the marginal tax rate is correlated with the error term in the share equation. To address this, an instrument for the marginal tax rate is constructed by replacing AGI in the calculation above with the following: AGI less taxable capital gains, dividends, and taxable interest, plus 5 percent of the value to total financial assets. This is intended to purge the marginal tax rate of the effects of endogenous financial asset allocation decisions. The 5% figure roughly approximates the average taxable return to financial assets over this period. Poterba and Samwick (1999) use a roughly similar measure of the marginal tax rate as a proxy for the actual marginal tax rate; I will use it as an instrument.[25]

Financial explanatory variables include *labor income*, *net worth*, *financial assets as a share of total assets*, and a *pension benefit* variable. All are measured in constant 1989 dollars. Net worth is a very comprehensive measure of total assets minus debts.[26] The "pension benefit" variable includes the value of any private pension benefits currently received by retirees, plus the expected value of pension benefits in the first year of retirement from defined benefit plans from a current job or any kind of pension from a past job. These kinds of pensions are the only major source of private wealth not included in the measures of net worth or retirement accounts.[27]

Education, *risk preference,* and *liquidity preference* variables include the following. Dummies for “some college,” “college graduate,” and “graduate degree” are included to control for education. Three dummy variables representing different responses to a 1983 question about “the amount of financial risk you are willing to take when you save or make investments” are included as proxies for risk preference. “Very high” risk preference, for example, indicates a willingness to “take substantial financial risks expecting to earn substantial returns.”[28] Liquidity preference is measured by three dummy variables based on responses to a 1983 survey question asking “how you feel about tying up your money in investments for a long period of time.” “Very high” liquidity preference, for example, corresponds to ”not willing to tie up money at all.”[29] All of these variables are excluded from the fixed-effects and correlated random-effects specifications, because they do not vary over time, so their influences are absorbed into the fixed-effect.[30]

Finally, demographic variables on *marital status*, *age*, and *age squared* are included. Age is omitted from the fixed-effects and correlated random-effects specifications because it becomes perfectly collinear with the time dummies when its cross-sectional variation is removed – it changes by the same amount for everyone.[31]

Descriptive statistics

Table 1 illustrates the unweighted means of all variables used in the econometric analysis. Detailed nationally representative descriptive statistics based on the SCF cross-sections have already been provided elsewhere (e.g., Poterba and Samwick 1999 and Samwick, 2000), so I focus here on the unweighted characteristics of the panel sub-sample used for my econometric estimation. The average household in the sample holds most financial assets in interest bearing accounts (43.5 percent) and retirement accounts (26.8 percent). Directly held equities comprise 16.6 percent of the average portfolio, followed by tax-free bonds (5.9 percent), taxable bonds (4.9 percent) and taxable mutual funds (2.4 percent). Total financial assets represent about a third of gross assets.

Given that the average marginal tax rate dropped sharply from 33 percent to 23 percent, one might expect to see increases in ownership of assets that are taxed relatively heavily, and declines for the more lightly taxed assets. Changes in asset allocations over time do not appear particularly consistent with this pattern; there is a decline in one relatively tax-favored asset (equities), but also a decline in a heavily taxed asset (interest-bearing assets) and an increase in a lightly taxed asset (retirement accounts). This does not necessarily tell us much, though, because marginal tax rates changed by very different amounts for different people, and supply of assets is important as well as demand. According to the general equilibrium model worked out by Auerbach and King (1983), we would expect that changes in asset allocations should depend largely on how one's marginal tax rate changed *relative* to other people's rates. The econometric analysis below focuses on responses to relative variation in tax rates by controlling for time dummies.

Average labor income in the sample is $108,000, and average net worth is $2.44 million, both in 1989 dollars. It is also a relatively old sample, averaging 53 years of age in 1983. While this sample is not very representative of the average person in the U.S., it is more representative of the average *dollar* of wealth in the U.S. For example, separate calculations indicate that the mean household labor income, weighted to be representative of the U.S. population, is only $38,000 in 1989. This is considerably lower than the unweighted sample mean of $121,000 shown in Table 1. The mean labor income weighted by *the value of financial assets* times the population weights, however, was pretty close to the unweighted sample mean, at $126,000. This suggests that the sample is particularly useful for estimating how changes in tax rates might affect allocations of aggregate wealth, because the typical person in the sample is fairly representative of the kinds of people who control most of that wealth.

Changes in tax rates

Figure 1 illustrates variation in the size of marginal tax rate cuts between 1983 and 1989, caused primarily by the Tax Reform Act of 1986. To focus on the exogenous variation arising from the reform, it depicts the average change in marginal rates that would occur for observations in each 1983 tax bracket, if the 1989 tax law were applied to their (inflation-adjusted) 1983 income. TRA86 collapsed 20 different tax brackets, ranging from zero to 50 percent, into just five brackets: 0 percent, 15 percent, 28 percent, a 33 percent “bubble” caused by the phase-out of personal exemptions, and then another 28 percent marginal rate bracket for those with the highest incomes. People at the very top of the income distribution, particularly those in the 48 and 50 percent brackets, received substantially larger reductions in marginal rates than others did. Qualifying for the 48 percent bracket required a taxable income of $85,600 for a joint filer in 1983 (about $143,200 in 1999 dollars). The size of marginal rate cuts across the rest of the population depended largely on how close one was to either the 28 percent or 15 percent threshold. Substantial variation in the magnitude of tax rate changes across the population provides crucial identification for my empirical estimation. Asset allocations across tax brackets and time

Table 2 and Figure 2 divide the sample into two groups: those facing marginal tax rates of 48 percent or above in 1983, and everyone else. The former group represents a large portion of the sample, with 313 observations. In the cross-section, there appears to be a strong relationship between high marginal tax rates and ownership of tax-preferred assets. Those in the highest tax brackets put a far smaller share of their portfolios in interest-bearing accounts and a much larger share in tax-preferred instruments such as corporate equities and tax-free bonds. The changes over time between the two groups again display no consistent pattern, although the substantial drop in equity shares among those in the highest tax brackets would be consistent with a tax explanation.

Probabilities of asset ownership

Table 3 illustrates the proportion of households in the sample who own each type of asset. Asset shares for all categories except for interest-bearing accounts exhibit a high degree of censoring at zero; none of the asset shares except interest-bearing accounts is censored at one (not shown). Interest-bearing accounts, on the other hand, are owned by all but 2 observations in the sample, and a substantial fraction of the observations hold 100 percent of their financial assets in interest-bearing accounts. For these reasons, I estimate single-limit tobits; the interest-bearing account equations correct for censoring from above at 1, equations for all other assets correct for censoring from below at zero.[32]

Changes in asset ownership over time follow some interesting patterns. The share of households owning tax-free bonds increases significantly, which is consistent with what we would expect in response to a compression of the marginal rate structure. Equation (1) above suggests that people will hold tax-exempt bonds if they offer a higher after-tax rate of return than do taxable bonds. Moreover, if taxable and tax-free bonds are not perfect substitutes (for risk reasons), some people will hold tax-free bonds even if they offer a lower after-tax return than taxable bonds. An individual may hold some of each, putting a larger share in tax-free bonds as one's tax rate rises relative to the implicit tax rate in tax-free bond interest. A compression of the marginal rate structure pushes people on both sides of the implicit tax rate closer to it, which we would expect to cause more people to hold the tax-free bonds, with a smaller share held by each. A large increase in ownership of taxable mutual funds seems consistent with other evidence on their increasing popularity.

Econometric methods

This paper's empirical strategy is to focus on approaches that have generally found a significant relationship between taxation and portfolio choice in the previous literature. These include linear and tobit models of the effect of the tax rate on asset shares, and probit models of the effect of the tax rate on the decision whether or not to own particular assets. My main contribution is to examine whether controlling for unobserved heterogeneity makes a difference in these frameworks, utilizing a new data set and a good natural experiment.

Unobserved heterogeneity is likely to be an important problem in the study of taxation and portfolio choice. In a cross-section, the marginal tax rate is just a nonlinear function of income and a few other variables, such as marital status, family size, and mortgage interest payments. Therefore, the tax rate may act as a proxy for omitted variables that influence portfolio choice and are related to income in a nonlinear way. A list of such omitted variables might include: degree of financial sophistication, preference for risk, preference for liquidity, and degree to which one's portfolio choices are constrained or influenced by social security or defined-benefit pensions. While it is true that self-reported proxies for some of these influences, such as risk and liquidity preferences, are included in my cross-sectional analysis, it is unlikely they would fully control for these factors. To the extent that the marginal tax rate is correlated with omitted variables like these, conditional on the other explanatory variables, estimates of the tax coefficient will be inconsistent.[33]

Consider the following simple reduced-form share equation (setting aside for the moment the limited dependent variable problem):

$$s_{jit} = \alpha_i + \beta_{0t} + \beta_\tau \tau_{it} + \beta_X' X_{it} + \varepsilon_{it} \qquad (2)$$

In this equation, s_{jit} is the share of financial assets held in asset *j* by individual *i* at time *t*, α_i is an individual-specific fixed effect, β_{0t} is a time-specific intercept, τ_{it} is an individual's marginal tax rate in year *t*. Other explanatory variables are denoted X_{it}, and ε_{it} is a random error term. The time-specific intercepts are included to capture, to the extent possible, the influence of any unobserved influences that are constant across people but vary over time, such as macroeconomic conditions, pre-tax rates of return on various assets, and average changes in tax rates across time. In the presence of time dummies and fixed effects, identification arises solely from changes in an explanatory variable *relative* to the average change. Any of the explanatory variables may be correlated with the fixed effect, α_i. Note that in equations (2) through (6), everything on the right hand side should also be indexed with a *j* subscript, to reflect the fact that the parameters vary across each asset equation. The *j* subscript is omitted for readability.

The standard method for estimating an equation like (2) in a linear framework is to remove the fixed effect by differencing the data from its individual means over time, or by first-differencing the data (the two approaches are identical when there are only two time periods). The first-differenced fixed-effects estimator for two periods (t=1 and t=2), would be:

$$s_{ji2} - s_{ji1} = (b_{02} - b_{01}) + b_t(t_2 - t_1) + b_X{}'(X_{i2} - X_{i1}) + (e_{i2} - e_{i1}) \qquad (3)$$

Chamberlain (1980, 1984) proposes an alternative approach. Allow the unobserved time-invariant effect to be freely correlated with the observed explanatory variables from both periods according to the following equation:

$$a_i = l_{t1}t_{i1} + l_{t2}t_{i2} + l_{X1}{}'X_{i1} + l_{X2}{}'X_{i2} + u_i\,, \qquad (4)$$

where u_i is a mean zero individual-specific random-effect. Substituting (4) into (2) yields:

$$s_{jit} = b_{0_t} + b_t t_{it} + b_X{}'X_{it} + l_{t1}t_{i1} + l_{t2}t_{i2} + l_{X1}{}'X_{i1} + l_{X2}{}'X_{i2} + (e_{it} + u_i)\,. \qquad (5)$$

When estimated on the pooled data from both periods, equation (5) is fully identified. Note that perfect collinearity is not a problem because τ_{it} is identical to τ_{i1} for only half of the sample, and is similarly identical to τ_{i2} for only the other half the sample (the same goes for each of the X variables). Efficiency can be enhanced by estimating (5) as a random-effects model, which takes into account the correlation of the error term across time for an individual, which is caused by u_i. A conventional random-effects model accounts for the correlation across time in an individual's error terms, but does not include the rest of equation (4) in the specification, so it is inconsistent if the explanatory variables are correlated with omitted time-invariant variables. Chamberlain's approach, as outlined in (5), is still consistent under these conditions, and is thus sometimes called *correlated* random-effects estimation.

In a linear model, fixed-effects and correlated random-effects estimation are not mathematically identical, but are extremely close. However, the correlated random-effects approach has two key advantages for the application in this paper. First, it allows the β coefficients to vary over time, while still controlling for unobserved heterogeneity. This is not possible in the fixed-effects estimator. As discussed above, the coefficient on the tax rate would be expected to vary over time because of statutory changes in *f*, the share of each asset that is subject to tax. In other words, β_τ is a function of *f* , which can change over time due to tax reforms (as it did for equities as a result of TRA86). We can let the coefficient on the tax rate change across years by estimating the following equation:

$$\begin{aligned} s_{jit} &= b_{0t} + b_{t1} d_1 t_{it} + b_{t2} d_2 t_{it} + b_X{}' X_{it} \\ &+ l_{t1} t_{i1} + l_{t2} t_{i2} + l_{X1}{}' X_{i1} + l_{X2}{}' X_{i2} + (e_{it} + u_i) \end{aligned} \quad , \qquad (6)$$

where d_1 and d_2 are dummy variables for the first and second periods, respectively. As with the other fixed-effects and correlated random-effects approaches, identification of the tax coefficients requires changes in tax rates over time that differ across the population. Intuitively, the identification for this model comes from the cross-sectional relationship between tax rates and the asset share in each period, partialing out the effect of this period's tax rate on the other period's share. To the extent that one period's tax rate has an effect on the other period's asset share, conditional on that period's tax rate, it is assumed to be due to the correlation with omitted time-invariant variables.[34]

A second important advantage of the correlated random-effects approach is that it enables us to control for unobserved heterogeneity in a non-linear estimation equation such as a tobit or probit. Because these functions are non-linear, fixed effects cannot be removed by differencing. In principle, one could estimate these models by including dummy variables for each individual, estimating α_i directly. In addition to being computationally intractable for a large number of observations, this approach will produce inconsistent parameter estimates when the number of time periods is small.[35] Estimation of a correlated random-effects model, by contrast, is both straightforward and consistent in a tobit or probit framework (Chamberlain, 1980 and 1984). In this paper, I estimate equation (6) by random-effects tobit and probit maximum-likelihood, which is computationally straightforward for two periods.[36] This approach can also be adapted to a Heckman two-step procedure (Woolridge, 1995). Unfortunately, I found multicollinearity to be a severe problem when the Heckman two-step approach was applied to my data, so I focus mainly on tobit and probit results in this paper.

