Essays in Labor Economics

Author: Geoffrey Todd Sanzenbacher

Boston College

The Graduate School of Arts and Sciences

Department of Economics

ESSAYS IN LABOR ECONOMICS

A dissertation

By

GEOFFREY SANZENBACHER

Submitted in partial fulfillment of the requirements

For the degree of

Doctor of Philosophy

May 2010

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Geoff Sanzenbacher

Advised by Shannon Seitz

Issues pertaining to low income workers are of the upmost interest to policy makers. In the mid 1990s, the issue of welfare recipients and work was at the forefront of public policy, as the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 was passed. One of the many goals of the policy was to "end the dependence of needy families on government benefits" by encouraging work and ultimately higher wages. The first paper of my dissertation explores the processes by which work leads to wage growth for welfare recipients. I find that welfare recipients have similar returns to tenure and experience as non-recipients and that tenure has higher returns than experience for these women. Because of this, policies that discourage leaving work, like a work requirement, are more effective encouraging wage growth than policies discouraging welfare use, like a time-limit.

A decade later, the low savings rates of low income workers has led policy makers within the Obama administration to consider making Individual Retirement Accounts (IRAs) available to all workers. The second paper of this dissertation examines how likely low individual workers are to participate in these plans. We find that low-income workers not currently offered voluntary retirement savings plans are less likely to participate than those currently offered those plans. The paper indicates policy makers should be wary of basing estimates of participation in the offered IRAs on current participation, as this may overestimate the participation rate by up to 25 percent.

Wage Growth and the Job Dynamics of Welfare Recipients

Geoffrey Sanzenbacher
Boston College

Abstract

The welfare reforms of 1996 were designed to encourage single mothers to become self-sufficient through employment. Yet, these women often end up in unstable, low-paying jobs. In this paper, I quantify the importance of (1) the returns to tenure and experience, (2) job mobility, and (3) job exit in leading to these employment outcomes. I estimate a model of full-time work, part-time work, and welfare use. To allow differences in wage growth between recipients and non-recipients, I incorporate heterogeneity in job offer arrival rates, the returns to experience and tenure, and the rate of job destruction. I show that, for welfare recipients, tenure is a more important source of wage growth than work experience. Thus, policies encouraging lengthy employment spells could encourage wage growth. Policy experiments indicate that a work requirement on welfare receipt encourages longer employment spells and four times as much wage growth for women between the ages of 18 and 33 as a five-year lifetime welfare receipt time limit.

1 Introduction

In 1996, the United States reformed welfare to "end the dependence of needy families on government benefits." Prior to these reforms, welfare recipients had low employment rates, short spells of employment, and low wages. Taking data from the NLSY79 between 1979 and 1993 for a sample of 16 to 34-year-old women who did not graduate from college we can see the following:

1. Women who did not receive welfare between 1979 and 1993 had an employment rate of 73 percent compared with 44 percent for women who received welfare.
2. Welfare recipients frequently exit employment. Of working welfare recipients, 25 percent were not employed the next year compared with 11 percent for nonrecipients. The average tenure of recipients is 1.5 years shorter than nonrecipients.
3. Many of these transitions to non-employment are accompanied by welfare use: 43 percent of welfare recipients who transited from work to nonemployment were using welfare in the first year out of work.
4. Welfare recipients have low wages and little wage growth. The average hourly wage of nonrecipients was \$9.16 compared with \$7.19 for recipients. For nonrecipients, wages grow 48 percent between age 20 and 30 compared with 28 percent for recipients.

These facts illustrate differences in employment rates, the length of employment spells, and wage growth between women who received welfare at least once between 1979 and 1993 and those who did not receive welfare over that period.

Given these facts, it is unclear whether work encouraged by welfare reform is stable or accompanied by valuable experience.[1] Since the goal of welfare reform is to encourage self-sufficiency, it is important to understand the processes that lead welfare recipients to end up in low-wage jobs. In this paper, I provide quantitative estimates of the importance of (1) the returns to tenure, (2) the returns to experience, and (3) job mobility in leading to these outcomes for welfare mothers. I also explore whether individual characteristics or

[1] While it is usually agreed the reforms were successful in encouraging welfare recipients to work [Blank, 2002], it is not clear this work led to stable, high-wage jobs [Johnson and Corcoran, 2003].

pre-reform policy led to infrequent employment and low wages for these women. Finally, I determine what policies are most effective for encouraging wage growth for welfare mothers. These tasks are accomplished by estimating a structural model of full-time work, part-time work, and welfare use.

The facts presented above are consistent with two possibilities captured by my model. The first is that pre-reform welfare discouraged employment and encouraged employment exits into welfare use. In this case, policies that limit welfare use will encourage recipients to work more frequently and transit from work to nonemployment less often. If welfare recipients have high returns to tenure and experience or opportunities for job mobility, then these policies will encourage higher wages.

The second possibility is that welfare recipients are offered low wages and work infrequently for reasons unrelated to welfare policy. For example, women with the lowest stigma attached to welfare use may have permanently lower wages, lower returns to tenure, or high preferences for nonemployment. In this case, welfare reform will encourage work, but the characteristics of the women affected will preclude higher wages and lengthy periods of employment. To incorporate this possibility into my model, I allow for heterogeneity across observable and unobservable characteristics in preferences, job offer arrival rates, wage offer distributions, and rates of job destruction. The structural model will allow me to separate the effects of individual characteristics from the effect of policy on wages and wage growth. I estimate the model using a sample of non-college-educated women from the National Longitudinal Survey of Youth 1979 cohort.

The estimates reveal that high school dropouts have returns to tenure of 3.7 percent per year and returns to experience of 1.7 percent per year. Graduates have returns to tenure and experience of 4.3 percent and 1.3 percent, respectively. Of the two unobserved types allowed in estimation, the second has a lower stigma for welfare but 0.6 percent higher returns to tenure and experience. Thus, recipients have similar or higher returns

to experience and tenure as other women. However, these types of women also have much lower wages: dropouts wages are 22 to 25 percent lower than graduates, and the second unobservable type has 18 percent lower wages than the first. Because tenure has higher returns than experience, long spells of employment are an important element of wage growth.

The estimates also indicate that extended employment is difficult to encourage for welfare mothers. Unemployed black dropouts receive offers in only 26 percent of years and the unobserved type with low welfare stigma has a high preference for nonemployment. Furthermore, mothers face a utility cost of $4,600 dollars for full-time work.[2] Finally, black dropouts, a group likely to use welfare, have the highest level of job destruction at an annual rate of 6.1 percent. The importance of tenure has important policy implications. Policies discouraging movement from work to nonemployment will be effective in encouraging wage growth.

I perform several experiments to examine how different policies affect wage growth. These policy changes include enforcing a time limit on welfare receipt, attaching a work requirement to welfare use, and subsidizing low wages. These experiments show pre-reform welfare policy was partially responsible for low employment rates, short spells of employment, and low wages amongst welfare recipients prior to welfare reform. For example, requiring women to work within two years of receiving welfare reduces voluntary exits from employment by 30 percent, increases average tenure by 32 percent, and leads to average wages that are 11 percent higher by age 33 compared with pre-reform policy. However, the characteristics of welfare recipients limit the effects of policy changes; frequent recipients' wages remain 20 percent below nonrecipients and they exit employment 80 percent more often even after changes in policy. Estimating a structural model allows the effects of policy

[2]Using the structure I impose, it is not possible to separate pecuniary costs of work from nonpecuniary costs. This cost most likely reflects child care, but may include other costs as well.

and individual characteristics on wages to be separated.

This paper contributes to three distinct literatures. The first examines processes of wage growth among young, unskilled women. The typical study focuses on how a specific avenue for wage growth - for example, job mobility or the returns to experience - differ across observable characteristics like race and education. This is a large literature and includes work by Gladden and Taber (1999), Holzer and LaLonde (1999), Loeb and Corcoran (2001), Alon and Tienda (2005), and Connolly and Gottschalk (2006).[3] Gladden and Taber show that work experience leads to wage growth, even among less skilled workers and women who have received welfare. Loeb and Corcoran similarly find that women who received welfare between 1978 and 1992 had comparable returns to experience as women who had not. Gladden and Taber also show the importance of job exits and mobility in wage growth: job loss is associated with drops in the wage and mobility with increases.

Given the importance of job mobility and exit, Holzer and LaLonde show less educated minorities are more likely to experience job exits than other women. Alon and Tienda show this group is also less likely to experience job-to-job mobility. Thus, this literature suggests mobility and extended employment are important elements of wage growth. It also offers evidence on heterogeneity in these elements between welfare recipients and non-recipients. As such, heterogeneity in opportunities for job-to-job mobility and extended employment are included in my model. Also emphasizing the importance of mobility and tenure, Connolly and Gottschalk estimate a model of wage growth through experience, tenure, and job mobility allowing for heterogenous wage growth between individuals with different levels of education. They find that women with less than a high school education have low levels of within-job wage growth absolutely and compared to more educated women. Motivated by their work, I allow for heterogeneity in the returns to job tenure and

[3]Although their work deals with young men, Topel and Ward (1992) was one of the major motivations behind the work of Gladden and Taber.

general experience across women with different levels of education.

The second literature this paper contributes to evaluates the potential for wage growth through work for welfare recipients.[4] This literature can be divided into two categories: one set of studies finds that work does not lead to wage significant wage growth for recipients while the other set suggests the opposite. The set of studies finding low returns to work for welfare recipients includes work by Burtless (1997), Gottschalk (2000), Card and Hyslop (2005), and Johnson and Corcoran (2003). Using pre-reform data, Burtless (1997) attempted to evaluate the potential of welfare recipients to find gainful employment. He concludes that the skill deficiencies faced by welfare recipients do not preclude employment but do have negative effects on the wages they can obtain. Reviewing research on the returns to experience for welfare recipients, Gottschalk (2000) finds that experimental evidence suggests only small, temporary gains to work for these women while non-experimental studies are more optimistic. A study by Card and Hyslop (2005) uses data from a time-limited wage subsidy program in Canada, the Self-Sufficency Project (SSP). Their study finds that welfare recipients experience little wage growth as a result of the work encouraged by the program. They also show that the wage growth caused by the subsidy is temporary and vanishes once the time limit is reached. Johnson and Corcoran (2003) use a data set collected from the state of Michigan to show that women who had received welfare prior to the reforms faced difficulties obtaining "good" jobs after the reforms.[5] These difficulties were shown to be the result of job instability.

The second set of papers in this literature suggest work is accompanied by growth in wages for welfare recipients. Grogger (2009) uses data from an experiment in Florida, the Family Transition Program (FTP), to evaluate the effects of a welfare time limit and a

[4]There is also an extensive theoretical and empirical literature evaluating the effects of welfare reform on work and welfare receipt. Excellent reviews of this literature can be found in Blank (2002) and Grogger and Karoly (2005).

[5]In their work, Johnson and Corcoran defined a good job as paying at least 7 dollars an hour, having health benefits, and being full time or voluntarily part time.

decrease in earned income reduction rates on experience and the wage. He finds that the type of welfare policy change implemented in Florida in 1994 can lead to 3.7% higher wages after four years relative to a control group.[6] Card, Michaloplous, and Robins (2001) find that women who left the SSP program in Canada had similar returns to experience as women who were never in the program, although at low wages. Gottschalk and Connolly (2004) also use SSP data to show that women in the program are able to select into jobs with a higher return to tenure and stay on those jobs longer than women not in the program. They also find that the returns to general work experience are low for women in SSP and the control group. The finding that reform policies can impact tenure and that tenure may have higher returns than experience motivate the model described above.

This paper also relates to the structural literature that predicts the behavioral effects of welfare reform. Two papers are Swann (2005) and Keane and Wolpin (2007).[7] This literature structurally models employment, schooling, marriage, and fertility decisions. Both find that welfare reform significantly increases the probability that welfare recipients work, with smaller policy effects on marriage and fertility. However, these papers do not incorporate job tenure or job mobility and do not model wage losses associated with movement from work to nonemployment. The literature above suggests these factors for wage growth are important for women who use welfare. My model incorporates tenure, mobility, and job exit into a framework that will allow predictions to be made regarding the effects of welfare reform on wage growth. This will allow greater insight into how the wages of welfare recipients grow and provide more detailed prescriptions as to which types of policies will encourage wage growth.

The paper is organized as follows: Section 2 describes the model, Section 3 describes the data, Section 4 provides the econometric specification, Section 5 provides results, Section

[6]Grogger uses data on reservation wages to control for effect of the policy and unobervables on selection into work.

[7]Other papers in this literature include Keane and Wolpin (2002a) and Keane and Wolpin (2002b).

6 provides policy simulations, and Section 7 concludes.

2 Model

In this section, I provide an outline of the model, with exact functional forms provided in Appendix A. In the model, each woman has a finite decision horizon beginning when she leaves school and ending at age 50. In each year, she makes joint decisions about full-time work, part-time work, and welfare receipt from her decision set, D_a. Because job offers are not available in every period, the decision set is stochastic.[8] There are six choices a woman can make at any time: full-time work with or without welfare, part-time work with or without welfare, and nonemployment with or without welfare. The employment decision is $e = \{0, p, f\}$ for nonemployment, part-time, and full-time work, respectively, and the welfare decision is $b = \{n, y\}$ for no welfare receipt and welfare receipt, respectively. The decision a woman makes at any age a is denoted $d_a = \{eb\}$.

I assume there are K types of women who differ from each other along observable and unobservable dimensions. Specifically, a woman's type contains observable information on her race, education, and whether she has children.[9] A woman's type also contains a component that allows correlation between permanent unobservable differences in preferences over choices, the wage distribution, and job offer arrival rates.

At any age a, a type k woman's objective is to maximize her expected lifetime utility, U_a, given the choice set she faces. A woman's yearly utility, u_a, is determined by her type, k, her current decision, $\{eb\}$, and her consumption, c_a. Women have preferences over each choice in the model. A woman's preference for work is allowed to vary depending on whether or not she has had a first child, which is a determinant of her type, k. Her

[8] It is important to note that education is not modeled as a choice variable. In the model, women begin making decisions once their education is completed and are assumed to remain at that level of education for the remainder of their time in the model.

[9] Endogenizing the fertility decision is being considered as an extension to the model presented here.

preference over the welfare state, which is allowed to vary by race, can be thought of as a stigma attached to welfare use. Each period, there is a shock to a woman's preferences for each one of the choices and her current job. Her period preferences are:[10]

$$u_{d_a}(c_a, k) = \alpha_{e,k} + \alpha_{b,k} + c_a + \epsilon^u_{a,j,e,b} \quad ; \quad d_a = \{eb\} \tag{1}$$

The utility shock, $\epsilon_{a,j,e,b}$, differs across choices and across job offers. The shocks are assumed to be normally distributed with a covariance matrix Σ_ϵ to be estimated. This assumption ensures that no observed choice/job-transition combination has zero probability. Consumption, c_a, in a period is constrained by the sum of the wage times hours worked, the welfare benefit if taken, and any exogenous non-labor income. Women who work full time are assumed to work 1870 hours a year while women who work part time work 792 hours a year.[11] Thus, the budget constraint is:

$$c_a = 1870 * w_a * I(e = f) + 792 * w_a * I(e = p) + B_a I(b = y) + y_a \tag{2}$$

where B_a is the welfare benefit received and y_a is exogenous unearned income. Given this specification, a woman's objective in any period is to:

$$\max_{d_a \in D_a} U_a = E\big[\sum_{a'=a}^{50} \beta^{a'-a} u_{d_{a'}}\big] \tag{3}$$

The choice set available to a woman depends on whether a job offer arrives, the type of offer that arrives, and whether an old job is still available. The utility of full-time and part-time work depends on the offered wage. The next section discusses these aspects of

[10]The absence of a coefficient on money income implies that the utility (disutility) from choosing work or welfare use is being expressed in dollar terms.

