

## ABSTRACT

Title of Dissertation:           ESSAYS ON THE IMPACT OF SOCIAL SECURITY  
ON THE RETIREMENT DECISION

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This dissertation investigates the impact of Social Security on the retirement decision. In the first half of the dissertation, I investigate the impact of the repeal of the Retirement Earnings Test (RET) for 65-to-69-year olds on the labor supply of older men. Under the RET, Social Security benefits are taxed if earnings exceed a stipulated amount. Using life tables from the National Vital Statistics Report, I first demonstrate that for white males, in expected value, the RET acts as a tax on beneficiary earnings. Thus, the repeal of the RET is effectively a tax cut for 65-to-69-year-old men.

Using data from the Outgoing Rotation Groups of the CPS, I estimate the effect of the repeal of the RET on the labor supply of older workers aged 55-75 years. In a life-cycle context, younger workers (55-to-64-year-olds) are expected to reduce labor supply, while workers in the age group targeted by the repeal (65-to-69-year-olds) should choose to increase labor supply.

The empirical results suggest that older workers work more because of an effective income tax cut, while younger workers reduce their labor supply. Further, the

increased labor supply among older workers stems primarily from older workers staying in the labor force longer rather than from older workers re-entering the labor force.

In the second half of the dissertation, I use data from the 1970 and 1980 Censuses and a reduced form model to examine joint retirement decisions in a household. The eligibility criteria for Social Security produce peaks in retirement at ages 62 and 65. Moreover, the retirement of one's spouse can also affect one's labor supply because of income pooling within the family or complementarities in leisure.

I find that the spouse's turning 62 and 65 affect one's own exit from the labor force. To test whether leisure complementarities are the reason for those cross effects, I use protection from age discrimination as an instrument for wages and find that, while own coverage by age discrimination laws sometimes increases own labor supply, spousal coverage has no effect. This finding suggests that the linkages in labor supply among spouses do not stem from leisure complementarities.

ESSAYS ON THE IMPACT OF SOCIAL SECURITY ON THE RETIREMENT  
DECISION

by

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To my grandmother, Tran Thi Phu.

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# CHAPTER ONE

## INTRODUCTION

The link between Social Security incentives and labor supply behavior has long been of interest to economists. The broader question of the impact of wages on labor supply is one of the oldest known questions in the field of economics. Indeed, as Pencavel (1986) notes, “conjectures about whether an increase in remuneration brought forth more work effort can be traced back at least to the mercantile economists.” Given the current Social Security debate, the link between Social Security incentives and labor supply is particularly worthy of investigation. With the solvency of the Social Security System in doubt and politicians clamoring for reform, it is important to determine how changing rules within Social Security affects the retirement decisions made by workers.

Over the past half-century, labor force participation among older men has declined dramatically. While in 1950 nearly 60% of men between the ages of 65 and 69 were in the labor force, by 1990, the number stood at only 26%. This large decrease has often been attributed to the growth of jobs covered by Social Security as well as the increasing share of household income Social Security benefits constitute among households with heads age 65 or older (Diamond and Gruber (2000)).

Another piece of evidence that suggests that Social Security has a significant effect on the labor supply decisions of older workers is the peak in the hazard rate of retirement at age 62, the earliest age at which one becomes eligible for Social Security retirement benefits. There is also a peak in retirement at age 65, but this finding is not as easily attributable to Social Security. Medicare eligibility, pension eligibility, and



mandatory retirement also become important at age 65 and are other feasible explanations. The peak in retirement at age 62, on the other hand, seems to have no other good explanation than Social Security eligibility (Hurd (1990)). Furthermore, as Burtless and Moffitt (1986) point out, there was no peak in retirement at age 62 in 1960, when men were not allowed to draw Social Security Retirement Benefits before age 65.

Despite this evidence, there is still no consensus that Social Security was the driving force behind the reduction in the labor supply of the elderly. For instance, Aaron (1982) offers a number of explanations for early retirement other than Social Security eligibility. As real income has risen over time, individuals may have simply chosen to consume more goods, including retirement. In addition, early retirement could be chosen because it might increase the present value of Social Security benefit streams, especially among workers who had an older spouse, a short life expectancy, or both. Finally, the author suggests the possibility that workers misperceive Social Security benefits because of a high discount rate.

In his survey, Hurd (1990) notes that other studies also find no impact of Social Security on retirement. Some papers attribute this result to the growth in company pension plans, which often have incentives built in to encourage workers to retire as early as age 55, thereby decreasing the importance of Social Security retirement benefits which cannot be claimed before age 62. Hurd also suggests that other papers possibly find no effect of Social Security on retirement because of the particular definition of retirement they use. He remarks that because retirement can be defined as either a permanent departure from the labor force, a self-assessed retirement, “a sudden and discontinuous

drop in the hours of work,” or merely the departure from a firm, the finding that results are contingent on the definition of retirement chosen is not surprising.

## **I. How Social Security Affects Labor Supply**

The Social Security system was established as a social insurance program to insure workers against loss of income due to old age, among other things. However, over the lifecycle, the first several cohorts of Social Security beneficiaries received more in benefits than they paid into the system. Because those cohorts were not expecting Social Security benefits, this essentially amounts to a wealth shock, as lifetime wealth has increased, possibly reducing labor supply because of this wealth effect.

Another way Social Security might impact the labor supply of beneficiaries is through program rules that directly affect the wage rate of older workers. The Retirement Earnings Test (RET) states that a Social Security old-age beneficiary may not earn more than a stipulated amount without having some of his or her Social Security benefits taxed away. This test was in place for 62-to-69-year-olds until 2000. It thus has the effect of holding down labor supply of the elderly. In a life-cycle context, the Social Security system tends to steepen the age-labor supply profile for two reasons—the price of time is relatively higher when young (because of the RET) and wealth is relatively lower when young (because of exacerbated liquidity constraints).

Further, Social Security forces workers to save more than they otherwise would. Workers must pay into the system while working, and receive a payout in the form of Social Security benefits from the system when retired. Thus, with a Social Security system in place, individuals who were already liquidity constrained have the constraint

worsen because they are forced to save more than they would optimally. Liquidity constrained persons would like to borrow from the future, but Social Security forces them to lend into the future. For liquidity constrained persons, turning 62 is like receiving a scheduled helicopter drop of money they would have preferred to receive earlier in life. Thus, at age 62, one would expect liquidity constrained people to work much less. The third chapter builds on this insight and includes the possibilities that spouses enjoy spending time together, i.e., there are complementarities in leisure, and that families pool their income, so shocks to either person's income or wealth also affect the other person's labor supply.

## **II. Repeal of the Retirement Earnings Test**

In the next chapter, I estimate the impact of the repeal of the Retirement Earnings Test on the labor supply of older people. The repeal of the RET only applied to beneficiaries aged 65 through 69. I evaluate the impact of the law using 70+-year-olds as my comparison group. I use labor supply data from the basic monthly Current Population Survey (CPS) from 1997-2002 for white men aged 55-75 who are in the fourth month of the CPS survey. I examine the impact from a lifecycle context, where workers can substitute labor supply intertemporally. In a lifecycle context, the wage rate when young has fallen in relative terms, because a tax on working in old age, the RET, has been removed. Thus, I would expect the age-labor supply profile to flatten, as younger workers work less, and older workers, namely the 65-to-69-year-olds, work more in response to the repeal. Using both hours worked last week and employment last week as my dependent variables, my results confirm that 65-to-69-year-olds work more,

while the young cohort works less in response to the repeal. I further find that the increase in labor supply by 65-to-69-year-olds is driven by an increase in full-time workers, while the share of part-time workers stays roughly constant. This result suggests that the increase was driven by the repeal of the RET. The RET does not discourage part-time work as much as full-time work as all workers could have earnings up to the same exempt amount. Furthermore, the share of 65-to-69-year-old workers who receive Social Security benefits increased after the repeal, suggesting that those 65-to-69-year-olds are not merely working longer in response to higher expected longevity or other reasons, but because of the repeal.

### **III. Joint Retirement**

In the second paper, I investigate the joint retirement decisions of married couples. While it is well known that one's own age affects labor force participation discontinuously at ages 62 and 65, little is known about a wife's labor supply when her husband reaches those ages. When an older spouse, often the husband, becomes eligible for Social Security retirement benefits at age 62, liquidity constraints within the household are relaxed, and household wealth is increased. If families pool income, either spouse turning 62 should relax liquidity constraints and affect the labor supply decisions of both spouses. I include dummy variables for either spouse turning 62 and 65 and both spouses turning 62 and 65 in the equations for husband's and wife's labor supply in a probit model to account for this effect. I find that there are not only the expected own-age effects for turning 62 and 65, but also cross-age effects. Those cross-age effects point to either complementarities in leisure or liquidity constraints as the reason for joint

retirement. To test whether cross-age effects come from complementarities in leisure, I instrument for wages by including two dummy variables, one for whether the person is covered by age discrimination legislation, and one for whether his or her spouse is covered. If the cross-age effects are driven by complementarities in leisure, then if age discrimination laws raise wages and thus the labor supply of older men, they should also raise the labor supply of their wives, whether or not the latter were covered (and vice versa). I find that while age discrimination laws raise labor supply among older workers (as Neumark and Stock (1999) found) a husband's coverage does not affect his wife's labor supply (and vice versa). This result suggests that complementarities in leisure are relatively modest and that the relaxing of liquidity constraints for either spouse affects the labor supply of both spouses as most models of family decision making suggest.

## CHAPTER TWO

### THE IMPACT OF THE REPEAL OF THE RETIREMENT EARNINGS TEST ON THE LABOR SUPPLY OF OLDER WORKERS

#### I. Introduction

Until 2000, all Social Security beneficiaries under the age of 70 were covered by the Retirement Earnings Test (RET). This less well-known provision of the Social Security Law stipulates that if an old-age beneficiary's earnings exceed a certain exempt amount, his benefits were reduced. The RET was included in the Social Security bill because the Social Security system was designed to provide economic security for the disabled, the unemployed, and the elderly. It was thought that only those who suffered an economic loss in income due to old age were to receive retirement benefits. Over time the RET has become less stringent. In 1983, the RET was repealed for those aged 70 and above, and in 2000, the RET was repealed for all beneficiaries who had reached at least the Normal Retirement Age (NRA), which was 65 at the time.

Some might argue that there should be little impact on the labor supply of the elderly in response to the repeal because of the existence of the Delayed Retirement Credit (DRC). The DRC increases the benefit amount if the worker delays claiming Social Security retirement benefits past the NRA. A delay in claiming can be either voluntary or involuntary. In the voluntary case, the worker does not start claiming until he is well past the NRA and receives credits for the months that he has waited. In the involuntary case, the worker has some of his Social Security benefits taxed away due to the RET, and he receives credits for the months of Social Security benefits lost. The

adjustment in benefits for delaying claiming Social Security started at one percent per year in 1972 and eventually will reach eight percent per year for those who turn 65 in 2008. The DRC was introduced as a measure of fairness to those who had to delay receipt of benefits because of the RET. Depending on the individual's life expectancy, the DRC might make the RET actuarially fair—beneficiaries would have to delay receipt of benefits but in return receive a higher benefit amount in the future.

However, individuals might still perceive the RET as a tax for a number of reasons. First, the RET is a tax in a cross-sectional sense. Furthermore, there is uncertainty over the time of death, so risk-averse individuals might prefer earlier receipt of a smaller benefit amount to later receipt of a larger benefit amount even if the values of the stream of benefits are equal in expected terms. Moreover, liquidity constraints that would lead the individual to prefer benefit receipt now to later also make the RET a tax. In addition, the adjustment might not be actuarially fair for the average worker. Finally, because of differences in life expectancies, even if the RET is actuarially fair for some workers, it might still not be fair for all of them.

Because it is unclear whether the RET should be viewed as a tax, it is not surprising that the literature on the impact of the Retirement Earnings Test on the labor supply of the elderly has produced mixed results. In his survey, Leonesio (1990) notes “virtually all of [the] research [on the RET] indicates that the effect is probably small and that eliminating the test would have a minor impact on the work activity of older Americans.” However, in a more recent paper, Friedberg (2000) claims that this lack of an impact found should be attributed to a lack of innovation in the parameters of the RET

over the time period of earlier studies. Friedberg finds that removing the RET would significantly increase labor supply of the affected group.

The repeal of the RET in 2000 is the “natural experiment” I exploit to analyze the impact of the Retirement Earnings Test on labor supply. I estimate changes in the employment rate and hours worked using a differences-in-differences model to see how the elderly responded to this policy change. This repeal is also a nice experiment that allows me to newly examine the impact Social Security rules have on the labor supply of the elderly.

The primary data for this paper is the Outgoing Rotation Groups (ORG) data on white men aged 55-75 in the fourth month of the Current Population Survey (CPS), from January 1996-December 2002. I find that hours increase for the group targeted by the repeal, 65-to-69-year-olds, and decrease for younger workers. Further, there is weaker evidence that the employment rate of the target group rises in response to the repeal, while the employment rate of younger workers decreases. The increase in labor supply among the target group seems to be driven mostly by increases in labor supply among 65- and 66-year-olds. In the presence of labor market transition costs, this finding suggests that the effect of the repeal on the target group might increase over time as those 65- and 66-year-olds age.

## **II. Brief History of the Retirement Earnings Test**

Individuals can start claiming Social Security old-age benefits after they reach age 62. Benefit amounts depend on the worker’s earnings over his lifetime and his age at retirement. The earlier a worker claims, the lower the benefit amount. In addition to this



adjustment, the RET, a provision of the Social Security Law, stipulated that no benefits were to be paid to an old-age beneficiary who had “regular employment.” Because the Social Security System was intended to be a social insurance program, retirement benefits were only to be received by those who had suffered a loss of income due to old age. In 1939, Congress established earnings of \$15 or more in a month as the amount that would constitute “regular employment.” By 1950, the RET had come under increasing criticism. Several problems were noted: The “all-or-none” nature of the RET was deemed excessive. A more gradual reduction in benefits was seen as more appropriate for workers who wanted to progress from full-time work to full retirement by way of part-time work. Furthermore, some workers never retired and thus would never receive benefits, which seemed unfair, so there were efforts to make Social Security retirement benefits an annuity, payable upon attaining a particular age.

These criticisms led to the gradual relaxation in the RET over time. First, the amount of income that beneficiaries were allowed to earn without having their benefits reduced has been increasing over time. Second, starting in 1954, beneficiaries aged 75 and above were no longer subject to the RET. Since then, increasingly younger beneficiaries have been exempted from the RET. Third, since 1960, the rate at which benefits were reduced for earnings above the exempt amount has been repeatedly lowered (Dewitt (1999)).

In 1999, the year before the RET was repealed for beneficiaries who had reached at least the NRA, 62-to-64 were allowed annual earnings of up to \$9,600. For labor income above that amount, their Social Security benefits were reduced by \$1 for each \$2 in excess of that amount, effectively resulting in a 50 percent marginal tax rate.

Beneficiaries 65 to 69 years old were allowed annual earnings of up to \$15,500. Income in excess of that amount resulted in a reduction of benefits by \$1 for every \$3 above the exempt amount, or a 33 percent marginal tax rate. In 2000, the RET was repealed for the NRA cohort—those who were at least 65 years of age—abruptly dropping the marginal tax rate on excess income from 33 percent to zero for a large share of that cohort.

A counterpart to the RET, the Delayed Retirement Credit (DRC), was introduced in 1972. The DRC increases a worker's benefit amount for each month that he delays receipt of benefits past the NRA, which was 65 at the time. When the DRC was first introduced, the benefit amount would increase by one percent for each year that the worker delayed receipt of Social Security benefits. In 1977, the DRC was increased to three percent per year. In 1983, Congress passed a law that set scheduled increases in the DRC, starting in 1990, and eventually increasing the DRC to eight percent per year.<sup>1</sup>

### **III. Literature on the Effect of the Retirement Earnings Test on the Labor**

#### **Supply of Older Workers**

Until recently, the general consensus in academic research had been that the RET has little effect on labor supply. Burtless and Moffitt (1985) developed a life-cycle model to jointly estimate the optimal retirement age and the hours of work immediately after retirement in the presence of the Social Security system and the earnings test. They define retirement as a discontinuous drop in labor supply. Estimating a life-cycle model via maximum likelihood estimator with panel data on men from the Retirement History Survey (RHS), they find the impact of Social Security benefits on retirement probabilities

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<sup>1</sup>Information from Social Security Administration Website. See Figures 2 and 3 for changes in the RET exempt amount and the DRC.

grows with age. First, examining the means, they find that retirement age clusters at age 65, as expected. Second, they also find that retirement working hours as a fraction of the hours worked needed to earn exactly the exempt amount clusters around 1, which is also expected. Their identification in their model comes from variation in wages and assets. In their model, they incorporate three kinks—one from the point on where the federal income tax applies, and the other two arising from the retirement earnings test (one at the exempt amount and one at the point where all Social Security benefits have been “taxed” away due to excess earnings). They do not account for changes in the marginal tax rate of the federal income tax, arguing that that effect is dominated by the “tax” on Social Security benefits that applies to excess earnings, rather holding it constant. Their simulations show that eliminating the earnings test would not affect hours of work much and would leave the optimal retirement age virtually unchanged. The authors also find that lowering benefits and increasing the normal retirement age would result in an increase in retirement age and hours of work. While the model incorporates the life-cycle aspect into it, jointly estimating age at retirement and post-retirement labor supply, they are estimating off limited variation in data that is quite dated.

Leonesio (1990) argues that there are a number of factors that mitigate the impact of the RET on labor supply. First, the DRC and benefit re-computation<sup>2</sup> offset much of the loss due to the earnings test. Second, the law has been relaxed over the years, so the elderly can earn substantial amounts before being subjected to benefit reductions. Third, the group of affected persons is small enough to not make an impact on the aggregate

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<sup>2</sup> Benefits are based on your Average Indexed Monthly Earnings. The AIME takes into account the 35 highest-earning years, so if your earnings later in life are greater than early in life, you might replace some low-earning years with higher ones, thus increasing your Social Security benefits.

labor supply. Fourth, many workers do not control the number of hours worked on the job, and thus reaction to changes in the law might be small in the short run. Leonesio suggests that other factors, such as private pensions, Social Security benefit levels, health, job opportunities, family circumstances, and personal preferences are more important in determining the date of retirement than the RET. His survey is unconvincing because he mostly appeals to rigidities that might hold down the effect in the short term, but he does not consider possible long term effects.

Honig and Reimers (1993) estimate the impact of the earnings test on labor supply by examining the re-entry behavior of older men who have retired from their career job. The authors contend that if workers can choose their hours freely, the earnings test should not impact the labor force participation rate. If workers earn more than the exempt amount, they can simply cut back on their hours to reduce their earnings to below the exempt amount. One would expect to see a reduction in the number of hours worked as workers perceive the RET as a tax, but there is no reason to expect that workers will drop out of the labor market entirely. However, in estimating a hazard model of re-entry into the labor force after retirement, they find that “older men perceive a discontinuity in the budget constraint” driven by fixed costs to working. Moreover, older workers act myopically with regard to the RET. Identification comes from variation in wages and other income and wealth. They do not account for the DRC and thus act as if there were a kink at the exempt income amount. As the authors note, myopia alone should not affect labor force participation, merely the number of work hours. Thus, their results imply myopia, a lack of good part-time jobs, and large fixed costs to labor market re-entry. Furthermore, using simulations, they find that the

scheduled increases in the DRC that passed as part of the Social Security Amendments of 1983 would not affect older men's retirement behavior. Instead, increasing the exempt amount would more effectively accomplish the goal of stimulating labor supply among the elderly than increasing the DRC. While the results are quite convincing, the study suffers from the lack of scope—the design allows the authors to examine re-entry behavior since retirement from a career job only, not the labor supply behavior of men still on their career job.

Friedberg (2000) examined whether changes in the RET exempt amount over time altered labor supply of the affected age group. She finds that earnings of older workers bunch just below the exempt amount for their respective age group, suggesting that workers are aware of the exempt amount and that they try to stay below that threshold. Furthermore, she shows that the clustering spot moves when the exempt amount changes. This evidence suggests that the RET reduces labor supply. She also estimates that eliminating the RET at age 65 would increase the average hours worked by 5.3 percent for beneficiaries in the 65-to-69 age group who are at or above the kink in the budget constraint. Her identification strategy involves exploiting the variation in wages and assets to estimate the impact on hours worked, using a model with a kinked budget constraint, where the only tax is the reduction in Social Security benefits due to excess earnings. Because she relies on the variation from wages, she can only estimate the effect on those already at work. Therefore, her model does not account for beneficiaries who might now enter the labor force. She argues that previous studies on the effect of the RET are outdated and rely on data from the 1970's, a period with little change in the RET. Using recent CPS data, she finds that the labor supply of older workers is sensitive

to the RET. It should be noted that she treats the RET as a mere reduction in benefits, ignoring the actuarial adjustment in future benefits via the DRC.

Gruber and Orszag (2000) use data from the March Supplement CPS for 1974-1999 and find that there are small labor supply responses among males to changes in the RET exempt amount, the tax on benefits, and the DRC. They estimate a simple reduced form model and find that when linear and quadratic age-specific time-trends are included as independent variables, changes in the RET law such as increasing exempt amounts have no effect on labor supply. However, they do find that changes in the law have affected the labor supply of women. Their findings also indicate that removing the RET might increase earlier claiming of Social Security benefits among both men and women. Their study unfortunately suffers from a lack of variation in the data necessary to find any responses in a cross-sectional model. The result among women might be attributable to a general upward trend in female labor supply.

Pingle (2002) uses the 1985 to 1996 panels of the Survey of Income and Program Participation to examine the impact of raising the DRC over this time period. He finds that the changes in the DRC do affect labor supply, although those changes seem restricted to individuals near the age of 65. Using the actual DRC amount to identify its effect on the labor supply of those around 65 years of age, he finds a significant and large effect, using a linear trend for birth year cohort. However, the result is sensitive to the specification of the trend term.

What is missing in the literature is a study of what happened when the RET was repealed for a large portion of Social Security beneficiaries. This chapter attempts to fill that gap. One major weakness previous studies suffered from was the lack of variation in

RET rules that they could employ in their analysis. In addition, the studies also suffered from selection bias in that they could only observe a limited sample (for example, those that re-entered the labor force or those that were working even before the repeal). In addition, they all had to rely on out-of-sample predictions in their analysis because the repeal had not taken place yet. The repeal of the RET in 2000 is a “natural experiment” that allows me to re-examine the issue.

#### **IV. The Retirement Earnings Test as a Tax in the Life-Cycle Context**

Whether the repeal of the RET has an impact on labor supply depends on whether or not the RET is a tax on earnings. In this section, I aim to demonstrate the manner in which the RET operates as a tax because I wish to treat the repeal of the RET as an income tax cut. In a cross-sectional sense, the RET is obviously a tax—workers lose benefits they would otherwise receive because they earn more than the exempt amount of income. However, it is more appropriate to examine the RET from a life-cycle perspective. In 1999, if the earnings of a Social Security beneficiary who had reached the NRA of 65 were above a specified exempt amount of earnings, his Social Security benefits were “taxed” at a 33 percent rate.<sup>3</sup> However, future benefits are adjusted via the Delayed Retirement Credit. If the future adjustment is actuarially fair and there are no liquidity constraints or risk aversion by individuals, the RET is not a tax in a life-cycle context and it should have no effect on the labor supply of 65-to-69-year-olds. On the

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<sup>3</sup> The worker can be on any of three segments of his budget constraint before the repeal. He can earn below the exempt amount and thus receive full Social Security benefits, and face a zero marginal tax rate due to the RET. Further, he can earn more than the exempt amount, but receive partial Social Security benefits, and thus face a 33 percent marginal tax rate due to the RET. Finally, he can earn so much in excess of the exempt amount that all of his Social Security benefits are taxed away by the RET, and thus again face a zero marginal tax rate due to the RET.

other hand, if the actuarial adjustment is not fair, and thus the expected value of the lifetime stream of Social Security benefits decreases, the RET always acts as a tax for some individuals.

With a known mortality date, one can calculate the value of the stream of Social Security benefits a worker will accumulate over his lifetime, conditional on when he starts claiming Social Security in the absence of the RET. Thus, one can determine when a worker should start claiming so as to maximize the stream of Social Security benefits. I will conduct just this exercise assuming zero inflation<sup>4</sup> and no changes in the Average Indexed Monthly Earnings (AIME)<sup>5</sup> upon which the amount of Social Security benefits is based. Furthermore, I consider only the worker's own Social Security benefits, not the secondary benefits a spouse might draw. Additionally, I make the simplifying assumption that workers start claiming at either age 62, 65, or 70 only. I choose these three ages because age 62 is the earliest age at which one can start claiming benefits, 65 is the Normal Retirement Age now, and it is not rational to start claiming after age 70 because one cannot accrue Delayed Retirement Credits past age 70. Note that the results I find are irrespective of the earnings level of the beneficiary under the above assumptions. I find that for a man who knows that he will die before age 77, claiming at age 62 maximizes his lifetime Social Security benefits. Moreover, assuming the DRC is five percent, only a worker who will die at age 90 or older would want to delay claiming benefits until age 70 in order to maximize the value of his stream of Social Security

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<sup>4</sup> Although that is accounted for by the Cost Of Living Adjustments (COLA).

<sup>5</sup> Changes are minimal. A case in which a worker replaces two years of average pay (\$2,683 in 2002) with high pay (\$4,145 in 2002) results in the monthly Primary Insurance Amount (PIA) increasing from \$1,559 to \$1,570. The monthly PIA is the amount of monthly benefits received by a beneficiary who retired at the Normal Retirement Age.



benefits. Thus, for anyone who knows that he will die before age 90 and who earns an income between the ages of 65 and 69 that substantially exceeds the exempt amount, the RET acts as a tax because the adjustment in future benefits is not actuarially fair.

Having shown that for an individual with a low life expectancy the RET is a tax in a life-cycle context with no borrowing constraints and with certainty about the time of death, it follows that the RET is always a tax if the individual dies at a sufficiently young age. Borrowing constraints, myopic viewpoints, and risk aversion exacerbate the effect of the tax further. Moreover, those three factors can lead individuals to perceive and experience the RET as a tax even if the adjustment via the DRC is actuarially fair.

Therefore, the repeal of the RET is an income tax cut in a life-cycle context even for risk-neutral individuals who face no liquidity constraints if they die at a sufficiently young age.<sup>6</sup> Beneficiaries aged 65-69 who work at the moment and who are subject to liquidity constraints or for whom the adjustment via the DRC is not actuarially fair perceive this repeal as an income tax cut. I would expect this tax cut to spur those aged 65-69 to work more because I believe the substitution effect dominates the income effect, especially at lower levels of income. The income effect should be relatively small at lower earnings levels. In fact, for beneficiaries aged 65-69 who do not work at all, the income effect is zero at the margin because if a worker earned less than the exempt amount prior to the repeal, the RET did not effectively tax him, so there is no income

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<sup>6</sup> Note that those who are most affected are those receiving partial Social Security benefits. Those who have all their benefits taxed away face essentially a lump-sum tax, so they experience only an income effect. Those who earn below the exempt amount should not be affected, as their marginal tax rate has not changed, and they do not experience an income effect, either. However, in the presence of discontinuities in the budget set, the latter might experience a substitution effect.

effect. Therefore, I would expect individuals aged 65-69 to increase labor supply in response to the repeal.

