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THE EARNED INCOME TAX CREDIT AND TRANSFER PROGRAMS: A STUDY OF LABOR MARKET AND PROGRAM PARTICIPATION

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EXECUTIVE SUMMARY

The cornerstone of the Clinton administration's welfare reform agenda is a large expansion of the earned income tax credit (EITC), a refundable tax credit directed primarily toward low-income taxpayers with children. This paper reviews existing studies and provides new evidence on the degree to which policies, like the EITC, that alter after-tax wages affect hours of work, labor market participation, and transfer program partici-

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pation. Simulations based on recent labor supply estimates suggest that the overall effect of the EITC expansion on hours of work from those in the labor market will be negative but fairly small. We then examine the effect of the EITC on labor market participation. We use a detailed SIPP-based microsimulation model of the tax and transfer system to accurately characterize families' budget constraints. Our empirical model relates labor market and program participation decisions to budget constraint variables and other characteristics. We find that the positive effect of the EITC on labor market participation offsets and, depending on the hours and weeks worked by new labor market participants, can exceed the negative effect of the EITC on hours worked by those already in the labor force. We also show that transfer program participation is negatively correlated with after-tax wages, which should, over time, lower the cost of the EITC.

1. INTRODUCTION

The cornerstone of the Clinton administration's welfare reform agenda is a large expansion of the earned income tax credit (EITC), a refundable tax credit directed primarily toward low-income taxpayers with children. In fiscal year 1998 the EITC is expected to cost the federal government $24.5 billion, $7 billion of which will result from expansions incorporated in the 1993 Omnibus Budget Reconciliation Act (OBRA93). In contrast, the federal share of the Aid to Families with Dependent Children (AFDC) program is expected to be $16 billion in 1998. Even though the EITC has been in the tax code since 1975 and now takes a central role in the nation's antipoverty policy, relatively little has been written about it.

Like the federal government, many states also have embarked on ambitious plans to reform welfare. From January 1, 1992, to June 14, 1994, 20 states received waivers from the Department of Health and Human Services to alter aspects of the AFDC program (Wiseman, 1993, 1994). Although the range of experiments is very broad, one of the most popular changes, adopted by California, Colorado, Florida, Illinois, Iowa, Michigan, Utah, Vermont, and Wisconsin, reduces the rate at which recipients lose AFDC benefits as they earn income.

The EITC and many of the state welfare waivers are part of a "make work pay" strategy of welfare reform. These policies attempt to reduce families' reliance on AFDC and other transfers by increasing the after-tax return to work, which in turn is expected to increase labor market participation and hours of work. The effectiveness of these policies depends, in part, on the degree to which people respond to change in incentives. This paper reviews existing studies and provides new evidence on the degree to which after-tax wages and other factors affect labor market participation, hours of work, and transfer program participation.

Many papers have been written on various aspects of these issues. In the following section, we survey portions of the literature, paying special attention to the studies' implications for the OBRA93 EITC expansions. Estimates from the literature imply that the 1993 EITC expansion will reduce hours of work by taxpayers already in the labor force because most EITC recipients have incomes in the phaseout range of the credit, where additional earnings reduce credit payments. Nevertheless, most labor supply estimates imply that the elasticity of hours of work with respect to the after-tax wage is small for both men and women, so the overall effect of the EITC expansion on hours is expected to be fairly small.

Less attention has been paid in the labor supply literature to the effects of wage rates, taxes, and transfers on labor market participation, particularly for women. Heckman (1993), for example, notes the lack of attention paid to participation decisions in the literature and then writes, "Participation (or employment) decisions generally manifest greater responsiveness to wage and income variation than do hours-of-work equations for workers." The participation margin is particularly important for the EITC, because the structure of the credit ensures that it will have its most beneficial labor market effects through participation. Consequently, we concentrate on participation rates in our empirical analysis. Because studies have also found that transfer programs affect the behavior of single-parent families differently than their two-parent counterparts, we also focus on differences between family types.

We examine the determinants of labor market and transfer program participation by first carefully modeling the budget constraints that families face. State income tax and AFDC rules vary across states, and all transfer programs have asset tests, income restrictions, rules on household composition, and complex interactions with other programs. These features, together with the intricacies of the federal income tax, make modeling these programs and their interactions a major undertaking. We develop a detailed microsimulation model, described in Section 3, that uses monthly data for the 1990 calendar year drawn from the Survey of Income and Program Participation (SIPP) to calculate benefits and taxes. The model is coded in the computer language C, runs on a personal computer, and contains detailed modules for SSI, AFDC, food stamps, the federal income tax, state income taxes, and payroll taxes. The current version of the model contains more than 10,000 lines of executable code, fully reflects tax and program interactions, and provides accurate estimates of program benefits and taxes. Section 3 also describes the policy variables (wage rates, taxes, and transfers) and outcome variables (transfer program participation, labor market participation, and hours) that are the focus of our research.

In Section 4 we describe a simple empirical model of labor market and program participation based on work by Moffitt and Wolfe (1992). The empirical model relates participation decisions to budget constraint variables and other characteristics. Estimates from the model suggest that wage rates are positively correlated and benefit guarantees are negatively correlated with labor market participation. In Section 5 we use our estimates to simulate the effect of the EITC on labor supply across family types. Our results imply that the OBRA93 EITC changes will increase the labor force participation of single-parent families in our sample by 3.3 percentage points, evaluated at the mean characteristics of the sample. The response of other family types is smaller. We find that the positive effect of the EITC on labor market participation offsets and, depending on the hours and weeks worked by new labor market participants, can exceed the negative effect of the EITC on hours worked by those already in the labor force. We also show that transfer program participation is negatively correlated with after-tax wages, which should, over time, lower the cost of the EITC.

2. THE LABOR SUPPLY LITERATURE AND THE EITC

The EITC is a credit on the federal income tax available to working poor families with children. Unlike most credits and deductions in the federal individual income tax system, the EITC is refundable—that is, if the amount of the credit exceeds what the taxpayer owes, he or she receives a payment from the U.S. Treasury for the difference.[1]

The EITC schedule can be divided into three ranges. In the subsidy range, the amount of the credit increases with every dollar of earned income. In 1994, for example, the credit equals 26.3 percent of earned income (wages, salaries, self-employment income, and farm income) for taxpayers with one child, up to an earned income of $7,750; hence, their maximum benefit is $2,038 (26.3 percent of $7,750). In the flat range—which, in 1994, is between $7,751 and $11,000 for taxpayers with one child—taxpayers receive the maximum credit. In the phaseout range, EITC benefits are reduced with every dollar of earned income. In 1994,

[1] Scholz (1994) discusses the EITC and shows that the participation rate—the percentage of eligible taxpayers who actually receive the credit—was 80 to 86 percent in 1990. Holtzblatt, McCubbin, and Gillette (1994) discuss the labor market incentives of the EITC. Alstott (1995) discusses the design of and policy checks surrounding the EITC in the context of the broader tax and transfer system.

taxpayers with one child and incomes exceeding $11,000 have their $2,038 credit reduced by 15.98 cents for every dollar of income until the credit is eliminated at an income of $23,760.[2] For the first time, the EITC is also available to childless taxpayers in 1994, though the maximum credit ($306, or 7.65 percent of $4,000) is considerably smaller than that available to other taxpayers.

The EITC, which was adopted in 1975, was originally promoted as a way to relieve the burden of the social security payroll tax on low-wage working parents.[3] The credit has grown dramatically in the last 20 years. The original EITC equaled 10 percent of earnings up to a maximum credit of $400 for taxpayers with children and was phased out at a rate of 10 cents per dollar of earnings (or adjusted gross income, whichever was higher) for incomes between $4,000 and $8,000. In 1996 when the OBRA93 changes are fully phased in, the credit rate will be 40 percent of earnings for families with two or more children and 34 percent for families with one child. The maximum credit (in 1994 dollars) for taxpayers with two or more children will be $3,370; for taxpayers with one child, $2,040; and for taxpayers with no children, $306. Table 1 summarizes EITC parameters for several years discussed in this paper.