[11]These were the median hours for women working more than 30 hours a week on their main job (full-time) and less than 30 hours a week on their main job (part time).

the model.

2.1 Job Dynamics and the Wage

In the model, job offers do not arrive in every period. The offer arrival rate varies across types of women and is dependent on a woman's previous year's decision. This captures the idea that welfare recipients may not be searching for employment or that contacts may occur at a different rate while employed. Given this, the offer arrival rate is denoted:[12]

$$\lambda_a = \lambda(d_{a-1}, k) \tag{4}$$

When a new offer arrives, the wage is determined by the woman's type and her accumulated full-time and part-time work experience, $X_{a,f}$ and $X_{a,p}$. I assume some jobs have different starting wages, which shifts the wage profile up. This shifter is denoted $\gamma_{0,j}$. As tenure, $T_{a,j}$, accumulates, the wage is assumed to grow at a rate $\gamma_{T,j}$, where the subscript j denotes heterogeneity across jobs in the wage profile.

Heterogeneity in the intercept and wage profile of specific offers allows for the possibility that welfare recipients end up in "dead-end" jobs: jobs that have low wages and little opportunity for wage growth. The agent knows whether the offer was part time or full-time, $\eta = \{f, p\}$.

The heterogeneity in offers is modeled as a multinomial logit that depends on an individual's type and the years of full-time or part-time experience she has accumulated, $X_{a,f}$ and $X_{a,p}$. The probability any offer j arrives with a given set of characteristics is:

$$Pr(\gamma_{0,j} = \gamma_0, \gamma_{T,j} = \gamma_T, \eta_j = \eta) = \lambda^j(X_{a,f}, X_{a,p}, k) \tag{5}$$

[12]For the econometric specification of the model, see Appendix A.

If welfare reform encourages work experience and experience is associated with the arrival of a higher starting wage or higher growth jobs, then the reforms may encourage substantial wage growth through mobility. On the other hand, if the type of woman who receives welfare also receives more "dead-end" offers, then the reforms may encourage work that is unaccompanied by wage growth and may not lead to self-sufficiency.

The wage in any period is:

$$\begin{aligned} w_a &= exp(\gamma^{\eta}_{0,j} + \gamma^{\eta}_{k} I(k = k') + \gamma^{\eta}_{X_f,1} X_{a,f} + \gamma^{\eta}_{X_p,1} X_{a,p} \\ &- \gamma^{\eta}_{X_f,2} X^2_{a,f} - \gamma^{\eta}_{X_p,2} X^2_{a,p} + \gamma^{\eta}_{T,j} T_{a,j}) \epsilon^{w}_{a,j} \end{aligned} \tag{6}$$

When an individual enters a period with a previously held job j, there is a chance she receives a new offer j' with a new intercept and slope parameter. In this case, she faces a decision over which job to accept. If an individual accepts the new offer, their tenure is reset to zero and she begins that period with the starting wage for that offer.

Finally, some offers that were held in the previous period are exogenously terminated at a rate δ. The rate at which that occurs depends on an individual's type:

$$\delta_a = \delta(k) \tag{7}$$

Wages grow through stable employment at one job if $\gamma^{\eta}_{T,j}$ is high and the job does not end. Wages also grow if there are substantial returns to general work experience. Wages grow through mobility if offers arrive frequently while employed or if "better" offers arrive more often as work experience accumulates. Job terminations or voluntary exits result in a loss of the job as an option and the loss of wage growth that occurred with tenure.

Aside from the wage, whether or not a woman has children also has an impact on the utility from working. Children also affect the value of work relative to the other model

choices by determining whether or not a woman can choose welfare. The next section discusses how children are included in the model.

2.1.1 Children

Children determine whether a woman can claim welfare and her utility from employment. Pregnancies at age a, denoted p_a, are assumed to occur exogenously and result in a birth one year later. The rate at which pregnancies occur is a function only of a woman's type.[13] This rate is estimated directly from the data and is expressed:

$$Pr(p_a = 1) = \phi(k) \tag{8}$$

A woman knows her probability of having a first child in future periods.[14] Only the first birth is recorded. After the first birth, a woman can choose to receive welfare and her preferences for work may be shifted. Prior to a birth, women do not have the option of welfare available.

2.2 The Welfare Benefit

The welfare guarantee level (the base amount a woman can claim assuming no other income) is assumed to be exogenous. Women who do not have children cannot choose to accept welfare. The grant level available to a woman with children is the average grant level between 1983 and 1991 as calculated by McKinnish, Sanders, and Smith (1998).

The welfare benefit available to a women in any period also has an endogenous portion. This is the amount her benefit is reduced when she decides to work, the earned income reduction rate. This effective rate is also taken from McKinnish, Sanders, and Smith

[13]Endogenizing the decision to have a child is a future model extension.

[14]Incorporating child age into the model is also a future model extension. The age of a child is an important aspect of the work/welfare decision, especially under welfare time limits [Grogger, 2002].

(1998). Finally, the benefit may be reduced if she has any unearned income. Thus, the welfare benefit available in a period is:

$$B_a = G - r_1 * 1870 w_a I(e = f) - r_1 * 792 w_a I(e = p) - r_2 y_a \tag{9}$$

where G is the grant level given a woman has children and r_1 and r_2 are the reduction rates for earned and unearned income. In the model, benefit reduction from earned income serves as the work disincentive of welfare.

The next section briefly discusses non-labor income, which decreases the value of the welfare benefit at the rate r_2.

2.3 Non-labor Income

Non-labor income is assumed to arrive exogenously from a distribution that depends on a woman's type and on her past period's full-time employment status. Non-labor income is included in the model, since some women may be less likely to use welfare because of income from sources other than work. These women are less likely to be affected by the reforms. Non-labor income is allowed to vary by previous full-time employment status because previously employed women may be collecting unemployment insurance simultaneously with welfare. Non-labor income is modeled as follows:

$$y_a = exp(\theta_{0,k} I(k = k') + \theta_1 I(d_{a-1} = \{f, b\}))\epsilon_a^y \quad ; \quad b = n, y \tag{10}$$

Unearned income encompasses child support, unemployment insurance, assistance from relatives, and income from farms and businesses that reduce the welfare benefit a woman can claim. Non-labor income makes a woman's welfare benefit smaller and thus discourages welfare use for certain women.

2.4 Dynamic Programming Problem

The lifetime utility maximization problem characterized by equation (3) can be written in terms of value functions. The value of any decision is a function of the period utility from that choice plus the expected value of future behavior given the choice. The expectation is over the distribution of future preference shocks, future wage shocks, and the probabilities job offers arrive or old jobs are terminated.

Given the elements described above, the utility of a choice depends on a woman's type, k, her child status, P_a, her tenure entering the period, $T_{a,j}$, her work experience, $X_{a,j,\nu}$, and the jobs she has available, J_a. Finally, utility depends on the set of utility and income shocks the woman receives. Denote the vector of state variables described above:

$$\{k, J_a, P_a, T_{j,a}, X_{f,a}, X_{p,a}, d_{a-1}, \epsilon\} = \mathbf{S_a} \tag{11}$$

The state space contains up to two different jobs: j_a is an old job that was not terminated and j'_a is a new offer that would begin today. The state variables evolve according to the following rule:

$$\begin{aligned} T_{a+1,j} &= T_a + I(j_{a+1} = j_a) \\ X_{a+1,f} &= X_{a,f} + I(d_a = \{fb\}) \;\; ; \;\; b = n, y \\ X_{a+1,p} &= X_{a,p} + I(d_a = \{pb\}) \;\; ; \;\; b = n, y \end{aligned} \tag{12}$$

The alternative specific value functions can be expressed as Bellman Equations:

$$\begin{aligned} V_a(\mathbf{S_a}) &= \max_{d_a \in D_a, j_a \in J_a} [V_{d_a,j_a}(\mathbf{S_a})] \\ V_{d_a,j_a}(\mathbf{S_a}) &= u_{d_a,j_a} + \beta E(V_{a+1}(\mathbf{S_{a+1}} | d_a \in D_a, j_a \in J_a, \mathbf{S_a})) \quad if \;\; a < 50 \\ V_{d_a,j_a}(\mathbf{S_a}) &= u_{d_a,j_a} \quad if \;\; a = 50 \end{aligned} \tag{13}$$

The problem is stochastic to the agent for two reasons: (1) she does not know the future values of the utility, wage, and non-labor income shocks and (2) she does not know what choice sets will be available in future periods. There are 15 shocks in the model, 10 utility shocks to each choice/job-offer combination, two wage shocks to the old full-time or part-time job, two wage shocks to the new full-time or part-time offer, and one shock to non-labor income. I assume all of the shocks are normally distributed with mean 0 and a covariance matrix to be estimated, Σ_ϵ.

The dynamic programming problem is solved recursively. In the last period, expected values of the optimal choice are calculated for each reachable state space point $\mathbf{S_A}$ and each potential choice set using Monte Carlo Simulation. Going back a period, this expected value is used to do the same calculation for period $A-1$. This is repeated until the first period is reached. This process is described in detail in Keane and Wolpin (1994).[15]

The next section describes the data used to estimate the model described above.

3 Data

Data come from the National Longitudinal Survey of Youth 1979 cohort (NLSY79). The NLSY is ideal for this application for several reasons. First, it contains dated job histories that allows me to identify ongoing jobs versus new jobs. Second, a series of questions on reasons for job exit allow involuntary and voluntary exits to be identified. Finally, the NLSY's longitudinal design allows tenure and experience to be observed from the time the individual enters the sample. The model is estimated using the survey years 1979 to 1994.[16]

The sample itself consists of women who have exited full-time school and who never

[15]For example, for a set of last period state variables $\mathbf{S_A}$, d draws of the wage and utility shocks are drawn and the highest valued choice for each possible decision set is recorded. The average over the d draws is the expected maximum value of arriving in A with that choice set available and state space $\mathbf{S_A}$. Going back a period, that expected value is used to do the same calculation for period a-1.

[16]The model is estimated on pre-reform data. After estimation, policy simulations will examine the effects of changes in welfare policy.

completed a college education.[17] Eliminating college graduates reduces the potential sample size from 6,283 to 5,124 women. Once a woman misses her first interview, she is eliminated from the sample from that point onwards. This implies the sample size is changing, as women enter the sample by leaving full-time school and others exit the sample through attrition. Table 1 contains descriptive statistics of the sample used for the estimation of the model.

Table 1: Sample Statistics, Estimation Sample by Age

	Age of Women		
	20	25	30
# In Sample	2,663	3,499	2,415
% Black	27.9	28.2	31.8
% High School Grad	82.7	83.9	84.4
% With Children	38.2	66.4	79.7
% Using Welfare	12.8	12.9	13.4
Average No. Years Using Welfare	0.4	1.0	1.8
Average Full-Time Experience	0.54	2.57	5.52
% Working Full Time	48.8	49.6	55.5
- Average Wage	\$7.1	\$8.9	\$10.3
Average Part-Time Experience	0.12	0.39	0.79
% Working Part Time	7.3	7.1	9.0
- Average Wage	\$6.6	\$7.9	\$8.8

3.1 Demographic Information and Children

Two observable aspects of a woman's type are her race and education. In the NLSY, individuals are asked during their first interview to identify their primary race or ethnicity. This question is used to determine whether a person was black or of Hispanic origin. There

[17] College graduates were dropped because they do not typically utilize welfare and thus are not relevant to this study. Only 3 percent of college graduates reported ever having used welfare.

are two levels of education in the model. A woman is said to have completed high school if she entered the sample having completed grade 12 or claimed to have a equivalence certificate or a diploma. Otherwise, she is coded as a dropout.

Another aspect of a woman's type is whether or not she has children. This affects her utility from work and the availability of welfare. Data on the existence and birth dates of children are collected from the Fertility and Relationship History section of the NLSY79, which contains dated birth histories. A woman is said to have a child if she listed at least one child's birth date on her child roster. Once a woman has her first child, it is assumed that the child lived with her for the remainder of the sample.

3.2 Jobs and the Wage

Data on jobs and wages are collected from the Work History section of the NLSY79. This section includes dated histories for all jobs worked in a given year. It also includes hours worked in a typical week on that job and its hourly wage. In the model, an individual is allowed to work only one job at a time. For women working more than one job at a time, a main job is identified from the data.[18] The main job is identified as part time if it required less than 30 hours in a typical week and as full time otherwise.

An important feature of the model is the role job-to-job and job-to-unemployment transitions play in wage growth.[19] A job change is said to occur if the main job in the period was not the last period's main job.

In the model, job exits are either voluntary or involuntary. Separate identification of offer arrival rates and preferences requires these two types of moves be distinguishable from one another. The NLSY asks the reason for any job exit. A job exit is said to be

[18] In practice, this involves merging information on up to 10 jobs from consecutive interviews to get the job history for any calendar year. The main job is the job in which the most hours were worked in a year.

[19] In each period, the NLSY assigns every job a number and collects what number the job was associated with the previous period.

nonvoluntary if the individual claimed she was fired, laid off, or if she claimed a program she was enrolled in ended. All other job exits are said to be voluntary.

3.3 Sources of Non-Labor Income

Data regarding an individual's welfare participation is collected from NLSY79-created variables on yearly welfare receipt. An individual is said to have used welfare in a given year if they received over $200 of AFDC in real 2000 dollars that year. Table 2 contains the proportion of women with children on welfare by type. The table indicates that black dropouts are the most likely to use welfare, followed by black graduates and white dropouts.

Table 2: Proportion of Women with Children Receiving AFDC, by Type

Type	Proportion on Welfare
White Graduates	0.103
White Dropouts	0.269
Black Graduates	0.286
Black Dropouts	0.529

Non-labor income is the sum of public assistance, including Social Security insurance, food stamps, disability, unemployment insurance, and income from farms and businesses.

4 Estimation Method

The solution to the dynamic programming problem serves as an input to perform Simulated Maximum Likelihood. This is accomplished in several steps. First, conditional on a set of model parameters, the dynamic programming problem is solved. This results in a set of expected maximum values for tomorrow's choices for an observation given $\mathbf{S_{i,a}}$. Given these expected values and the stochastic elements of the model, the value of any

choice is known to the agent. However, the econometrician does not observe the stochastic shocks to utility or the wage. Thus, the probability any choice and wage is observed takes the form of a multivariate integral over the shocks. Because errors are assumed to be jointly normal, this integral does not have a closed form. To form the simulated likelihood function, the next step in estimation involves calculating these probabilities by Monte Carlo Integration.

For each observation in the data, R wage and utility shocks are taken and the value function for each choice calculated. Given these value functions, the maximum valued choice and the simulated wage are recorded.[20] The probability any choice/wage combination is observed is the proportion of draws that it was the maximum valued combination. For each observation, these probabilities are calculated for any decision set, D_a, the individual could have faced. If $n_a(d_a \in D_a, w_a)$ is the number of times a simulated choice/wage combination is the maximum, then:

$$Pr(d_a \in D_a, w_a|\mathbf{S_{i,a}}) = \frac{n_a(d_a \in D_a, w_a)|\mathbf{S_{i,a}})}{R} \tag{14}$$

There are two issues to note regarding equation (15). The first is that it implies that the wage will not enter the likelihood function separately from the observed choice but through a joint probability. This is necessary because I have assumed the agent observes the shocks associated with both their current wage and with any wage offer she receives. Ignoring the worker's selection into jobs would bias the estimates of the wage parameters if they were estimated via OLS.[21] Second, the probability any specific wage is observed is a near-zero probability event. If an individual's observed wage is not produced in the simulations constructing their likelihood contribution, then their observation will have zero

[20]The wage is recorded only if the choice was full-time or part-time work. If the choice is nonemployment, the wage shocks are integrated out in estimation.

