The Rise in SSI Participation among Children: Assessing the Impact on Poverty and Labor Supply*

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Abstract

From 1989 to 1996 the number of children receiving benefits from the federal Supplemental Security Income (SSI) program increased by 260 percent to 955,000. Some recent work has documented the shift of children from welfare to SSI during this period but there has been little work examining the consequences for families. In this paper we aim to present new information about the rise in child participation in SSI and its determinants and to empirically investigate the effects on family poverty and maternal labor supply. We utilize an Instrumental Variables/Triple Differences approach that exploits the fact that boys were 85 percent more likely than girls to enroll in SSI during the 1990s. Additionally, we focus on children living in female-headed families because they were several times more likely than other children to receive SSI. Our findings suggest that among female-headed families, the increase in SSI participation resulted in an approximately 2 percentage point decline in child poverty and a 2 percentage point increase in maternal labor supply during the 1990s.

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I. INTRODUCTION

Since its inception in the early 1970s, the federal government's Supplemental Security Income (SSI) program has provided cash benefits and Medicaid health insurance to low-income aged, blind, and disabled individuals. Both children and adults have always been eligible for this means-tested program, but throughout the 1970s and 1980s adults were several times more likely to be receiving SSI. This disparity was partly attributable to the stricter medical eligibility criteria that the Social Security Administration (SSA) used in its disability determinations for children. For an adult to meet the disability standard, he must be deemed unable to engage in any substantial work. In the 1990 Supreme Court case of *Sullivan v. Zebley*, the Court ruled that in order to meet the program's legislative standard of equal treatment, a functional limitation component comparable to that for adults must be included in the determination process for children. This decision had the effect of lowering the level of severity required for children to be eligible for SSI benefits [U.S. GAO 1994, 1995].

At the time of the *Zebley* ruling, there were 265,000 children receiving SSI benefits, with this number having increased by less than 3 percent annually in the preceding three years. Seven years after the Supreme Court decision, the number of children on SSI had increased by 260 percent to more than 955,000. This growth came to a halt after the passage of the 1996 Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA), which tightened child eligibility criteria and resulted in the termination of benefits for nearly 100,000 SSI children who were found to be "no longer disabled". While PRWORA effectively put an end to further expansion of SSI, the fraction of children receiving benefits remained substantially higher than it was in 1989 (Figure 1A).

In this paper we aim to estimate the effect of the growth in SSI enrollment among children on family poverty and maternal labor supply. Identifying a causal relationship is difficult for at least three reasons. First, because the program is means-tested, there is an endogenous relationship between family resources and SSI participation. Second, SSI is a federal program with very little variation in program parameters across states. While some states do supplement their child SSI benefits, these state-level payments account for just three percent of SSI payments to children. And third, the growth in child SSI enrollment occurred during the same period as welfare reform, the expansion of the Earned Income Tax Credit (EITC), growth in the adult SSI program, and other factors that make it difficult to disentangle its effect from the effect of these other factors.

We exploit two key sources of variation to overcome these difficulties: (1) variation across households in the number of boys and (2) variation across states in welfare benefit levels. The growth in SSI participation over the nineties was 85 percent greater for boys than for girls, and by 1999 the ratio of boys to girls on the program was 1.74. This disparity appears to be driven by differential rates of mental disorder diagnoses between boys and girls. Boys living in female-headed families were particularly likely to enroll in SSI after *Zebley*. This differential participation provides the basis for a reduced form “triple differences” approach and a Two Stage Least Squares (2SLS) approach to identifying the causal effect of SSI participation on family outcomes. The key assumption of our identification strategy is that families headed by single mothers with boys would have experienced the same change in outcomes during the 1990s as families headed by single mothers with girls were it not for the differential participation in SSI. Our empirical strategy also exploits heterogeneity in welfare benefit generosity across states, which influenced both the probability of an SSI application and the increment to the family’s transfer income if the child enrolled. The advantage of this identification strategy is that our results are unlikely to be biased by welfare reform, expansions in the EITC, or the increase in adult SSI enrollment.

While dozens of papers have studied the effects of AFDC and TANF on family outcomes, few have studied the consequences of child participation in SSI. This is no doubt partially because there were fifteen times more families on AFDC than there were families with a child on SSI before the *Zebley* ruling. But by the end of 2003, owing to the concurrent rise in child SSI enrollment and the contraction of the welfare caseloads, that ratio had fallen to 2.1. If one additionally considers SSI receipt by adults, there are now slightly more children in households receiving SSI payments than in households receiving any benefits from TANF.

Previous work has demonstrated that SSI is to some extent a substitute for AFDC or TANF (Kubik, 2003, 1999; Garrett and Glied, 2000; and Sevak and Schmidt, forthcoming). In the first empirical section of our paper, we build on this previous work using annual state-level data on welfare caseloads and SSI application, award, and enrollment rates among children. Our findings demonstrate that prior to the *Zebley* ruling there existed a weak positive relationship between state-level changes in SSI and AFDC enrollment. This is perhaps not surprising given that an economic downturn would tend to increase demand for both programs, with this more than offsetting the substitution described above. But during the seven years after the *Zebley* ruling, there is a significant negative relationship between changes in AFDC and SSI enrollment, suggesting that a large fraction of those newly eligible for SSI would otherwise have received AFDC benefits. We further demonstrate that children in states with low AFDC benefits at the time of the *Zebley* ruling were significantly more likely to apply for and be awarded SSI benefits than those in high AFDC benefit states. We supplement this set of results with county-level specifications and reach similar conclusions.

Theoretically, the effect of SSI on poverty is ambiguous. If SSI leads to a reduction in other transfer income or in earnings then it is possible that family income could decline as a result of a successful SSI application. The effect on labor supply is even more difficult to predict given that SSI is more generous than welfare in most states and yet the program's labor supply incentives are much stronger. Given that the income and substitution effects will typically go in opposite directions both the sign and the magnitude of the effect of child SSI is thus ultimately an empirical question.

Our results for poverty suggest that enrollment in SSI substantially lowers the probability that a child will be poor, with the estimates indicating a much larger effect in low AFDC benefit states. Our point estimates imply that for every five children made eligible for SSI, three are lifted above the poverty line. Our findings also suggest that the expansion of child SSI has increased maternal labor supply, with strong and significant effects for the probability of work, the number of weeks worked last year, and the usual hours worked per week. Taken together, our results strongly suggest that the expansion of child SSI during the 1990s contributed to the decline in child poverty and to the increase in labor supply among single mothers.

The remainder of the paper proceeds as follows. Section 2 provides background information about the SSI program and the rise in the SSI child caseload. Section 3 discusses previous work and new empirical evidence on the switching from AFDC to SSI in the 1990s. Section 4 presents our empirical approach to identifying the effect of child participation in SSI on family poverty and then describes the results. Section 5 discusses the effects on maternal labor supply. And Section 6 provides a concluding discussion.

II. BACKGROUND

The first SSI payments were disbursed in January of 1974, when 51 state-level programs that had assisted low-income aged, blind, and disabled adults were consolidated into one federal program. In its first year, most SSI recipients were above the age of 65 and total benefits paid out were 34 percent lower than in the federal-state AFDC program. Virtually no children were transferred from the state programs though more than 70,000 children were receiving benefits by the end of this first year. Fifteen years later, the number of children age 0 to 17 receiving SSI benefits had increased to 265,000, with most of the increase occurring during the 1970s.

From 1989 to 2002, the number of SSI recipients increased by 48 percent while the corresponding number receiving welfare fell by more than 54 percent. SSI is now the largest means-tested cash-assistance program in the country, with the $34.6 billion in 2002 expenditures exceeding welfare spending by 242 percent.[1] At the end of 2002, the SSI caseload consisted of roughly 4 million adults age 18 to 64, nearly 2 million age 65 or greater, and just short of 1 million below the age of 18. Comparing these figures with the corresponding enrollment data from December of 1989, the number of recipients age 18 to 64 increased by 73 percent, the number of elderly declined by 2 percent, and the number of children increased by 265 percent.

II.A. Trends in child SSI caseloads

The number of children receiving SSI increased substantially after the 1990 Supreme Court decision in *Sullivan v. Zebley*, rising from approximately 265,000 in 1989 to over 950,000 just seven years later. In percentage terms, this represented an increase from 0.4 to 1.4 percent of all children between the ages of 0 and 17. This period of rapid growth represented a sharp break in the slight upward trend prior to *Zebley*: during the four years from 1985 to 1989 the number of

[1] The actual outlay of benefits associated with the federal SSI program is much higher: Medicaid spent more than $150 billion on medical care for SSI recipients in 2003.

children on SSI increased by just 37,500. Figure 1A plots the percentage of children on SSI from 1985 to 2003 and Figure 1B plots the percentage of children applying for or awarded SSI. As revealed by this latter figure, there was a noticeable spike in applications and awards after the *Zebley* ruling, with these two series peaking in 1993 and 1994, respectively.

The period of rapid growth in child SSI caseloads ended in 1996 with the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA). This legislation eliminated the "comparable severity" standard for individuals under age 18 and replaced it with the requirement that a child must have a medically determinable impairment that results in "marked and severe functional limitations" and meets the existing statutory duration requirement. The law also eliminated references to "maladaptive behaviors" in the Listing of Impairments for children and discontinued the use of individualized functional assessments for children. Figure 1C shows that there was a spike in 1997 in the fraction of child SSI recipients who had their benefits suspended for the reason "no longer disabled", with more than 10 percent of SSI children dropped from the rolls during this year. This annual suspension rate has remained substantially greater than the pre-1997 level since the passage of PRWORA.

The expansion of child participation in the SSI program has made it much more relevant as a source of assistance for poor families with children. As shown in Figure 2, the ratio of AFDC/TANF families to SSI children has fallen substantially since 1985, with an especially rapid decline occurring from 1989 to 1996. Given that few SSI children have a sibling under the age of 18 also on the program, this is a reasonable approximation to the ratio of families on welfare to the number of families with a child on SSI. The change from 14.2 to 4.6 during this seven-year period was entirely driven by the expansion in SSI enrollment, as the number of

AFDC families actually increased by 17 percent. This ratio fell to 2.2 during the next six years, with this decline explained by changes in welfare caseloads rather than in SSI receipt.

There are many different medical conditions with which children qualify for SSI and the frequency of these changed substantially after the *Zebley* ruling. In December of 1989 the most common condition was mental retardation, which accounted for almost 42 percent of child SSI cases. As shown in Table 1, an additional 6 percent of SSI recipients in this base year qualified because of some other mental disorder. In the seven years following the Supreme Court decision, the number of children qualifying in this category grew by 1324 percent (from 16,495 to 234,935) with all other conditions growing by a still substantial 190 percent (from 248,395 to 720,239). Because of the stricter standards resulting from the PRWORA legislation the number of children on SSI declined by 11 percent from 1996 to 1999. Figure 1A reveals that despite this decline, the fraction of children receiving SSI benefits was approximately three times greater in 1999 than in 1989.

Growth in SSI receipt was substantially greater among boys than girls in the years following the *Zebley* ruling. The change from 1989 to 1999 in the number of boys on SSI was 85 percent greater than the corresponding change in the number of girls. Though boys were more likely to be on the program in the base year, their share of SSI child cases increased from 58.5 percent in 1989 to 63.4 percent by 1999.[2] The differential growth among boys was to a large extent caused by their greater likelihood of qualifying in the "other mental disorder" category, the most rapidly growing diagnosis category during this time period. The ratio of boys to girls in this category was 3.01 in 1999 versus 1.47 for all other conditions.[3]

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[2] The number of males age 22 to 29 on the SSI program is just 11 percent higher than the corresponding number of female SSI recipients.

[3] The claim that boys are more likely than girls to experience mental disorders is widely supported in the clinical psychology literature. Gender has been identified as the most consistently documented risk factor for Conduct

II.B. Program Parameters and Rules

Eligibility requirements and federal payment standards for SSI are nationally uniform, though states have the option to supplement the federal SSI payment.[4] Since 1975, these rates have been increased by the same percentage as the cost-of-living increases in Social Security benefits. In 2004 the maximum federal SSI payment was $564 monthly for an individual and $846 monthly for a couple with both the husband and wife eligible. Children on SSI are significantly more likely than adults to receive the maximum SSI payment,[5] with the average payment to an SSI child in May of 2004 equal to $507.40. This average payment reflected an average federal payment of $491.30 and an average state supplement of $16.10. As the only source of income, these payments are not sufficient to lift a family out of poverty. However, they could do so if the family has other sources of income, such as earnings or transfers from other programs. SSI payments could also be effective in lowering the poverty deficit.

