A RANDOM EFFECTS LOGIT MODEL OF WORK-WELFARE TRANSITIONS

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Early studies of welfare duration (Bane and Ellwood, 1983; and Plotnick, 1983) showed that the majority of new spells on welfare lasted less than two years, thus calling into question the popular view that welfare was a trap from which most recipients never escaped. On the other hand, these studies also found that a substantial minority of recipients received assistance for protracted periods and made up a majority of the caseload at any moment in time.

This paper adds a new dimension to welfare dynamics by focusing on transitions in and out of work for welfare recipients. We develop a competing risk, discrete time duration model of transitions between three states: receiving Aid to Families with Dependent Children (AFDC) while not working, receiving AFDC while working, and not receiving AFDC. A random effects multinomial logit model for panel data is applied to data from administrative records from the Wisconsin Computerized Reporting Network (CRN).

The contribution of this paper is both substantive and methodological. Substantively, we provide information useful to policymakers who have recently stressed work oriented approaches to welfare reform. A central issue in the debate is the degree to which persons who continue

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to receive AFDC move in and out of work. Methodologically, we follow the descriptive tradition of earlier studies. However, we provide a much richer specification of the statistical model generating the data.

Figure 1 illustrates the six transitions which fully describe the work/welfare dynamics. A person receiving AFDC starts from one of two states--working or not working. From either of these states the recipient may leave welfare or change work status. Likewise, a person not receiving AFDC may return to welfare, either to welfare/work or to welfare/no work. Previous studies, which have focused almost solely on exits from welfare, have collapsed two transitions (from welfare/work and welfare/no work to no welfare) and ignored an additional four.

In this paper we focus specifically on exits and entries into the state of work and welfare. While relatively few AFDC recipients work at any moment in time (15 percent in Wisconsin in June, 1985) the hazard of leaving AFDC is considerably higher if the recipient is working. Furthermore, changes in federal law enacted in 1981 severely reduced the benefits to working while receiving aid. In this paper we ask whether those changes affected the hazard of starting or ending a spell of welfare and work.

In this paper we first review the institutional background and the methodology for estimating a discrete-time, competing-risk model with unobserved heterogeneity. We then describe the data, which we use in the following section to estimate the model. The final section summarizes our results.

INSTITUTIONAL BACKGROUND

In 1966 Congress passed the so-called $30 and a third provision which allowed AFDC recipients to keep the first $30 they earned in every month plus a third of all additional earnings. This change was instituted to reduce the statutory benefit reduction rate on earnings for welfare recipients from 100 percent to 67 percent in order to encourage work effort.

In 1981 the Congress passed the Omnibus Reconciliation Act (OBRA) which included major changes in the AFDC eligibility rules. Some rule changes, such as the counting of step parent income, led to immediate administrative termination of cases. Others led not only to initial case closings but also changes in the rules affecting future behavior. One of the key changes in the rules was to limit the $30 and a third provision to the first four months of a work spell. This change was instituted in order to target welfare payments to recipients with no other sources of income. The result was that the total income of an AFDC recipient would be no higher whether she worked or not in the fifth or subsequent months of a work spell.

A purely static model of labor supply would predict that no AFDC recipient would work more than four months after the passage of OBRA. This would lead to a jump in the hazard rates of leaving welfare and of quitting work in month five. However, longer term considerations and the ease with which women can adjust their labor market behavior affect the response to the jump in the tax rate that occurs in the fifth month. For example, if the value of keeping a job includes expected future wage increases, some women might be willing to bear the costs of work related expenses although all current earnings are taxed away. Caseworkers may

also impose implicit sanctions on welfare recipients who quit their job, since work lowers costs to the agency even if it does not raise the welfare recipient's net income.

STATISTICAL MODEL

General Framework. The statistical model we estimate is a multinomial logit model with a two factor heterogenity distribution, each with two points of support. Its objective is to model the monthly sequence of observations on states and observed covariates, which are given by

$$(\{S_{it}, X_{it}\}, t = 1, \ldots, T)$$

where S_{it} indicates the state of person i in month t and X_{it} is a sequence of covariates. X can include spell duration and time-invariant and time-varying variables which are strictly exogeneous. In our application, S_{it} can take three values:

S_{it} = 1 if receiving welfare and receive no earnings
2 if receiving welfare and receive earnings
3 if not receiving welfare

The probability of person i making a transition from state j to state k is assumed to follow a Markov process which depends on a vector of observable covariates, X_{it}, and a vector of unobserved, time invariant, person specific random effects, Z_i, which are assumed to be independent of X_{it}.

