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WELFARE FLOWS AND CASELOAD DYNAMICS

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Data from the 1990, 1991, 1992, 1993 and 1996 panels of the Survey of Income and Program Participations are used to estimate AFDC/TANF entry and exit rates. These estimates of AFDC/TANF entry and exit rates are used to conduct simulations aimed at determining the roles of economic conditions and welfare reform in explaining AFDC caseloads changes during the 1990s. The results of these simulations indicate that economic conditions were the engine driving the run-up in caseloads during the early 1990s and the decrease in caseloads following 1993.

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I. Introduction

The mid-1990s was a time of unprecedented efforts at welfare reform. In the early-1990s the Department of Health and Human Services began granting states waivers to run experimental welfare reform programs. Some of the reforms implemented under these waivers time limited benefit receipt, required recipients to work, reduced the effective tax rate on earnings, and reduced the benefit increase available to recipients upon the birth of a child. Others reduced the number of recipients who were exempt from training associated with Job Opportunity and Basic Skills Program (JOBS), and made it easier for states to sanction recipients for failing to comply with JOBS requirements.

Efforts at welfare reform culminated in August of 1996 when President Clinton signed the Personal Responsibility and Work Opportunity Reconciliation Act

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(PRWORA). This law replaced the Aid to Families with Dependent Children program (AFDC) with Temporary Assistance to Needy Families (TANF) block grants. In order to receive TANF block grants states must limit the lifetime receipt of TANF benefits to 60 months and require TANF recipients to work within 2-years of receipt. Although the PRWORA legislation mandated minimum time limits and minimum work requirements, it gave states much greater latitude in designing their public assistance programs.

Between 1992, when states began implementing statewide pre-PRWORA waivers, and mid-1997, when TANF was implemented nationwide, AFDC-Total caseloads fell by nearly 40 percent. Particularly important in light of these dramatic caseload declines is the extent to which welfare reform and economic growth hastens exits from, and slows entry to, the AFDC or TANF (AFDC/TANF) program(s). This paper uses data from multiple panels of the Survey of Income and Program Participation (SIPP) merged with state-level program and unemployment data to estimate the rates of exit from and entry to the AFDC/TANF program(s). The sample period covered by the data used in this analysis is November 1989 to February 2000. The objectives of this exercise are to assess the direct effect of the pre-PRWORA waivers, early TANF implementation, and economic conditions on the rate of entry to and exit from the AFDC/TANF program(s) and to document how these entry and exit rate effects translate to the AFDC/TANF caseload.

A number of studies have examined the relative contributions of economic growth versus welfare reform in explaining the decline in AFDC/TANF caseloads using state-level caseload data. Among these studies there is wide disagreement as to the effect of pre-PRWORA waivers on caseloads. Previous studies of the effects of welfare reform rely on some form of time series data and, therefore, run the risk of measuring spurious correlation. This paper contributes a new approach. Although duration models abstract from the variable of interest (the AFDC/TANF caseload) they have the advantage of not being as subject to the problem of spurious correlation. Furthermore, estimates of the rate of entry to and exit from AFDC/TANF can be used to simulate the time path of aggregate caseloads under various counterfactual histories of economic conditions and welfare reform implementation.

The organization of the remainder of this paper is as follows. In Section II the previous literature on the dynamics of AFDC/TANF participation and caseload change is reviewed. In Section III the SIPP data is discussed. In Section IV the proportional hazard model is presented and specifications for the baseline hazard model are discussed. The results are presented in Section V. Section VI concludes this paper.

II. Literature review

Corresponding to the time that many states began to implement their pre-PRWORA waivers and continuing through implementation of TANF, aggregate welfare caseloads decreased dramatically. Between 1994 and mid-1997, welfare (AFDC-Total) caseloads fell by more than 40 percent. Numerous studies have examined the impact of pre-PRWORA waivers on AFDC participation during the 1990s using aggregate state level data. Among these studies a consensus on the magnitude of the effects of the pre-PRWORA waivers on AFDC/TANF participation has not emerged (Levine and Whitmore 1998; Wallace and Blank 1999; Figlio and Ziliak 1999; Ziliak et al. 2000; Blank 2001; Klerman and Danielson 2004). In particular some studies find negligible effects of pre-PRWORA waivers on AFDC/TANF participation while other studies find non-negligible and statistically significant negative effects.

Underlying the lack of consensus on the effects of welfare reform on AFDC participation is the coincidence of welfare reforms and a period of unprecedented economic growth.[1] To complicate matters further, there were other significant policy changes during this period that are thought to have substantial implications for the welfare population.[2] Separating the effects of welfare reform from economic growth and other policy changes with state-level panel data has proved difficult.

Concerns about the accuracy of estimates of the impact of welfare reform from state-level panel data suggest that it would be useful to assess whether welfare reform had a measurable impact on entry to and exit from the AFDC/TANF program(s). Although the mechanism by which pre-PRWORA waivers and TANF

[1] Between 1992 and early 2000 the United States underwent the longest sustained period of economic growth in recorded history. During this period of time over 20 million new jobs were created, the national unemployment rate fell from 7.5 percent to 4.0 percent, and real wages increased nearly 7 percent.

[2] This included the expansion of the Earned Income Tax Credit as well as increases in the minimum wage.

implementation affect caseloads has not been studied closely, it seems likely that these reform efforts would affect initial AFDC/TANF take-up behavior and the decision to continue receiving AFDC benefits.

The only evidence on the relative contributions of entry and exit to AFDC caseload changes comes from Grogger et al. (2003). They attribute 40 percent of the decline in SIPP AFDC participation rates to declining entry rates and 60 percent of the decline to increasing AFDC exit rates. While they do not explicitly address the causes of the decline in entry and increase in exit, it seems likely that welfare reform and/or the macroeconomy played strong roles in driving these trends. Thus, the impact of pre-PRWORA waivers, TANF implementation, and changes in economic conditions should be observable in data describing spells of AFDC/TANF receipt and eligibility without receipt. Because the AFDC/TANF caseload is a function of both take-up and exit behavior, estimates of the effects of welfare reform and unemployment rates on entry and exit from AFDC/TANF can be used to simulate caseloads under alternative policy and economic environments.

In a recent study Klerman and Haider (2004) pursued this approach. Using a very large sample of administrative data from California, they estimate AFDC entry and AFDC continuation rates using specifications that control for monthly county unemployment rates (and lags) and monthly period effects. These estimates of AFDC entry and continuation rates are then used to simulate the response of AFDC caseloads to alternative histories of the county unemployment rates. In their preferred simulation, Klerman and Haider determine that strong labor market performance, as manifested by declining county unemployment rates, accounts for roughly 45 percent of the rather substantial decline in California's per-capita AFDC caseload between March of 1995 and December of 1998.

Because Klerman and Haider use administrative data from California, their paper does not provide much information on the relative contribution of changes in local labor market conditions to the reduction in AFDC caseloads nationally; nor does it provide any direct information on the relative contribution of welfare reform to caseload declines in California or elsewhere. While their work does not warrant a firm conclusion regarding the effect of improving economic conditions on AFDC caseloads nationally, it does suggest that the economy played a substantial role.