The EITC has different labor supply incentives depending on the taxpayer's income relative to the subsidy, flat, or phaseout range of the credit. For taxpayers with no earned income, the substitution effect associated with higher wages will provide an unambiguous incentive to enter the labor market.[4] For taxpayers with incomes in the subsidy range, the substitution effect provides an incentive to increase hours of work, whereas the income effect provides an incentive to decrease hours of work. The net effect is ambiguous. There is only an income effect in

[2] Taxpayers with two or more children in 1994 are entitled to a larger credit ($2,528, or 30.0 percent of income up to $8,425). This credit is phased out at a 20.22-percent rate for taxpayers with incomes between $11,000 and $26,000.

[3] The credit also has been defended as an income security program for low-income families, a work incentive for welfare recipients, a subsidy to take into account the child care and health care needs of children in low-income families, and an efficient mechanism for offsetting the effects of regressive federal tax proposals. See Yin et al. (1994) for further discussion of the credit.

[4] Price changes can be decomposed into two effects. A wage subsidy increases the return to labor, making leisure more expensive. The substitution effect suggests that as leisure becomes more costly, people take less (work more) holding utility constant. The income effect suggests that with a higher wage rate, people have more income for given hours of work. With higher income, people buy more of everything they like, including leisure, which implies they will work less. For people just entering the labor market, there is no income effect. Hence, the EITC, by offering a higher wage, provides an unambiguous incentive to enter the labor market.

TABLE 1.
EITC Parameters under Law Prior to OBRA93 and under OBRA93, Selected Years

		Flat region			Phaseout region	
	Credit rate (percent)	*Beginning income*	*Ending income*	*Maximum credit*	*Phaseout rate (percent)*	*Income cutoff*
Pre-OBRA93 Law						
1990						
1+ Children	14	$6,800	$10,750	$953	10.00	$20,264
1993						
1 Child	18.5	7,750	12,200	1,434	13.21	23,050
2+ Children	19.5	7,750	12,200	1,511	13.93	23,050
Young child[a]	5	7,750	12,200	388	3.57	23,050
Health credit[b]	6	7,750	12,200	465	4.285	23,050
Omnibus Budget Reconciliation Act of 1993 (OBRA93)						
1994						
1 Child	26.3	7,750	11,000	2,038	15.98	23,760
2+ Children	30.0	8,425	11,000	2,528	17.68	25,300
No child[c]	7.65	4,000	5,000	306	7.65	9,000
1996 and beyond						
1 Child	34.0	6,000	11,000	2,040	15.98	23,760
2+ Children	40.0	8,425	11,000	3,370	21.06	27,000
No child[c]	7.65	4,000	5,000	306	7.65	9,000

Source: Figures for the August 1993 budget agreement (OBRA93) were kindly provided by Janet Holtzblatt at the Office of Tax Analysis, U.S. Department of Treasury. The other figures are from U.S. Congress (1993).

Note: Figures for 1994 and beyond are in 1994 dollars.

[a] The young child (or "wee tots") credit was for taxpayers who had a child under the age of one in the tax year and incomes in the ranges designated in the table.

[b] The supplemental health insurance credit goes to taxpayers with incomes in the ranges designated in the table who paid health insurance premiums that included coverage for one or more qualifying children. The taxpayer cannot take advantage of the supplemental health insurance credit on expenses used for the medical expense deduction or health insurance deduction for the self-employed (and vice versa).

[c] The taxpayer must be between the ages of 25 and 65.

the flat range of the credit, which provides an incentive to decrease hours of work. In the phaseout range, the substitution and income effects work in the same direction and both provide an incentive to decrease hours of work.

The labor market effects of the credit depend on the distribution of taxpayers within the credit's ranges and the degree to which people in and out of the labor market respond to incentives. Scholz (1994) estimates that in 1996, 77 percent of EITC recipients will have incomes that fall in the flat or phaseout range of the credit, which raises the concern that the EITC may lead to a net reduction in the labor supplied by low-income workers. The large literature on the determinants of hours of work yields labor supply elasticities that enable us to estimate the effects of the EITC on hours of work. We discuss the literature—conventional labor supply models, models that incorporate kinked budget constraints, evidence from the income maintenance experiments, and other studies of transfer programs—in the following subsections. Fewer papers provide guidance for thinking about the effects of the EITC on labor market participation. Consequently, the primary focus of our empirical work is on the determinants of labor market and transfer program participation.

2.1 Empirical Models of Men's and Women's Hours of Work

A common empirical model of labor supply (see Pencavel, 1986, p. 52) is

$$H_i = \alpha_0 + \alpha_1 W_i + \alpha_2 Y_i + \alpha_3 A_i + \epsilon_i, \qquad (1)$$

where H_i is the hours of work, W_i is the market after-tax wage rate, Y_i is nonlabor income, A_i is a vector of exogenous household characteristics, α_i's are parameters to be estimated, and ϵ_i is a normally distributed error term. This specification, along with the assumption that a worker with a specific set of characteristics faces a horizontal demand curve for his or her services, is sufficient to identify a labor supply function consistent with utility maximizing behavior. Because U.S. labor force participation rates for prime-age males are very high,[5] and, hence, wages and hours for virtually all prime-age males are observed, Equation (1) can be estimated to study the labor supply of men directly.

Models similar to equation (1) are also a common specification of female labor supply. Although women's labor force participation rates have risen sharply over the last 100 years, labor force participation rates for women are still not nearly as high as they are for men.[6] For familiar reasons of sample selection, estimating equation (1) with a sample of working women can lead to biased estimates of all the parameters of interest. Consequently, empirical models of female labor supply frequently adopt a two-stage estimation approach (Heckman, 1976, 1979).

[5] In data from the Survey of Income and Program Participation (described below), 93.0 percent of males age 25 to 64 work.

[6] The labor force participation rate for women age 25 to 64 is 59.3 percent in our data from the Survey of Income and Program Participation.

In the first stage, parameters governing the decision to work or not to work are estimated, generally through probit regression. In the second stage, an equation similar to equation (1) is estimated, augmented with a term reflecting the conditional mean of ϵ_i (and often using instrumental variables for the wage rate).

Labor supply models derived from equation (1) have considerable appeal.[7] Models for both men and women are easily estimated using common statistical software, the linear labor supply equation results from a somewhat unusual but nevertheless well-defined utility function,[8] and the parameter estimates are easily interpreted as income and substitution effects (see footnote 4). Specifically, the income effect is equal to $H\alpha_2$, while the substitution effect is $\alpha_1 - H\alpha_2$.

Estimated elasticities from equation (1) for the labor supply of men show a fairly consistent pattern. Pencavel (1986) surveys a number of studies of male labor supply up to 1982; he reports that estimates of the elasticity of hours worked with respect to wages (derived from α_1) cluster between -0.17 and -0.08. Estimates of the income effect, again in elasticity form, range from -0.63 to 0.08. In only 7 of the 12 studies, however, is the substitution effect positive as the conventional labor supply model suggests it should be, which leads Pencavel to question the empirical relevance of the conventional model.

Estimated elasticities of the labor supply of women are considerably more varied than those for men. Killingsworth and Heckman (1986, p. 179) write, "There has been a consensus of relatively long standing that compensated and uncompensated female labor supply wage elasticities are positive and larger in absolute value than those of men." They then summarize 31 studies with 93 sets of estimates, where uncompensated wage elasticities of annual hours (α_1 in elasticity form) range from -0.3 to more than 14.0. Estimates of the income effect appear to be more precisely estimated than the uncompensated wage elasticities and range (with a few exceptions) from -0.02 to -0.48. The view that women's labor force behavior differs significantly from that of men has shifted

[7] A number of complications that arise with empirical implementation of this model are described below. We do not discuss dynamic models of labor supply (MaCurdy, 1981; Blundell and Walker, 1986) or household bargaining models of family labor supply (McElroy, 1990).

[8] The utility function consistent with equation (1) is

$$U(X,H;A,\epsilon) = \left(\frac{\alpha_2 H - \alpha_1}{\alpha_2^2} \right) \exp \left\{ \frac{\alpha_2(\alpha_0 + \alpha_2 X + \alpha_3 A + \epsilon) - \alpha_1}{\alpha_2 H - \alpha_1} \right\}$$

(Deaton and Muellbauer, 1981; Hausman, 1981; Pencavel, 1986).

over the last 10 years, however, driven in part by an influential paper by Mroz (1987). In that study, Mroz presents a detailed sensitivity analysis of the economic and statistical assumptions used to estimate the conventional model of female labor supply. Unlike much of the previous literature, Mroz's paper finds in his most reliable specifications that the compensated and uncompensated wage elasticities of women workers are similar to those of men.