Another econometric issue is that, as discussed above, the marginal tax rate is endogenous because it depends in part on the allocation of financial assets between taxable and tax-preferred categories. I address this problem using an instrumental variable approach; the instrument is described above in the data section.[37] The tobit and probit models are adapted to an instrumental variables framework using a procedure outlined in Newey (1987).[38]

Before presenting results for the tobits and probits, I will show the results of some linear two-stage-least-squares specifications, including pooled cross-section, fixed-effects, and correlated random-effects approaches. Linear models have one advantage over tobits in this application, which is that the linear models automatically satisfy the adding-up constraint that the marginal effects of each variable should sum to zero across all six equations, because the net change in asset shares is zero. A marginal effect is the effect of a marginal increase in a given explanatory variable on the expected value of the dependent variable, conditional on the other variables, which is generally what is of economic interest. In the linear model, the marginal effect of a variable is simply its coefficient β, and there is no need to impose restrictions across the equations, as the data imposes them automatically in this case.[39] In the tobit model, on the other hand, the marginal effect is the estimated coefficient times the predicted probability of being a non-limit observation, which is a nonlinear function of all the variables and parameters.[40] When there is censoring at zero (grouping all explanatory variables into the vector **X** for notational simplicity), the marginal effect of a variable x_k on the expected value of the asset share is:

$$\frac{\partial E[s_i \mid X_i]}{\partial x_k} = b_k \Phi\left(\frac{b' X_i}{s}\right) \quad , \tag{7}$$

where σ is the standard deviation of the tobit regression, and Φ(•) is the cumulative density function of the standard normal distribution. For censoring at 1, the marginal effect is:

$$\frac{\partial E[s_i \mid X_i]}{\partial x_k} = b_k \Phi\left(\frac{1 - b' X_i}{s}\right). \tag{8}$$

Imposing a set of nonlinear constraints across all six tobit equations for each explanatory variable poses serious computational challenges that are not addressed in this paper.[41] However, there is reason to believe that the marginal effects from the linear models are likely to be very close to the marginal effects that would be obtained by imposing the adding-up constraints on the tobit model. It is a well established empirical regularity that, although they are not in general consistent estimators of the marginal effects, linear least squares coefficients very closely approximate tobit marginal effects evaluated at the means of the data.[42]

Econometric results

Results of linear models

Table 4 presents the tax coefficients from linear 2SLS estimates. Roughly speaking, assets are laid out with more heavily taxed assets towards the left side of the table, and more tax-preferred assets towards the right. If high marginal tax rates influence people to invest in tax preferred assets, we should expect to see negative coefficients on the left side of the table, as people respond to the higher tax rates by reducing their allocations to heavily taxed assets. As we move to the right, we should see positive coefficients, as the higher tax rate induces more investment in tax preferred assets.

The first row of Table 4 displays results of treating the panel as a pooled cross-section, with no allowance for unobserved heterogeneity. These cross-sectional coefficients generally suggest a strong and statistically significant impact of the marginal tax rate on portfolio allocations. Both the marginal tax rate and the portfolio shares are measured on a 0-1 scale. So the interpretation of the first coefficient is that a one percentage point increase in the tax rate would decrease the percentage of assets held in interest-bearing accounts by 0.906 percentage points. Taxes are found to have a significant positive impact on tax preferred assets including mutual funds, equities, tax-free bonds, and retirement accounts. Taxable bond holdings appear to have little correlation with taxes. Overall, these results appear consistent with the cross-sectional estimates of Poterba and Samwick (1999), although the effects are stronger in my estimates; the source of differences from Poterba and Samwick's results will be discussed further below.

Table 4's second row shows that fixed-effects estimation wipes out much of the impact of taxation on portfolio choices. Although tax effects remain statistically significant and in the expected direction for interest-bearing accounts, mutual funds, and equities, the magnitude of the effect is much smaller in almost all cases. The tax effect on interest-bearing accounts, for example, is reduced from –0.906 to –0.186. Notable exceptions are mutual funds, which appear largely unaffected by the fixed-effects innovation, and taxable bonds, for which the tax effect changes signs to the expected negative direction, but remains statistically insignificant. The third row of Table 4 illustrates that, at least in the linear model, the correlated random-effects approach produces almost exactly identical results to a fixed-effects model.

The bottom half of Table 4 shows the effects of allowing the tax coefficients to vary over time in the linear framework. Rows 4 and 5 display the tax coefficients for 1983 and 1989 if we treat the panel as a pooled cross-section, with no adjustment for unobserved heterogeneity. In general, the results are consistent with more restrictive estimates in row 1. The most notable differences are that the coefficient on equities declines sharply in 1989 relative to 1983, while the coefficients on tax-free bonds and retirement accounts increase. This may be related to the fact that equities became relatively less tax favored after TRA86, when the 60 percent exclusion for capital gains realizations was eliminated. As indicated in equation (1) above, a large change in *f* could cause a change in the coefficient on τ.

Rows 6 and 7 present linear correlated random effects estimates allowing tax coefficients to change over time. Again, controlling for unobserved heterogeneity greatly reduces the magnitude of most tax effects overall, although in some cases the tax effect increases. The effect of taxes on interest-bearing accounts falls sharply but remains statistically significant in 1983, and disappears in 1989. Exceptions to the general pattern of diminished tax effects include a larger negative effect of taxes on taxable bonds which emerges in 1989, and effects on mutual funds remain statistically significant and grow a bit. Overall these results suggest that, after controlling for unobserved heterogeneity, some influence of taxes on portfolio choice remains, but in general the size of the effect is greatly reduced.

To aid in the interpretation of these results, consider the following exercise. If we make some assumptions about the degree to which various assets are subject to tax, we can construct an estimate of the responsiveness of the overall taxable share of one's portfolio to the marginal tax rate:

$$\frac{\partial q_t}{\partial t_t} = \sum_{j=1}^{J} f_{jt} b_{tjt} \ , \qquad (9)$$

where θ_t is the share of one's portfolio subject to tax, f_{jt} is the fraction of the return to asset *j* subject to tax under year *t* tax law, b_{tjt} is the tax coefficient for asset *j* in year *t*, and J is the number of assets (six). For illustrative purposes, I make the following assumptions about *f* for each asset, which are based on commonly used assumptions in the literature and combined with information from the SCF. Returns to interest-bearing accounts and taxable bonds are assumed to be 100 percent taxable, while returns to tax-free bonds and retirement accounts are assumed to face no tax on their returns. Equities are assumed to be 36.1 percent taxable in 1983 and 46.8 percent taxable in 1989. Mutual funds are assumed to be 65.9 percent taxable in 1983 and 78.7 percent taxable in 1989 (the sources of all these assumptions are provided in the endnotes).[43]

The implied impact of the marginal tax rate on the overall taxable share of assets is shown in Table 5.[44] In 1983, a pooled cross-section approach suggests that a one percentage point increase in the MTR would reduce the share of one's portfolio that is subject to tax by 0.701 percentage points. The correlated random-effects approach reduces the estimate of this impact to just 0.152. In 1989, a one percentage point increase in the marginal tax rate would reduce the taxable portfolio share by 0.910 percentage points under the pooled cross-section approach, and only 0.117 percentage points under correlated random effects. Controlling for unobserved heterogeneity thus appears to cause a major reduction in the estimated overall responsiveness of the taxable share of the portfolio to taxation.

Consider the implications of the results in Table 5 for government revenue. Assume that revenue from taxing capital income is given by: $R = tq(t)rA$

where R is revenue, $q(t)$ is the taxable share expressed as a function of t, and rA is a fixed level of pre-tax capital income. Further assume that when people respond to the tax increase by substituting away from taxable assets and towards nontaxable ones, the taxable assets they no longer hold do not end up in the hands of other taxable entities. This could be the case, for example, if the supply of each asset is perfectly elastic, or if the taxable assets are transferred to tax-free accounts or bought up by tax-exempt entities, such as non-profit organizations or pension funds. Under these assumptions, we can construct an estimate of the maximum degree to which the responsiveness of θ to τ reduces revenue gains from a tax increase:

$$\frac{\partial R / \partial t}{\partial R / \partial t_{|behavior}} = \frac{t \dfrac{\partial q}{\partial t} rA + qrA}{qrA} = \frac{t \dfrac{\partial q}{\partial t} + q}{q} \tag{10}$$

The numerator represents the derivative of revenue with respect to the tax rate, incorporating the behavioral response of θ, and the denominator represents the derivative of revenue with respect to the tax rate assuming no behavioral response (i.e. $\partial\theta/\partial\tau = 0$). Thus (10) is an estimate of the increase in revenues resulting from a tax increase, as a fraction of what it would be in the absence of any behavioral response.

Based on the unweighted means from Table 1, combined with the estimates of f given above, in 1983, τ=0.33, and θ=.578.[45] Evaluating equation (10) at these values using the pooled cross-section estimate of $\partial\theta/\partial\tau$ for 1983 (-0.701) implies that an increase in capital income taxation would only bring in 60 percent as much revenue as it would if there were no behavioral response. Evaluating the same equation using the correlated random effects estimate of $\partial\theta/\partial\tau$ for 1983 (-0.152) implies that the tax increase would bring in 91 percent as much revenue as it would in the absence of behavioral response. Of course, this should be viewed as an upper bound on the degree to which behavioral responses reduce revenue growth from a tax increase. Contrary to the assumption made at the outset of this exercise, part of the response to changes in relative tax rates among people would be a reallocation of taxable assets away from people with relative increases in their tax rates, towards people with relative decreases in their tax rates. The overall effect on revenue would depend on how much higher the tax rates of the people divesting are than the tax rates of the people investing. But in any event, this gives some sense of the economic importance of the difference between the cross-sectional and panel estimates.

One final point worth noting is that controlling for unobserved heterogeneity greatly reduces the tax effects even though the cross-sectional estimates control for the answers to risk- and liquidity-preference questions in the 1983 SCF. Although these survey questions probably do not fully capture people's risk and liquidity preferences, this does suggest that other unobserved factors, such as financial sophistication, may be responsible for the apparent bias in the cross-sectional marginal tax rate estimates.

Tobit results

Table 6 displays the marginal effects of the marginal tax rate on asset shares based on tobit estimation. A conventional random-effects (RE) tobit is compared with a correlated random-effects tobit; only the latter controls for omitted variable bias. The conventional RE model requires the same restrictive assumptions about orthogonality between unobservables and explanatory variables as does the pooled cross-section estimator, but it is more efficient. Overall, the most striking result from this table is its similarity to the results in Table 4. Marginal effects from the conventional random-effects tobit are very close to the pooled cross-section estimates in the bottom half of Table 4, and marginal effects from the correlated random-effects tobits are quite close to the linear correlated random-effects estimates in the last two rows of Table 4. In both the linear and tobit approaches, controlling for unobserved heterogeneity reduces the magnitude of the estimated tax effects to a similar degree. Most of the other qualitative conclusions from linear analysis in the last section carry over to the tobit case as well. If anything, the evidence for tax effects is weaker in the correlated random effects tobit framework; only the marginal effects for mutual funds are statistically significant.

Marginal effects in the top half of the table are calculated based on the unweighted means of the data. The bottom half of the table shows a weighted-sum of individual marginal effects, where the weights are the population weights multiplied by financial assets. These latter marginal effects are meant to give a better picture of how tax rate changes would affect aggregate asset shares. It turns out that marginal effects are largely similar under either method of calculating them, although generally a bit larger under the second approach. The tax effect on tax-free bonds in particular is much larger under the weighted-sum approach.

A few additional points are worth noting about Table 6. When evaluated at the unweighted means of the data, the tobit marginal effects come reasonably close to satisfying the constraint that they sum to zero across the equations. In the case of the correlated random-effects model, they are only off by 0.013 in 1983 and 0.010 in 1989. The weighted sum of the individual marginal effects does not do as well in that regard. One other notable result is that in the conventional random-effects tobit estimates, the tax effect on equities remains strong in both years, unlike in the linear model, where they were estimated to weaken significantly in 1989.

Tables 7 provides full results for the conventional random-effects Tobit model, presented as marginal effects evaluated at the unweighted sample means. The results in Table 7 generally conform to expectations. For example, a greater tolerance for risk is found to have a strong impact on demand for equities and to reduce demand for interest-bearing accounts and bonds. Education has a significant positive impact on all assets aside from interest-bearing accounts, and the effect is particularly pronounced for equities. Net worth has a strong positive effect on equity demand and a strong negative impact on demand for interest-bearing accounts. Conditional on all the other variables including the tax rate, labor income does not appear to have a strong effect on any asset shares except taxable bonds. The pension benefit appears to have little impact on any equation.