The federal SSI payment for an adult recipient is based on the individual's countable income. The first $20 of unearned or earned income is excluded, as is the first $65 of monthly earnings plus one-half of any earnings above $65. Individuals generally are not eligible for SSI if they have resources in excess of $2,000. Certain resources are excluded, most commonly a home, an automobile, household goods and personal insurance of reasonable value, burial plots and spaces, and life insurance. In the case of a child SSI recipient, some of the income and assets of

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Disorder (CD) (Robins, 1991). Through childhood, boys greatly outnumber girls with respect to diagnosed CD, with ratios of four to one; Oppositional Defiant Disorder (ODD) is also believed to be more common in boys than girls, at least through preadolescence; Rates of Antisocial Personality Disorder (APD) and Substance Use Disorders are also far higher among boys, as are the predictive risks from early to later forms of antisociability (American Psychiatric Association, 1987). Autism occurs more frequently in male individuals with approximately three or four males for every one female with autism (Bryson et al., 1988; Steffenburg & Gillberg, 1986; Cokmar, Szatmari & Sparrow, 1992). Childhood-onset cases of Schizophrenia also appear to be in excess among males (Green, Padron-Gyol, Hardesty, & Bassiri, 1992; Werry, 1992).

[4] From 1975 to 1996, 23 states have continuously provided supplemental SSI payments. No state increased supplements faster than inflation; states have allowed inflation to erode supplements or have reduced them in the face of state fiscal problems.

[5] The main reason for this seems to be that many adults are dually eligible for social security benefits which will lower their SSI benefits below the maximum.

certain family members living in the same household are “deemed” to the recipient. Payments from AFDC/TANF to other household members are excluded from deeming, as are foster care payments, food stamps, and EITC benefits to anyone in the household. Household income that is used by another public assistance program to determine the payment amount to someone other than the SSI recipient is also excluded from deeming. There is an allowance for each ineligible child as well as a parental living exclusion.

As an example, consider a family in 2000 comprised of two children and one parent that had only two sources of family income, an SSI payment to one child and $1,300 in monthly earnings. Starting with the $1,300 in earnings, the SSA first deducts $256 for the other child’s monthly allowance, $20 for the general income exclusion, and $65 for the earnings exclusion. One-half of the remaining $959 in earnings is excluded, bringing deemed income to $479.50. The parental living allowance of $513 would then be subtracted leaving zero income deemed to the child and thus he would qualify for the maximum benefit. Consider instead that the family consisted only of the parent and eligible child. Then the deeming calculation would exclude the monthly allowance of $256 and thus $94.5 would be deemed to the child resulting in an SSI payment of $417.50. With only earned income, one parent, and one SSI eligible child, for the child’s SSI benefit to fall below the maximum, monthly earnings would need to be above $1,111 if no other child, $1,367 if one other child, and $1,623 if two other children in the household.

II.C. SSI benefits compared to welfare benefits: Heterogeneity across states

In early 1990 just prior to the *Zebley* ruling, there was substantial heterogeneity across states in AFDC benefit generosity. For example, a woman with one child in the state of Alabama could receive $98 per month from AFDC while her counterpart in California could take in almost six times as much at $580. If the child became eligible for SSI in that year then he would no longer receive benefits from AFDC. While technically the mother could still be eligible for AFDC, there appear to have been virtually no "zero-child" AFDC cases since the *Zebley* ruling. Given, a $386 maximum SSI benefit in 1990, the family in Alabama would experience a $288 increase in monthly transfer income if the child became eligible for SSI. Though California did supplement its child SSI benefit, the family with one child would have experienced an $81 decline in monthly transfer income if the child became eligible for SSI.

Things are somewhat more complicated for a family on welfare with two or more children. In that case, if one of the children qualified for SSI, the other two family members could continue to receive AFDC benefits. For these families, SSI was still more financially attractive for families in low-benefit states because the decline in welfare benefits moving from N to N-1 AFDC recipients was smaller. For example, in Alabama the difference in monthly AFDC benefits for a two and three-person unit was just $20 versus $125 in the state of California.

Figure 3 suggests that the generosity of AFDC benefits was strongly related to the probability that children applied for SSI as a result of the *Zebley* ruling. The figure displays annual SSI application rates among children in the two states with the lowest AFDC benefits (Alabama and Mississippi) and the two with the highest (Alaska and California). As is clear from the figure, there was a much larger response to the liberalized eligibility criteria in the low-benefit states, though these states also had higher application rates to begin with. We proceed to a more systematic analysis of this relationship in the next section.[6]

[6] State governments had some financial incentives to facilitate the shift from welfare-to-disability because they paid an average of 46 percent of AFDC benefits versus less than 5 percent of child SSI benefits in 1990. To the extent that differences across states in this financial incentive were important, they would tend to bias against a finding that child SSI applications, awards, and enrollment increased more in low AFDC-benefit states for two reasons. First, high-AFDC benefit states would experience a larger fall in AFDC spending after moving a family or an individual from welfare to disability. Second, high-AFDC benefit states tended to be wealthier and thus paid a higher fraction of benefits for their AFDC recipients (prior to PRWORA the AFDC match rate was inversely related to a state's per-capita income). Thus while Alabama's share of AFDC spending would have fallen by $26 if it moved a two-person

III. SWITCHING FROM WELFARE TO SSI

The SSI program serves a population that overlaps to a significant extent with the welfare population. Children from female-headed households constitute a disproportionate share of the SSI child caseload. As shown in Table 3, 57 percent of children on SSI in 1999 lived in a female-headed household compared to just 21 percent of all children. Nineteen percent of children on SSI lived in poverty in 1999 though this is substantially lower than the corresponding share of 67% for children with some TANF income in their household.[7]

An individual cannot receive benefits from both of these programs, but as shown in Table 2, many families do. This is possible, for example, by having a parent on SSI and a child-only welfare case, or a child on SSI and a parent and sibling(s) on welfare. According to the 1990 *SIPP*, 39.2 percent of households with children and receiving any SSI payment also received some TANF income while 11.5 percent of households with children and with some TANF income also received SSI.[8] The corresponding percentages a decade later were 19.6 and 22.3.

Table 2 demonstrates the extent to which the growth in SSI has offset the decline in welfare caseloads. From 1989 to 2001 the fraction of children receiving AFDC/TANF fell by more than half from 10.2% to 4.8%. But the corresponding decline in the share receiving either AFDC or SSI was much less marked, falling from 11.7% to 8.5%. Thus for every 100 children in households that no longer have any AFDC/TANF income, there are 41 additional children in households with some SSI payments. It is interesting to note that, while in 1989 there were

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family off of this program in 1990, California's would have declined by more than ten times as much. The fact that California and other high AFDC-benefit states supplemented SSI would to some extent offset this difference.

[7] These demographic characteristics come from the 2002 *SSI Annual Statistical Report*, which is based on 1999 SSA administrative records matched to wave 12 of the 1996 *SIPP*.

[8] Unfortunately the 1990 SIPP does not allow us to determine whether a child or her parent/guardian is the SSI recipient.

almost four times more children in households with AFDC as with SSI, by 2001 the numbers were approximately equal.

III.A. Previous Research

Previous research has presented empirical evidence on the substitute nature of AFDC/TANF and SSI. Given the difference in funding sources, states have always had an incentive to move needy individuals from welfare to SSI.[9] By expanding the definition of disability for children, the ruling in the *Zebley* case increased the ability of states to move children from AFDC to SSI. Kubik (2003) presents evidence that in the five years following *Zebley*, states that experienced negative fiscal shocks in the early 1990s were more likely to encourage the move of children from the AFDC program to SSI.

Schmidt and Sevak (*forthcoming*) focus on variation across states in welfare reform policies, which created heterogeneity in the incentives facing individuals to switch from welfare to SSI. As pointed out by the authors, as AFDC became more restrictive, the SSI program's lack of time limits and work requirements became increasingly attractive. Their paper provides an empirical test of the hypothesis that welfare reform led to increased shifting from AFDC to SSI. The authors use March CPS data from 1988 through 1997 to examine the probability that a woman or her children receive SSI. The explanatory variable of interest is the interaction of an indicator variable for being a female-headed household and an indicator variable for living in a state with an approved welfare waiver. The estimated coefficient on this variable implies that female-headed households in states pursuing welfare reform are 21.6 percent more likely to participate in the SSI program. The authors suggest that their findings have implications for the

[9] The change in funding structure from AFDC to TANF increased the incentive of states to move individuals from welfare to SSI, but as described above, the PRWORA legislation tightened child eligibility for SSI and thereby made it more difficult to enroll children in the SSI program.

well-being of families affected by welfare reform time limits, though they do not explicit address outcomes other than program participation in their paper.

Garrett and Glied's (2000) examination of state-level SSI and AFDC caseloads focuses on variation across states in AFDC benefit levels. The authors use state-level data from 1987 to 1994, excluding 1990 and 1991, to estimate the effect of the *Zebley* ruling on SSI child participation and how that effect varies with AFDC benefit level. Using Ordinary Least Squares (OLS) techniques and controlling for state fixed effects, the authors estimate the effect of the AFDC benefit amount and SSI state supplementation amount, interacted with an indicator variable indicating a year post-*Zebley.* They find that the positive impact of *Zebley* on child SSI participation is more pronounced in states with lower AFDC payments and higher state SSI supplementation payments.

Similar evidence is provided by Kubik (1999). Using data from several years of the National Health Interview Survey (NHIS), he finds that after the 1990 liberalization of child SSI, the likelihood that an AFDC-eligible family identifies their child as suffering from a health impairment increases in the amount of extra SSI benefits they could receive by moving a child from AFDC onto SSI. The most commonly reported medical condition for these children is mental illness. Using March CPS data from 1987-1989 and 1991-1994, he finds that the interaction of net SSI benefit and a low education indicator is positively related to the likelihood that a family receives any SSI payment in the years after the *Zebley* decision, but not before. This interaction is also negatively related to the likelihood that the head of the family works, but there is no clear evidence of a differential impact pre- and post-*Zebley*.

III.B. New results on the shift from welfare to SSI

We present additional evidence regarding the relationship between AFDC and child SSI caseloads in the 1990s. Our major empirical contribution, presented in Sections IV and V below, is an analysis of poverty and labor supply outcomes that exploits variation in gender composition to identify a causal effect of SSI child participation on family outcomes. Here we build on the previous literature regarding the nature of caseload shifting in three ways: (1) we examine state-level SSI application and award data, in addition to state-level caseload data; (2) we examine county-level SSI and AFDC caseload data, which allows us to exploit within-state variation; and (3) we investigate how county population demographics determine county-level growth in SSI child caseloads. These analyses yield new insights into the nature of the shift from welfare to SSI.

Table 4 makes explicit the negative relationship between the fraction of children on welfare and on SSI after the 1990 liberalization of SSI. A negative relationship would result from families moving an eligible child from one program to the other. Countering this negative relationship is the positive correlation that would result from economic shocks causing income eligibility to change for both programs. Column (1) reports the results of an OLS regression of the change in the fraction of children on SSI on the change in the fraction of children on welfare in the years 1986 to 2001, controlling for year effects. The coefficient estimate is -0.0175 with a robust standard error of 0.0052. Column (2) reports that when state effects are controlled for the in the regression the coefficient estimate changes only slightly to -0.0159. Column (3) reports the results of allowing the relationship to vary by three periods: the pre-*Zebley* period (1985-1989), the expansion period (1990-1996), and the post-PRWORA period (1997-2001). There is no significant relationship between the change in the fraction of children on welfare and SSI in either the pre-*Zebley* or post-PRWORA period. During the expansion years, a one-percentage point reduction in the fraction of kids on AFDC is associated with an increase of 0.06 percentage points in the fraction of kids on SSI. Compared to the baseline fraction of 0.4% in 1989, this represents a 15 percent increase. A five-percentage point reduction in the fraction of kids on AFDC would increase the fraction of children on SSI by 75% in the average state.

We corroborate these results using data obtained from the Social Security Administration on SSI child awards and applications. Here we use the fraction of children applying for or awarded benefits as the dependent variable given that both are flow measures. As shown in Columns (4)–(9) of Table 4 the empirical results are qualitatively similar to those obtained using enrollment data. There is an insignificant relationship between changes in the fraction of children on welfare and the fraction of children awarded or applying for SSI in the periods before and after the SSI expansion. This presumably reflects the offsetting substitute and complementary nature of the two programs. However, during the expansion period, substitution dominates, with a one-percentage point decrease in the fraction of children on welfare associated with increases of 0.05% and 0.08% in the percentage of children awarded and applying for SSI, respectively. Given pre-*Zebley* means for these variables of just 0.08% and 0.20%, this would represent a 62% increase in awards and a 40% increase in applications above their baseline rates.