Recent work by Ribar (2005) addresses some of the shortcomings of Klerman/Haider. Using data from the 1992 and 1993 panels of the geo-coded SIPP, Ribar estimates AFDC entry and continuation rates on a sample of unmarried women with children. AFDC entry and continuation rates are modeled as functions of individual demographic characteristics, a comprehensive set of local labor market characteristics, year effects, the maximum AFDC benefit level, and a dummy variable for whether a statewide welfare waiver was in effect.

Ribar's estimates indicate that local labor market factors and AFDC benefit levels are strong determinants of AFDC entry and exit, but pre-PRWORA waiver implementation was not. There are several possible explanations for the lack of statistically significant welfare reform effects in Ribar's analysis. First, most of the states that implemented waivers did so in 1994, 1995, and 1996. For the period of time covered by Ribar's sample (October of 1991 through December of 1995), many of the state welfare reform programs were in their infancy. It may take a few months or even a few years for states to fully implement their programs or for recipients to adjust their behavior in the face of a new policy environment.

Secondly, some of the details surrounding the construction of the sample may factor in the lack of discernable welfare reform effects in Ribar's study. As noted above, Ribar's estimates are obtained using a sample of all unmarried women. Given this case selection criteria, it is not surprising that welfare reform does not have a statistically significant affect on AFDC entry and exit. The stock sampling scheme will lead to an underestimate of the rate of transition from a particular state. This downward bias in the rate of transition estimates occurs because long spells are over sampled in stock samples. A sample consisting of unmarried women who are not receiving welfare at a particular time will be composed of a large number of women who will never be on welfare, and are unlikely to be affected by welfare reform. Conversely, a sample consisting of unmarried women who are on welfare at a particular time will be composed of a large number of long-term welfare cases. These long-term welfare cases are the cases least likely to be affected by welfare reform.

Both the work by Ribar and Haider and Klerman are extensions of an earlier literature on the dynamics of AFDC participation. Papers by Blank (1989), Fitzgerald (1991 and 1995), Blank and Ruggles (1996), and Hoynes (2000), as well as earlier papers by Hutchens et al.(1981), Plotnick (1983), and O'Neill et al. (1987) were concerned with estimating the determinants of AFDC exit and/or entry. The early studies in this literature primarily highlight the determinants of AFDC entry and exit without addressing focused research questions (Hutchens 1981; Plotnick 1983;

O'Neill et al. 1987). Later studies deal with a varied set of questions including the degree of duration dependence in AFDC spells (Blank 1989), the effect of local labor market indicators on AFDC exit (Fitzgerald 1995, Hoynes 2000), the dynamics of eligibility versus participation (Blank and Ruggles 1996), and the importance of marriage market conditions in determining welfare exits (Fitzgerald 1991).

This study is designed as both a substantial extension of the dynamic AFDC/ TANF participation literature, as well as a synthesis of this literature with the caseload literature described earlier. Relative to earlier papers in the dynamics AFDC/TANF participation literature, I am using more recent data that fully spans the period of welfare reform in the US. Additionally, because I pool 5 SIPP panels there is enough data to estimate models with the full compliment of controls.[3] Due to the coverage and size of the sample, I am able to estimate the effects of pre-PRWORA waivers, TANF implementation, and state unemployment rates on AFDC/ TANF entry and exit using interstate variation in the timing of implementation. The estimates obtained using this approach are used to simulate the time path of aggregate caseloads under different policy and economic environments in a way that makes the results presented here easily comparable with those presented in the AFDC/TANF caseload literature.

III. Data

The data used in this analysis were constructed from the 1990, 1991, 1992, 1993, and 1996 panels of the SIPP, along with state-level monthly unemployment rates and welfare reform implementation and dates.[4]

The SIPP is a series of overlapping longitudinal surveys administered and published by the United States Census Bureau. Once every wave (4 months) SIPP participants are asked about their income, earnings, and program participation over the prior 4 months. The 1990 and 1991 SIPP panels ran for 32 months and the

[3] Virtually all of the studies of AFDC/TANF entry and exit conducted using data with national coverage are estimated without state effects. Given the demonstrated importance in providing controls for state effects in other welfare related research and the small sample sizes cited in many of the studies, I am left to conclude that state effects were omitted because of insufficient data in most cases.

[4] I obtained the state unemployment rate data from the Bureau of Labor Statistics. The welfare reform implementation dates used in this analysis came from the CEA (1999).

1992, 1993, and 1996 SIPP panels ran for 40, 36, and 48 months respectively. Together all of the SIPP panels provided monthly information on approximately 330,000 individuals.

The first step in constructing the data was identifying a sub-sample of women who were categorically eligible to receive AFDC/TANF during some part of the time they appeared in a SIPP panel. The determination of categorical eligibility for AFDC/TANF is complex as the programs' categorical eligibility standards were changing over the period of study.

Historically, benefits associated with the AFDC program were available only to single parents with dependent children, with the vast majority of the AFDC caseload composed of female-headed families with children. Beginning in the 1970s some states began running AFDC Unemployed Parent (AFDC-UP) programs. These programs provided AFDC benefits to two parent families if the principle wage earner was unemployed and certain other conditions were met.[5] Following the adoption of the Family Support Act of 1988 in October 1990 all states were required to run AFDC-UP programs. Because eligibility rules were different for AFDC-UP, cases associated with this program were separately identified from the more traditional single parent (or child-only) AFDC-Basic cases in state-level caseload data.

In setting up their TANF programs most states did away with any distinction between single parent and other types of families in determining eligibility for aid.[6] These changes in eligibility determination ended the practical distinction between the Basic and UP portions of the caseload. Thus, in the period covered by the data categorical eligibility for the AFDC/TANF changed dramatically. In the early part of the decade eligibility in most states was restricted to single parents with dependent children. By the end of the decade eligibility in most states was based solely on a family's financial status.

I abstract from these changes in eligibility by focusing on modeling flows into and out of receipt and eligibility of the Basic component of AFDC/TANF (AFDC/

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[5] The principle wage earner was required to have a recent history of employment.

[6] As of 1999 33 states used the same eligibility guidelines for 2 parent families and single parent families. The remaining states allow 2 parent families to receive benefits only if the principle wage earner is incapacitated or unemployed.

TANF-Basic).[7] The definition of categorical eligibility that I adopt in this analysis corresponds loosely with categorical requirements for receipt of AFDC-Basic. Any female sample respondent who is above the age of 17, single (not married with a present spouse), and living with at least one of her children (or 7+ months pregnant) is coded as categorically eligible for receipt of AFDC/TANF-Basic.[8]

To record first observed spells of AFDC/TANF receipt, SIPP respondents were tracked and their transitions onto and off of AFDC/TANF were noted. An AFDC/TANF spell begins when a sample member transitions to AFDC/TANF receipt after a period of known non-receipt, and ends when they stop receiving AFDC/TANF for at least one month, they leave the sample, or their SIPP panel completes. Spells ending with a transition to non-receipt are coded as complete. Spells that are not completed by the time information on the individual is exhausted are coded as right censored. To be sure that I am coding spells of receipt that will allow me to simulate changes in AFDC-Basic caseloads, I excluded cases in which the

[7] Following the widespread implementation of state TANF programs in 1997, the Basic component of the AFDC/TANF caseload was not separately identifiable in administrative caseload data. For the months covered by the SIPP data used in this study, and for which the Basic component of the AFDC caseload can be separately identified (1990 through June 1997), the ratio of the national AFDC-Basic caseload (receipt units) to the AFDC-Total caseload (receipt units) ranged between 0.92 and 0.95. Although most of AFDC-Basic caseload is composed of female headed families with children, child-only cases (cases in which the adult does not appear on the grant) are also included in the count. In 1990 these child-only cases account for approximately 10 percent of the AFDC-Basic caseload, but by 1996 they accounted for over 20 percent of the AFDC-Basic caseload.