Table 8 shows full results for the correlated random-effects tobit model. The top half of the table presents the marginal effects of the explanatory variables on the asset shares. Controlling for unobserved heterogeneity wipes out most of these marginal effects. Mutual funds are once again an exception, with small but significant effects for a number of variables. The bottom half of the table shows the estimated correlation between the explanatory variables and the unobserved fixed-effects, multiplied by the probability of holding a positive share. Most of the tax effects apparent in the cross-section are absorbed into these lambda coefficients.

Comparison to Poterba and Samwick

Aside from the use of panel data and controls for unobserved heterogeneity, the basic econometric model used in this paper is broadly similar to that used in Poterba and Samwick (1999). In both cases, the shares of financial assets held in various forms are regressed on the marginal tax rate and a variety of control variables using a tobit framework. There are differences in the details, though. Among the major differences in the specification used by Poterba and Samwick, compared to mine, are as follows. Their sample includes all households in the SCF cross-sections; households with low levels of financial assets or with very young heads are not excluded. Their list of dependent variables includes an “other financial assets” category, which largely consists of trust accounts and the cash value of certain life insurance policies (both of which are relatively tax-favored, as they can accumulate interest tax-free). Dummy variables for ranges of income, net worth, and age are used in the Poterba and Samwick specification, as opposed to continuous variables (and a quadratic in age) in my specification.[46] Their list of explanatory variables also includes dummies for various occupations and industries, and excludes the SCF variables on risk and liquidity preference. Their algorithm for calculating the marginal tax rate also differs from mine in a few details. For example, instead of using the AGI variable reported in the SCF, they construct their own estimate of AGI based on various components of income reported in the SCF, and the procedure for imputing itemized deductions is somewhat different.[47]

Table 9 presents an approximate replication of the results from Poterba and Samwick (1999) based on the 1983 cross-section of the SCF. The first row presents marginal effects reported by Poterba and Samwick for a tobit model estimated on the full 1983 SCF cross section.[48] The second row represents the results of my attempt to replicate their study using a specification as close to theirs as I could construct, using the same data. I was unable to replicate their results exactly because they do not provide sufficient information in their paper to exactly replicate their tax calculation algorithm.[49] Using a somewhat different estimate of the marginal tax rate than they use, but an otherwise identical specification, the basic pattern of results in very similar but not quite the same. The main difference is that the magnitudes of the effects of interest-bearing accounts, retirement accounts, and other financial assets are moderately larger in my replication.

Both the Poterba and Samwick results and my replication suggest that marginal tax rates have a statistically significant impact, in the expected direction, on interest-bearing accounts and retirement accounts. Poterba and Samwick's estimates for other years of the SCF were generally similar, although other years also produced statistically significant estimates for mutual funds and tax-exempt bonds. The magnitudes of these effects, however, are strikingly smaller than those found in my "pooled cross-section" specifications in Tables 4 and 6. This difference is driven almost entirely by the use of dummy variables for income and net worth ranges in the Poterba and Samwick specification. A dummy variable approach allows for a more flexible non-linear relationship between income (or net worth) and asset shares, and this greatly reduces the estimated independent effect of the marginal tax rate. The third row of results in Table 9 shows the results of a specification identical to my replication of Poterba and Samwick, except that income and net worth are included as continuous variables rather than in dummy-variable form. These results are very similar to those reported for my conventional random-effects tobit model (not controlling for unobserved heterogeneity) in the first row of Table 6. Taken together, all of the other differences in specification and sample selection between Poterba and Samwick and my paper, aside from unobserved heterogeneity and the functional form of income and net worth, do not appear to significantly affect the marginal tax rate coefficients. This is reassuring for a number of reasons. For example, it suggests that the large attrition in the SCF panel does not substantially bias the tax-parameter estimates; they are very similar whether estimated on the full 1983 cross-section, or on the sub-sample of observations that chose to respond again in 1989.

The main message of Table 9, however, is that allowing income and net worth to have a more flexible, non-linear relationship with asset shares in cross-sectional data reduces the magnitude of tax effects almost as much as controlling for unobserved heterogeneity in panel data. One drawback of the dummy variable approach is that we do not know the true functional form of the relationship of asset shares with income or net worth (or their possibly time-invariant unobserved correlates). So any attempt to capture this through a non-linear function of income or net worth, such as dummy variables for income ranges, must be somewhat arbitrary. As an extreme example, if we broke income down into enough categories and interacted it with a few other characteristics such as marital status, there would be almost perfect collinearity with the marginal tax rate. We would find no statistically significant effect of taxes on portfolio choice, but this would not be particularly informative.[50] In addition, when dummy variables for broad ranges of income and net worth are used in place of continuous measures, the marginal tax rate coefficient may be biased because of its correlation with the now omitted variation in income or net worth within these ranges. Focusing on the effect of exogenous *changes* in the tax schedule caused by tax reforms, as I do in this study, should be more robust to mis-specification of the income and net worth terms (or exclusion of variables that are related to income or net worth in a non-linear fashion). As it turns out, it appears that the specific flexible functional form for income and net worth used in the Poterba and Samwick study leads to estimates of the impact of marginal tax rates that are coincidentally not far from those I find when controlling for unobserved heterogeneity. For example, compare the first two rows of Table 9 with the third row of results in Table 6.

Probit results

Table 10 shows marginal effects estimates for probit models of the decision whether or not to own each particular asset. Interest-bearing accounts are omitted because they are almost universally owned. The estimation approaches and layout of the table are identical to those for the tobit tax effects. Note that there are no adding-up constraints for the probit models, as the decision to purchase one type of asset does not require someone to stop owning another type altogether. Conventional random-effects estimates find a very strong positive relationship between the marginal tax rate and the ownership probability for each type of asset. The interpretation of the upper left hand entry, for example, is that a one percentage point increase in the marginal tax rate would increase the probability of owning taxable bonds by 0.499 percent. Tax rates are found to have a particularly strong influence on ownership probabilities for equities and retirement accounts. Once again, controlling for unobserved heterogeneity through a correlated random-effects approach sharply reduces the estimated marginal effects and renders most of them statistically insignificant. Taxes have a marginally significant positive impact on the probability of owning tax-free bonds in 1983.

Table 11 provides complete marginal effects estimates for the conventional random-effects probit. Education, self-reported risk tolerance, and net worth are found to have a strong positive influence on the probability of owning equities. Education and risk tolerance also have a significant positive impact on mutual fund ownership probabilities. Table 11 shows the full results for the correlated random effects probit, which again reduces most of the marginal effects.[51]

Significance tests

Table 12 presents significance levels for a variety of hypothesis tests on the linear, tobit, and probit correlated random effects models. First, the correlated random-effects approach is tested via a test of the hypothesis that the right-hand-side variables are jointly uncorrelated with the fixed-effect. In other words, it tests whether the λ's are jointly zero. The λ's are found to be jointly statistically significant in all cases except mutual funds. This result suggests that controlling for unobserved heterogeneity is indeed important, and effectively rejects the hypothesis that the pooled cross-section or conventional random-effects estimates are consistent.

The next test regards the joint significance of the tax rate λ's, to see whether the marginal tax rate suffers from omitted variable bias in conventional estimation. The null hypothesis that the tax λ's are jointly zero is strongly rejected for interest-bearing accounts, equities and mutual funds, and for retirement accounts in the probit and tobit models. We can clearly not reject it for mutual funds in the linear and tobit cases. In general, this test provides strong support for the need to control for unobserved heterogeneity in almost all the asset equations.

Finally, a test of the null hypothesis that the tax β's are constant over time produces more mixed results. The hypothesis cannot be rejected in the tobit and probit models, but can be rejected at a reasonable significance level for most assets in the linear equations.

Distinguishing unobserved heterogeneity from state dependence

Thus far, my estimates have demonstrated that estimators that control for unobserved heterogeneity significantly reduce the estimated impact of taxes on portfolio choices. This result essentially derives from the fact that large changes in tax incentives, particularly those introduced in 1986, did not appear to cause significant changes in asset allocations within the three years immediately following the change. One alternative explanation would be that portfolios adjust slowly, in part due to transaction costs involved in asset sales. This hypothesis implies some degree of state dependence in the dependent variable; i.e. a lagged value of the dependent variable may be an important explanatory variable in the asset share equations. In general, it is difficult to distinguish state dependence from unobserved heterogeneity. A regression which controls for state dependence may find a large significant positive effect of the lagged dependent variable on the current value of the dependent variable, and interpret it as evidence of slow adjustment. On the other hand, even if the lagged dependent variable really has no independent effect on the current value, it may enter significantly because all values of the dependent variable across time are correlated with any time-invariant unobserved heterogeneity. Conversely, a fixed-effects approach may find that changes in some explanatory variable over time have a weak effect on changes in the dependent variable, but this could merely reflect a slow adjustment process.

Distinguishing these two hypotheses is impossible with only two periods of data, but a third period is available if the 1986 wave of the SCF panel is added in. The balanced three-year panel reduces the cross-sectional sample size to 796 households, while increasing the total number of observations to three times that (2,388). Adding the 1986 panel necessitates some important changes, because the 1986 survey is missing some variables and combines a number of other variables into single categories. The major changes are as follows. Tax-free and taxable bonds must be combined into a single category, and so must equities and mutual funds. No information is available on defined contribution pension plans, so they are removed from net worth and retirement accounts. Insufficient information is available to construct the pension benefit variable, so it is omitted. Finally, some modifications are made in the procedure for calculating tax rates to account for somewhat less complete information in 1986.[52]

Because the linear and tobit models in the previous section led to qualitatively similar results, and because distinguishing unobserved heterogeneity and state dependence is much more econometrically tractable in the linear case, this section will focus on linear estimates.[53] In that case, the following approach is appropriate. Suppose that asset shares are determined by the following equation:

$$s_{jit} = a_i + b_{0t} + b_t t + b_X X + g s_{jit-1} + e_{it} \tag{11}$$

In this framework, a change in t or an X variable in one period will lead to further changes in later periods through its influence on the lagged value of s in future equations. A large positive value of γ would indicate a slow adjustment process. If γ is between zero and one in absolute value, the long-run effect of τ on s will be $\beta_\tau/(1-\gamma)$.[54]

The effects of unobserved heterogeneity can be removed, while still controlling for state dependence, simply by taking the first difference of equation (11).[55] I follow this approach, regressing the change in s between 1986 and 1989 on changes in the explanatory variables over that period, and the change in s between 1983 and 1986. One additional problem is that the lagged value of s will be endogenous if the error term is serially correlated. To address this, I use lagged values of the changes in τ and X as instruments for the lagged change in s, which will be consistent provided the lagged changes in τ and X are uncorrelated with current change in ε.[56] For comparison purposes, I also estimate standard cross-sectional and fixed effects approaches, as well as a lagged dependent variable model without fixed effects.

Table 14 illustrates the results. The first row shows the pooled cross-section results including data from all three years. Results are quite similar to previous cross-sectional findings, with a strong influence of taxes on asset shares. The second through fourth rows show the results of a lagged dependent variable model estimated on the 1986 and 1989 data. Marginal tax rate coefficients are reduced substantially compared to the first specification. But the long run effects in the case of interest-bearing accounts and equities/mutual funds are restored to high levels. The lagged asset shares all enter positively and significantly, suggesting a slow adjustment process (or, possibly, an important role for persistent unobserved heterogeneity). For retirement accounts, the lagged share exceeds one, which would cause the long-run effects of all variables to diverge to infinity.

The fifth row presents conventional fixed effects estimates based on all three years of data. Similar to earlier findings, the tax coefficients are significantly smaller than in the cross section. However, the tax effects are stronger than was found in the fixed-effects estimates on the panel including only 1983 and 1989, back in the second row of Table 4. All of the tax effects except for bonds are significantly different from zero. So there does appear to be some support in the data for the existence of a tax effect, albeit still a considerably smaller one than indicated by cross-sectional data.

The last three rows present estimates of the full model for distinguishing state dependence from unobserved heterogeneity. The marginal tax rate coefficients are similar to those in the previous fixed effects specification, with significant negative effects this time on interest-bearing accounts and bonds, and a significant positive effect on equities / mutual funds. Surprisingly, the lagged tax share coefficients all turn negative in this specification. This would suggest the opposite of slow adjustment – rather, the long-run effects of a permanent change in the marginal tax rate would be *smaller* than the short run effects. If equation (11) is a reasonable approximation of the adjustment process for asset shares[57], then this suggests that the primary finding of this paper, that unobserved heterogeneity significantly biases the marginal tax rate coefficients upwards, is robust. However, it is also possible that a richer model of the adjustment process of portfolios may be in order, which provides a potential avenue for future research.

Conclusion

This paper has provided new estimates of the effect of taxation on asset demands, utilizing a panel data set spanning a major change in tax law. The estimates suggest that unobserved heterogeneity plays an important role, and that cross-sectional estimates can be biased upwards to a large degree as a result, although a very flexible specification of income and net worth in a cross-sectional model appears to mitigate this bias. Overall, there is some evidence of an effect of marginal tax rates on portfolio choices in some of the panel specifications, but the size of the effect is considerably smaller than suggested by my cross-sectional estimates. Efforts to distinguish unobserved heterogeneity from state dependence do not change this conclusion substantially.