We next investigate the baseline determinants of the 1989 to 1999 change in the fraction of children on SSI, again using state-level data. After *Zebley*, we should see families in low AFDC benefit states applying for SSI at a higher rate than families in high AFDC benefit states. This follows from the heterogeneity in potential financial gain facing an AFDC-eligible individual. Figure 3 shows the pronounced difference in child application rates in Mississippi and Alabama relative to California and Alaska. In the four years following the *Zebley* ruling, the child application rate increased by 1.79 percentage points and 1.38 percentage points, respectively, in the two low-benefit states versus just 0.23 percentage points and 0.20 percentage points in the high benefit states. While 1989 application rates were higher in Mississippi and Alabama, the increases of 334 percent and 398 percent in application rates there were much larger than the corresponding increases of 175 percent and 159 percent in California and Alaska.

One caveat to this prediction is that AFDC benefit level is significantly and positively associated with the fraction of children on AFDC. We might expect that children already enrolled in AFDC would be more likely to apply for SSI than other children, as they are acquainted with public assistance programs and have access to social workers and government officials who could provide information about the SSI program. Thus one would expect that, after controlling for a state's AFDC benefit level, the fraction of children applying for or enrolling in the SSI program during the 1990s would be an increasing function of the fraction on AFDC at the time of the *Zebley* ruling.

The Ordinary Least Squares results listed in Table 5 confirm these predictions. Column (1) reports that a state's fraction of children on AFDC in 1989 is a positive and significant predictor of the change in the fraction of children on SSI – the coefficient estimate is 0.025 with a standard error of 0.013. Column (2) shows that the growth in SSI enrollment among children is negatively related to a state's baseline AFDC benefit, as predicted. Column (3) reports the results of including both variables in the regression, which substantially increases the magnitude of the two estimates. This suggests that because of the positive relationship between AFDC benefit generosity and AFDC enrollment including one of the variables without the other will bias the estimates toward zero. Coefficient estimates from this third specification indicate that a five percentage point increase in the fraction of children on AFDC just prior to the *Zebley* decision is associated with a 0.3 percentage point, or 75 percent, increase in the fraction of children newly eligible for SSI. The coefficient estimate on the maximum AFDC benefit (scaled by .01) for a family of three suggests that an additional $100 in AFDC benefits is associated with a 0.14 percentage point smaller increase from 1989 to 1999 in the fraction of children enrolled in SSI. In column (4) we include as an explanatory variable the initial fraction of children enrolled in SSI. The inclusion of this variable substantially reduces the two coefficient estimates of interest, suggesting that the very factors influencing SSI receipt after the *Zebley* decision had an effect prior to *Zebley* as well. This is not surprising given the substantial differences pre-*Zeble*y in application rates shown in Figure 3.

Specifications (5) through (12) report the results from estimating analogous specifications for the average fraction of children awarded SSI and the average fraction of children who applied for SSI in each of the ten years from 1990 to 1999. The results here are quite similar to those in the first four specifications, suggesting that applications and awards were significantly higher in states with low AFDC benefits and with a large fraction of children receiving AFDC just prior to the *Zebley* ruling. As with the change in enrollment, the inclusion of the 1989 award and application rates reduce the estimates for the AFDC benefit and the fraction of children on AFDC, though all four remain significant at the one percent level.

Figure 4 provides additional evidence of the extent to which the effect of the *Zebley* ruling varied across states. In 1989, the ratio of SSI child expenditures to total AFDC spending in the U.S. was just 0.05. Twelve years later, this ratio had increased by more than a factor of ten to 0.54. The magnitude of this change differed substantially between high and low benefit states. For example, in Alabama the ratio increased from 0.33 to 5.04 while in California the ratio grew from 0.05 to just 0.16.

In the results summarized in Table 6 we use county-level data on AFDC and SSI enrollment to investigate further the determinants of SSI growth among children. Because we have AFDC data for just half of all counties, the number of observations in specifications that include this variable is 1,541. The coefficient estimates in column (1) are almost identical to the analogous state-level estimates. In column (2) we add state fixed effects to the first specification - thus dropping the state-level measure of AFDC generosity. The coefficient estimate for the fraction of children on AFDC is unchanged. In the third column we control for the fraction of children in the county who were poor in 1989. This variable is significant and reduces the estimate for the AFDC enrollment rate variable by almost 40 percentage points. This estimate declines further once we control for the fraction of children living with just one parent, which is itself strongly positively related with the fraction of children on welfare. In the subsequent specifications we drop the AFDC enrollment variable so that we can include all counties in our sample. Our findings still suggest that counties with a large fraction of poor children and with relatively many children living in female-headed households experienced much greater increases in SSI receipt. The final specification shows that the growth in SSI was significantly lower in poor counties that had relatively many Hispanic children.

IV. The Effect of Child Participation in SSI on Family Poverty

In this section, we consider the impact that enrolling a child in SSI has on the likelihood that his family lives in poverty. The effect is theoretically ambiguous. Enrolling a child in SSI could increase or decrease total transfer income, depending on whether the family previously received income from AFDC (or TANF) and if so, the potential net financial change in moving a child to SSI. Enrolling a child in SSI could also lead to either an increase or decrease in earned family income. As shown in Table 2, data from the 2001 *SIPP* reveals that children in households with some SSI income are 48 percent less likely than children in households with TANF income to be living in poverty and are 73 percent less likely to be living below 50 percent of the poverty threshold.[10] These differences are consistent with a positive impact of SSI on household resources, but certainly cannot be interpreted as causal.

IV.A. Predicted effects

Let us denote total family income as $Y_j = \bar{Y} + w_j L_j$, where w_j is net-of-tax wage, L_j is hours of work, and $\bar{Y}$ is unearned income, which for simplicity we assume consists only of transfer income. The change in total family income experienced by a family who moves one child from AFDC to SSI can thus be represented as $\Delta Y_j = (\bar{Y}^S - \bar{Y}^A) + (w_j^S L_j^S - w_j^A L_j^A)$, where the superscript *S* refers to conditions under SSI and superscript *A* refers to conditions under AFDC. In this expression, $\bar{Y}^S - \bar{Y}^A$ represents the difference in transfer income a family receives if they move a child from AFDC to SSI. If the family consists of more than one child, $\bar{Y}^S$ might include some AFDC income. This difference is a function of the difference in SSI and AFDC benefit amounts as well as the difference in the AFDC payment to a family of size *N* and a family of size *N-1* (because the SSI-eligible child is removed from the AFDC unit). If a family enrolls a child on SSI who was not previously receiving public assistance income, then the change in family income is $\Delta Y_j = (\bar{Y}^S - 0) + (w_j^S L_j^S - w_j L_j)$.

There are two reasons why families in low AFDC benefit states who enroll a child in SSI are more likely to see increases in transfer income than families in high AFDC benefit states. First, $\bar{Y}^S - \bar{Y}^A$ will be more positive in almost all cases. And second, families in these states are more likely to be receiving no transfer income initially. For example, in 1990 the ratio of kids on

[10] This does not differentiate between adult and child SSI receipt. According to SIPP data matched to SSA administrative records, households with an adult on SSI are more than twice as likely as households with a child on SSI to be below the poverty line (44% vs. 19%). Thus SSI children are even less likely than TANF children to be in poverty than our comparison suggests. While the 2001 *SIPP* does differentiate between adult and child SSI receipt, it appears to overstate the former and understate the latter. Thus for the comparison in Table 2 we only consider whether any person in the household is receiving SSI.

AFDC to number of kids in poverty was 0.93 in California and 0.68 in Alaska compared to 0.37 in Alabama and 0.52 in Mississippi (Overview of Entitlement Programs, 1998). If there is no offsetting change in earned income, then the increase in family income and corresponding reduction in poverty will be greater for families living in low AFDC benefit states.

IV.B. Data

To investigate the effect of SSI receipt on family poverty, we would ideally use individual-level census data. Unfortunately, data limitations preclude the use of IPUMS data for both the first stage analysis and an analysis of family poverty. Census data for the year 2000 does not have information on SSI receipt for children under the age of 15. Thus a family with a child on SSI but no adults on SSI would be reported as having zero SSI income. Additionally, SSI payments to children would not be captured in family income or poverty measures in the 2000 census.[11] We therefore use data from the March CPS, also known as the Annual Demographic Supplement, to examine the effect of SSI receipt on poverty. The March CPS offers the advantage of having family-level data on poverty and SSI receipt that incorporates SSI payments to children.[12] However, if there are some individuals under the age of 15 in the household then it cannot be reliably determined from the CPS who in the family is the actual SSI recipient.

The CPS has a monthly sample of approximately 50,000 households. We combine data from the 1986-1990 and 1996-2000 files, referencing the years 1985-1989 and 1995-1999, in

[11] The 1990 Census does not separately identify SSI payments to any family members. These payments would instead show up in public assistance income or perhaps in social security income.
[12] In the CPS file, families and unrelated individuals are classified as being above or below the poverty level using a poverty index adopted by a Federal Interagency Committee in 1969 and slightly modified in 1981. The modified index provides a range of income cutoffs or "poverty thresholds" adjusted to take into account family size, number of children, and age of the family householder or unrelated individual. The poverty cutoffs are updated every year to reflect changes in the Consumer Price Index.

order to increase sample size and include years before and after the expansion of child SSI. We retain all children under the age of 18 except for those who are either the reference person or the spouse of the reference person, thus retaining individuals who live with both parents, the mother only, the father only, or with neither parent.

IV.C. Empirical Approach

The structural equation of interest for family *i* is the following:

$$Y_i = \alpha + \boldsymbol{\beta} \mathbf{X}_i + \gamma SSI_i + \varepsilon_{i,}$$

where Y is an indicator for whether the family lives in poverty, $\mathbf{X}$ is a vector of family demographics, and *SSI* is an indicator for whether the family receives SSI income based on a child's enrollment in the program. If properly estimated, γ captures the causal effect of child participation in SSI on the probability that a family lives in poverty. This effect is likely to vary across individuals though we consider here the baseline case of a common treatment effect.

There are three key obstacles to reliable identification of this parameter. First, SSI receipt is not exogenous to family income, which introduces a simultaneity problem. Estimating γ with OLS techniques would therefore yield an upward-biased coefficient estimate. Second, SSI is a federal program with very little variation across states in program rules. There is therefore no obvious control group. Third, the growth in child SSI enrollment occurred during the same period as welfare reform, the EITC expansion, the growth in the adult SSI program, and other factors that make it difficult to disentangle its effect from the effect of these other factors.

In an effort to surmount these and other obstacles to identification, we use an instrumental variable that influences SSI receipt but should not be otherwise related with changes in family poverty or labor supply during the 1990s. Our candidate instrument is an interaction of three variables: *momonly*numboys *post*. The variable *post* is equal to one in the years after the *Zebley* decision, 1996-2000 in our sample, and zero before. SSI receipt among children several years after the *Zebley* decision was three times greater than before. The variable *momonly* is equal to one if the child lives with the mother but the father is not present and zero otherwise. Children in female-headed families were almost five times more likely than other children to enroll in SSI during the 1990s. The variable *numboys* is equal to the number of boys in the family. From 1989 to 1999 there was an increase of 1.02 percentage points in the fraction of boys receiving SSI while for girls the increase was approximately half as large at 0.55 percent.[13]

While a number of other changes occurred during this period that would differentially affect female-headed families, we rely on the fact that SSI growth among children differentially affected families with relatively many boys whereas other policies did not. And although the greater prevalence of disabilities among boys may influence poverty or labor supply, this would have been true prior to the *Zebley* decision as well. As long as there are not differential changes in health among boys and girls from the late 1980s to the late 1990s, our strategy should provide a reliable estimate for the causal effect of SSI on family outcomes.[14]

The first-stage equation predicts the likelihood that family i in state j in year t receives any SSI income as a function of this triple interaction. The SSI indicator is set equal to a one if a family reports non-zero SSI income in the year and is equal to zero otherwise.[15] As discussed above, families in states with high AFDC benefits have less of a financial incentive to switch a child from AFDC to SSI. The predicted impact of *Zebley* on SSI participation is thus lower in

[13] In 1989 there were 4.72 boys per 1000 and 3.51 girls per 1000 on SSI. By 1999 these enrollment rates had increased to 14.93 per 1000 and 9.03 per 1000, respectively, with both having declined from the 1996 highs of 17.17 and 10.10.

[14] Our identification strategy makes the additional assumption that child SSI receipt does not influence whether the child lives only with his mother. We intend to investigate this issue in the next version of this paper.