[21]This issue is discussed at length Aguirregabiria and Mira (2008).

probability. To address this issue, I follow Keane and Wolpin (1997, 2001, 2008) and assume the wage is measured with error. Specifically, I assume the measurement error is distributed log-normally so that:

$$w_a^{obs} = w_a^{true} exp(\epsilon_{m,a}) \tag{15}$$

where $\epsilon_{m,a} \sim N(0, \sigma_m^2)$. Given this assumption, the probability any choice wage combination is observed can be written:

$$Pr(d_a \in D_a, w_a^{obs}) = \int \phi(\epsilon_{m,a}) Pr(d_a \in D_a, ln(w_a^{obs}) - \epsilon_{m,a}) d\epsilon_{m,a} \tag{16}$$

where ϕ indicates the normal density and the probability of jointly making decision d_a and accepting the true wage comes from equation (14).

While equation (14) yields a choice probability for a woman making decision d_a facing decision set D_a, this is not sufficient to construct her likelihood. Since offers are unobserved, the econometrician cannot be certain which decision set a woman faced in any period. The probability a woman faces decision set D'_a can be written:

$$Pr(D'_a = D_a) = \Lambda_{D'_a}(\mathbf{S_a}) \tag{17}$$

Denote $\tilde{D}$ the set of decision sets an observed choice can come from. For example, full-time or part-time work can come from a decision set containing one or two jobs but not from a choice set without a job available. Given this, the likelihood contribution associated with

an individual observed making choice d_a with the associated wage w_a can be written:[22]

$$l_{i,a} = \sum_{k_u}^{K_u} p_{k_u} \sum_{D'_{i,a} \in \tilde{D}} \Lambda_{D'_{i,a}}(\mathbf{S_{i,a}}) \frac{n_{i,a}(d_{i,a} \in D'_{i,a}, w_{i,a} | \mathbf{S_{i,a}})}{R} \tag{18}$$

where p_k is the probability the woman is of unobserved type k. An individual's contribution to the likelihood at age a is a mixture of the contribution she would make for each type. The likelihood function for the sample is the product over ages and over individuals of the individual likelihood contributions, à la Heckman and Singer (1984).

The likelihood function is constructed and maximized through an iterative process: (1) a set of model parameters, Γ, is chosen, (2) the model is solved recursively conditional on Γ, (3) R draws for each individual are used to simulate the choice/wage probability in (15) given this solution, (4) the likelihood function is evaluated and the process begins again at (1). The likelihood function is maximized using the Simplex Algorithm of Nelder and Mead.[23]

4.1 Identification

4.1.1 Separate Identification of Offer Rates and Utility

One problem with identification occurs because elements of an individual's state space enter the probability a particular choice is available and the probability it is the maximal choice. Because only accepted offers are observed the econometrician does not know which decision set a woman is facing in any period. Thus, a woman who does not work in a period either did not receive an offer or received an offer that was not her maximum valued option. The econometrician cannot distinguish between these two events.

[22] k_u is the unobserved portion of an individual's type. This is part of her type k, which is an element of $\mathbf{S_a}$.

[23] Currently, this algorithm is used for its relative speed. The solution of the Nelder and Nead algorithm may be tested by using other optimization routines.

Two approaches are taken to deal with this identification issue. The first is to use plausible exclusion restrictions to identify the effects of individual type characteristics on model parameters. For example, the offer arrival rate is different for women with different levels of education but the preferences for the six states are assumed to be the same. Thus, if a high school dropout is observed to work infrequently, it would have to be because of a low offer arrival rate, not a high preference for nonemployment. This approach is taken where appropriate. However, it does not help to identify the level of offer rates.

The second approach for identification is to use an NLSY question on the reason a job was exited. This is used to determine whether the job exit was voluntary or exogenous. When a job is exited voluntarily, the econometrician knows a job was available and the expected offered wage. This information can be used to identify the utility of the non-employed state as it was chosen over an offered wage. As long as each type of woman is observed making voluntary transitions from work to nonemployment, then the preference for the unemployed state can be obtained for that type. The offer rate can be separately identified by observed transitions from unemployment to employment.

4.1.2 Identification of Wage Parameters

The goal of this paper is to determine whether work experience encouraged by welfare reform will lead to wage growth for the affected population. However, the estimation of the wage distribution as specified in equation (6) uses those individuals who select into work. Welfare recipients work much less frequently than other individuals. If this difference is driven by elements not captured by an individual's observed type, then estimates based on those who work frequently will tend to overstate the gains to work experience for welfare recipients. Identification of the true parameters in equation (6) requires that the unobserved shocks be uncorrelated with tenure, experience, and the individual's type. This seems unlikely, since most studies find permanent unobserved heterogeneity plays a large role

in determining wage offers and preferences.[24] Ignoring this permanent component to the unobservables implies that the variables of the wage equation and other model features will be correlated with contemporaneous shocks.

To account for selection into work and welfare, I allow for unobserved permanent heterogeneity. I allow cross-correlation between the preference for welfare, the wage level, and the returns to experience and tenure. Identification of the wage parameters requires that conditional on an individual's observed and unobserved characteristics, the contemporaneous shocks are uncorrelated with variables of interest.

The objective of the estimation is to identify the type proportions, p_k, and the type shifters for the model elements affected by the heterogeneity. Identification of the parameters and type proportions follows from cross-group variation in model outcomes.

5 Results

In this section, I present results from estimation of the model described above using data from the 1979-1994 years of the NLSY79. The sample was stopped at 1994 because this was the last period before significant welfare reform began in the form of state waivers. Post-1994 data will be used to see how well the model describes individual behaviors and outcomes following the reforms. The parameterization of the wage equation, job offer rates, utility function, and destruction rates can be found in Appendix A together with parameter estimates and standard errors. In this estimation, there are two types of jobs, full time and part time. There were two unobserved types.

[24] Keane and Wolpin (2007) is especially relevant to the work presented here. Swann (2005) does not include unobserved heterogeneity, although he suggests his estimates imply it is present.

5.1 Parameter Estimates

5.1.1 Wage Equations

The specification and estimates for the wage equations can be found in Appendix A.1. There are several results worth noting, beginning with the estimates for the full time wage equation for a type 1 individual.[25] The estimates indicate both graduates and high school dropouts have higher returns to tenure than experience. For graduates, tenure has returns of 4.3 percent and full-time experience returns of 1.3 percent. For dropouts, the returns to tenure are 3.7 percent with returns to full-time experience of 1.6 percent. These estimates imply encouraging long spells of employment will encourage wage growth for welfare recipients. They also imply dropouts, a type likely to use welfare, will see similar levels of wage growth as graduates when working.

The estimated parameters also indicate that part time jobs pay between 5 and 20 percent less than full-time work depending on a woman's type. Furthermore, part time work experience decreases the level of full-time wage offers. Encouraging part time work will reduce opportunities for more lucrative full-time work. Black dropouts, the most likely group to use welfare, are the only group for which part time wage offers are higher than full-time offers. They may be especially likely to end up in these "dead-end" part time jobs.

Finally, the unobserved type shifters indicate that type 2, the type that has a high preference for welfare use, has 18 percent lower offered wages than type 1 individuals, but a 0.6 percentage point higher return to tenure and experience. Thus, the type likely to use welfare will experience wage growth as they work, but from much lower absolute levels than nonrecipients. This is consistent with results presented in Gladden and Taber (1999), although they only examined the returns to experience. Gladden and Taber find that the

[25]There were two unobserved types in this estimation. Type 1 was normalized to have no unobserved type shifters.

coefficient in a wage equation on the interaction between welfare use and an instrument for work experience is positive. Their estimates indicate that for white women the wage profile is 5.2% steeper for welfare recipients than nonrecipients, with a smaller positive coefficient for black women.

A major contribution of this paper is to model the returns to tenure and experience separately. The estimates imply tenure has higher returns than experience. Two remaining questions are: (1) are my estimates of tenure and experience consistent with other papers in this literature and (2) what does separating tenure and experience add to the understanding of welfare recipients' wage growth?

The answer to the first question seems to be that they are generally consistent. For example, for female black dropouts, Gladden and Taber estimate returns to full-time experience of about 4.6 percent.[26] In my estimation, black dropouts have returns to full time tenure of 3.7 percent and returns to full time experience of 1.7 percent. Data from the NLSY79 indicates that each average year of experience is accompanied by 0.5 years of tenure accumulation for these women. Thus, my estimates indicate the typical return to a year of experience is 3.6 percent for black dropouts.[27] On the other hand, Connolly and Gottschalk (2006) find much lower returns to within-job experience for dropouts, 0.8%, than I find here.[28] The result that women of the type likely to receive welfare have higher returns to tenure and experience is consistent with the literature. While Loeb and Corcoran find lower returns to experience for welfare recipients compared to non-recipients,

[26]To obtain this estimate, they used potential experience as an instrument for actual experience, interacting it with a race and gender dummy. A separate estimate on the interaction between experience and education allowed me to infer the returns to experience of black dropouts.

[27]The estimate I provide is slightly lower, because in my model, the returns to mobility are not captured by full-time experience. The wage is estimated through the behavioral model, implying the returns to mobility are not confounded with the returns to experience.

[28]The reason for this discrepency is unclear. However, Connolly and Gottschalk use a sample of women taken from the SIPP with an average age is 36. This paper and the work of Gladden and Taber both use the NLSY 1979 cohort, where women range from 14 to 36 between 1979 and 1994. It seems possible interactions between age and tenure could play a role in these differences.

Gladden and Taber find that the interaction between welfare receipt and work experience is positive.

Including tenure also has important implications for the wage growth of recipients. Consider two women with five years of full-time experience. A typical working non-recipient will have 2.6 years of tenure while a typical working frequent recipient will have 1.7 years of tenure on their current job. Thus, despite having the same years of experience, the nonrecipient will have 3.3 percent lower wages, conditional on their type. If welfare reform can increase the employment spell length of frequent recipients, then it can increase the returns to experience of their work. This would not be captured by models that include only the returns to experience.

5.1.2 Preference Parameters

Appendix A.2 contains the specification of the utility function and the estimates of the preference parameters. The estimates show that women with children face a cost to full-time work of $4,600 relative to women without children.[29] This estimate is in line with outside estimates of child care costs. According to the National Association of Child Care Resource and Referral Agencies, the cost of child care in the median U.S. state is $4,803. Thus, women with children will be more likely to leave work voluntarily for any given wage. Furthermore, individuals of the unobserved type with the lowest stigma for welfare use, type 2, also have the higher utility value of nonemployment,$2,629 over type 1 individuals. Both of these results indicate welfare recipients may be unlikely to experience long spells of employment. Finally, the estimates indicate that blacks have a much lower welfare stigma than whites, a result consistent with previous studies.[30]

[29]This cost likely reflects child care, but may reflect other costs of working as well.
[30]For example, Keane and Wolpin (2008).

5.1.3 Offer Rates

The specification of the yearly job offer rate and parameter estimates can be found in Appendix A.3. The estimates indicate that the type of women most likely to use welfare, unemployed black dropouts, has the lowest annual offer rate, receiving offers in 23 percent of all years. This is followed by white dropouts at 30 percent and then both black and white graduates at nearly 42 percent. However, once these women are employed, the job offer arrival rate increases substantially. For example, black dropouts who were employed the previous period receive new offers 58 percent of the time. Being employed leads to more job offers and opportunities for mobility. Encouraging consistent employment will encourage wage growth both through the returns to tenure and through job mobility.

The estimates also indicate that being on welfare greatly reduces the probability a job offer arrives. This reduction is between 18 and 35 percentage points depending on the woman's type. Reducing welfare's availability, through time limits or work requirements, will increase the number of offers recipients receive by discouraging welfare use.

5.1.4 Job Offer Types

There are two types of jobs in the model: full time and part time. The specification and parameter estimates for the probability an offer was full time can be found in Appendix A.4. Part-time jobs have lower starting wages than full-time jobs and part time experience lowers full-time wage offers. If the types of women who receive welfare also receive few full time offers, the ability of policy to encourage work in high-wage jobs will be limited.

The estimates reveal this is occurring, but to a limited degree. Black dropouts have the highest probability of a job offer being part time, at 25 percent. This is followed by black high school graduates at 23 percent, white dropouts at 19 percent, and white graduates

at 18 percent.[31] The estimates also reveal that part-time work experience decreases the probability a full-time offer arrives. For example, 37 percent of new offers to a black dropout with four years of part-time experience are part time, 13 percentage points higher than if she had no part time experience.[32] Thus, part time work decreases the full-time wage and the probability a full-time offer arrives.

5.1.5 Job Termination Rates

Appendix A.5 contains the parameterization and estimates of the job termination rates, which were highest for black dropouts and lowest for white high school graduates. Black dropouts lose their jobs at an annual rate of 6 percent, compared with 5.4 percent for white dropouts, 3.4 percent for white graduates, and 4.4 percent for black graduates. A job termination rate of 6 percent implies black dropouts have only a 73 percent chance of accumulating 5 years of tenure without a job termination whereas a white graduate will have an 85 percent chance. Thus, long spells of employment may be difficult to encourage for women who are likely to use welfare.

5.1.6 Unobserved Type Proportions

The estimates indicate that the second unobserved latent type has the following characteristics: (1) 17 percent lower wage levels than the first type, (2) 0.6 percentage point higher returns to tenure and experience, (3) a high preference for nonemployment, and (4) a low stigma for welfare use. This type comprises 30 percent of the population. In the simulations discussed below, the second type comprises a disproportionately high number of frequent welfare recipients.

[31] Because of the difficulty interpreting the logit parameters estimated, these percentages are for women without any full-time or part-time experience.

[32] This is for a woman with no full-time experience.

5.2 Model Fit

5.2.1 Internal Validation

The estimated parameters were used to create a simulated sample of 5,500 individuals to compare the results of the model against the actual data. The sample was simulated to match the observable type proportions, entrance times, and ages observed in the actual data. Table 3 presents statistics from the actual and simulated sample for the total proportion of individuals in full-time and part-time work, their wages in those jobs, and the percent of women with children on welfare. These are basic statistics the model should fit.

Table 3: Selected Statistics for Women Age 18-33, NLSY and Simulated Samples

	Actual	Simulated
Statistic		
Proportion Working Full Time	0.513	0.534
- Mean Full-Time Wage	\$8.65	\$8.45
Proportion Working Part Time	0.079	0.086
- Mean Part-Time Wage	\$7.72	\$7.29
Proportion Eligible Women on AFDC	0.205	0.175

The model fits the average choice proportions over the sample both qualitatively and quantitatively. It overstates the proportion of years spent in full-time work by 4 percent and in part time work by 9 percent. As a result, it underestimates the percent of years spent on welfare by 14 percent. The model fits the average full-time wage well, underestimating the average part time wage by 5 percent.

Aside from choice proportions and wages, the model should also fit conditional transition probabilities from unemployment to work, from job to job, and from work to unemployment. Table 4 contains a comparison of these probabilities across the actual and simulated samples. The model fits these conditional transition probabilities well; none of the simu-

lated transition probabilities are more than 5 percent away from the actual probabilities.