Let $\Pr(S_{it+1} \mid S_{it}, X_{it}, Z_i, \theta)$ be the probability of person i making a transition in period t from state S_{it} to S_{it+1}, conditional on observables and unobservables, where θ is a vector of parameters to be

estimated. The probability of person i experiencing the sequence of states $\{S_{it}\}$ is therefore given by

$$(1) \qquad P_{it} = \prod_{t=1}^{T-1} \Pr(S_{it+1} \mid S_{it}, X_{it}, Z_i, \theta)$$

which is the person's contribution to the likelihood function.

Since Z_i is unobserved it must be integrated out of the likelihood function. Let f(Z) be the probability that Z_i equals Z and assume Z takes R discrete values. The likelihood for person i is, therefore,

$$(2) \qquad L_i(\theta) = \prod_{t=1}^{T} \sum_{r=1}^{R} f(Z_r) \Pr(S_{it+1} \mid S_{it}, X_{it}, Z_r, \theta)$$

Given a random sample of N individuals the likelihood of the sample is

$$(3) \qquad L = \prod_{t=1}^{N} L_i(\theta)$$

which can be maximized through numerical methods over θ.

Specific Functional Form Assumptions. In order to estimate θ we specify the distribution of the unobserved heterogeneity and specify a functional form for the conditional probability of exiting. Recent developments in the literature on single risk duration analysis (Heckman and Singer, 1984) have explored the inherent tradeoff between the parametrization of duration dependence and unobserved heterogenity. They demonstrate that a duration model can be identified by parameterizing either duration dependence or the unobserved heterogeneity distribution. We choose a simple parametric form for the distribution of unobserved

heterogeneity. This allows us to examine the correlation between the heterogeneity components for the transitions out of any given state. We approximate nonparametric duration dependence with a sequence of dummies in the early months of a spell. After two years in a state, we constrain the hazard to not exhibit any further duration dependence.[1]

For the unobserved heterogeneity distribution we assume a two factor model, each with two points of support. Let Z_{ijk} and Z_{ijl} be the values of the unobservables affecting the probability that person i will make a transition from state j to state k or from state j to state l, respectively. Each can take two values:

$$Z_{ijk} \in \{Z^1_{jk}, Z^2_{jk}\}$$

$$Z_{ijl} \in \{Z^1_{jl}, Z^2_{jl}\}$$

The probability of the realizations of these two random varaibles is given by:

$$(4) \qquad \text{Prob}(Z_{ijk} = Z^h_{jk},\ Z_{ijl} = Z^g_{jl}) = \pi^{hg}_j \quad \text{for} \quad h,g = 1,2$$

$$\text{where} \quad \sum_{h=1}^{2} \sum_{g=1}^{2} \pi^{hg}_j = 1 \quad \text{for } j = 1,2,3$$

We place no restriction on the covariance between Z_{ijk} and Z_{ijl}.

For the conditional probability of making a transition from state j to state k (conditional on X_{it}, Z_{ijk} and Z_{ijl}) shown in equation (1), we assume a multinomial logistic distribution:

$$\text{Prob}(S_{it+1} = k \mid S_{it} = j,\ X_{it},\ Z_{ijk},\ Z_{ijl},\ \theta) \equiv P_{it}(k \mid j)$$

$$= \frac{[\exp(X_{it}\beta_{jk} + Z_{ijk})]^{I(k \neq j)}}{1 + \sum_{k \neq j} \exp(X_{it}\beta_{jkt} + Z_{ijk})} \qquad j,\ k = 1,2,3$$

where $I(k \neq j)$ is an indicator function, equal to one if k is not equal to j and equal to zero otherwise.[2]

With this statistical framework we can estimate the conditional probability of leaving any one of our three states. In the following section we describe the data used to accomplish this task.

DATA

The data come from the Computer Reporting Network (CRN) of the State of Wisconsin. This computerized system, which became fully operational at the end of 1980, contains detailed information on all AFDC recipients in the State of Wisconsin. We took a random sample of 5000 cases which were either on AFDC or AFDC-U, a program open to income eligible two parent families with an unemployed spouse. Of these, 848 women were in the state of work and welfare during at least one month--they had earnings while receiving AFDC during at least one month. We chose two similar-sized subsamples of women who were in the other two states, welfare/no work and off welfare.