#### A. When to Start Claiming Social Security Benefits

With perfect information and no liquidity constraints, the optimal age to start claiming Social Security benefits depends on one's time of death (TD). In the following illustration, I make the simplifying assumption that workers only claim at ages 62, 65, and 70. Also, only retired workers' own benefits are considered, not spouse's or other secondary benefits that might factor into the decision on when to start claiming benefits. Assuming there is no discounting and that there are no changes in the AIME amount due to the replacing of lower-earnings years with higher ones, the answer comes down to the following:  $80*(TD-62) > 100*(TD-65)$ ?<sup>7</sup>

If the inequality holds, claiming early (at age 62) is better than waiting until one reaches the NRA. This statement is true if  $TD < 77$ . Factoring in discount rates moves the maximum TD up. However, if there are no liquidity constraints, this factor should not matter given COLA. Factoring in a higher AIME due to benefit re-computation moves the maximum TD down, but the AIME should not change much for most workers. Next, considering the question of whether a worker should start claiming at age 65 or age 70 is the same as asking oneself if the following inequality is true, assuming the DRC is five percent:  $100*(TD-65) > 125*(TD-70)$ ?

If this inequality holds, claiming at age 65 is better than claiming at the latest possible (while still rational) date, age 70, as workers may not accrue additional Delayed

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<sup>7</sup> Receive 80% of the PIA for (TD-62) years if you claim at age 62, or receive 100% of PIA for (TD-65) years

Retirement Credits after age 70. This inequality holds if  $TD < 90$ . Thus, for all workers whose TD is less than 90 years (at a DRC of five percent per year), the RET acts as a tax, as benefits are only partially replaced.<sup>8</sup> I show what the expected value of the stream of SS benefits is, not accounting for discounting, in a simple numerical example, where a worker's annual PIA is 100, conditional on his time of death in Table 3.1. Clearly the RET acts as a tax for many workers and it should not be surprising that the repeal of the RET should have an effect on the labor supply of the affected age group (see Figure 1 for a graphical illustration of this exercise).<sup>9</sup>

#### B. What About Spouse's Benefits?

Note that the DRC does not increase spouse's benefits. If a worker in my sample of white men is married and his wife receives spouse's benefits, the RET is more likely to act as a tax for him in a family decision-making context than if he were single. Assuming the primary earner is a man who retired at the NRA or later and his wife receives spouse's benefits, the wife receives 50 percent of the Primary Insurance Amount (PIA)<sup>10</sup> if she starts claiming at her NRA or later. Her benefit amount is adjusted downward if she starts claiming earlier. If the NRA is 65 and she starts claiming at age 62, she receives 75 percent of 50 percent, 37.5 percent, of her husband's PIA, assuming he retired after

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<sup>8</sup> Indeed, even at 6.5 percent per year, assuming a discount rate of zero and adjustments in the AIME,  $TD > 85(85.4)$ , to be precise) for workers to not perceive the RET as a tax.

<sup>9</sup> Note that although the DRC eventually reaches 8 percent, I cannot repeat these simple exercises as easily because by the time the DRC is 8 percent, the NRA will have increased to 67+ years.

<sup>10</sup> The PIA is the benefit amount that, in this case, a retired worker receives monthly if he retires at the Normal Retirement Age (NRA). The annual PIA refers to the benefit amount a retired worker receives annually if he retires at the NRA.

attaining the NRA. On the other hand, if the husband delays retirement until age 70 and the wife does not claim until she has attained at least her NRA, she still receives only 50% of his PIA, not 50% of his benefit amount as adjusted via the DRC. However, if the husband dies, the widow receives his full benefit amount, including the adjustment via the DRC.

For example, consider a 65-year-old married man whose earnings in 1999 are \$42,500, \$27,000 over the exempt amount of \$15,500, and whose 65-year-old wife draws spouse's benefits, and they both started claiming after attaining the NRA. Then, the family's Social Security benefits are reduced by a total of \$9,000, one-third of \$27,000. His benefits are reduced by \$6,000 and his wife's are reduced by \$3,000. Assuming his annual PIA is \$12,000 and he retired at the NRA, he is considered to have lost six months' worth of benefits (\$6,000 is 6/12 of his annual PIA, \$12,000). Thus, he receives a credit of  $5.5*(6/12) = 2.75$  percent (with the DRC at 5.5 percent per year) for losing six months' worth of Social Security benefits. However, the loss in his wife's spouse's benefits is not replaced. She continues to receive spouse's benefits of 50 percent of his PIA (\$12,000), \$6,000, rather than 50 percent of what her husband now receives,  $1.0275*PIA$  (12,330), \$6,165.

#### C. How Much Money was Taxed Away by the RET Prior to its Repeal?

Here I calculate the expected value of the stream of Social Security benefits for a typical white male at age 62 conditional on when he first claims, using life tables from 1999 (see Table 3.2.). To simplify the calculation, I assume that workers claim only at exact ages in years, e.g. they may claim at exactly 62 years of age but not at 62 years and

3 months, and that full annual benefits are received at the end of each year. I find that an average white male maximizes the value by first claiming benefits at age 65 with a DRC of five percent per year. For the average white male who would want to start claiming at age 65, being forced to delay receipt of any Social Security benefits until age 70 is equivalent to having about 12 percent of his lifetime Social Security benefits taxed away (see Table 3.1.). This finding suggests that for the average white male, the RET is a tax. The actual tax rate depends on his lifetime earnings profile and life expectancy. Note that this statement does not imply that the tax affects the average white male, as he might or might not choose to work past age 65, even in the absence of the RET.

To put a dollar figure on this loss, note that the mean annual PIA in the middle quintile of PIA's was \$11,700 in 2000. The annual PIA is the amount of benefits a retired worker receives annually if he retires at the Normal Retirement Age, 65 at the time. Thus, for the average white male, assuming he had about average earnings<sup>11</sup> throughout his life, the difference in being able to start claiming and receiving benefits at age 65 and being forced to defer receipt of benefits until age 70 was \$21,539.70—he received \$152,404.20 instead of the \$173,943.90 he would have received if he had been allowed to receive his full benefits from age 65 on.

## **V. Expected Economic Consequences of the Repeal of the RET**

Absent liquidity constraints and risk aversion and if there is no Retirement Earnings Test in place, a worker will start claiming Social Security benefits at the age where the expected present value of the stream of Social Security benefits is maximized. The labor supply decision will be made independently from the decision about when to

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<sup>11</sup> Here, “average” refers to the average beneficiary, not the average white male.

start claiming. Without the RET, there is no limit on earnings even when benefits are claimed, so there is no link between the labor supply decision and the decision about when to start claiming benefits.<sup>12</sup>

With a RET in place, however, one would expect two things to happen: a reduction in labor supply at the ages where the RET applies, and an increase in labor supply at other ages,<sup>13</sup> assuming the RET acts as a tax. The intuition behind this prediction is that because of the tax on working from age 62 to age 69, there is a substitution effect that leads to people working more when they are paid relatively more. There is also an income effect that induces individuals to work more over their lifetime because lifetime wealth has decreased. The size of this effect depends on the effective tax rate of the RET.

After the repeal of the RET, labor supply at the ages for which it has been repealed should increase, as relative wages have risen for those groups (substitution effect). The income effect would tend to reduce lifetime labor supply, but it would not necessarily result in exit from the labor force, but rather should result in decreased intensity, e.g., reduced hours.

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<sup>12</sup> I abstract from the income effect from Social Security benefits that might arise because of liquidity constraints. With liquidity constraints, it could be rational to start claiming at an age that does not maximize the present value of the stream of Social Security benefits.

<sup>13</sup> Absent labor force transition costs. With labor market entry/exit costs, one would expect labor supply to drop after the ages for which the RET applies as well. If the affected age range for the RET is 62-69, then, in the absence of labor market entry/exit costs, one would expect labor supply to rise before age 62 and after age 69. However, with labor force transition costs present, it is possible that one finds that labor supply after age 69 also declines because workers who exit the labor market prior to age 69 find it too costly to re-enter the labor market.

The RET did not apply to a worker who is 70 or older even prior to 2000, so the repeal of the RET has no effect on workers in that age group if it was unanticipated. Thus, workers aged 70 and older would be a suitable comparison group in the time period immediately after the repeal, as long as the repeal was unexpected. Any change in their labor supply is not a result of the repeal of the RET.

A worker between the ages of 65 and 69 experiences an income effect if he had some of his benefits taxed away prior to the repeal because this tax cut increases his lifetime wealth, which in return depresses lifetime labor supply. However, at the same time, the reduction in the marginal tax rate from 33 percent to zero for some workers also means that there is a substitution effect working in the opposite direction. The return to working has increased, which spurs labor supply. The likely net result is an increase in labor supply because the income effect is likely to be minimal unless the worker had substantial earnings before the repeal. If the worker was not working at all or had labor income below the exempt amount, there is no income effect because he was not taxed prior to the repeal.

For a worker younger than age 65, income and substitution effects work in the same direction. The income effect that arises if the worker was going to earn more than the exempt amount between the ages of 65 and 69 results in lower lifetime labor supply. Further, working between the ages of 65 and 69 becomes relatively more attractive than before the repeal because compensation at those ages has increased in relative terms, so there is a substitution effect that induces a reduction in labor supply before age 65.

One way to examine the effect on labor supply is to examine the change in the employment rate. One would expect more 65-to-69-year-olds at work and fewer younger

workers at work in any given week, while the old group, aged 70 and older, should be unaffected by the repeal and thus make a good comparison group. Additionally, hours should decrease for younger workers (ages 55-64), increase for the group that is affected (ages 65-69), and not change as a result of the repeal for the oldest workers (ages 70+).

Labor market entry/exit costs might also affect the behavior exhibited by workers. If entry costs into the labor market are high, once workers retire, many do not re-enter the labor market, even after the repeal. Instead, the increase in the employment rate of 65-to-69-year-olds would come about because workers stay in the labor force longer. One would observe an increase in the rate at which workers stay in the labor force, but one would not necessarily find an increase in the rate of labor market entry.

## **VI. Econometric Model**

As I argued in the previous section, different groups will react in different ways to the repeal of the RET. First, those aged 64 and younger are expected to reduce labor supply in response to the repeal. I split that group into those aged 61 and younger, the young group, and those aged 62 to 64, the intermediate group. These two groups differ in that 62-to-64-year-olds may claim Social Security retirement benefits subject to the RET while those younger than age 62 may not. Thus, the effects for those two groups might differ. An additional reason why those two groups should be treated separately is that when the RET was repealed for those who had attained the NRA in 2000, which was 65 at the time, there were two other major changes for Social Security beneficiaries who had not yet attained the NRA but would reach the NRA in a given calendar year: the exempt amount was substantially increased and the marginal tax rate on earnings in excess of the



exempt amount was lowered from 50 percent to about 33 percent. Thus, it is not clear whether those aged 62 to 64 as a whole increase or decrease labor supply. Those aged 65 to 69, the target group, are expected to increase labor supply in response to the repeal (see Table 3.2. for an illustration of income and substitution effects for each of the three groups).

The data I have for my model are individual level data on age, educational attainment, race, gender, hours worked last week, labor force status last week, and date of survey. I opt to estimate a differences-in-differences model with the old group, those aged 70 and older, as my comparison group. The old group does not experience a change in response to the repeal because they were already exempt from the RET prior to 2000.

The equation I estimate is the following:

$$\begin{aligned}
 Y_{it} = & \phi_0 + (\text{Target Group})_{it} \phi_{\text{target}} + (\text{Young Group})_{it} \phi_{\text{young}} & (1) \\
 & + (\text{Intermediate Group})_{it} \phi_{\text{intermediate}} + \sum_{j=31 \text{ to } 46} \text{Edu}_{i,j} \phi_{\text{edu},j} \\
 & + \sum_{k=55 \text{ to } 75} \text{Age}_{i,k} \phi_{\text{age},k} + \sum_{l=1997 \text{ to } 2002} \text{Year}_{t,l} \phi_{\text{year},l} \\
 & + \sum_{m=1 \text{ to } 12} \text{Month}_{t,m} \phi_{\text{month},m} + \varepsilon_{it}
 \end{aligned}$$

$Y_{it}$  is a measure of labor supply for individual  $i$  at time  $t$ . I use both an indicator variable for whether an individual worked last week and hours worked last week as a labor supply measure. The treatment variables are  $(\text{Target Group})_{it}$ , which equals 1 if the respondent is between 65 and 69 years of age and the year is 2000 or later, 0 otherwise,  $(\text{Young Group})_{it}$ , which equals 1 if the respondent is younger than or equal to 61 years of age and the year is 2000 or later, 0 otherwise, and  $(\text{Intermediate Group})_{it}$ , which equals 1 if the respondent is between 62 and 64 years of age and the year is 2000 or later, 0 otherwise.

$\sum_j \text{Edu}_{i,j}$  is a set of dummy variables for educational attainment.  $\sum_k \text{Age}_{i,k}$  is a set of

dummy variables for age in years,  $\sum_l \text{Year}_{t,l}$  is a set of dummy variables for year of survey, and  $\sum_m \text{Month}_{t,m}$  is a set of dummy variables for month of survey.  $\varepsilon_{it}$  is assumed a normally distributed error term.

As shown earlier, workers with a short life expectancy are most affected by the RET and should modify their labor supply the most in response to the repeal of the RET. Unfortunately, I do not observe a person's life expectancy in the CPS data, so I cannot control for life expectancy directly. However, Kitagawa and Hauser (1968) find that there is a strong inverse relationship between educational attainment and mortality among white males. More recently, Christenson and Johnson (1995) confirmed this result using actual data from death certificates after education was added to the U.S. death certificate.

Further, Hurd and McGarry (1995) find that not only do low-educated individuals die younger, they also expect to die younger (Table 3.5.). Using data from the Health and Retirement Study (HRS), the authors find that individuals' subjective life expectancies vary according to various risk factors in the same direction macro data predicts those risk factors to act. Hurd and McGarry's findings include low-educated respondents' knowing that their mortality rate is higher than that of higher-educated individuals.

Thus, because low-educated individuals know that they die younger, and the RET taxes those who die young the most, low-educated workers should react most strongly to the repeal of the RET. One might argue that low-educated workers might not understand that their lower life expectancy makes the RET more of a tax for them than for higher-educated workers. However, low-educated workers are also more likely to perceive the RET as a tax if one makes the assumption that they are also less informed about Social Security law, and might, in fact, not know about the DRC. In addition, low-educated

workers are more likely to be subject to liquidity constraints and thus perceive the RET as a tax.

Therefore, I also run equation (1) with subsamples where I take workers from one of two groups: those who have a high school diploma or less, and those who have a college degree or more. The former is expected to change labor supply more drastically than the latter. I again use employment and hours as my measures of labor supply.

Finally, I would like to determine whether the change in the employment rate came about because of change in entry into the labor force, change in exit from the labor force, or change in continuation in the labor force. Entry indicates that a person was not at work last year but is at work now. Exit indicates that a person was at work last year but is not at work now. Continuation indicates that a person was at work last year and is still at work. For this last set of regressions, I use entry, exit, and continuation as my measures of labor supply in equation (1).

The parameters I estimate via my differences-in-differences model are the initial responses to the repeal of the RET. Note that as I add more data in a few years, this same analysis could not be done because 70+-year-olds would no longer be a suitable comparison group. Those 70+-year-olds would have been affected by the repeal earlier when they were still 65 years old, so 70+-year-olds will not be a suitable comparison group in the future any more.

#### A. Notes on the Comparison Group

There is some concern over whether 70+-year-olds are a suitable comparison group. In fact, 70+-year-olds probably behave very differently from the young group. A

differences-in-differences model has several drawbacks: Among the young, it might understate the effect, absent general equilibrium effects. For example, one might believe that the change in the labor market environment captured by the comparison group of 70+-year-olds understates the change among the 55-to-61-year-olds because the latter are more flexible and can react more strongly to positive changes in the labor market environment. If the labor market situation among the elderly improves, as it did as evidenced by the increase in labor force participation among 70+-year-olds, the young might benefit more than 70+-year-olds. In that case, the effect captured by the model is too small, and a lower bound on the real effect. However, if there are general equilibrium effects at work here, then increased labor supply by 70+-year-olds could lower demand for labor by 55-to-61-year-olds even with no repeal of the RET. General equilibrium effects are probably small, however, considering that few 70+-year-olds work. Overall, if I find that the young decrease their labor supply, the result is probably a lower bound on the true effect.

Further, 70+-year-olds are probably still slightly different from 65-to-69-year-olds, for two reasons. First, they are a little older. Second, under Social Security rules it might be optimal to claim anytime from age 65-69, depending on the DRC amount and changes in the AIME, but it can never be optimal to claim after age 70, so all eligible 70+-year-olds should have claimed and should currently receive Social Security retirement benefits. Again, assuming that the change in the labor market environment captured by the comparison group of 70+-year-olds understates the change among the 65-to-69-year-olds, this time the results represent the upper bound on the coefficients. As 70+-year-olds increase labor supply, 65-to-69-year-olds probably increase it even more,

even absent the repeal. On the other hand, the general equilibrium story would point in the opposite direction—as 70+-year-olds worked more, they made the effect of the repeal appear smaller than it really is. Overall, the effect on the labor supply of 65-to-69-year-olds might not be quite as large as the results would indicate.

#### B. Caveats with a Differences-in-Differences Model

One problem that might arise from using a differences-in-differences model is that I am imposing a linear structure on the effect. Essentially, the underlying assumption is that each age group experiences a change in the intercept in response to the repeal. The intercept is assumed not to change for the comparison group, but all other coefficients are assumed constant. This assumption might be too restrictive, especially when I find that breaking up the target group by individual ages gives me different effects for different ages.

In addition, the differences-in-differences model captures some unspecified average effect for each treatment group. In reality, each individual treatment group contains a great deal of heterogeneity because each individual faces a different effective tax rate from the RET due to differing life expectancies and lifetime earnings profiles. Further, it matters greatly where a worker is situated on his budget constraint. Even in a simplified case where the RET is the only tax that is applied to a worker's earnings, there are three different segments a worker can be located at along his budget constraint. He could be on the first segment, earn less than the exempt amount, receive his full Social Security benefits, and face a zero marginal tax rate. Alternatively, he could be located on the second segment, face a marginal tax rate between zero and 33 percent, depending on

his life expectancy and other factors, and receive part of his Social Security benefits. Finally, he could be located on the third segment, face a zero marginal tax rate and receive no Social Security benefits.<sup>14</sup> Workers on each segment experience different effects. At first glance, workers on the first segment face a change only in an irrelevant part of their budget set as they are in the region where their marginal tax rate was already zero before the repeal. However, if workers are not be able to choose the exact desired amount of hours worked, they might experience the repeal as a tax cut. Workers on the second segment could theoretically increase or decrease labor supply, though the substitution effect should dominate for most of them. On the one hand, there is a large substitution effect as the marginal tax rate jumps from 33 percent to zero, inducing an increase in labor supply. At the same time, those workers can now claim full benefits while working, so there is also an income effect, lowering labor supply, though that effect should be rather small. Workers on the third segment experience purely an income effect. Their marginal tax rates stay unchanged, but now they can claim full benefits at the same time.

While it is arguable that workers on the first and second segment are probably quite similar in their response to the repeal, workers on the third segment behave entirely differently. In reality, the effect captured by my model is a weighted average of workers on those three segments, where I do not know the weights, and the effect I find is really a lower bound on the effect of the repeal on the first and second segment, tempered by the negative effect on those on the third segment. Among 65-to-69-year-olds, I estimate that

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<sup>14</sup> He could also be on either of two kinks, either facing a 33 percent marginal tax rate and still receiving full benefits or facing a zero marginal tax rate and receiving no benefits as on segment three.

between 8.5 percent and 13 percent were on the second segment before the repeal, while between 3.5 percent and 8 percent of them were located on the third segment.<sup>15</sup>

## **VII. Data Set**

I use data from the Outgoing Rotation Groups (ORG) from the Current Population Survey (CPS) from January 1996-December 2002. The CPS is a monthly survey of about 50,000 households conducted by the Census Bureau for the Bureau of Labor Statistics. The sample is representative of the noninstitutionalized civilian population. The CPS interviews households the same four months in a two-year period. Households in the fourth and eighth months are part of the ORG, and contain detailed data on earnings. Observations from the fourth (eighth) month in the CPS will be referred to as ORG 1 (ORG 2). The two dependent variables I use from the CPS are hours worked last week and a discrete indicator that equals 1 if a person is at work and 0 otherwise.

My sample includes white men aged 55-75 in the fourth month of the CPS. Age in the CPS is age in years in the interview week. This imprecise measure is unfortunate because Social Security rules are contingent on exact age. The only sample restriction is that none of the white men in my sample be disabled, so only persons for whom working is a choice variable are included. When I eliminate the disabled, my sample size is reduced from 76,864 to 71,311.

Table 2.5.2. presents means of work and hours variables for observations on ORG 1 before the repeal and after the repeal. From that table, I can see that both the employment rate and hours worked are much higher after the repeal than before the

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<sup>15</sup> Those numbers are based on the annual PIA using March Supplement CPS data 1996-2002.

repeal for 65-to-69-year-olds. My test for whether the repeal of the RET changed labor supply behavior among the elderly is to compare employment rates and hours worked for workers aged 55-75 before and after the repeal.

For the equations where I use employment and hours worked as my measures of labor supply, I use observations from ORG 1 of the CPS. For the equations where I use entry, exit, and continuation as my measures of labor supply, I use observations from ORG 2 of the CPS because I know whether those individuals were at work a year ago or not, and can construct entry, exit, and continuation variables. Sample 2 consists of ORG 2 observations from January 1997-December 2002. Note that Sample 2 contains 56-76-year olds, so the young group consists of 56-61-year-olds, and the comparison group consists of 70-76-year-olds for this sample.

I construct two different samples containing observations from ORG 1. The first ORG 1 sample, Sample 1A, spans January 1996-December 2001 and together with Sample 2 builds my panel from which I can construct entry, exit, and continuation variables. I match individuals across time by household identification number, line number, race, and sex. Attrition does not seem to be too big a problem in my panel, as I retain 59,127 observations for Sample 2 of originally 68,204 observations in Sample 1A. Table 2.6. shows the means in ORG 1 for those who attrit from the sample and those who do not.

The other ORG 1 sample, Sample 1B, spans January 1997-December 2002. Sample 1B is used to test the effect of the RET on hours worked last week and employment last week. Its advantage over Sample 2 is that attrition is much lower from the perspective of the original CPS interview.



## VIII. Results

Table 2.7. shows the results from a simple before-after means test. I find that employment of 62-to-64-year-olds and 65-to-69-year-olds increased by a statistically significant 2.34 and 3.39 percentage points, respectively, after the repeal. On the other hand, employment of 55-to-61-year-olds remained unchanged, and while employment was higher post-repeal among 70-to-75-year-olds, the difference is only marginally significant. Looking at hours, 65-to-69-year-olds worked more hours after the repeal, as did 70-to-75-year-olds. On the other hand, the increase in hours among 62-to-64-year-olds and the decrease in hours among 55-61-year-olds are only marginally significant.

Tables 2.8.1.-2.8.5. present results for equation (1) with employment as the labor supply measure using Sample 1B in a probit model. Here, as I do subsequently for the probit models, I report the mean of the marginal effect for each coefficient. I estimate models for the full sample and two subsamples, limited to those who have attained at most a high school diploma and those who have at least a four-year college degree, respectively. I use the subsamples to examine whether the low-educated react more strongly to the repeal than the high-educated. Recall that the low-educated know that their life expectancy is shorter than that of the high-educated, so the RET affects them the most. I find that among the control covariates, older age is associated with lower labor supply across all education groups. In addition, employment tends to increase in education, both unsurprising results.

As shown in Table 2.8.4., in the full sample, the results suggest that 65-to-69-year-olds increase employment by 2.21 percentage points and the coefficient is

statistically significant. In addition, 55-to-61-year-olds reduce employment by 1.60 percentage points, but the coefficient is only marginally significant in the probit model. Among the low-educated, those workers who should react more strongly to the repeal, only the coefficient for the target group is statistically significant, suggesting that they increase employment by 3.31 percentage points. Among the high-educated, only the coefficient on the young is statistically significant, suggesting that they reduce employment by 4.00 percentage points.

Table 2.8.5. presents the results again with equation (1) using employment as a measure of labor supply with Sample 1B, but now the target group has been broken up into individual age groups to see if 65- and 66-year-olds are more affected by the repeal than 67-to-69-year-olds. Again, I estimate probit models for the full sample and subsamples by educational attainment. If there are labor force transition costs and entry is costly, I would expect 65- and 66-year-olds to increase labor supply much more than 67-to-69-year-olds. 65- and 66-year-olds are more likely to still be in the labor force and thus less likely to have transition costs imposed on them, so it is less costly for them to increase labor supply. Indeed, I find that while there is no increase in the employment rate for 66-to-69-year olds, 65-year-olds increase their employment rate by a statistically significant 4.91 percentage points in response to the repeal. Among the low-educated, I find that 65-, 67-, and 68-year-olds increase employment. By 5.93, 4.55, and 5.25 percentage points, respectively, and all but the coefficient for 67-year-olds is statistically significant, and that one is marginally significant. The results in a linear model are similar and are presented in Tables 2.9.1.-2.9.5.

Using hours as my dependent variable, I again find that labor supply is decreasing in age and decreasing in education. In Tables 2.10.1.-2.10.5., I present results for equation (1) using hours worked last week as my measure of labor supply with Sample 1B. Recall that in a life-cycle context younger workers should reduce their hours if they expect to increase hours between the ages of 65 and 69. I estimate models using a linear specification for the full sample and subsamples by educational attainment. In the full sample, I find that the target group has increased hours worked by a statistically significant 0.98 hours a week. The number might seem small at first, but the average hours worked by the target group before the repeal is a mere 9.09 hours a week, so this represents roughly an eleven percent increase. In addition, I find that 55-to-61-year-olds decrease their hours worked per week by a statistically significant 1.25 hours. Among the low-educated, 65-to-69-year-olds increase labor supply by 1.54 hours a week, and 55-to-61-year-olds decrease hours by 0.99 hours per week, and both coefficients are statistically significant. Among the high-educated, only the coefficient for the young group is statistically significant, and it suggests that 55-to-61-year-olds decrease labor supply by 1.71 hours a week.