Table 4: Yearly Transition Probabilities for Women Age 18-33, NLSY and Simulated Samples

	Actual	Simulated
Proportion With Transition From:		
Unemployment to Employment	0.289	0.304
One Job to Another	0.247	0.233
Job to Unemployment	0.154	0.147
- Voluntarily	0.113	0.109
- Non-Voluntarily	0.041	0.038

In the introduction to this paper, the series of stylized facts presented suggest that frequent welfare recipients earn low wages, have low rates of employment, and make frequent movements from employment to nonemployment. Can the model as estimated fit these stylized facts? This is an important test of the model's ability to capture differences in employment outcomes and dynamics across recipients and nonrecipients.

Table 5 compares mean wages for nonrecipients, infrequent welfare recipients, and frequent welfare recipients. Nonrecipients did not receive welfare at all between ages 18 and 33, infrequent recipients received welfare in one to three years between ages 18 and 33, and frequent recipients received welfare for more than three years between ages 18 and 33.[33] I have divided the sample by years of experience to compare both the levels and growth paths of wages.

Table 5 shows the model fits the growth paths of wages across recipients and nonrecipients well, overstating the growth path for frequent recipients slightly. However, the wage levels for nonrecipients are too low and are too high for infrequent recipients. The importance of including unobservable types in the estimation can be seen by examining Figure 1. This shows how well the model fit the wage profile of frequent recipients before

[33]The welfare recipiency groups were divided to roughly equalize the size of the two welfare receiving groups.

Table 5: Mean Wage for Women Aged 18-33, Actual and Simulated Samples, by Years of Welfare Receipt.

	Lifetime Years of Welfare Receipt					
	0 Years		1-3 Years		Over 3 Years	
Years of Experience	Actual	Sim	Actual	Sim	Actual	Sim
0-2 Years	7.52	7.35	6.70	7.00	6.33	6.31
3-5 Years	9.09	8.40	7.78	8.02	7.26	7.21
5+ Years	11.44	10.56	9.20	9.58	8.09	8.74

and after including a second unobserved type.

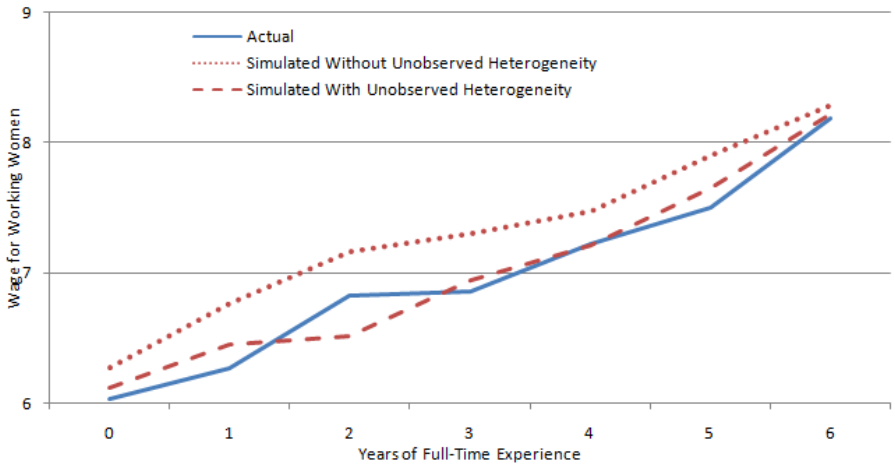

Figure 1: Mean Wage for Frequent Welfare Recipients

The model also fits differences in employment transition probabilities, employment rates, and rates of tenure accumulation across welfare recipient types. Recipients are much less likely to work and less likely to accumulate high levels of tenure due to their frequent moves away from employment and slow movement from unemployment to employment.

Table 6 indicates the model is able to capture differences in employment transitions and tenure accumulation. However, it does overstate transitions from nonemployment to employment across all recipiency groups, and also overstates the number of frequent recipients leaving work voluntarily. Still, the model fits the qualitative features of the data

Table 6: Selected Statistics, Actual and Simulated Sample

	Lifetime Years of Welfare Receipt					
	0 Years		1-3 Years		Over 3 Years	
	Actual	Sim	Actual	Sim	Actual	Sim
Probability of Transition from:						
Nonemployment to Employment	0.38	0.44	0.34	0.29	0.17	0.23
One Job to Another	0.25	0.21	0.28	0.26	0.28	0.30
Job to Unemployment	0.11	0.09	0.19	0.19	0.32	0.34
- Voluntarily	0.09	0.06	0.13	0.14	0.22	0.28
- Nonvoluntarily	0.02	0.03	0.06	0.05	0.10	0.06
Statistic:						
Employment Rate	0.72	0.78	0.61	0.55	0.33	0.38
Average Tenure	2.44	2.63	1.75	1.60	1.06	0.95
Mean Highest Tenure	5.22	5.59	3.86	3.80	1.95	2.19

well across recipiency groups.

5.2.2 External Validation

The previous section demonstrates that the model can reproduce the behavior, wage profiles, and employment transitions of the NLSY sample across welfare recipiency groups. Yet, one of the major motivations of the empirical approach taken in this paper, preforming policy experiments, requires taking the analysis out of sample. Can the model as estimated replicate the effects of a change in welfare policy?

To answer this question, I utilize a well documented experiment called the "Florida Family Transition Program (FTP)." This program randomly assigned women who received welfare between May 1994 and February 1995 to either a control group or an experimental group. The control group remained under AFDC guidelines. The experimental group faced two primary policy changes: (1) a 24 month time limit over any 60 month period and (2) a 50% decrease in the earned income reduction rate. The experiment lasted five years. Grogger (2009) used reservation wage data to examine the effects of the experiment on

wages while controlling for selection into work. Bloom et al. (2000) documented the effects of the experiment on employment and welfare use.

I compare the effects of the experiment as documented by Grogger and Bloom et al. with the effects generated by my model. To do this, I simulate a sample of women to age 33 under AFDC rules. I then flag a group of women who received welfare at age 32 and divide them into two groups, a group who continues for the next five periods under AFDC and a group who continues under Florida's FTP as described above. The results of the simulations, Grogger's, and Bloom's analysis is contained in Table 7.

	Actual	Simulated
Bloom et al. (2000)		
% Δ in Employment, Control to FTP	10.3	13.2
% in FTP Reaching Time Limit	17.0	12.0
Grogger (2009)		
Increase in Experience, FTP Over Control	0.25	0.21
% Δ in Wage caused by FTP	3.7	2.9

The simulations produce larger effects on employment than occurred in the actual experiment. As a result, fewer women reach the two-year time limit than did in Florida's FTP. The effects on additional work experience and the wage of FTP participants are comparable to the experiment.[34]

6 Policy Experiments

The estimates indicate that long spells of employment will lead to wage growth, even for frequent recipients. However, it is possible their low wage levels, high costs of employment,

[34]The effect of the policy on the wage was estimated by Grogger using reservation wage data to perform a singly censored bivariate probit regression. I control for selection into work using the reservation wage generated by the model and the simulated woman's unobserved type, which is observed in the simulations, to perform a standard Heckman two-step regression.

and high rates of job termination will preclude extended spells of employment and thus substantial wage growth. Can welfare reform lead to wage growth for women likely to be affected by the policy changes? Are certain policies more effective than others?

To answer these questions, I performed four sets of policy experiments to examine how frequent welfare recipients' wages respond to welfare reform. Three of these policy experiments involved changes to welfare policy while the fourth involved a wage subsidy.

6.1 Changes in Welfare Policy

The three sets of policy experiments focused on welfare rules were:

1. A Five-Year Time Limit on Benefit Receipt: In this experiment, women can receive welfare for five periods over the course of their lifetimes. Once their five years of receipt is up, they are ineligible for welfare. This experiment is similar to the federally mandated time limit imposed in 1996 as part of welfare reform.
2. A Two Year Work Requirement: In this experiment, a woman has to work within two years of nonemployed welfare receipt. Once she receives welfare for two consecutive periods without working, she is ineligible for a benefit.
3. A Two Year Work Requirement with a Child Care Subsidy for Full-Time Work: This experiment is the same as the second experiment, but supports full time working mothers with a $2,000 child care subsidy. This kind of policy is common in many states, for example, California.

Implementing the policy experiments required resolving the model given the estimates described above after making changes to the model's structure. Simulating a welfare receipt time limit required adding an extra variable to $\mathbf{S_a}$: total time spent on welfare. Each year a woman chose welfare receipt this variable increased by one year. Once total time spent on welfare exceeded five years, a woman was no longer able to choose welfare receipt. Enforcing a work requirement also required an additional state variable: the number of consecutive years spent nonemployed and on welfare. For this policy, women who both did not work and received welfare for two consecutive periods could not receive welfare in the

third period. However, once a woman worked full or part time or did not receive welfare this variable was reset to zero and welfare receipt was an available choice again. Adding a child subsidy to this policy involved increasing consumption by $2,000 for women who both worked full-time and had children.

The effects of these policies on annual choice proportions for frequent welfare recipients are described in Table 8.[35] The experiments indicate that a work requirement is better encouraging work than a time limit. The proportion of women in full time work increased 30 percent under a time limit, 41 percent under a work requirement, and 68 percent if a subsidy was added to support the work requirement.

Table 7: Annual Choice Proportions, Frequent Recipients Age 18-33 Under Baseline Policy and Three Experiments

		Experiment					
	Baseline	(1)		(2)		(3)	
	Level	Level	% Δ	Level	% Δ	Level	% Δ
Working Full Time	0.28	0.37	29.7	0.41	46.1	0.48	68.4
Working Part Time	0.10	0.10	0.0	0.12	20.0	0.10	0.0
With Children on Welfare	0.37	0.22	-40.7	0.24	-35.1	0.24	-35.4

All three of these policies also reduce welfare use: a time limit reduces welfare use by 40 percent and a work requirement by 35 percent with or without a subsidy. The smaller reduction in welfare use under a work requirement is largely due to workers mixing work and welfare; mixing work and welfare meets the work requirement while adding a year of lifetime receipt under a time limit.

Clearly these policies encourage work and discourage welfare use. Will they encourage long spells of employment, job-to-job mobility, and ultimately wage growth? Table 9 indicates a work requirement is more effective than a time limit in encouraging long spells of employment, higher levels of tenure, and job mobility. This is an interesting result, given

[35]In these experiments, "frequent welfare recipients" are women who would have received welfare for more than three years under the baseline policy. Frequent recipients were the target of the 1996 reforms.

both policies reduce welfare use by similar amounts without requiring any subsidization of work. A time limit increases average tenure by 22 percent and increases tenure on the longest held job by 25 percent. For a work requirement, these increases are 32 and 37 percent. Adding a child care subsidy to the work requirement enhances the requirement's effects, increasing average tenure by 60 percent and tenure on the longest held job by 59 percent.

Table 8: Annual Choice Proportions, Frequent Recipients Age 18-33 Under Baseline Policy and Three Experiments

		Experiment					
	Baseline	(1)		(2)		(3)	
	Level	Level	% Δ	Level	% Δ	Level	% Δ
Prob. of Transition:							
Non-Emp. to Emp.	0.23	0.28	20.2	0.32	39.1	0.34	46.4
Job to Job	0.30	0.31	3.3	0.32	6.6	0.31	3.3
Vol. Job to Non-Emp.	0.27	0.22	-20.5	0.19	-29.6	0.16	-43.5
Statistic:							
Average Tenure	0.94	1.16	22.6	1.24	31.9	1.52	60.9
Average Highest Tenure	2.19	2.74	25.3	3.00	37.0	3.48	59.2

Given this result, it seems likely a work requirement will be more effective than a time limit in encouraging wage growth. Figure 2 shows this is the case. Adding a child care subsidy enhances the effect of the work requirement, decreasing moves from work to nonemployment more than the work requirement alone.

Figure 2 illustrates the dynamics of wage growth caused by the policy changes. At first, the average wage is similar under the baseline and each policy. However, as frequent recipients increase their time spent working and decrease their job exits they accumulate both additional tenure and experience relative to the baseline policy.

Two conclusions can be reached from these simulations. The first is that low employment rates, short spells of employment, and low wages of recipients were caused in part by pre-

reform policy. Altering welfare policy encourages higher employment rates and longer spells of employment. The workers affected by the policy are as mobile as workers who were employed before the policy. Thus, as Figure 2 shows, work encouraged by the reforms is accompanied by wage growth.

The second is that exogenous characteristics still play a large role in determining the wages and employment dynamics of recipients. Even under the work requirement/subsidy policy change, frequent recipients leave their jobs voluntarily 80 percent more often than nonrecipients and have average tenure that is a year below nonrecipients. Under a subsidy and work requirement, frequent recipients' wages remain 20 percent below the wages of nonrecipients. The difference is more pronounced for the other policies.

The policy simulations also provide insight into the gains of modeling employment spells and mobility. Tables 7 and 8 indicate that the policies above both encourage work and discourage job exit. For frequent recipients, under the baseline welfare policy, each year of full-time experience was associated with a 0.2 year increase in tenure on the job. Under these policies, each year of experience was accompanied by a 0.32 year increase in tenure.[36]

[36] A 0.2 year increase in tenure means, on average, a frequent recipient with 10 years of full-time experience would have two years of tenure on her current job.

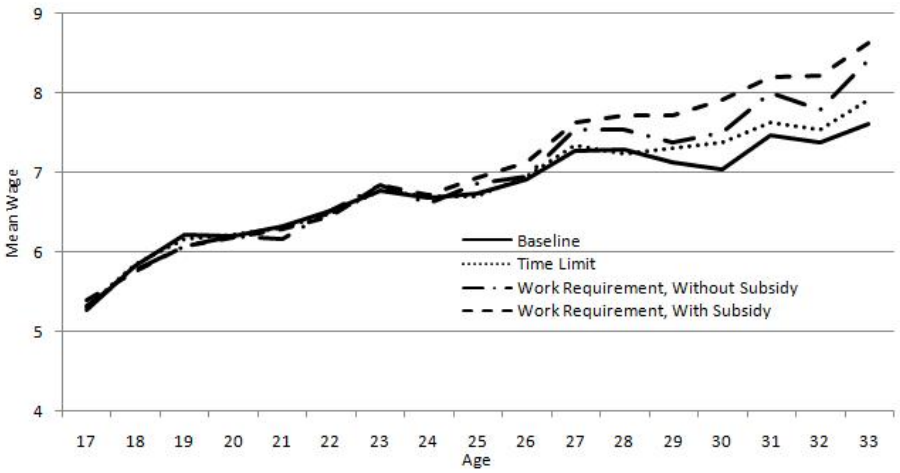

Figure 2: Mean Wage for Frequent Recipients, by Policy

This is a feature of welfare reform that could not be captured without modeling job dynamics; welfare not only encourages additional experience, but encourages longer employment spells and thus more tenure with that experience. This implies the returns to experience under the reforms are greater than under the baseline policy. Under the baseline policy, a year of full-time experience for a frequent recipient had an average return of 2.5 percent. Under the typical reform, it was 2.9 percent - a 15 percent increase. This effect of the reforms would not be captured without modeling employment spell length.

Given the substantial change in behavior and wages caused by the policy changes, it is natural to wonder how the utility of the frequent recipients is altered. While it is impossible for lifetime utility to increase under a time limit or work requirement, it is possible that later in life the higher wages caused by the reforms lead to higher period utility.[37] The next section investigates this possibility.

6.1.1 Welfare Analysis

To perform welfare analysis I record the period utility associated with each simulated individual's maximum valued choice under each of the policies described above. Figure 3 shows the utility of frequent recipients between the ages of 18 and 33. Both the time limit and work requirement policies remain below the baseline policy for the entirety of the sample. A work requirement has a more immediate effect on utility, as workers must alter their decisions away from their optimal choice earlier. The utility effect of this policy on young women is particularly severe, as they have not had time to accumulate experience and tenure in response to the policy change. However, toward the end of the sample period, women affected by a work requirement remain closer to the baseline policy, as their constraint applies only if they haven't worked within two years.