[15] Unlike the one in the decennial Census the family SSI income variable in the CPS does include child SSI payments.

high AFDC benefit states. This additional source of variation is used to create a second instrument in the first-stage equation. The estimated equation is as follows:

$$\begin{aligned}
&AnySSI_{ijt} = \beta_o + \beta_1 post_t + \beta_2 momonly_i + \beta_3 numboys_i + \beta_4 numkids_i + \beta_5 (momonly * post)_{it} + \\
[1]\ &\beta_6 (numboys * post)_{it} + \beta_7 (numkids * post)_{it} + \beta_8 (momonly * numboys)_{it} + \\
&\beta_9 (momonly * numboys * post)_{it} + \beta_{10} (momonly * numboys * post * AFDCmax_{90})_{itj} + \theta_j + \varepsilon_{ijt}
\end{aligned}$$

The unit of observation in this regression is the family. All regressions are estimated with observations weighted by the sum of the individual CPS sample weights for each child in the family. In addition to the dichotomous variables *post, momonly, numboys,* and interactions thereof, the regression equation includes state fixed effects and controls for the total number of children in the family and an interaction of this variable with the *post* indicator. The predicted sign of $\hat{\beta}_9$ is positive and of $\hat{\beta}_{10}$ is negative.

The reduced-form relationship between family poverty status and the exogenous variables can be expressed as follows:

$$\begin{aligned}
&PovertyStatus_{ijt} = \gamma_o + \gamma_1 post_t + \gamma_2 momonly_i + \gamma_3 numboys_i + \gamma_4 numkids_i + \gamma_6 (momonly * post)_{it} + \\
[2]\ &\gamma_7 (numboys * post)_{it} + \gamma_8 (numkids * post)_{it} + \gamma_9 (momonly * numboys)_{it} + \\
&\gamma_{10} (momonly * numboys * post)_{it} + \gamma_{11} (momonly * numboys * post * AFDCmaxben90)_{itj} + \lambda_j + \varepsilon_{ijt}
\end{aligned}$$

The estimated coefficient on *momonly*numboys*post* has the interpretation of the triple-differences estimate of the impact of SSI receipt on family poverty status. The estimated coefficient on *momonly*numboys*post*AFDCmaxben90* captures the differential impact of SSI receipt on poverty for families living in states with relatively greater AFDC benefit levels, as proxied for by the 1990 maximum benefit for a family of three (scaled by .01). The predicted sign of $\hat{\gamma}_9$ is negative and of $\hat{\gamma}_{10}$ is positive.

The second-stage equation relating poverty status to predicted SSI receipt is as follows:

$$
\begin{aligned}
&PovertyStatus_{ijt} = \mu_o + \mu_1 \hat{A}nySSI + \mu_2 post_t + \mu_3 momonly_i + \mu_4 numboys_i + \mu_5 numkids_i + \\
[3]\quad &\mu_6(momonly * post)_{it} + \mu_7(numboys * post)_{it} + \mu_8(numkids * post)_{it} + \\
&\mu_9(momonly * numboys)_{it} + \rho_j + \varepsilon_{ijt}
\end{aligned}
$$

The excluded instruments used to predict *AnySSI* are *momonly*post*numboys* and the quadruple interaction m*omonly*post*numboys*AFDCmaxben90*. The identifying assumption underlying a causal interpretation of μ_1 is that families headed by single mothers with boys would have experienced the same change in outcomes during the 1990s as families headed by single mothers with girls, were it not for the differential participation in SSI.

IV.D. Empirical Results

Table 8 reports the results of estimating this first-stage equation. Column (1) lists the estimated coefficients when the equation is estimated on the full sample of families. The mean value of *AnySSI* in the full sample is 0.026, with an average of .018 in the pre-*Zebley* period and of .034 several years later. As shown, female-headed families are significantly more likely than other families to receive SSI income and this difference is significantly more pronounced in the post-*Zebley* years. In particular, $\hat{\beta}_8$ implies that an additional boy in the family increases the likelihood of a female-headed household receiving SSI income in the post-*Zebley* period by 0.0078 percentage points. This coefficient estimate is significantly different from zero, with a standard error of 0.0017. Given that .034 is the baseline probability of SSI receipt in female-headed families, this represents an increase of 22 percent in the probability that a family receives any SSI. While CPS data does not identify whether the SSI payment is awarded to a child or an adult in the house, the number of boys in a family should only have predictive value for child SSI receipt. We thus interpret this estimated difference is SSI receipt as being driven by differential participation in SSI among children.

Column (2) adds the maximum AFDC benefit for a family of three in 1990 to the regression. As discussed above, families in high AFDC benefit states have less financial incentive to switch a child from AFDC to SSI. The predicted impact of *Zebley* on SSI participation is thus lower in high AFDC benefit states. In this specification the estimated coefficient on *momonly*post*numboys* is 0.0196, with an associated standard error of 0.0024. This represents a 58 percent increase in the likelihood that a family headed by a single mother receives SSI income. The estimated coefficient on the interaction of *momonly*post*numboys*AFDCmaxben90* is -0.0029, with a standard error of 0.0004. This is consistent with our prediction that the differential SSI participation of families with relatively more boys is less pronounced in states with generous AFDC benefits.

Columns (3)-(5) report the results obtained from estimating equation [1] separately for families of one, two, and three or more children. This substantially lowers our sample size and thus our statistical precision, but the coefficient estimates are very similar to those obtained in the full sample. Having boys in a female-headed family appears to increase the participation response to the *Zebley* ruling for families consisting of one, two, or three or more children. This confirms that the interaction of *momonly*post*numboys* is capturing the differential impact of *Zebley* on families with larger numbers of boys, as opposed to a differential impact on female-headed families with greater numbers of children.

The odd-numbered columns in Table 9 report the results of estimating the reduced-form relationship described by equation [2]. The estimated main effects of the right-hand side variables are as expected. The coefficient on *post* suggests that the likelihood of a family living in poverty decreases over the nineties. The coefficient on *momonly* implies that female-headed families are 36 percentage points more likely than other families to live in poverty in the pre-

Zebley period with this difference declining to 29 percentage points several years after the *Zebley* decision. The coefficient of particular interest is the estimated coefficient on *momonly*numboys*post*, which is assumed to be correlated with poverty only to the extent that it is correlated with SSI participation. The estimated coefficient is -0.0134 (standard error of 0.0053). As predicted, the reduction in poverty associated with SSI participation is negatively related to the baseline AFDC benefit level. The estimated coefficient on *momonly*numboys*post*AFDCmaxben90* is 0.0016, with a standard error of 0.0009.

Two-stage least squares estimation of equation [3] relates predicted SSI participation at the family level to family poverty status. Recall that the motivation for using a 2SLS approach is that OLS will likely yield a biased estimate of the causal relationship between SSI participation and poverty status due to the endogeneity of family characteristics and SSI eligibility. Table 7 reports the results of OLS estimation of equation [3], where *AnySSI* is not predicted, but rather indicates whether the family does in fact receive any SSI payments. The estimated coefficient on *AnySSI*, reported in Column (1), implies that a family who receives SSI income is 20.2 percentage points more likely than a family who does not to live in poverty, controlling for female-headed status, the difference in poverty rates between time periods, number of children in the family, and the other variables listed in equation [3]. Columns (2) and (3) report that families with SSI receipt are 1.92 percentage points less likely to live below 50 percent of the poverty threshold and 22.4 percentage points more likely to live between 50 and 99 percent of the poverty threshold. Both coefficient estimates are statistically significant. Taken together, these estimates might be interpreted as suggesting that SSI participation raises families out of severe poverty, but not out of poverty entirely. Alternatively, they might suggest that SSI participation is targeted among those below the poverty line, but does not serve those in severe poverty.

The even-numbered columns in Table 9 report the results obtained from 2SLS estimation of equation [3]. The estimated coefficients differ noticeably from those obtained from OLS. In particular, the results of this analysis suggest that SSI participation is associated with significant reductions in poverty. The effect on the likelihood of a family living in poverty is estimated to be a reduction of 64.4 percentage points, with a standard error of 27.8. The estimated effect of SSI participation on the likelihood of a family living below 50 percent of the poverty threshold is a 48.7 percentage point reduction, with a standard error of 19.8. The estimated effect on the probability of living between 50 and 99 percent of the poverty threshold is not statistically significant: the estimated coefficient is -0.158 with a standard error of 0.229. The *F*-statistic of the first-stage regression is 247.25, suggesting that the instruments are strongly correlated with the predicted endogenous regressor. These results suggest that child participation in the SSI program reduces family poverty. Furthermore, this reduction in poverty appears to be driven primarily by a reduction in severe poverty.

V. The Effect of Child Participation in SSI on Maternal Labor Supply

The effect that enrolling a child on SSI will have on family resources is determined both by the change in transfer income received by the family and by any change in earnings caused by a behavioral labor supply response. The empirical analyses presented in Section IV offer evidence that family poverty declines as a result of child SSI enrollment. This suggests that the increase in transfer income experienced by a family that enrolls a child on SSI is not completely offset by a reduction in earnings. In other words, there is a net gain in family resources. In this section, we empirically address the question of how maternal labor supply responds to SSI participation. We focus on single mothers because SSI receipt is approximately five times greater among children living with a single mother than among all other children.

V.A. Predicted effects

The effect of SSI participation on labor supply is theoretically ambiguous. If a family moves a child from welfare to SSI, there is most likely an increase in transfer income.[16] Any increase in unearned income could lead to an income effect that would tend to reduce labor supply. If the child is the only child in the family, then earned income in the family is subject to the lower SSI benefit reduction rate (50 percent versus 67 or 100 percent under AFDC), which could lead to a positive substitution effect. For a family that was not previously on AFDC or TANF, both income and substitution effects are negative. However, given the rules regarding deeming of parental income in the 2000 calendar yaer, the SSI benefits for a child living with a single mother would not fall until his mother's monthly earnings exceed $1,111 if the child has no ineligible sibling, $1,367 if one ineligible sibling, and $1,623 if two ineligible siblings. In other words, the marginal tax rate that would lead to a negative substitution effect does not apply until relatively high levels of monthly earnings.

An additional complicating factor in that the labor supply elasticity of mothers with disabled children is likely to be different from this same parameter among all other women. For example while the traditional labor supply model predicts that a worker will consume some unearned income in the form of leisure (reduced labor), in this instance we might expect a mother to consume an increase in transfer income in the form of care for her disabled child, which would allow her to increase hours of labor

There is no clear prediction as to how the effect of enrolling a child on SSI will vary between high and low AFDC benefit states. The difference depends on the likelihood that a family was previously on AFDC (higher in high AFDC benefit states, so a more positive

[16] In some instances the family's amount of transfer income could decline, but given the relative generosity of SSI benefits and the fact that a family would only choose to switch if it were financially beneficial, we consider this to be an exceptional case.

substitution effect), the likelihood that a family still has a child on AFDC (higher in high AFDC benefit states, so a less positive substitution effect), the increase in transfer income (higher in low AFDC benefit states, so a more negative income effect), the sign and relative magnitudes of the substitution and income effects (ambiguous), and how these vary across states (ambiguous).

V.B. Data and Empirical Approach

We take advantage of the large sample size in the census IPUMS 5% micro sample to examine the effect of child SSI participation on maternal labor supply. We drop children who do not live with their mother – thus restricting attention to families with both parents and those headed by a single mother. We further restrict our sample to families with just one or two children because maternal labor supply varies substantially with the number of children and because – unlike with the CPS - we have a large sample size when using the IPUMS.

As noted above, census data for the year 2000 does not include information about SSI receipt for family members under the age of 15. IPUMS data therefore do not reliably identify families with a child SSI recipient. This precludes estimation of a 2SLS model using IPUMS data. Instead we estimate the reduced-form relationship:

$$\begin{aligned}[4]\quad & LaborSupply_{ijt} = \theta_o + \theta_1 twokids + \theta_2 post_t + \theta_3 momonly_i + \theta_4 numboys_i + \theta_5 (twokids * post) + \\ & \theta_6 (twokids * momonly) + \theta_7 (twokids * numboys) + \theta_8 (momonly * post) + \theta_9 (numboys * post) + \\ & \theta_{10} (momonly * numboys) + \theta_{11} (twokids * momonly * post) + \theta_{12} (twokids * numboys * post) + \\ & \theta_{13} (twokdids * momonly * numboys) + \theta_{14} (momonly * numboys * post)_{it} + \varepsilon_{ijt}\end{aligned}$$

The estimated coefficient on *momonly*numboys*post* in equation [4] has the interpretation of the triple-differences estimate of the impact of SSI receipt on maternal labor supply. The identifying assumption underlying a causal interpretation of θ_{14} is that families headed by single mothers with boys would have experienced the same change in labor supply during the 1990s as families headed by single mothers with girls, were it not for the differential participation in SSI. Although the greater prevalence of disabilities among boys may affect maternal labor supply, this would have been true prior to the *Zebley* decision as well, and is controlled for in the regression with the interaction of *momonly*numboys*.