Table 2.10.5. presents results using equation (1) using hours worked last week as a measure of labor supply with Sample 1B again, but now the target group has been broken up into individual age groups. I again estimate linear models using the full sample and subsamples. In the full sample, I find that the increase in hours in the target group is mainly driven by 65-year-olds who increase labor supply by a statistically significant 2.38 hours a week. For 65-year-olds, the average number of hours worked a week prior to the repeal is 11.19, so this increase translates to about a 21 percent increase.

None of the other coefficients are statistically significant. When I estimate the model using the subsamples, I find that, among the low-educated, 65-, 67-, and 68-year-olds increase the number of hours by 2.64, 1.83, and 2.31, respectively, and the coefficients are all statistically significant.

Tables 2.11.1.-2.11.5. present results with equation (1) using employment, entry, exit, and continuation as my measures of labor supply with Sample 2. I present the results for the control covariates first and find that again, employment is increasing in education and decreasing in age. Entry is decreasing in age, but there is not systematic pattern between entry and education. Exit peaks at age 62 and is high at ages 64 and 65, but then decreases, not surprising considering that the exit rate cannot be high if employment is low. Continuation is decreasing in age but increasing in education. I use entry, exit, and continuation to examine whether the change in employment came about because more or fewer workers enter, exit, or continue in the labor force than before. In those tables, I again estimate probit models. I re-run equation (1) with employment as my measure of labor supply with this new sample to ensure that the results from Sample 2 do not differ too much from Sample 1B because of attrition and a slightly different age composition. Using employment as the labor supply measure, I find that I get the expected signs for the young and the target group, negative and positive, respectively. The coefficient on the target group is statistically significant and the magnitude of the coefficients is similar to that of the results from Sample 1B in Table 2.8.4. When I use entry, exit, and continuation as my measures of labor supply, respectively, I find that the repeal affects the continuation rate of the target group, as the likelihood of 65-to-69-year-olds' continuing in the labor force has increased by a statistically significant 2.43

percentage points in response to the repeal. Neither entry nor exit of 65-to-69-year-olds is statistically significantly affected by the repeal. Table 2.11.5. presents results with equation (1) using employment, entry, exit, and continuation as my measures of labor supply with Sample 2 with the target group broken up into individual age groups. In those tables, I again estimate probit models. I find that the continuation rate of 65-year-olds has increased by a statistically significant 4.78 percentage points. I get similar results using a linear probability model, shown in Tables 2.12.1.-2.12.5.

I return to the issue of whether the low-educated are differentially affected from the high-educated, using Wald tests, reported in Table 2.13. I find that the coefficients for the low-educated are statistically significantly different from the high-educated. The tests do indicate that jointly the model for the low-educated is different from that for the high-educated. However, allowing all control covariates to differ between the low- and the high-educated, I find that the coefficients of interest are not statistically significantly different in the low-educated sample from the high-educated sample.

#### A. Robustness Tests

I run several robustness tests. First, I use April 2000 as the date of law change, when the law was passed and announced to the public instead of January 2000, the date for which the law was effective. The results do not change much. All coefficients of significance still point in the same direction. The only changes are that the coefficient on the target group in Sample 1B is now only marginally significant in the estimation on employment rate in the probit model. Also, the coefficient on age 68 when I break up the target group by individual age is now no longer marginally significant. Finally, in the

low-educated sample, the coefficient on age 65 is now only marginally significant. In the hours equation, all coefficients stay virtually the same. In the conditional equation using Sample 2, the coefficient on the young is now statistically significant and the coefficient on the target group is now just marginally significant in the employment equation. Further, the effect on exit among the young is no longer marginally significant. Also, the coefficient on age 66 for the continuation and employment estimations is now marginally significant. Finally, the coefficient on age 68 is now no longer marginally significant (see Tables 2.14.1.1.-2.14.5.2.).

When I add disabled persons back to the sample, again, results do not change much. Now, in the employment equation for the modified Sample 1B I get a statistically significant coefficient for the young in the linear and probit models instead of a marginally significant one. Furthermore, the coefficient on the target group is only marginally significant in the probit model for the low-educated. Additionally, the coefficient on age 68 is no longer marginally significant in the overall sample in the linear model overall, and now only marginally significant in both linear and probit model among the low-educated. Finally, among the low-educated, the coefficient on 67-year-olds is now no longer marginally significant in the probit model. For the hours equation, I find that among the low-educated, the coefficients on ages 67 and 68 are no longer statistically significant. In the conditional equation, the coefficient on exit among the young is no longer marginally significant. In addition, the coefficient on the target group in the probit estimation is now only marginally significant. Finally, the effects I find for the young in the employment equation are now statistically and marginally significant, respectively, in the linear and probit model, respectively (see Tables 2.15.1.1.-2.15.5.2.).

Overall, all significant (or marginally significant) coefficients retain the same sign, only the level of significance changes, if that.

## B. Auxiliary Results

If the repeal was the impetus for the increase in labor supply among the target group, I would expect the number of full-time workers to increase in the target group. If the increase in labor supply were concentrated among the part-time workers, one would think that the repeal had little to do with the increase, because part-time work was an option even before the repeal, as long as one stayed beneath the exempt amount. As shown in Figures 3.4. and 3.5., however, full-time work increased among the target group, while part-time work stayed roughly constant. Conversely, full-time work stayed relatively constant among the young and intermediate groups. The comparison group, the 70+-year-olds, experienced a slight increase in full-time work as well.

The repeal of the RET could have two effects. It could increase labor supply among those who were going to claim benefits by the normal retirement age regardless of any changes to the law. It could also entice individuals who were going to work from ages 65-69 to claim earlier. However, because the DRC is undergoing a number of increases before and after the repeal of the RET, the increase in labor supply that I found in the regression analysis could have happened for two different reasons. Those older men might be working now while also claiming benefits instead of simply claiming benefits and not working. Or those men could be working now and they are delaying claiming benefits because of more generous DRC's. Using 1997-2002 March CPS data, I also estimate simple correlations between earnings and Social Security income among

65-to-69-year-olds. I find that the correlation coefficient increases from -0.2477 to -0.0095, the negative correlation virtually disappearing. It not surprising that the coefficient was large and negative before the repeal. Because of the earnings test, there was a mechanical relationship prior to the repeal. However, what is of interest is that the coefficient becomes virtually zero after the repeal. Furthermore, I find that, before the repeal, among 4,347 white men aged 65-69, 3,741 (86.1%) received Social Security Benefits. Out of the 1,494 who worked that year, 1,148 (76.8%) received benefits. After the repeal, 1,682 out of 1,873 white men (89.8%) received benefits. Noteworthy is that out of the 678 that worked, now 582 (85.8%) received benefits.

## **IX. Discussion**

Using data from the Outgoing Rotation Groups of the CPS, I find evidence that the repeal of the RET spurred labor supply among 65-to-69-year-olds as measured by employment, especially among the low-educated. The coefficients are statistically significant in both probit and linear probability models. In addition, I also find that 65-to-69-year-olds increase labor supply as measured by hours worked last week.

Further, I find that younger workers, those between 55 and 61 years of age, reduce their labor supply in response to the repeal. This finding is not surprising when one examines the issue from a life-cycle perspective—younger workers are substituting work at a later age for work now because a penalty for work at a later age has been lifted. They are reacting to a reduction in their relative wage by lowering labor supply.

I further find that the increase in labor supply is largely driven by continuers, not re-enterers. The rate of 65-to-69-year-olds who were at work a year ago and still are at



work has increased by about two percentage points, while the rate of 65-to-69-year-olds who entered the work force over the past year has remained unchanged. People who were 67 at the time of the repeal are often already out of the labor force and do not re-enter. On the other hand, people who were 65 at the time of repeal are often still in the labor force, and their continuing to stay in the labor force drives the increase in the employment rate of the target group. Therefore, I expect the long-term effect of the repeal to be greater as those 65-year-olds who continue to stay in the labor force age. This result leads me to believe that the coefficient for the target group found here is actually a lower bound on the long-run impact of the repeal of the RET on labor supply.

The results I get when I break up the target group into individual ages suggest that the long-term effect will be greater. I find that continuation increases dramatically among 65-year-olds. Therefore, I would expect the effect of the repeal on the labor supply of older workers to increase over time as more workers opt to continue to stay in the labor force until a later age than now as workers re-optimize their labor supply behavior over their lifetime.

There is also some weaker evidence that the low-educated react more strongly than the high-educated. This finding suggests that those who are taxed the most by the RET, the low-educated who die youngest, are impacted the most by the repeal. Coefficients on regressions with the low-educated group tend to be greater in magnitude than those on regressions with the high-educated group. Moreover, coefficients among the low-educated group also tend to be significant more often. These results seem to support the theory that the RET taxes low-educated workers the most, so they should change their labor supply behavior most strongly after it was repealed. Wald-tests

confirm that the models are statistically significantly different for the low-educated from the high-educated.

I further find that the increase in labor supply by 65-to-69-year-olds is driven by an increase in full-time workers, while the share of part-time workers stays roughly constant in that age group. This result suggests that the increase in labor supply was driven primarily by full-time workers and that the repeal of the RET is the impetus for the increase in labor supply. The RET was not discouraging part-time work as much as full-time work, as workers could still work part time with no penalty as long as they earned less than the exempt amount. Furthermore, when looking at simple means and correlations between labor supply and Social Security benefits receipt, I find that, after the repeal, more 65-to-69-year-old men receive benefits while working. This result suggests that the increase in labor supply was primarily driven by men who were going to claim benefits by the normal retirement age regardless of any changes to the law, rather than by men who are working longer because of the increase in the DRC.

## CHAPTER THREE

### WHY IS HUSBAND'S LABOR SUPPLY STRONGLY CORRELATED WITH WIFE'S LABOR SUPPLY? THE CASE OF OLDER COUPLES

#### I. Introduction

Presently, there are many proposals to extend the solvency of the Social Security program. For the most part, research on the effect of Social Security on labor supply has focused on the retirement behavior of men. But there is increasing recognition that to fully understand the implications of policy changes one needs to consider the retirement behavior of men and women jointly. In fact, because of support from survivor's benefits and the greater longevity of women, the majority (60 percent) of adults receiving Social Security benefits in old age are women. And given the historical rise in the labor force participation rate of women over the last 40 years, policy effects on women's retirement choices will have increasing importance.

This paper focuses on the joint labor supply decisions of husbands and wives. There are large literatures on male labor supply and retirement, and there is a moderately-sized literature on family labor supply for prime-age couples. By comparison, very few studies examine the family labor supply of older couples. This is perhaps surprising in light of evidence that retirement frequently occurs for both spouses within one calendar quarter (see Blau (1988)).

As Hurd (1990) points out, there are three reasons why exit from the labor force may be positively linked among spouses. First, in equilibrium, assortative mating may lead spouses to behave similarly for unobservable reasons. For example, spouses may

have common preferences for leisure, and the inherent unobservability of these preferences is one reason why simple tabulations may not identify causal relationships. But assortative mating can lead to correlation in wages, too. For example, suppose spousal education is positively correlated and more educated men and women work for more of their lives. Then, even if each partner's labor supply decision is made independently of the other's, cross-sectional analyses will show a positive correlation between spouses' labor supply. To the degree that factors associated with both marriage and labor supply can be observed, one can control for them directly. However, offered wages for retired people must be estimated, again raising the possibility of bias due to unobserved heterogeneity.

Many unobservable factors likely affect both marriage and work. While assortative mating provides one example of problematic unobserved heterogeneity, it is not difficult to think of other sources of bias. Married couples share their lives in many ways, and it is likely that joint unobserved factors affecting husbands' labor supply affect wives' labor supply as well. For example, because husbands and wives live together, unobserved local economic conditions will simultaneously affect labor supply and retirement decisions of both spouses.

A promising approach in investigating the extent of leisure complementarities involves finding changes in retirement incentives that face one spouse but plausibly do not affect the other. This variation is then used to construct instrumental variables for the retirement status of the directly affected spouse.

Baker (1999) provides a good example of this approach in studying a policy change in the Canadian Income Security system. The Spouse's Allowance extended

retirement benefits to a spouse aged 60 to 65 (the previous minimum was 65) provided they have a spouse aged 65 or older. Hence there is variation in eligibility because of differences in spouses' ages. As long as these age differences can be treated as exogenous, spousal coverage by the program becomes exogenous, conditional on own age. Since this program largely targeted women, it provides an opportunity to examine husbands' response to wives' incentives.

In this paper, I examine the joint labor supply of husbands and wives in a family context with separability of utility from consumption and leisure. My approach accounts for the joint labor supply decisions of married couples. I explicitly investigate the role of liquidity constraints as well as the effects of wage variation on labor force participation. I do allow separate effects of own age on each spouse's labor supply decision. I model both exogenous changes in liquidity at a known age and exogenous changes in wages due to variation across states in age discrimination laws.

I introduce two empirical tests of the underlying model using two sources of variation. My first source of variation in economic variables concerns liquidity constraints. It is well-known that there is a spike in retirement at age 62. One explanation for this spike is the fact that Social Security benefits cannot be claimed until age 62. As Rust and Phelan (1997) and others have argued, this restriction means that before a spouse turns 62, some families will be liquidity constrained with respect to benefits available to that spouse. Hence for liquidity constrained couples, a spouse's turning 62 acts like a scheduled "helicopter drop" of money. This increase in accessible wealth will lead the family to increase consumption over the balance of the lifecycle and will reduce the marginal utility of wealth. As noted above, it will also lead to reductions in optimal

labor supply for both spouses, much like increases in nonlabor income do in a static model. The key point to note is that each spouse's turning 62 has an effect on wealth for liquidity constrained households. I can thus investigate the effects of liquidity constraints associated with each spouse on the other's behavior.

This argument suggests the benefit of a very simple strategy for establishing the role of liquidity constraints on spousal labor supply—investigating the correlation between own labor supply and spouse's age. My key assumption (as in Baker (1999)) is that age differences between husbands and wives are otherwise exogenous to the labor supply decision once I have controlled for own age. Once this assumption is made, the presence of liquidity constraints suggests that not only should labor supply of one's spouse labor supply be discontinuous at ages 62 and 65, but when a spouse turns 62 and 65, one's own labor supply should be affected as well. Despite its simplicity, to my knowledge this approach has not been investigated to date.

I test the null hypothesis that there are discrete changes in neither preferences nor wages at age 62. Even after controlling for a quartic in each spouse's age, this hypothesis is strongly and consistently rejected in my data, which come from the U.S. Censuses for 1970-1980. The results imply that either there is a discrete increase in preferences for own leisure at age 62, or effective (perceived) wages rise at that age.

The second source of variation focuses on anticipated changes in wages. In MaCurdy's (1981) terminology, this amounts to investigating changes in labor supply arising from movement along a spouse's lifetime wage profile. The effect of such changes in the price of a husband's time on his wife's consumption of leisure (and vice-versa) is a test of direct complementarity of spousal leisure. I argue that state-level

variation in age discrimination laws effectively shift a covered person's wage. Neumark and Stock (1999) show that protection from age discrimination increases the fraction of older men in the labor force while steepening age-earnings profiles. If differences across states in spousal age differences are exogenous to labor supply decisions, then I can exploit the fact that own protection from age discrimination is not perfectly collinear with spousal protection and use the same variation while controlling for both age and state effects. In light of Neumark and Stock's (1999) results, I interpret these laws as effectively increasing wages when workers are older. Controlling for own coverage by the laws, I can thus use spousal coverage as a source of variation to identify cross-price effects, and hence the sign of any complementarities in spousal leisure.

In section II, I briefly review some relevant papers. In section III, I discuss age discrimination laws in place between 1970 and 1980 and discuss the effects of these laws on wages. The details of the laws will be important in the specification of the econometric model. In section IV, I describe my data. In section V, I present the econometric model. In section VI, I report my main empirical results. I find that labor force participation of both spouses changes significantly when either turns 62, even after controlling for demographic variables. Unless this finding is to be attributed to discrete changes in either preferences or wages occurring at exactly age 62, this evidence suggests that the labor supply choice of each spouse is significantly related to liquidity effects deriving from either spouse's age. I also show that protection from age discrimination raises own labor force participation for both spouses in 1970, but only for husbands in 1980, extending Neumark and Stock's (1999) results to women as well as men. However, I find no evidence that spousal coverage either raises or lowers own labor force

participation. This finding suggests that husbands' leisure and wives' leisure are not complements. Finally, in section 7, I offer some conclusions.

## **II. Brief Review of the Literature**

Hurd (1990) was one of the first to document the link in the retirement behavior between husbands and wives. Using the New Beneficiary Survey (NBS), he finds that the retirements of spouses occur as a joint process, as spouses often retire within a short time of each other. Further, he finds that age differences explain large differences in retirement age. However, he offers no explanation as to why joint retirement takes place, merely that it could be due to unobservables, assortative mating, or complementarities in leisure, leaving the question open for future research.

Blau (1998) confirms this link using the RHS. He finds that the probability of exiting the labor force in a given quarter is greater for the individual if the spouse is out of the labor force. His identification strategy exploits the fact that husbands and wives can be in one of four different labor force states—both can work, husband works, wife does not, or both do not work. For example, a wife whose husband is already out of the labor force exits the labor force the following calendar quarter with probability 3.53 percent. On the other hand, the probability of her exiting the labor force if her husband still is in the labor force is a mere 2.40 percent. For the husband the story is similar: with the wife already out of the labor force, the probability of his exiting is 2.61 percent. If she still is in the labor force, the probability of exiting is 1.98 percent. Conversely, if the spouse is already out of the labor force, the individual is also less likely to enter the labor force. In addition, if the spouse is out of the labor force, the individual is also less likely



to enter the labor force. For the wife, if the husband is out of the labor force, her likelihood of entering the labor force in the following quarter is 0.52 percent. If the husband is employed, that probability increases to 1.11 percent. The results for the husband are similar. If the wife is out of the labor force, the probability of the husband's entering is 1.44 percent. If the wife is in the labor force, the probability of the husband's entry jumps to 2.24 percent. Regression analysis confirms this finding that he obtained from simply examining the descriptive statistics. His study provides compelling evidence for a link between spouses' labor supplies.

Baker (2002) examines the impact of the introduction of the Spouse's Allowance (SPA) in Canada on the labor supply behavior of couples. The SPA made it possible for individuals to receive Canadian Income Security (IS) benefits as early as age 60, provided their spouse was at least 65 years of age. This implicitly made joint retirement less costly for spouses, as the younger spouse was immediately eligible for IS benefits. Exploiting age differences as the identification strategy, he found an increase in the rate of being out of the labor force among males aged 65 to 75 with SPA-eligible spouses, while the control groups (males aged 65 to 75 with spouses aged 65 to 75, males aged 65 to 75 with spouses aged 55 to 59, and single males aged 65 to 75) did not. Women eligible for SPA (females aged 60 to 64 with spouses aged 65 to 75), on the other hand, did not share in the increase in labor force participation those in the comparison cohorts, which comprised of women aged 60 to 64 with spouses either aged 50 to 59 or aged 60 to 64. Thus, there seems to be a joint response to the change in policy, which directly affected the wives only, suggesting some joint retirement behavior.

My approach builds on Baker (2002), using age differences between husbands and wives to identify the effect of each turning 62 and 65, respectively. Because the SPA induced wives to reduce labor supply, it is not possible to determine whether husbands reduced labor supply because of income pooling or complementarities in leisure. However, using variation in age discrimination laws across states and the variation in age differences, where coverage by those age discrimination laws induces increased labor supply, I can examine whether the joint response arises from complementarities in leisure or not, a question previously unanswered in the literature.

### **III. Age Discrimination Laws Across States**

Protection from age discrimination laws restricts firms' ability to dismiss, deny employment to, or reduce wages of workers on the basis of age. The ages for which protection is effective varies from state to state, but many states have maximum ages above which workers are no longer protected from age-related job actions.

Three states passed laws against age discrimination before 1940: Colorado, Louisiana, and Massachusetts. Six more states followed before 1960: Connecticut, New York, Oregon, Pennsylvania, Rhode Island, and Wisconsin. Between 1960 and 1980, 29 more states, including the District of Columbia, followed suit. In addition, many of the states which already had anti-age discrimination laws expanded coverage of their laws. From 1980 on, nine more states enacted their first state laws against age discrimination, in addition to the expansion of those laws in other states. Today, only four states have never passed a state law against age discrimination: Alabama, Arkansas, Mississippi, and South Dakota. Note that not only does the existence of such laws vary by state, but so

does enforcement of those laws. Therefore, state fixed effects in the econometric analysis will account not only for varying local economic conditions and other unobservables but also for the varying enforcement in each state. Coverage by those laws varies by state, but not by occupation within a state.

In 1968, Congress enacted the Age Discrimination in Employment Act (ADEA) to “promote employment of older persons based on their ability rather than age; to prohibit arbitrary age discrimination in employment; to help employers and workers find ways of meeting problems arising from the impact of age on employment.” In 1979, the ADEA was strengthened when enforcement authority was transferred to the Equal Opportunity Commission (EEOC). At the time, a number of states also had state laws banning age discrimination or mandatory retirement, with the laws generally covering people whose age falls in a specific range.

Aging out of age discrimination protection is essentially a reduction in wages, even when employment is the relevant margin. When a worker is dismissed because he or she is too old, the worker’s opportunity set of jobs has shrunk. Under the assumption that workers can always find a job at some lower wage, dismissal is effectively a wage reduction for workers with job-specific human capital. These assertions suggest that adjusting for age, workers covered by age discrimination laws should have higher wages than unprotected workers and should also be in the labor force more often; this is exactly what Neumark and Stock (1999) find.

If workers have perfect foresight concerning their likelihood of experiencing age discrimination, then the implied wage change is an evolutionary one and will not affect the marginal utility of wealth. By contrast, if there is some uncertainty over whether a

worker will be subject to age discrimination, then the wage change is discontinuous, and the marginal utility of wealth will have to adjust.

The cross-sectional differences I want to exploit in my analysis are those that arise from different age differences between the wife and the husband and geographic differences between couples, which determine whether there are any age discrimination laws in a state.

#### **IV. Data**

In my empirical work, I use the 1970 and 1980 Public Use Micro Samples from the Decennial Census long form surveys. In 1970, the census implemented two surveys, one conducted on 15 percent of households and the other on 5 percent of households. From each of these surveys, I use two independent samples representing 1 percent of the population, so that the combined files are a 4 percent sample of the U.S. population in 1970. The 1980 public use data files provide data on 5 percent of the U.S. population. I limit my sample to men and women who married far in advance of retirement age. Except when the underlying variables are unavailable, all couples in my sample satisfied the following selection criteria:

1. The husband was no older than 35 when the couple was married.
2. The marriage is the first one for the husband.
3. The husband is 55 to 74 years old.
4. The wife's age is within 15 years of the husband's.
5. The husband and wife both worked within a ten year time period prior to the Census year.

6. Neither the husband nor the wife were on active military duty in the week the household was surveyed.
7. Neither the husband nor the wife has a work-limiting disability.
8. No data used in the analysis or sample selection is allocated by the Census Bureau.

Table 3.1. presents summary statistics for the 1970 Census sample. In 1970, husbands' average age is 61.06 years, about 4 years older than their wives. On average, couples in my sample got married when the husband was 25.69 years old. This implies that the couples in my 1970 sample have been married an average of 35.37 years. 63.65 percent of couples live together with no other individual in their household. In 1970, 57.93 percent of husbands 55-74 and 47.31 percent of their wives had not received a high school diploma. In 1980, husbands' average age is 60.95 years, about 3 years older than their wives. On average, couples in 1980 married when the husband was 25.12 years old. This implies that the couples in my 1980 sample have been married an average of 35.84 years. 61.44 percent of couples live together with no other individual in their household. Of those that live with other people, 5.14 percent lived with a person 6 years of age or younger, 35.10 percent lived with a person aged 18 or younger, and 10.36 percent live with a person 70 years of age or older.

Individuals are recorded as "Out of the Labor Force (OLF)" if this was their labor force status in the week before responding to the Census. People are recorded as "In the Labor Force" if they were at work, not at work but with a job or were unemployed in the Census week. According to Table 3.1., 20.31 percent of men and 34.44 percent of women in my sample were not in the labor force in the Census week in 1970.

Table 3.2. presents a cross-tabulation of labor force participation rates for both spouses in 1970. This table highlights the central correlation on which I focus: correlation between the husband's choice and his wife's choice to participate in the labor force. Table 3.2. makes clear the strong correlation in spouses' choices: in 1970, 70.2 percent of the wives of working husbands worked, while only 44.7 percent of the wives of non-working husbands worked. Therefore, when a husband is not in the labor force, the likelihood that his wife is also out of the labor force is 25 percentage points higher than when the husband is working.

In 1970, I start out with 527,165 men aged 55 to 74. Out of those, 448,634 were married. When I impose the qualifier that spouses' ages be within 15 years of each others', the sample shrinks to 429,400. Out of those, 410,123 were married by the age of 35. Of those, 373,989 were still in their first marriage. Of those, there were 354,638 couples where neither spouse had a disability that prevented them from working. Among them, 171,111 were couples where both spouses had worked sometime in the past ten years. Finally, 123,728 had no relevant variables allocated for either husband or wife. All 123,728 of those couples had no missing variables for the employment variable, but only 123,581 had no missing variables among the relevant ones.

In 1980, I start out with 792,855 men aged 55 to 74. Of those, 679,616 were married, and 656,350 were within 15 years of their wife's age. 611,924 were married by the age of 35, and 505,871 of those were married just once. 388,395 of those are couples where neither is disabled so as to prevent any work. 228,613 of those were couples where both spouses worked at some time during the past ten years. 132,414 of them had no relevant variables allocated, and 132,310 had no missing values for labor force status.

## V. Econometric Model

I address the question of whether, after controlling for own age and other characteristics, spouse's age still affects one's own labor supply. I estimate a reduced form probit, with a dummy for labor force participation as the dependent variable. The dummy equals 1 if the person is out of the labor force, and 0 if the respondent was at work with a job, not at work with a job, or looking for a job.