[37]Placing a constraint on behavior can never lead to a lifetime utility increase in the type of model estimated in this paper.

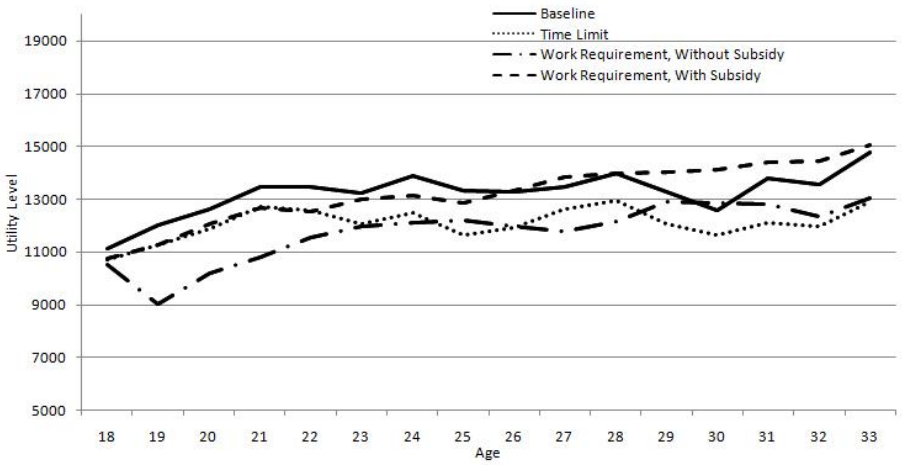

Figure 3: Utility Level for Frequent Recipients, by Policy

On the other hand, the utility from a time limit is similar to the baseline policy at first, dropping as time goes on as the time limit directly affects more women. From a utility standpoint, the best policy is clearly the work requirement with a child care subsidy. This policy both increases the consumption of women forced to work in response to the policy change and encourages wage growth during the period.

Enforcing a time limit or work requirement on welfare recipients has negative effects on frequent recipients' utility. On the other hand, a wage subsidy would not have negative utility effects and may encourage long spells of employment, less welfare use, and wage growth without constraining the choice set of recipients and without forcing government agencies to keep track of recipients' behavior. The next section focuses on the effects of this type of policy.

6.2 Effects of Subsidizing Low Wages

Wage subsidies have become popular in recent years, both to combat poverty and to encourage work for lower skill workers. For example, the federal government of the United States provides a wage subsidy for low-income families through the earned income

tax credit. Nearly half of the states have followed suit, implementing their own version of the EITC. To examine how a counter-factual wage subsidy alters the behavior of frequent welfare recipients I introduced a subsidy of the following form to both full-time and part-time wages:

$$s_a = max(\frac{1}{2}(12 - w_a), 0)$$

This subsidy was implemented by resolving the model given the parameters described above and adjusting the consumption from work to reflect the added income from the subsidy. For example, a woman who was offered a full-time job at $8 an hour would receive an additional two dollars an hour in consumption from that job. Can a wage subsidy of this sort have similar effects to welfare policy changes without placing constraints on an individual's choice set?

Tables 10 and 11 indicate that a wage subsidy of this size can have comparable effects to welfare policy changes.[38]

Table 9: Annual Choice Proportions, Frequent Recipients Age 18-33 Under Baseline Policy and Three Experiments

		Experiment			
	Baseline	Work Requirement		Wage Subsidy	
	Level	Level	% Δ	Level	% Δ
Working Full Time	0.28	0.41	46.1	0.43	53.6
Working Part Time	0.10	0.12	20.0	0.10	0.0
With Children on Welfare	0.37	0.24	-35.1	0.30	-18.9

The subsidy increases years spent in full-time work by 54 percent and decreases years spent on welfare by 19 percent. The policy causes a smaller drop in welfare use than a work requirement because the subsidy does not place any restrictions on receipt. Table 11 shows how the wage subsidy affects job dynamics relative to the work requirement and

[38]For comparison sake, I included the results for the work requirement without a child care subsidy in Tables 9 and 10.

baseline policies.

Table 10: Annual Choice Proportions, Frequent Recipients Age 18-33 Under Baseline Policy and Three Experiments

		Experiment			
	Baseline	Work Requirement		Wage Subsidy	
	Level	Level	% Δ	Level	% Δ
Prob. of Transition:					
Non-Emp. to Emp.	0.23	0.32	39.1	0.31	34.7
Job to Job	0.30	0.32	6.6	0.33	10.0
Vol. Job to Non-Emp.	0.27	0.19	-29.6	0.17	-37.0
Statistic:					
Average Tenure	0.94	1.24	31.9	1.26	34.0
Average Highest Tenure	2.19	3.00	37.0	2.98	36.1

Table 11 shows that a wage subsidy is similar to a work requirement with respect to encouraging long periods of employment. However, a subsidy accomplishes this in a very different way. Instead of eliminating nonwork options from the individual's choice set, it makes work more valuable directly through the subsidy. This means the policy has no negative effect on the workers' utility. Figure 4 shows that individuals under the subsidy program have higher utility than under the baseline policy or the work requirement.

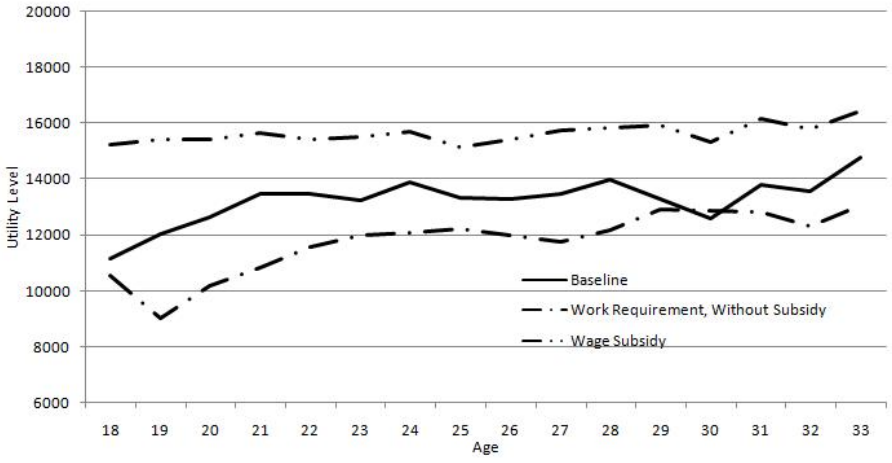

Figure 4: Utility Level for Frequent Recipients, By Policy

Finally, a subsidy has slightly different effects on the average wage of frequent recipients than a work requirement. While a wage subsidy affects employment spell length and employment rates similarly to a work requirement, it has a slightly different effect on observed wages. A wage subsidy that disproportionately benefits low-wage jobs encourages workers to accept lower wages than under the baseline or work requirement policies. At young ages, the average wage of frequent recipients who work is lower than under the baseline policy. However, as experience and tenure accumulate, this difference dissipates and is ultimately overcome by the wage growth this causes. It is worth noting that while a subsidy has similar effects on work behaviors and wages as policies that place restrictions on welfare use the subsidy may be much more costly to the government and thus less desirable in practice.

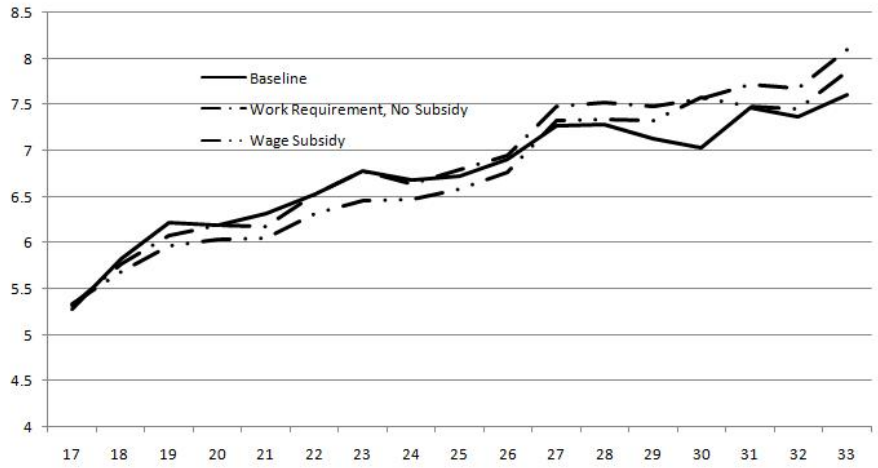

Figure 5: Mean Wage for Frequent Recipients, By Policy

7 Conclusion

The estimates reveal that women most likely to use welfare have returns to tenure and experience that are similar to other women. More importantly, women with a high

preference for welfare have slightly higher returns to tenure and experience than other women. Yet, the success of welfare reform in leading to high wages will be tempered by the low relative starting wages of women who use welfare. The type of women with a high preference for welfare have 18 percent lower wages than women who do not. Encouraging these women to work leads to wage growth, but not high wages. This result is consistent with the conclusions of Gladden and Taber (1999), who find that welfare recipients have five percent steeper wage profiles but much lower wage levels than non-recipients.

This model provides an opportunity to analyze the types of policies that can be implemented to encourage wage growth. For welfare recipients, tenure has substantially higher returns than does general work experience. Policies that encourage not only employment, but lengthy employment, will be most successful in encouraging wage growth. Policies that specifically penalize unemployment or offer work support are effective in encouraging long spells of employment and thus substantial wage growth.

These results also raise familiar questions, specifically regarding unobserved heterogeneity. What causes some women to have such a high preference for welfare? Why do these women have such low wage levels when compared with other women? This heterogeneity does not preclude welfare recipients from experiencing wage growth with employment, but seems to preclude the achievement of high wages. The policy experiments indicate that even after 15 years of welfare policy designed to encourage the accumulation of experience and tenure, the wages of frequent recipients are 10 percent lower than infrequent recipients and 22 percent lower than non-welfare recipients. Without addressing the root cause of these low wages, welfare reform will likely encourage wage growth and employment without leading to the high wages envisioned by the reforms.

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9 Appendix A: Econometric Specification and Estimates

9.1 The Wage Equation

The wage equation was specified as follows:

$$\begin{aligned} w_{i,a} = exp[&\gamma_1^\eta(1-r_a)(1-g_a) + \gamma_2^\eta r_a(1-g_a) + \gamma_3^\eta(1-r_a)g_a \\ &+ \gamma_4^\eta r_a g_a + \gamma_5^\eta(1-g_a)X_{i,a,f} + \gamma_6^\eta(1-g_a)X_{i,a,p} + \gamma_7^\eta g_a X_{i,a,f} \\ &+ \gamma_8^\eta g_a X_{i,a,p} + \gamma_9^\eta X_{i,a,f}^2 + \gamma_{10}^\eta X_{i,a,p}^2 + \gamma_{11}^\eta(1-g_a)T_{i,a,j} + \gamma_{12}^\eta g_a T_{i,a,j} \\ &+ \gamma_{13} I(type=2) + \gamma_{14} I(type=2)(X_{i,a,f} + X_{i,a,p}) + \gamma_{14} I(type=2)T_{i,a,j}]\epsilon_{i,a,j} \end{aligned}$$

where g_a is an indicator that takes a value of 1 if the individual graduated high school and r_a is an indicator if the individual is black. In the table, σ_1 was the standard deviation of

	$\eta = f$		$\eta = p$	
Parameter	Estimate	Std. Error	Estimate	Std. Error
γ_1	1.6870	0.0135	1.5574	0.0863
γ_2	-0.0151	0.1024	0.0722	0.5001
γ_3	0.2085	0.0165	0.2366	0.1005
γ_4	0.2002	0.0278	0.1380	0.2268
γ_5	0.0168	0.0081	0.0794	0.0751
γ_6	-0.0330	0.0542	0.0269	0.1178
γ_7	0.0126	0.0135	0.0568	0.0236
γ_8	-0.0254	0.0132	0.0493	0.0378
γ_9	0.0014	0.0035	-0.0017	0.0036
γ_{10}	0.0020	0.0174	-0.0023	0.0103
γ_{11}	0.0372	0.0135	-0.0439	0.0864
γ_{12}	0.0426	0.0086	0.0380	0.0428
γ_{13}	-0.1876	0.0498	-0.1876	0.0498
γ_{14}	0.0066	0.0034	0.0066	0.0034
σ_1^w	0.3295	0.0122	0.3678	0.0143
σ_2^w	0.3555	0.0159	0.4032	0.0165
σ^m	0.3831	0.0162	0.3831	0.0162

the wage shock to previously held jobs, σ_2 the standard deviation of new offers, and σ^m the standard deviation of the measurement error associated with wages.

9.2 Utility Specification and Estimates

The utility function was specified in the following way:

$$\begin{aligned} u_{i,a} &= \alpha_1 k_{i,a} I(e_{i,a} = f) + \alpha_2 (1 - k_{i,a}) I(e_{i,a} = p) + \alpha_3 k_{i,a} I(e_{i,a} = p) \\ &+ \alpha_4 (1 - k_{i,a}) I(e_{i,a} = 0) + \alpha_5 k_{i,a} I(e_{i,a} = 0) + \alpha_6 (1 - r_{i,a}) I(b_{i,a} = 1) \\ &+ \alpha_7 r_{i,a} I(b_{i,a} = 1) + \alpha_8 I(type = 2) I(e_{i,a} = 0) + \alpha_9 I(type = 2) I(b_{i,a} = 1) + \epsilon_{i,a,e,b,j} \end{aligned}$$

where e_a is the employment decision, b_a is the welfare decision, and k_a is an indicator equal to 1 if the individual has children. In this specification, full-time work without kids was used as the baseline utility status. The estimates of the nine utility parameters and shock variances are provided in the table below:[39]

Parameter	Estimate	Std. Error
α_1	-4600.52	581.23
α_2	4994.47	2357.02
α_3	2583.93	1221.69
α_4	6315.28	1186.78
α_5	3284.74	549.64
α_6	-9496.17	743.29
α_7	-4514.28	773.82
α_8	2629.29	1170.41
α_9	8320.24	1186.78
$\sigma_{a,0,0,0}$	22359.35	912.50
$\sigma_{a,0,0,1}$	30638.11	5677.39
$\sigma_{a,2,f,b}$	10301.04	1610.41
$\sigma_{a,1,p,b}$	3294.46	1486.42
$\sigma_{a,2,p,b}$	3808.53	1936.09

9.3 Job Offer Arrival Rates

The job offer arrival rate was:

$$\lambda_{i,a} = \frac{1}{1 + exp[\lambda Z^O_{i,a}]}$$

[39]The variance of the utility shock associated with full-time work at an old job was normalized to zero. The utility shocks for the old job have subscript 1, while new offers have subscript 2. Finally, because few people combined work and welfare before welfare reform, the utility shocks for mixing work and welfare were were constrained to have the same variance as those for work alone.

where:

$$\begin{aligned}\lambda Z^{O}_{i,a} &= \lambda_1(1-r_{i,a})(1-g_{i,a})I(e_{i,a-1}=0) + \lambda_2(1-r_{i,a})g_{i,a}I(e_{i,a-1}=0) \\ &+ \lambda_3 r_{i,a}(1-g_{i,a})I(e_{i,a-1}=0) + \lambda_4 r_a g_a I(e_{i,a-1}=0) \\ &+ \lambda_5(1-r_{i,a})(1-g_{i,a})I(e_{i,a-1}\neq 0) + \lambda_6(1-r_{i,a})g_{i,a}I(e_{i,a-1}\neq 0) \\ &+ \lambda_7 r_{i,a}(1-g_{i,a})I(e_{a-1}\neq 0) + \lambda_8 r_a g_a I(e_{i,a-1}\neq 0) + \lambda_9 I(b_{i,a-1}=1) + \lambda_{10} I(type=2)\end{aligned}$$