Studies based on equation (1) are poorly suited for estimating the labor supply effects of the EITC. For taxpayers in the subsidy range of the credit, the coefficient on wages, α_1, can be used to assess the effects of the credit as long as nonlinearities caused by the tax and transfer system are ignored. The EITC's effect on labor supply for households in the flat range of the credit, however, cannot be determined from α_2 in equation (1), because Y in the canonical model is exogenous nonlabor income (such as dividends and other capital income), while the EITC depends on labor market choices. A similar problem arises in the phase-out range of the credit, where the EITC will reduce net wages by 15.98 percent for taxpayers with one child and 21.06 percent for taxpayers with two or more children in 1996. The general problem arises because the standard model of labor supply treats the tax system as being proportional, while the EITC creates nonlinearities in the budget constraint.

Nonlinearities that arise from the tax and transfer system are addressed in an alternative approach to the labor supply problem, popularized by Hausman (1981). Each (linear) segment of the kinked budget constraint faced by taxpayers can be fully characterized by two parameters: the slope (the after-tax wage rate) and the intercept (the "virtual income").[9] Optimal hours are found for each linearized segment of the budget constraint (replacing Y in equation [1] with virtual income), with hours being determined by the segment or kink point that yields the highest utility (see Hausman, 1981, or Moffitt, 1986, 1990 for details).

Applying the kinked budget set approach to the labor supply problem, Hausman (1981) estimated an uncompensated wage elasticity for males that was close to zero. However, he also estimated a large negative income elasticity. Because each segment of the nonlinear budget constraint is characterized by a virtual income term that generally exceeds zero, Hausman's results implied that progressive income taxation in the United States generates large reductions in male labor supply and large efficiency losses. Hausman's work was the first to incorporate

[9] Virtual income is the income a household would have if the given linear segment of the budget constraint were extended to the vertical axis at zero hours of work in the typical leisure-consumption diagram.

rigorously the effects of the tax system in the empirical model, which raised the possibility that the small behavioral effects found in the earlier literature resulted from ignoring nonlinearities caused by the tax system.

This possibility is examined by MaCurdy, Green, and Paarsch (1990) and Triest (1990), who estimate kinked budget set models of labor supply.[10] Both studies conclude that the hours decisions of prime-aged married men are relatively invariant to net wages and virtual incomes. Triest finds that the hours decisions of married women are only slightly more sensitive to changes in taxation than are the hours decisions of men, but he raises the possibility that their participation decisions may be quite sensitive to changes in the net wage.

Estimates from kinked budget set studies provide a natural way to discuss the EITC's effects on hours of work, given the nonlinearities caused by the credit. To simulate the effects of the OBRA93 EITC expansion, we use labor supply parameters from the kinked budget constraint literature and data from the 1990 Survey of Income and Program Participation. Specifically, we simulate the EITC that each family in our sample would receive in 1993 and in 1996. We then calculate the percentage change in real net wages (assuming that all tax rates stay the same except for the EITC rate) and estimate the change in virtual income for each family.[11] We then apply representative estimates of the wage and income elasticities from the kinked budget set literature to simulate the effect on hours worked.

The first three columns of Table 2 summarize the implications of the kinked budget set literature for the OBRA93 EITC expansions in 1996, relative to the law that was in effect in 1993. We use elasticities estimated by Triest (1990) as the central parameters. Triest finds uncompensated wage elasticities of around 0.05 for men and 0.25 for women, and his estimates of the virtual income elasticity are 0.0 for men and −0.15 for women. These parameters imply that the EITC increases hours of work by 3.9 percent in the credit's subsidy range and imply a modest negative effect on labor supply for taxpayers in the phaseout range of the credit. Given that EITC recipients are disproportionately in the phaseout range of the credit and these households work more than other EITC recipi-

[10] MaCurdy, Green, and Paarsch (1990) and MaCurdy (1992) show that the kinked budget set approach requires that the substitution effect, $\alpha_1 - H\,\alpha_2$, be nonnegative for all interior kink points on all individuals' budget constraints. In practice, they argue that this requirement rules out the possibility that labor supply is backward-bending.

[11] For secondary wage earners, the changes in net wages and virtual incomes that arise from the EITC are calculated if one assumes the hours of the primary earner fixed at their observed value.

TABLE 2.
Simulated Labor Supply Responses to Changes in EITC Law from 1993 to 1996

	Estimated percent change in annual hours worked[a]					
	Kinked budget set simulations[b]			Simulations using NIT parameters[c]		
	MaCurdy et al.	*Triest*	*Hausman*	*Johnson and Pencavel*	*Mean Parameters*	*Robins and West*
All recipients	−0.09	−0.54	−4.04	−1.16	−1.17	−1.63
By credit range						
Subsidy	1.88	3.92	13.46	6.88	2.44	2.25
Flat	−0.09	−0.19	−1.79	−0.60	−1.08	−1.64
Phaseout	−0.53	−1.11	−4.73	−1.46	−1.50	−1.63
By marital status						
Husbands	0.00	−0.34	−3.17	−1.44	−1.32	−1.47
Wives	−1.47	−3.03	−11.36	−1.43	−2.64	−4.09
Single female heads	−0.53	−1.11	−4.02	−0.79	−1.08	−1.63
Single male heads	0.00	−0.18	−1.56	—	—	—
By sex						
Male	0.00	−0.34	−3.15	−1.44	−1.32	−1.47
Female	−0.57	−1.17	−4.33	−0.93	−1.08	−1.63

Note: The estimates given for the kinked budget set simulations are median percentage changes. Medians are presented instead of means because a small number of very low-income single parents in the subsidy range have extremely high marginal tax rates and, therefore, extremely large simulated wage effects.

[a] The median monthly hours for the sample is 160. If one reads down the rows of the table, median monthly hours are 80 (for the subsidy range), 148, 160, 180, 140, 160, 160, 180, and 160.

[b] The wage and virtual income elasticities for the kinked budget set simulations are as reported in Triest (1994). The elasticities from Triest (1990) are presented as the central estimates; MaCurdy, Green, and Paarsch (1990), the low estimates; and Hausman (1981), the high estimates.

[c] Parameters from the NIT studies are as reported in GAO (1993). The estimates from Johnson and Pencavel (1984) imply the least negative labor supply. The parameters from Robins and West (1983) imply the largest negative effects. These parameters along with the arithmetic mean have been adjusted to 1994 dollars.

ents, the overall effect of the EITC on the labor supply of working recipients, based on Triest's estimates, is negative, but small (−0.54 percent).

Triest's estimated elasticities imply that the EITC will affect the hours worked by men and women differently. Men are relatively unresponsive to the EITC's incentives to alter hours. Only the small wage effect is important, since the estimated income effect is zero. Consequently, Triest's parameters imply that men will reduce their hours by only 0.34 percent due to the EITC expansion. For women in the subsidy range of the credit, the wage elasticity of 0.25 implies a larger positive effect of the EITC on hours than his estimate for men. In all three ranges of credit, women will reduce hours of work due to the negative virtual income effect of −0.15, and women in the phaseout range will further reduce hours due to the positive uncompensated wage elasticity. The specific amount of the reduction in hours depends on the size of the maximum EITC relative to virtual income in the absence of the EITC. Overall, the Triest parameters imply that the EITC will reduce the labor supplied by women workers by 1.2 percent.[12] Estimates of MaCurdy, Green, and Paarsch (1990) imply slightly lower responses of both men and women to increases in the EITC, while the estimates of Hausman (1981), imply much higher effects relative to those using the Triest parameters, particularly for women and taxpayers in the subsidy range of the credit.