The baseline model is a univariate probit model with the following equations:

$$l_{ht}^{*i} = \theta_{h0}^i + D_{62ht}^i \theta_{h62h} + D_{62wt}^i \theta_{h62w} + (D_{62ht}^i * D_{62wt}^i) \theta_{h,both62} + D_{65ht}^i \theta_{h65h} \quad (2)$$

$$+ D_{65wt}^i \theta_{h65w} + (D_{65ht}^i * D_{65wt}^i) \theta_{h,both65} + X_{ht}^i \theta_{hh} + X_{wt}^i \theta_{hw} + \varepsilon_{ht}^i$$

$$l_{wt}^{*i} = \theta_{w0}^i + D_{62ht}^i \theta_{w62h} + D_{62wt}^i \theta_{w62w} + (D_{62ht}^i * D_{62wt}^i) \theta_{w,both62} + D_{65ht}^i \theta_{w65h} \quad (3)$$

$$+ D_{65wt}^i \theta_{w65w} + (D_{65ht}^i * D_{65wt}^i) \theta_{w,both65} + X_{ht}^i \theta_{wh} + X_{wt}^i \theta_{ww} + \varepsilon_{wt}^i$$

where  $l_{ht}^{*i}$  and  $l_{wt}^{*i}$  are desired leisure (latent) of the husband and the wife, respectively, i.e., dummy variables for whether husband and wife, respectively, are out of the labor force.  $D_{62ht}^i$  is a dummy variable that equals 1 if the husband is at least 62 years old,  $D_{62wt}^i$  is a dummy variable that equals 1 if the wife is at least 62 years old,  $D_{65ht}^i$  is a dummy variable that equals 1 if the husband is at least 65 years old, and  $D_{65wt}^i$  is a dummy variable that equals 1 if the wife is at least 65 years old. The coefficients on the dummy variables are the coefficients of interest. The dummy variables for being at least 62 and 65 account for the well known discontinuous drops in labor force participation when one turns 62 and 65. The more interesting variables are the dummy variables for the spouse if he or she is at least 62 and 65 years old, and the interaction terms for both spouses being at least 62 and 65 years of age. In a family labor supply context with

income pooling and liquidity constraints, either spouse's turning 62 results in an increase in family wealth and should thus affect labor supply of both spouses. I also include interaction terms because, for example, the wife might be drawing spouse's benefits<sup>16</sup>, so there might be an additional wealth effect for both being at least 62. At age 65, the earliest age for Medicare eligibility and often the earliest age for pension receipt, there might again be wealth effects at work. In addition to the wealth effects, one might imagine that there are complementarities in leisure, so if turning 62 or 65 induces one spouse to retire (because of mandatory retirement or other reasons), the other one might retire with the spouse, not because of wealth effects, but because they might enjoy spending time together.  $X_{ht}^i$  is a vector control for the demographics of the husband, which includes a dummy variable for whether the husband is white, a vector of dummy variables for educational attainment for the husband, a quartic in the demeaned age variable, and a dummy variable for whether the observation is drawn from the 15 percent sample for 1970,  $X_{wt}^i$  is a vector control for the demographics of the wife, analogous to the one for the husband plus whether she has had any children, and if yes, how many, and  $\varepsilon_{ht}^i$  and  $\varepsilon_{wt}^i$  are the error terms for the equations for the husband and the wife, respectively. I also run an alternative model where I include state fixed effects to account for local labor market conditions and other unobservables. The one problem with the model is that one might suggest that  $\theta_{h0}^i$  and  $\theta_{w0}^i$  are actually endogenous and depend on the marginal utility of wealth, which in turn depends on the income over the couple's lifetime. If there are cohort-specific changes in wages or Social Security policy, the marginal utility of wealth will be correlated with the age-62 (and age-65) dummy variables. Because of the

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<sup>16</sup> Spouse's benefits are benefits one receives on account of a spouse's Social Security earnings history, up to 50% of what the working spouse receives.



focus on the liquidity effects at age 62 (and age 65), I believe the problem is not too severe in a reduced form model. However, to better capture unobservables that might affect the marginal utility of wealth, I run a bivariate probit model as well that will help control for unobservables that jointly affect both spouses, i.e., local labor market conditions and wealth.

The second part of the analysis uses coverage by age discrimination laws as an instrument for wages. I use an instrumental-variables approach to determine whether the apparent joint retirement comes about because of complementarities in leisure. Coverage by age discrimination laws should serve well as an instrument because in itself coverage does not affect the labor supply decision itself. However, as Neumark and Stock (1999) have shown, coverage does affect the offer wage, and via that channel, labor supply. Furthermore, while coverage by age discrimination laws effectively increases one's own wage in old age, it should have no effect on the spouse's wage. If I find that own coverage increases labor supply while spouse's coverage does not, this result would point toward the absence of complementarities in leisure. Specifically, I use the following equations for my bivariate probit analysis:

$$\begin{aligned}
 I_{ht}^{*i} = & \theta_{h0}^i + D_{62ht}^i \theta_{h62h} + D_{62wt}^i \theta_{h62w} + (D_{62ht}^i * D_{62wt}^i) \theta_{h,both62} + D_{65ht}^i \theta_{h65h} \\
 & + D_{65wt}^i \theta_{h65w} + (D_{65ht}^i * D_{65wt}^i) \theta_{h,both65} + D_{hpadt}^i \theta_{h hpad} + D_{wpadt}^i \theta_{h wpad} \\
 & + X_{ht}^i \theta_{hh} + X_{wt}^i \theta_{hw} + \varepsilon_{ht}^i
 \end{aligned} \tag{4}$$

$$\begin{aligned}
 I_{wt}^{*i} = & \theta_{w0}^i + D_{62ht}^i \theta_{w62h} + D_{62wt}^i \theta_{w62w} + (D_{62ht}^i * D_{62wt}^i) \theta_{w,both62} + D_{65ht}^i \theta_{w65h} \\
 & + D_{65wt}^i \theta_{w65w} + (D_{65ht}^i * D_{65wt}^i) \theta_{w,both65} + D_{hpadt}^i \theta_{w hpad} + D_{wpadt}^i \theta_{w wpad} \\
 & + X_{ht}^i \theta_{wh} + X_{wt}^i \theta_{ww} + \varepsilon_{wt}^i
 \end{aligned} \tag{5}$$

where  $D_{\text{hpadt}}^i$  is a dummy variable that equals 1 if the husband is covered by age discrimination laws,  $D_{\text{wpadt}}^i$  is a dummy variable that equals 1 if the wife is covered by age discrimination laws, and the remaining variables are as defined for equations (2) and (3). I would expect that own coverage by age discrimination laws increases one's employment rate (or decreases the rate at which one is out of the labor force), as Neumark and Stock (1999) find. I have no priors for spousal coverage by age discrimination laws. Should I find that spousal coverage increases the rate at which one is out of the labor force, this result would point toward substitution in leisure, while a decrease in the rate at which one is out of the labor force would point toward complementarities in leisure. I again run an alternative model where I include state fixed effects. Recall that with age discrimination laws, state fixed effects account not only for local labor market conditions and other unobservables as before, but also for the difference in enforcement of those laws across states. In the first-stage regression on wages of husbands and wives, respectively, I find F-values of 7.30 and 22.78, suggesting that coverage by age discrimination laws is highly correlated with wages. Therefore, the identifying variation in my model comes from the variation in age differences between husbands and wives and the variation in laws that applies to the couples based on the state they live in.

## **VI. Results**

I report results from using a bivariate probit model with state fixed effects in Tables 3.4.1 and 3.4.2. Here and subsequently, results reported are the means of the marginal effects. I find that the husband's being white increases the probability of being

out of the labor force for both husband and wife. While hard to interpret, my highly stratified sample is more than 90 percent white, so the result might not be meaningful. Further, I find that own education decreases the likelihood of being out of the labor force for both husbands and wives, as expected. On the other hand, spouse's education increases the probability of being out of the labor force for both husbands and wives. Having no children increases the probability of being out of the labor force for both, but conditional on having children, having more children increases the likelihood of being out of the labor force compared to fewer for both. Own age increases the probability of being out of the labor force for both. On the other hand, spouse's age has no effect on the wife, and a small but statistically significant effect on the husband, increasing his probability of being out of the labor force.

Among the parameters of interest in Table 3.4.2., all coefficients are of the expected sign. In the husband's equation, his being at least 62 increases his probability by a statistically significant 3.23 percentage points. In addition, his being at least 65 increases the probability by a statistically significant 10.38 percentage points, and his wife's being at least 65 increases it by a marginally significant 1.49 percentage points, while the interaction term for age 65 implies a statistically significant decrease by 1.78 percentage points if both are 65 or older.

For the wife, both her being at least 62 increases her probability by a statistically significant 4.75 percentage points, while her and her husband's being at least 62 increases the probability by a statistically significant 1.87 percentage points. Moreover, her being at least 65 increases the probability by a statistically significant 6.69 percentage points, and her husband's being at least 65 increases it by a marginally significant 1.37

percentage points. The results are quite similar without state fixed effects and using a univariate probit, while most of the coefficients are similar using a linear probability model (see Tables 3.4.1. and 3.4.2. for results without state fixed effects and Tables 3.5.1. and 3.5.2. for the results using a univariate probit model).

For 1980, I present the results in Tables 3.6.1. and 3.6.2. For the control covariates, there are only a few major differences compared to 1970. Wives' being white decreases the probability of husbands' being out of the labor force. Also, the wife's education has no effect on husband's labor force status now. Furthermore, conditional on having children, the number of children does not affect the wife's labor force status, while it reduces the probability of the husband being out of the labor force.

As for the parameters of interest, those results are also quite similar to those from 1970, with all own age effects again of the expected sign and statistically significant, as presented in Table 3.6.2. I find that the husband's being at least 62 increases his probability by a statistically significant 7.08 percentage points, and his wife's being at least 62 increases the probability by a statistically significant 2.03 percentage points. In addition, his being at least 65 increases his probability by a statistically significant 6.39 percentage points. For the wife, her being at least 62 increases her probability by a statistically significant 6.43 percentage points, and their both being at least 62 increases her probability by a statistically significant 3.12 percentage points. Further, her being at least 65 increases her probability by a statistically significant 2.88 percentage points, and their both being at least 65 increases her probability by a statistically significant 5.25 percentage points. Results without state fixed effects, presented in Tables 3.6.1. and

3.6.2., and using a univariate probit model, presented in Tables 3.7.1. and 3.7.2., are quite similar.

I now focus my attention on the role of protection from age discrimination. Because age discrimination laws were being implemented in the years surrounding 1970, there is a good degree of variation in the laws across states for the 1970 data. The 1970 estimates of own- and cross-effects related to age discrimination laws seem to confirm the absence of complementarities in leisure. Again I find strong evidence that coverage by age discrimination laws lowers the rate at which both husbands and wives are out of the labor force in 1970. Tables 3.8.1 and 3.8.2. present results for 1970 with protection from age discrimination as an instrument for wages. The control covariates are very similar to those from the estimation without age discrimination. Most of the coefficients for the age-dummy variables are similar as well, with one exception: both being at least 65 years of age now increases the wife's probability of being out of the labor force by a marginally significant 2.00 percentage points. As for the parameters of interest in this estimation, own coverage by age discrimination laws and spousal coverage by age discrimination laws, the only statistically significant coefficients are that the husband's probability of being out of the labor force decreases by 1.94 percentage points if he is covered, while the wife's probability decreases by 2.48 percentage points if she is covered. I find no evidence that spousal coverage affects labor force status of either husband or wife. Removing state fixed effects, the effect of the wife's coverage on her own labor force status disappears, but there are no cross effects, as shown in Tables 3.8.1. and 3.8.2.

In 1980, results for the control covariates are similar to the estimation without age discrimination laws. Tables 3.9.1 and 3.9.2. presents the results for equations (4) and (5) for 1980. The coefficients of interest are statistically significant only in the husband's equation. In the bivariate probit model with state fixed effects, his probability of being out of the labor force decreases by 1.50 percentage points if he is covered. If I remove state fixed effects, as also shown in Tables 3.9.1. and 3.9.2. the results suggest that his probability increases by 1.73 percentage points if his wife is covered, and decreases by 1.03 percentage points if he is covered. I believe that the model with state fixed effects better reflects the differences in laws and enforcement across states, but neither the results with or those without state fixed effects suggest the existence of complementarities in leisure.

## **VII. Discussion**

In this paper I attempt to carefully investigate the channels through which joint labor supply decisions will operate among older couples. For 1970 and 1980, I find consistent evidence that linkages in labor force participation exist beyond those due to unobserved heterogeneity in preferences and wages in the form of cross and interaction effects. I find particularly strong evidence in 1980 that turning 62 has a strong effect on one's spouse's labor force participation. In addition there is evidence that turning 65 has such effects.

These results suggest two conclusions. First, liquidity constraints appear to play a role in labor force participation decisions among older couples, even after controlling for the marginal utility of wealth. This finding confirms some research, e.g. Rust and Phelan

(1997). Second, effects of relaxing liquidity constraints related to one spouse's age are present for both spouses, suggesting some degree of intrahousehold budgeting.

Finally, I find evidence that own protection from age discrimination raises labor force participation for both men and women. Despite the strength of this effect, there is no evidence that spousal coverage by age discrimination laws affects own labor force participation. This finding suggests an absence of spousal leisure complementarities, contrary to what has often been suggested as a major reason for joint retirement. A caveat, however, should be mentioned. Because the data used are from the 1970 and 1980 Censuses, this paper looks at an historical event. Large changes in female labor supply and family structures, especially given the sample selection criteria, limit the applicability of the findings.

## CHAPTER FOUR

### CONCLUSION

Whether Social Security affects labor supply has long been a topic of interest in economics. The peak in retirement at age 62 has often been attributed to age 62 being the earliest age at which one can claim Social Security old-age benefits. Furthermore, the peak at age 65 has long been attributed to the earliest age at which one is eligible for Medicare and pension benefits. In this dissertation, I find evidence that supports the hypothesis that Social Security rules heavily influence the retirement decisions of older Americans.

In chapter two, I examine the effect of the Repeal of the Retirement Earnings Test (RET) on the labor supply of older men. After demonstrating that the RET acted effectively as a tax on earnings in old age for most men, I use a differences-in-differences model to estimate whether labor supply, as measured alternately by hours of work last week and employment last week, increased for the group directly affected by the repeal, 65-to-69-year-olds. I use 70+-year-olds as my comparison group and find that indeed 65-to-69-year-olds increased labor supply in response to the repeal. Additionally, the young cohort in my sample, 55-to-61-year-olds, decreased labor supply in anticipation of increased labor supply in old age. I find that continuing in employment increases dramatically among 65-year-olds suggesting that the long-term effect will be greater. Therefore, I would expect the effect of the repeal on the labor supply of older workers to increase over time as more workers opt to stay in the labor force until a later age than now as workers re-optimize their labor supply behavior over their lifetime.



When looking at simple means and correlations between labor supply and Social Security benefits receipt, I find that, after the repeal, more 65-to-69-year-olds receive benefits while working. In addition, the negative correlation between Social Security income receipt and labor supply disappears. This result suggests that the increase in labor supply was primarily driven by men who were going to claim benefits by the normal retirement age regardless of any changes to the law, and the result is not because men are working longer now because of the increase in the DRC.

Furthermore, the share of 65-to-69-year-olds who are full-time workers, as defined by either 20+ hours worked last week or 30+ hours worked last week, increases dramatically for 65-to-69-year-olds after the repeal while it stays relatively constant for younger cohorts. This result suggests that the RET was effectively reducing labor supply of the target group. Previous studies have tried to examine the impact of a repeal of the Retirement Earnings Test on the labor supply of the elderly. However, those models relied on both explicit functional forms and out-of-sample predictions to get their results in the absence of the repeal. With data after the repeal now available, I find that, in fact, there are large labor supply effects.

Drawbacks are that without a structural model, I cannot identify the channels through which the reduction in labor supply comes about. Avenues for future research include examining the labor supply of non-whites and women as well. Also, when more data become available, the results should become stronger than they are now, and this should be verifiable. Additionally, one might consider incorporating the changing DRC and RET exempt amounts into the model, especially if a data set with more precise age measures becomes available. Furthermore, when a suitable panel data set is released for

this time span, one might re-examine the issue of entry, exit, and continuation over a longer time frame. Finally, it is worth investigating why the labor supply of the 70+-year-olds also increased during this time period.

In chapter three, I use a spouse's turning 62 and 65 as instruments for the relaxing of liquidity constraints. I find that, even after controlling for own age and various other covariates, there are discontinuities in labor supply when one turns 62 and 65. This result is well-documented in the literature. However, I also find that in both 1970 and 1980, the coefficient on the interaction term for wife's and husband's being at least age 62 is positive and significant in the wife's labor supply equation. This result suggests that there might be income pooling or complementarities in leisure. Husband's and wife's being at least 62 means both are eligible for Social Security old-age and/or spouse's benefits, relaxing the liquidity constraint the couple might have faced earlier. In addition, even if only one spouse were driven to retire by his or her turning 62, the other might join him or her in retirement because they might enjoy spending time together. In 1970, I also find a marginally significant coefficient for husband's being at least 65 on wife's labor force status, as she is more likely to be out of the labor force if her husband is at least 65. For the husband, I find that his wife's being at least 65 increases his probability of being out of the labor force. However, the interaction term for both being 65 or older is statistically significant and suggests that he works more when both are at least 65. In 1980, I find that the wife's being at least 62 years old reduces the husband's labor supply significantly. For the wife's labor supply, I find that the interaction terms for both 62 years and over and 65 years and over are statistically significant.

To test whether there are complementarities in leisure, I introduce coverage by age discrimination laws as an instrument for wages. I confirm Neumark and Stock's (1999) results who find that coverage by age discrimination laws increases the probability of employment. However, I find that spousal coverage by those laws has no effect on one's labor supply. This result suggests that complementarities in leisure are not important in the joint labor supply decision. Previous studies have often cited complementarities in leisure as a possibly major component in the incidence of joint retirement of spouses. However, for the sample I examined, complementarities do not seem to play a major role.

This dissertation provides more evidence that Social Security rules affect the labor supply behavior of the elderly greatly. While other studies had attempted to examine the effect of the retirement earnings test before, the data to make this analysis possible did not really exist. Thus, previous studies had to rely on out-of-sample predictions and strict functional form assumptions to predict what the impact would be. My second chapter fills that void in the literature. With three years of data after the repeal, I can look at the data itself to analyze the impact of the repeal on labor supply of the elderly.

The third chapter examines the issue of joint retirement in an historical context. I find that indeed, husbands and wives do seem to make joint labor supply decisions even accounting for unobservable characteristics and preferences. In addition, I find that complementarities in leisure seem to be less of a factor in driving this decision. Rather, income pooling seems a more plausible explanation. Identification strategies of previous studies examining joint retirement did not allow the authors to separate income pooling

from complementarities in leisure. Using age discrimination laws as instruments allows me to distinguish between complementarities in leisure and income pooling.

I find that Social Security rules affect the labor supply of the elderly. Major changes such as the repeal of the Retirement Earnings Test have great implications not only for beneficiaries but also for those near retirement age. In addition, the issue of joint retirement should be taken into account when proposing Social Security rules changes, possibly due to income pooling within the family, less so because of complementarities in leisure. As Social Security will be reformed in the coming years, policymakers should be mindful of the dynamic nature of changing the rules to avoid unintended negative consequences of changes in the law.

## APPENDIX

Table 2.1.

Social Security Benefits Conditional on Age at Which They Were First Claimed

		Claim At Age	62	65	70
Die At Age	DRC				
80	3%		1,440	1,500	1,150
85	3%		1,840	2,000	1,725
90	3%		2,240	2,500	2,300
80	5%		1,440	1,500	1,250
85	5%		1,840	2,000	1,875
90	5%		2,240	2,500	2,500
80	6.5%		1,440	1,500	1,325
85	6.5%		1,840	2,000	1,987.5
90	6.5%		2,240	2,500	2,650

Table 2.2.

Expected Value of the Stream of Social Security Benefits for White Males,  
 Conditional on When Benefits Were First Claimed<sup>17</sup>

Claiming Age	Index of Value of Stream of Social Security Benefits
62	1421.8
63	1454.9
64	1476.4
65	1486.7
66	1463.1
67	1432.4
68	1395.1
69	1351.6
70	1302.6

<sup>17</sup>The life tables from the National Vital Statistics Report, Vol. 50, No. 6, Mar 21, 2002 were used to compute the expected value of the stream of SS benefits. The DRC is assumed to be five percent per year. The amounts above are relative to the annual Primary Insurance Amount (PIA), the benefit amount a worker receives annually if he starts claiming at the NRA, which was 65 at the time. A value of 100 corresponds to the annual PIA. In other words, a white male claiming at age 62 can expect to receive 14.218 times the amount of his annual PIA over his lifetime. I implicitly assume that one can either receive the full amount of benefits or none each year, but not partial benefits that would be received if one had some of the benefits taxed away. Also, full annual benefits are claimed at the end of the year (for simplicity), and benefits can only be claimed at exact ages in years. There is no discounting or inflation here.

Table 2.3.

## Income, Substitution, and Overall Effects on Labor Supply for Each Age Group

Ages	Income Effect <sup>18</sup>	Substitution Effect	Overall Effect
55-61	Negative	Negative	Negative
62-64 <sup>19</sup>	Negative	Ambiguous	Ambiguous
65-69	Negative	Positive	Positive
70+	None	None	None

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<sup>18</sup> Note that the income effect comes into play only if the worker was already earning more than the exempt amount. It will only reduce hours of workers already working a lot of hours, not push part-time workers out of the labor force or reduce hours of workers who were below the exempt amount anyway.

<sup>19</sup> The effect for 62-to-64 is unclear because some of the workers in this group also experience a reduction in the marginal tax rate on excess income from 50 percent to 33 percent and a substantial increase in the exempt amount.

Table 2.4.1.  
Variables Used in Regressions

Variable	Definition
Hours	Hours worked last week at all jobs
Work	Respondent at work or not
Education	Dummy variables for each level of education attainment in the CPS
Year	Year of survey
Month	Month of survey
Age	Dummy variables for each age
Target Group	Dummy variable for treatment group: age $\geq$ 65 and age $\leq$ 69 and year $\geq$ 2000
Young Group	Dummy variable for treatment group: age $\leq$ 61 and year $\geq$ 2000
Intermediate Group	Dummy variable for treatment group: age $\geq$ 62 and age $\leq$ 64 and year $\geq$ 2000

Table 2.4.2.  
Values for Education Variable

Value	Educational Attainment
31	Less Than 1st Grade
32	1st, 2nd, 3 <sup>rd</sup> , or 4th Grade
33	5th or 6th Grade
34	7th or 8th Grade
35	9th Grade
36	10th Grade
37	11th Grade
38	12th Grade, No Diploma
39	High School Grad-Diploma or Equiv. (GED)
40	Some College But No Degree
41	Associate Degree-Occupational/Vocational
42	Associate Deg.-Academic Program
43	Bachelor's Degree(ex.: BA, AB, BS)
44	Master's Degree(ex.: MA, MS, MEng, MEd, MSW)
45	Professional School Deg(ex.: MD, DDS, DVM)
46	Doctorate Degree(ex.: PhD, EdD)



Table 2.5.1.

Means for Work and Hours Variables in ORG 1, January 1997-December 2002

Sample	Obs.	Full Sample	
		Fraction Work	Avg. Hours
Young (55-61)	28,747	0.7790 (0.4149)	33.79 (21.83)
Intermediate (62-64)	9,885	0.4906 (0.4999)	19.37 (22.67)
Target (65-69)	15,864	0.2924 (0.4549)	9.88 (17.94)
65	3,371	0.3551 (0.4786)	12.68 (19.81)
66	3,313	0.3148 (0.4645)	10.87 (18.73)
67	3,061	0.2806 (0.4494)	9.32 (17.47)
68	3,146	0.2727 (0.4454)	8.82 (16.86)
69	2,973	0.2294 (0.4205)	7.32 (15.79)
Old (70-75)	16,815	0.1688 (0.3746)	5.19 (13.57)
All	71,311	0.4869 (0.4998)	19.73 (23.01)

Table 2.5.2.

## Means for Work and Hours Variables in ORG 1,

## Pre-Repeal and Post-Repeal Observations

Sample	Pre-Repeal Observations			Post-Repeal Observations		
	Obs.	Fraction Work	Avg. Hours	Obs.	Fraction Work	Avg. Hours
Young (55-61)	13,238	0.7807 (0.4138)	34.05 (21.94)	15,509	0.7775 (0.4159)	33.56 (21.73)
Intermediate (62-64)	4,728	0.4784 (0.4996)	18.91 (22.69)	5,157	0.5018 (0.5000)	19.78 (22.65)
Target (65-69)	7,768	0.2751 (0.4466)	9.09 (17.33)	8,096	0.3090 (0.4621)	10.65 (18.48)
65	1,561	0.3203 (0.4667)	11.19 (19.04)	1,810	0.3851 (0.4867)	13.97 (20.36)
66	1,621	0.2992 (0.4580)	9.90 (17.94)	1,692	0.3298 (0.4703)	11.80 (19.41)
67	1,509	0.2717 (0.4450)	9.02 (17.34)	1,552	0.2893 (0.4536)	9.61 (17.60)
68	1,617	0.2548 (0.4359)	8.17 (16.46)	1,529	0.2917 (0.4547)	9.51 (17.24)
69	1,460	0.2260 (0.4184)	7.03 (15.23)	1,513	0.2327 (0.4227)	7.60 (16.31)
Old (70-75)	8,349	0.1636 (0.3699)	4.82 (12.89)	8,466	0.1739 (0.3790)	5.55 (14.21)

Table 2.6.

Means of ORG 1 Variables of Individuals who Leave the Sample Before

ORG 2 Interview and of Individuals who Stay in the Sample

	Observations	Work	HS Diploma or more	Married	Hours
Leavers	9,077	0.4554 (0.4980)	0.7928 (0.4053)	0.7098 (0.4539)	18.63 (22.83)
Stayers	59,127	0.4818 (0.4997)	0.8109 (0.3916)	0.8233 (0.3814)	19.47 (23.01)
Young (55-61)	3,213	0.7482 (0.4341)	0.8534 (0.3538)	0.6972 (0.4596)	32.56 (22.43)
Leavers					
Young (55-61)	20,090	0.7997 (0.4002)	0.8790 (0.3261)	0.8291 (0.3765)	34.97 (21.50)
Stayers					
Intermediate (62-64)	1,324	0.5544 (0.4972)	0.8369 (0.3696)	0.7311 (0.4435)	22.61 (23.00)
Leavers					
Intermediate (62-64)	8,430	0.5777 (0.4940)	0.8422 (0.3645)	0.8434 (0.3634)	23.58 (23.40)
Stayers					
Target (65-69)	1,963	0.2909 (0.4543)	0.7718 (0.4198)	0.7346 (0.4417)	10.65 (18.79)
Leavers					
Target (65-69)	13,634	0.3297 (0.4701)	0.7975 (0.4019)	0.8281 (0.3773)	11.44 (19.03)
Stayers					
Old (70-75)	2,577	0.1649 (0.3712)	0.7105 (0.4536)	0.6958 (0.4602)	5.30 (13.17)
Leavers					
Old (70-75)	16,973	0.1801 (0.3843)	0.7256 (0.4462)	0.8027 (0.3979)	5.55 (13.99)
Stayers					

Table 2.7.

Testing the Differences in the Means for the four Age Groups,  
before the Repeal and after the Repeal

Sample	Dependent Variable	Obs.	Difference in Means	P-Value
Young (55-61)	Employment	28,747	-0.0032	0.512
Intermediate (62-64)	Employment	9,885	0.0234	0.020
Target (65-69)	Employment	15,864	0.0339	0.000
Old (70-75)	Employment	16,815	0.0103	0.076
Young (55-61)	Hours	28,747	-0.4834	0.061
Intermediate (62-64)	Hours	9,885	0.8691	0.057
Target (65-69)	Hours	15,864	1.5612	0.000
Old (70-75)	Hours	16,815	0.7250	0.001

Table 2.8.1.