Parameter	Estimate	Std. Error
λ_1	0.8285	0.2170
λ_2	0.3066	0.0832
λ_3	1.1993	0.3809
λ_4	0.3857	0.1398
λ_5	-0.8924	0.5840
λ_6	-1.1143	0.2106
λ_7	-1.5231	1.0492
λ_8	-0.8886	0.3387
λ_9	1.6676	0.2982
λ_{10}	-2.3593	0.9468

9.4 Job Type Rates

When a new offer arrives, it is either a full-time or part-time job. This determines the distribution of its starting wage and return to experience and tenure. The probability a job was full time was specified as:

$$Pr(\eta = f) = \frac{1}{1+\eta Z^{\eta}_{i,a}}$$

where:

$$\begin{aligned}\eta Z^{\eta}_{i,a} &= \eta_1(1-r_{i,a})(1-g_{i,a}) + \eta_2(1-r_{i,a})g_{i,a} + \eta_3 r_{i,a}(1-g_{i,a}) \\ &+ \eta_4 r_{i,a}g_{i,a} + \eta_5 X_{i,a,f} + \eta_6 X_{i,a,p}\end{aligned} \tag{19}$$

Parameter	Estimate	Std. Error
η_1	-1.3952	0.9468
η_2	-1.2180	0.2576
η_3	-1.1079	1.3473
η_4	-1.3105	0.4923
η_5	-0.0168	0.0584
η_6	0.1521	0.1446

9.5 Job Destruction Rates

Each year, there is a probability than an old job is exogenously terminated. That job is no longer available to the agent. The probability this occurs was specified as:

$$\delta_{i,a} = \frac{1}{1 + \delta Z^{\delta}_{i,a}}$$

where:

$$\delta Z^{\delta}_{i,a} = \delta_1(1 - r_{i,a})(1 - g_{i,a}) + \delta_2(1 - r_{i,a})g_{i,a} + \delta_3 r_{i,a}(1 - g_{i,a}) + \delta_4 r_{i,a} g_{i,a}$$

Parameter	Estimate	Std. Error
δ_1	2.8862	0.6204
δ_2	3.4160	0.2982
δ_3	2.7415	1.0472
δ_4	3.2708	0.4300

9.6 A Note on the Gains to Job Switching

One caveat regarding the model and estimation should be noted. The model specified above makes the restrictive assumption that the gains to switching jobs lasts only one period. This restriction is non-trivial, since existing literature has shown a large fraction of wage growth comes from the returns to switching jobs. For example, in a study of young men, Topel and Ward (1992) show that during the first ten years of labor force participation one third of wage growth can be attributed to job changing. While incorporating returns to job switching that last the entire tenure of a job is possible without changing the model specified, it would add greatly to the computational burden of estimation. This is because introducing a permanent gain to job switches would require keeping track of the starting wage in the state space of the model. This could be accomplished by discretizing the starting wage to a limited number of points, expanding the state space by a factor of the number of points chosen. This extension is important but is left to future work for the time being. Because I don't allow for permanent increases in the wage through job switching, all wage growth in this model is attributed to job experience and tenure. If wage growth through job switching is important, this wage growth will incorrectly be attributed to tenure and experience.

Pension Participation and Uncovered Workers

Nadia Karamcheva
Boston College

Geoffrey Sanzenbacher*
Boston College

Abstract

Would extending tax-deferred pensions to uncovered, low income workers result in high rates of pension participation? Presently, 61 percent of low income workers eligible for a voluntary tax-deferred pension plan choose to participate. However, it is unclear whether individuals currently eligible for these plans are representative of low income workers in general. As the Obama administration considers extending tax-deferred pensions to uncovered workers, it is important to understand how likely these workers are to participate. This requires controlling for selection into jobs offering tax-deferred pensions, an exercise not yet attempted in the pension participation literature. In this paper, we control for selection into jobs offering tax-deferred pension plans using the fraction of jobs in a worker's state offering these plans as an exclusionary restriction. The results suggest workers currently offered a tax-deferred pension are more likely than otherwise similar individuals to participate. Thus, current estimates over predict the fraction of workers who would participate if voluntary retirement plans were extended to them. Our results indicate that ignoring selection overestimates the percent of all low income workers that would participate in a tax-deferred savings plan by 25 percent.

*Address: Department of Economics, Boston College, Chestnut Hill, MA, 02467. The authors would like to thank Norma Coe, Peter Gottschalk, Alicia Munnell, Shannon Seitz, and participants at Boston College's Labor Economics Lunch Seminars for their helpful comments. We would also like to thank Kelly Haverstick for her help obtaining the Form 5500 data. All mistakes are the authors' responsibility.

1 Introduction

Over the last thirty years, defined contribution (DC) pension plans have become the norm in the private sector, replacing more traditional defined benefit (DB) plans. Unlike DB plans, workers offered a DC plan are not required to participate. While a majority, 61 percent, of low-income workers eligible for a DC plan choose participation, this rate is substantially lower than for high-income individuals. Low participation rates, coupled with low pension coverage mean only a third of low-income workers are actually enrolled in a pension.[1] Low pension enrollment together with future declines in Social Security's generosity imply that 60 percent of low-income workers are at risk of retirement income that cannot maintain their standard of living.[2] Thus, low pension enrollment amongst low-income workers has become a concern of policymakers; the Obama administration has suggested requiring employers to automatically enroll employees in IRAs. Indeed, they cite majority participation among low-income workers as evidence that such a plan would be effective in ensuring pension coverage. However, caution should be taken when using the participation rates of workers currently covered to infer how workers not covered would respond to the policy change. If searching for a job is costly, then only workers who suspect they will participate in a DC pension would look for a job offering one. For low-income workers, for whom search costs and pension contributions represent a large share of their income, this selection effect may be especially strong. The current literature on pension participation amongst eligible workers does not control for this selection effect.

In departure from previous work on pension participation, we allow that workers offered a defined contribution plan may select into jobs offering plans based on an unobserved propensity to participate. To accomplish this we estimate two Heckman (1979) type selection models, one with participation in the plan as the outcome and one with the individual's contribution to the plan as the outcome. The first model has been referred to in the literature as bivariate probit with sample selection.[3] The second specification replaces the participation variable with the contribution rate of the worker to their defined contribution

[1] Karamcheva and Sanzenbacher (2010)

[2] Munnell et al. (2009)

[3] See Green (2008). This model consists of two latent variable equations, whose errors have joint normal distribution.

pension. A person's contribution rate is a censored continuous variable; the contribution rate is censored below at 0. We estimate this model using the two-step Heckman procedure where in addition we control for censoring of the dependent variable.[4] In both specifications, the ratio of defined contribution plans to defined benefit plans in an individual's state of residence is used as the required exclusionary restriction. Thus, the key assumption is that individuals do not move to a state to be more likely to obtain a defined contribution pension plan.[5] Our analysis indicates that ignoring selection has two consequences: (1) it overestimates the percent of non-covered workers who would participate in a DC type plan and (2) it underestimates the effect of income on participation and thus the gap in participation between low and high-income non-covered individuals.

Low participation in DC plans amongst eligible low-income workers is not surprising; participation removes income from the worker's pay check and requires decisions on risk allocation that low-income workers may have difficulty making. Previous literature has found a positive relationship between participation and income for workers offered a defined contribution plan. For example, Huberman et al. (2007) show that an increase in compensation of $10,000 leads to a 3.7 percent increase in the probability of participating in an offered defined contribution plan. Similarly, Bassett et al. (1998) find that a $1,000 increase in family income yields a 0.3 percent increase in the probablity of participation in an offered defined contribution plan. Munnell et al. (2009) utilize data from the Survey of Consumer Finances and find a similarly positive relationship between an individual's income and the decision to participate in an offered defined contribution pension. This can be interpreted as the effect of income on pension participation for the group of individuals offered defined contribution plans. This result may not be useful to policy makers who wonder how individuals not being offered plans respond when enrolled in an IRA.[6]

[4]In particular, we specify a probit model on whether an individual was offered a defined contribution plan, obtain the inverse mills ratio, and use this ratio as a regressor in a standard tobit model on the contribution rate.In this specification, standard errors are obtained by bootstrapping.

[5]This variable was obtained using the IRS Form 5500. See 5500-CRR Data, Panel of Current and Usable Form 5500 Data, Center for Retirement Research at Boston College, available at: http://crr.bc.edu/frequently_requested_data/data_on_the_form_5500_annual_reports.html

[6]A crucial difference between the Obama Administration's plan and employer provided DC plans is automatic enrollment. In the Administration's plan, the default is participation while for the employer provided plans we use for estimation the default is usually non-participation. Madrian and Shea (2001) examine how automatic enrollment increases participation rates. Our estimates will not account for this,

The results indicate that ignoring the selection into defined contribution jobs leads to estimates for the pension participation equation that do not apply to the population at large. In a probit on individuals offered a defined contribution plan, the marginal effect of log income on participation is 6.7 percent; once controlling for selection that estimate increases to 11.9 percent. Furthermore, the estimate of the correlation between unobservables in the selection equation and those leading to participation equation is positive. Thus, individuals not offered a defined contribution plan will be less likely to participate than observably similar individuals who were offered a plan. Our estimates indicate that a policy extending matched tax-deferred savings plan, similar to defined contribution plans, to all low-income individuals would result in participation rates of 42 percent. Without controlling for selection that estimate is much higher, around 56 percent.[7]

The paper proceeds as follows. In section 2, we describe recent trends in pension offer rates and participation across income groups. In section 3 we discuss the econometric specification we use to determine the effect of different factors on pension participation. In section 4 we discuss results and in section 5 we conclude.

2 Trends in Pension Access and Participation

Our interest, and the interest of policy makers, in the pension participation of low-income workers follows from the drop in pension participation for low-income workers that has accompanied the shift to voluntary DC pensions. Pension participation is the result of two events: 1) access to a retirement plan, and 2) enrollment in a plan. Figure 1 shows the share of individuals working for an employer that sponsors a plan over the last three decades. Plan sponsorship clearly differs by earnings group. Only about one-third of individuals in the bottom third work for an employer that sponsors a plan, compared with over 70 percent for the highest earnings group.[8] Overall, pension sponsorship has remained

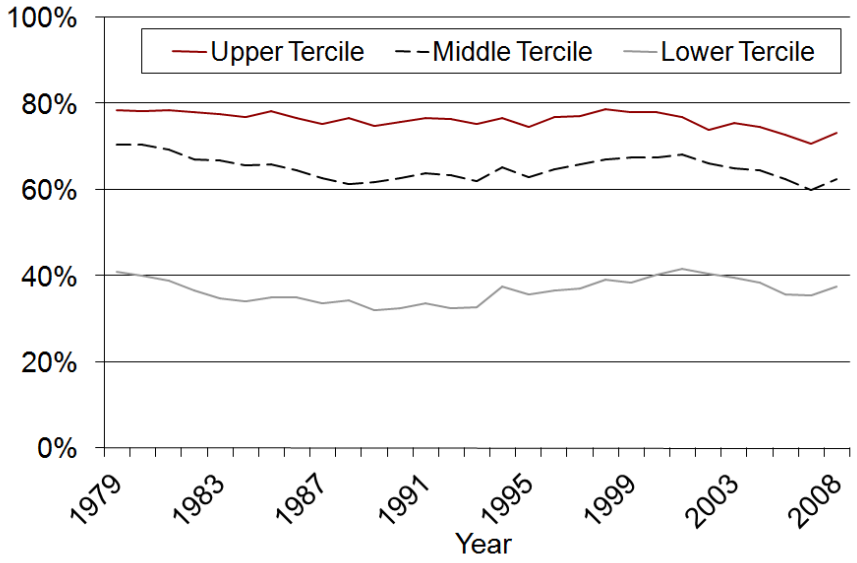
but the selection effect we find indicates current non-covered workers may be more likely to opt out of a plan.

[7]The 32 percent gure for low-income participation should be viewed as a lower bound on participation in the plan. As mentioned earlier, the purposed plan includes automatic enrollment which has been shown to increase participation. Our data does not include information on automatic enrollment. For example, see Madrian and Shea (2001).

[8]Earnings were defined as the reported monthly earnings on the first listed job.

relatively stable.

Figure 1: Pension Sponsorship, all Private Sector Male Workers Age 25-64, by Earnings Tercile, 1979-2008

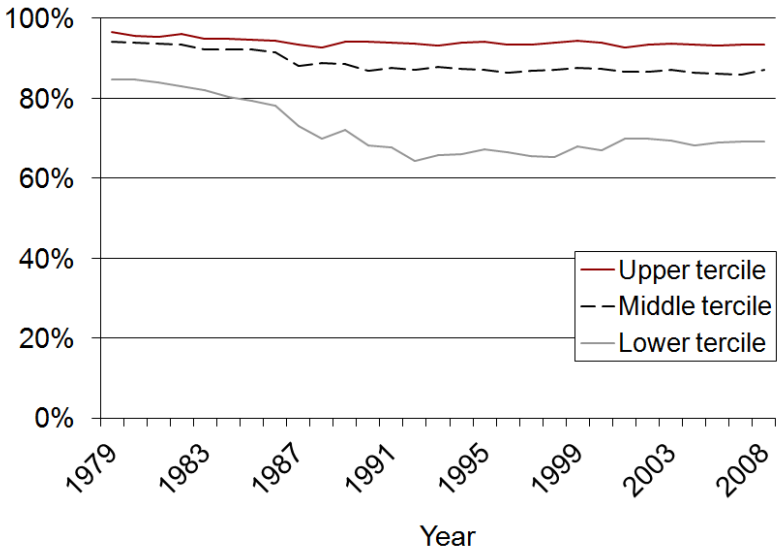


Source: Authors' calculations from U.S. Bureau of Labor Statistics, Current Population Survey (CPS), 1980-2008.

In contrast, the participation rates for workers whose employers provide a plan have shown considerable divergence among earnings groups over time (see Figure 2). While workers in the top third have had a nearly constant participation rate over the past 25 years, the rate for the middle third declined considerably – from 94 to 86 percent – and for the lowest third fell sharply – from 85 to 69 percent. These drops could be the result of a number of factors, ranging from ineligibility to misinformation about the plans to an inability to contribute due to budget constraints.

The data on pension access and participation together determine the overall participation rate, as shown in Figure 3. The biggest drops in overall participation occurred among middle and low earners, where the rate fell by 22 and 29 percent, respectively. Decreasing participation rates among low earners at sponsoring employers is the main driver of the group's overall decline in participation rather than any dramatic change in its access to pensions.

Figure 2: Pension Participation Rate for Private Sector Male Workers Age 25-64 at Employers with Pensions, by Earnings Tercile, 1979-2008

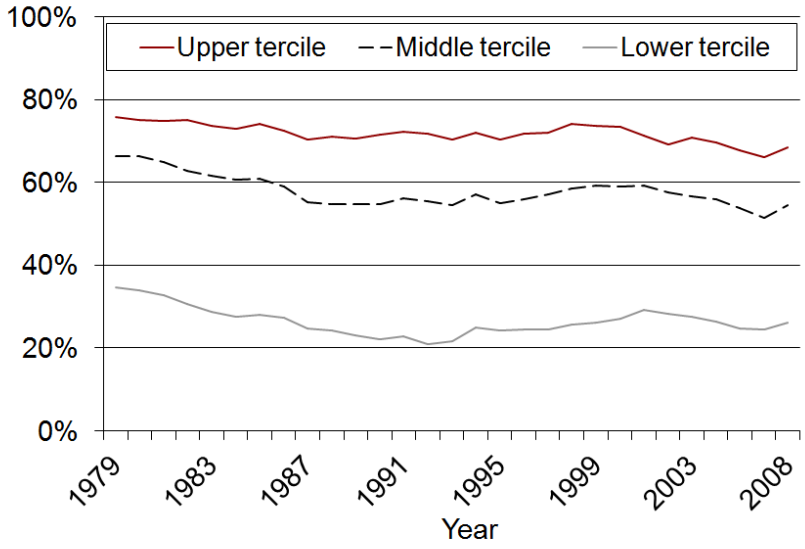


Source: Authors' calculations from 1980-2008 CPS.