There are two major advantages to these data. First, using administrative records greatly reduces the problem of measurement error inherent in surveys which gather information from recipients. Second, the CRN keeps information on persons who leave AFDC, allowing us to look at recidivism.

There are, however, several disadvantages to this data set. First, it contains limited information on demographic characteristics. Second, only the Wisconsin caseload is covered. Inasmuch as Wisconsin is a high-benefit state with liberal rules, one must be careful in extrapolating results from this study to the nation as a whole.[3] The use of state data has the further disadvantage that some persons who exit AFDC in Wisconsin were going to programs in other states. However, the results of previous studies using this data base have been remarkably similar to results based on data for other states (see Hutchens, 1986, for a comparison of studies). The fourth limitation of the data set is that it, like other case record data, cannot be used to estimate the probability of entering an initial welfare spell, since we do not have information on potential recipients who were never on the rolls.

Our data covers the sixty months between January 1981 and December 1985. Our sampling frame was the universe of all cases which experienced at least one transition among the three states shown in Figure 1 during this period.[4] This scheme, which samples cases regardless of their current status (instead of sampling spells in progress), avoids length biased sampling (See Ondrich, 1985).

Tables 1 and 2 provide summary statistics which begin to describe the changes occurring over the OBRA period. While these tables are suggestive, they neither control for changes in observed nor unobserved characteristics. For this, a statistical model is estimated in the next section.

The first four columns of Table 1 show the empirical hazard rates (conditional probabilities) for exiting work and welfare in the pre- and post-OBRA periods.[5] In the pre-OBRA period the hazard of stopping work

while remaining on welfare (column 1) and the hazard of leaving welfare (column 3) each fluctuated in the neighborhood of 10 percent during the first five months.[6] This resulted in only 34.9 percent of persons working while receiving welfare for five or more months (column 5). In the post-OBRA period the hazard of a welfare/work spell stopping through an exit from AFDC (column 4) increased by more than 30 percent over the pre-OBRA level in each of the first five months. Exits to welfare/no work also increased sharply during the first two months. The result was that the length of welfare/work spells was clearly shorter after the implementation of the OBRA changes--by the end of the fourth month only 27.0 percent of the spells were still in progress, compared to 42.3 percent in the pre-OBRA period.

There is, however, no evidence of an unusually sharp increase in the proportion stopping work in the fifth month, when the $30 and a third disregard is eliminated in the post-OBRA period. The hazard, in fact, declines from 10.3 percent to 8.0 between the fourth and fifth months. Thus, while persons may have ended welfare/work spells in anticipation of the termination of $30 and a third, they were not maximizing current incomes by truncating spells precisely at the end of the fourth month.

Even more surprising is the finding that a substantial proportion of recipients continue to work while facing a 100 percent statutory benefit reduction rate. In fact 12.4 percent of welfare/work spells are still in progress after eight months.

Table 2 presents the pre- and post-OBRA empirical hazard rates for the transitions into the welfare/work state. The small number of pre-OBRA observations obscure any patterns that might be present,

although, it is notable that the pre-OBRA hazard rate for the transition from no welfare to welfare/work is less than or equal to the post-OBRA hazard in each of the first eight months.

While suggestive, the data in Tables 1 and 2 does not control for observed or unobserved characteristics of the caseload, which changes with duration and calendar timing or for changes in the environment. To test for significant changes in the hazard, conditional on observables and unobservables, it is necessary to estimate the statistical model developed earlier. This is done in the following section.

ESTIMATION

The model we estimate includes race, education and age at the first transition.[7] Control variables which vary with time include the presence of a spouse, the requirement instituted in December, 1982 that recipients report their income monthly, the quarterly state unemployment rate, and three dummy variables indicating the age of the youngest child. We also include three dummy variables which identify the loss of eligibility for the $30 and a third provision; we include one dummy for the last month of eligibility, one for the first month of ineligibility, and one for all subsequent months of ineligibility. A dummy variable for the months during the implementation of OBRA and a dummy for the months after the OBRA changes capture other effects of the OBRA rule changes.

The parameters of the unobserved heterogeneity are estimated jointly with the multinomial logit coefficients. We assume that the unobservables for transitions from one state are independent of the

unobservables for transitions from the other states. This implies that the likelihood function for a sequence of transitions factors into three likelihood functions--one for transitions out of each of the three states. These three likelihoods can be estimated separately.

The two values of Z_{j1} which define the points of support for the unobserved heterogeneity distribution cannot be identified separately from the intercept of the multinomial logit equation for the corresponding transition. Therefore, we constrain the intercepts to be zero. The appendix describes the estimation procedure and presents the estimates for the unobserved heterogeneity distribution.