Like the older literature on labor supply, the kinked budget set papers provide little guidance for the way the EITC might affect labor force participation, though the aggregate effect of the EITC on labor supply will depend, in part, on the credit's effect on labor market participation. In addition, the observed responses to taxes, benefits, and wages by low-income households may differ from those of other households in ways not reflected by the common empirical specifications. For example, the transfer system sharply alters the budget constraints faced by low-income households. Papers in the labor supply literature typically either ignore or incorporate only simplified representations of the transfer system in their analyses.

2.2 *Income Maintenance Experiments*

Between 1968 and 1982, the United States sponsored two rural income maintenance experiments in North Carolina and Iowa, and two urban income maintenance experiments, one in Gary, Indiana, and one split between Seattle, Washington, and Denver, Colorado. Robins (1985) reports that $225 million (in 1984 dollars) was spent on these experiments, of which $63 million represented direct payments to families. The main purpose of these experiments was to measure the work effort and earnings effects of higher transfers, including a negative income tax (NIT), on the low-income population. Pencavel (1986) surveys NIT labor supply estimates for men, and Robins (1985) does the same for all family types.

Hoffman and Seidman (1990, Chap. 3) and the U.S. General Account-

[12] Triest (1994) simulates the welfare effects of several policies that would increase the progressivity of the U.S. individual income tax using estimated labor supply elasticities from the literature. He concludes that expanding the EITC is a particularly efficient way of increasing progressivity.

ing Office (1993, Chap. 3) simulate the effects of the EITC on hours of work, using behavioral parameters estimated from studies of the income maintenance experiments. These studies' results are difficult to compare with our simulations based on the kinked budget set literature because the underlying data differ and because the OBRA93 EITC expansion is more generous than the policies examined by Hoffman and Seidman and the GAO. Therefore, we extend the GAO simulations using data from the 1990 SIPP to examine the OBRA93 EITC expansion. Following the GAO's approach, we apply estimates of the wage effects on hours from the NIT experiments to the changes in EITC subsidy and clawback rates, and estimates of the income effects on hours to the change in the level of the EITC between 1993 and 1996. We use an arithmetic mean of estimates from several NIT studies for the central behavioral parameters. For comparison, we also use estimates from Johnson and Pencavel (1984), since those imply the least negative labor supply effects, and estimates from Robins and West (1983), since those imply the largest negative labor supply effects.

The results are summarized in the last three columns of Table 2. The labor supply responses implied by the NIT are somewhat larger than those implied by the low and medium kinked budget set estimates (the first two columns of Table 2). The discrepancy between the two sets of estimates primarily arises from differences in the underlying behavioral parameters for female heads of households. The NIT experiments allowed wage and income elasticities for female family heads to be estimated separately from the elasticities for other women. In the kinked budget set simulations, we used elasticities for women that mix what appear to be disparate labor market responses of married and unmarried women. Though the estimates differ between the two sets of elasticity estimates, the labor market simulations show that the large change in the EITC between 1993 and 1996 is expected to have a fairly small negative overall effect on the hours worked by EITC recipients, under all but the Hausman (1981) parameter estimates.

Like the rest of the literature, the NIT experiments provide little evidence about the likely effects of the EITC on labor market participation. In addition, an important qualification to both sets of labor market simulations is that we assume that workers perceive the EITC as an increase in their after-tax wage. The design of the income maintenance experiments made clear the links between transfer payments, earned income, and the benefit reduction rate. In contrast, 99.5 percent of EITC recipients receive the credit in a lump sum after filing a tax return (U.S. General Accounting Office, 1992), so the links between earnings, benefits, and the phaseout may be less clear to recipients. To the extent that workers do not associate the EITC with higher net-of-tax wages, our simulations presumably overstate the effects of the credit on hours.[13]

2.3 Microeconometric Studies of Transfer Programs

Several papers estimate structural models of the effects of income transfers on the labor market behavior of female-headed households with children.[14] Blank (1985) estimates a labor supply model, similar to equation (1), jointly with a welfare participation model to examine how cross-state variation in wages, welfare benefits, demographic characteristics, and taxes affects labor market behavior. AFDC participants are given a state-specific implicit tax rate, which, as we discuss below, masks variation in taxes within a state that depends on family size and structure, unearned income, and "disregards."[15] Tax rates for nonparticipants include federal and state income taxes and payroll taxes. Blank finds that wages are positively correlated with hours of work and that other income and welfare benefits are negatively correlated with hours, but the economic effects of these variables are relatively small for her sample.[16] She emphasizes that factors other than income and taxes are very important in influencing welfare decisions.

Fraker and Moffitt (1988) examine the effects of food stamps and AFDC on the labor supply of female-headed households with children. Their empirical approach models AFDC participation, food stamp participation, and a discrete hours decision (0, 20, and 40 hours) and addresses a number of complications that arise in the empirical labor supply literature. Wages for women not in the labor force are calculated in a manner consistent with Heckman and MaCurdy (1981).[17] Wages are also allowed

[13] Because almost all recipients receive the EITC as a lump-sum payment, we simulated the effects of the 1993 expansion on hours worked if recipients perceive only income effects. Using the parameters from Triest (1990), we find that the overall median change in hours is −0.13 percent, which is clearly smaller than the combined wage and income effects described above.

[14] Danziger, Haveman, and Plotnick (1981) provide a detailed, wide-ranging discussion of the research on income transfer programs on labor supply and saving up to 1981. Moffitt (1992) provides a more recent comprehensive survey of the incentive effects of transfer programs.

[15] Disregards are deductions from earnings used to calculate transfer program benefits. These deductions generally apply to child care expenses, work expenses, and, in some cases, general expenses.

[16] For example, a $1 increase in wages in her sample (the mean wage was $2.91 an hour) would increase hours by one per week.

[17] Potential wages of those who do not work are not observed. Often, wage rates for nonworkers are imputed from auxiliary wage rate regressions, and actual wage rates are used for workers (Triest, 1990; Moffitt and Wolfe, 1992). This is appropriate only if the

to increase with hours of work, and a "stigma" term is included in the model so not all families eligible for benefits actually receive them (Moffitt, 1983). Fraker and Moffitt find no detectable effects of benefit reduction rates on program participation. Their estimated uncompensated wage elasticities vary from 0.26 to 0.35, and their income elasticities vary from −0.07 to −0.11. The wage and income elasticities are comparable with those reported in the earlier literature. The study by Fraker and Moffitt is one of only a few that examines the effect of net wages on program participation. They find that the difference in net income at 20 hours of work and at zero hours of work has a statistically significant positive effect on the labor market participation of female heads of households.

Keane and Moffitt (1991) consider the effects of the tax rates imposed by AFDC, food stamps, and housing programs on the hours and participation decisions of female heads of households using simulation estimation methods. Along with addressing the complications raised by Fraker and Moffitt (1988), they account for the expected value of Medicaid benefits and the expected value of private health insurance in the budget constraint using calculations from Moffitt and Wolfe (1992).[18] Keane and Moffitt estimate an uncompensated wage elasticity with respect to hours of 0.66 and a total income elasticity with respect to hours of −0.24 for single-parent households using data from the 1984 Survey of Income and Program Participation. Thus, relative to the literature described above, their results suggest that the EITC would have considerably larger positive effects on hours in the credit's subsidy range and larger negative effects on hours in its phaseout range.

When Keane and Moffitt simulate the labor market effect of raising the gross wage rate by $1, they find that hours increase by only 38 percent of what their estimates of the uncompensated wage elasticity of hours imply. This occurs for two primary reasons. First, a change in the gross wage has only a small effect on the net-of-tax wage because they calculate that low-income women face extremely high cumulative marginal tax rates, often exceeding 100 percent. Even reducing the AFDC tax rate to 50 percent from 100 percent leaves cumulative tax rates of 60 to 80

wage rates of nonworkers are predicted without error. An alternative is to use the predicted wage for both workers and nonworkers (Blank, 1985; Hoynes, 1993). When the estimated model is nonlinear, this is appropriate only if all families base their decisions on the econometrician's predictions instead of their actual wage rates. Fraker and Moffitt (1988) rigorously account for missing wages by using the predicted wage for nonworkers and then integrating out over the error distribution of the unobserved wage rate.