[8] The only way to determine the number and ages of children born to, and living with, women in the sample is to match the children's birth dates to their mother's records on the basis of a parent number variable attached to each child's record. Unfortunately, the records for many children in the 1990, 1991, 1992 and 1993 SIPP panels appear to be erroneous. The most common problem is children dropping out of the SIPP and subsequently reappearing in the same sample unit with a different person number. A literal interpretation of this pattern would be that one child left the sample unit after which time a second child with the same birth date moved in. Because of these errors it is impossible to construct accurate information on the number and ages of the minor biological children that are living with each woman during each month she appears in the SIPP. The best that can be done is to determine the birth dates of children that can be linked with each woman over the *entire* time she appears in the SIPP. I then assume that all minor children are living with their mother during each month, making adjustments for the type of recording errors described in the previous paragraph by separately identifying twins from probable recording errors. Twins are identified as a pair of children with the same birth date that overlap for at least one wave of the SIPP. Recording errors are identified as two or more children with the same birth dates that do not overlap for at least one wave of the SIPP.

respondent does not meet the requirements for categorical eligibility within the first three months of a spell beginning.

First observed spells of eligibility without AFDC/TANF receipt were coded in a similar way. A spell of eligibility without receipt begins when an AFDC/TANF recipient leaves the program for one or more months (while remaining categorically eligible), or when a previously categorically ineligible female becomes categorically eligible because of near a full term pregnancy (7+months) for a resultant birth, an 18th birthday, or a marital disruption of 2 or more months.[9] A spell of AFDC/TANF eligibility without receipt ends when a categorically eligible woman begins receiving welfare, leaves the sample, her SIPP panel completes, or she no longer meets requirements for categorical eligibility. Spells of eligibility without receipt that end with a transition to AFDC/TANF are coded as complete. Spells ending for other reasons are coded as right censored.[10]

Because a spell of eligibility without receipt can be initiated by spell of receipt ending, many respondents with spells of receipt also have spells of eligibility without receipt. To be more precise, there are 2,483 women who have spells of AFDC/TANF receipt, 4,863 women who have spells of categorical eligibility without receipt, and 1,755 women who are observed as having had a spell of receipt and a spell of eligibility without receipt during their time in the SIPP.

Means and standard deviations for the variables used in this analysis of AFDC/ TANF entry and exit are shown in Table 1. The figures shown in Table 1 and the

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[9] As might be expected there are a large number of cases where categorical eligibility status changes for only a short period of time. For example, it is quite common for women that are married for much of a SIPP panel to have one month where they report not living with their spouse. It is doubtful that these women are at significant risk for receipt of the Basic component of AFDC during these one month spells of categorical eligibility. As such there is a good reason these spells should be excluded from the sample. The requirement that a marital disruption last 2 or more months to begin a spell of eligibility without receipt is designed to exclude spells caused by one month marital disruptions. Occasionally, a one month marital disruption is followed by a more lengthy marital disruption. In these cases the first, one month marital disruption is ignored, but the second longer disruption is treated as a spell of eligibility without receipt, assuming all of the other requirements for categorical eligibility are met.

[10] For the type of empirical model that I am using the coefficients associated with each risk are additively separable in the log likelihood function. One implication of this is that estimates of AFDC/TANF entry rate are not affected by whether spells of eligibility without receipt ending with a loss in eligibility are coded as censored or treated as a separate risk.

Table 1. Sample means and standard deviations

Variable	AFDC/TANF recipients (N=2,843)	Eligible non-recipients (N=4,863)
Age_t	29.4073	30.6891
	(8.7172)	(8.7694)
Less than high $school_t$ (=1)	0.3580	0.2445
	(0.4795)	(0.4298)
High school $graduate_t$ (=1)	0.3899	0.3924
	(0.4878)	(0.4883)
Some $college_t$ (=1)	0.2521	0.3632
	(0.4343)	(0.4810)
White (=1)	0.4575	0.5828
	(0.4983)	(0.4932)
Black (=1)	0.3516	0.2655
	(0.4776)	(0.4416)
Hispanic (=1)	0.1909	0.1518
	(0.3931)	(0.3588)
Number of children under 18_t	1.8131	1.5696
	(1.1407)	(1.2073)
$Preschool_t$ (=1 if respondent has a child less than 6-years old)	0.6283	0.4300
	(0.4834)	(0.4951)
$Baby_t$ (=1 of respondent has a child less than 2-years old)	0.3355	0.1524
	(0.4723)	(0.3594)
Older $children_t$ (=1 if respondent's youngest child is 16 or 17)	0.0314	0.0448
	(0.1745)	(0.2069)
AFDC/TANF prior to baseline (=1)		0.4491
		(0.4975)
Real maximum $benefit_t$ (the maximum monthly AFDC/TANF benefit for a 3-person family in hundreds of US $\$_{2000}$)	4.5361 (1.8413)	4.3737 (1.7902)
$Waiver_t$ (=1)	0.1857	0.1729
	(0.3889)	(0.3782)
$TANF_t$ (=1)	0.1740	0.2572
	(0.3792)	(0.4372)
Unemployment $rate_t$ (state monthly unemployment rate, seasonally adjusted)	6.1969 (1.5897)	5.9406 (1.5849)

Note: Sources are 1990, 1991, 1992, 1993, and 1996 panels of the SIPP (all individual level variables), CEA (1999) (welfare reform variables), and the Bureau of Labor Statistics (unemployment rate). Means and standard deviations are unweighted.

estimates that follow were computed without the use of SIPP weights.[11] The first column shows the means for the sample of women who experienced spells of AFDC/TANF receipt and the second column shows the means for the sample of women who experienced spells of categorical eligibility with non-receipt. Time-varying covariates are denoted by a subscript *t*. Time invariant variables were recorded during the first month of a spell. Most of the variables in Table 1 require little explanation. There are, however, some variables that will require additional comment.

Baby indicates whether an AFDC/TANF recipient is at least 3 months pregnant or has a child less than 2 years of age. All things being equal, one would expect that women with very young children are less likely to leave welfare. *Unemployment Rate* is the monthly unemployment rate (seasonally adjusted) from the state in which the respondent resided. *Real maximum benefit* is the maximum monthly AFDC/TANF benefit for a three-person family in the state in which the respondent resided in 2000 dollars.