Table 3 presents the results of the estimates for the transitions out of the welfare/work state. The coefficients on the control variables show that older, more educated recipients are less likely to stop working and more likely to leave welfare. Higher unemployment rates are associated with exits, both by jobs coming to an end and by direct exits from welfare. Little duration dependence is apparent in the data after controlling for unobserved heterogeneity.

The effect of the OBRA rule changes is represented by the coefficients on the OBRA variables and on the $30 and a third variables. Consistent with several other studies we find that the probability of a working recipient leaving welfare was substantially increased during the period when the OBRA changes were implemented. The coefficient of .574 on post-OBRA for the transition from welfare/work to no welfare indicates that, after OBRA, working recipients were also more likely to leave AFDC during the first three months of a job. However, as the exhaustion of the $30 and a third provision approaches in the fourth month, the

coefficient of -.344 indicates a lower tendency to leave AFDC. After the fifth month, the hazard rate falls below the pre-OBRA rate.

The analogous coefficients in the first column of Table 3 indicate that OBRA did not have a significant effect on the probability that a woman would quit her job and stay on welfare. The negative effect of the exhaustion of the $30 and a third provision undoubtedly reflects compositional effects--after cases were eliminated from the roles for excessive earnings and recipients quit their jobs because of lack of financial incentives, those who remained had lower probabilities of going back to welfare/no work.

Table 4 presents the estimates for the transitions into the welfare/work state. The most striking result is the coefficient on post-OBRA in the entrance from welfare/work. The large magnitude of the coefficient shows that the odds of starting a job while remaining on AFDC were much lower in the post-OBRA period. The analogous coefficient in the second column shows no significant change on the probability that a former recipient would start a welfare/work spell.

Tables 5 and 6 provide the predicted hazard rates for transitions in and out of the welfare/work state using the coefficients from all six transitions in the pre- and post-OBRA periods. These tables are similar to Tables 1 and 2 but they hold constant for observable characteristics. The predicted hazards are for women with particular observed characteristics but the patterns are typical. The predicted hazard rates reflect the estimated distribution of unobserved heterogeneity; the composition of the population at risk changes as individuals with zero probability of exiting by one or the other transitions make up an increasing proportion of the risk set.

Table 5 illustrates the combined effect of the post-OBRA dummy and the $30 and a third variables. It shows that while the survivor rate of work/welfare is initially lower in the post-OBRA period, for persons with spells over six months the survivor rates are comparable pre- and post-OBRA. This is the result of more negative duration dependence indicated by the negative $30 and a third coefficients.

The first two columns of Table 6 shows the dramatic effect of the post-OBRA coefficient on the probability that a recipient will start working. While the probability that a recipient would start working while continuing to receive aid was 4-5 percent in the pre-OBRA period, this dropped to 1 percent in the post-OBRA period. Similarly, columns 3 and 4 show that the probability that a former recipient would start a new welfare spell while working was cut by over a third. The fact that this effect was not apparent in Table 2 shows the importance of controlling for observable characteristics.

CONCLUDING REMARKS

Classifying AFDC recipients according to their work status and following recipients through multiple spells aids our examination of changes in AFDC participation associated with the OBRA rule changes. While we cannot completely separate the administrative effects from the behavioral responses, we find that OBRA affected both the probability of leaving welfare while working and the probability of starting work while on welfare.

The limitation of the $30 and a third provision to the first four months of any episode of welfare/work is asociated with an increase in

transitions from welfare/work to no welfare during the first five months of the spell. Individuals continuing to work and receive AFDC after the fifth month have a lower exit rate than individuals in similar pre-OBRA months. While we find the increase in early exits to be consistent with the search for higher paying jobs in anticipation of the tax increase, we are at a loss for a behavioral explanation of the lower hazard rates in subsequent months. If this is due to sorting on unobserved characteristics in the post-OBRA period, it suggests the need for a richer specification of unobserved heterogeneity.

We also find a dramatic decrease in the probability that a woman will start a job while remaining on AFDC. This is consistent with the view that women are less likely to start jobs when there is a 100% tax rate after four months.

An interesting extension of the framework would allow for correlation of unobservable characteristics across states of origin. For example, it would be possible to examine whether individuals with a greater tendency to move from welfare/work to no welfare also had a greater tendency to make the transition back to welfare/work. Allowing for such correlation and allowing the distribution of unobserved heterogeneity for new entrants to change over the timeframe would control for the sorting we suspect is occurring in the post-OBRA period.