Appendix: Derivation of optimal asset share equation for two risky assets

This derivation is an adaptation of the approach in Friend and Blume (1975) to the case of two risky assets and differential taxation. Under the assumptions stated in the text, the progress of wealth over time is described by the equation:

$$W_{t+dt} = W_t\{1 + [s_1 r_1^e + (1 - s_1) r_2^e]dt + [s_1\sigma_1 + (1 - s_1)\sigma_2]z(t)\sqrt{dt} \qquad \text{(A.1)}$$

A Taylor-series approximation of $U(W_{t+dt})$ around W_t, dropping terms invloving dt raised to the 2nd power or higher, yields:

$$\begin{aligned} E[U(W_{t+dt})] &= U(W_t) + U'(W_t)W_t[s_1 r_1^e + (1 - s_1) r_2^e]dt \\ &+ \frac{1}{2}U''(W_t)W_t^2[s_1^2\sigma_1^2 + (1 - s_1)^2\sigma_2^2 + 2s_1(1 - s_1)\sigma_{12}]dt \end{aligned} \qquad \text{(A.2)}$$

Maximizing (2) with respect to s_1 yields the first order condition:

$$U'(W_t)W_t(r_1^e - r_2^e) + U''(W_t)W_t^2[s_1\sigma_1^2 - (1 - s_1)\sigma_2^2 + \sigma_{12} - 2s_1\sigma_{12}] = 0 \qquad \text{(A.3)}$$

Solving for s_1 yields the optimal asset share equation:

$$s_1 = \frac{1}{c_t}\frac{\left(r_1^e - r_2^e\right)}{\left(\sigma_1^2 + \sigma_2^2 - 2\sigma_{12}\right)} + \frac{\sigma_2^2 - \sigma_{12}}{\left(\sigma_1^2 + \sigma_2^2 - 2\sigma_{12}\right)} \qquad \text{(A.4)}$$

c_t is the coefficient of relative risk aversion, $W_t[-U''(W_t)/U'(W_t)]$.

If there is an income tax rate τ, and each asset j has a fraction f_j subject to tax, then the expected rate of return is replaced by the after-tax expected rate of return, $(1-f_j\tau)r_j^e$. The variance becomes:

$$E\left[(1 - f_j\tau)r_j - (1 - f_j\tau)r_j^e\right]^2 = (1 - f_j\tau)^2 E\left[r_j - r_j^e\right]^2 = (1 - f_j\tau)^2\sigma_j^2 \qquad \text{(A.5)}$$

The covariance between assets j and k becomes:

$$\begin{aligned} &E\left[(1 - f_j\tau)(r_j - r_j^e)(1 - f_k\tau)(r_k - r_k^e)\right] = (1 - f_j\tau)(1 - f_k\tau)E\left[(r_j - r_j^e)(r_k - r_k^e)\right] \\ &= (1 - f_j\tau)(1 - f_k\tau)\sigma_{jk} \end{aligned} \qquad \text{(A.6)}$$

Substituting the tax-adjusted returns, variances, and covariances into (A.4) yields equation (1) in the text.

Table 1. Unweighted means of all variables, 1983-89 panel

	Pooled	1983	1989
Share of financial assets held in:			
Interest-bearing accounts	0.435	0.454	0.416
Taxable bonds	0.049	0.048	0.050
Taxable mutual funds	0.024	0.015	0.033
Corporate equities	0.166	0.182	0.149
Tax-free bonds	0.059	0.053	0.065
Retirement accounts	0.268	0.248	0.287
Explanatory variables			
Marginal tax rate	0.28	0.33	0.23
Labor income (100,000s)	1.08	0.96	1.21
Net worth (millions)	2.44	1.99	2.88
Financial assets / total assets	0.33	0.32	0.35
Pension benefit (thousands)	1.73	1.80	1.65
Some college	0.15		
College graduate	0.27		
Graduate degree	0.26		
Risk preference: Average	0.48		
High	0.21		
Very high	0.07		
Liquidity preference: Average	0.39		
High	0.33		
Very high	0.11		
Married	0.79	0.80	0.78
Age /100	0.56	0.53	0.59
(Age/100) squared	0.33	0.30	0.37
Number of observations	1968	984	984

All dollar amounts are in constant 1989 dollars. Variables shown only once are time invariant. Sample excludes households with heads below age 25 in 1983, and households with less than $2,000 in financial assets in either year.

Table 2. Portfolio shares by tax group and year

	1983	1989	Change
MTR >= 48% in 1983 (n=313)			
Interest-bearing accounts	0.227	0.236	0.009
Taxable bonds	0.061	0.067	0.006
Taxable mutual funds	0.022	0.034	0.012
Corporate equities	0.351	0.275	-0.076
Tax-free bonds	0.121	0.146	0.026
Retirement accounts	0.218	0.242	0.024
MTR < 48% in 1983 (n=671)			
Interest-bearing accounts	0.560	0.499	-0.060
Taxable bonds	0.042	0.042	0.000
Taxable mutual funds	0.012	0.032	0.020
Corporate equities	0.103	0.091	-0.012
Tax-free bonds	0.022	0.027	0.006
Retirement accounts	0.262	0.308	0.047

Based on unweighted data from 1983-89 SCF panel.

Table 3. Share of households owning each type of asset

	Pooled	1983	1989
Share of observations owning:			
Interest-bearing accounts	0.999	0.998	1.000
Taxable bonds	0.387	0.417	0.357
Taxable mutual funds	0.161	0.125	0.196
Corporate equities	0.506	0.519	0.493
Tax-free bonds	0.253	0.235	0.271
Retirement accounts	0.672	0.665	0.679
Share holding interest-bearing accounts only	0.134	0.127	0.140

Based on unweighted data from 1983-89 SCF panel.

Table 4. Marginal effect of tax rate on portolio share, linear 2SLS regressions

	Dependent variable					
	Interest-bearing accounts	Taxable bonds	Mutual funds	Equities	Tax-free bonds	Retire-ment accounts
Time-invariant tax coefficient						
Pooled cross-section	-0.906	0.017	0.063	0.265	0.263	0.298
	(0.075)	(0.035)	(0.023)	(0.055)	(0.035)	(0.073)
Fixed effects	-0.186	-0.056	0.064	0.130	-0.035	0.084
	(0.108)	(0.054)	(0.036)	(0.072)	(0.048)	(0.101)
Correlated random effects	-0.186	-0.056	0.064	0.130	-0.035	0.084
	(0.108)	(0.054)	(0.036)	(0.071)	(0.048)	(0.101)
Time-varying tax coefficient						
Pooled cross-section						
1983	-0.873	0.031	0.043	0.311	0.235	0.254
	(0.078)	(0.036)	(0.024)	(0.057)	(0.037)	(0.076)
1989	-1.029	-0.034	0.138	0.096	0.370	0.460
	(0.136)	(0.063)	(0.042)	(0.100)	(0.064)	(0.133)
Correlated random effects						
1983	-0.174	-0.066	0.071	0.116	-0.028	0.082
	(0.109)	(0.055)	(0.036)	(0.072)	(0.048)	(0.102)
1989	-0.001	-0.205	0.166	-0.089	0.075	0.054
	(0.192)	(0.096)	(0.063)	(0.127)	(0.085)	(0.179)
Means of dependent variables	0.435	0.049	0.024	0.166	0.059	0.268

Standard errors in parentheses.

Table 5. Marginal effect of tax rate on overall taxable portfolio share, linear models with time-varying tax coefficients

Year	Specification	
	Pooled cross-section	Correlated random effects
1983	-0.701 (0.074)	-0.152 (0.105)
1989	-0.910 (0.128)	-0.117 (0.183)

Standard errors in parentheses.

Table 6. Marginal effect of tax rate on portolio share, tobit models

	Dependent variable						Sum of marginal effects
	Interest-bearing accounts	Taxable bonds	Mutual funds	Equities	Tax-free bonds	Retire-ment accounts	
Marginal effects evaluated at unweighted means of X							
Random effects							
1983	-0.815	0.040	0.027	0.263	0.080	0.329	-0.077
	(0.088)	(0.022)	(0.006)	(0.046)	(0.010)	(0.077)	
1989	-1.024	0.003	0.040	0.309	0.112	0.486	-0.074
	(0.140)	(0.037)	(0.010)	(0.088)	(0.018)	(0.130)	
Correlated random effects							
1983	-0.176	-0.012	0.021	0.060	0.015	0.079	-0.013
	(0.137)	(0.034)	(0.008)	(0.067)	(0.011)	(0.118)	
1989	-0.030	-0.073	0.025	0.014	0.012	0.042	-0.010
	(0.223)	(0.058)	(0.014)	(0.123)	(0.024)	(0.202)	
Weighted sum of individual marginal effects							
Random effects							
1983	-0.798	0.047	0.089	0.355	0.350	0.284	0.328
	(0.086)	(0.026)	(0.020)	(0.062)	(0.042)	(0.067)	
1989	-1.008	0.003	0.136	0.414	0.493	0.440	0.478
	(0.138)	(0.043)	(0.034)	(0.118)	(0.077)	(0.117)	
Correlated random effects							
1983	-0.171	-0.013	0.076	0.081	0.077	0.066	0.115
	(0.133)	(0.039)	(0.028)	(0.091)	(0.056)	(0.098)	
1989	-0.029	-0.093	0.103	0.018	0.060	0.038	0.097
	(0.219)	(0.073)	(0.056)	(0.165)	(0.127)	(0.182)	
Means of dependent variables	0.435	0.049	0.024	0.166	0.059	0.268	

Notes: The weighted sum of individual marginal effects is weighted by the product of the SCF population weights and the amount of financial assets in each year. Standard errors in parentheses.

Table 7. Full results for random-effects tobit: marginal effects evaluated at sample means

	Interest-bearing accounts		Taxable bonds		Mutual funds	
	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error
MTR, 1983	-0.815	(0.088)	0.040	(0.022)	0.027	(0.006)
MTR, 1989	-1.024	(0.140)	0.003	(0.037)	0.040	(0.010)
Constant, 1983	0.677	(0.124)	-0.052	(0.034)	-0.044	(0.008)
Constant, 1989	0.595	(0.125)	-0.043	(0.034)	-0.040	(0.008)
Labor income	-0.201	(0.348)	0.128	(0.059)	0.002	(0.012)
Net worth	-0.423	(0.151)	0.036	(0.026)	-0.008	(0.008)
Financial/total assets	-0.222	(0.033)	0.037	(0.009)	0.007	(0.002)
Pension benefit	0.026	(0.163)	-0.082	(0.067)	0.009	(0.008)
Some college	-0.104	(0.026)	0.012	(0.007)	0.001	(0.002)
College graduate	-0.093	(0.027)	0.014	(0.007)	0.003	(0.002)
Graduate degree	-0.115	(0.027)	0.009	(0.007)	0.003	(0.002)
Risk pref.: Avg.	-0.091	(0.023)	-0.001	(0.006)	0.005	(0.002)
High	-0.076	(0.032)	-0.014	(0.008)	0.004	(0.002)
Very high	-0.112	(0.040)	-0.014	(0.011)	0.001	(0.003)
Liquidity pref.: Avg.	-0.008	(0.028)	0.007	(0.007)	0.002	(0.001)
High	0.052	(0.028)	0.002	(0.007)	0.003	(0.002)
Very high	0.010	(0.033)	0.010	(0.010)	0.001	(0.002)
Married	-0.107	(0.019)	0.005	(0.005)	-0.001	(0.001)
Age /100	-2.578	(0.427)	-0.019	(0.122)	0.025	(0.028)
(Age/100) squared	2.486	(0.373)	-0.018	(0.107)	-0.012	(0.024)
Scale factor	0.978		0.275		0.039	

Notes: Net worth is in hundreds of millions, labor income is in tens of millions, and pension benefits are in hundreds of thousands. "Scale factor" is the factor by which the Tobit coefficients are multiplied to find the marginal effects.

Table 7 (continued). Full results for random-effects tobit: marginal effects evaluated at sample means

	Equities		Tax-free bonds		Retirement accounts	
	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error
MTR, 1983	0.263	(0.046)	0.080	(0.010)	0.329	(0.077)
MTR, 1989	0.309	(0.088)	0.112	(0.018)	0.486	(0.130)
Constant, 1983	-0.489	(0.078)	-0.109	(0.019)	-0.828	(0.123)
Constant, 1989	-0.514	(0.081)	-0.107	(0.019)	-0.743	(0.127)
Labor income	0.052	(0.134)	0.002	(0.015)	-0.041	(0.366)
Net worth	0.336	(0.066)	0.033	(0.012)	-0.274	(0.161)
Financial/total assets	0.164	(0.017)	0.016	(0.003)	-0.031	(0.033)
Pension benefit	0.039	(0.078)	-0.008	(0.011)	-0.046	(0.122)
Some college	0.052	(0.016)	0.012	(0.004)	0.046	(0.024)
College graduate	0.072	(0.016)	0.018	(0.003)	0.000	(0.024)
Graduate degree	0.071	(0.016)	0.019	(0.003)	0.035	(0.025)
Risk pref.: Avg.	0.065	(0.015)	0.007	(0.003)	0.013	(0.020)
High	0.120	(0.019)	0.003	(0.004)	-0.027	(0.027)
Very high	0.133	(0.023)	0.000	(0.005)	-0.004	(0.036)
Liquidity pref.: Avg.	-0.001	(0.015)	-0.004	(0.003)	0.035	(0.024)
High	-0.035	(0.016)	-0.005	(0.003)	0.005	(0.026)
Very high	-0.013	(0.023)	-0.017	(0.006)	-0.005	(0.031)
Married	0.037	(0.012)	0.003	(0.002)	0.087	(0.018)
Age /100	0.612	(0.266)	0.065	(0.061)	3.676	(0.449)
(Age/100) squared	-0.354	(0.231)	0.001	(0.051)	-4.097	(0.418)
Scale factor	0.480		0.075		0.697	

Notes: Net worth is in hundreds of millions, labor income is in tens of millions, and pension benefits are in hundreds of thousands. "Scale factor" is the factor by which the Tobit coefficients are multiplied to find the marginal effects.