Probit Estimates of Work Equation, Sample 1B, ORG 1,

January 1997-December 2002, Age Covariates

	Education Group	All	HS & Less	BA & More
Marginal Effect (Standard Error) on Dummy Variable	Age 56	-0.0242 (0.0103)	-0.0019 (0.0154)	-0.0316 (0.0185)
	Age 57	-0.0438 (0.0103)	-0.0318 (0.0151)	-0.0568 (0.0189)
	Age 58	-0.0524 (0.0102)	-0.0617 (0.0143)	-0.0433 (0.0193)
	Age 59	-0.0916 (0.0099)	-0.0750 (0.0140)	-0.1064 (0.0189)
	Age 60	-0.1372 (0.0094)	-0.1180 (0.0129)	-0.1613 (0.0184)
	Age 61	-0.1491 (0.0092)	-0.1508 (0.0124)	-0.1469 (0.01840)
	Age 62	-0.2469 (0.0089)	-0.2285 (0.0116)	-0.2781 (0.01870)
	Age 63	-0.3047 (0.0081)	-0.2892 (0.0101)	-0.3092 (0.0185)
	Age 64	-0.3181 (0.0079)	-0.2986 (0.0098)	-0.3241 (0.0184)
	Age 65	-0.3715 (0.0065)	-0.3495 (0.0077)	-0.3992 (0.01490)
	Age 66	-0.3891 (0.0061)	-0.3626 (0.0073)	-0.4138 (0.0146)
	Age 67	-0.4040 (0.0057)	-0.3798 (0.0066)	-0.4259 (0.0143)
	Age 68	-0.4055 (0.0057)	-0.3712 (0.0069)	-0.4534 (0.0132)
	Age 69	-0.4293 (0.0050)	-0.3949 (0.00590)	-0.4625 (0.0128)
	Age 70	-0.4335 (0.0050)	-0.3973 (0.0060)	-0.4859 (0.0118)
	Age 71	-0.4463 (0.0046)	-0.4142 (0.0052)	-0.4861 (0.0118)
	Age 72	-0.4398 (0.0048)	-0.4072 (0.0055)	-0.4812 (0.0122)
	Age 73	-0.4567 (0.0042)	-0.4200 (0.0049)	-0.5076 (0.0106)
	Age 74	-0.4594 (0.0042)	-0.4186 (0.00500)	-0.5222 (0.0100)
	Age 75	-0.4637 (0.0040)	-0.4216 (0.00490)	-0.5249 (0.0097)
	N	71,311	36,352	20,125

Robust standard errors used.

Table 2.8.2.

Probit Estimates of Work Equation, Sample 1B, ORG 1,

January 1997-December 2002, Education Covariates

	Education Group	All	HS & Less	BA & More
Marginal Effect (Standard Error) on Dummy Variable	Education=32	-0.0102 (0.0357)	-0.0110 (0.0351)	N/A
	Education=33	0.0014 (0.0339)	0.0011 (0.0334)	N/A
	Education=34	-0.0067 (0.0324)	-0.0060 (0.0319)	N/A
	Education=35	-0.0054 (0.0333)	-0.0060 (0.0328)	N/A
	Education=36	-0.0227 (0.0326)	-0.0228 (0.0319)	N/A
	Education=37	0.0122 (0.0332)	0.0119 (0.0329)	N/A
	Education=38	-0.0023 (0.0357)	-0.0049 (0.0351)	N/A
	Education=39	0.0176 (0.0315)	0.0143 (0.0312)	N/A
	Education=40	0.0342 (0.0317)	N/A	N/A
	Education=41	0.0278 (0.0331)	N/A	N/A
	Education=42	0.0222 (0.0336)	N/A	N/A
	Education=43	0.0651 (0.0318)	N/A	-6.3702 (0.5693)
	Education=44	0.0534 (0.0322)	N/A	-7.6324 (0.6117)
	Education=45	0.1954 (0.0310)	N/A	Omitted
	Education=46	0.1772 (0.0315)	N/A	-0.0182 (0.7485)
	N	71,311	36,352	20,125

Robust standard errors used.

Table 2.8.3.

Probit Estimates of Work Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Month and Year Covariates

	Education Group	All	HS & Less	BA & More	
Marginal Effect (Standard Error) on Dummy Variable	Year=1997	Omitted	0.0014 (0.0085)	Omitted	
	Year=1998	0.0020 (0.0062)	0.0000 (0.0084)	0.0007 (0.0122)	
	Year=1999	0.0041 (0.0061)	Omitted	0.0046 (0.0119)	
	Year=2000	0.0112 (0.0082)	0.0149 (0.0127)	0.0218 (0.0183)	
	Year=2001	0.0067 (0.0081)	0.0038 (0.0126)	0.0291 (0.0182)	
	Year=2002	0.0010 (0.0080)	-0.0004 (0.0124)	0.0158 (0.0180)	
	Month =February	0.0088 (0.0085)	0.0099 (0.0118)	0.0031 (0.0164)	
	Month =March	-0.0032 (0.0087)	0.0066 (0.01210)	-0.0121 (0.0166)	
	Month =April	0.0154 (0.0087)	0.0241 (0.0120)	-0.0031 (0.0167)	
	Month =May	0.0091 (0.0086)	0.0057 (0.0119)	0.0026 (0.0164)	
	Month =June	0.0041 (0.0086)	0.0026 (0.0119)	-0.0236 (0.0165)	
	Month =July	-0.0104 (0.0087)	-0.0021 (0.0121)	-0.0592 (0.0170)	
	Month =August	-0.0046 (0.0087)	0.0114 (0.0119)	-0.0373 (0.0167)	
	Month =September	0.0118 (0.0086)	0.0090 (0.0120)	-0.0029 (0.0160)	
	Month =October	0.0238 (0.0086)	0.0271 (0.0120)	0.0108 (0.0160)	
	Month =November	0.0110 (0.0085)	0.0127 (0.0118)	-0.0089 (0.0162)	
	Month =December	0.0143 (0.0086)	-0.0004 (0.0118)	0.0134 (0.0162)	
	N		71,311	36,352	20,125

Robust standard errors used.

Table 2.8.4.

Probit Estimates of Work Equation, Sample 1B, ORG 1,

January 1997-December 2002, Repeal Covariates

	Education Group	All	HS & Less	BA & More
Marginal Effect (Standard Error) on Repeal of RET Variable	Target Group	0.0221 (0.0106)	0.0331 (0.0143)	-0.0101 (0.0209)
	Young Group	-0.0160 (0.0095)	-0.0036 (0.0132)	-0.0400 (0.0180)
	Intermediate Group	0.0078 (0.0117)	-0.0040 (0.0157)	-0.0059 (0.0230)
N		71,311	36,352	20,125

Robust standard errors used.

Table 2.8.5.

Probit Estimates of Work Equation, Sample 1B, ORG 1, January 1997-December 2002,

Repeal Covariates, With Age Group 65-69 Broken Up Into Individual Age Groups

	Education Group	All	HS & Less	BA & More
Marginal Effect (Standard Error) on Repeal of RET Variable	Age 65	0.0491 (0.0170)	0.0593 (0.0237)	0.0086 (0.0325)
	Age 66	0.0092 (0.0174)	-0.0037 (0.0236)	0.0198 (0.0337)
	Age 67	0.0244 (0.0182)	0.0455 (0.0249)	-0.0272 (0.0356)
	Age 68	0.0284 (0.0182)	0.0525 (0.0244)	-0.0355 (0.0371)
	Age 69	-0.0062 (0.0189)	0.0092 (0.0256)	-0.0276 (0.0369)
	Young Group	-0.0160 (0.0095)	-0.0036 (0.0132)	-0.0400 (0.0180)
	Intermediate Group	0.0078 (0.0117)	-0.004 (0.0157)	-0.0060 (0.0230)
N		71,311	36,352	20,125

Robust standard errors used.



Table 2.9.1.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1,

January 1997-December 2002, Age Covariates

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Dummy Variable	Age 56	-0.0202 (0.0086)	-0.0016 (0.0135)	-0.0241 (0.0138)
	Age 57	-0.0374 (0.0090)	-0.0294 (0.0142)	-0.0442 (0.0148)
	Age 58	-0.0461 (0.0092)	-0.0606 (0.0146)	-0.0342 (0.0148)
	Age 59	-0.0860 (0.0097)	-0.0758 (0.0148)	-0.0898 (0.0164)
	Age 60	-0.1390 (0.0102)	-0.1294 (0.0153)	-0.1456 (0.0177)
	Age 61	-0.1544 (0.0104)	-0.1758 (0.0160)	-0.1306 (0.0172)
	Age 62	-0.2942 (0.0126)	-0.3037 (0.0185)	-0.2864 (0.0226)
	Age 63	-0.3904 (0.0129)	-0.4176 (0.0186)	-0.3282 (0.0238)
	Age 64	-0.4142 (0.0130)	-0.4362 (0.0185)	-0.3504 (0.0244)
	Age 65	-0.5115 (0.0118)	-0.5425 (0.0166)	-0.4656 (0.0221)
	Age 66	-0.5458 (0.0116)	-0.5717 (0.0164)	-0.4895 (0.0226)
	Age 67	-0.5779 (0.0116)	-0.6124 (0.0160)	-0.5110 (0.0230)
	Age 68	-0.5806 (0.0115)	-0.5902 (0.0160)	-0.5601 (0.0228)
	Age 69	-0.6326 (0.0113)	-0.6517 (0.0156)	-0.5766 (0.0228)
	Age 70	-0.6367 (0.0108)	-0.6491 (0.0148)	-0.6177 (0.0218)
	Age 71	-0.6657 (0.0104)	-0.6883 (0.0141)	-0.6224 (0.0221)
	Age 72	-0.6545 (0.0108)	-0.6762 (0.0146)	-0.6140 (0.0229)
	Age 73	-0.6922 (0.0104)	-0.7064 (0.0141)	-0.6659 (0.0217)
	Age 74	-0.6989 (0.0103)	-0.7028 (0.0141)	-0.7004 (0.0222)
	Age 75	-0.7087 (0.0102)	-0.7097 (0.0140)	-0.7078 (0.0217)
N		71,311	36,352	20,125

Robust standard errors used.

Table 2.9.2.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1,

January 1997-December 2002, Education Covariates

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Dummy Variable	Education=32	-0.0090 (0.0340)	-0.0092 (0.0339)	N/A
	Education=33	0.0011 (0.0324)	0.0016 (0.0323)	N/A
	Education=34	-0.0083 (0.0309)	-0.0069 (0.0308)	N/A
	Education=35	-0.0072 (0.0319)	-0.0069 (0.0318)	N/A
	Education=36	-0.0242 (0.0312)	-0.0238 (0.0311)	N/A
	Education=37	0.0100 (0.0318)	0.0106 (0.0317)	N/A
	Education=38	-0.0035 (0.0343)	-0.0053 (0.0342)	N/A
	Education=39	0.0148 (0.0302)	0.0127 (0.0301)	N/A
	Education=40	0.0324 (0.0303)	N/A	N/A
	Education=41	0.0251 (0.0317)	N/A	N/A
	Education=42	0.0209 (0.0325)	N/A	N/A
	Education=43	0.0636 (0.0304)	N/A	-0.1335 (0.0115)
	Education=44	0.0513 (0.0308)	N/A	-0.1443 (0.0125)
	Education=45	0.1973 (0.0319)	N/A	Omitted
	Education=46	0.1752 (0.0318)	N/A	-0.0217 (0.0148)
N		71,311	36,352	20,125

Robust standard errors used.

Table 2.9.3.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1,

January 1997-December 2002, Month and Year Covariates

	Education Group	All	HS & Less	BA & More	
Coefficient (Standard Error) on Dummy Variable	Year=1997	Omitted	0.0012 (0.0083)	Omitted	
	Year=1998	0.0020 (0.0062)	0.0003 (0.0083)	0.0011 (0.0122)	
	Year=1999	0.0041 (0.0061)	Omitted	0.0040 (0.0118)	
	Year=2000	0.0112 (0.0082)	0.0130 (0.0105)	0.0212 (0.0180)	
	Year=2001	0.0067 (0.0081)	0.0020 (0.0104)	0.0288 (0.0180)	
	Year=2002	0.0010 (0.0080)	-0.0015 (0.0102)	0.0153 (0.0176)	
	Month =February	0.0088 (0.0085)	0.0113 (0.0117)	0.0045 (0.0163)	
	Month =March	-0.0032 (0.0087)	0.0065 (0.0120)	-0.0109 (0.0165)	
	Month =April	0.0154 (0.0087)	0.0244 (0.0118)	-0.0015 (0.0166)	
	Month =May	0.0091 (0.0086)	0.0057 (0.0119)	0.0039 (0.0162)	
	Month =June	0.0041 (0.0086)	0.0029 (0.0119)	-0.0210 (0.0164)	
	Month =July	-0.0104 (0.0087)	-0.0031 (0.0120)	-0.0578 (0.0170)	
	Month =August	-0.0046 (0.0087)	0.0121 (0.0118)	-0.0353 (0.0166)	
	Month =September	0.0118 (0.0086)	0.0091 (0.0118)	-0.0028 (0.0158)	
	Month =October	0.0238 (0.0086)	0.0278 (0.0119)	0.0111 (0.0159)	
	Month =November	0.0110 (0.0085)	0.0126 (0.0117)	-0.0076 (0.0159)	
	Month =December	0.0143 (0.0086)	0.0003 (0.0117)	0.0142 (0.0161)	
	N		71,311	36,352	20,125

Robust standard errors used.

Table 2.9.4.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Repeal Covariates

Education Group		All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Target Group	0.0268 (0.0101)	0.0378 (0.0130)	-0.0065 (0.0222)
	Young Group	-0.0140 (0.0083)	-0.0017 (0.0114)	-0.0364 (0.0174)
	Intermediate Group	0.0126 (0.0128)	-0.0018 (0.0173)	-0.0022 (0.0259)
N		71,311	36,352	20,125
R-Squared		0.2875	0.2961	0.2513

Robust standard errors used.

Table 2.9.5.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1, January 1997-  
December 2002, Repeal Covariates, With Age Group 65-69 Broken Up  
Into Individual Age Groups

Education Group		All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Age 65	0.0615 (0.0192)	0.0723 (0.0255)	0.0171 (0.0386)
	Age 66	0.0138 (0.0190)	-0.0019 (0.0251)	0.0317 (0.0398)
	Age 67	0.0282 (0.0190)	0.0484 (0.0242)	-0.0270 (0.0407)
	Age 68	0.0322 (0.0190)	0.0586 (0.0249)	-0.0383 (0.0405)
	Age 69	-0.0041 (0.0179)	0.0098 (0.0229)	-0.0265 (0.0397)
	Young Group	-0.0140 (0.0083)	-0.0017 (0.0114)	-0.0364 (0.0174)
	Intermediate Group	0.0127 (0.0128)	-0.0018 (0.0173)	-0.0022 (0.0259)
N		71,311	36,352	20,125
R-Squared		0.2876	0.2963	0.2515

Robust standard errors used.

Table 2.10.1.

Linear Estimates of Hours Equation, Sample 1B, ORG 1,

January 1997-December 2002, Age Covariates

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Dummy Variable	Age 56	-0.8775 (0.4622)	0.3230 (0.7055)	-1.1643 (0.7832)
	Age 57	-1.8967 (0.4848)	-1.5019 (0.75150)	-2.8294 (0.8231)
	Age 58	-2.5736 (0.4854)	-2.9025 (0.73910)	-2.5848 (0.8380)
	Age 59	-4.2089 (0.5063)	-2.7758 (0.7558)	-5.4942 (0.8785)
	Age 60	-6.8598 (0.5285)	-5.8300 (0.7764)	-7.6280 (0.9498)
	Age 61	-8.3565 (0.5235)	-8.8712 (0.7729)	-7.8491 (0.9270)
	Age 62	-14.7809 (0.5652)	-14.3930 (0.8775)	-15.2860 (1.1463)
	Age 63	-19.3350 (0.5619)	-20.1416 (0.8498)	-17.1766 (1.1649)
	Age 64	-20.7584 (0.5510)	-21.3043 (0.8372)	-18.7396 (1.1450)
	Age 65	-25.3517 (0.5282)	-26.0411 (0.7484)	-24.3736 (1.0165)
	Age 66	-27.1265 (0.5123)	-27.5528 (0.7171)	-25.4262 (1.0393)
	Age 67	-28.7065 (0.4984)	-29.3266 (0.6830)	-27.1849 (1.0202)
	Age 68	-28.9059 (0.4923)	-28.6003 (0.6883)	-28.3998 (1.0365)
	Age 69	-30.6866 (0.4857)	-30.6424 (0.6694)	-30.1658 (1.0180)
	Age 70	-30.6835 (0.4635)	-30.3477 (0.6472)	-31.1363 (0.9315)
	Age 71	-31.7212 (0.4575)	-31.3948 (0.6353)	-31.5905 (0.9637)
	Age 72	-31.8432 (0.4489)	-31.3294 (0.6321)	-32.2489 (0.9179)
	Age 73	-32.9695 (0.4389)	-32.5191 (0.6129)	-33.7504 (0.8627)
	Age 74	-32.6086 (0.4617)	-31.5546 (0.6560)	-34.8674 (0.8977)
	Age 75	-33.0521 (0.4500)	-32.0397 (0.6413)	-34.8581 (0.9002)
N		71,311	36,352	20,125

Robust standard errors used.

Table 2.10.2

Linear Estimates of Hours Equation, Sample 1B, ORG 1,

January 1997-December 2002, Education Covariates

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Dummy Variable	Education=32	0.1466 (1.3811)	0.1404 (1.3770)	N/A
	Education=33	0.4441 (1.3211)	0.4577 (1.3173)	N/A
	Education=34	0.6434 (1.2569)	0.6355 (1.2527)	N/A
	Education=35	-0.1296 (1.2972)	-0.1081 (1.2928)	N/A
	Education=36	-0.0385 (1.2767)	-0.0338 (1.2723)	N/A
	Education=37	1.1460 (1.2946)	1.1596 (1.2903)	N/A
	Education=38	0.4063 (1.4103)	0.3727 (1.4059)	N/A
	Education=39	1.6247 (1.2285)	1.6356 (1.2243)	N/A
	Education=40	2.0214 (1.2370)	N/A	N/A
	Education=41	2.0678 (1.3107)	N/A	N/A
	Education=42	1.4459 (1.3372)	N/A	N/A
	Education=43	3.8994 (1.2393)	N/A	-6.3702 (0.5693)
	Education=44	2.6470 (1.2600)	N/A	-7.6324 (0.6117)
	Education=45	10.2570 (1.3318)	N/A	Omitted
	Education=46	10.2528 (1.3322)	N/A	-0.0182 (0.7485)
N		71,311	36,352	20,125

Robust standard errors used.

Table 2.10.3.

Linear Estimates of Hours Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Month and Year Covariates

	Education Group	All	HS & Less	BA & More	
Coefficient (Standard Error) on Dummy Variable	Year=1997	Omitted	-0.0345 (0.3598)	Omitted	
	Year=1998	-0.0987 (0.2715)	-0.1123 (0.3593)	-0.6918 (0.5635)	
	Year=1999	0.2473 (0.2725)	Omitted	0.0145 (0.5570)	
	Year=2000	0.8152 (0.3241)	0.7260 (0.4133)	0.9374 (0.7089)	
	Year=2001	0.5967 (0.3231)	0.4180 (0.4106)	0.8796 (0.7109)	
	Year=2002	0.2777 (0.3197)	0.1110 (0.4042)	0.2893 (0.6956)	
	Month =February	0.6343 (0.3775)	0.6725 (0.4962)	0.7760 (0.7774)	
	Month =March	0.0013 (0.3834)	0.2708 (0.5057)	-0.0486 (0.7945)	
	Month =April	0.8810 (0.3829)	0.9477 (0.4969)	0.6363 (0.7933)	
	Month =May	0.9789 (0.3878)	1.0543 (0.5173)	0.5702 (0.7749)	
	Month =June	0.7844 (0.3922)	1.0470 (0.5279)	-0.6435 (0.7915)	
	Month =July	-0.0274 (0.3871)	0.5123 (0.5154)	-2.2849 (0.7978)	
	Month =August	0.2080 (0.3885)	1.0757 (0.5183)	-1.3662 (0.7845)	
	Month =September	0.5434 (0.3768)	0.3650 (0.5034)	-0.0602 (0.7461)	
	Month =October	1.1025 (0.3796)	0.9945 (0.4992)	1.0970 (0.7626)	
	Month =November	0.4662 (0.3754)	0.4394 (0.4967)	-0.1872 (0.7583)	
	Month =December	0.5616 (0.3790)	-0.0437 (0.4966)	0.6813 (0.7643)	
	N		71,311	36,352	20,125

Robust standard errors used.

Table 2.10.4.

Linear Estimates of Hours Equation, Sample 1B, ORG 1,

January 1997-December 2002, Repeal Covariates

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Target Group	0.9820 (0.3789)	1.5418 (0.4759)	-0.0542 (0.8767)
	Young Group	-1.2444 (0.3601)	-0.9896 (0.5027)	-1.7074 (0.7264)
	Intermediate Group	0.1004 (0.5399)	-0.3361 (0.7250)	-0.4308 (1.1183)
N		71,311	36,352	20,125
R-Squared		0.3126	0.3197	0.2788

Robust standard errors used.

Table 2.10.5.

Linear Estimates of Hours Equation,

Sample 1B, ORG 1, January 1997-December 2002, Repeal Covariates

With Age Group 65-69 Broken Up Into Individual Age Groups

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Age 65	2.3840 (0.7770)	2.6392 (1.0297)	1.4223 (1.6198)
	Age 66	0.9602 (0.7418)	0.4808 (0.9481)	1.6797 (1.7016)
	Age 67	0.6719 (0.6950)	1.8258 (0.8442)	-1.7503 (1.5987)
	Age 68	0.9683 (0.6969)	2.3052 (0.9020)	-1.7692 (1.6598)
	Age 69	-0.2255 (0.6582)	0.3807 (0.7885)	-0.3904 (1.6011)
	Young Group	-1.2444 (0.3601)	-0.9896 (0.5027)	-1.7075 (0.7265)
	Intermediate Group	0.1003 (0.5400)	-0.3363 (0.7250)	-0.4314 (1.1185)
N		71,311	36,352	20,125
R-Squared		0.3127	0.3198	0.2790

Robust standard errors used.



Table 2.11.1.

Probit Estimates of Work and Labor Force Transition Equations,

Sample 2, ORG 2, January 1997-December 2002, Age Covariates

	Dependent Variable	Employment	Entry	Exit	Continuation
Marginal Effect (Standard Error) on Dummy Variable	Age	0.0044	0.0064	-0.0072	-0.0039
	57	(0.0115)	(0.0052)	(0.0064)	(0.0106)
	Age	-0.0529	-0.0030	0.0084	-0.0426
	58	(0.0108)	(0.0047)	(0.0072)	(0.0102)
	Age	-0.0575	-0.0013	-0.0006	-0.0484
	59	(0.0109)	(0.0048)	(0.0069)	(0.0102)
	Age	-0.1006	-0.0029	0.0101	-0.0857
	60	(0.0104)	(0.0048)	(0.0075)	(0.0098)
	Age	-0.1312	0.0077	0.0126	-0.1227
	61	(0.0098)	(0.0056)	(0.0076)	(0.0092)
	Age	-0.1975	-0.0137	0.0774	-0.1683
	62	(0.0099)	(0.0046)	(0.0117)	(0.0096)
	Age	-0.2552	-0.0045	0.0371	-0.2311
	63	(0.0088)	(0.0052)	(0.0099)	(0.0084)
	Age	-0.2912	0.0048	0.0210	-0.2722
	64	(0.0081)	(0.0063)	(0.0093)	(0.0075)
	Age	-0.3402	-0.0107	0.0395	-0.3170
	65	(0.0067)	(0.0047)	(0.0100)	(0.0072)
	Age	-0.3486	-0.0013	0.0038	-0.3252
	66	(0.0065)	(0.0055)	(0.0081)	(0.0069)
	Age	-0.3687	-0.0100	-0.0016	-0.3355
	67	(0.0060)	(0.0047)	(0.0078)	(0.0066)
	Age	-0.3847	-0.0109	0.0033	-0.3527
	68	(0.0055)	(0.0047)	(0.0081)	(0.0060)
	Age	-0.3940	-0.0205	-0.0083	-0.3605
	69	(0.0052)	(0.0038)	(0.0076)	(0.0057)
	Age	-0.3999	-0.0222	-0.0223	-0.3641
70	(0.0052)	(0.0037)	(0.0066)	(0.0049)	
Age	-0.4003	-0.0073	-0.0257	-0.3727	
71	(0.0051)	(0.0051)	(0.0064)	(0.0046)	
Age	-0.4170	-0.0194	-0.0258	-0.3829	
72	(0.0046)	(0.0042)	(0.0065)	(0.0042)	
Age	-0.4231	-0.0302	-0.0174	-0.3833	
73	(0.0043)	(0.0030)	(0.0072)	(0.0042)	
Age	-0.4296	-0.0279	-0.0302	-0.3912	
74	(0.0041)	(0.0033)	(0.0064)	(0.0039)	
Age	-0.4315	-0.0275	-0.0429	-0.3934	
75	(0.0040)	(0.0034)	(0.0055)	(0.0038)	
Age	-0.4307	-0.0312	-0.0411	-0.3907	
76	(0.0041)	(0.0031)	(0.0057)	(0.0039)	
N		59,127	59,127	59,127	59,127

Robust standard errors used.

Table 2.11.2.

Probit Estimates of Work and Labor Force Transition Equations,  
 Sample 2, ORG 2, January 1997-December 2002, Education Covariates

	Dependent Variable	Employment	Entry	Exit	Continuation
Marginal Effect (Standard Error) on Dummy Variable	Education=32	-0.0475 (0.0407)	-0.0242 (0.0195)	0.0345 (0.0268)	-0.0252 (0.0400)
	Education=33	-0.0241 (0.0385)	-0.0169 (0.0181)	0.0216 (0.0255)	-0.0061 (0.0380)
	Education=34	-0.0151 (0.0369)	-0.0088 (0.0172)	0.0177 (0.0246)	-0.0058 (0.0363)
	Education=35	-0.0239 (0.0380)	-0.0172 (0.0179)	0.0139 (0.0253)	-0.0054 (0.0374)
	Education=36	-0.0222 (0.0373)	-0.0096 (0.0175)	0.0026 (0.0250)	-0.0122 (0.0367)
	Education=37	-0.0029 (0.0378)	-0.0102 (0.0177)	0.0230 (0.0251)	0.0087 (0.0372)
	Education=38	0.0068 (0.0407)	0.0070 (0.0188)	0.0004 (0.0274)	-0.0023 (0.0405)
	Education=39	0.0044 (0.0361)	-0.0144 (0.0169)	0.0146 (0.0241)	0.0216 (0.0355)
	Education=40	0.0200 (0.0363)	-0.0130 (0.0170)	0.0182 (0.0242)	0.0358 (0.0356)
	Education=41	-0.0086 (0.0377)	-0.0112 (0.0176)	0.0171 (0.0251)	0.0061 (0.0371)
	Education=42	0.0357 (0.0381)	0.0004 (0.0177)	0.0206 (0.0252)	0.0375 (0.0375)
	Education=43	0.0542 (0.0363)	-0.0124 (0.0170)	0.0134 (0.0242)	0.0695 (0.0356)
	Education=44	0.0363 (0.0367)	-0.0103 (0.0172)	0.0117 (0.0245)	0.0499 (0.0360)
	Education=45	0.2039 (0.0380)	-0.0146 (0.0177)	0.0156 (0.0252)	0.2173 (0.0373)
	Education=46	0.1542 (0.0378)	-0.0183 (0.0178)	0.0297 (0.0250)	0.1726 (0.0371)
	N		59,127	59,127	59,127

Robust standard errors used.