18 Moffitt and Wolfe (1992) examine the effects of AFDC and Medicaid on the welfare and labor market participation of female heads of families. They find both Medicaid and private health insurance have substantial effects on labor market and welfare participation.

percent. They write, "Such tax rates, especially at the low gross wage rates faced by women participating in welfare programs, imply very small income gains from working." Second, many families not in the labor force are not on the margin of working. This seems to indicate that labor supply studies of low-income households that do not model the labor force participation decision may give a misleading impression of the effect of economic variables on labor market behavior.

Hoynes (1993) presents the only structural empirical model of two-parent families in the literature on labor supply and transfer program participation. She restricts her sample, taken from the 1984 SIPP (but applying to 1986), to two-parent families that in the absence of labor market participation would be eligible for AFDC-UP (the AFDC program for intact families with an unemployed parent). Hoynes finds that two-parent families are considerably more responsive to changes in AFDC-UP program parameters than would be expected from studies of single-parent households. For example, she estimates that eliminating AFDC-UP would increase men's labor supply by 46.9 hours and women's labor supply by 31.6 hours per month, while Moffitt (1983) estimates that eliminating AFDC would increase hours worked by female heads of household by roughly four hours per week. Hoynes also finds fairly large effects on program participation and hours of work from changes in benefit reduction rates, in contrast to much of the literature. She speculates that the more elastic behavior may reflect behavioral differences between one- and two-parent families, but reconciling the results is beyond the scope of her study.

2.4 *Lessons from the Literature and Empirical Questions*

The literature yields consistent evidence that the labor supply of prime-age males is fairly insensitive to changes in wages and incomes. In the standard literature and the income maintenance experiments, uncompensated wage elastics are often estimated to be small and negative (around −0.1), while income elasticities are small and positive (around 0.15). Using the kinked budget set appraoch, Triest (1990) estimates an income elasticity of 0.0 and an uncompensated wage elasticity of 0.05. As shown in Table 2, estimates in this range imply that the EITC will have only minor effects on the hours worked by prime-age males. Because the labor force participation rate for this group is very high, it is also clear that the EITC will not have an important effect on labor force participation.[19]

[19] Even if taxes do not affect hours of work or participation, they may affect other dimensions of labor market behavior such as commuting distances, stress on the job, and forms of compensation (Feldstein, 1993).

Labor supply estimates from the literature on married women show considerably more variability, but recent estimates from the traditional literature (Mroz, 1987), the NIT experiments (Robins, 1985), and the kinked budget set literature (Triest, 1990) suggest that the elasticities of women's hours of work to changes in wages and incomes are similar to those estimated for men. Triest's estimates of an uncompensated wage elasticity of 0.25 and an income elasticity of −0.15 are representative of recent studies. These estimates imply that substantial increases in the EITC are likely to have negative, but not particularly large, effects on hours worked. A sophisticated, careful study by Keane and Moffitt (1991) generates wage and income elasticities that are somewhat larger than those mentioned above, so additional work in this area could still be valuable.

Keane and Moffitt (1991) and Giannarelli and Steuerle (1994) use a microsimulation approach to calculate detailed tax rates on households caused by the tax and transfer system. Their results have pessimistic implications for the effect of the EITC on labor market behavior. Keane and Moffitt find the tax rates faced by a representative low-income, female-headed family range from 75 percent (in Ohio) to 124 percent (in California).[20] Even large increases in after-tax wages, such as those offered by the OBRA93 expansions of the EITC, are likely to have only modest effects on labor market participation given these tax rates. At the same time, the relatively large uncompensated wage and income elasticities they estimate imply that the EITC will have a larger, negative effect on the hours of taxpayers that work than suggested by much of the literature. Thus, Keane and Moffitt's results would imply a fairly substantial reduction in aggregate labor supply by those already in the labor market, which would not be offset by increases in labor market participation.

3. TAXES, TRANSFERS, AND LOW-INCOME FAMILIES

The central focus of our empirical work is on the determinants of labor force and transfer program participation. One difficulty with examining the effects of policies that may affect participation is that a broad range of programs exist that, in conjunction with state and federal tax systems, create a complex set of incentives that are difficult for analysts to characterize. In much of the previous literature, for example, research-

[20] The reported tax rates are calculated from changes in after-tax and after-transfer income associated with increasing hours of work from 0 to 20 hours, and from 20 to 40 hours.

ers use uniform tax rates for all transfer program recipients in a given state. Even within a state, however, taxes and benefits vary depending on sources of income, program participation, assets, demographic characteristics, and income exemptions. To characterize the budget sets facing households, we have developed a simulation model that accurately represents transfer program rules, tax systems, and their complex interactions.

3.1 The Simulation Model and Data

Detailed modules for AFDC, food stamps, Supplemental Security Income (SSI), the federal income tax, state income taxes, and the payroll tax are the building blocks for the microsimulation model. All modules have a common structure: Each defines the unit of analysis for tax and transfer programs, performs income and asset tests or determines adjusted gross income and taxable income, and determines benefits or taxes.[21]

The model uses monthly data from the 1990 SIPP for the period January to December 1990. Using SIPP allows us to make fewer imputations than would be required if we used annual data, such as the Current Population Survey (CPS).[22] In addition, SIPP provides a high level of detail on sources of income, particularly for low-income households. The main advantage of SIPP, however, is that it provides data that allow us to calculate program eligibility and benefits in every month. Over time frames longer than a month, variation in income, labor force participation, assets, and family composition leads to an ambiguous definition of program eligibility and participation. Thus, when we refer to labor market participation, transfer program participation, or other point-in-time concepts, we are referring to the last month in wave 3 of the 1990 SIPP panel, which was collected from September through December 1990.[23] State and federal taxes, which are calculated on an annual basis,

[21] The simulation model is described in more detail in Dickert, Houser, and Scholz (1994). The current version of the model also incorporates data on child care expenses from the SIPP topical modules. These data are used in all the transfer program modules, for federal tax credits, and, in some cases, for state tax credits or deductions. We also use asset information from the SIPP topical modules to implement asset tests associated with transfer programs. Data on alimony payments are also used to assess income eligibility in transfer programs (through "deeming" rules) and as an adjustment to income in the federal income tax module.

[22] SIPP's sample includes roughly one-third as many households as the CPS's sample. SIPP does not separately identify all 50 states because nine smaller states are combined into three groups for confidentiality reasons.

[23] There are four rotation groups in SIPP whose interviews are staggered across months. A "wave" is completed when each of the four groups has been interviewed once. Each SIPP panel generally consists of seven or eight waves.

are appropriately accounted for by adding incomes and benefits over the calendar year.

We restrict our sample in several ways that are common in the literature on transfer programs. We exclude families in which the head or spouse (1) is older than 65 or has a disability, (2) would not be eligible for program benefits even if he or she did not work, (3) does not have children under 18, or (4) is self-employed. The first restriction arises from our focus on labor market behavior.[24] The second restriction, which primarily excludes families with assets that exceed program asset tests, arises from our focus on transfer program participation. The third restriction arises from our focus on the EITC. In 1990 a taxpayer had to support a child to be eligible for the EITC.[25] The fourth restriction arises from the difficulty of calculating wage rates for the self-employed.

The simulation model allows us to identify families eligible for benefits, which is necessary for participation rate analyses. In the following subsection we use the model to describe patterns of eligibility and participation in the food stamp and AFDC programs, focusing on three different groups: single-parent families, primary earners in married couples, and secondary earners in married couples. The data described in the remaining parts of Section 3 are summarized in Table 3.

3.2 Program Eligibility and Participation

AFDC eligibility rules and benefits vary sharply by family structure and across states. Roughly 15 percent of the individuals in the sample are eligible for AFDC. Almost all of these are single parents, of whom 43.8 percent are eligible. Average AFDC benefits, conditional on eligibility, are \$372 per month for single-parent families and are somewhat larger for two-parent families. In 1990, the maximum AFDC benefit for a three-person household in the continental U.S. ranged from \$120 in Mississippi to \$694 in California (U.S. Congress, 1990). The model shows similar variation in AFDC benefits across large states, ranging from an average of \$115.53 per month in Mississippi to an average of \$652.06 per month in

24 The labor market behavior of the disabled population is clearly an interesting topic of research (see, e.g., Haveman, de Jong, and Wolfe, 1991), but SIPP does not have information on the severity of the disability. Because the type of disability may inordinately influence labor market behavior, we have chosen to drop the disabled from our sample.