These figures illustrate the importance of understanding the relationship between pension participation and income. The participation rate of low-income workers has dropped not because employers are offering them plans with less frequency but because the plans they are being offered are increasingly voluntary. Indeed, low-income workers have a variety of reasons for choosing not to participate in an offered pension. The most frequent reason is their low-income itself. While defined benefit plans are funded through employer contributions that do not directly decrease a worker's paycheck, participation in a defined contribution plan lowers a worker's take home pay. Thus, low-income workers may be less willing to trade off today's consumption for tomorrow's than a similar high-income individual. Low-income workers may be less likely to participate for other reasons, including lack of knowledge about the benefits of participation or inability to take the steps to enroll. This is evident when examining figure 4, which shows that 27 percent of low-income workers voluntarily decline participation in offered defined contribution plans, with 16 percent doing so because of their low-income.[9]

[9]An additional 33 percent of low income workers were ineligible for the offered defined contribution plan

Figure 3: Pension Participation Rate for Private Sector Male Workers Age 25-64, by Earnings Tercile, 1979-2008



Source: Authors' calculations from 1980-2008 CPS.

The concern of our study is that a voluntary non-participation rate of 27 percent should be viewed as a lower bound for low-income workers in general. Put another way, workers offered defined contribution plans may be the most likely to participate in such plans, having selected into a job that offers a mechanism for deferred compensation. If this is the case, policies aimed at providing voluntary savings plans amongst non-pensioned individuals would have lower participation rates than are seen in the pensioned population. In the next section, we present our empirical strategy for uncovering what determines pension participation.

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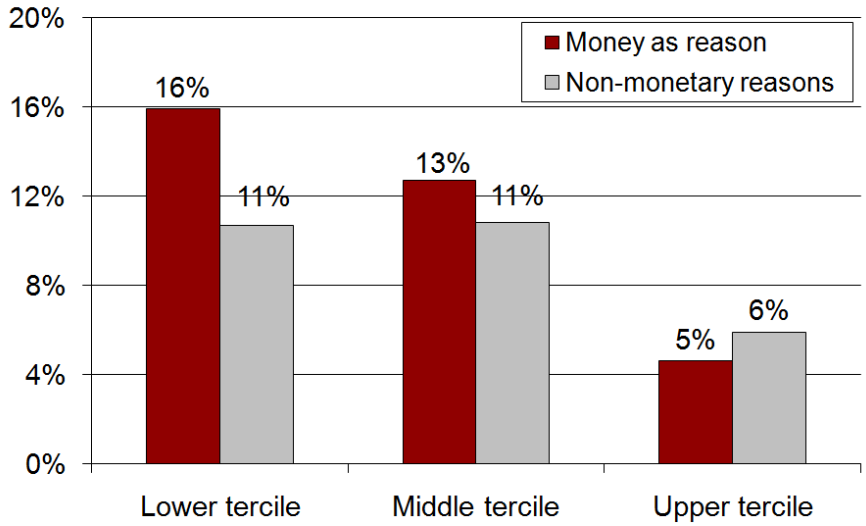
and thus did not voluntarily decline participation.

Figure 4: Percent of Eligible Private Sector Male Workers Age 25-64 Declining Defined Contribution (DC) Plans, 2007



Source: Authors' calculations from 2007 SIPP.

3 Empirical Strategy

3.1 Data

We use data from the 1996, 2001, and 2004 panels of the Survey of Income and Program Participation (SIPP). In each panel of the SIPP, workers were asked a topical module entitled "Retirement Expectations and Pension Plan Coverage." This topical module posed a series of questions on whether or not their present employer provided a pension, whether or not the individual participated in that pension, the type of pension the individual was offered, the individual's contribution rate if the pension was a defined contribution plan[10] and whether the employer provides a matching contribution.[11] This information, combined

[10]In practice, determining an individual's contribution rate required examining a series of variables. Some individuals responded to a question on their percent of earnings contributed to the defined contribution pension. For these individuals, this rate was defined as their contribution rate. Some individuals did not answer this question but instead gave a contribution amount and frequency. This was used together with their income to determine a contribution rate.

[11]In the 1996 panel, the "availability of employer match" question was not asked to non-participants in the DC plan. Instead, we imputed the missing values of the variable by using observations from the 2004

with the SIPP's core information on an individual's demographic characteristics and employer characteristics make the SIPP a good data set for estimating the relationships we have in mind.

For people offered a pension plan, we divide individuals into two categories: individuals offered a defined contribution plan and individuals offered a defined benefit plan. For workers who participated in their plan, individuals who claimed their benefit was based on earnings or years on the job are classified as defined benefit workers while workers who claimed they had an individual account plan are classified as defined contribution workers. If a worker chose not to participate in their plan a follow-up question asks if the plan they declined was a tax-deferred plan. If they answered "yes" to this question then they are classified as having been offered a defined contribution plan. Otherwise, we assume they were simply ineligible for an available defined benefit plan.[12]

Aside from pension plans, the core data of the SIPP provides information on individual and employer characteristics that are likely to be associated with pension offers and participation. This information includes an individual's age, race, education, marital status, state of residence, and child status. The data also include an individual's income from work and their net worth. On the employer side, the size of the worker's employer, whether or not the employee is covered by a union, the worker's tenure at their current firm, and the industry of employment were also obtained from the SIPP. Tables 1 through 4 present descriptive statistics of the workers in our sample. Table 1 examines pension coverage by type of plan across income groups. Overall 46 percent of the workers in our sample have access to a pension plan at their job.[13] About 26 percent of those plans are Defined Contribution and 20 percent are Defined Benefit. In addition, pension coverage increases by income terciles

and 2007 panel and STATA's hotdeck routine. Hotdeck stochastically imputes observations by matching individuals on user-specified variables. The ones that we used included firm size, industry, union status and the ratio of DC to DB plans in worker's state. More info about "hotdeck" can be found at: http://ideas.repec.org/c/boc/bocode/s366901.html

[12]In general, workers cannot decline participation in a defined benefit plan if they are eligible. Thus, we assume if they did not participate in an offered defined benefit plan that they were ineligible for that plan.

[13]Notice that this number is lower than what other studies usually report because our definition of "offered a pension plan" requires sponsorship on the part of the employer *and* worker eligibility. In that respect, workers who are currently not eligible to participate in the plan (for whatever reason) are classified as not being offered one. This definition is necessitated by the fact that in our estimation we want to explore the decision to voluntary participate in plan - which is only viable if the worker is already eligible.

- 19.3 percent for the bottom income group compared with 69.7 percent for the top one.

Table 2 focuses on workers who have access to a defined contribution pension and examines their conditional contribution and participation rates. Among workers who are already with a firm that provides a DC type plan, 79.7 percent choose to participate - 62.2 percent for those in the lowest income tercile, compared with 89 percent for those in the highest. Even though participation rates correlate highly with income, once we control for participation, contribution rates as a percent of salary do not seem to have a clear relationship with income. Among workers already participating (i.e. contributing a positive amount), the mean contribution rate is 7.4 percent; 8.8 percent for the bottom income tercile compared with 7.4 percent for the top income tercile.[14]

Our intuition is that workers who are currently at DC sponsoring jobs are potentially different from those who are not, due to unobservable differences in tastes or constraints, which makes them more or less likely to partipate in an offered DC plan. They are also different in terms of observable characteristics. Tables 3 and 4 compare how different these groups of workers are in terms of observable characteristics. Workers in our sample who have access to DC plans are older, more likely to be married, more likely to be white, and have more education than other workers. In addition, their mean annual income is about $15,000 higher and their net worth is $18,000 higher than that of workers who currently don't have access to a DC plan. In addition, those workers who decide to participate in the DC plan are more likely to be male and married, are older, have more education and have longer tenure with their current employer compared with workers who choose to not participate. They also have higher income, wealth and net worth than the non-participating group. In addition, 87 percent of the participating workers have an employer match at the job, compared to 79 percent of those not participating. This is consistent with previous studies that find a positive link between the existence of employer match and the likelihood of participation.

[14]We examined the 401(k) participation and contribution rates by income brackets as well, even though we did not include them in the paper. The results are largely consistent with previous literature that uses SIPP data. See for example Engen et al. (1994), Poterba et al. (1994) and Poterba et al. (1995). It should be noted that overall 401(k) eligibility and conditional participation rates are higher in our sample as compared to what those previous studies report. This could be due to the fact that we focus on individuals, rather than families and that our data comes from more recent panels of the SIPP when DC plans have already become the most popular type of employer-provide pensions.

Finally, we use plan level data to obtain information on the proportion across states of defined contribution versus defined benefit plans that employers provide. The data come from the IRS Form 5500. The Form 5500 is a tax form filled out by employers who offer a pension plan. Importantly, the form asks for characteristics of the plan that can be used to determine whether a plan was defined benefit or defined contribution. The form also asks how many workers are covered by the plan and how many individuals work at the offering firm. We use this information to determine the fraction of individuals offered a pension plan who were offered a defined contribution plan by state.[15] This ratio differs by state and over time.

3.2 Modeling the Participation Decision

The goal of our empirical framework is to derive consistent estimates of the effect of income (and other factors) on DC participation. This will inform us as to what DC participation rates to expect as a result of a policy changes extending coverage. We also want to know how the policy effect will differ by income groups. In order to achieve these two goals we need an empirical specification which controls for the possible self-selection of workers into jobs with pension plans. Our hypothesis is that workers who are currently in DC sponsoring jobs differ systematically and in unobservable ways from workers not offered a DC plan. Failure to control for selection into jobs providing defined contribution plans will lead to parameter estimates of the decision to participate that do not apply to the average individual and will lead to biased conclusions in terms of policy effects.

The decision to participate in an offered 401(k) plan is discrete. Whether the worker is offered a 401(k) plan is also a discrete outcome. In addition, we want to allow the unobservable that affects the probability of being offered a 401(k) plan to be correlated with the unobservable that affects the worker's decision to choose participation. The empirical setup that achives both is a bivariate probit model with sample selection as described in Green (2008). This formulation was first presented by Van de Ven and Van Pragg (1981)

[15] In practice, we could obtain similar information from the SIPP. However, the sample size of the SIPP makes getting an estimate of the defined contribution ratio difficult for some states.

and applied to our question of interest has the following basic set up:

$$p_1 = \Pr[\text{Offered DC} = 1] = \Phi(\mathbf{z}'\gamma) \quad (1)$$

$$p_2 = \Pr[\text{Participate} = 1|\text{Offered DC} = 1] = \Phi(\mathbf{x}'\beta) \quad (2)$$

where

$$\mathbf{x} = \{\text{demographics, tenure at current job, annual income, wealth, ...}\}$$

$$\mathbf{z} = \{\text{demographics, annual income, wealth, } \textit{union status, DC ratio by state, ...}\}$$

We make the usual assumption that the errors in the two equations have the standard normal distribution and ρ is a correlation parameter denoting the extent to which the two error covary. When $\rho \neq 0$, standard probit techniques applied to equation (2) yield biased results. To achieve consistent and asymptotically efficient estimate of β we need to account for the sample selection. In addition, for the model to be identified, the selection equation should have at least one variable that is excluded from the outcome equation. Otherwise, the model is identified only by functional form, and the coefficients have no structural interpretation. The exclusionary restriction should alter the probability an individual is offered a defined contribution pension plan but not the probability he participates in that plan. In our empirical specification union status and the proportion of defined contribution to define benefit plans as offered by employers in the given state, serve as exclusionary restrictions. The identifying assumption is that workers do not move into states because they want a better/worse chance of being offered a defined contribution pension plan and that being at an union job affects one's probability of having access to a 401(k) plan but not his propensity to save. We believe that both of these assumptions are reasonable as long as they capture variation in the availability of 401(k) plans coming form the employer side and are exogenous factors in the workers' saving decisions. On the other hand, one could argue that given enough time employers would respond to changes in the environment (such as the introduction of new government policy) and it would be no longer appropriate to assume such factors are exogenous. Because our empirical specification is not able to capture such dynamics, it is advisable to interpret the results as indicative of short-term

rather than long-term effects of the policy.

3.3 Modeling the Contribution Rate

From a policy point of view it is interesting to determine what effect the extension of 401(k) plan availability will have on the contribution rates of people who decide to participate. In our second empirical specification, we are interested in how much the employees chose to contribute to their 401(k) plan. Similarly to before, we need to control for sample selection. In addition, we need to take into account the fact that even though continuous, the outcome variable is censored from below at 0. This is due to the fact that workers cannot choose to contribute less than 0 percent. In the absence of censoring, it would have been appropriate to use the standard Heckman 2-step selection model. Ignoring the censoring, however, would lead to biased results.

Instead we refer to our approach as "censored regression with sample selection". We proceed in the following way:

Step 1: Estimate a probit equation on being offered a 401(k) plan, using union status and state-level DC to DB ratio as exclusionary restricitons:

$$\begin{aligned} p &= \Pr[\text{Offered DC} = 1] = \Phi(\mathbf{z}\gamma) \qquad (3) \\ \mathbf{z} &= \{\text{demographics, annual income, wealth, } \textit{union status, DC ratio by state, ...}\} \end{aligned}$$

Then compute the estimate of the inverse Mills ratio $\lambda(\mathbf{z}'\widehat{\gamma}) = \frac{\phi(\mathbf{z}'\widehat{\gamma})}{\Phi(\mathbf{z}'\widehat{\gamma})}$

Step 2: Via maximum likelihood estimate a Tobit[16] model with censoring from below at 0, where the latent variable is linear in regressors with an additive error; the inverse Mills ratio is added as one of the regressors:

$$\textit{Contribution Rate}^* = \mathbf{x}'\beta + \sigma_{12}\lambda(\mathbf{z}'\widehat{\gamma}) + \epsilon_2$$

$\textit{Contribution Rate} = \max(0, \textit{Contribution Rate}^*)$ if Offered DC$= 1$

$\textit{Contribution Rate} = .$ if Offered DC$= 0$

where

$\mathbf{x} = \{$demographics, tenure at current job, annual income, wealth, ...$\}$

[16] See Tobin (1958).

Contribution Rate* is the worker's desired contribution rate and *Contribution Rate* is the one observed in the data, which is missing for individuals not offered a 401(k) plan. The coefficient in front of $\lambda(\mathbf{z}'\widehat{\gamma})$ is the covariance between the errors in the selection and outcome equations. A test of whether $\sigma_{12} = 0$ is a test of whether or not the errors are correlated and sample selection correction is needed. Since the the tobit model does not take into account the estimation error in deriving $\widehat{\lambda}$, the standard errors would be incorrect. We derive the correct standard errors by bootstrapping 10,000 times.[17] Table 8 compares results achieved by running a regression on the contribution rate ignoring selection, a censored regression model, and a selection model with censoring.

4 Results

4.1 Participation Decision

Table 5 compares estimation of a probit model on participation without controlling for selection and while controlling for selection, the "Heckman Probit." The right hand column provides the estimates of the selection equation.