Table 2.11.3.

Probit Estimates of Work and Labor Force Transition Equations,  
Sample 2, ORG 2, January 1997-December 2002, Month and Year Covariates

	Dependent Variable	Employment	Entry	Exit	Continuation	
Marginal Effect (Standard Error) on Dummy Variable	Year=1996	Omitted	Omitted	Omitted	Omitted	
	Year=1997	-0.0076 (0.0066)	-0.0069 (0.0029)	0.0016 (0.0042)	-0.0004 (0.0066)	
	Year=1998	-0.0039 (0.0066)	-0.0080 (0.0028)	-0.0014 (0.0041)	0.0047 (0.0066)	
	Year=1999	-0.0062 (0.0095)	-0.0017 (0.0046)	0.0039 (0.0063)	-0.0041 (0.0098)	
	Year=2000	0.0026 (0.0095)	-0.0057 (0.0045)	0.0002 (0.0062)	0.0085 (0.0098)	
	Year=2001	-0.0032 (0.0095)	-0.0045 (0.0045)	0.0021 (0.0063)	0.0016 (0.0097)	
	Year=2002	N/A	N/A	N/A	N/A	
	Month =February	-0.0054 (0.0097)	-0.0024 (0.0047)	-0.0012 (0.0062)	-0.0029 (0.0097)	
	Month =March	-0.0139 (0.0098)	0.0036 (0.0051)	-0.0012 (0.0063)	-0.0168 (0.0098)	
	Month =April	0.0103 (0.0098)	0.0033 (0.0050)	0.0030 (0.0065)	0.0075 (0.0098)	
	Month =May	0.0156 (0.0097)	0.0058 (0.0051)	-0.0069 (0.0060)	0.0107 (0.0097)	
	Month =June	0.0050 (0.0098)	0.0094 (0.0054)	0.0059 (0.0065)	-0.0030 (0.0097)	
	Month =July	-0.0086 (0.0098)	0.0083 (0.0053)	-0.0010 (0.0062)	-0.0155 (0.0098)	
	Month =August	-0.0166 (0.0097)	0.0148 (0.0057)	0.0165 (0.0069)	-0.0293 (0.0096)	
	Month =September	0.0086 (0.0096)	0.0103 (0.0055)	0.0032 (0.0063)	-0.0006 (0.0096)	
	Month =October	0.0105 (0.0098)	0.0007 (0.0049)	0.0031 (0.0064)	0.0105 (0.0098)	
	Month =November	0.0028 (0.0097)	-0.0009 (0.0048)	0.0050 (0.0064)	0.0042 (0.0097)	
	Month =December	-0.0009 (0.0097)	-0.0012 (0.0047)	0.0093 (0.0066)	0.0008 (0.0097)	
	N		59,127	59,127	59,127	59,127

Robust standard errors used.

Table 2.11.4.

Probit Estimates of Work and Labor Force Transition Equations,  
 Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates

	Dependent Variable	Work	Entry	Exit	Continuation
Marginal Effect (Standard Error) on Repeal of RET Variable	Target	0.0232	0.0009	0.0055	0.0243
	Group	(0.0109)	(0.0055)	(0.0075)	(0.0113)
	Young	-0.0133	-0.0048	0.0085	-0.0074
	Group	(0.0101)	(0.0047)	(0.0069)	(0.0102)
	Intermediate	-0.0017	-0.0020	-0.0080	-0.0001
	Group	(0.0120)	(0.0059)	(0.0071)	(0.0122)
N		59,127	59,127	59,127	59,127

Robust standard errors used.

Table 2.11.5.

Probit Estimates of Work and Labor Force Transition Equations,  
 Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates,  
 With Age Group 65-69 Broken Up Into Individual Age Groups

	Dependent Variable	Work	Entry	Exit	Continuation
Marginal Effect (Standard Error) on Repeal of RET Variable	Age 65	0.0456	0.0003	0.0024	0.0478
		(0.0185)	(0.0094)	(0.0115)	(0.0190)
	Age 66	0.0191	-0.0041	-0.0042	0.0249
		(0.0183)	(0.0081)	(0.0115)	(0.0189)
	Age 67	0.0073	0.0047	0.0001	0.0034
		(0.0186)	(0.0100)	(0.0123)	(0.0191)
	Age 68	0.0137	0.0039	0.0273	0.0118
		(0.0193)	(0.0100)	(0.0156)	(0.0200)
	Age 69	0.0307	0.0011	0.0059	0.0328
		(0.0197)	(0.0100)	(0.0140)	(0.0204)
	Young	-0.0133	-0.0048	0.0085	-0.0074
	Group	(0.0101)	(0.0047)	(0.0070)	(0.0102)
	Intermediate	-0.0017	-0.0020	-0.0080	-0.0001
	Group	(0.0120)	(0.0059)	(0.0071)	(0.0122)
N		59,127	59,127	59,127	59,127

Robust standard errors used.

Table 2.12.1.

Linear Probability Estimates of Work and Labor Force Transition Equations,

Sample 2, ORG 2, January 1997-December 2002, Age Covariates

	Dependent Variable	Employment	Entry	Exit	Continuation
Coefficient	Age	0.0037	0.0076	-0.0075	-0.0040
(Standard Error)	57	(0.0097)	(0.0059)	(0.0068)	(0.0108)
on Dummy	Age	-0.0494	-0.0035	0.0086	-0.0459
Variable	58	(0.0105)	(0.0057)	(0.0073)	(0.0114)
	Age	-0.0546	-0.0012	-0.0007	-0.0534
	59	(0.0107)	(0.0058)	(0.0072)	(0.0116)
	Age	-0.1036	-0.0034	0.0105	-0.1002
	60	(0.0114)	(0.0058)	(0.0076)	(0.0121)
	Age	-0.1430	0.0091	0.0131	-0.1521
	61	(0.0117)	(0.0064)	(0.0076)	(0.0124)
	Age	-0.2384	-0.0167	0.0780	-0.2217
	62	(0.0140)	(0.0065)	(0.0098)	(0.0144)
	Age	-0.3351	-0.0055	0.0375	-0.3296
	63	(0.0140)	(0.0065)	(0.0091)	(0.0144)
	Age	-0.4005	0.0058	0.0215	-0.4063
	64	(0.0142)	(0.0073)	(0.0089)	(0.0143)
	Age	-0.4958	-0.0130	0.0409	-0.4828
	65	(0.0131)	(0.0063)	(0.0090)	(0.0132)
	Age	-0.5125	-0.0014	0.0038	-0.5111
	66	(0.0129)	(0.0067)	(0.0081)	(0.0129)
	Age	-0.5542	-0.0121	-0.0018	-0.5420
	67	(0.0126)	(0.0063)	(0.0080)	(0.0126)
	Age	-0.5902	-0.0133	0.0031	-0.5769
	68	(0.0124)	(0.0063)	(0.0081)	(0.0124)
	Age	-0.6102	-0.0254	-0.0088	-0.5849
	69	(0.0121)	(0.0058)	(0.0080)	(0.0122)
	Age	-0.6207	-0.0270	-0.0215	-0.5937
	70	(0.0118)	(0.0056)	(0.0072)	(0.0119)
	Age	-0.6211	-0.0091	-0.0247	-0.6120
	71	(0.0118)	(0.0064)	(0.0070)	(0.0116)
	Age	-0.6622	-0.0237	-0.0249	-0.6385
	72	(0.0114)	(0.0061)	(0.0072)	(0.0113)
	Age	-0.6790	-0.0371	-0.0169	-0.6419
	73	(0.0111)	(0.0054)	(0.0076)	(0.0113)
	Age	-0.6967	-0.0343	-0.0291	-0.6624
	74	(0.0111)	(0.0055)	(0.0072)	(0.0112)
	Age	-0.6997	-0.0337	-0.0410	-0.6660
	75	(0.0109)	(0.0056)	(0.0069)	(0.0110)
	Age	-0.7018	-0.0384	-0.0394	-0.6634
	76	(0.0112)	(0.0055)	(0.0070)	(0.0113)
	N	59,127	59,127	59,127	59,127

Robust standard errors used.

Table 2.12.2.

Linear Probability Estimates of Work and Labor Force Transition Equations,

Sample 2, ORG 2, January 1997-December 2002, Education Covariates

	Dependent Variable	Employment	Entry	Education	Continuation	
Coefficient (Standard Error) on Dummy Variable	Education=32	-0.0434 (0.0373)	-0.0228 (0.0203)	0.0321 (0.0232)	-0.0206 (0.0348)	
	Education=33	-0.0213 (0.0354)	-0.0158 (0.0199)	0.0184 (0.0207)	-0.0055 (0.0332)	
	Education=34	-0.0144 (0.0338)	-0.0090 (0.0193)	0.0136 (0.0193)	-0.0055 (0.0315)	
	Education=35	-0.0243 (0.0350)	-0.0172 (0.0197)	0.0105 (0.0202)	-0.0071 (0.0327)	
	Education=36	-0.0221 (0.0342)	-0.0094 (0.0196)	0.0000 (0.0195)	-0.0127 (0.0319)	
	Education=37	-0.0043 (0.0348)	-0.0106 (0.0198)	0.0192 (0.0202)	0.0064 (0.0325)	
	Education=38	0.0060 (0.0376)	0.0102 (0.0221)	-0.0020 (0.0216)	-0.0042 (0.0356)	
	Education=39	0.0030 (0.0332)	-0.0144 (0.0190)	0.0104 (0.0188)	0.0174 (0.0308)	
	Education=40	0.0193 (0.0333)	-0.0130 (0.0191)	0.0140 (0.0190)	0.0323 (0.0310)	
	Education=41	-0.0103 (0.0350)	-0.0116 (0.0198)	0.0133 (0.0202)	0.0013 (0.0328)	
	Education=42	0.0354 (0.0355)	0.0022 (0.0203)	0.0170 (0.0205)	0.0332 (0.0334)	
	Education=43	0.0546 (0.0334)	-0.0126 (0.0191)	0.0085 (0.0190)	0.0671 (0.0311)	
	Education=44	0.0360 (0.0338)	-0.0105 (0.0193)	0.0068 (0.0193)	0.0465 (0.0315)	
	Education=45	0.2083 (0.0352)	-0.0154 (0.0197)	0.0104 (0.0202)	0.2238 (0.0331)	
	Education=46	0.1569 (0.0350)	-0.0182 (0.0197)	0.0258 (0.0204)	0.1751 (0.0330)	
	N		59,127	59,127	59,127	59,127

Robust standard errors used.

Table 2.12.3.

Linear Probability Estimates of Work and Labor Force Transition Equations,  
Sample 2, ORG 2, January 1997-December 2002, Month and Year Covariates

	Dependent Variable	Employment	Entry	Exit	Continuation
Coefficient (Standard Error) on Dummy Variable	Year=1996	0.0070 (0.0082)	0.0021 (0.0039)	-0.0032 (0.0049)	0.0049 (0.0078)
	Year=1997	-0.0003 (0.0081)	-0.0055 (0.0038)	-0.0020 (0.0050)	0.0052 (0.0078)
	Year=1998	0.0035 (0.0082)	-0.0064 (0.0037)	-0.0049 (0.0049)	0.0098 (0.0078)
	Year=1999	Omitted	Omitted	Omitted	Omitted
	Year=2000	0.0085 (0.0068)	-0.0041 (0.0032)	-0.0036 (0.0044)	0.0126 (0.0068)
	Year=2001	0.0031 (0.0067)	-0.0029 (0.0033)	-0.0017 (0.0044)	0.0060 (0.0067)
	Year=2002	N/A	N/A	N/A	N/A
	Month =February	-0.0047 (0.0097)	-0.0022 (0.0044)	-0.0009 (0.0061)	-0.0025 (0.0097)
	Month =March	-0.0132 (0.0098)	0.0034 (0.0047)	-0.0014 (0.0062)	-0.0165 (0.0098)
	Month =April	0.0100 (0.0098)	0.0029 (0.0046)	0.0032 (0.0063)	0.0071 (0.0098)
	Month =May	0.0158 (0.0097)	0.0052 (0.0046)	-0.0073 (0.0059)	0.0106 (0.0097)
	Month =June	0.0050 (0.0098)	0.0087 (0.0047)	0.0051 (0.0061)	-0.0037 (0.0098)
	Month =July	-0.0089 (0.0098)	0.0074 (0.0047)	-0.0011 (0.0061)	-0.0164 (0.0098)
	Month =August	-0.0161 (0.0098)	0.0138 (0.0048)	0.0165 (0.0064)	-0.0299 (0.0097)
	Month =September	0.0093 (0.0096)	0.0097 (0.0048)	0.0034 (0.0061)	-0.0004 (0.0096)
	Month =October	0.0107 (0.0097)	0.0005 (0.0046)	0.0028 (0.0061)	0.0103 (0.0097)
	Month =November	0.0030 (0.0097)	-0.0011 (0.0045)	0.0049 (0.0061)	0.0041 (0.0097)
	Month =December	-0.0003 (0.0098)	-0.0013 (0.0045)	0.0091 (0.0063)	0.0010 (0.0098)
	N	59,127	59,127	59,127	59,127

Robust standard errors used.

Table 2.12.4.

Linear Probability Estimates of Work and Labor Force Transition Equations,

Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates

	Dependent Variable	Work	Entry	Exit	Continuation
Coefficient	Target	0.0270	0.0011	0.0063	0.0260
(Standard Error)	Group	(0.0105)	(0.0047)	(0.0064)	(0.0098)
on Repeal of RET	Young	-0.0125	-0.0056	0.0094	-0.0069
Variable	Group	(0.0090)	(0.0045)	(0.0056)	(0.0090)
	Intermediate	-0.0011	-0.0019	-0.0102	0.0008
	Group	(0.0134)	(0.0059)	(0.0084)	(0.0131)
N		59,127	59,127	59,127	59,127
R-Squared		0.2886	0.0056	0.0108	0.2757

Robust standard errors used.

Table 2.12.5.

Linear Probability Estimates of Work and Labor Force Transition Equations,

Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates,

With Age Group 65-69 Broken Up Into Individual Age Groups

	Dependent Variable	Work	Entry	Exit	Continuation
Coefficient	Age 65	0.0561	0.0004	0.0046	0.0557
(Standard Error)		(0.0210)	(0.0090)	(0.0144)	(0.0201)
on Repeal of RET	Age 66	0.0239	-0.0045	-0.0042	0.0285
Variable		(0.0205)	(0.0100)	(0.0119)	(0.0194)
	Age 67	0.0102	0.0049	0.0007	0.0053
		(0.0199)	(0.0091)	(0.0117)	(0.0188)
	Age 68	0.0151	0.0040	0.0256	0.0111
		(0.0193)	(0.0090)	(0.0126)	(0.0181)
	Age 69	0.0305	0.0007	0.0055	0.0298
		(0.0189)	(0.0074)	(0.0119)	(0.0179)
	Young	-0.0125	-0.0056	0.0094	-0.0069
	Group	(0.0090)	(0.0045)	(0.0056)	(0.0090)
	Intermediate	-0.0011	-0.0019	-0.0102	0.0008
	Group	(0.0134)	(0.0059)	(0.0084)	(0.0131)
N		59,127	59,127	59,127	59,127
R-Squared		0.2887	0.0056	0.0109	0.2758

Robust standard errors used.



Table 2.13.

Test whether Certain Coefficients are Different for the High-Educated from the Low-Educated, Allowing all Coefficients to Differ

Dependent Variable	Coefficients Tested	P-Value	Model
Employment	All Except Education	0.000	Probit
Employment	All Treatment Groups	0.168	Probit
Employment	Target Group	0.081	Probit
Employment	All Except Education	0.000	Linear
Employment	All Treatment Groups	0.195	Linear
Employment	Target Group	0.085	Linear
Hours	All Except Education	0.000	Linear
Hours	All Treatment Groups	0.350	Linear
Hours	Target Group	0.191	Linear

Table 2.14.1.1.

Probit Estimates of Work Equation, Sample 1B, ORG 1, January 1997-December 2002,

Repeal Covariates, Treatment Date is April 2000

	Education Group	All	HS & Less	BA & More
Marginal Effect (Standard Error) on Repeal of RET Variable	Target Group	0.0180 (0.0096)	0.0288 (0.0129)	-0.0113 (0.0188)
	Young Group	-0.0158 (0.0084)	-0.0104 (0.0116)	-0.0375 (0.0158)
	Intermediate Group	0.0089 (0.0107)	-0.0113 (0.0144)	0.0037 (0.0210)
N		71,311	36,352	20,125

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.1.2.

Probit Estimates of Work Equation, Sample 1B, ORG 1, January 1997-December 2002,

Repeal Covariates, With Age Group 65-69 Broken Up Into Individual Age Groups,

Treatment Date is April 2000

	Education Group	All	HS & Less	BA & More	
Marginal Effect (Standard Error) on Repeal of RET Variable	Age 65	0.0356 (0.0164)	0.0375 (0.0226)	0.0134 (0.0311)	
	Age 66	0.0058 (0.0168)	-0.0061 (0.0228)	0.0160 (0.0327)	
	Age 67	0.0256 (0.0176)	0.0481 (0.0241)	-0.0292 (0.0345)	
	Age 68	0.0258 (0.0177)	0.0459 (0.0238)	-0.0338 (0.0361)	
	Age 69	-0.0069 (0.0184)	0.0189 (0.0250)	-0.0346 (0.0356)	
	Young Group	-0.0158 (0.0084)	-0.0105 (0.0116)	-0.0373 (0.0158)	
	Intermediate Group	0.0089 (0.0107)	-0.0114 (0.0144)	0.0039 (0.0210)	
	N		71,311	36,352	20,125

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.2.1.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Repeal Covariates, Treatment Date is April 2000

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Target Group	0.0215 (0.0097)	0.0329 (0.0126)	-0.0097 (0.0205)
	Young Group	-0.0145 (0.0077)	-0.0091 (0.0108)	-0.0342 (0.0151)
	Intermediate Group	0.0137 (0.0124)	-0.0109 (0.0170)	0.0083 (0.0245)
N		71,311	36,352	20,125
R-Squared		0.2875	0.2961	0.2514

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.2.2.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1, January 1997-  
December 2002, Repeal Covariates, With Age Group 65-69 Broken Up Into Individual  
Age Groups, Treatment Date is April 2000

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Age 65	0.0453 (0.0191)	0.0473 (0.0256)	0.0217 (0.0377)
	Age 66	0.0093 (0.0188)	-0.0045 (0.0250)	0.0254 (0.0392)
	Age 67	0.0287 (0.0188)	0.0511 (0.0241)	-0.0313 (0.0399)
	Age 68	0.0290 (0.0189)	0.0518 (0.0251)	-0.0371 (0.0397)
	Age 69	-0.0065 (0.0177)	0.0181 (0.0228)	-0.0373 (0.0385)
	Young Group	-0.0145 (0.0077)	-0.0091 (0.0108)	-0.0339 (0.0151)
	Intermediate Group	0.0137 (0.0124)	-0.0110 (0.0170)	0.0086 (0.0245)
N		71,311	36,352	20,125
R-Squared		0.2876	0.2962	0.2515

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.3.1.

Linear Estimates of Hours Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Repeal Covariates, Treatment Date is April 2000

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Target Group	0.7934 (0.3828)	1.570 (0.4786)	-0.7506 (0.8838)
	Young Group	-1.1830 (0.3571)	-0.9967 (0.4968)	-1.9121 (0.7112)
	Intermediate Group	0.3543 (0.5370)	-0.4448 (0.7226)	-0.0771 (1.1047)
N		71,311	36,352	20,125
R-Squared		0.3126	0.3197	0.2789

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.3.2.

Linear Estimates of Hours Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Repeal Covariates, With Age Group 65-69 Broken  
Up Into Individual Age Groups, Treatment Date is April 2000

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Age 65	2.0679 (0.7818)	2.2835 (1.0393)	1.2966 (1.6240)
	Age 66	0.5666 (0.7494)	0.5774 (0.9578)	0.1436 (1.7260)
	Age 67	0.8034 (0.6989)	2.1189 (0.8555)	-2.2047 (1.5920)
	Age 68	0.8802 (0.7057)	2.1349 (0.9172)	-1.8914 (1.6643)
	Age 69	-0.4849 (0.6602)	0.6785 (0.7970)	-1.5970 (1.5866)
	Young Group	-1.1832 (0.3571)	-0.9988 (0.4969)	-1.9023 (0.7112)
	Intermediate Group	0.3539 (0.5370)	-0.4470 (0.7226)	-0.0675 (1.1047)
N		71,311	36,352	20,125
R-Squared		0.3127	0.3198	0.2791

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.4.1.

Probit Estimates of Work and Labor Force Transition Equations, Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates, Treatment Date is April 2000

	Dependent Variable	Work	Entry	Exit	Continuation
Marginal Effect (Standard Error) on Repeal of RET Variable	Target Group	0.0194 (0.0102)	-0.0003 (0.0051)	0.0009 (0.0067)	0.0212 (0.0105)
	Young Group	-0.0209 (0.0092)	-0.0063 (0.0042)	0.0053 (0.0062)	-0.0139 (0.0092)
	Intermediate Group	-0.0137 (0.0113)	-0.0037 (0.0054)	-0.0068 (0.0067)	-0.0110 (0.0114)
	N	59,127	59,127	59,127	59,127

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.4.2.

Probit Estimates of Work and Labor Force Transition Equations, Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates, With Age Group 65-69 Broken Up Into Individual Age Groups, Treatment Date is April 2000

	Dependent Variable	Work	Entry	Exit	Continuation
Marginal Effect (Standard Error) on Repeal of RET Variable	Age 65	0.0384 (0.0181)	-0.0012 (0.0089)	-0.0008 (0.0109)	0.0415 (0.0185)
	Age 66	0.0259 (0.0179)	-0.0032 (0.0081)	-0.0110 (0.0104)	0.0309 (0.0184)
	Age 67	0.0014 (0.0183)	0.0040 (0.0097)	-0.0024 (0.0117)	-0.0025 (0.0187)
	Age 68	0.0117 (0.0190)	0.0061 (0.0101)	0.0145 (0.0140)	0.0074 (0.0196)
	Age 69	0.0185 (0.0193)	-0.0073 (0.0083)	0.0075 (0.0139)	0.0274 (0.0200)
	Young Group	-0.0209 (0.0092)	-0.0063 (0.0042)	0.0053 (0.0062)	-0.0139 (0.0092)
	Intermediate Group	-0.0137 (0.0113)	-0.0037 (0.0054)	-0.0068 (0.0067)	-0.0110 (0.0114)
	N	59,127	59,127	59,127	59,127

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.5.1.

Linear Probability Estimates of Work and Labor Force Transition Equations,

Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates,

Treatment Date is April 2000

	Dependent Variable	Work	Entry	Exit	Continuation
Coefficient	Target	0.0242	0.0003	0.0021	0.0239
(Standard Error)	Group	(0.0102)	(0.0047)	(0.0064)	(0.0097)
on Repeal of RET	Young	-0.0199	-0.0070	0.0068	-0.0129
Variable	Group	(0.0087)	(0.0044)	(0.0055)	(0.0088)
	Intermediate	-0.0143	-0.0034	-0.0072	-0.0109
	Group	(0.0132)	(0.0059)	(0.0083)	(0.0130)
N		59,127	59,127	59,127	59,127
R-Squared		0.2887	0.0056	0.0108	0.2758

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.14.5.2.

Linear Probability Estimates of Work and Labor Force Transition Equations,  
 Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates, With Age Group  
 65-69 Broken Up Into Individual Age Groups, Treatment Date is April 2000

	Dependent Variable	Work	Entry	Exit	Continuation	
Coefficient (Standard Error) on Repeal of RET Variable	Age 65	0.0496 (0.0209)	-0.0009 (0.0090)	0.0014 (0.0144)	0.0505 (0.0201)	
	Age 66	0.0339 (0.0205)	-0.0027 (0.0100)	-0.0111 (0.0118)	0.0366 (0.0194)	
	Age 67	0.0049 (0.0199)	0.0047 (0.0091)	-0.0014 (0.0117)	0.0002 (0.0188)	
	Age 68	0.0138 (0.0193)	0.0065 (0.0090)	0.0151 (0.0128)	0.0073 (0.0181)	
	Age 69	0.0191 (0.0189)	-0.0061 (0.0073)	0.0074 (0.0120)	0.0252 (0.0180)	
	Young Group	-0.0199 (0.0087)	-0.0070 (0.0044)	0.0068 (0.0055)	-0.0129 (0.0088)	
	Intermediate Group	-0.0143 (0.0132)	-0.0034 (0.0059)	-0.0072 (0.0083)	-0.0109 (0.0130)	
	N		59,127	59,127	59,127	59,127
	R-Squared		0.2887	0.0056	0.0108	0.2759

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.1.1.

Probit Estimates of Work Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Repeal Covariates, Includes Disabled Persons

	Education Group	All	HS & Less	BA & More
Marginal Effect (Standard Error) on Repeal of RET Variable	Target Group	0.0211 (.0110)	0.0289 (0.0149)	-0.0097 (0.0214)
	Young Group	-0.0207 (0.0095)	-0.0177 (0.0129)	-0.0426 (0.0182)
	Intermediate Group	0.0104 (0.0119)	0.0004 (0.0160)	-0.0049 (0.0232)
N		76,864	40,612	20,568

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.1.2.

Probit Estimates of Work Equation, Sample 1B, ORG 1, January 1997-December 2002,  
Repeal Covariates, With Age Group 65-69 Broken Up Into Individual Age Groups,  
Includes Disabled Persons

	Education Group	All	HS & Less	BA & More	
Marginal Effect (Standard Error) on Repeal of RET Variable	Age 65	0.0531 (0.0176)	0.0647 (0.0246)	0.0126 (0.0331)	
	Age 66	0.0086 (0.0180)	-0.0080 (0.0243)	0.0217 (0.0343)	
	Age 67	0.0202 (0.0188)	0.0331 (0.0259)	-0.0260 (0.0362)	
	Age 68	0.0227 (0.0189)	0.0421 (0.0255)	-0.0381 (0.0377)	
	Age 69	-0.0062 (0.0196)	0.0088 (0.0266)	-0.0320 (0.0377)	
	Young Group	-0.0207 (0.0095)	-0.0177 (0.0129)	-0.0426 (0.0182)	
	Intermediate Group	0.0104 (0.0119)	0.0004 (0.0160)	-0.0049 (0.0232)	
	N		76,864	40,612	20,568

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.