As described above, the labor supply response to SSI enrollment theoretically depends on whether the family has another child participating in the AFDC or TANF program. For a family with one child on SSI and another on AFDC/TANF, the program would be expected to reduce labor supply because of the negative income effect. Alternatively if a family switched from welfare to SSI then the program's effect is ambiguous, with the income effect and substitution effect going in opposite directions.

V.C. Empirical Results

Table 10 reports the results of estimating equation [4] for four different measures of maternal labor supply: an indicator for being in the labor force, an indictor for working, weeks worked last year, and usual hours of work. The estimated coefficient on the triple interaction of interest implies that an additional boy in the family increases the likelihood that a single mother works post-*Zebley* by 0.39 percentage points, with a standard error of 0.24. The estimated effects on the other three labor supply measures are all statistically significant. As shown in column (2), the coefficient estimate in the equation describing probability of working is 0.0058, with a standard error of 0.0026. As shown in column (3), the coefficient estimate in the equation describing weeks worked last year is 0.241, with a standard error of 0.118. And as shown in column (4), the coefficient estimate in the equation describing usual hours worked is 0.224, with a standard error of 0.099.[17]

[17] We estimated each of these four regressions separately for families with one and two children. All eight coefficient estimates are positive, though just three of them are statistically significant. We therefore group the one and two child families to increase our precision.

We can use these reduced-form coefficient estimates to construct Wald Estimators of the effect of SSI participation on maternal labor supply. Using our CPS data, we estimate a regression that is identical to the ones summarized in Table 10 for any SSI receipt and obtain an estimate of 0.0056 for the *momonly*numboys*post* coefficient (with a standard error of .0025). Dividing each of the estimates from Table 10 by .0056, our point estimates suggest a substantial effect on maternal labor supply. For example, for each 100 women with children enrolling in SSI an additional 70 enter the labor force and an additional 100 are working.

While our estimates are compatible with a wide range of effects on maternal labor supply, taken together they strongly suggest that the welfare to disability shift has increased employment and labor force attachment among single mothers.

VI. Discussion

During the 1990s there were several important changes in tax and expenditure programs that affected the material well-being and labor supply of single-mothers and their children. One important but understudied change was the expansion of SSI eligibility among low-income children resulting from a 1990 Supreme Court decision. This paper has presented new evidence about the rise in child participation in SSI. It has also presented empirical evidence about the impact this had on poverty and maternal labor supply. In order to identify a causal relationship between program participation and family outcomes, we exploit the fact that boys were 85 percent more likely than girls to enroll in SSI during this expansion period. We focus on children living only with their mother because of their substantially greater rate of SSI receipt.

The results of our empirical analysis suggest that the rise in SSI participation among children led to statistically significant declines in family poverty and statistically significant increases in maternal labor supply. Given that SSI receipt increased by approximately three percentage points among female-headed families, our estimated point estimates imply that this expansion can explain a two percentage point drop in child poverty and a two percentage point increase in maternal labor force participation among these families. It therefore appears that the expansion of SSI benefits for children made an important contribution to the decline in child poverty and the increase in labor supply among single mothers that occurred during the 1990s.

Future work that examines the long-term effect of the shift from welfare to SSI is clearly warranted. This is especially true given that average durations on SSI are substantially longer than those on AFDC/TANF. It would also be interesting to extend the analysis in this paper to other outcome variables including family structure, consumption, educational outcomes, and health.

Our findings also underscore the extent to which the effect of this policy change varied across states. Disadvantaged children from low AFDC benefit states derived a much greater benefit from the expansion of SSI than children in states with more generous welfare benefits. Thus while low-income children in states like Alabama and Mississippi once received transfer income that was several times lower than their counterparts in California and New York, this disparity is now much less pronounced. In addition, as our county-level analyses demonstrate, low-income Hispanic children were significantly less likely than other low-income children to enroll in the program during the expansion years. Thus, while Louisiana and Texas had equally generous AFDC benefits and similar child poverty rates at the time of the *Zebley* decision, there were 4.5 times more applicants in the former than in the latter during the 1990s. In future work we intend to examine the causes and consequences of this low take-up among Hispanic children.

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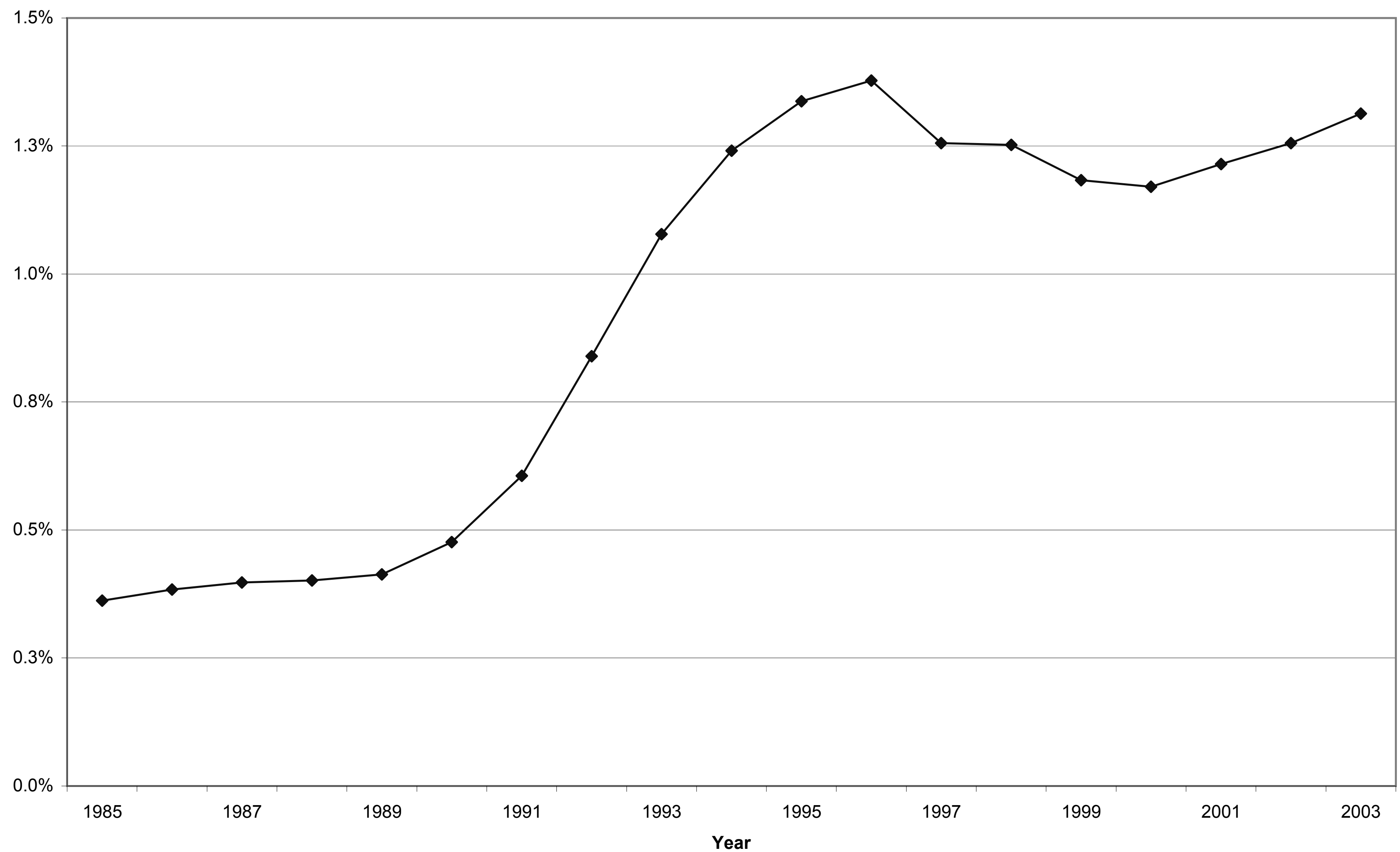
Figure 1A: Percentage of Children on SSI from 1985-2003



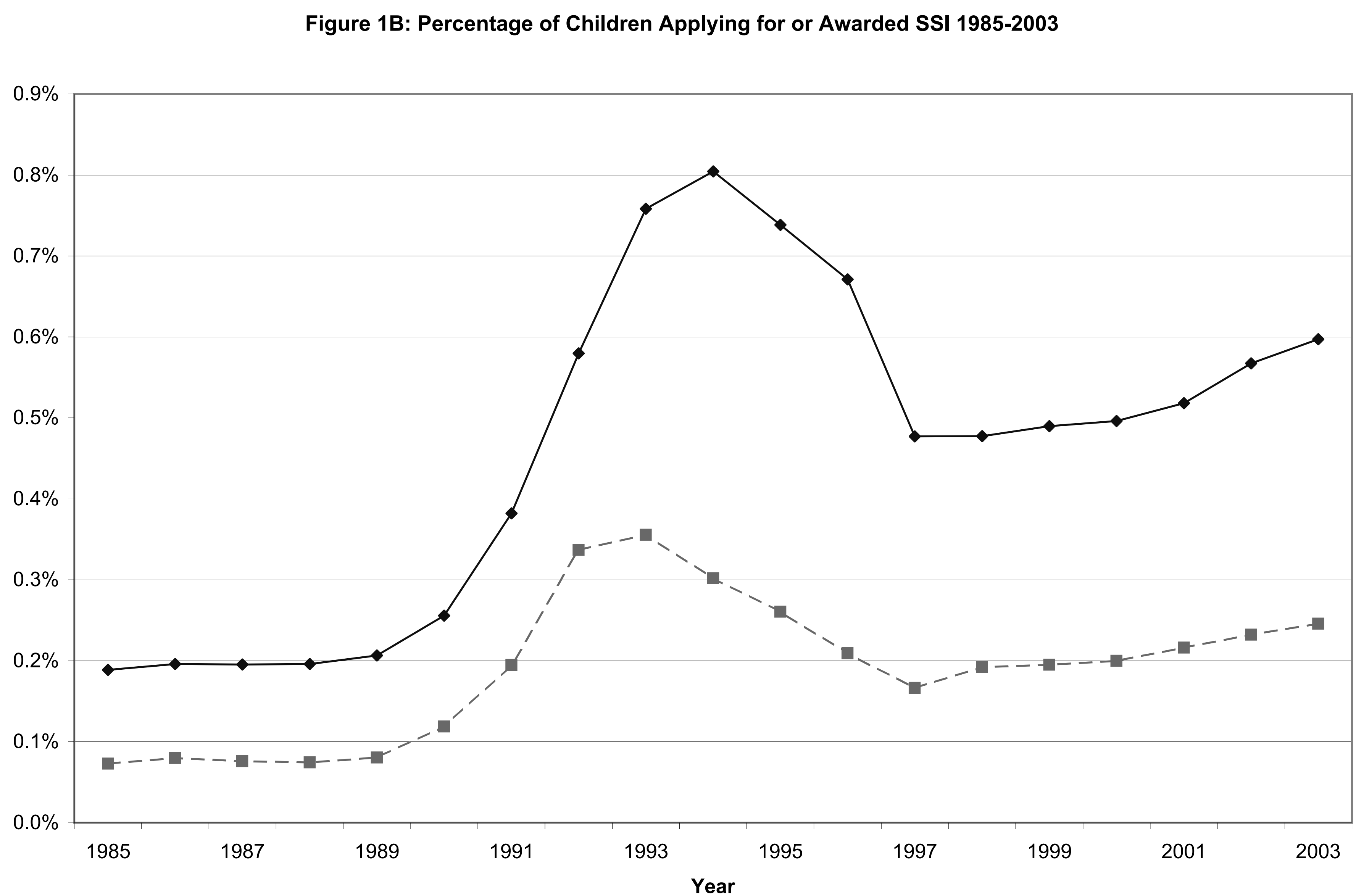


Figure 1B: Percentage of Children Applying for or Awarded SSI 1985-2003
0.9%
0.8%
0.7%
0.6%
0.5%
0.4%
0.3%
0.2%
0.1%
0.0%
1985
1987
1989
1991
1993
1995
1997
1999
2001
2003
Year