Table 8. Full results for correlated random-effects tobit: marginal effects evaluated at sample means

	Interest-bearing accounts		Taxable bonds		Mutual funds	
	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error
Marginal effect on S (βΦ)						
Marginal tax rate, 1983	-0.176	(0.137)	-0.012	(0.034)	0.021	(0.008)
Marginal tax rate, 1989	-0.030	(0.223)	-0.073	(0.058)	0.025	(0.014)
Constant, 1983	0.681	(0.147)	-0.051	(0.040)	-0.046	(0.010)
Constant, 1989	0.365	(0.145)	-0.038	(0.040)	-0.046	(0.010)
Labor income	-0.183	(0.308)	0.147	(0.121)	0.003	(0.014)
Net worth	-0.343	(0.306)	0.158	(0.072)	-0.005	(0.024)
Financial/total assets	-0.104	(0.055)	0.016	(0.015)	0.002	(0.003)
Pension benefit	0.236	(0.292)	-0.062	(0.085)	0.027	(0.014)
Married	-0.015	(0.040)	-0.016	(0.011)	-0.003	(0.002)
(Age/100) squared	3.488	(0.782)	-0.111	(0.212)	0.075	(0.045)
Correlation with fixed effect (times probability>0) (λΦ)						
Marginal tax rate, 1983	-0.805	(0.114)	0.083	(0.032)	0.003	(0.006)
Marginal tax rate, 1989	-0.750	(0.191)	0.012	(0.053)	0.020	(0.013)
Labor income, 1983	0.378	(0.965)	-0.348	(0.226)	0.000	(0.046)
Labor income, 1989	0.098	(0.453)	-0.039	(0.119)	0.002	(0.013)
Net worth, 1983	0.263	(0.389)	-0.248	(0.063)	-0.008	(0.019)
Net worth, 1989	-0.122	(0.315)	0.027	(0.087)	0.002	(0.014)
Financial/total assets, 1983	-0.013	(0.055)	0.003	(0.014)	0.003	(0.003)
Financial/total assets, 1989	-0.149	(0.053)	0.030	(0.015)	0.005	(0.003)
Pension benefit, 1983	-0.293	(0.236)	0.014	(0.072)	-0.012	(0.013)
Pension benefit, 1989	-0.133	(0.228)	-0.070	(0.071)	-0.024	(0.016)
Married, 1983	-0.014	(0.038)	0.011	(0.010)	0.001	(0.002)
Married, 1989	-0.082	(0.037)	0.016	(0.010)	0.001	(0.003)
(Age/100) squared, 1983	13.827	(4.564)	0.539	(1.292)	-0.546	(0.297)
(Age/100) squared, 1989	-15.532	(4.085)	-0.419	(1.152)	0.425	(0.262)
Scale factor	0.981		0.271		0.035	

Notes: Net worth is in hundreds of millions, labor income is in tens of millions, and pension benefits are in hundreds of thousands. "Scale factor" is the factor by which the Tobit coefficients are multiplied to find the marginl effects.

Table 8 (continued). Full results for correlated random-effects tobit: marginal effects evaluated at sample means

	Equities		Tax-free bonds		Retirement accounts	
	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error
Marginal effect on S (βΦ)						
Marginal tax rate, 1983	0.060	(0.067)	0.015	(0.011)	0.079	(0.118)
Marginal tax rate, 1989	0.014	(0.123)	0.012	(0.024)	0.042	(0.202)
Constant, 1983	-0.417	(0.097)	-0.101	(0.019)	-1.008	(0.141)
Constant, 1989	-0.386	(0.098)	-0.104	(0.018)	-0.710	(0.139)
Labor income	0.097	(0.118)	-0.009	(0.024)	-0.169	(0.401)
Net worth	0.092	(0.159)	-0.007	(0.018)	0.087	(0.366)
Financial/total assets	0.098	(0.024)	0.004	(0.004)	0.013	(0.050)
Pension benefit	-0.047	(0.098)	-0.017	(0.016)	-0.199	(0.170)
Married	0.032	(0.022)	0.002	(0.004)	0.048	(0.032)
(Age/100) squared	-0.605	(0.416)	0.131	(0.082)	-3.800	(0.755)
Correlation with fixed effect (times probability>0) (λΦ)						
Marginal tax rate, 1983	0.368	(0.065)	0.076	(0.011)	0.256	(0.111)
Marginal tax rate, 1989	0.239	(0.114)	0.071	(0.025)	0.454	(0.170)
Labor income, 1983	0.287	(0.495)	-0.012	(0.074)	-0.038	(0.795)
Labor income, 1989	-0.145	(0.158)	0.017	(0.019)	0.106	(0.271)
Net worth, 1983	0.085	(0.105)	0.008	(0.019)	-0.228	(0.261)
Net worth, 1989	0.244	(0.145)	0.026	(0.017)	-0.437	(0.298)
Financial/total assets, 1983	0.072	(0.031)	0.005	(0.005)	-0.093	(0.050)
Financial/total assets, 1989	0.031	(0.028)	0.012	(0.005)	-0.008	(0.048)
Pension benefit, 1983	0.224	(0.156)	-0.025	(0.022)	-0.007	(0.144)
Pension benefit, 1989	0.033	(0.114)	0.018	(0.021)	0.336	(0.190)
Married, 1983	-0.023	(0.025)	0.004	(0.004)	0.001	(0.032)
Married, 1989	0.027	(0.021)	-0.003	(0.005)	0.033	(0.032)
(Age/100) squared, 1983	-1.275	(3.059)	-0.421	(0.515)	-31.348	(4.559)
(Age/100) squared, 1989	1.848	(2.742)	0.294	(0.476)	31.041	(4.124)
Scale factor	0.476		0.066		0.696	

Notes: Net worth is in hundreds of millions, labor income is in tens of millions, and pension benefits are in hundreds of thousands. "Scale factor" is the factor by which the Tobit coefficients are multiplied to find the marginl effects.

Table 9. Replication of Poterba and Samwick tobit results, using full 1983 SCF cross-section

Marginal effects evaluated at unweighted means of explanatory varibales	Dependent variable: Interest-bearing accounts	Taxable bonds	Mutual funds	Equities	Tax-free bonds	Retire-ment accounts	Other financial assets	Sum of marginal effects
Poterba & Samwick	-0.183*	0.002	-0.001	-0.015	0.009	0.172*	0.017	0.012
Replication	-0.309 (0.107)	0.010 (0.023)	-0.008 (0.014)	-0.036 (0.029)	0.001 (0.002)	0.209 (0.048)	0.092 (0.073)	-0.041
Replication, replacing income and net worth dummies with continuous variables	-0.831 (0.059)	0.090 (0.013)	0.030 (0.008)	0.208 (0.018)	0.025 (0.004)	0.331 (0.027)	0.152 (0.040)	0.004
Means of dependent variables	0.600	0.031	0.006	0.074	0.017	0.121	0.151	

Notes: Standard errors in parentheses. Results in first row are from Poterba and Samwick (1999). Results in second row represent my approximate replication of Poterba and Samwick on the full 1983 SCF cross-section. See text for details.
* indicates statistical significance at the 5% level (marginal effect standard errors are not reported by Poterba and Samwick).

Table 10. Marginal effect of tax rate on probability of asset ownership, probit models

	Asset				
	Taxable bonds	Mutual funds	Equities	Tax-free bonds	Retire-ment accounts
Marginal effects evaluated at unweighted means of X					
Random effects					
1983	0.499	0.216	1.253	0.674	1.126
	(0.187)	(0.060)	(0.227)	(0.078)	(0.186)
1989	0.408	0.399	1.715	0.848	1.277
	(0.325)	(0.111)	(0.411)	(0.152)	(0.321)
Correlated random effects					
1983	0.215	0.107	0.271	0.154	0.360
	(0.269)	(0.078)	(0.364)	(0.110)	(0.295)
1989	0.110	0.199	0.392	0.102	0.054
	(0.462)	(0.155)	(0.632)	(0.255)	(0.483)
Weighted sum of individual marginal effects					
Random effects					
1983	0.478	0.557	0.671	1.060	0.702
	(0.179)	(0.156)	(0.122)	(0.123)	(0.116)
1989	0.400	1.043	0.926	1.388	0.813
	(0.319)	(0.289)	(0.222)	(0.249)	(0.204)
Correlated random effects					
1983	0.201	0.301	0.137	0.255	0.203
	(0.251)	(0.220)	(0.184)	(0.182)	(0.166)
1989	0.107	0.580	0.200	0.174	0.031
	(0.450)	(0.453)	(0.322)	(0.433)	(0.283)
Share of observations owning asset	0.387	0.161	0.506	0.253	0.672

Notes: The weighted sum of individual marginal effects is weighted by the product of the SCF population weights and the amount of financial assets in each year. Interest-bearing accounts not shown because 99.9 percent of observations hold them. Standard errors are in parentheses.

Table 11. Full results for random-effects probit: marginal effects evaluated at sample means

	Taxable bonds		Mutual funds		Equities		Tax-free bonds		Retirement accounts	
	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error
MTR, 1983	0.499	(0.187)	0.216	(0.060)	1.253	(0.227)	0.674	(0.078)	1.126	(0.186)
MTR, 1989	0.408	(0.325)	0.399	(0.111)	1.715	(0.411)	0.848	(0.152)	1.277	(0.321)
Constant, 1983	-0.197	(0.284)	-0.445	(0.081)	-2.120	(0.358)	-1.028	(0.151)	-2.001	(0.290)
Constant, 1989	-0.191	(0.287)	-0.432	(0.083)	-2.227	(0.376)	-0.988	(0.157)	-1.816	(0.300)
Labor income	0.948	(0.698)	0.026	(0.088)	0.857	(0.534)	-0.044	(0.161)	0.324	(0.693)
Net worth	-0.086	(0.324)	0.002	(0.067)	2.058	(0.358)	0.445	(0.085)	-0.047	(0.335)
Financial / total assets	0.363	(0.073)	0.095	(0.019)	0.780	(0.093)	0.229	(0.030)	0.252	(0.076)
Pension benefit	-0.721	(0.505)	0.077	(0.069)	0.114	(0.325)	-0.235	(0.105)	0.245	(0.426)
Some college	0.039	(0.068)	0.009	(0.018)	0.195	(0.071)	0.092	(0.033)	0.087	(0.059)
College graduate	0.081	(0.061)	0.032	(0.016)	0.278	(0.068)	0.158	(0.029)	0.042	(0.056)
Graduate degree	0.105	(0.063)	0.043	(0.016)	0.269	(0.068)	0.173	(0.028)	0.182	(0.059)
Risk preference										
Average	0.030	(0.056)	0.046	(0.016)	0.220	(0.061)	0.073	(0.026)	0.060	(0.052)
High	-0.056	(0.070)	0.044	(0.018)	0.412	(0.082)	0.044	(0.030)	0.013	(0.064)
Very high	-0.127	(0.099)	0.020	(0.024)	0.353	(0.106)	0.041	(0.039)	0.042	(0.084)
Liquidity pref.										
Average	0.063	(0.058)	0.021	(0.014)	0.014	(0.066)	-0.030	(0.021)	0.138	(0.056)
High	-0.002	(0.062)	0.027	(0.015)	-0.098	(0.070)	-0.032	(0.024)	0.023	(0.059)
Very high	0.058	(0.082)	-0.001	(0.026)	-0.021	(0.095)	-0.114	(0.044)	-0.072	(0.077)
Married	0.120	(0.049)	0.002	(0.012)	0.175	(0.052)	0.051	(0.020)	0.223	(0.047)
Age /100	-0.861	(1.014)	0.341	(0.277)	3.049	(1.233)	0.798	(0.500)	7.114	(1.040)
(Age/100) squared	0.244	(0.896)	-0.222	(0.241)	-2.194	(1.087)	-0.251	(0.431)	-7.796	(0.974)
Scale factor	0.352		0.073		0.398		0.134		0.270	

Notes: Net worth is in hundreds of millions, labor income is in tens of millions, and pension benefits are in hundreds of thousands. "Scale factor" is the factor by which the probit coefficients are multiplied to find the marginal effects.