Table 2.15.2.1.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Repeal Covariates, Includes Disabled Persons

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Target Group	0.0238 (0.0097)	0.0299 (0.0122)	-0.0057 (0.0220)
	Young Group	-0.0193 (0.0083)	-0.0174 (0.0113)	-0.0391 (0.0174)
	Intermediate Group	0.0157 (0.0121)	0.0044 (0.0159)	-0.0004 (0.0256)
N		76,864	40,612	20,568
R-Squared		0.2415	0.2194	0.2342

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.2.2.

Linear Probability Estimates of Work Equation, Sample 1B, ORG 1, January 1997-  
December 2002, Repeal Covariates, With Age Group 65-69 Broken Up Into Individual  
Age Groups, Includes Disabled Persons

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Age 65	0.0609 (0.0182)	0.0696 (0.0235)	0.0218 (0.0383)
	Age 66	0.0124 (0.0183)	-0.0062 (0.0236)	0.0330 (0.0394)
	Age 67	0.0220 (0.0181)	0.0325 (0.0226)	-0.0246 (0.0403)
	Age 68	0.0237 (0.0182)	0.0427 (0.0235)	-0.0400 (0.0401)
	Age 69	-0.0038 (0.0171)	0.0084 (0.0214)	-0.0304 (0.0395)
	Young Group	-0.0193 (0.0083)	-0.0174 (0.0113)	-0.0391 (0.0174)
	Intermediate Group	0.0157 (0.0121)	0.0044 (0.0159)	-0.0004 (0.0256)
N		76,864	40,612	20,568
R-Squared		0.2416	0.2195	0.2344

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.3.1.

Linear Estimates of Hours Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Repeal Covariates, Includes Disabled Persons

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Target Group	0.8751 (0.3637)	1.2295 (0.4472)	-0.0454 (0.8688)
	Young Group	-1.4525 (0.3550)	-1.5808 (0.4891)	-1.8659 (0.7253)
	Intermediate Group	0.2549 (0.5073)	-0.0719 (0.6606)	-0.3371 (1.1002)
N		76,864	40,612	20,568
R-Squared		0.2676	0.2443	0.2625

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.3.2.

Linear Estimates of Hours Equation, Sample 1B, ORG 1,  
January 1997-December 2002, Repeal Covariates, With Age Group 65-69  
Broken Up Into Individual Age Groups, Includes Disabled Persons

	Education Group	All	HS & Less	BA & More
Coefficient (Standard Error) on Repeal of RET Variable	Age 65	2.3532 (0.7329)	2.4630 (0.9399)	1.6096 (1.6072)
	Age 66	0.8917 (0.7115)	0.2676 (0.8883)	1.7047 (1.6813)
	Age 67	0.5121 (0.6630)	1.2871 (0.7905)	-1.6734 (1.5839)
	Age 68	0.6779 (0.6680)	1.7492 (0.8463)	-1.8337 (1.6416)
	Age 69	-0.2473 (0.6314)	0.2773 (0.7382)	-0.6193 (1.5981)
	Young Group	-1.4524 (0.3550)	-1.5809 (0.4891)	-1.8658 (0.7254)
	Intermediate Group	0.2549 (0.5073)	-0.0721 (0.6606)	-0.3376 (1.1003)
N		76,864	40,612	20,568
R-Squared		0.2677	0.2444	0.2628

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.4.1.

Probit Estimates of Work and Labor Force Transition Equations, Sample 2, ORG 2,  
January 1997-December 2002, Repeal Covariates, Includes Disabled Persons

	Dependent Variable	Work	Entry	Exit	Continuation
Marginal Effect (Standard Error) on Repeal of RET Variable	Target Group	0.0206 (0.0112)	0.0007 (0.0053)	0.0050 (0.0072)	0.0215 (0.0115)
	Young Group	-0.0180 (0.0099)	-0.0051 (0.0044)	0.0053 (0.0065)	-0.0118 (0.0100)
	Intermediate Group	0.0036 (0.0121)	0.0001 (0.0058)	-0.0073 (0.0069)	0.0036 (0.0121)
	N	63,786	63,786	63,786	63,786

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.4.2.

Probit Estimates of Work and Labor Force Transition Equations, Sample 2, ORG 2,  
January 1997-December 2002, Repeal Covariates, With Age Group 65-69 Broken Up  
Into Individual Age Groups, Includes Disabled Persons

	Dependent Variable	Work	Entry	Exit	Continuation
Marginal Effect (Standard Error) on Repeal of RET Variable	Age 65	0.0443 (0.0189)	0.0016 (0.0091)	0.0000 (0.0109)	0.0451 (0.0192)
	Age 66	0.0178 (0.0186)	-0.0042 (0.0077)	-0.0024 (0.0113)	0.0240 (0.0190)
	Age 67	0.0056 (0.0190)	0.0049 (0.0097)	0.0007 (0.0120)	0.0011 (0.0193)
	Age 68	0.0115 (0.0197)	0.0047 (0.0098)	0.0260 (0.0151)	0.0083 (0.0202)
	Age 69	0.0226 (0.0202)	-0.0027 (0.0089)	0.0049 (0.0135)	0.0275 (0.0207)
	Young Group	-0.0180 (0.0099)	-0.0051 (0.0044)	0.0053 (0.0065)	-0.0118 (0.0100)
	Intermediate Group	0.0036 (0.0121)	0.0001 (0.0058)	-0.0074 (0.0069)	0.0036 (0.0121)
	N	63,786	63,786	63,786	63,786

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.5.1.

Linear Probability Estimates of Work and Labor Force Transition Equations,

Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates,

Includes Disabled Persons

	Dependent Variable	Work	Entry	Exit	Continuation
Coefficient	Target	0.0222	0.0008	0.0057	0.0213
(Standard Error)	Group	(0.0100)	(0.0045)	(0.0062)	(0.0094)
on Repeal of RET	Young	-0.0185	-0.0059	0.0064	-0.0126
Variable	Group	(0.0090)	(0.0043)	(0.0054)	(0.0089)
	Intermediate	0.0054	0.0003	-0.0087	0.0051
	Group	(0.0126)	(0.0054)	(0.0078)	(0.0122)
N		63,786	63,786	63,786	63,786
R-Squared		0.2446	0.0051	0.0094	0.2354

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 2.15.5.2.

Linear Probability Estimates of Work and Labor Force Transition Equations,  
 Sample 2, ORG 2, January 1997-December 2002, Repeal Covariates, With Age Group  
 65-69 Broken Up Into Individual Age Groups, Includes Disabled Persons

	Dependent Variable	Work	Entry	Exit	Continuation
Coefficient (Standard Error) on Repeal of RET Variable	Age 65	0.0505 (0.0197)	0.0015 (0.0085)	0.0012 (0.0134)	0.0490 (0.0189)
	Age 66	0.0204 (0.0194)	-0.0047 (0.0096)	-0.0021 (0.0113)	0.0251 (0.0182)
	Age 67	0.0078 (0.0190)	0.0050 (0.0087)	0.0014 (0.0113)	0.0028 (0.0179)
	Age 68	0.0115 (0.0183)	0.0049 (0.0086)	0.0241 (0.0121)	0.0067 (0.0171)
	Age 69	0.0206 (0.0181)	-0.0024 (0.0072)	0.0048 (0.0115)	0.0230 (0.0171)
	Young Group	-0.0185 (0.0090)	-0.0024 (0.0043)	0.0064 (0.0054)	-0.0126 (0.0089)
	Intermediate Group	0.0054 (0.0126)	0.0003 (0.0054)	-0.0087 (0.0078)	0.0051 (0.0122)
N		63,786	63,786	63,786	63,786
R-Squared		0.2447	0.0052	0.0095	0.2354

Unreported covariates include dummy variables for age, educational attainment, month of survey, and year of survey. Robust standard errors used.

Table 3.1.

## Descriptive Statistics, Census Data, 1970 and 1980

	1970		1980	
	Husband	Wife	Husband	Wife
Age 62 or Older	0.4009 (0.4901)	0.2240 (0.4169)	0.3884 (0.4874)	0.2460 (0.4307)
Age 65 or Older	0.2416 (0.4281)	0.1182 (0.3228)	0.2319 (0.4221)	0.1321 (0.3386)
Age	61.06 (4.88)	57.09 (5.95)	60.95 (4.82)	57.85 (5.63)
Fraction out of the Labor Force	0.2031 (0.4023)	0.3444 (0.4752)	0.2259 (0.4182)	0.3607 (0.4802)
Age at First Marriage <sup>20</sup>	25.69 (4.12)	21.88 (4.06)	25.12 (3.88)	21.72 (3.82)
Protected by Age Discrimination Law	0.2549 (0.4358)	0.2860 (0.4519)	0.6895 (0.4627)	0.7297 (0.4441)
Fraction White	0.9398 (0.2379)	0.9398 (0.2378)	0.9380 (0.2412)	0.9373 (0.2424)
SS Income=0	0.7795 (0.4146)	0.8626 (0.3443)	0.7610 (0.4265)	0.8290 (0.3765)
SS Income (cond.)	4,682 (2,276)	3,138 (1,812)	6,128 (2,334)	3,919 (1,891)
<HS Diploma	0.5793 (0.4937)	0.4731 (0.4993)	0.3666 (0.4819)	0.2874 (0.4525)
HS Diploma	0.3196 (0.4663)	0.4383 (0.4962)	0.4528 (0.4978)	0.5885 (0.4921)
BA or Advanced Degree	0.1011 (0.3015)	0.0885 (0.2841)	0.1806 (0.3847)	0.1241 (0.3297)
Alone	0.6365 (0.4810)		0.6144 (0.4867)	
Youngest<6 (cond.)	0.6651 (0.4720)		0.0514 (0.2208)	
Youngest 7-18 (cond.)	0.1583 (0.3651)		0.3510 (0.4773)	
Oldest>70 (cond.)	0.0438 (0.2046)		0.1036 (0.3048)	
N	123,581		132,310	

Standard deviations in parentheses.

<sup>20</sup> Note that observations on age of first marriage in 1970 are available only for the 5% Sample.

Table 3.2.1.

Own Labor Supply Conditional on Spouse's Labor Supply, Census Data, 1970

1970	Wife in Labor Force	Husband in Labor Force
Spouse in Labor Force	0.702	0.850
Spouse out of Labor Force	0.447	0.662
Difference	0.255	0.188
Percentage Change	44.4%	24.9%

Table 3.2.2.

Own Labor Supply Conditional on Spouse's Labor Supply, Census Data, 1980

1980	Wife in Labor Force	Husband in Labor Force
Spouse in Labor Force	0.697	0.845
Spouse out of Labor Force	0.439	0.649
Difference	0.233	0.196
Percentage Change	33.4%	23.2%

Table 3.3.

Values for Education Variable

Values	Educational Attainment
1	Less than High School
2	High School Graduate or Equivalent
3	Some College
4	College Graduate
5	Advanced Degree

Table 3.4.1.

Bivariate Probit, no Age Discrimination Laws, Census Data 1970, other Covariates

1970	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband's Race	0.0421 (0.0229)	0.0343 (0.0160)	0.0461 (0.0227)	0.0371 (0.0158)
Wife's Race	0.0074 (0.0239)	-0.0220 (0.0187)	0.0051 (0.0239)	-0.0208 (0.0186)
Husband's Education=1	-0.0703 (0.0064)	0.0510 (0.0050)	-0.0700 (0.0064)	0.0513 (0.0050)
Husband's Education=2	-0.0583 (0.0063)	0.0219 (0.0055)	-0.0597 (0.0063)	0.0214 (0.0054)
Husband's Education=3	-0.0490 (0.0067)	0.0091 (0.0058)	-0.0514 (0.0067)	0.0074 (0.0057)
Husband's Education=5	0.0109 (0.0083)	-0.0286 (0.0062)	0.0113 (0.0083)	-0.0283 (0.0062)
Wife's Education=1	0.1294 (0.0065)	0.0031 (0.0048)	0.1315 (0.0065)	0.0045 (0.0048)
Wife's Education=2	0.0760 (0.0065)	-0.0130 (0.0047)	0.0767 (0.0065)	-0.0121 (0.0047)
Wife's Education=3	0.0753 (0.0072)	-0.0174 (0.0050)	0.0741 (0.0072)	-0.0171 (0.0049)
Wife's Education=5	-0.0818 (0.0084)	-0.0035 (0.0069)	-0.0830 (0.0084)	-0.0040 (0.0068)
Dummy for no Children	0.0534 (0.0044)	0.0435 (0.0033)	0.0517 (0.0044)	0.0421 (0.0033)
Number of Children	0.0074 (0.0008)	0.0029 (0.0006)	0.0071 (0.0008)	0.0028 (0.0006)
Demeaned Husband's Age	-0.0013 (0.0018)	0.0230 (0.0015)	-0.0016 (0.0018)	0.0228 (0.0015)
Demeaned Husband's Age Squared	0.0000 (0.0002)	0.0004 (0.0001)	0.0000 (0.0002)	0.0004 (0.0001)
Demeaned Husband's Age Cubed	0.0000 (0.0000)	-0.0001 (0.0000)	0.0000 (0.0000)	-0.0001 (0.0000)
Demeaned Husband's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age	0.0128 (0.0006)	0.0020 (0.0005)	0.0129 (0.0006)	0.0021 (0.0005)
Demeaned Wife's Age Squared	0.0008 (0.0001)	0.0002 (0.0001)	0.0007 (0.0001)	0.0002 (0.0001)
Demeaned Wife's Age Cubed	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Dummy for 15% Sample	0.0354 (0.0027)	0.0589 (0.0020)	0.0340 (0.0027)	0.0579 (0.0020)
N		123,581		123,581

Bivariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.



Table 3.4.2.

Bivariate Probit, no Age Discrimination Laws, Census Data 1970, Covariates of Interest

1970	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband 62+	0.0088 (0.0075)	0.0322 (0.0061)	0.0070 (0.0074)	0.0323 (0.0061)
Wife 62+	0.0494 (0.0104)	0.0105 (0.0079)	0.0475 (0.0102)	0.0090 (0.0078)
Husband&Wife 62+	0.0179 (0.0097)	-0.0015 (0.0072)	0.0187 (0.0095)	-0.0005 (0.0072)
Husband 65+	0.0147 (0.0078)	0.1042 (0.0065)	0.0137 (0.0077)	0.1038 (0.0065)
Wife 65+	0.0826 (0.0122)	0.0159 (0.0081)	0.0669 (0.0122)	0.0149 (0.0080)
Husband&Wife 65+	0.0190 (0.0116)	-0.0188 (0.0069)	0.0141 (0.0115)	-0.0178 (0.0069)
Rho	0.2583 (0.0058)		0.2520 (0.0058)	
N	123,581		123,581	

Bivariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.5.1

Univariate Probit, no Age Discrimination Laws, Census Data 1970, other Covariates

1970	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband's Race	0.0427 (0.0229)	0.0374 (0.0159)	0.0465 (0.0227)	0.0399 (0.0156)
Wife's Race	0.0065 (0.0239)	-0.0261 (0.0189)	0.0044 (0.0240)	-0.0246 (0.0188)
Husband's Education=1	-0.0706 (0.0064)	0.0509 (0.0050)	-0.0703 (0.0063)	0.0513 (0.0050)
Husband's Education=2	-0.0584 (0.0063)	0.0218 (0.0055)	-0.0598 (0.0063)	0.0212 (0.0054)
Husband's Education=3	-0.0491 (0.0067)	0.0092 (0.0058)	-0.0515 (0.0067)	0.0075 (0.0057)
Husband's Education=5	0.0108 (0.0083)	-0.0285 (0.0062)	0.0112 (0.0082)	-0.0282 (0.0062)
Wife's Education=1	0.1292 (0.0065)	0.0020 (0.0048)	0.1313 (0.0065)	0.0035 (0.0048)
Wife's Education=2	0.0758 (0.0065)	-0.0137 (0.0047)	0.0766 (0.0065)	-0.0129 (0.0047)
Wife's Education=3	0.0751 (0.0072)	-0.0179 (0.0049)	0.0740 (0.0072)	-0.0177 (0.0049)
Wife's Education=5	-0.0817 (0.0084)	-0.0026 (0.0069)	-0.0829 (0.0084)	-0.0031 (0.0068)
Dummy for no Children	0.0536 (0.0044)	0.0439 (0.0033)	0.0519 (0.0044)	0.0424 (0.0033)
Number of Children	0.0074 (0.0008)	0.0027 (0.0006)	0.0071 (0.0008)	0.0027 (0.0006)
Demeaned Husband's Age	-0.0013 (0.0018)	0.0233 (0.0015)	-0.0015 (0.0018)	0.0231 (0.0015)
Demeaned Husband's Age Squared	-0.0000 (0.0002)	0.0004 (0.0001)	0.0000 (0.0002)	0.0004 (0.0001)
Demeaned Husband's Age Cubed	-0.0000 (0.0000)	-0.0001 (0.0000)	0.0000 (0.0000)	-0.0001 (0.0000)
Demeaned Husband's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age	0.0129 (0.0006)	0.0021 (0.0005)	0.0129 (0.0006)	0.0022 (0.0005)
Demeaned Wife's Age Squared	0.0008 (0.0001)	0.0002 (0.0001)	0.0007 (0.0001)	0.0002 (0.0001)
Demeaned Wife's Age Cubed	-0.0000 (0.0000)	-0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age to the Fourth Power	-0.0000 (0.0000)	-0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Dummy for 15% Sample	0.0351 (0.0027)	0.0587 (0.0020)	0.0027 (0.0027)	0.0576 (0.0020)
N	123,581		123,581	

Univariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.5.2.

Univariate Probit, no Age Discrimination Laws, Census Data 1970, Covariates of Interest

1970	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband 62+	0.0085 (0.0075)	0.0317 (0.0061)	0.0084 (0.0075)	0.0317 (0.0060)
Wife 62+	0.0495 (0.0104)	0.0102 (0.0078)	0.0476 (0.0103)	0.0088 (0.0078)
Husband&Wife 62+	0.0182 (0.0097)	-0.0011 (0.0072)	0.0205 (0.0097)	-0.0002 (0.0072)
Husband 65+	0.0143 (0.0078)	0.1035 (0.0065)	0.0135 (0.0078)	0.1031 (0.0065)
Wife 65+	0.0826 (0.0124)	0.0162 (0.0080)	0.0818 (0.0124)	0.0151 (0.0080)
Husband&Wife 65+	0.0181 (0.0115)	-0.0198 (0.0069)	0.0193 (0.0115)	-0.0185 (0.0069)
N	123,581		123,581	

Univariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.6.1.

Bivariate Probit, no Age Discrimination Laws, Census Data 1980, other Covariates

1980	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband's Race	0.0411 (0.0190)	0.0272 (0.0146)	0.0420 (0.0189)	0.0284 (0.0145)
Wife's Race	0.0159 (0.0194)	-0.0313 (0.0165)	0.0155 (0.0194)	-0.0234 (0.0162)
Husband's Education=1	-0.0550 (0.0051)	0.0634 (0.0045)	-0.0493 (0.0052)	0.0711 (0.0045)
Husband's Education=2	-0.0362 (0.0050)	0.0550 (0.0043)	-0.0336 (0.0050)	0.0591 (0.0043)
Husband's Education=3	-0.0209 (0.0053)	0.0382 (0.0047)	-0.0225 (0.0053)	0.0386 (0.0047)
Husband's Education=4	-0.0020 (0.0059)	0.0226 (0.0051)	-0.0006 (0.0059)	0.0244 (0.0051)
Wife's Education=1	0.1637 (0.0068)	0.0055 (0.0051)	0.1667 (0.0068)	0.0067 (0.0051)
Wife's Education=2	0.1144 (0.0061)	0.0034 (0.0048)	0.1155 (0.0061)	0.0029 (0.0048)
Wife's Education=3	0.0995 (0.0067)	-0.0038 (0.0050)	0.0969 (0.0067)	-0.0054 (0.0049)
Wife's Education=4	0.0886 (0.0077)	0.0029 (0.0057)	0.0893 (0.0077)	0.0038 (0.0057)
Dummy for no Children	0.0239 (0.0051)	0.0246 (0.0039)	0.0219 (0.0051)	0.0245 (0.0039)
Number of Children	0.0011 (0.0007)	-0.0039 (0.0006)	0.0009 (0.0007)	-0.0036 (0.0006)
Demeaned Husband's Age	0.0007 (0.0018)	0.0266 (0.0014)	0.0003 (0.0017)	0.0263 (0.0014)
Demeaned Husband's Age Squared	0.0002 (0.0002)	0.0006 (0.0001)	0.0002 (0.0002)	0.0006 (0.0001)
Demeaned Husband's Age Cubed	0.0000 (0.0000)	-0.0002 (0.0000)	0.0000 (0.0000)	-0.0002 (0.0000)
Demeaned Husband's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age	0.0183 (0.0007)	0.0050 (0.0005)	0.0183 (0.0007)	0.0048 (0.0005)
Demeaned Wife's Age Squared	0.0009 (0.0001)	0.0000 (0.0001)	0.0009 (0.0001)	0.0000 (0.0001)
Demeaned Wife's Age Cubed	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
N	132,310		132,310	

Bivariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.6.2.

Bivariate Probit, no Age Discrimination Laws, Census Data 1980, Covariates of Interest

1980	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband 62+	0.0114 (0.0072)	0.0701 (0.0063)	0.0120 (0.0072)	0.0708 (0.0063)
Wife 62+	0.0660 (0.0104)	0.0210 (0.0078)	0.0643 (0.0103)	0.0203 (0.0078)
Husband&Wife 62+	0.0301 (0.0096)	-0.0060 (0.0069)	0.0312 (0.0096)	-0.0062 (0.0068)
Husband 65+	0.0031 (0.0075)	0.0646 (0.0059)	0.0020 (0.0075)	0.0639 (0.0059)
Wife 65+	0.0320 (0.0114)	0.0045 (0.0074)	0.0288 (0.0113)	0.0027 (0.0073)
Husband&Wife 65+	0.0499 (0.0115)	-0.0063 (0.0069)	0.0525 (0.0115)	-0.0050 (0.0069)
Rho	0.3819 (0.0052)		0.3711 (0.0053)	
N	132,310		132,310	

Bivariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.7.1.

Univariate Probit, no Age Discrimination Laws, Census Data 1980, other Covariates

1980	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband's Race	0.0415 (0.0190)	0.0247 (0.0150)	0.0423 (0.0189)	0.0264 (0.0148)
Wife's Race	0.0159 (0.0195)	-0.0287 (0.0167)	0.0158 (0.0194)	-0.0214 (0.0163)
Husband's Education=1	-0.0551 (0.0051)	0.0638 (0.0045)	-0.0493 (0.0052)	0.0717 (0.0046)
Husband's Education=2	-0.0363 (0.0050)	0.0557 (0.0043)	-0.0337 (0.0050)	0.0597 (0.0044)
Husband's Education=3	-0.0208 (0.0053)	0.0389 (0.0047)	-0.0224 (0.0053)	0.0392 (0.0047)
Husband's Education=4	-0.0022 (0.0059)	0.0219 (0.0051)	-0.0007 (0.0059)	0.0236 (0.0051)
Wife's Education=1	0.1631 (0.0068)	0.0007 (0.0051)	0.1662 (0.0068)	0.0021 (0.0050)
Wife's Education=2	0.1138 (0.0061)	-0.0003 (0.0048)	0.1151 (0.0061)	-0.0007 (0.0048)
Wife's Education=3	0.0988 (0.0067)	-0.0075 (0.0049)	0.0963 (0.0067)	-0.0090 (0.0049)
Wife's Education=4	0.0879 (0.0077)	0.0006 (0.0056)	0.0888 (0.0077)	0.0017 (0.0056)
Dummy for no Children	0.0238 (0.0051)	0.0244 (0.0039)	0.0217 (0.0051)	0.0243 (0.0039)
Number of Children	0.0011 (0.0007)	-0.0042 (0.0006)	0.0008 (0.0007)	-0.0038 (0.0006)
Demeaned Husband's Age	0.0007 (0.0018)	0.0266 (0.0014)	0.0003 (0.0017)	0.0262 (0.0014)
Demeaned Husband's Age Squared	0.0002 (0.0002)	0.0006 (0.0001)	0.0002 (0.0002)	0.0006 (0.0001)
Demeaned Husband's Age Cubed	0.0000 (0.0000)	-0.0002 (0.0000)	0.0000 (0.0000)	-0.0002 (0.0000)
Demeaned Husband's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age	0.0183 (0.0007)	0.0051 (0.0005)	0.0183 (0.0007)	0.0050 (0.0005)
Demeaned Wife's Age Squared	0.0009 (0.0001)	0.0000 (0.0001)	0.0009 (0.0001)	0.0000 (0.0001)
Demeaned Wife's Age Cubed	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
N	132,310		132,310	

Univariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.7.2.

Univariate Probit, no Age Discrimination Laws, Census Data 1980, Covariates of Interest

1980	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband 62+	0.0112 (0.0072)	0.0702 (0.0063)	0.0118 (0.0072)	0.0709 (0.0063)
Wife 62+	0.0664 (0.0103)	0.0216 (0.0078)	0.0646 (0.0103)	0.0209 (0.0078)
Husband&Wife 62+	0.0302 (0.0096)	-0.0072 (0.0068)	0.0315 (0.0096)	-0.0075 (0.0068)
Husband 65+	0.0030 (0.0075)	0.0645 (0.0059)	0.0018 (0.0075)	0.0636 (0.0058)
Wife 65+	0.0314 (0.0114)	0.0053 (0.0074)	0.0281 (0.0113)	0.0035 (0.0073)
Husband&Wife 65+	0.0493 (0.0115)	-0.0082 (0.0068)	0.0521 (0.0115)	-0.0067 (0.0068)
N	132,310		132,310	

Univariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.8.1.