Appendix

For each state j = 1,2,3, the probabilities of each woman having any pair (Z^h_{jk}, Z^g_{jl}) are constrained to lie inside the unit interval:

$$\pi^{hg}_j = \frac{\exp(\alpha^{hg}_j)}{\sum_{h=1}^{2}\sum_{g=1}^{2}\exp(\alpha^{hg}_j)} \; ; \quad \alpha^{22}_j \equiv 0 \text{ without loss of generality.}$$

Initital attempts to maximize the likelihood for all of the states led to increasingly negative values for Z^2_{jk} for all j and k. In order to obtain convergence, Z^2_{jk} was constrained to equal negative infinity. This implies that some women will never make the transition from state j to state k.

There is also the possibility that one of the cells of the Z_{jk} is empty. For example, if there are no women with the pair (Z^2_{kl}, Z^2_{jl}) then π^{22}_j will equal zero, and α^{22}_j will equal negative infinity. This was the case for women in the welfare/no work state; to obtain convergence, π^{22}_1 was constrained to equal zero.

The parameter estimates and the implied probabilities are listed below. A parameter which was constrained during estimation is indicated by a standard error of zero.

State of Origin 1: Welfare, No Work

Probability Parameter	Estimate	Standard Error
α_1^{11}	1.30	(.294)
α_1^{12}	-1.10	(.483)
α_1^{21}	0	(0)
α_1^{22}	$-\infty$	(0)

Estimated points of support and implied probabilities:

		z_{13}	
		-3.60 (.410)	$-\infty$ (0)
z_{12}	-4.93 (.613)	$\pi_1^{11} = .734$	$\pi_1^{12} = .067$
	$-\infty$ (0)	$\pi_1^{21} = .200$	$\pi_1^{22} = 0$

State of Origin 2: Welfare, Work

Probability Parameter	Estimate	Standard Error
α_2^{11}	5.66	(2.37)
α_2^{12}	3.41	(2.53)
α_2^{21}	4.14	(2.39)
α_2^{22}	0	(0)

Estimated points of support and implied probabilities:

		z_{23}	
		-4.23 (.433)	$-\infty$ (0)
z_{21}	-1.21 (.487)	$\pi_2^{11} = .754$	$\pi_2^{12} = .079$
	$-\infty$ (0)	$\pi_2^{21} = .165$	$\pi_2^{22} = .003$

State of Origin 3: No Welfare

Probability Parameter	Estimate	Standard Error
α_3^{11}	1.67	(.777)
α_3^{12}	1.98	(.806)
α_3^{21}	1.10	(.897)
α_3^{22}	0	(0)

Estimated points of support and implied probabilities:

		z_{32}	
		-3.67 (.932)	$-\infty$ (0)
z_{31}	-.931 (.478)	$\pi_3^{11} = .321$	$\pi_3^{12} = .437$
	$-\infty$ (0)	$\pi_3^{21} = .181$	$\pi_3^{22} = .060$

Notes

[1]It should be noted that if the unobserved heterogeneity distribution is not sufficiently flexible, the duration dependence coefficients will be biased downward. In fact, improper specification of either unobserved heterogeneity or duration dependence will lead to inconsistent estimates of all parameters. However, we find that our estimates of the slope parameters are insensitive to changes in the specification of unobserved heterogeneity.

[2]Note that B_{jj}, $j = 1, 2, 3$, can be set equal to zero, without loss of generality.

[3]For example, recidivism in Wisconsin may be higher than for the nation as a whole, since Food Stamp recipients who had previously received AFDC are automatically reenrolled in AFDC when their earnings drop to a point making them eligible.

[4]Excluding cases which were on AFDC and in the same work status all sixty months potentially causes selection problems. This problem is mitigated if, as we assume, unobserved heterogeneity is independent across states of origin.

[5]We divide the timeframe into the following periods:

pre-OBRA	January 1981 - December 1981
during OBRA	January 1982 - April 1982
post-OBRA	May 1982 - December 1985

[6]Fluctuations reflect sampling variability resulting from the small number (206) in the risk set in month 1.

[7]To improve computational accuracy, education was divided by ten and age was divided by one hundred prior to estimation.