[11] The SIPP is a stratified probability sample of the US population. Given the nature of the SIPP sampling criteria, the use of sampling weights is warranted. The SIPP files contain longitudinal weights and calendar year weights. Calendar year weights are computed for all sample members who were interviewed (self or proxy) during January of the year corresponding to the weight. Each SIPP panel contains a calendar year weight variable for each year it completely covers. For example, the 1993 SIPP panel (First Rotation Group: October 1992-September 1995, Fourth Rotation Group: January 1993-December 1995) contains calendar year weights for 1993 and 1994. Calendar weights are not computed for 1992 and 1995 because not all rotation groups were interviewed during each wave covering these years. The SIPP calendar weights are designed to allow SIPP users to produce estimates of population parameters that are unbiased for the US population in each full year covered by a panel. Calendar weights may be used as longitudinal weights for samples that cover a period no longer than one calendar year, but are not appropriate for other types of longitudinal samples. The longitudinal weights (or modified version of them) would be appropriate for this analysis but for the fact that they are not computed for a substantial fraction of the sample. SIPP longitudinal weights are only computed for original sample members who were interviewed (self or proxy) during each wave of a panel. This means that longitudinal weights are missing for more than 25 percent of the women in my sample. In any type of weighted analysis these women would be dropped, reducing the effective sample size substantially. Because of these concerns about sample size, I decided not to show the results based on the weighed numbers. In practice the weighted estimates were similar to the un-weighted estimates. None of the conclusions reached in this study are affected by the decision not to use the longitudinal weights.

A. Measuring welfare reform

The welfare reform variables (*waiver* and *TANF*) were coded using implementation dates obtained from the Council of Economic Advisors (CEA) (1999). *Waiver* is a binary variable that reflects whether or not the state in which a respondent resided had a major statewide pre-PRWORA waiver in place. *TANF* is a binary variable that reflects whether the state in which a respondent resided had a TANF program in place. Because all of the states that implemented statewide waivers eventually implemented TANF, *Waiver* turns off (is equal to zero) after a state's TANF program has been implemented.

It is important to note that *waiver* and *TANF* do not reflect the degree of diversity in state welfare reform. The CEA (1999) identifies dates for the implementation of 6 major types of statewide waivers. Some of these waivers reduced the number of recipients who were exempt for activities related to the Job Opportunities and Basic Skills Program (JOBS exemption waivers) or strengthened the sanctions for not complying with JOBS program requirements (JOBS sanction waivers). Other waivers changed the formula used to compute benefits, either by altering the relationship between the benefit and earned income (earnings disregard waivers) or by reducing incremental increases in benefits for families that had additional children on welfare (family cap waivers). Some states applied for and were granted waivers that allowed them to set limits on the amount of time recipients could receive aid (time limit waivers). Yet another type of waiver allowed states to mandate work for recipients after a specified number of months of welfare receipt (work requirement waivers).

The net effect of pre-PRWORA waiver implementation is not to unambiguously decrease the caseload. Some waivers may actually lead to caseload increases in the short-run. For instance, increased earnings disregards, which make it possible for recipients to earn more income and retain eligibility for cash assistance, should lead to caseload increases in the short-run. The parameter estimates of the waiver effect in this study will reflect average effect of waiver implementation. The magnitude of this effect will depend on the mix of waivers in place over the period of time spanned by the sample.[12]

[12] Researchers have had limited success separating the effects of the various waiver provisions on AFDC/TANF caseload measures. Grogger (2003), Ziliak et al.(2000), and Klerman and Danielson (2004) have had the most success.

After PRWORA was passed in August of 1996, the welfare reform picture simplified slightly. PRWORA mandated that states set up cash assistance programs that complied with the federal guidelines. These guidelines stipulated that state programs had to include a lifetime limit on welfare receipt of no more than 60 months. States were also required to mandate work for recipients who were on welfare for more than 24-months continuously, but exercised a great deal of autonomy over the remaining components of their cash assistance programs.

States that expanded earnings disregards prior to TANF kept these programs in place and some states that did not have expanded earnings disregards prior to TANF expanded them as part of their TANF programs. Some states made family caps part of their TANF programs even though they were not part of their AFDC programs. While some states implemented time limits that were much shorter than those required by PRWORA, several states refused to impose time limits, opting instead to pay the cost of time limited recipients out of state funds.

For the vast majority of states, TANF implementation signifies that time limits are in place, that some work or work-related activities are mandated for recipients, that there are extensive sanctions in place for recipients who fail to comply with these requirements, and that earned income is taxed at or below AFDC benefit reduction rates. In some states TANF implementation may also signify that family caps are in place and that child-care is extensively subsidized.

IV. Estimation

The primary approach taken toward estimation in this analysis is attributable to Prentice and Gloeckler (1978). They provide an extension to the Cox proportional hazard model that allows for estimation of grouped or discrete duration data in a manner that does not require making functional form assumptions concerning the parametric form of the baseline hazard.

Consider the chance that an individual's spell of AFDC/TANF receipt (or non-receipt) ends in interval $(t, t+h]$ conditional on the individual having remained on AFDC/TANF (eligible without receipt) until t. The limit of this conditional probability as $h \to 0$ is known as the hazard rate. More formally,

$$\lim_{h \to 0} \frac{\mathrm{P}\left(t < \tau < t+h \mid \tau > t\right)}{h} = \lambda\left(t\right). \tag{1}$$

The proportional hazard assumption amounts to a functional form restriction on $\lambda(t)$. More specifically, the proportional hazards model assumes that

$$\lambda(t) = \lambda_0(t) \cdot \exp\left(x(t) \cdot \beta\right), \tag{2}$$

where the function $\lambda_0(t)$ is known as the baseline hazard, $x(t)$ is a vector of covariates that are allowed to depend on time, and β is a vector of parameters. Next, consider the probability that an individual has an AFDC/TANF spell (a spell of eligibility without receipt) of at least length t. This probability is known as the survivor function. In the proportional hazard model the survivor function takes the form

$$S(t) = \exp\left(-\int_0^t \lambda(u)\, du\right). \tag{3}$$

Assuming the vector of covariates is constant over the interval $[\, j-1, j)$, $S(t)$ can be rewritten as

$$S(t) = \exp\left(-\sum_{j=1}^{t} \int_{j-1}^{j} \lambda(u)\, du\right) = \exp\left(-\sum_{j=1}^{t} \exp\left[\gamma_j + x_j \cdot \beta\right]\right), \tag{4}$$

where $\gamma(j) = \ln\left[\int_{j-1}^{j} \lambda_0(u)\, du\right]$.