25 The second and third restrictions imply that we treat asset accumulation and fertility decisions as exogenous to transfer programs. Hubbard, Skinner, and Zeldes (1993) discuss a dynamic, stochastic, life-cycle simulation model where the transfer system has a large effect on asset accumulation. Moffitt (1992) surveys the literature on transfer programs and fertility decisions.

TABLE 3.
Means of Selected Sample Characteristics for Single Parents and Primary and Secondary Earners in Two-Parent Families (standard deviations in parentheses)

	Total sample	Single parents	Primary wage earner	Secondary wage earner
% Eligible AFDC	14.87	43.80	1.74	1.85
Conditional AFDC benefit	372.26	367.88	442.00	408.58
	(214.20)	(206.86)	(313.55)	(272.25)
AFDC participation rate (%)	71.79	76.10	25.00	23.53
% Eligible food stamps	39.94	70.04	26.33	26.33
Conditional food stamp	207.64	204.12	221.58	222.09
benefit	(96.64)	(96.29)	(110.08)	(90.84)
Food stamp participation (%)	51.59	66.67	32.23	34.71
Wage ($ per hour)	7.52	6.55	8.62	7.30
	(4.16)	(3.19)	(3.75)	(4.99)
Tax rate (0 to 20 hours)	40.93	49.62	31.92	42.07
	(25.57)	(28.69)	(24.43)	(16.22)
% Working	70.10	56.44	91.62	60.94
Usual hours worked/week	27.91	20.49	39.19	23.33
	(20.68)	(19.66)	(15.70)	(21.18)
Family size	3.89	3.01	4.28	4.28
	(1.30)	(1.16)	(1.16)	(1.16)
Number of dependents	2.08	1.91	2.15	2.15
	(1.11)	(1.10)	(1.10)	(1.10)
% with children < 6	52.72	46.81	55.39	55.39
% w/ poor or fair health	10.94	17.45	7.07	8.92
% Women	63.84	94.46	26.00	73.99
% Nonwhite	27.84	44.52	20.57	20.02
% Urban	74.19	78.22	72.36	72.36
Region—East	17.57	21.30	15.89	15.89
Region—Midwest	20.91	23.10	19.91	19.91
Region—South	46.50	39.83	49.51	49.51
Region—West	15.40	15.76	15.23	15.23
Number of observations	2669	831	919	919
% of Total sample		31.14	34.43	34.43

Notes: Data are from the 1990 Survey of Income and Program Participation. The mean age is 33.7 years, and this does not vary greatly among the types. The mean level of education is 11.67 grades, and this also does not vary greatly among the types.

California.[26] The variation in AFDC benefits across states dwarfs the variation required to adjust for cost-of-living differences.[27]

Not every family eligible for program benefits actually receives them. (Moffitt [1983] presents a well-known discussion of this issue.) With our model simulations of eligibility and SIPP data on participation, we can examine the participation rate—the fraction of eligible families that report receiving program benefits. In our sample, the ADFC participation rate is 71.8 percent, and 76.1 percent for single-parent families. These rates are comparable to the 62- to 72-percent rates reported by Blank and Ruggles (1993) in their study of participation rates using the 1986 and 1987 SIPP panels, but somewhat higher than earlier results in the literature (Moffitt, 1983). The empirical model estimated in Section 4 examines factors correlated with transfer program participation.[28]

Uniform national rules govern food stamp eligibility and benefits. In our sample, 39.9 percent of families are eligible for food stamps. By family type, 70.0 percent of single-parent families and 26.3 percent of two-parent families are eligible. Average food stamp benefits, conditional on eligibility, are $208 per month, and there is little difference across family types. Because of uniform national standards, there is little cross-state variation in food stamp benefits. The average benefit of $140.62 per month in California, however, is considerably below the sample mean of $207.64. This occurs because California's generous AFDC system reduces the amount of food stamps a family is eligible to receive as AFDC is included in food stamp "countable income."

Studies typically find the food stamp participation rate to be lower than the AFDC rate. We find the participation rate in the food stamps program is 51.6 percent, which is higher than an older estimate of 38 percent by MacDonald (1977) but similar to recent estimates of 54 to 66 percent by Blank and Ruggles (1993). The average participation rate masks sharp variation across family types. The participation rate is 66.7

26 We define a large state as one with at least 50 observations in the sample, which includes 22 states.

27 For example, cost-of-living indices reported in the *New York Times* (8/5/94, A.1) give Mississippi an index value of 86.7 and California a value of 112.3; Los Angeles had a value of 127.9, while Dothan, Alabama, had a value of 87.4. (In both the state and urban indices, 100 denotes an average value.) If cross-state AFDC differences solely reflected cost-of-living differences, maximum benefits in Mississippi would exceed $500, given California's benefit.

28 Both primary earners and secondary earners in two-parent families respond to questions about program participation and benefits, which may account for the slight difference in reported participation and conditional benefits between primary and secondary earners in Table 3. In addition, stepparents may not be eligible for benefits in some states.

percent for single-parent families and around 33 percent for two-parent families.

3.3 *Taxes*

The simulation model also allows us to calculate families' marginal and average tax rates corresponding to actual or counterfactual situations. In our empirical work, for example, we need a tax rate measure that is exogenous to labor market and transfer program participation decisions. Like Keane and Moffitt (1991), we calculate an exogenous tax rate that is based on the change in after-tax and transfer income that would result from increasing hours of work from 0 to 20 (or 40) hours per week. An inevitable problem in labor market studies, however, is that wages are only observed for people who work.

We follow Triest (1990) and Moffitt and Wolfe (1992) and use predicted wages for adults that are not in the labor market.[29] Predictions are based on log wage regressions estimated separately for men and women, dropping the self-employed and adjusting for selection. Covariates include quadratics in age and education, an interaction of age and education, the state unemployment rate for women (in the male wage regressions we use the general state unemployment rate), and dummy variables for Hispanic, African-American, married, and residence in a metropolitan area. We use number of children and a dummy variable for the presence of children younger than six as exclusion restrictions in the probit regression for being in the labor force. Like Hoynes (1993), we calculate median wages from our empirical model because of the skewness of the wage distribution. For both men and women, the regression estimates are similar to other studies in the literature, and predicted wages closely match actual wages. Descriptive statistics for the samples used to estimate wages and the regression results for men and women are available from the authors on request.

Taxes in the model arise from benefit reduction rates associated with income transfer programs (often called implicit taxes) and from payroll, state income, and federal income taxes (often called explicit taxes). When the combined effect of the tax and transfer system is considered, it is clear that tax burdens on low-income families can be very high.

A family receiving AFDC that has earned income is entitled to a number of disregards before AFDC benefits are reduced. A recipient with earnings can subtract $90 per month for work expenses, as much as $175

[29] See footnote 17 for a more detailed discussion of unobserved wages. In every empirical model that we estimate, we examine the sensitivity of our results to using predicted wage for each adult in the sample.

per child per month for child care expenses, and, for the first four months of work, \$30 plus an additional one-third of remaining earnings. For earned income exceeding these disregards, AFDC benefits are reduced by $66\frac{2}{3}$ cents for every dollar of income for the first four months of work. After four consecutive months, the tax rate increases to 100 percent after deducting work and child care expenses (and \$30 for eight additional months).[30]

Food stamp recipients who have earned income are also entitled to a number of disregards before food stamp benefits are reduced. A recipient with earnings receives a standard deduction of \$127, a deduction for dependent-care expenses up to \$160, and an additional deduction of 20 percent of earned income.[31] Participating households are expected to contribute 30 percent of their income exceeding deductions toward food purchases. Thus, the implicit marginal tax rate on income exceeding deductions (including AFDC and SSI benefits) is 30 percent.