Bivariate Probit, with Age Discrimination Laws, Census Data 1970, other Covariates

1970	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband's Race	0.0417 (0.0229)	0.0341 (0.0160)	0.0459 (0.0227)	0.0372 (0.0158)
Wife's Race	0.0085 (0.0239)	-0.0206 (0.0187)	0.0051 (0.0239)	-0.0210 (0.0186)
Husband's Education=1	-0.0706 (0.0064)	0.0505 (0.0050)	-0.0701 (0.0064)	0.0513 (0.0050)
Husband's Education=2	-0.0583 (0.0063)	0.0217 (0.0054)	-0.0597 (0.0063)	0.0213 (0.0054)
Husband's Education=3	-0.0491 (0.0067)	0.0087 (0.0057)	-0.0514 (0.0067)	0.0074 (0.0057)
Husband's Education=5	0.0109 (0.0083)	-0.0289 (0.0062)	0.0112 (0.0083)	-0.0285 (0.0062)
Wife's Education=1	0.1297 (0.0065)	0.0035 (0.0048)	0.1315 (0.0065)	0.0044 (0.0048)
Wife's Education=2	0.0766 (0.0065)	-0.0123 (0.0047)	0.0768 (0.0065)	-0.0121 (0.0047)
Wife's Education=3	0.0756 (0.0072)	-0.0170 (0.0050)	0.0741 (0.0072)	-0.0171 (0.0049)
Wife's Education=5	-0.0816 (0.0085)	-0.0033 (0.0069)	-0.0829 (0.0084)	-0.0040 (0.0068)
Dummy for no Children	0.0532 (0.0044)	0.0433 (0.0033)	0.0517 (0.0044)	0.0422 (0.0033)
Number of Children	0.0073 (0.0008)	0.0028 (0.0006)	0.0071 (0.0008)	0.0028 (0.0006)
Demeaned Husband's Age	-0.0014 (0.0018)	0.0226 (0.0015)	-0.0015 (0.0018)	0.0224 (0.0015)
Demeaned Husband's Age Squared	0.0000 (0.0002)	0.0004 (0.0001)	0.0000 (0.0002)	0.0004 (0.0001)
Demeaned Husband's Age Cubed	0.0000 (0.0000)	-0.0001 (0.0000)	0.0000 (0.0000)	-0.0001 (0.0000)
Demeaned Husband's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age	0.0128 (0.0006)	0.0021 (0.0005)	0.0128 (0.0006)	0.0021 (0.0005)
Demeaned Wife's Age Squared	0.0008 (0.0001)	0.0002 (0.0001)	0.0007 (0.0001)	0.0002 (0.0001)
Demeaned Wife's Age Cubed	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Dummy for 15% Sample	0.0353 (0.0027)	0.0587 (0.0020)	0.0341 (0.0027)	0.0578 (0.0020)
N	123,581		123,581	

Bivariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.



Table 3.8.2.

Bivariate Probit, with Age Discrimination Laws, Census Data 1970,

## Covariates of Interest

1970	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband 62+	0.0087 (0.0075)	0.0326 (0.0061)	0.0088 (0.0075)	0.0328 (0.0061)
Wife 62+	0.0492 (0.0104)	0.0101 (0.0078)	0.0475 (0.0103)	0.0089 (0.0078)
Husband&Wife 62+	0.0181 (0.0097)	-0.0013 (0.0072)	0.0203 (0.0097)	-0.0005 (0.0072)
Husband 65+	0.0143 (0.0078)	0.1019 (0.0065)	0.0139 (0.0078)	0.1021 (0.0065)
Wife 65+	0.0819 (0.0125)	0.0168 (0.0081)	0.0790 (0.0125)	0.0146 (0.0080)
Husband&Wife 65+	0.0188 (0.0116)	-0.0191 (0.0069)	0.0200 (0.0116)	-0.0179 (0.0069)
Husband Protected	-0.0040 (0.0070)	-0.0243 (0.0045)	0.0010 (0.0074)	-0.0194 (0.0048)
Wife Protected	-0.0066 (0.0066)	0.0060 (0.0042)	-0.0248 (0.0093)	-0.0048 (0.0059)
Rho		0.2579 (0.0058)		0.2519 (0.0058)
N		123,581		123,581

Bivariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.9.1.

Bivariate Probit, with Age Discrimination Laws, Census Data 1980, other Covariates

1980	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband's Race	0.0414 (0.0190)	0.0271 (0.0147)	0.0421 (0.0189)	0.0286 (0.0145)
Wife's Race	0.0157 (0.0194)	-0.0314 (0.0165)	0.0153 (0.0194)	-0.0238 (0.0162)
Husband's Education=1	-0.0551 (0.0051)	0.0637 (0.0045)	-0.0493 (0.0052)	0.0711 (0.0045)
Husband's Education=2	-0.0361 (0.0050)	0.0551 (0.0043)	-0.0336 (0.0050)	0.0592 (0.0043)
Husband's Education=3	-0.0209 (0.0053)	0.0383 (0.0047)	-0.0225 (0.0053)	0.0386 (0.0047)
Husband's Education=4	-0.0021 (0.0059)	0.0228 (0.0051)	-0.0006 (0.0059)	0.0244 (0.0051)
Wife's Education=1	0.1635 (0.0068)	0.0056 (0.0051)	0.1666 (0.0068)	0.0065 (0.0051)
Wife's Education=2	0.1146 (0.0061)	0.0032 (0.0048)	0.1155 (0.0061)	0.0028 (0.0048)
Wife's Education=3	0.0994 (0.0067)	-0.0037 (0.0050)	0.0969 (0.0067)	-0.0056 (0.0049)
Wife's Education=4	0.0885 (0.0077)	0.0029 (0.0057)	0.0893 (0.0077)	0.0037 (0.0057)
Dummy for no Children	0.0239 (0.0051)	0.0248 (0.0039)	0.0219 (0.0051)	0.0245 (0.0039)
Number of Children	0.0011 (0.0007)	-0.0039 (0.0006)	0.0009 (0.0007)	-0.0036 (0.0006)
Demeaned Husband's Age	0.0006 (0.0018)	0.0263 (0.0014)	0.0001 (0.0017)	0.0259 (0.0014)
Demeaned Husband's Age Squared	0.0002 (0.0002)	0.0006 (0.0001)	0.0002 (0.0002)	0.0005 (0.0001)
Demeaned Husband's Age Cubed	0.0000 (0.0000)	-0.0002 (0.0000)	0.0000 (0.0000)	-0.0002 (0.0000)
Demeaned Husband's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age	0.0183 (0.0007)	0.0051 (0.0005)	0.0183 (0.0007)	0.0048 (0.0005)
Demeaned Wife's Age Squared	0.0009 (0.0001)	0.0001 (0.0001)	0.0009 (0.0001)	0.0000 (0.0001)
Demeaned Wife's Age Cubed	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
Demeaned Wife's Age to the Fourth Power	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)	0.0000 (0.0000)
N	132,310		132,310	

Bivariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Table 3.9.2.

Bivariate Probit, with Age Discrimination Laws, Census Data 1980,

## Covariates of Interest

1980	No State Fixed Effects		With State Fixed Effects	
	Wife	Husband	Wife	Husband
Husband 62+	0.0116 (0.0072)	0.0706 (0.0063)	0.0122 (0.0072)	0.0713 (0.0063)
Wife 62+	0.0660 (0.0104)	0.0205 (0.0078)	0.0643 (0.0103)	0.0203 (0.0078)
Husband&Wife 62+	0.0301 (0.0096)	-0.0059 (0.0069)	0.0311 (0.0096)	-0.0063 (0.0068)
Husband 65+	0.0024 (0.0075)	0.0634 (0.0059)	0.0009 (0.0075)	0.0618 (0.0059)
Wife 65+	0.0313 (0.0114)	0.0076 (0.0075)	0.0280 (0.0114)	0.0026 (0.0073)
Husband&Wife 65+	0.0500 (0.0115)	-0.0065 (0.0069)	0.0525 (0.0115)	-0.0051 (0.0069)
Husband Protected	-0.0054 (0.0058)	-0.0103 (0.0038)	-0.0089 (0.0063)	-0.0150 (0.0043)
Wife Protected	-0.0032 (0.0059)	0.0173 (0.0038)	-0.0042 (0.0078)	-0.0007 (0.0050)
Rho		0.3821 (0.0052)		0.3711 (0.0053)
N		132,310		132,310

Bivariate probit marginal effects reported (standard error in parentheses). Dependent variable is a dummy variable that equals 1 if person is out of labor force, 0 otherwise.

Figure 2.1.  
 Value of Benefit Stream Contingent on Claiming Age,  
 DRC=5% per year

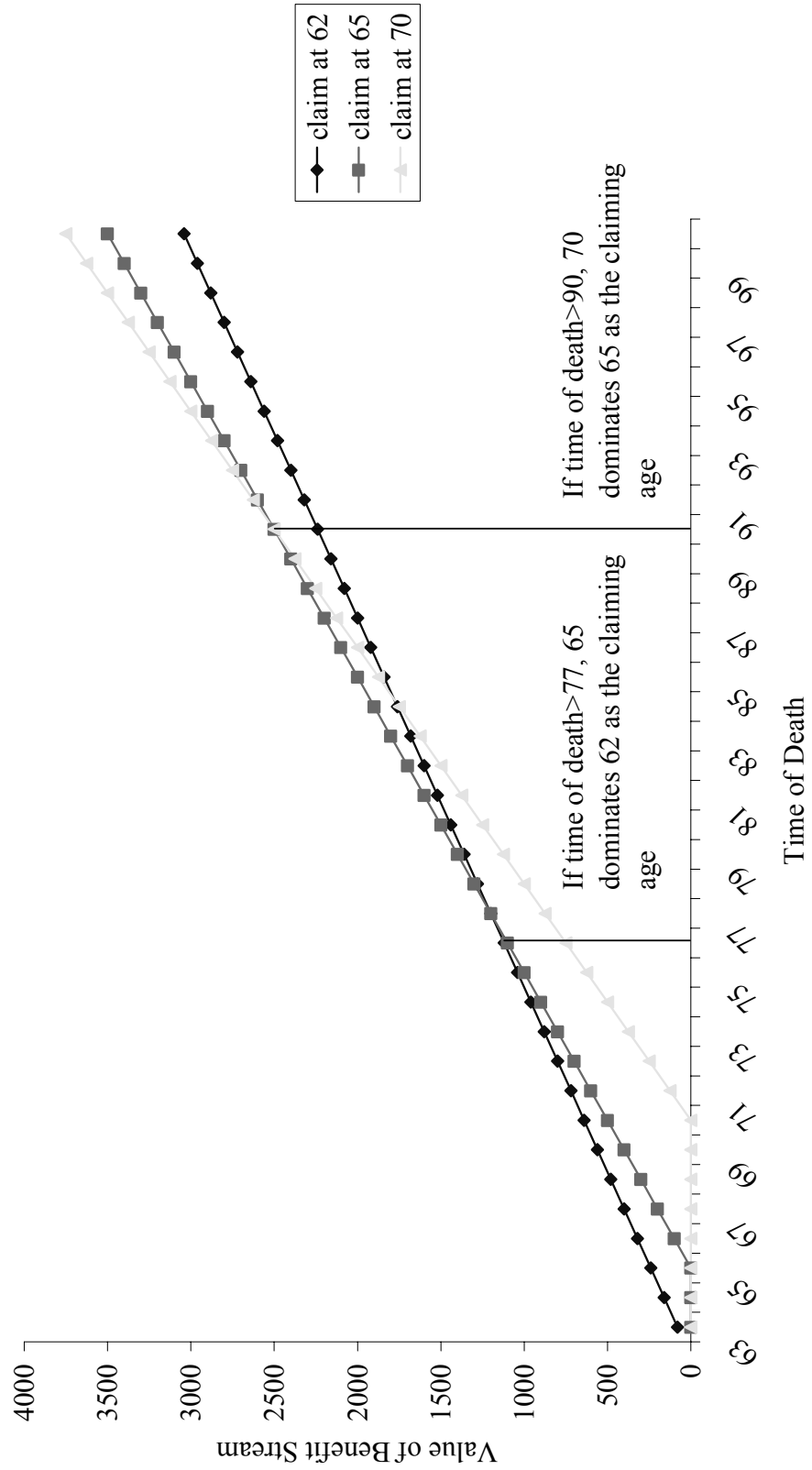


Figure 2.2.  
Annual Exempt Amounts for RET, 1940-1999

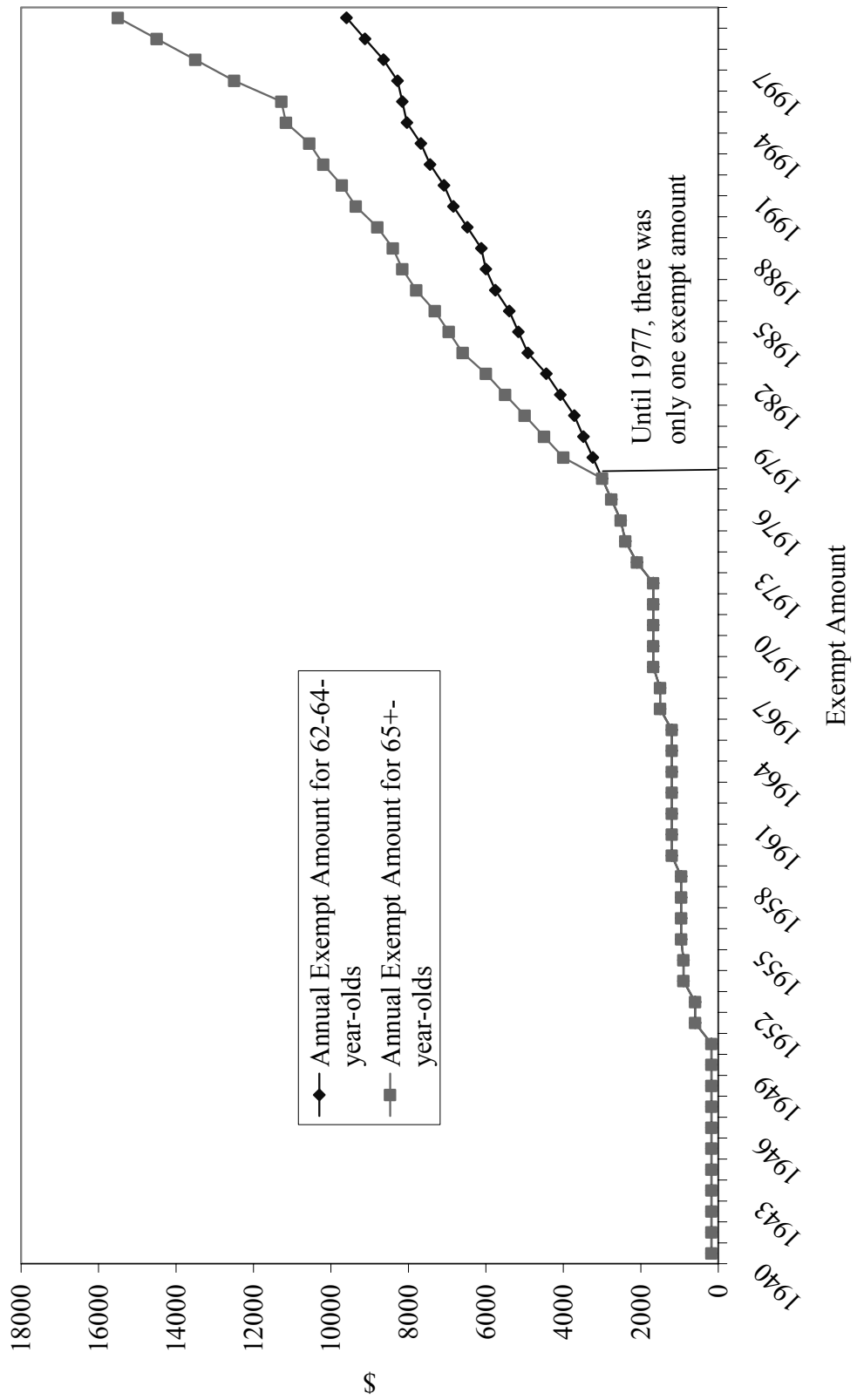


Figure 2.3.  
 DRC amount for various birth year cohorts,  
 as specified in the 1983 Social Security Amendments

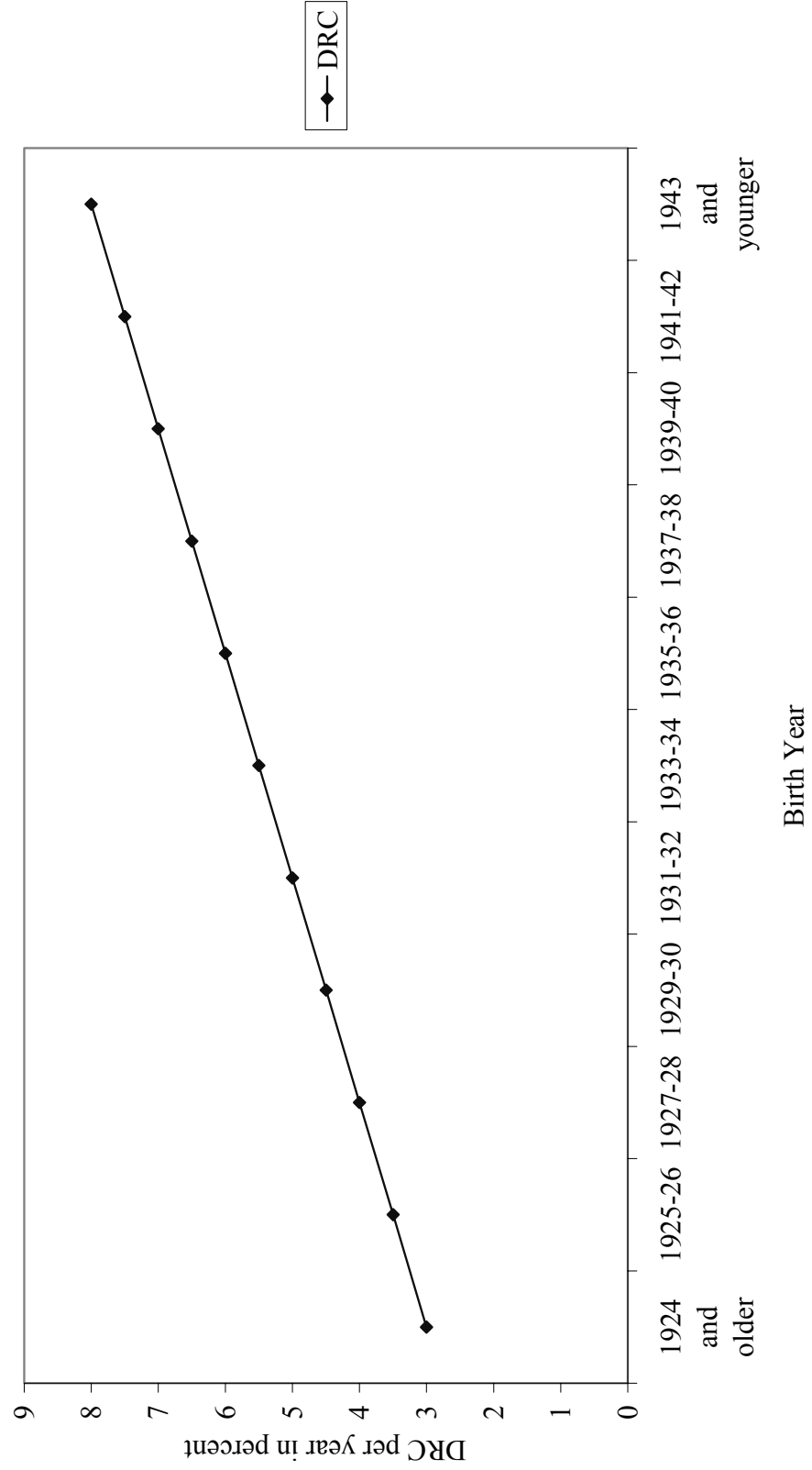


Figure 2.4.  
Share of White Males who Worked more than 30 Hours Last Week

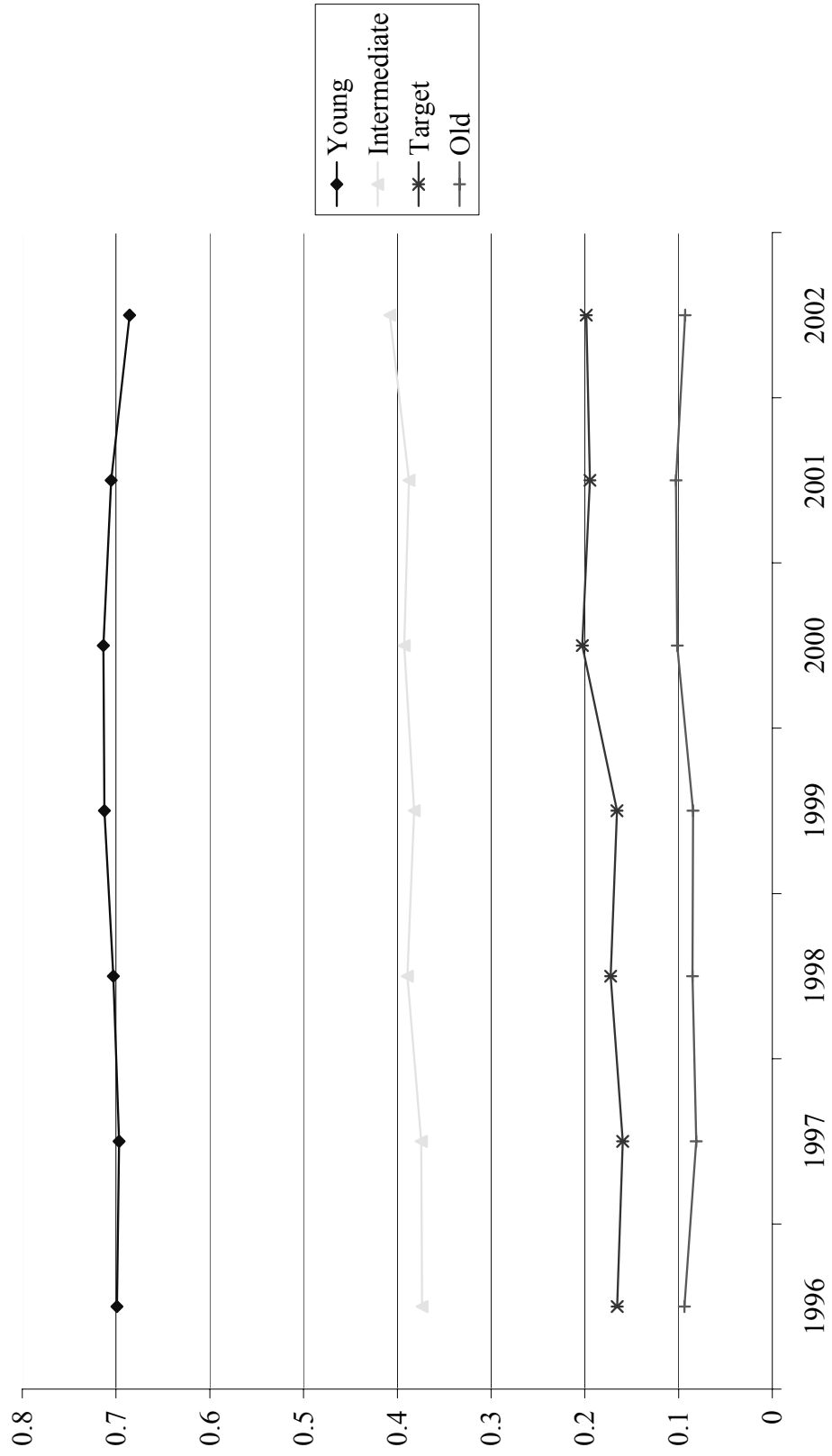
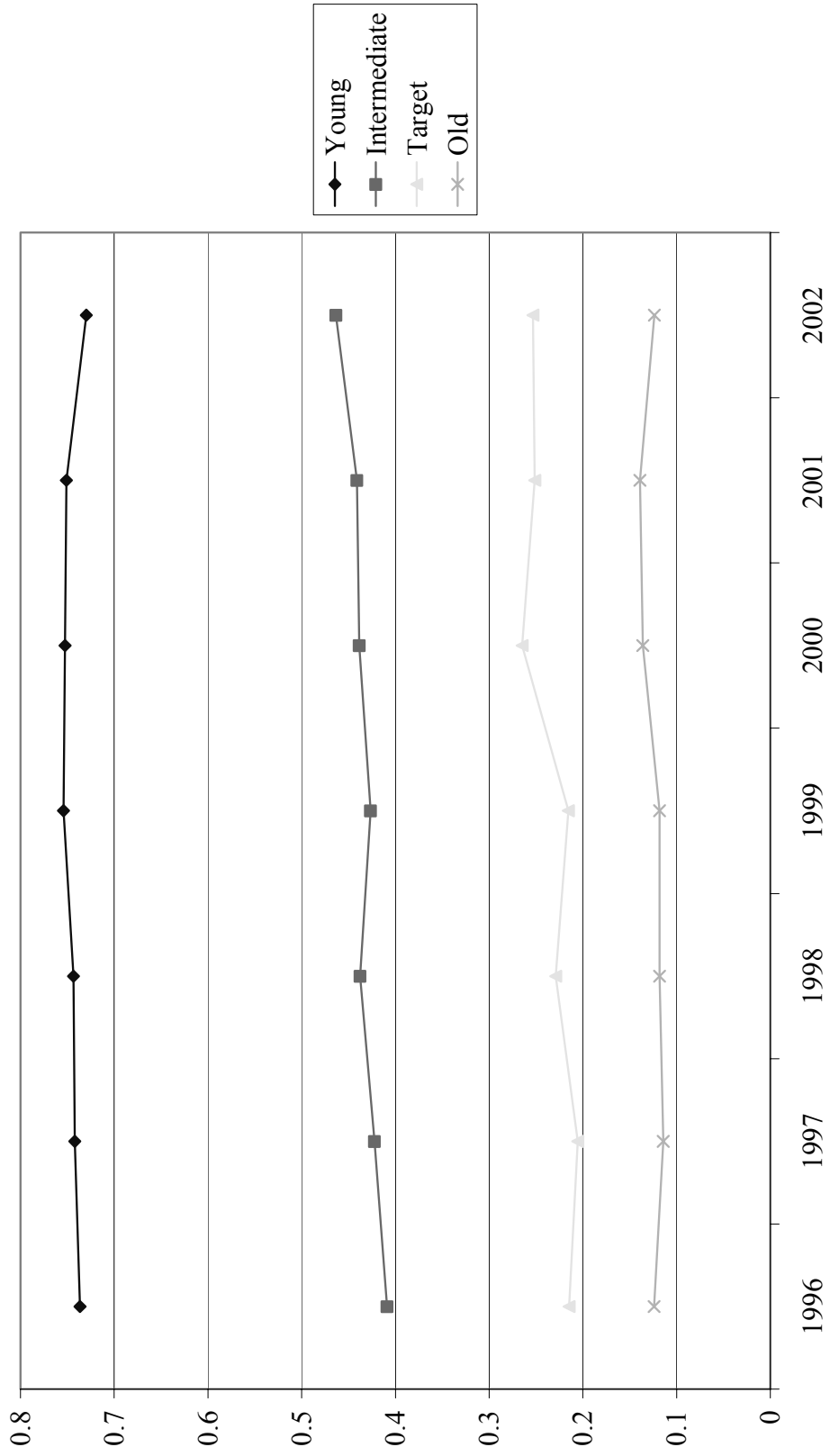


Figure 2.5.  
Share of White Males who Worked more than 20 Hours Last Week





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