Figure 1C: Percent of SSI Children with Benefits Suspended: 1988-2002

Figure 2: Ratio of AFDC Families to SSI Children 1985-2002

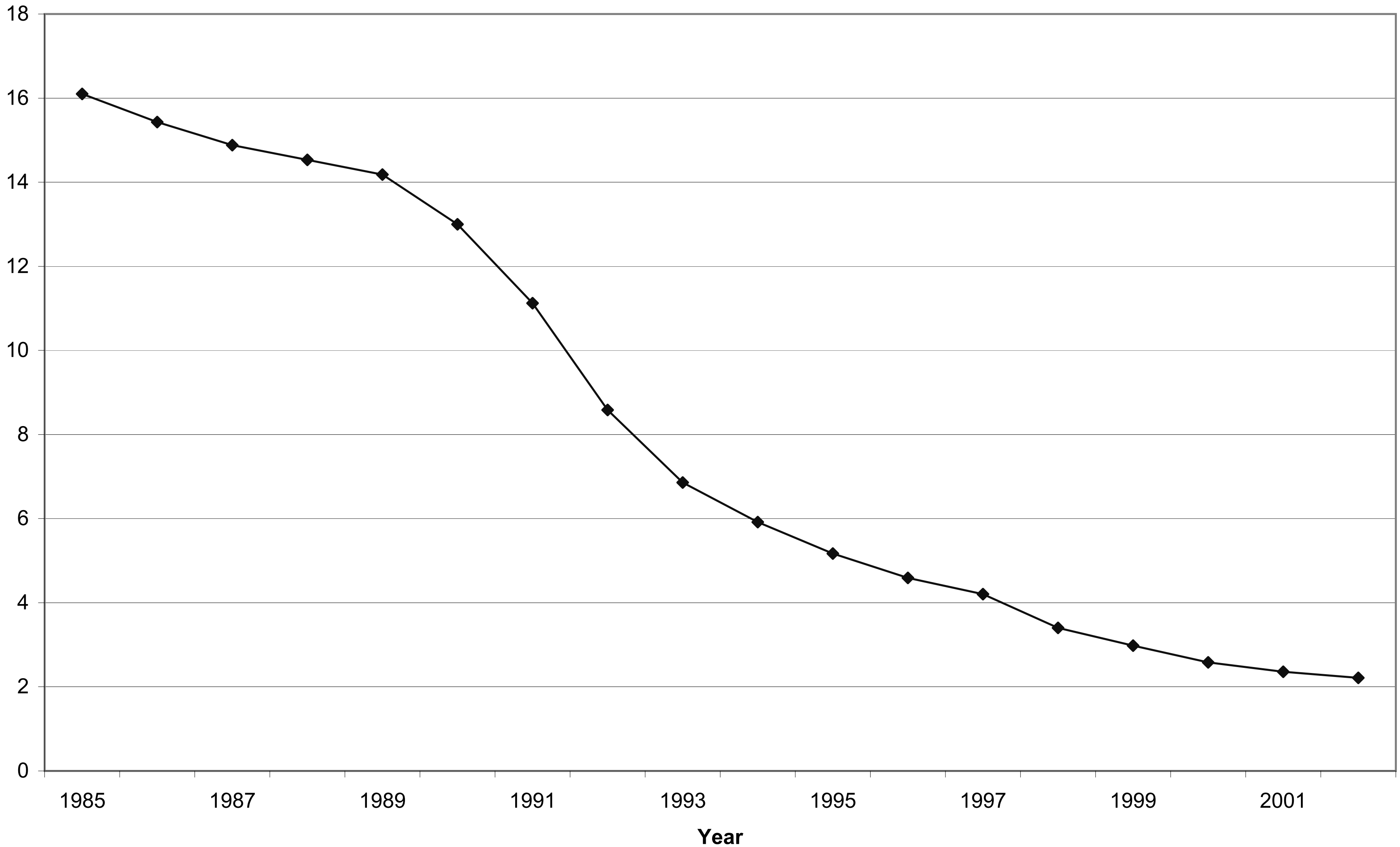


Figure 3: SSI Child Application Rates in Low / High AFDC Benefit States

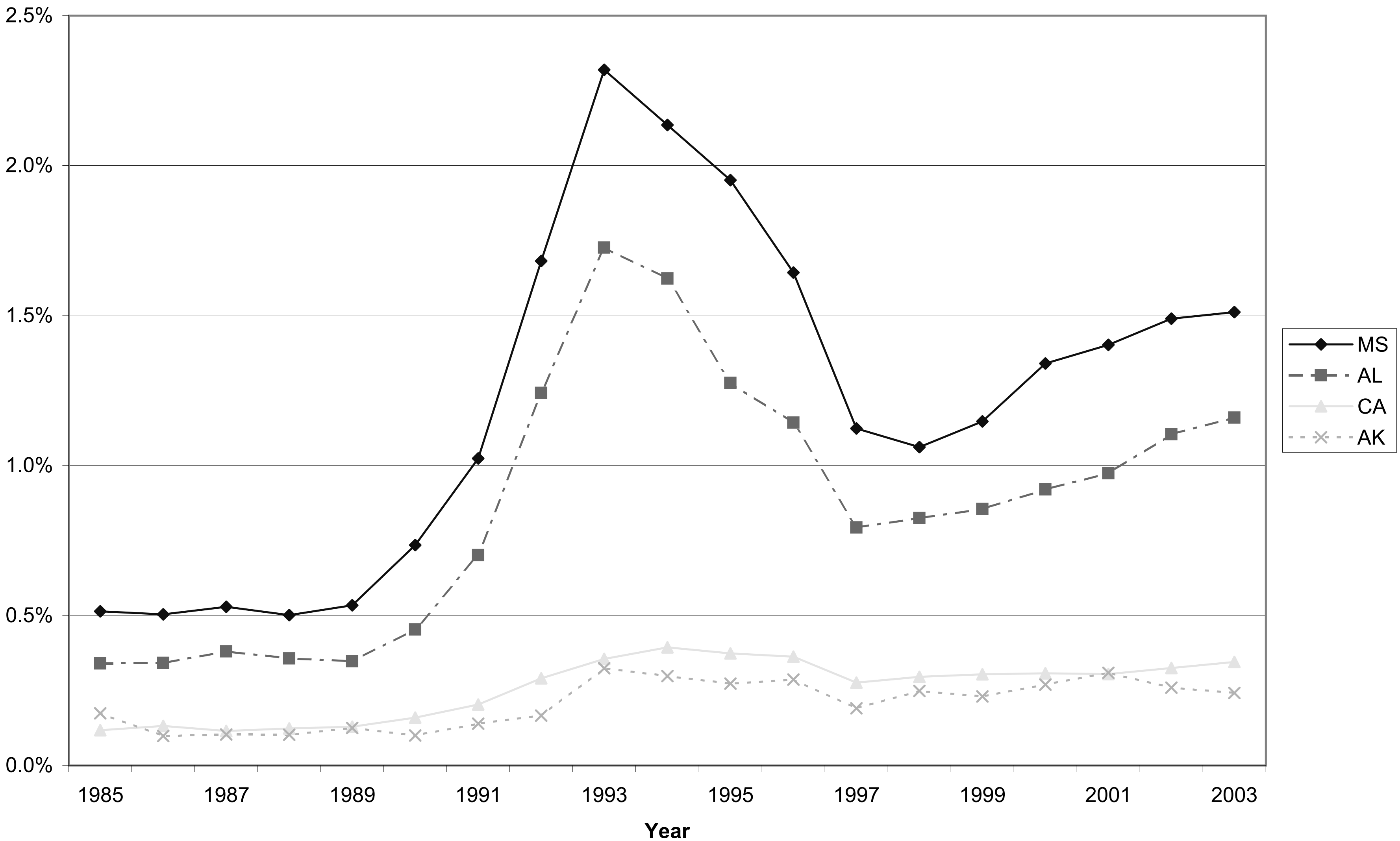


Figure 4: Ratio of SSI Kid $ to AFDC Dollars in Low vs. High Benefit States 1989 vs. 2001