Once the survivor function is known, the likelihood function follows straightforwardly. Consider a sample consisting of n spells of AFDC/TANF receipt (or non-receipt). Let T_i denote the length of the *i'th* individual's AFDC/TANF spell (spell of non-receipt). These spells can either be uncensored or censored. For censored spells I adopt the convention that the person is known to have remained on (off) the program until the end of the T_i'th period. Letting $\delta_i = 0$ for all right censored observations and $\delta_i = 1$ for all observations with completed spells of AFDC/TANF receipt (or non-receipt) the likelihood function is

$$L(\gamma,\beta)=\prod_{i=1}^{n}\left\{\left(1-\delta_i\cdot\exp\left[-\exp\left\{\gamma_{T_i}+x_{i,T_i}\cdot\beta\right\}\right]\right)\times\right.$$
$$\left.\times\exp\left(-\sum_{j=1}^{T_i-\delta_i}\exp\left\{\gamma_j+x_{i,j}\cdot\beta\right\}\right)\right\}. \tag{5}$$

Note that the first part of equation (5) is the probability of leaving AFDC/TANF (or entering AFDC/TANF) in period T_i, assuming the spell is ongoing at the end of T_i - 1, while the second part of equation (5) is the probability of remaining (surviving) on AFDC/TANF (off AFDC/TANF) for T_i - δ_i months. Given (5) the log likelihood function is provided by

$$l(\gamma,\beta)=\sum_{i=1}^{N}\left\{\ln\left(1-\delta_i\cdot\exp\left[-\exp\left\{\gamma_{T_i}+x_{i,T_i}\cdot\beta\right\}\right]\right)-\sum_{j=1}^{T_i-\delta_i}\exp\left\{\gamma_j+x_{i,j}\cdot\beta\right\}\right\}. \tag{6}$$

Given sufficient data, it is straightforward to use (6) to estimate the sequence of duration effects $\left\{\gamma_t\right\}_{t=1}^{46}$ and the vector of coefficient parameters β. Estimation of this sequence requires the completion of multiple spells at each possible duration. Due to limited sample sizes, there are not enough completed spells to allow me to estimate the full sequence of γ_t terms. Specifically, there are only 24 *completed* spells of AFDC/TANF receipt, and only 21 *completed* spells of eligibility without receipt, that last longer than 24 months. Additionally, for both types of spells there are a number of durations shorter than 24 month for which there are only a small number of completions.

To deal with these data limitations I impose some restrictions on the model outlined above. First, $\gamma(t)$ continues to take the form of a step function, but with a potential step every 4 months, rather than every month. The assumption here is that AFDC/TANF entry and exit rates are constant over 4 month intervals. Because there are very few completed spells longer than 24 months, all spells lasting longer than 24 months are coded as right censored at 24 months. Artificially censoring the data in this way avoids the problem of having to impose very restrictive assumptions on the shape of $\gamma(t)$ beyond 24 months. It also avoids the problem of non-random sampling of long completed spells due to differences in the SIPP panel lengths.[13]

V. Results

Estimates of the coefficient vector β for AFDC/TANF exit and entry rates are shown in Table 2. Because the estimates shown in Table 2 correspond to a proportional hazards model, they may be interpreted as the marginal fraction changes in the AFDC/TANF entry (exit) rate associated with a small change in the associated independent variable. In addition to the variables shown in Table 2, the specifications also include 6 duration dummies that allow the AFDC/TANF entry and exit rates to shift every four months,[14] 50 state effects, 11 monthly dummy variables, and 3 seam dummy variables.[15]

Overall, the estimates are largely consistent with expectations. Age does not seem to have a large effect on AFDC/TANF exit, but is a significant determinant of AFDC/TANF entry, with older women being much less likely to enter AFDC/TANF than younger women. More educated women have lower AFDC/TANF entry rates, but educational attainment does not have a statistically significant effect on AFDC/TANF exit rates. Relative to white women, black and Hispanic women appear to have lower AFDC/TANF exit rates and higher AFDC/TANF entry rates.

Turning to the family composition variables, the estimated effects of *number of children under 18* on AFDC/TANF entry and exit are not large or statistically

[13] Because the 1996 panel runs for 48 months, while the longest pre-1996 SIPP panel runs for only 40 months, not artificially censoring the data in the manner described above might lead to over-sampling of long spells from the 1996 SIPP panel, relative to other earlier SIPP panels.

[14] In the model describing the AFDC/TANF entry rate these duration dummy variables are interacted with *on welfare prior to baseline*.

[15] Because SIPP respondents are asked every 4 months about whether they were on AFDC during the previous 4 months, there is some tendency for them to report starting spells of AFDC receipt during the first month covered by an interview and/or report leaving AFDC during the last month covered by an interview. This so called “seam bias” problem with the SIPP has the potential to cause biased estimates of the time varying covariates and the baseline hazard parameters if not addressed properly. To deal with the seam bias present in the SIPP, I include seam dummy variables in the models. These seam dummy variables indicate whether a person-month corresponds to the second, third, or fourth month covered by an interview.

Table 2. AFDC/TANF exit and entry rate estimates (robust standard errors in parentheses)

Variable	AFDC/TANF exit rate (N=2,843, N*T=22,102)	AFDC/TANF entry rate (N=4,863, N*T=51,617)
Age_t	0.0302	-0.0955 **
	(0.0246)	(0.0181)
Age_t squared (/100)	-0.0423	0.1002 **
	(0.0354)	(0.0255)
High school $graduate_t$ (vs. <high school)	0.0357	-0.1942 **
	(0.0676)	(0.0717)
Some $college_t$ (vs. <high school)	0.0968	-0.4420 **
	(0.0753)	(0.0841)
Black (vs. white)	-0.2193 **	0.2844 **
	(0.0687)	(0.0769)
Hispanic (vs. white)	-0.1831 **	0.2004 **
	(0.0889)	(0.0973)
Number of children under 18_t	-0.0321	0.0109
	(0.0286)	(0.0337)
$Preschool_t$ (=1 if respondent has a child less than 6-years old)	-0.1613 **	-0.0353
	(0.0783)	(0.0858)
$Baby_t$ (=1 of respondent has a child less than 2-years old)	-0.1100	0.3144 **
	(0.0732)	(0.0800)
Older $children_t$ (=1 if respondent's youngest child is 16 or 17)	-0.0410	-0.3300 **
	(0.1647)	(0.1999)
Real maximum $benefit_t$ (the maximum monthly AFDC/TANF benefit for a 3-person family in hundreds of US $\$_{2000}$)	-0.0450	-0.2330 **
	(0.967)	(0.1066)
$Waiver_t$ (=1)	0.0434	-0.1884
	(0.1057)	(0.1207)
$TANF_t$ (=1)	0.2039 **	-0.3753 **
	(0.1230)	(0.1425)
Unemployment $rate_t$ (state monthly unemployment rate, seasonally adjusted)	-0.0394	0.1754 **
	(0.374)	(0.0398)
Duration effects	Included	Included
State effects	Included	Included[a]
Panel year effects	Included	Included[b]
2 ln (*likelihood*)	9,120.25	9,182.29

Note: All estimates are unweighted. [a] The duration effects in the entry rate specification are interacted with the *AFDC/TANF prior to baseline*. [b] All specifications included 1992, 1992 and 1996 panel indicators. ** Statistically significant at the 5-percent level.
* Statistically significant at the 10-percent level.

significant, but some of the child age variables do appear to be important determinants of AFDC/TANF entry and/or exit. In particular, it is estimated that welfare recipients with a preschool aged child are approximately 16 percent less likely than observational equivalent women with older children to exit AFDC/TANF. The estimated effects of the ages of the children in the family on entry are even larger. Non-recipients with babies (children less than 2-years old) are 31 percent more likely to enter AFDC/TANF, relative to women with older minor children. Additionally, non-recipients whose youngest minor child is age 16 or older are 33 percent less likely to enter AFDC/TANF, relative to women with younger children.