The layering of programs is complicated and may lead to even higher benefit reductions than would occur when programs are examined in isolation. The model captures these program interactions. For example, a person receiving benefits from both AFDC and food stamps and who has earned income exceeding the AFDC disregards would face a $66\frac{2}{3}$-percent marginal tax rate on AFDC. From the perspective of the food stamps program, earned income increases by 80 cents for every dollar of earnings (due to the 20-percent disregard for earnings), but unearned income falls by $66\frac{2}{3}$ cents for every dollar of earnings (the reduction in AFDC benefits). The net of these amounts, $13\frac{1}{3}$ cents, is the increase in food stamp "countable income" for every dollar of earnings, and is taxed at a 30-percent rate, which adds 4 percentage points to the AFDC marginal tax rate. Thus, the implicit tax rate on earnings (beyond the disregards) in the first four months of employment for a person receiving AFDC and food stamps is $70\frac{2}{3}$ percent; after four months it is 94 percent.[32]

30 In the calculations below, we do not account for the additional one-third earned income deduction because we cannot assess the duration of AFDC receipt in our data. Thus, our AFDC tax rate is akin to a "long-run" tax. Of AFDC households with earned income, 50.6 percent have the \$30 plus one-third disregard, while roughly 13 percent of all AFDC households have earned income (U.S. Congress, 1993).

31 There are additional deductions for shelter expenses that exceed 50 percent of countable income after the other deductions and for medical expenses for households that include an elderly or disabled member. SIPP does not have data on shelter and medical expenses, so we do not account for these deductions in the model.

32 After four months of earnings, the AFDC tax rate on earnings increases to 100 percent so AFDC benefits fall dollar for dollar. This 100-percent marginal tax is slightly offset by the food stamps program because, although food stamp earned income still increases by 80 cents for every dollar of earnings, unearned income would have fallen by \$1. Food stamp

Explicit taxes also affect low-income families. We assume that the combined 15.3-percent employee and employer shares of the payroll tax are borne by employees. The federal income tax is modeled as precisely as possible given available information in SIPP, except that we assume all taxpayers use the standard deduction. Special attention is paid to accurately modeling the EITC (see Scholz, 1994, for details). Average effective tax rates from the model closely match data from the Internal Revenue Service (1993), though average tax rates on low-income households in our model are somewhat lower than those shown in income tax data. Relative to the IRS data, our low-income tax-reporting units are more likely to be families as opposed to children working part-time jobs and, thus, are more likely to receive the EITC, which leads to lower average effective rates in our model.

We calculate state income taxes using tax forms collected from the 41 states (including the District of Columbia) with state income taxes in 1990. Nine states have credits for low-income filers (including state EITC's in Iowa, Maryland, Rhode Island, Vermont, and Wisconsin); eight states have credits for the elderly and disabled. In addition, we incorporate many other special provisions for dependents and exemptions. Credits vary among states with regard to eligibility, generosity, and whether the credits are refundable. New Mexico, for example, has a low-income credit that uses "modified Adjusted Gross Income (AGI)." Modified AGI includes AFDC, SSI, and food stamp benefits that are calculated in our simulations. Average effective state income tax rates vary in the model from 0.0 percent (several states) to 5.5 percent in the District of Columbia and are strikingly similar to those reported in the Bureau of the Census (1992, Tables 463 and 687).[33]

3.3.1 Cumulative Average and Marginal Tax Rates The layering of implicit and explicit taxes can lead to high average and marginal tax rates on low-income households. To illustrate these tax rates, we present a series of figures showing marginal and average tax rates along several dimensions: a median- versus high-tax family, a low- versus high-benefit state, and single-parents versus primary earners in two-parent families.

Figure 1 shows marginal tax rates for four single-parent families in Texas (a low-benefit state) and New York (a high-benefit state). The families shown are the median and 90th percentile families in each state

countable income, therefore, falls by 20 cents and food stamp benefits would increase by six cents for every dollar of earnings, leading to a cumulative tax rate of 94 percent.

[33] Dickert, Houser, and Scholz (1994) provide comparisons of the simulation model's output with administrative sources.

when ranked according to the cumulative tax rate distribution (going from 0 to 20 hours of work). We calculate marginal tax rates for these representative families from the incremental change in after-tax and after-transfer income resulting from an increase in one hour of work at the earner's market wage. The marginal tax rate is relevant for incremental labor supply decisions.

Each graph has several shaded bars. The dark gray bottom bar in each figure (at low hours of work) shows the marginal federal income tax rate for each family. It is −14 percent at low hours because each family is eligible for the EITC subsidy, which was 14 percent in 1990 (see Table 1). The large, unshaded portion of each bar reflects the marginal tax rate from AFDC. The marginal rate on AFDC is initially zero because of disregards. As hours increase, the AFDC marginal rate is 100 percent until benefits are fully "clawed back." The light gray bars reflect the food stamp marginal tax rate, which is 24 percent when AFDC is untaxed. When AFDC is taxed, the food stamp marginal tax rate is actually negative (see footnote 32). There is a uniform 15.3-percent payroll tax on all families. Low-income families in New York also pay a modest state income tax, shown in black.

Several features are apparent in Figure 1. First, until benefits are fully clawed back, AFDC is the primary source of high tax rates on low-income families. Second, Figure 1 illustrates the progression of federal marginal tax rates for each family. The tax rate is −14 percent when families are in the EITC subsidy range, zero when families are in the flat part of the credit, and 10 percent in the phaseout range. The two New York families reach the 15-percent federal tax bracket (along with the EITC phaseout) at 38 or more hours of work. Third, Figure 1 shows the effect of a "notch" for the median Texas family. Moving from 36 to 37 hours of work, this family would lose $1.50 of food stamp benefits for every $1 of earned income. Families eligible for food stamps must have "gross" income less than 130 percent of the poverty line and must also meet a "net" income test. Our median Texas family, while still meeting the net income test at 37 hours, fails the gross income test and, hence, loses all remaining benefits. Notches are common in the income transfer system due to various asset and income tests.[34] Fourth, cumulative marginal tax rates in the figures are equal to the sum of the positive and negative rates, and they can be very high. Thus, for example, the cumulative marginal tax rate for the median family in Texas working two hours per week is 95.3 percent (the 100-percent AFDC rate + the 15.3-percent

[34] Lyon (1994) and Giannarelli and Steuerle (1994) also discuss notches in income transfer programs.