Table 12. Full results for correlated random-effects probit: marginal effects evaluated at sample means

	Taxable bonds		Mutual funds		Equities		Tax-free bonds		Retirement accounts	
	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error	marg. effect	std. error
Marginal effects on S (βΦ)										
MTR, 1983	0.215	(0.269)	0.107	(0.078)	0.271	(0.364)	0.154	(0.110)	0.360	(0.295)
MTR, 1989	0.110	(0.462)	0.199	(0.155)	0.392	(0.632)	0.102	(0.255)	0.054	(0.483)
Constant, 1983	-0.218	(0.347)	-0.479	(0.100)	-1.806	(0.416)	-0.916	(0.165)	-2.588	(0.367)
Constant, 1989	-0.324	(0.350)	-0.497	(0.100)	-1.655	(0.410)	-0.915	(0.156)	-2.129	(0.360)
Labor income	1.190	(0.950)	0.011	(0.107)	1.377	(1.009)	-0.163	(0.229)	-0.028	(1.014)
Net worth	0.376	(0.755)	0.037	(0.237)	1.376	(1.008)	0.016	(0.228)	0.329	(1.217)
Financial/total assets	0.177	(0.118)	0.042	(0.033)	0.561	(0.147)	0.093	(0.041)	0.309	(0.125)
Pension benefit	-0.482	(0.654)	0.329	(0.141)	-1.158	(0.844)	-0.411	(0.185)	-0.345	(0.509)
Married	-0.030	(0.090)	-0.012	(0.023)	0.132	(0.101)	0.060	(0.034)	0.152	(0.094)
(Age/100) squared	0.682	(1.670)	0.652	(0.463)	-3.587	(2.279)	0.951	(0.832)	-5.236	(1.712)
Correlation with fixed effect (times prob. >0) (λΦ)										
MTR, 1983	0.667	(0.287)	0.101	(0.056)	1.577	(0.351)	0.693	(0.097)	1.023	(0.269)
MTR, 1989	-0.148	(0.448)	0.231	(0.130)	0.933	(0.546)	0.522	(0.238)	1.059	(0.404)
Labor income, 1983	-2.637	(1.926)	0.049	(0.415)	2.442	(2.740)	-0.020	(0.676)	1.211	(1.742)
Labor income, 1989	-0.160	(0.720)	0.059	(0.098)	-1.029	(0.880)	0.162	(0.129)	0.095	(0.550)
Net worth, 1983	-1.349	(0.600)	-0.107	(0.149)	-0.920	(0.713)	-0.074	(0.212)	-0.378	(0.629)
Net worth, 1989	0.329	(0.628)	0.017	(0.137)	1.215	(1.136)	0.388	(0.174)	-0.699	(0.742)
Financial / assets, '83	0.104	(0.125)	0.038	(0.028)	0.165	(0.142)	0.058	(0.045)	0.012	(0.122)
Financial / assets, '89	0.249	(0.117)	0.041	(0.031)	0.182	(0.140)	0.133	(0.044)	-0.136	(0.113)
Pension benefit, '83	-0.224	(0.681)	-0.199	(0.120)	1.629	(0.984)	-0.162	(0.194)	0.487	(0.598)
Pension benefit, '89	-0.425	(0.523)	-0.294	(0.157)	1.692	(0.918)	0.296	(0.158)	0.875	(0.417)
Married, 1983	0.132	(0.087)	0.011	(0.023)	-0.098	(0.107)	0.017	(0.041)	0.097	(0.085)
Married, 1989	0.094	(0.086)	0.014	(0.027)	0.126	(0.098)	-0.039	(0.039)	-0.023	(0.080)
(Age/100) sqrd, '83	6.967	(11.228)	-5.904	(2.959)	-6.703	(12.612)	-3.601	(4.484)	-70.872	(11.347)
(Age/100) sqrd, '89	-7.503	(10.052)	4.738	(2.625)	9.814	(11.477)	2.655	(4.158)	67.273	(10.258)
Scale factor	0.350		0.067		0.397		0.118		0.261	

Notes: Net worth is in hundreds of millions, labor income is in tens of millions, and pension benefits are in hundreds of thousands.
"Scale factor" is the factor by which the Tobit coefficients are multiplied to find the marginal effects.

Table 13. Significance levels for selected hypothesis tests

Null hypothesis	Dependent variable					
	Interest-bearing accounts	Taxable bonds	Mutual funds	Equities	Tax-free bonds	Retire-ment accounts
Linear correlated random effects						
All X variables jointly uncorrelated with fixed effect	0.000	0.028	0.706	0.000	0.000	0.000
Marginal tax rates jointly uncorrelated with fixed effect	0.000	0.167	0.396	0.000	0.000	0.113
Marginal tax rate betas constant over time	0.200	0.040	0.033	0.022	0.082	0.827
Tobit correlated random effects						
X variables jointly uncorrelated with fixed effect	0.000	0.001	0.199	0.000	0.000	0.000
Marginal tax rates jointly uncorrelated with fixed effect	0.000	0.021	0.201	0.000	0.000	0.001
Marginal tax rate betas constant over time	0.359	0.121	0.637	0.619	0.871	0.808
Probit correlated random effects						
X variables jointly uncorrelated with fixed effect		0.075	0.070	0.000	0.000	0.000
Marginal tax rates jointly uncorrelated with fixed effect		0.061	0.027	0.000	0.000	0.000
Marginal tax rate betas constant over time		0.756	0.420	0.801	0.808	0.387

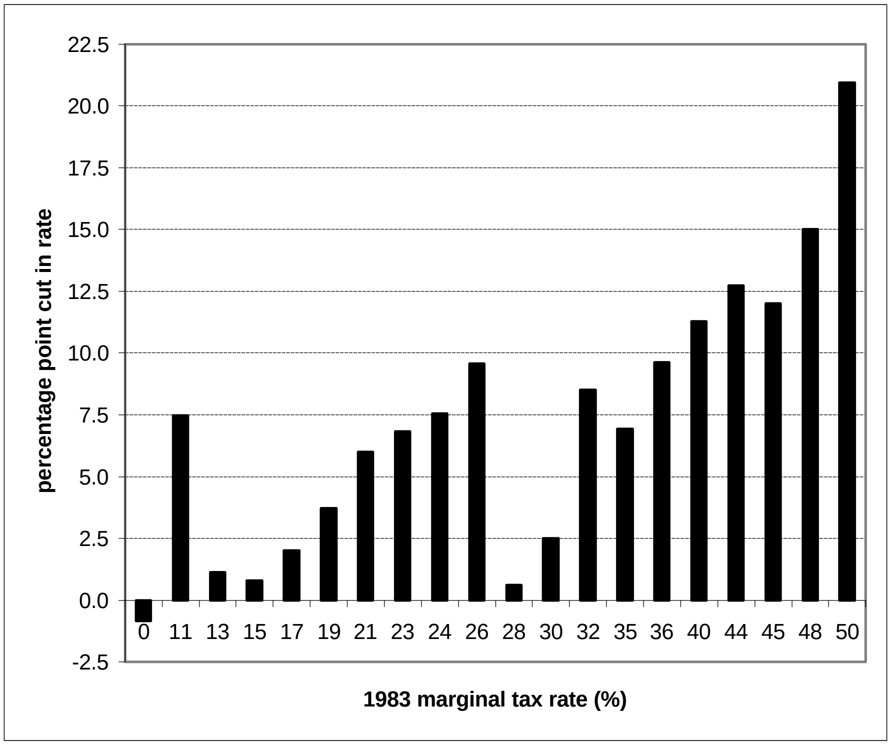
Note: Probabilities for linear model are based on F-tests of the joint restrictions. Probabilities for Tobit and probit models are based on Wald chi-squared tests of the joint restrictions.

Table 14. Results of linear 2SLS regressions on three-year panel including 1986: distinguishing unobserved heterogeneity from state dependence

		Dependent variable			
		Interest-bearing accounts	All bonds	Equities & mutual funds	Retire-ment accounts
Pooled cross section					
Marginal tax rate		-1.001	0.323	0.392	0.286
		(0.081)	(0.054)	(0.065)	(0.074)
Pooled cross section,with lagged share					
Marginal tax rate	(β)	-0.507	-0.020	0.246	-0.123
		(0.139)	(0.133)	(0.080)	(0.130)
Lagged share	(γ)	0.444	0.857	0.487	1.304
		(0.099)	(0.278)	(0.097)	(0.291)
Marginal tax rate, long run	(β/(1−γ))	-0.911	-0.139	0.480	*
		(0.165)	(1.171)	(0.137)	
Fixed effects					
Marginal tax rate		-0.284	-0.118	0.262	0.198
		(0.137)	(0.094)	(0.108)	(0.051)
Fixed effects with lagged share					
Marginal tax rate	(β)	-0.232	-0.171	0.382	0.126
		(0.151)	(0.102)	(0.116)	(0.221)
Lagged share	(γ)	-0.696	-0.138	-0.462	-1.227
		(0.215)	(0.318)	(0.270)	(0.728)
Marginal tax rate, long run	(β/(1−γ))	-0.137	-0.150	0.261	*
		(0.087)	(0.104)	(0.102)	
Means of dependent variables		0.446	0.211	0.112	0.230

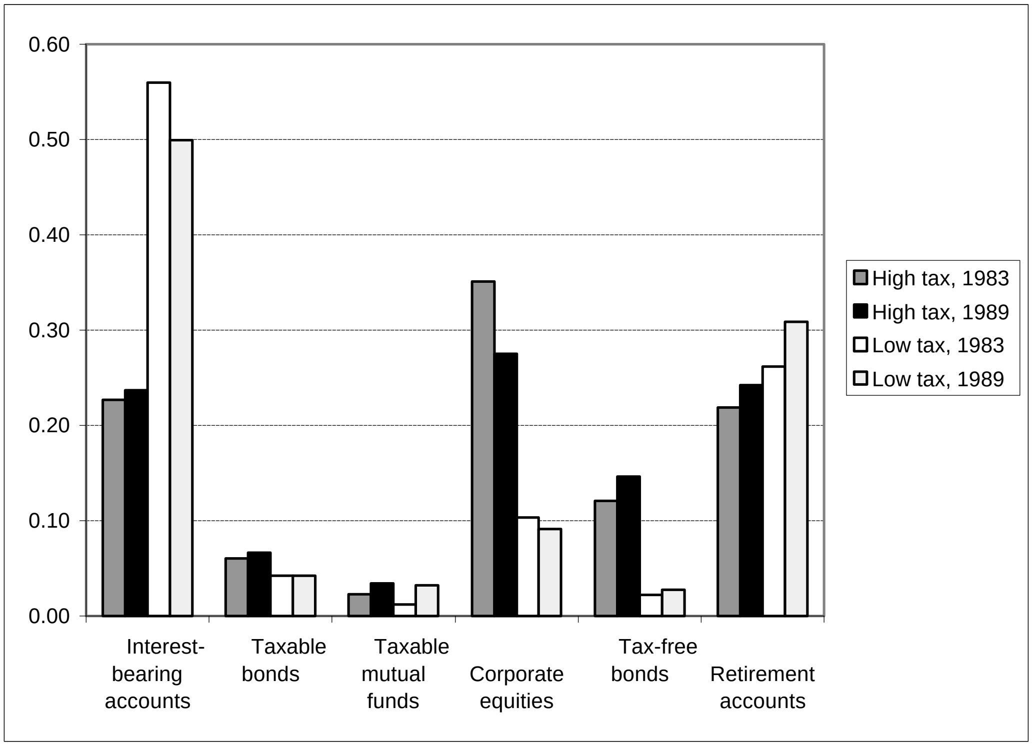
Standard errors in parentheses. Note that defined contribution pensions are excluded from retirement accounts, taxable and tax-free bonds are combined, and equities and mutual funds are combined because of limited information in the 1986 data. Lagged values of X are used as instruments for lagged dependent variable.
* If gamma is greater than or equal to one in absolute value, the long run effects diverge to infinity.

Figure 1. Average percentage point cut in marginal tax rate, 1983-89, by 1983 marginal tax rate



Notes: chart depicts average reduction in marginal tax rate in 1989 law compared to 1983 law, holding real income constant at 1983 level, for observations in 1983-89 SCF panel.

Figure 2. Share of financial assets held in each asset type, by tax group and year



Notes: "High tax" group had a marginal tax rate of 48% or above in 1983. "Low tax" group had a marginal tax rate below 48% in 1983. Data are unweighted.

Endnotes

[1] See, for example, Joint Committee on Taxation (1995).

[2] This is an adaptation of the approach in Friend and Blume (1975) to the case of two risky assets with differential taxation. Friend and Blume consider the choice between one risky and one risk-free asset, each of which is subject to the same rate of tax. This simplifies the optimal portfolio share equation somewhat. However, in reality, no asset is risk free – even treasury bills are subject to inflation risk (inflation indexed government bonds were not yet available during the period covered by this study). See Friedman and Roley (1979) for the derivation of the optimal portfolio share for *n* risky assets.

[3] Stiglitz (1969) examines how limited loss offsets affect optimal portfolio choices in a CAPM framework.

[4] See Gordon and Slemrod (1983) for a general equilibrium model of portfolio choice in which the government returns revenues to taxpayers as lump-sum transfers.

[5] Laroque (1987) and Constantinides (1986) develop portfolio choice models with transaction costs.

[6] In principle, it would be possible to hold negative shares of certain assets, for example by selling equities short. However, in practice, we do not observe this happening to a great extent, perhaps due to transaction costs and other constraints on short sales. Similarly, household debt might be viewed as a negative holding of interest-bearing assets such as bonds. Empirical work in this area has usually focused on the share of positive assets held in particular forms, without considering negative asset shares, and this study follows that convention.

[7] When applied to a dependent variable censored below at zero, this approach is also sometimes referred to as the Cragg (1971) model.