Most of the state-level variables have the expected effects. Both of the welfare reform variables have positive effects on AFDC/TANF exit and negative effects on AFDC/TANF entry. While the effect of waiver implementation is not statistically distinguishable from zero in the exit or entry specification, the effect of TANF implementation is statistically significant in both. It is estimated that TANF implementation increases the exit rate by 20 percent and decreases the entry rate by 38 percent. *Unemployment rate* has a small negative effect on AFDC/TANF exit and a large positive effect on AFDC/TANF entry. Based on the estimates reported in Table 2 a 1-point increase in the state monthly unemployment rate would increase AFDC/TANF entry rates by 18 percent.

The only state-level variable with coefficient estimates that differs substantially from anticipated values is *real maximum benefit*.[16] The estimated effect of the real value of the maximum AFDC/TANF benefit for a 3-person family on AFDC/TANF exit is small, negative, and not statistically distinguishable from zero. While *real maximum benefit* does not appear to affect the process determining welfare exit, it

[16] In interpreting the coefficients associated with *real maximum benefit* it is important to note they are identified on the basis of dubious sources of variation. The problem in trying to identify the effect of changes in maximum AFDC/TANF benefits during the 1990s is that most states did not alter their benefit schedules very often during this period. When they did alter their schedules, the changes tended to be modest. This fact means that once the variation in maximum AFDC/TANF benefits between states and between panels is removed (by virtue of state effects and panel year effects being included in the models) there is not much meaningful variation left over on which to base the coefficient estimate. What little variation that remains is primarily attributable to within panel variation in the Consumer Price Index, small incremental changes in nominal maximum AFDC/TANF benefits, and larger changes in AFCD/TANF benefits schedules in a select few states.

does have a modest and statistically significant estimated effect on welfare entry. It is estimated that a $100 increase in the maximum AFDC/TANF benefit for a 3-person family would *decrease* the AFDC/TANF entry rate by over 23 percent.

In addition to the results shown in Table 2, a number of alternative specifications were run. One set of alternative specifications attempted to tease out the effect of various waiver provisions by coding them separately in the entry and exit rate equations. These results were somewhat unsatisfactory as none of the disaggregated waiver variables had large and/or statistically significant effects in either the entry or exit rate specifications. Another set of alternative specifications was run to gauge the sensitivity of the results to using SIPP panel weights. Overall, estimates obtained using weighted data are substantively similar to those reported in Table 2. I also ran other specifications which excluded fixed effects (in favor of a South region dummy variable) and experimented with different ways to control for secular time effects. Not surprisingly, the estimated effect of variables measured at the state-level tended to be larger in magnitude and significance in specifications that omitted state and/or time effects.

A. Explaining caseload changes

The estimates reported in Table 2 imply that there should be a measurable impact of macroeconomic conditions and welfare reform on AFDC/TANF caseloads, but they do not give any indication as to how large these effects actually are. To garner insights about the contributions of welfare reform and macroeconomic performance in explaining the AFDC caseload changes of the early to mid 1990s, I use data from the 1991 through 1998 March Current Population Surveys (CPS) along with the estimates reported in Table 2 to determine the impact of unemployment rate changes, the implementation of statewide pre-PRWORA welfare reform waivers, and TANF implementation on the AFDC/TANF-Basic caseload share.[17]

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[17] The definition of the AFDC/TANF-Basic caseload share used in this study is the number of AFDC/TANF-Basic cases (assistance unit count) denominated by the number of single women with children. Estimates of the number of single women children were sourced from the 1991-1998 CPSs. This definition of the AFDC/TANF-Basic caseload differs slightly from that used in other studies. Most other studies define the caseload share as the AFDC/TANF Basic or Total caseload denominated by the total population or the number of females aged 15 to 44. In so-much-as changes in the denominators of the respective caseload measures do not change

This analysis proceeds in two parts. The first set of results relates changes in state unemployment rates and welfare reform variables to changes in the long-run AFDC/TANF-Basic caseloads share. These estimates will give some indication as to whether the estimates reported in Table 2 are broadly consistent with the estimates reported in the caseload literature, but will not give any indication as to how well the dynamics implied by the Table 2 estimates reflect actual caseload dynamics. To determine how well entry and exit rate estimates do at explaining the timing of actual caseload changes, a series of simulations are conducted. These simulations are aimed at addressing commonly asked questions in the caseload studies: What percentage of the run up in caseloads during the early 1990s is attributable the deteriorating macroeconomic conditions and how much of the sharp decrease in caseloads is attributable to improving economic conditions versus welfare reform?

Table 3 shows the effect of increasing state unemployment rates by 1-percentage point, implementation of statewide welfare reform waivers, and national TANF implementation on average AFDC/TANF entry and exit rates, and the long-run AFDC-Basic caseload share. These estimates were obtained by applying the parameter estimates of the AFDC/TANF entry and exit rate in Table 2 to a sample of AFDC/TANF-Basic recipients and eligible non-recipients from the 1994 March CPS. Depending upon whether individual respondents reported receiving AFDC in 1993, an AFDC/TANF entry or exit rate was imputed.[18] The first two rows of Table 3 show the effects of increases to the state unemployment rate and welfare reform implementation on these imputed entry and exit rates. The third row of Table 3 shows how changes in the imputed entry and exit rates corresponded to changes in the long-run AFDC-Basic caseload share.[19]

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rapidly, all caseload share measures will indicate very similar patterns over time. I use the caseload share measure described above because, for the purposes of the simulations that follow, it is important that the AFDC/TANF-Basic caseload standardized by the population at immediate risk for receipt. Roughly speaking, the population immediately at risk for AFDC/TANF-Basic receipt is the population of female family heads.

[18] The March CPS asks respondents about their program participation during the previous calendar year. In particular, the March 1994 CPS indicates a respondent's recipient status for the calendar year beginning in January of 1993.

[19] The long-run AFDC/TANF-Basic caseload share can be computed rather straightforwardly provided information on the average entry and exit rates. Assuming a steady-state, the AFDC caseload share is given by $\overline{C} = \overline{In} / (\overline{In} + \overline{Out})$, where $\overline{In}$ and $\overline{Out}$ represent the average rates of AFDC/TANF-Basic entry and exit.