The results of the probit and the probit with selection are largely consistent with the literature with respect to sign and significance. Similarly to previous studies,[18] we find that individuals who are married, well educated, have high tenure at their firm, and work at firms with a employer match[19] are all more likely to participate than others. Blacks

[17]We do the estimation and the bootstrapping in STATA. We use STATA's 'tobit' command which we imbed in an user-written bootstrap routine. In essence we sample with replacement 10,000 times and estimate the model on each of these sub-samples of the data. The variation in the estimates gives us our confidence intervals.

[18]See for example Munnell et al. (2009) and Bassett et al. (1998).

[19]Most previous studies find that employees respond positively to the existence of an employer match. Munnell et al. (2009) find a significant positive effect of the employer match on contribution rates, although the relationship is concave with respect to the size of the match. Similarly, Bassett et al. (1998) find that workers with employer matches are more likely to participate in 401(k) plans than workers without such matches. No evidence is found that the level of the employer match has a positive impact on employee participation, however. Kusko et al. (1998) found little change in either participation or contributions in response to large changes over time in matching provisions. Papke (1995) showed that participation increases with the level of the match rate, with smaller marginal effects at higher match rates, and that contributions increase markedly as the employer moves from a zero to a positive match rate, with a negative effect at very high match rates. Papke and Poterba (1995) concluded that participation increases with the match rate but found no significant effect on contributions. Clark and Schieber (1998) found a positive

and younger individuals are less likely to participate. Also consistent with expectations, individuals with high networth and high-income are more likely to participate in an offered defined contribution plan than other individuals.

These results are not in any way surprising - previous literature has already documented similar (in terms of direction and significance) relationships between the independent variables and 401(k) plan participation. What is interesting is the change in the magnitudes of the effects once we control for selection.

Our intuition led us to two testable implications: (1) workers' selection on unobservables into defined contribution jobs is positively correlated with unobservables related to participation; and (2) the effect of income on participation without controlling for selection will be underestimated. The estimates suggest these implications hold in the data. The estimated correlation between unobservables in the outcome and participation equations is .251. A likelihood ratio test shows this correlation is significantly different from 0. This suggests individuals in defined contribution jobs are more likely to participate than similar individuals in jobs not offering these pension plans. The effect of income is also smaller in the equation estimated without controls for selection. In the simple probit model the marginal effect of log income on pension participation is 6.7 percent, when controlling for selection it is 11.9 percent. This shows us that in the population low-income workers are much less likely than high-income workers to participate in a defined contribution type pension plan. However, in the selected sample this difference is mitigated by selection into the plan. If one is to predict the effect of policies targeted at the whole population but ignores the selection effect, he is likely to understate the importance of income and the disparity of the policy effect among income groups.

Let us illustrate this point with the recently proposed government plan to extend DC-type coverage to workers not currently covered by a plan. Our estimates indicate that the predicted participation rate for this population taken from the current literature would be an overestimate. Tables 6 and 7 indicate the size and magnitude of this overestimation under two different assumptions: (1) all defined contribution plans offer a match (employer and government) at the same rate as the average firm in our sample, and (2) none of the

effect of the match rate on both participation and contributions, but their sample contained no firms without a match rate.

defined contribution plans provide a match.[20]

Table 6 indicates that if matched defined contribution plans were provided to all individuals a standard probit would suggest that 74.5 percent of individuals would participate.[21] This number is lower than the percent that participate in actual offered plans only because of observable differences between those offered and those not offered a defined contribution pension. Once controlling for selection, the predicted participation rate drops to 65.0 percent. This drop occurs because the group of people not already offered plans is substantially less likely to participate based on unobserved characteristics. This factor is not captured by the probit alone. If we examine the low-income group, the group commonly associated with the extension of defined contribution plans, we see an even larger difference. Under the probit estimation, 56.0 percent of individuals in the lower tercile of the income distribution would participate if all were offered a defined contribution pension plan. Once controlling for selection, this drops to 42.4 percent, a drop of 13.6 percentage points. In addition, the difference is larger for low-income individuals than individuals in the middle and higher terciles. For those groups the difference is 9.9 and 5.5 percentage points respectively. Ignoring selection seems to have especially large ramifications for low-income workers, the group most likely to be affected by any potential effort to expand coverage. In addition, ignoring selection would make us expect a smaller gap in participation rates between low and high-income groups. Specifically, the probit model would suggest a difference of 32.4 percentage points in participation rates after the expansion of the policy. In contrast, when controlling for selection would expect a much higher pension inequality of 40.5 percentage points.

4.2 Contribution Decision

Our second model uses information on contribution rates instead of on participation. In this model, an individual offered a defined contribution plan participated if they contributed greater than 0 percent of their income. If they contributed nothing, then they were classified

[20]In President Obama's 2010 State of the Union Address the plans were proposed to be similar to IRAs and thus would not provide a match.

[21]The differences discussed below are similar if the plans are not matched. These differences can be calculated from table 7.

as a non-participant. Thus, the contribution rate is a censored dependent variable that is the result of selection into a defined contribution job. The results are presented in table 8. The results from a simple OLS, a standard tobit, and a tobit controlling for selection are included in the table.

The results in table 8 are largely consistent with the results of the participation model. An individual's age, the existence of an employer match, and a higher level of education are all associated with higher contribution rates. On the other hand, children and being black are associated with lower contribution rates. Individuals with higher income and with higher net worth have higher contribution rates than others. The positive and significant coefficient on the inverse Mills ratio indicates a positive correlation between unobservables in the selection and contribution equations. As in the previous model, the effect of income is underestimated in the equations not controlling for selection. In the OLS estimation, the effect of log income is actually negative, in the tobit without selection it is insignificant, in the tobit controlling for selection it is positive and significant. The analysis of the contribution rates confirms the analysis of the participation equations: there is positive selection into defined contribution jobs and ignoring this selection leads to underestimation of the effect of income on pension participation. It also provides some insight on contribution rates should defined contribution plans be extended to more individuals. Controlling for selection in the participation equation revealed that basing estimates of expected participation rates on those currently offered pensions will overestimate the overall participation rate. A similar result holds for pension contribution rates, as can be seen in tables 9 and 10. We will focus on estimates based on the assumption that all defined contribution plans, including those provided by the government, are matched at the average rate. The standard tobit predicts that the average participating individual from the population would contribute 4.6 percent of their income to their defined contribution plan. Once controlling for selection, the contribution percentage drops to 4.3 percent. This decline is of similar magnitude across income groups.[22]

The results from the participation and contribution equations suggest that ignoring se-

[22]The tobit with selection yielded a percent participating that was likely too low. Only 41 percent of individuals had a positive predicted contribution rate above 0 percent, which is well below 57.5 percent suggested by the probit.

lection will have two effects: (1) it will overestimate the participation rate and overestimate it the most for low-income workers; and (2) it will overestimate contribution rates similarly for all income groups. The implication of these results should be considered when making policies that extend defined contribution plans to low-income individuals. The estimates suggest both participation rates and contribution rates will be lower than is suggested by the current literature.

5 Conclusion

As defined contribution plans have expanded over the last three decades, pension participation amongst low-income individuals has fallen more than for any other group. This decline has been driven by the voluntary decision of many low-income individuals to decline participation in offered pension plans. This decline, combined with perceived future decreases in Social Security's generosity, have led many to believe extending tax-deferred savings plans to low-income individuals is a way to ensure them an acceptable retirement income. Given that 60 percent of eligible low-income workers offered a defined contribution pension participate in the plan, on the surface at least, this seems like an effective strategy. Our analysis suggests this picture may be too optimistic.

Our intution suggested, and our estimation confirms, that workers may select into defined contribution jobs based on some unobserved propensity to participate in the plan. The implication of this selection is that individuals not currently at jobs offering defined contribution plans may be especially unlikely to contribute to a tax-deferred savings plan. This may be especially true of the low-income population for whom the selection effect may be strong. Our estimates suggest that the selection effect is non-trivial. Controlling for selection leads to the prediction that only 42 percent of individuals in the lower income tercile are likely to participate in an offered tax-deferred savings plan. Those who do participate are likely to contribute 3.1 percent of their income to the plan. Offering these kinds of savings vehicles to individuals not covered by a pension plan may be helpful to those who participate. However, by controlling for the selection of workers into pensioned jobs we believe our estimates show that fewer individuals will participate than policymakers

might hope.

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Table 1: Percent of Workers with Pension Coverage by Type of Plan and by Income

	Offered a Pension*	DC type	DB type
All	46.0%	25.9%	20.1%
Bottom Income Tercile	19.3%	11.0%	8.3%
Middle Income Tercile	49.0%	28.5%	20.5%
Top Income Tercile	69.7%	38.3%	31.4%

* Where'offered a pension' means the worker is with an employer who sponsors a pension plan and the worker is eligible to participate

Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 2: 401(k) Sponsorship, Participation and Contribution Rates by Income

	DC Sponsorhip Rate	Conditional Participation Rate	Conditional Contribution Rate
All	25.9%	79.7%	7.4%
Bottom Income Tercile	11.0%	62.2%	8.8%
Middle Income Tercile	28.5%	74.0%	6.9%
Top Income Tercile	38.3%	89.0%	7.4%

Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 3: Characteristics of Workers by 401(k) Sponsorship

Variable	Not Offered DC	Offered DC
Female	48.8%	46.6%
% Married	52.3%	63.1%
% with Children	44.3%	43.9%
% White	82.7%	85.9%
% Black	11.6%	8.5%
Age	37.4	40.3
Years of Education	12.9	13.8
Tenure	6.1	7.7
Mean Income	$28,522	$43,783
Mean Wealth	$157,167	$176,621
Mean Net Worth	$148,434	$166,831

Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 4: Characteristics of Workers Offered 401(k) Plans by Participation Status

Variable	Do Not Participate	Participate
Female	53.8%	44.7%
% Married	49.5%	66.6%
% with Children	43.6%	43.9%
% White	80.1%	87.4%
% Black	13.7%	7.2%
Age	36.5	41.2
Years of Education	13.1	14.0
Tenure	4.3	8.5
Mean Income	$28,769	$47,609
Mean Wealth	$102,333	$ 195,552
Mean Net Worth	$92,620	$185,741
Employer Provides a Match	79.0%	87.1%

Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 5: 401(k) Participation Models

	Probit			Probit with Selection				
				Outcome eq-n			Sel. eq-n	
Variable	Coeff	SE	Marg Prob	Coeff	SE	Marg Prob	Coeff	SE
Female	-.069**	.023	-.017**	-.050	.023	-.018*	.068**	.011
Age	.119**	.032	.029**	.171**	.033	.063**	.288**	.014
Age^2	-.002**	.001	-.001**	-.003**	.001	-.001**	-.007**	.000
White	.060	.046	.015	.066	.045	.025	.059**	.023
Black	-.147*	.057	-.038*	-.158**	.055	-.059**	-.059**	.028
Married	.112**	.025	.028**	.112**	.025	.041**	.012	.013
Has kids	-.042	.025	-.010	-.050*	.025	-.018*	-.042**	.012
Years educ	.046**	.005	.011**	.048**	.005	.018**	.015**	.002
Tenure	.040**	.002	.010**	.038**	.002	.014**		
DB at old job	.039	.064	.010	.056	.063	.020	.085**	.030
DC at old job	.008	.053	.002	.050	.053	.018	.230**	.028
Provides match	.406**	.029	.106**	.405**	.028	.149**		
Log Networth	.075**	.006	.019**	.076**	.006	.029**	.015**	.003
Year 2004	-.013	.030	-.002	.003	.030	.001	.059**	.014
Year 2007	-.243**	.027	-.061**	-.227**	.027	-.084**	-0.059**	.014
Log Income	.274**	.018	.067**	.324**	.020	.119**	.278**	.008
DC ratio							1.952**	.079
Union Status							-.210**	.018
ρ							.251	.053
LR test of indep. eqns. ($\rho = 0$): chi2(1) = 19.99 Prob > chi2 = 0.0000								
N uncensored				19,582				
N censored	19,582			53,941				
N all	19,582			73,523				

* Significantly different from 0 at .05 level

** Significantly different from 0 at .01 level

Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 6: Predicted Probabilities of Participation in 401(k) if All 401(k) Plans Provide and Employer Match

	Probit	Probit with Selection
All	74.5%	65.0%
Bottom Income Tercile	56.0%	42.4%
Middle Income Tercile	75.6%	65.7%
Top Income Tercile	88.4%	82.9%

Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 7: Predicted Probabilities of Participation in 401(k) if No 401(k) Plans Provide an Employer Match

	Probit	Probit with Selection
All	64.7%	52.5%
Bottom Income Tercile	41.8%	52.5%
Middle Income Tercile	62.8%	29.4%
Top Income Tercile	79.7%	72.2%

Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 8: 401(k) Contribution Rate Models

	Least Squares		Censored Regression		Censored Regression with Sample Selection			
					Outcome eq-n		Sel. eq-n	
Variable	Coeff	SE	Coeff	SE	Coeff	SE	Coeff	SE
Female	-.616**	.093	-.703**	.120	-.710*	.120	.066**	.011
Age	.557**	.138	1.338**	.194	1.315**	.193	.283**	.014
Age^2	-.011**	.003	-.028**	.005	-.028**	.005	-.006**	.000
White	-.054	.191	-.024	.244	-.001	.244	.065**	.023
Black	-.885**	.240	-1.311**	.310	-1.304**	.310	-.054**	.028
Married	.521**	.105	.733**	.134	.732**	.134	.015	.012
Has kids	-.664**	.102	-.925**	.130	-.921**	.130	-.042**	.012
Years educ	.227**	.021	.278**	.027	.277**	.027	.016**	.002
Tenure	.070**	.007	.106**	.008	.106**	.008		
DB at old job	.075	.241	.086	.303	.076	.303	.090**	.030
DC at old job	.524**	.205	.621*	.262	.603**	.262	.239**	.028
Provides match	.727**	.124	1.419**	.162	1.415**	.162		
Log Networth	.354**	.025	.530**	.033	.529**	.033	.014**	.003
Year 2004	.074	.118	.091	.150	.085	.150	.055**	.014
Year 2007	-.703**	.111	-1.118**	.142	-1.122**	.142	-.064**	.014
Log Income	-.302**	.077	.193	.105	.170*	.117	.274**	.008
DC ratio							1.997*	.078
Union Status							-.210**	.018
ρ					.053*	.043		
N uncensored			12,757		12,757			
N left-censored			3,941		3,941			
N all	16,698		16,698		16,698		73,785	

* Significantly different from 0 at .05 level

** Significantly different from 0 at .01 level

Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 9: Predicted Conditional Contribution Rates* in 401(k) plans if All 401(k) Plans Provide an Employer Match

	Least Squares	Censored Regression	Censored Regression with Sample Selection
All	5.36%	4.59%	4.31%
Bottom Income Tercile	4.42%	3.30%	3.06%
Middle Income Tercile	5.20%	4.13%	3.80%
Top Income Tercile	6.27%	5.79%	5.46%

* Mean percent of salary contributed to 401(k) plan for workers participating in the plan.
Source: Authors' calculations, based on data from the Survey of Income and Program Participation.

Table 10: Predicted Conditional Contribution Rates* in 401(k) plans if No 401(k) Plans Provide an Employer Match

	Least Squares	Censored Regression	Censored Regression with Sample Selection
All	4.67%	3.60%	3.28%
Bottom Income Tercile	3.77%	2.59%	2.32%
Middle Income Tercile	4.50%	3.11%	2.75%
Top Income Tercile	5.55%	4.50%	4.11%

* Mean percent of salary contributed to 401(k) plan for workers participating in the plan.
Source: Authors' calculations, based on data from the Survey of Income and Program Participation.