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Table 1

Hazard and Survivor Rates for Exits
FROM Work and Welfare--Pre- and Post-OBRA

Duration	Hazard Rate of Exiting to: Welfare/No Work		Hazard Rate of Exiting to: No Welfare		Survivor Rate	
	Pre	Post	Pre	Post	Pre	Post
1	.121	.157	.078	.170	.801	.673
2	.068	.122	.101	.163	.666	.481
3	.130	.098	.096	.166	.515	.355
4	.115	.103	.064	.136	.423	.270
5	.070	.080	.105	.131	.349	.213
6	.089	.060	.156	.116	.263	.175
7	.000	.058	.100	.064	.237	.154
8	.115	.097	.154	.097	.173	.124
Number at risk in month 1:	206	1099	206	1099	206	1099

Table 3

Hazard Rate Estimates for
Exiting from Welfare/Work

Multinominal Logit Coefficients	Exit to: Welfare/No Work		No Welfare	
Age	-.313	(.818)	.0098	(.640)
White	-.0786	(.138)	.0172	(.0956)
Education	-.460	(.324)	1.31	(.300)
AFDC Subprogram	-.0989	(.222)	-.158	(.170)
Spouse present	-.165	(.199)	.492	(.159)
Monthly reporting	.390	(.210)	.167	(.177)
Unemployment rate	.0487	(.0268)	.0537	(.0217)
Youngest child age 5-6	-.0457	(.166)	-.0012	(.136)
Youngest child age 7-17	-.318	(.126)	.0860	(.102)
Youngest child age >17	-.217	(.735)	.414	(.747)
During OBRA	-.306	(.270)	.477	(.227)
Post-OBRA	.0114	(.257)	.574	(.227)
30&1/3 - final month	-.335	(.307)	-.344	(.272)
No 30&1/3 - first month	-.614	(.336)	-.391	(.288)
No 30&1/3 - remaining months	-.727	(.276)	-.677	(.243)
Duration: 2 months	-.243	(.128)	.0226	(.117)
3	-.401	(.155)	.0641	(.132)
4-6	-.138	(.251)	.123	(.230)
7-12	-.0859	(.282)	.0198	(.241)
13-24	-.0464	(.395)	-.215	(.313)
>24	-1.51	(1.48)	-.409	(.686)

Number of women in welfare/work sample = 848

(Standard errors in parentheses)

Note: Unobserved heterogeneity distribution parameter estimates are in the Appendix.

Table 4

Hazard Rate Estimates for
Exiting to Welfare/Work

Multinominal Logit Coefficients	Exit from: Welfare/No Work		No Welfare	
Age	-.370	(1.02)	.319	(1.47)
White	.720	(.168)	.131	(.205)
Education	.978	(.358)	.195	(.487)
AFDC Subprogram	.503	(.271)	.261	(.290)
Spouse present	-1.22	(.238)	-.672	(.235)
Monthly reporting	2.68	(.139)	1.69	(.237)
Unemployment rate	.0562	(.0302)	.0788	(.0428)
Youngest child age 5-6	-.237	(.214)	-.197	(.252)
Youngest child age 7-17	.450	(1.59)	-.0416	(.210)
Youngest child age >17	.545	(.765)	-1.27	(.947)
During OBRA	-.640	(.412)	.0545	(.443)
Post-OBRA	-1.58	(.247)	-.512	(.413)
Duration: 2 months	.027	(.222)	.117	(.177)
3	.373	(.212	-.946	(.288)
4-6	.0568	(.191)	-1.27	(.219)
7-12	-.263	(.200)	-2.24	(.271)
13-24	.0354	(.201)	-2.50	(.261)
>24	.0427	(.310)	-3.64	(.565)

Number of women in:
welfare/no work sample = 797
no welfare sample = 900

(Standard errors in parentheses)

Note: Unobserved heterogeneity distribution parameter estimates are in the Appendix.

Table 6

Predicted Hazard Rates for Exits
TO Work and Welfare Pre- and Post-OBRA

Duration	Hazard Rate of Exiting from: Welfare/No Work		No Welfare	
	Pre	Post	Pre	Post
1	.043	.009	.027	.017
2	.043	.010	.033	.020
3	.059	.013	.012	.007
4	.044	.010	.009	.006
5	.043	.010	.009	.006
6	.043	.010	.009	.006
7	.032	.007	.003	.002
8	.031	.007	.004	.002

Note: Hazard rates are evaluated for women with the following characteristics: 25 years old, white, 11 years education, 7.0% unemployment. The post-OBRA columns reflect a value of one for the post-OBRA dummy. All other dummies are set to zero.

Figure 1