Table 3. Effects of a change in AFDC benefit levels, unemployment rates, and welfare reform

	1- point increase in URATE	Waiver implementation	TANF implementation
Percentage change in the average AFDC/TANF entry rate	17.49	-18.79	-37.10
Percentage change in the average AFDC/TANF exit rate	-3.94	4.34	20.32
Percentage change in caseload share	11.27	-10.87	-24.20

Overall, the results reported in Table 3 compare favorably with the estimates from caseload literature. A 1-point increase in the unemployment rate leads to an increase in the caseload share of 11 percent. This estimate is slightly larger than those from the caseload literature that suggest that a 1-point increase in the unemployment rate would lead to a 3 to 8 percent increase in the aggregate caseload share (Levine and Whitmore 1998; Wallace and Blank 1999; Figlio and Ziliak 1999; Ziliak et al. 2000; Blank 2001). The Table 3 estimate indicates that waiver implementation will reduce the caseload share by 11 percent. This estimate is similar to the estimated waiver effects from most of the caseload studies that estimate that the effect of waiver implementation is to reduce the caseload share by 8 to 10 percent (Levine and Whitmore 1998; Wallace and Blank 1999; Blank 2001), and is larger than the waiver effect reported in some studies that add lagged caseload values as dependent variables (Figlio and Ziliak 1999; Ziliak et al. 2000). The few studies in the caseload literature that explicitly model TANF effects estimate that TANF implementation decreases the caseload share by 30 to 40 percent (CEA 99, Wallace and Blank 1999). The estimated TANF effect of -24 percent from Table 3 is reasonably consistent with these estimates.

The results in Table 3 indicate that the AFDC/TANF entry and exit rate estimates reported in Table 2 imply a long-run relationship between unemployment rates, welfare reform, and AFDC/TANF caseloads that is somewhat consistent with trends in caseloads throughout the 1990s, but do not indicate whether the path of adjustment to shocks in macroeconomic conditions or welfare reform are consistent with observed trends in caseloads.

To determine how well the entry and exit rate estimates in Table 2 predict changes in the actual AFDC/TANF-Basic caseload I perform a series of simulations.

In performing these simulations I use data from the 1991 to 1998 CPSs. From the CPSs, I extracted a sample of AFDC/TANF-Basic and potential AFDC/TANF-Basic recipients (unmarried women with at least one child). Using the coefficients estimates from Table 2 along with the CPS extracts, I calculated the weighted average of the AFDC/TANF entry and exit rates for each month between January 1990 and June 1997. These estimates of the average monthly AFDC/TANF entry and exit rates were used to simulate the effect of changes in state unemployment rates, waiver implementation, and early TANF implementation on the aggregate AFDC/TANF-Basic caseload share.[20]

Before making any attempt to determine the extent to which changes in the AFDC/TANF-Basic caseload share are explained by changes in macroeconomics conditions or welfare reform, it is useful to examine how well the monthly AFDC/TANF entry and exit rates implied by the Table 2 estimates fit the actual caseload data. Figure 1 plots the actual AFDC/TANF-Basic caseload share against the Predicted (Simulated) AFDC/TANF-Basic caseload share. Considering the fact that the predicted caseload series in Figure 1 was constructed by taking estimates from the SIPP and applying these estimates to data from the CPS, the projected caseload share fits the actual caseload share remarkably well. The entry and exit rate estimates in Table 2 do predict the sharp rise in the caseload share between 1990 and 1993 and much of the decrease in the caseload share following 1993.

Table 4 contains tabulations of the actual and predicted change in the AFDC/TANF-Basic caseload share from January 1990 to January 1993 and from January 1993 through June of 1997. The simulations shown in Table 4 take into account the effects of unemployment rate changes, waiver implementation, TANF implementation, changes in the generosity of AFDC/TANF benefits, as well as the changing demographic composition of the AFDC/TANF-Basic caseload and pool of eligible non-recipients. Between January 1990 and January 1993 the AFDC-Basic caseload share increased by 0.055. Changes in state unemployment rates explained 130 percent of this increase in the caseload share. Waiver implementation reduced the caseload share by 2 percent below its projected level in absence of reform. All else held constant, changes in the demographic composition of the caseload, the composition of the pool of eligible non-participants, and benefit levels would have increased the caseloads share by nearly 13 percent.

[20] Details of these simulations are provided in the Appendix.

Figure 1. AFDC-Basic caseload share versus the predicted AFDC-Basic caseload share

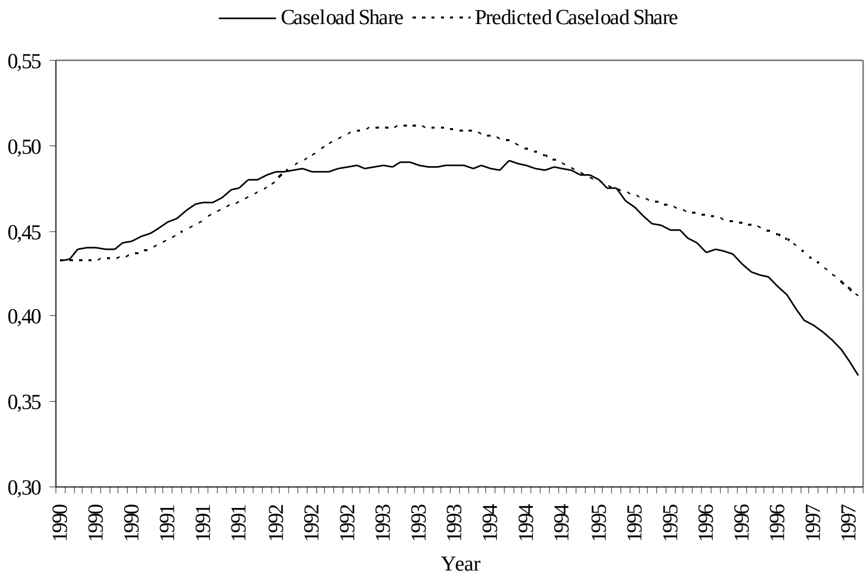

Table 4. Actual versus predicted change in the log AFDC/TANF-Basic caseload share

	1990 to 1993	1993 to mid-1997
Actual change	0.055	-0.122
Percentage change due to unemployment	130.30	74.32
Percentage change due to waiver implementation	-2.32	7.00
Percentage change due to TANF implementation	—	28.76
Percentage change due to other factors	12.65	-20.18
Percentage change explained	140.46	89.90

Between January of 1993 and June 1997 the AFDC/TANF-Basic caseload share declined by 0.12 (a 25-percent decrease). Seventy-four percent of this decrease in the caseload share can be attributed to improvement in economic conditions as measured by state monthly unemployment rates. Waiver implementation accounted for approximately 7 percent of the caseload decline between January 1993 and mid-1997, while TANF implementation also accounted for an estimated 29 percent of decline in the caseload share over the same period. Changes in maximum AFDC/TANF benefit levels and the demographic composition of the pool of actual and potential AFDC/TANF recipients would have, in absence of welfare reform and changing economic conditions, led to an increase in the caseload share.

Compared to studies that use annual state-level panel data the model that is used to conduct the simulations in Table 4 predicts a much larger share of both the run-up of the caseload share during the early 1990s and the decrease in the caseload share between 1993 and mid-1997. Using annual state-level panel data, Blank (2001) attributes about 23 percent of the caseload share to changes in annual state unemployment rates. Among the studies that examine the decline in caseloads after 1993, most conclude that 30 to 40 percent of the decline in caseloads can be explained by improving economic conditions. In these studies, waiver implementation accounts for 15 to 20 percent of the decline in the caseload share. The exceptions are papers by Figlio and Ziliak (1999) and Ziliak et al. (2000) which attribute 75 and 66 percent of the caseload decline respectively to improvements in economic conditions, and a negligible share of the caseload decline to waiver implementation.