[8] Most studies using the Heckman two-step approach also transform the dependent variable into log form. Since a large number of observations have a share of zero, and the log of zero is undefined, this functional form does not work in a Tobit or linear estimate using the full sample. One could use an arbitrary approximation, such as adding .001 to each share, but this is somewhat arbitrary and the results can be very sensitive to the approximation used. In the Heckman two-step, the undefined log-zeros are not included in the share equation, and are included in the selection equations as zeros or ones. This is usually motivated by a desire to reduce heteroscedasticity, although it is unclear whether the problem is worse or better in a log-linear framework. Many people hold asset shares very close to zero, and the variance of the log share becomes arbitrarily large near zero. The log-linear model is also sometimes motivated by a version of the CAPM model with one riskless asset. In that model, the asset equations take the form $s=(1/c)f(t)$, where c is the coefficient of risk aversion and $f(t)$ is a function of t including variances, covariances, rates of return, and degrees of taxability (see King and Leape 1998 for an example). Taking the log of each side makes the coefficient of relative risk aversion additively separable from the tax rate. This is a desirable property for estimation if we believe the coefficient of relative risk aversion varies across individuals, but it depends on the unrealistic assumption of a riskless asset.

[9] See Vella (1998) for a discussion of these issues.

[10] To aid in identification, Hubbard includes some variables, such as education, in the selection equation but not the level equation. This identification strategy is somewhat arbitrary; for example, if education serves as a proxy for financial sophistication, it is very likely to influence both decisions.

[11] Unlike other data sets with portfolio information, such as the SCF, the data used by King and Leape contain information on state of residence, enabling them to calculate the combined federal and state marginal income tax rate. This aids in identification by providing a source of tax rate variation that is independent of income.

[12] An implication of equation (1) is that both the intercepts and slope parameters of the share equations may vary depending on the combination of assets owned. Given the very large number of possible asset combinations and the generally small number of available observations, incorporating making the model flexible along both dimensions would render estimation impossible.

[13] Samwick (2000) also performs a Tobit analysis similar to that in Poterba and Samwick (1999), finding similar results.

[14] Ioannides restricts his sample to households whose head is in the labor force in 1983, to maintain comparability with another study he was doing. This accounts for the difference in sample size between Guell and Ionnides.

[15] In Ioannides' study, the set of explanatory variables differs greatly between the probit and level equations; no rationale is provided for why variables belong in one equation but not the other.

[16] Kennickell and Starr-McCluer (1996) present some descriptive statistics from the 1983-1989 SCF panel, including information on portfolio allocations at different points in the income distribution over time. No econometric analysis of portfolio shares is performed. Bertaut (1998) utilizes the 1983-1989 SCF panel to examine the determinants of whether someone owns any corporate equities. She estimates random-effects probit model, which will be inconsistent if the explanatory variables are correlated with unobserved time-invariant effects. Tax rates are not included in the analysis.

[17] The SCF surveys are conducted in the spring of the year they are named for. The Tax Reform Act of 1986 was enacted in October, and did not begin to be implemented until 1987.

[18] The 1983-84 tax cut was also offset to some extent by inflationary bracket creep, as indexation of tax brackets did not begin until after 1984.

[19] Poterba and Samwick's "other financial assets," including life insurance and trust accounts, are not counted as part of financial assets in my analysis.

[20] U.S. savings bonds, which are included in this category, benefit from deferral of tax on interest until they reach maturity (usually in five years). They represent only a very small portion of the aggregate value of bonds in this category -- 1.4 percent in 1989.

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[21] The SCF asks respondents a specific question about AGI in 1989 and later years. In 1983, the AGI variable is an estimate constructed by the SCF staff based on answers to various other questions on income in the SCF.

[22] Internal Revenue Service, various years.

[23] Itemized deductions are assumed to equal 14.1 percent of AGI in 1983 and 11.4 percent of AGI in 1989. These represent aggregate itemized deductions (other than interest) as a percentage of aggregate AGI earned by itemizers in that year, calculated from IRS data. Calculated using the IRS Individual Model File public use data sets for 1983 and 1989. Taxpayers are assumed to itemize if imputed plus actual itemized deductions are greater than the standard deduction (or zero bracket amount in 1983). The SCF includes information on charitable contributions in 1989, but not in 1983, so I do not use them in the calculation.

[24] State of residence is available in 1983 for observations that are not part of the high-income over-sampling group. It is completely omitted in other years.

[25] It is possible that itemized deductions for interest payments could be endogenously related to financial asset shares as well. I also tried estimates where interest paid deductions were replaced with imputed values in the tax rate instrument. (The imputations were done in a similar way as for other, unobserved itemized deudctions.) This had little effect on the results.

[26] The major components of net worth, expressed as an unweighted average percentage of gross assets in my sample, are: financial assets (33.3 percent); owner occupied housing (34.7 percent); other real estate (9.9 percent); business assets (10.8 percent); other miscellaneous assets (11.3 percent); and total debt (12.1 percent).

[27] No attempt is made to include and estimate of social security wealth, due to the limited information available on past earnings.

[28] “Average” risk preference indicates willingness to “take average financial risks expecting to earn average returns.” “High” risk preference indicates willingness to “take above average financial risks expecting to earn above average returns.” The omitted category is “not willing to take any financial risks.”

[29] “Average” liquidity preference corresponds to a willingness to “tie up money for an intermediate period of time to earn above average returns.” “High” liquidity preference corresponds to a willingness to “tie up money for a short period of time to earn average returns.” The omitted category is willing to “tie up money for a long period of time to earn substantial returns.” A small number of observations (21 in the case of risk and 39 in the case of liquidity) did not answer the questions, and are assigned to the “average” group.

[30] It would be possible to include the time-invariant variables in the correlated random- effects equations, but not to separately identify the β’s and λ’s. Including these variables in the correlated random-effects specifications made no significant difference to the coefficients on the other variables.

[31] Age is the age of the person who was listed as head of household in 1983, even if the head of household changed in the intervening years, for example due to death.

[32] Estimating a two-limit tobit would have no affect at all on the estimates for any of the assets except for interest-bearing accounts, and even in that case the effect is imperceptible. Given the other complications in estimation, a one-limit tobit makes computation considerably easier.

[33] See Feenburg (1987) for further discussion of the potential inconsistency of cross-sectional marginal tax rate coefficients.

[34] An alternative explanation to be explored below would be that lagged tax rates influence later period's shares because portfolios adjust slowly over time, so that a lagged dependent variable belongs in the equation.

[35] See Maddala (1987) or Hsiao (1989) for a discussion of this issue.

[36] I use the random-effects probit and tobit estimators programmed into LIMDEP (Greene, 1998). Jakubson (1983, 1988) and Guell (1991) do not estimate this model by random-effects maximum-likelihood. Rather, they estimate a minimum-distance estimator that converts reduced-form parameters from tobits on the individual cross-sections into the structural parameters. The advantage of their approach is that it allows the correlation among an individual's error terms to vary across time. Maximum likelihood estimation of a random-effects tobit or probit model is straightforward if the correlation of an individual's error terms across time is restricted to be constant. Otherwise, computation would become very difficult, as it would require evaluation of a T-fold integral, where T is the number of time periods. However, the assumption of equicorrelation across an individual's error terms is not restrictive when there are only two periods. In that case, maximum-likelihood estimation is fully general and feasible, and more efficient than the approach used by Guell and Jakubson.

[37] Guell (1991) does not make any adjustments for the endogeneity of the tax rate. Poterba and Samwick (1999) and Samwick (2000) use a proxy for the marginal tax rate intended to remove the endogeneity. The proxy is the marginal tax rate calculated by replacing actual dividend, interest, and capital gains income with 5% of the value of total financial assets. Even if this variable were exogenous, consistent estimation would require that it be used as an instrument for the actual marginal tax rate, rather than just being included as a proxy.

[38] Newey's approach involves the following steps. The first stage is a set of linear regressions of the endogenous explanatory variables on the full set of exogenous instruments. The second stage is an estimate of the reduced form equation for the dependent variable (in this case, the asset share) by maximum likelihood tobit, including the set of exogenous instruments and the residuals from the first stage on the right-hand-side. Random effects are incorporated at this stage in my estimation. Consistent estimates of the structural parameters and the asymptotic covariance matrix are then calculated by applying a minimum-distance estimator to the reduced-form parameters. Newey demonstrates that this approach is consistent and asymptotically efficient for a general class of nonlinear limited dependent variable models including both the tobit and probit. It should be noted that the conditions required for consistency of this estimator are more restrictive than for linear 2SLS; for instance, the error terms in both the share equations and the first stage equations must be normally distributed.

[39] Because the right-hand-side variables are identical in all of the asset share equations, there is also nothing to be gained by estimating them jointly as a system of equations – the results will be identical (see Greene, 2000, pp. 616-617).

[40] See Greene (2000, p. 909).

[41] Poterba and Samwick (1999) do succeed in imposing the adding-up constraints on the marginal effects in their tobit estimates. Implementing this in a panel framework is an avenue for future research.

[42] See Greene (2000, pp. 911-912). Greene (1981) shows that under some restrictive assumptions (i.e., normally distributed explanatory variables), OLS estimates will be a consistent estimator of the tobit coefficient times the probability of non-limit observations.

[43] For equities, $f=[(1-d)ep+d]$, where d is the ratio of dividends to total returns, e is the portion of realized capital gains included in AGI, and p is the ratio of the present value of capital gains tax payments under deferral and exclusion from tax at death to what it would be under accrual taxation. Set d at 0.29, the share of the total returns to the S&P500 between 1980 and 1989 (assuming dividend reinvestment) that was paid out in dividends. Source: *Economic Report of the President*, Table B-95. In 1983, e is 0.4, and in 1989, e is 1. A common rule of thumb for p is 0.25 (see, for example, Feldstein, Dicks-Mireaux, and Poterba (1983), and Gordon and MacKie-Mason (1992, p. 94). This yields an f for equities of 0.361 in 1983, and 0.468 in 1989. For mutual funds, I assume $f=[(1-b)f_e+b]$, where b is the share of mutual funds held in bonds (estimated at 0.4 based on information in the 1989 SCF), and f_e is the fraction of equity mutual fund returns subject to tax, calculated in the same way as for regular equities except that p is arbitrarily doubled to 0.5 to reflect the more frequent gains realizations generated by such funds (Dickson and Shoven, 1995). This yields an f of 0.659 in 1983 and 0.787 in 1989 for mutual funds.

[44] Calculation of standard errors for Table 5 requires knowledge not only of the variances of the tax coefficients for each equation, but also of the covariances of the coefficients across the equations. These are obtained by estimating the equations jointly by three stage least squares. As mentioned above, this does not affect the parameter estimates or standard errors, because in the case where all of the right hand side variables and instruments are identical for each equation, 3SLS is identical to 2SLS. However, 3SLS does produce estimates of the covariances of the parameters across the equations.

[45] For 1983, based on the unweighted averages,
$\theta=0.454+0.048+(0.015)(0.659)+(0.182)(0.361)=0.578$.

[46] The income ranges (in 1995 dollars) used in Poterba and Samwick are: 0 - $15,000; $15,000 - $25,000; $25,000 - $50,000; $50,000 - $75,000; $75,000 - $100,000; $100,000 - $250,000; and greater than $250,000. The ranges for net worth (also in 1995 dollars) are: 0 - $50,000; $50,000 - $100,000; $100,000 - $250,000; $250,000 - $1 million; and greater than $1 million. Age ranges are less than 25; 25 – 34; 35 – 44; 45 – 54; 55 – 64; and 65 or above.

[47] See Poterba and Samwick (1999, appendix) for details.

[48] Poterba and Samwick present separate results for bonds held in retirement accounts and equities held in retirement accounts. Since no information is provided on asset distributions within retirement accounts in the 1983 SCF, this is based entirely on an imputation. I do not attempt to replicate this imputation. The "retirement accounts result" for Poterba and Samwick presented in Table 7 represents the sum of the marginal effects for these two components of retirement accounts. In addition, their mutual fund category is for equity mutual funds only; bond mutual funds are put in their taxable bond category. Again, there is no information in the 1983 SCF on whether mutual funds contain bonds or stocks, so Poterba and Samwick must again construct these variables by imputation. In my replication, mutual funds include both stocks and bonds.

[49] The marginal tax rate I use in my replication of Poterba and Samwick is calculated as follows. AGI is calculated as the sum of: wages and salaries; income from a professional practice, business, or farm; rent, trust, and royalty income; income from private pensions; and alimony received. Exclusions subtracted from AGI include alimony payments and IRA deductions (assumed to be the maximum allowable deduction, as long as this was less than the amount reported held in IRA accounts). Itemized deductions are calculated in the same way as described above for my algorithm. Interest-paid deductions are derived from answers to questions on interest payments in the SCF, all other itemized deductions are imputed based on the average percent of AGI offset by non-interest itemized deductions in IRS data for that year. Taxpayers are assumed to take the maximum of itemized deductions or the zero-bracket amount. The marginal tax rate is then calculated as the tax rate on the increment of income equal to 5 percent of financial assets.

The major features of Poterba and Samwick's algorithm that I did not attempt to replicate are as follows. First, they impute state and local property tax deductions for each individual based on the values of real estate and personal property reported in the SCF, but no information is provided on how this imputation is carried out. Second, they use information on charitable contributions provided in the SCF in 1989 and later years to estimate the charitable deduction; however, there is no information on charity in the 1983 data, and it is not clear what Poterba and Samwick do with charity in that year. Third, other itemized deductions are assumed to be zero in Poterba and Samwick's algorithm. Fourth, they impute the taxpayer's alternative minimum tax liability, which can have an effect on the marginal tax rate. The SCF is missing much of the information needed to calculate the alternative minimum tax, and some assumption would need to be made about the share of capital income that comes in the form of capital gains; it is unclear what assumptions they used to calculate the alternative minimum tax.