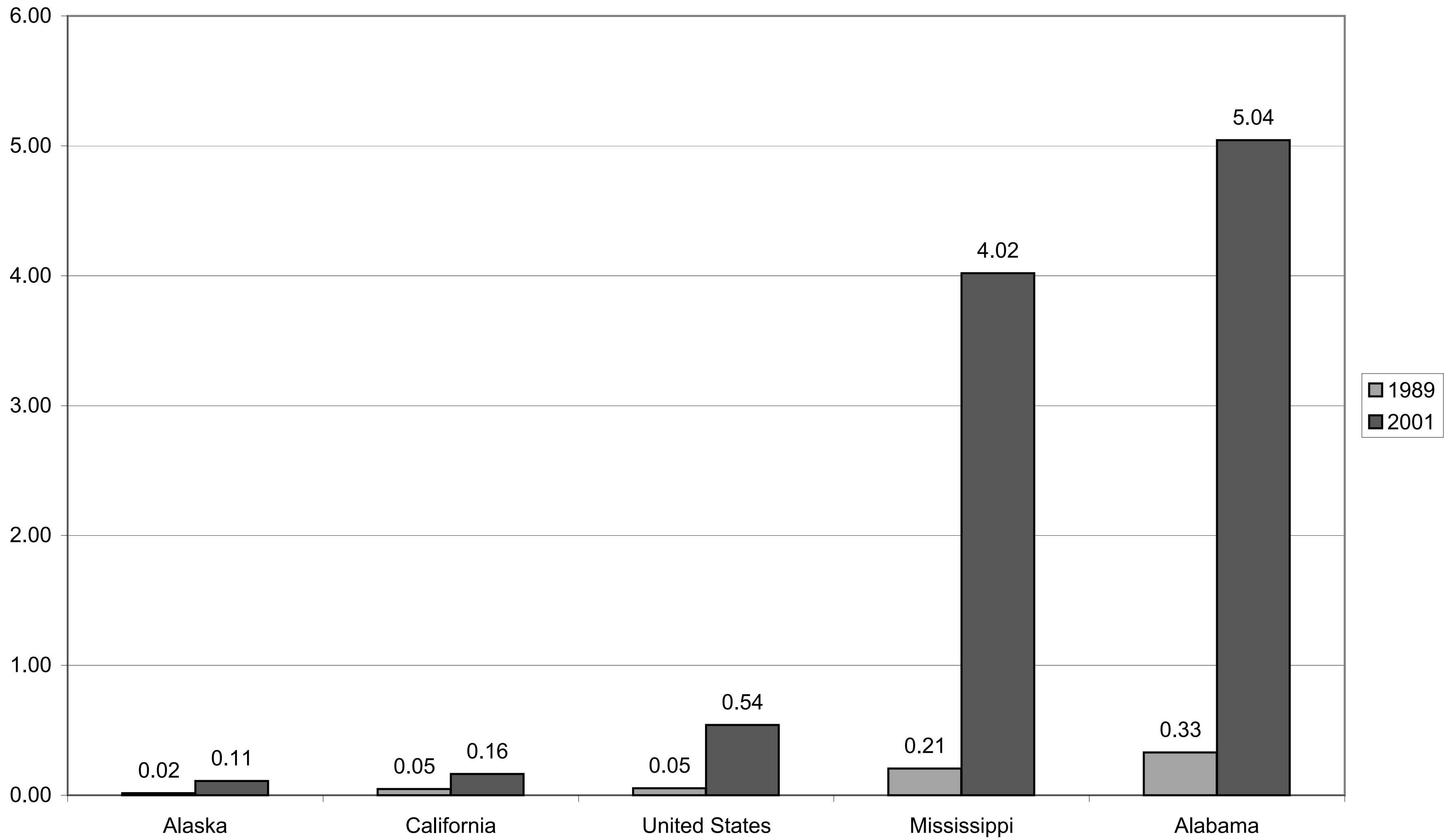


Figure 5: Labor Supply Incentives for Single Mother with Child SSI vs. AFDC

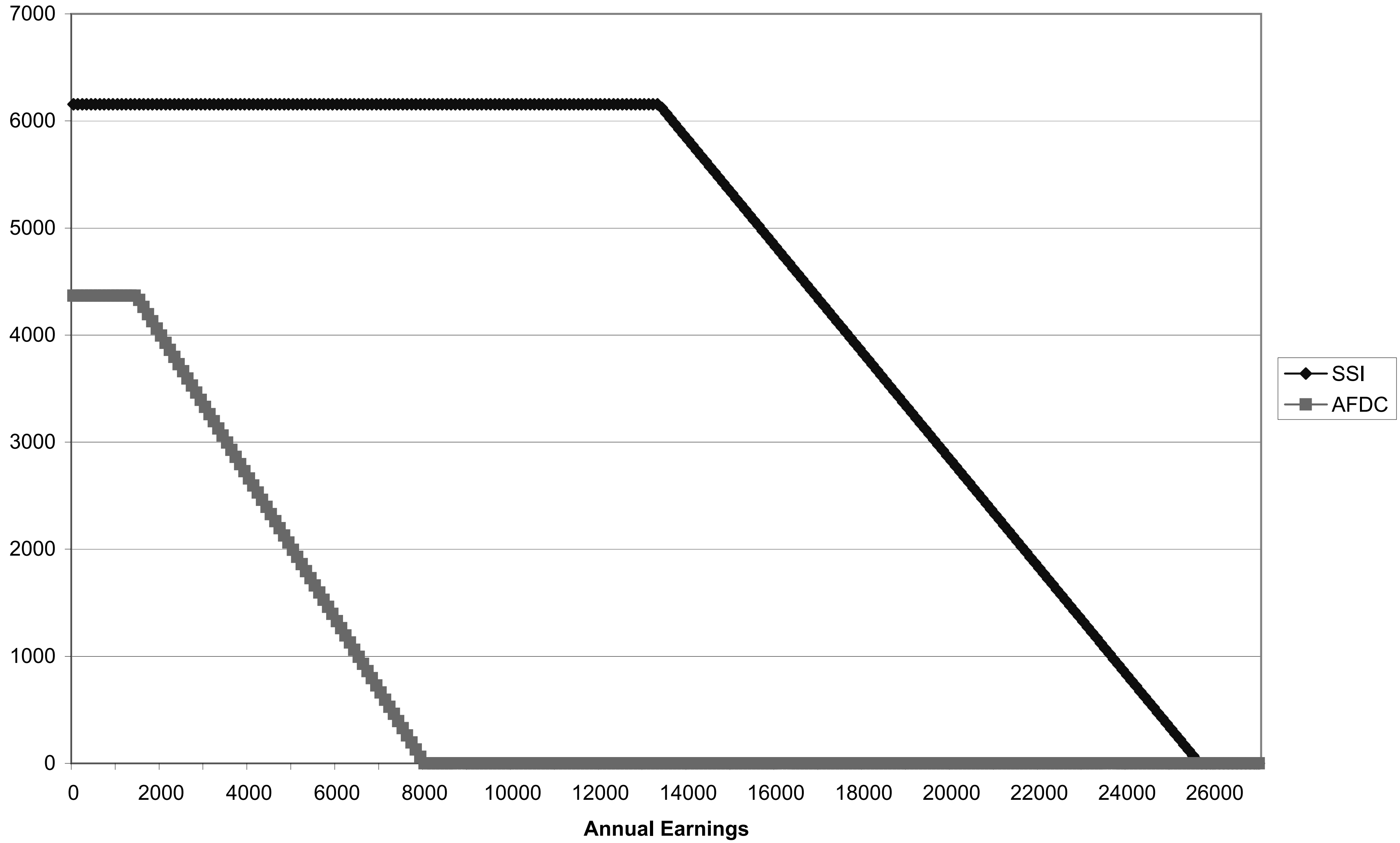


Table 1: Primary Diagnosis of Kids on SSI in 12/89 and 12/99

	Boys89	Boys99	Girls89	Girls99	Total89	Total99
Mental Retardation	66,993	179,454	43,062	113,650	110,055	293,104
	43.2%	33.4%	27.8%	36.7%	41.5%	34.6%
Other mental disorder	11,146	160,879	5,264	53,469	16,410	214,347
	7.2%	29.9%	3.4%	17.3%	6.2%	25.3%
Nervous system	37,573	57,575	30,164	45,211	67,737	102,786
	24.2%	10.7%	19.5%	14.6%	25.6%	12.1%
Congenital anomalies	12,981	23,974	11,464	21,340	24,445	45,314
	8.4%	4.5%	7.4%	6.9%	9.2%	5.3%
All other	26,167	115,484	19,953	76,028	46,120	191,512
	16.9%	21.5%	12.9%	24.5%	17.4%	22.6%
Total	154,983	537,366	109,907	309,697	264,890	847,063
# on SSI per 1000 Boys or Girls	4.72	14.93	3.51	9.03	4.13	12.05

Table 2: Children with SSI or AFDC/TANF Income in Household: Data from the 1990 and 2001 SIPP

	Any AFDC	Any SSI	All Other HH	Any TANF	Any SSI	All Other HH
# Households in SIPP Sample (unweighted)	808	221	7661	549	617	12009
# Kids in these sampled HH (unweighted)	1971	483	14171	1319	1306	22596
% of Households	8.5%	2.5%	90.0%	3.9%	4.4%	92.6%
% of Kids in these Households	10.2%	2.8%	88.3%	4.8%	4.7%	91.5%
Avg # Persons in HH	4.08	4.64	3.92	4.14	4.37	3.92
Avg # children age 0-17 in HH	2.37	2.16	1.84	2.37	2.06	1.86
Avg # persons age 18-64 in HH	1.65	2.23	2.06	1.70	2.13	2.02
Avg # persons age 65+ in HH	0.06	0.24	0.03	0.07	0.18	0.04
% HH with AFDC/TANF	100.0%	39.2%	0.0%	100.0%	19.6%	0.0%
% HH with SSI	11.5%	100.0%	0.0%	22.3%	100.0%	0.0%
% Female-Headed Household	74.6%	48.4%	17.1%	71.9%	52.3%	22.6%
% Kids in HH headed by Female	70.5%	46.2%	14.5%	68.7%	50.5%	20.4%
% Kids Black	45.0%	39.5%	12.0%	41.2%	38.3%	14.1%
% Kids White	47.5%	50.4%	84.3%	53.3%	52.9%	80.5%
% Kids Hispanic	18.7%	10.9%	10.2%	27.8%	19.5%	16.3%
Avg Monthly HH Income	$978	$1,833	$3,366	$1,438	$2,750	$5,179
Avg from AFDC/TANF	$378	$149	$0	$300	$52	$0
Avg from SSI	$44	$343	$0	$131	$554	$0
Avg from Social Security	$53	$176	$48	$84	$206	$67
Avg Earnings	$441	$1,024	$3,077	$820	$1,717	$4,854
Avg Other Income	$62	$141	$241	$103	$221	$258
% HH with Monthly Income 0-50% of Poverty	35.9%	6.0%	3.1%	33.7%	9.1%	4.1%
% HH with Monthly Income 50-100% of Poverty	40.6%	36.0%	6.3%	33.9%	26.0%	7.4%
% HH with Monthly Income 100-150% of Poverty	11.8%	23.9%	9.7%	14.9%	22.2%	9.9%
% HH with Monthly Income 150-200% of Poverty	3.8%	10.1%	11.7%	9.1%	16.1%	10.8%
% HH with Monthly Income > 200% of Poverty	7.9%	24.0%	69.3%	8.4%	26.6%	67.7%

notes: (1) HH counted in AFDC or SSI group must receive aid in one or more of the 4 months of wave
(2) Data weighted by sampling weight, unless otherwise noted

Table 3: Characteristics of SSI Child Recipients in December 1999

Demographic characteristics		
Male	588407	64.7%
Black	354785	39.0%
Hispanic	200161	22.0%
Under 5	137261	15.1%
Ages 5-9	264523	29.1%
Ages 10-14	326336	35.9%
Ages 15-17	180894	19.9%
Household type	Number	Percentage
Married couple present	353412	38.9%
No married couple, female householder	518804	57.1%
No married couple, male householder	36798	4.0%
Mother not present	113906	12.5%
Father not present	573206	63.1%
Grandchild of household head	114904	12.6%
Percentage Distribution of Family Income		
Earnings		40.8%
SSI		28.8%
Other public assistance		16.7%
Social Security		8.1%
Other		5.6%
Poverty status		
0-49% of poverty	25883	2.8%
50-99% of poverty	146631	16.1%
100-149% of poverty	291722	32.1%
150-199% of poverty	224286	24.7%
200-299% of poverty	186509	20.5%
300% or more of poverty	33983	3.7%
Household Size		
2 persons	76717	8.4%
3-4 persons	363680	40.0%
5+ persons	468617	51.6%
Family Size		
1 person	24037	2.6%
2 persons	89366	9.8%
3-4 persons	343404	37.8%
5+ persons	452208	49.7%
Household Living Quarters		
Owned	396400	43.6%
Public housing	83100	9.1%
Other not owned	429515	47.3%
Other Government Programs		
Food stamps	345127	38.0%
Other public assistance income	200272	22.0%
Housing assistance	133591	14.7%
Energy assistance	88780	9.8%

Source: 2002 SSI Annual Statistical Report

Table 4: The Relationship between SSI and Changes in AFDC Child Recipients

	Δ % Kids on SSI			% Kids Awarded SSI			% Kids Applying for SSI		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Δ % Kids on AFDC/TANF	-0.0175	-0.0159	0.0106	-0.0220	-0.0153	0.0027	-0.0490	-0.0268	-0.0055
	(.0033)	(.0033)	(.0107)	(.0043)	(.0029)	(.0094)	(.0121)	(.0065)	(.0212)
*Δ % Kids on AFDC/TANF * (Year90-96)*			**-0.0588**			**-0.0498**			**-0.0790**
			(.0124)			(.0108)			(.0244)
*Δ % Kids on AFDC/TANF * (Year97-01)*			-0.0145			-0.0051			0.0036
			(.0114)			(.0100)			(.0225)
Year Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
State Effects	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
# Observations	816	816	816	816	816	816	816	816	816
Years Included	1986-2001	1986-2001	1986-2001	1986-2001	1986-2001	1986-2001	1986-2001	1986-2001	1986-2001
R-squared	0.698	0.739	0.752	0.542	0.817	0.829	0.442	0.864	0.870

Note: Unit of observation is state*year. Regressions are weighted by share of kids in U.S. in state in each year.

Table 5: State-Level Determinants of the Growth in Child SSI Enrollment

	Δ % Kids on SSI 1989-99			
	(1)	(2)	(3)	(4)
% of Kids on AFDC in 1989	0.0248		0.0618	0.0224
	(.0134)		(.0123)	(.0134)
Max 1990 AFDC Benefit FamUnit=3		-0.00077	-0.00144	-0.00042
		(.00026)	(.00025)	(.00031)
% Kids on SSI in 1989				1.6115
				(.3504)
Constant	0.0052	0.0102	0.0051	0.0005
	(.0016)	(.0009)	(.0012)	(.00172)
# Observations	51	51	51	51
R-squared	0.066	0.148	0.441	0.615

	Avg % Kids Awarded SSI 1990-99			
	(5)	(6)	(7)	(8)
% of Kids on AFDC in 1989	0.0033		0.0140	0.0078
	(.0033)		(.0026)	(.0021)
Max 1990 AFDC Benefit FamUnit=3		-0.00026	-0.00042	-0.00017
		(.00006)	(.00005)	(.00005)
% Kids Awarded SSI in 1989				2.3953
				(.3414)
Constant	0.0020	0.0031	0.0019	0.00021
	(.0004)	(.0002)	(.0003)	(.00037)
# Observations	51	51	51	51
R-squared	0.020	0.300	0.558	0.784

	Avg % Kids Apply for SSI 1990-99			
	(9)	(10)	(11)	(12)
% of Kids on AFDC in 1989	0.0106		0.0442	0.0157
	(.0102)		(.0079)	(.0055)
Max 1990 AFDC Benefit FamUnit=3		-0.00083	-0.00131	-0.00042
		(.00018)	(.00016)	(.00013)
% Kids Applied SSI in 1989				2.0382
				(.2102)
Constant	0.0044	0.0079	0.0043	0.00141
	(.0012)	(.0006)	(.0008)	(.00066)
# Observations	51	51	51	51
R-squared	0.021	0.312	0.582	0.861

Table 6: County-Level Determinants of the Growth in Child SSI Receipt

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
State 1990 AFDC Benefit	-0.002							
	(.0001)							
Feb. 90 Fraction Kids on AFDC	0.054	0.054	0.035	0.010				
	(.006)	(.006)	(.011)	(.005)				
% Kids Poor in 89			0.024	0.016	0.020	0.018	0.029	0.032
			(.010)	(.004)	(.004)	(.002)	(.004)	(.005)
% Kids with Mom Only in 89				0.036	0.042	0.038	0.029	0.026
				(.004)	(.003)	(.002)	(.005)	(.004)
% Kids Black and Poor in 89							0.008	0.008
							(.005)	(.005)
% Kids Hisp and Poor in 89							-0.023	-0.024
							(.004)	(.004)
% Pop Rural in 89								-0.001
								(.001)
# Observations	1541	1541	1541	1541	1541	3121	3121	3121
R-squared	0.57	0.76	0.79	0.84	0.83	0.80	0.83	0.83
State Fixed Effects	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Unit of observation is the county. Dependent variable is equal to the change from 1989 to 1999 in the fraction of children in the county on SSI. Regressions are weighted by the number of children in the county. Robust standard errors are included in parentheses. The unit of observations increases from 1541 to 3121 when we drop the AFDC enrollment variable because we have this for just one-half of all counties. The weighted mean of the dependent variable is equal to .0101.