VI. Conclusions

One question that has received a great deal of attention concerns the extent to which implementation of pre-PRWORA waivers and an improving economic climate have contributed to the dramatic caseload declines that took place between 1993 and TANF implementation. Among the studies that use annual state-level panel data to address this question, a consensus as to the relative contributions of welfare reform and economic growth has not emerged. Some studies conclude that both macroeconomic factors and waiver implementation were important determinants of the caseload decline while other studies emphasize the importance of macroeconomic factors at the expense of waiver implementation.

In this paper simulations are presented that provide alternative estimates of the effects of welfare reform and unemployment rate changes on the caseload changes of the early to mid 1990s. These simulations indicate that increases in unemployment rates accounted for all increase in the AFDC caseload share between 1990 and 1993, and that both decreases in unemployment rates and early TANF implementation were important in explaining the decline in the AFDC caseload share between 1993 and mid-1997. State unemployment rate decreases accounted for approximately 74 percent of the decline in the caseload share between 1993 and mid-1997, while early TANF implementation accounted for 29 percent of the decrease over the same period; pre-PRWORA waiver implementation accounted for a mere 7 percent of the caseload share decline.

In relation to some of the caseload studies (Levine and Whitmore 1998; Wallace and Blank 1999, Blank 2001), the results presented in this paper place more importance on the macroeconomy and less importance on waiver implementation in explaining the caseload decline. The results presented in this paper are largely consistent with the conclusions of the subset of caseload studies that estimate specifications with lagged dependent variables and conclude that most of the decrease in the caseloads share post-1994 was due to economic conditions (Figlio and Ziliak 1999; Ziliak et al. 2000).

Appendix

The purpose of this appendix is to provide the details of the caseload change simulations presented in Table 4 of this paper. Note that the AFDC/TANF caseload share, denoted C_t, evolves according to the first order difference equation

$$C_{t+1} = C_t + \overline{In}_{t+1}(1 - C_t) - \overline{Out}_t(C_t) \qquad \text{(A1)}$$

where $\overline{In}_{t+1}$ is the average rate of inflow into the AFDC/TANF program from the pool of eligible non-recipients in month t+ 1 and $\overline{Out_t}$ is the average AFDC/TANF exit rate in month t. Provided estimates of the $\overline{In}_k$ through $\overline{In}_n$ and $\overline{Out}_k$ through $\overline{Out}_{n-1}$ and knowledge of the caseload share in the base month k, it is possible to simulate the path of the caseloads share between month $k + 1$ and period n.

Estimates of the sequence of average monthly AFDC/TANF entry and exits rates were calculated by obtaining predicted AFDC/TANF entry and exit rates using the coefficients from Table 2 applied to samples of categorically eligible non-recipients and AFDC/TANF recipients drawn from the CPS. These individual predicted entry and exit rates were then averaged over the appropriate CPS sample to produce estimates of the sequence of average AFDC/TANF entry and exit rates.[21]

[21] Variable definitions from the CPS match up very well with the definitions from the SIPP, so applying coefficient estimates obtained from SIPP data to samples from the CPS should not be a problem. Table A1 provides the means and standard deviations of the CPS sample used to conduct the caseload simulations.

Most of the coefficients in Table 2 can be applied to the CPS sample straightforwardly. There are, however, a number of variables that are only applicable to the SIPP sample, and for which values must be imputed in the CPS sample prior to applying the Table 2 coefficients. In particular, seam, duration and month effects are included in the model in Table 2, but the CPS does not have seams, and there is no way of knowing how long women in the CPS sample have been receiving AFDC/TANF (or were categorically eligible without receipt); nor is it possible to determine which months they were receiving AFDC/TANF (or were categorically eligible without receiving AFDC/TANF).

The basic approach for dealing with the absence of seam and month variables in the CPS samples was to use an average of the seam and month effects to compute predicted values of AFDC/TANF entry and exit. Thus, the month and seam effects for the entry and exit rates used in the Table 4 simulations are computed as the average of the 11 month and 3 seam effects in each specification.

To impute duration effects for the purpose of obtaining predicted AFDC/TANF exit rates, I first tabulated the fraction of current (ongoing) AFDC/TANF spells in a stock sample of AFDC/TANF recipients from the 1996 SIPP Wave 1 Topical Module (W2TM1) that fell into categories corresponding to the duration effect groupings. The AFDC/TANF exit rate for each CPS recipient was then computed as a weighted linear combination of exits rate at each of the 6 duration length groupings (see Section III), where the fraction of ongoing spells from the W2TM1 that fall into each of the groupings were used as weights

In absence of any information about the length of spells of categorical eligibility without receipt, I calibrated the duration effects in the entry specification so that when they were applied to women in a CPS sample of categorically eligible non-recipients, averaged, and used in conjunction with the average AFDC/TANF exit rates described above, the resultant caseload share was in a steady state at the beginning of each simulation period (January 1990 and January 1993). This calibration ensures that there will be no change in the simulated caseload share going forward unless there are changes in the variable values (excluding the month, seam, and duration effects).

To calculate the percent change due to unemployment shown in Table 4 I fixed the characteristics of the sample and allowed only the state unemployment rate to vary with time. This produces a sequence of estimated monthly AFDC/TANF entry and exit rates. These sequences of estimated entry and exit rates were used to simulate a sequence of caseload share values. These simulated values of caseload share were then compared with the actual caseload share to produce the percentage changes in the caseload share due to unemployment reported in Table 4.

The values of the percentage change due to waiver implementation, real benefit levels, and other factors were calculated similarly. The total percent of the caseload change numbers in the last row of Table 4 were computed by allowing all of the variables in the model to change simultaneously. Because the model is non-linear, the sum of the percentage change in the caseload due to waiver implementation, TANF implementation, changes in benefit levels, and demographic factors need not equal the total percent of the caseload share explained.

The means and standard deviations of the CPS Sample used to conduct the simulations are provided in the Appendix Table below.

Table A1. CPS sample means and standard deviations

Variable	AFDC/TANF recipients (N=10,639)	Eligible non-recipients (N=26,603)
Age	30.2728	34.7741
	(7.7865)	(8.6553)
Less than high school (=1)	0.4007	0.1751
	(0.4901)	(0.3800)
High school graduate (=1)	0.3794	0.3921
	(0.4853)	(0.4882)
Some college (=1)	0.2198	0.4328
	(0.4142)	(0.4955)
White (=1)	0.3998	0.5777
	(0.4877)	(0.4939)
Black (=1)	0.4327	0.2958
	(0.4955)	(0.4564)
Hispanic (=1)	0.1774	0.1265
	(0.3820)	(0.3324)
Number of children under 18	2.1290	1.6082
	(1.2163)	(0.8384)
Preschool (=1 if respondent has a child less than 6-years old)	0.6358	0.3870
	(0.4812)	(0.4871)
Baby (=1 of respondent has a child less than 2-years old)	0.2715	0.1330
	(0.4447)	(0.3395)
Older Children (=1 if respondent's youngest child is 16 or 17)	0.0257	0.0946
	(0.1636)	(0.2927)

Note: Source is 1991 through 1998 March Current Population Survey. Means and standard deviations are computed using weights.

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