[50] This problem is reduced to some extent when the estimates use pooled cross-sections from a number of different years that span major changes in tax law. If the functional form of the relationship between income and asset shares remains unchanged over time, changes in the functional-form relationship between marginal tax rates and income induced by tax reform provide identification. Poterba and Samwick use this identification strategy for some of their estimates.

[51] In addition to the probit and tobit estimates, I also ran Heckman two-step equations, with either share or log share as the dependent variable. Woolridge (1995) adapts the correlated random effects approach to the Heckman two-step framework. In the cross-sectional regressions, I found the results to be similar in some cases to the cross-sectional tobit and linear models, but in many

cases to be subject to very severe multicollinearity. The multicollinearity grew worse in many of the correlated random-effects equations.

[52] Information on AGI and capital income is unavailable in 1986. A "total income" variable is available in all three years of the panel. I assume that the ratio of AGI to total income is the same for each individual in 1983 and 1986. In addition, the absence of capital income means that the instruments must be constructed differently. In place of "AGI less dividends, interest, and capital gains, plus 5% of financial assets," I use "labor income plus currently received pension benefits, plus 5% of financial assets." I find that the changes necessitated by adding 1986 do not have much of an impact on the correlation between the instrument and the marginal tax rate.

[53] The correlated random-effects framework is not adaptable to the case of both unobserved heterogeneity and state dependence. Its method for removing unobserved heterogeneity requires including values of each explanatory variable from all time periods on the right-hand side. In the case of a lagged dependent variable, this would mean also including the contemporaneous value of the dependent variable on the right-hand side, rendering estimation impossible. If there is any unobserved heterogeneity at all, the lagged dependent variable, of all variables, would certainly be correlated with it.

[54] If $s_t=\beta x_t+\gamma s_{t-1}$, then $s_t=\beta x_t+\gamma(\beta x_{t-1}+\gamma(\beta x_{t-2}+\gamma(\beta x_{t-3}\ldots)))$. The long run effect of a permanent 1 unit change in x will be $\beta(1+\gamma+\gamma^2+\gamma^3+\gamma^4+\ldots)=\beta\sum_{i=0}^{\infty} g^{i}=\beta/(1-\gamma)$.

[55] See Maddalla (1987) and Hsiao (1989) for further discussion.

[56] I also tried including the lagged change in *s* directly, and found qualitatively similar results.

[57] Constantinides (1986) works out a theoretical model of portfolio choice with transactions costs which cause slow adjustment of portfolios; however, it does not lead to empirically tractable demand equations. Sargent (1978) discusses the implicit assumptions which are necessary for the lagged adjustment cost model used here to apply to a linear regression equation.

Auerbach, Alan J. and Mervyn A. King. 1983. “Taxation, Portfolio Choice, and Debt-Equity Ratios: A General Equilibrium Model.” *Quarterly Journal of Economics*. Vol. 98, No. 4 (November): 587-609.

Avery, Robert B., and Arthur B. Kennickell. 1988. *1986 Survey of Consumer Finances: Technical Manual and Codebook*. Washington DC: Federal Reserve Board of Governors.

Bertaut, Carol C. 1998. "Stockholding Behavior of U.S. Households: Evidence from the 1983-89 Survey of Consumer Finances." *Review of Economics and Statistics*. Vol. 80, No. 2 (May): 263-275.

Chamberlain, Gary. 1984. "Panel Data" in Z. Grilliches and M.D. Intriligrator, eds. *Handbook of Econometrics* Vol. II. New York: Elseivier Science Publishers.

Chamberlain, Gary. 1982. "Multivariate Regression Models for Panel Data." *Journal of Econometrics*. Vol. 18, No. 1 (January): 5-46.

Chamberlain, Gary. 1980. "Analysis of Covariance with Qualitative Data." *Review of Economic Studies.* Vol. 47, No. 1 (January): 225-38.

Constantinides, G. M. 1986. “Capital Market Equilibrium with Transactions Costs.” *Journal of Political Economy*. Vol. 94, No. 4 (August): 842-62.

Cragg, J. 1971. “Some Statistical Models for Limited Dependent Variables with Application to the Demand for Durable Goods.” *Econometrica*. Vol. 39, No. 5 (September): 829-44.

Domar, E.D. and R.A. Musgrave. 1944. "Proportional Income Taxation and Risk-Taking." *Quarterly Journal of Economics*. Vol. 58, No. 3 (May): 388-422.

Dickson, Joel, and John Shoven. 1995. “Taxation and Mutual Funds: An Investor Perspective.” in J. Poterba, ed.. *Tax Policy and the Economy*, Vol. 9. Cambridge: MIT Press: 151-181.

Economic Report of the President. 2000. Washington, DC: Government Printing Office.

Feenberg, Daniel. 1987. "Are Tax Price Models Really Identified: The Case of Charitable Giving." *National Tax Journal*. Vol. 40, No. 4 (December): 629-633.

Feldstein, Martin. 1969. "The Effects of Taxation on Risk Taking." *Journal of Political Economy*." Vol. 77, No. 5 (Sep.- Oct.): 755-764.

Feldstein, Martin, L. Dicks-Mireaux, and James Poterba. 1983. "The Effective Tax Rate and the Pretax Rate of Return." *Journal of Public Economics*. Vol. 21, No. 2 (July):129-158.

Friedman, Benjamin M. and V. Vance Roley. 1979. "A Note on the Derivation of Linear Homogeneous Asset Demand Functions." National Bureau of Economic Research Working Paper. No. 345 (May).

Friend, Irwin, and Marshall E. Blume. 1975. "The Demand for Risky Assets." *American Economic Review*. Vol. 65, No. 5 (December): 900-922.

Gordon, Roger H., and Joel Slemrod. 1983. "A General Equilibrium Simulation Study of Subsidies to Municipal Expenditures." *Journal of Finance*. Vol. 38, No. 2 (May): 585-594.

Greene, William H. 2000. *Econometric Analysis*, 4th Edition. Saddle River, NJ: Prentice Hall.

Greene, William H. 1998. *Limdep Version 7.0 User's Manual*. Plainview, NY: Econometric Software, Inc.

Greene, William H. 1981. "On the Asymptotic Bias of the Ordinary Least Squares Estimator of the Tobit Model." *Econometrica*. Vol. 49. Issue 2 (March): 505-513.

Grossman, S. and G. Laroque. 1990. "Asset pricing and optimal portfolio choice in the presence of illiquid durable consumption goods." *Econometrica*. Vol. 58, No. 1 (January): 25-51.

Guell, Robert. 1991. *Portfolio Composition: A Theoretical and Econometric Examination of the Effect of Age and Taxes*. Doctoral Thesis, Syracuse University.

Haliassos, Michael, and Carol Bertaut. 1995. "Why Do So Few Hold Stocks?" *Economic Journal.* Vol. 105, No. 5 (September): 1110-1129.

Heckman, James. 1979. "Sample Selection Bias as a Specification Error." *Econometrica*. Vol. 47, No. 1 (January):153-61.

Heckman, James. 1976. "The Common Structure of Statistical Models of Truncation, Sample Selection, and Limited Dependent Variables and a Simple Estimator for Such Models." *Annals of Economic and Social Measurement*. 5/4. 475-492.

Hines, James. 2000. "Comment on 'Portfolio Responses to Taxation: Evidence from the End of the Rainbow.'" in Joel Slemrod, ed. *Does Atlas Shrug: The Economic Consequences of Taxing the Rich*. Cambridge: Harvard University Press.

Hsiao, Cheng. 1989. *Analysis of Panel Data*. Cambridge: Cambridge University Press.

Hubbard, Glenn. 1985. "Personal Taxation, Pension Wealth, and Portfolio Composition." *Review of Economics and Statistics*. Vol. 67, No. 1 (February): 53-60.

Hubbard, Glenn. 1983. "The Financial Impacts of Social Security: A Study of Effects on Household Wealth Accumulation and Allocation." *Monograph Series in Finance and Economics*. New York: Salomon Brothers Center for the Study of Financial Institutions.

Internal Revenue Service. Various Years. *Statistics of Income: Individual Income Tax Returns*. Washington, DC: Government Printing Office.

Ioannides, Yannis M. 1992. "Dynamics of the Composition of Household Asset Portfolios and the Life Cycle." *Applied Financial Economics*. Vol. 2, No. 3 (September): 145-159.

Jakubson, George. 1983. *Unobserved Heterogeneity in Labor Economics: Applications of Panel Data Techniques.* Dissertation, University of Wisconsin-Madison.

Jakubson, George. 1988. "The Sensitivity of Labor-Supply Parameter Estimates to Unobserved Individual Effects: Fixed Effects and Random Effects Estimates in a Nonlinear Model Using Panel Data." *Journal of Labor Economics.* Vol. 6, No. 3 (July): 302-329.

Joint Committee on Taxation. 1995. *Methodology And Issues In The Revenue Estimating Process*. Scheduled for Hearing Before the Senate Committee on Finance on January 24, 1995. Publication No. JCX-2-95

Kennickell, Arthur B. 1996. *Codebook for the 1983-89 Panel Survey of Consumer Finances*. Washington, DC: Federal Reserve Board of Governors.

Kennickell, Arthur B. and Martha Starr-McCluer. 1996. "Household Saving and Portfolio Change: Evidence from the 1983-89 SCF Panel." Federal Reserve Board, Finance and Economics Discussion Series, working paper No. 96-18.

Kennickell, Arthur B. and Louise Woodburn. 1997. "Weighting Design for the 1983-89 SCF Panel." Washington, DC: Federal Reserve Board of Governors.

King, Mervyn and Jonathan Leape. 1998. "Wealth and Portfolio Composition: Theory and Evidence." *Journal of Public Economics*. Vol. 69, No. 2 (August): 155-193.

Lee, Lung-Fei, Maddala, G.S. and Trost, R.P. 1980. "Asymptotic Covariance Matrices of Two-Stage Probit and Two-Stage Tobit Methods for Simultaneous Equations Models with Selectivity." *Econometrica.* Vol. 48, No. 2 (March): 491-503.

Maddala, G.S. 1987. "Limited Dependent Variable Models Using Panel Data." *Journal of Human Resources*. Vol. 22, No. 3 (Summer): 307-38.

Merton, Robert C., 1973. "An Intertermporal Capital Asset Pricing Model." *Econometrica.* Vol. 41, No. 5 (September): 867-87.

Merton, Robert C. 1971. "Optimum Consumption and Portfolio Rules in a Continuous-Time Model." *Journal of Economic Theory.* Vol. 3, No. 4 (December): 373-413.

Merton, Robert C. 1969. 'Lifetime Portfolio Selection under Uncertainty: The Continuous Time Case." *Review of Economics and Statistics,* Vol. 51, No. 3 (August): 247-257.

Miller, Merton. 1977. "Debt and Taxes." *Journal of Finance*. Vol. 32, No. 2 (May): 261-275.

Moffitt, Robert and Mark Wilhelm. 1998. "Taxation and the Labor Supply Decisions of the Affluent," Office of Tax Policy Research Working Paper No. 98-8, Ann Arbor.

Newey, Whitney. 1987. "Efficient Estimation of Limited Dependent Variable Models with Endogenous Explanatory Variables." *Journal of Econometrics*. Vol. 36, No. 3 (November): 231-50.

Poterba, James and Andrew Samwick. 1999. "Taxation and Portfolio Composition: U.S. Evidence From the 1980s and 1990s." National Bureau of Economic Research, Working Paper No. 7392.

Samuelson, Paul A. 1969. “Lifetime Portfolio Selection by Dynamic Stochastic Programming.” *American Economic Review.* Vol. 51, No. 3 (August): 239-46.

Samwick, Andrew. 2000. "Portfolio Responses to Taxation: Evidence from the End of the Rainbow." In Joel Slemrod, ed. *Does Atlas Shrug? The Economic Consequences of Taxing the Rich.* Cambridge: Harvard University Press.

Sandmo, Agnar. 1985. "The Effects of Taxation on Savings and Risk Taking." In A. Auerbach and M. Feldstein, eds. *Handbook of Public Economics*. New York: Elsevier Science Publishers.

Sargent, Thomas. 1978. “Estimation of Dynamic Labor Demand Schedules under Rational Expectations.” *Journal of Political Economy*. Vol 86, No. 6 (December): 1009-1044.

Scholz, John Karl. 1994. "Tax Progressivity and Household Portfolios: Descriptive Evidence from the Surveys of Consumer Finances." In Joel B. Slemrod, ed. *Tax Progressivity and Income Inequality*. Cambridge: Cambridge University Press, 219-267.

Stiglitz, Joseph. 1969. "The Effects of Income, Wealth, and Capital Gains Taxation on Risk Taking." *Quarterly Journal of Economics*. Vol. 83, Issue 2 (May): 263-283.

Vella, Francis. 1998. “Estimating Models with Sample Selction Bias: A Survey.” *Journal of Human Resources*. Vol 33, No. 1 (Winter): 127-169.

Wooldridge, Jeffrey M. 1995. “Selection Corrections for Panel Data Models Under Conditional Mean Independence Assumptions.” *Journal of Econometrics*. Vol. 68, No. 1 (July): 115-32.