Table 7: OLS Estimates of Impact of SSI on Child Poverty

	(1)	(2)	(3)
	Poverty	0-49% Pov	50-99% Pov
Any SSI	0.2015	-0.0192	0.2207
	(.0048)	(.0035)	(.0042)
Post	-0.0117	-0.0040	-0.0076
	(.0035)	(.0026)	(.0030)
MomOnly	0.3613	0.1939	0.1674
	(.0034)	(.0025)	(.0029)
*MomOnly * Post*	-0.0857	-0.0596	-0.0262
	(.0036)	(.0027)	(.0032)
NumBoys	-0.0098	-0.0111	0.0012
	(.0014)	(.0011)	(.0012)
*NumBoys * Post*	-0.0099	-0.0041	-0.0059
	(.0020)	(.0014)	(.0017)
NumKids	0.0671	0.0351	0.0320
	(.0012)	(.0008)	(.0010)
*NumKids * Post*	0.0075	0.0031	0.0045
	(.0016)	(.0012)	(.0014)
*MomOnly * NumBoys*	0.0592	0.0571	0.0021
	(.0018)	(.0014)	(.0016)
# Observations	208200	208200	208200
R-squared	0.24	0.16	0.08

Sample includes all families with children under the age of 18 from the 1986-1990 and 1996-2000 March CPS. Unit of observation is the child though only one child per family is included. Regressions are weighted by the sum of person weights for each child in the family divided by the sum of these weights in each year (thus each year gets equal weight). Dependent variable equals one if child's family is in poverty, less than 50% of poverty, or 50-99% of poverty. Robust standard errors are included in parentheses and all regressions include

Table 8: The Impact of Gender Composition on SSI Receipt in Female-Headed Families

Dependent Var = I(Any SSI Income in Family)

	(1)	(2)	(3)	(4)	(5)
Post	-0.0035	-0.0036	0.0088	0.0090	-0.0318
	(.0017)	(.0017)	(.0018)	(.0020)	(.0069)
MomOnly	0.0210	0.0210	0.0129	0.0194	0.0281
	(.0018)	(.0018)	(.0024)	(.0031)	(.0060)
*MomOnly * Post*	0.0112	0.0109	0.0076	0.0101	0.0226
	(.0026)	(.0026)	(.0033)	(.0043)	(.0083)
NumBoys	0.0019	0.0019	0.0023	0.0005	0.0027
	(.0007)	(.0007)	(.0017)	(.0011)	(.0014)
*NumBoys * Post*	-0.0015	-0.0015	-0.0007	-0.0004	-0.0005
	(.0010)	(.0010)	(.0025)	(.0016)	(.0020)
NumKids	0.0032	0.0031			0.0065
	(.0005)	(.0005)			(.0014)
*NumKids * Post*	0.0071	0.0071			0.0139
	(.0008)	(.0007)			(.0020)
*MomOnly * NumBoys*	-0.0009	-0.0009	0.0010	0.0051	-0.0053
	(.0012)	(.0012)	(.0033)	(.0025)	(.0028)
*MomOnly * NumBoys * Post*	0.0078***	0.0196***	0.0146**	0.0112**	0.0178***
	(.0017)	(.0024)	(.0066)	(.0046)	(.0050)
*MomOnly * NumBoys * Post * AFDC Ben*		-0.0029***	-0.0007	-0.0020**	-0.0033***
		(.0004)	(.0012)	(.0008)	(.0008)
# Observations	208200	208200	90471	76005	41724
R-squared	0.0169	0.0171	0.0086	0.0142	0.0261
Number of Kids	All	All	1	2	3 or more

Sample includes all families with children under the age of 18 from the 1986-1990 and 1996-2000 March CPS. Unit of observation is the child though only one child per family is included. Regressions are weighted by the sum of person weights for each child in the family divided by the sum of these weights in each year (thus each year gets equal weight). Dependent variable equals one if family SSI income is greater than zero and is otherwise equal to zero. Robust standard errors are included in parentheses and all regressions include state fixed effects.

Table 9: IV Estimates of the Impact of SSI Receipt on Family Poverty

	Poverty		0-49% of Poverty		50-99% of Poverty	
	(1)	(2)	(3)	(4)	(5)	(6)
Any SSI		-0.6598		-0.4983		-0.1615
		(.2822)		(.2014)		(.2324)
Post	-0.0148	-0.0167	-0.0054	-0.0068	-0.0095	-0.0099
	(.0037)	(.0041)	(.0027)	(.0029)	(.0032)	(.0034)
MomOnly	0.3606	0.3754	0.1907	0.2017	0.1699	0.1736
	(.0040)	(.0059)	(.0029)	(.0042)	(.0035)	(.0048)
*MomOnly * Post*	-0.0738	-0.0684	-0.0542	-0.0499	-0.0196	-0.0184
	(.0056)	(.0069)	(.0041)	(.0049)	(.0049)	(.0057)
NumBoys	-0.0103	-0.0089	-0.0116	-0.0106	0.0013	0.0017
	(.0015)	(.0016)	(.0011)	(.0011)	(.0013)	(.0013)
*NumBoys * Post*	-0.0084	-0.0097	-0.0030	-0.0040	-0.0054	-0.0058
	(.0021)	(.0021)	(.0016)	(.0015)	(.0019)	(.0017)
NumKids	0.0677	0.0698	0.0350	0.0366	0.0327	0.0332
	(.0012)	(.0015)	(.0008)	(.0011)	(.0010)	(.0013)
*NumKids * Post*	0.0091	0.0137	0.0030	0.0065	0.0061	0.0072
	(.0016)	(.0027)	(.0012)	(.0019)	(.0014)	(.0022)
*MomOnly * NumBoys*	0.0632	0.0619	0.0595	0.0585	0.0037	0.0033
	(.0027)	(.0022)	(.0019)	(.0015)	(.0008)	(.0018)
*MomOnly * NumBoys * Post*	-0.0134**		-0.0101		-0.0033	
	(.0053)		(.0039)		(.0046)	
*MomOnly * NumBoys * Post * AFDC Ben*	0.0016*		0.0013		0.0004	
	(.0009)		(.0007)		(.0008)	
# Observations	208200	208200	208200	208200	208200	208200
R-squared	0.229	-	0.158	-	0.070	-

Sample includes all families with children under the age of 18 from the 1986-1990 and 1996-2000 March CPS. Unit of observation is the child though only one child per family is included. Regressions are weighted by the sum of person weights for each child in the family divided by the sum of these weights in each year (thus each year gets equal weight). Dependent variable in columns 1 and 2 is equal to one if family is in poverty and zero otherwise. The next two dependent variables differentiate between 0-49% and 50-99% of poverty line. Robust standard errors are included in parentheses and all regressions include state fixed effects.

Table 10: Reduced Form Estimate of Effect of Child SSI on Maternal Labor Supply

	In Lab Force	Working	Wks Last Yr	Usual Hours
Two Kids	-0.0412	-0.0391	-3.119	-3.183
	(.0016)	(.0017)	(.077)	(.0639)
Post	0.0003	0.0090	1.785	0.863
	(.0014)	(.0015)	(.068)	(.057)
*Two * Post*	0.0024	0.0013	0.485	0.331
	(.0022)	(.0022)	(.105)	(.087)
Mom Only	0.0629	0.0234	1.424	2.634
	(.0017)	(.0019)	(.088)	(.074)
*Two * Mom Only*	-0.0262	-0.0436	-1.369	-0.166
	(.0029)	(.0031)	(.145)	(.122)
*Mom Only * Post*	0.0149	0.0192	1.544	1.110
	(.0022)	(.0023)	(.108)	(.091)
*Two * Mom Only * Post*	0.0388	0.0391	1.715	1.407
	(.0031)	(.0033)	(.155)	(.130)
# Boys	0.0009	0.0005	-0.026	0.038
	(.0014)	(.0014)	(.067)	(.056)
*Two * # Boys*	-0.0006	-0.0011	0.025	-0.028
	(.0017)	(.0017)	(.081)	(.067)
*# Boys * Post*	-0.0006	-0.0010	0.042	0.009
	(.0019)	(.0019)	(.090)	(.075)
*Two * # Boys * Post*	0.0014	0.0024	0.045	0.058
	(.0022)	(.0023)	(.105)	(.088)
*Mom Only * # Boys*	-0.0080	-0.0085	-0.353	-0.375
	(.0022)	(.0024)	(.112)	(.093)
*Two * Mom Only * # Boys*	0.0031	0.0053	0.311	0.333
	(.0024)	(.0026)	(.119)	(.100)
*Mom Only * # Boys * Post*	**0.0039***	**0.0058****	**0.241****	**.224****
	(.0024)	(.0026)	(.118)	(.099)
# Observations	2591852	2591852	2591852	2591852
R-squared	0.0135	0.0111	0.0193	0.0243
Mean of Dep. Var	0.7155	0.6763	33.373	28.070

Sample includes all mothers living with one or more children under the age of 18 from the 1990 and 2000 IPUMS. Unit of observation is the mother and regressions are weighted by the average weight for all children in the family. The first two dependent variables capture labor supply at the time of the survey while the latter two refer to labor supply during the previous year. Robust standard errors are included in parentheses and all regressions include state fixed effects.

Figure A: SSI Applicants per 100 Children in AL, AR, LA, MS

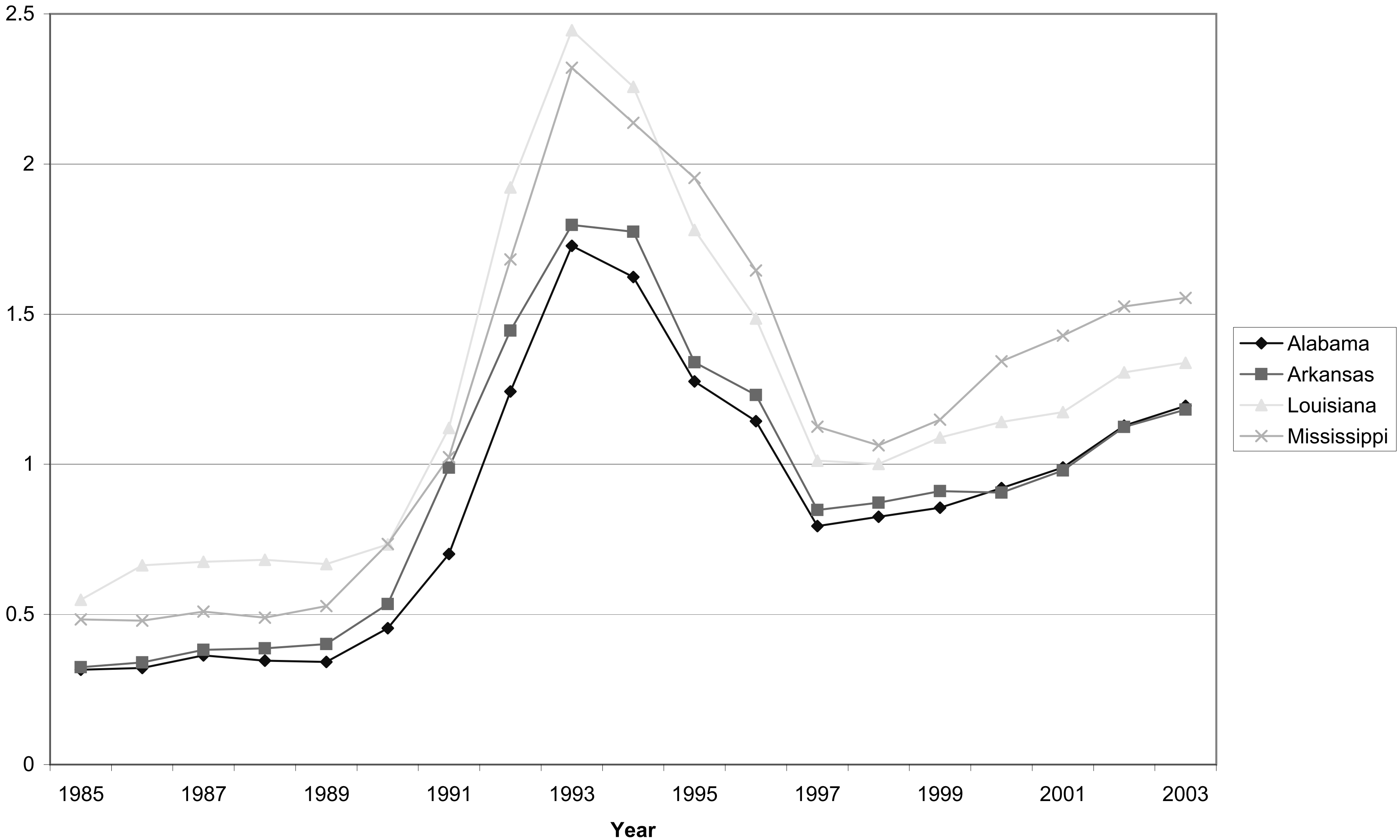


Figure B: SSI Applicants per 100 Children in AZ, CA, NM, TX