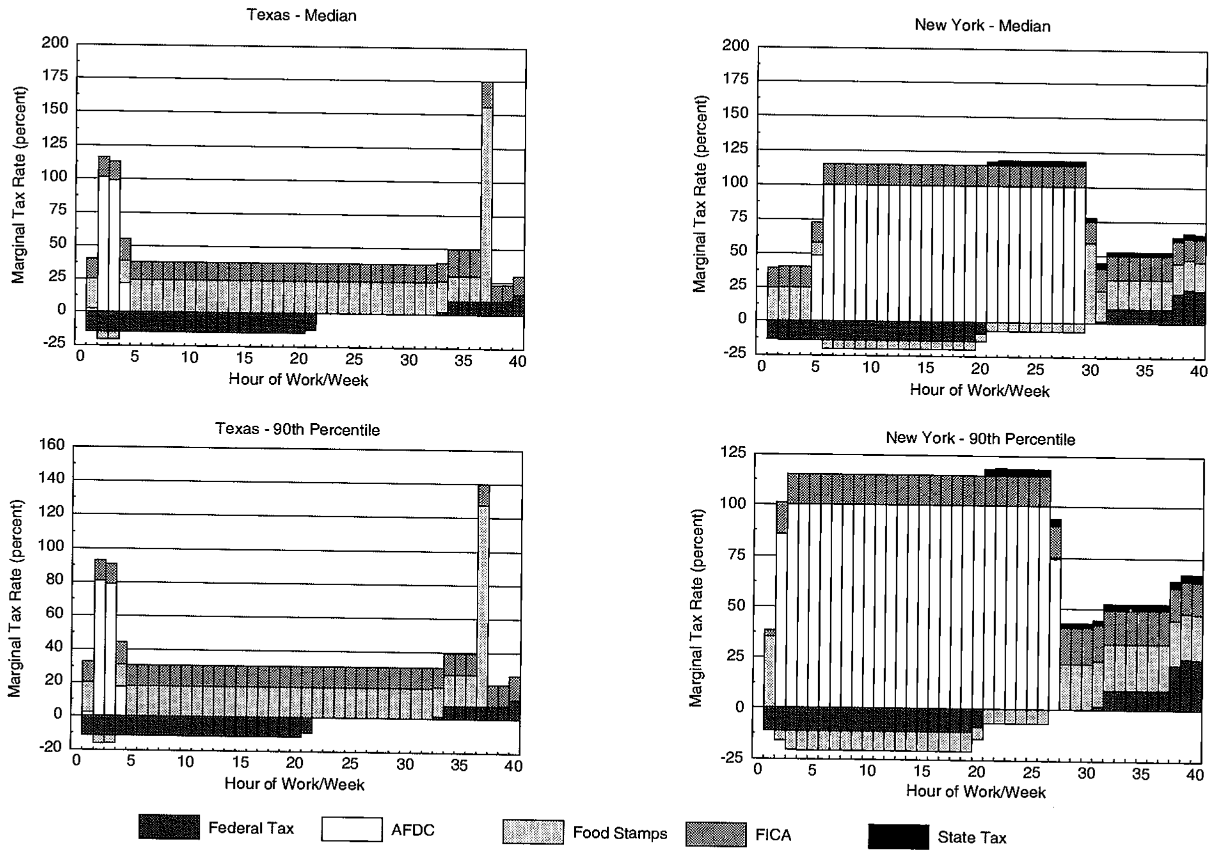


FIGURE 1. ***Marginal Tax Rates for Low-(TX) and High-(NY) Benefit States—Single-Parent Families.***

FICA rate + the −14-percent EITC rate + the −6-percent food stamp rate).

The marginal tax rates for the four families lead to the average tax rates shown in Figure 2. Average tax rates are defined from the change in after-tax and after-transfer incomes when weekly hours of work increase from zero to the number of hours shown on the horizontal axis. We view these average tax rates as being most relevant when families are making labor force participation decisions. The average tax figures tend to have a concave shape that is driven by AFDC. The AFDC average tax rate increases at low hours of work as disregards dissipate. Once disregards are exhausted, average AFDC rates are pulled up by the high marginal rates on AFDC. After benefits are fully clawed back, average AFDC rates again fall. Because the AFDC guarantee in Texas is low, cumulative tax rates in Texas tend to be considerably lower than in New York, particularly when one is looking at the rates from 0 to 20, or 0 to 40 hours of work, which are the relevant participation margins for most families.[35]

Cumulative average tax rates, which again are the sum of the positive and negative bars, can also be extremely high. For example, the average tax rate exceeds 85 percent for the high-tax New York family that enters the labor market and works anywhere from 8 to 35 hours per week. This implies that this family, when making labor market participation decisions, will receive no more than 15 cents for every dollar earned in the labor market over a broad range of hours. Tax rates like these undoubtedly discourage labor market participation. Average tax rates tend to be somewhat lower in low-benefit states and are much lower for families that do not receive transfers, particularly AFDC.

Appendix Figures 1 and 2 show similar figures for the median and 90th percentile two-parent families in Texas and New York. Tax rates apply to the earnings of the primary earner and assume the secondary earner is not in the labor market. The marginal tax rate figure shows several food stamp notches and marginal tax rates that tend to be considerably lower than those faced by one-parent families. This occurs because very few two-parent families are eligible for AFDC-UP in our data. The 90th percentile two-parent family in New York received AFDC-UP and, hence, faced marginal and average tax rates that are similar to those shown in Figures 1 and 2. Otherwise, marginal rates rarely exceed 50 percent, and average rates rarely exceed 40 percent for two-parent families.

The tax rates presented in the top panel of Figure 2 are lower—in

[35] We have made similar graphs for families in Illinois, a medium benefit state, that are available on request. For the median single-parent family the figures show patterns similar to those in Texas, while tax rates on the high-tax family are similar to those in New York.

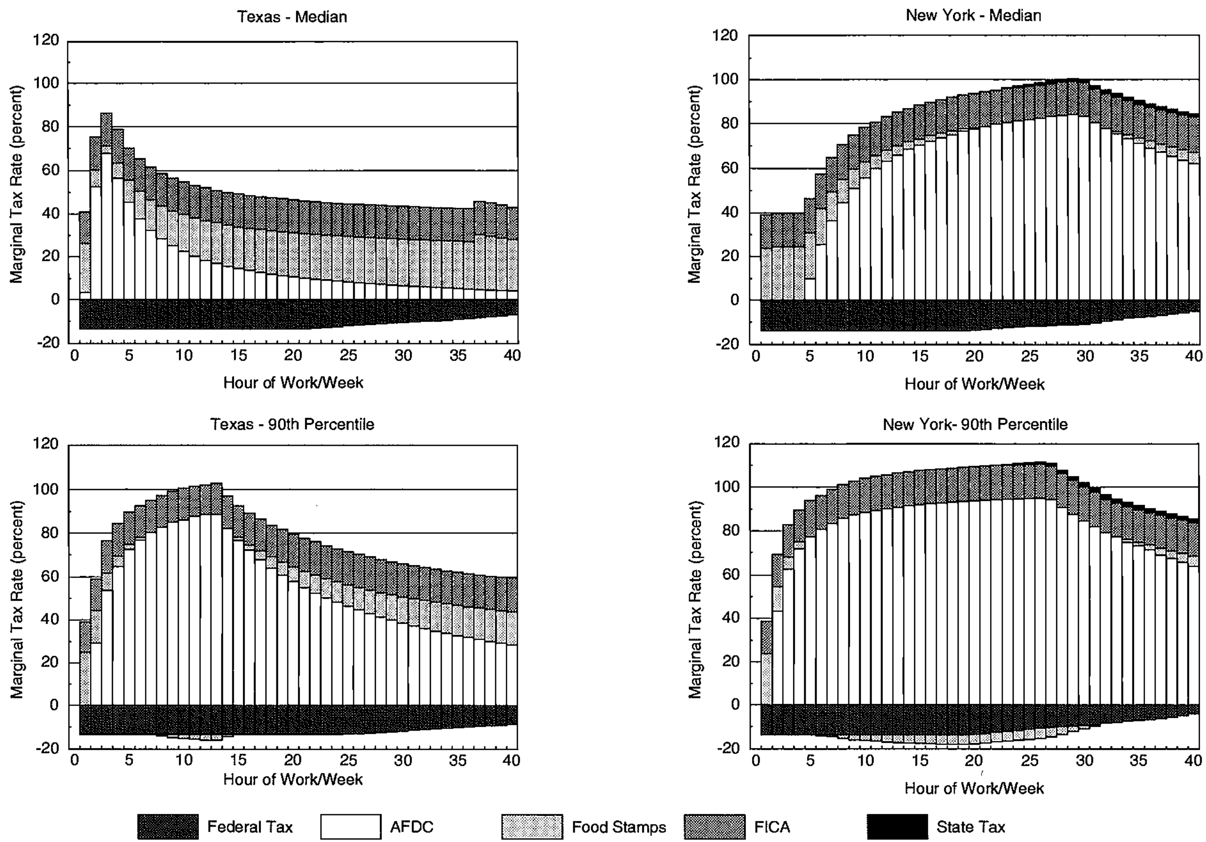


FIGURE 2. *Average Tax Rates for Low-(TX) and High-(NY) Benefit States—Single-Parent Families.*

some cases significantly lower—than those described in Keane and Moffitt's (1991) Table B-1.[36] Their 20-hour average rates, for example, are 95 percent in Texas, 112 percent in Ohio (a medium-benefit state), and 124 percent in California (a high-benefit state). Several factors make our estimated rates lower. The EITC was considerably more generous in 1990 than it was in 1984, the year of their calculations. The wage rate for the representative family in the Keane and Moffitt simulations is lower than the wage rates received by our families, which means AFDC benefits tend to be clawed back over a longer hours interval leading to higher average rates. Keane and Moffitt also assume all families receive housing benefits, which we do not model. Benefit reductions associated with housing increase cumulative rates by 4 to 33 percentage points in their calculations, depending on the state.[37] Finally, Keane and Moffitt include a fixed cost of working that adds 20 percentage points to all families' tax rates. We have no way of assessing the size of work-related expenses, but their existence means our calculations underestimate the change in net disposable income that would result from working. Two modeling differences make our rates higher than Keane and Moffitt's. We incorporate state taxes and treat the combined employer and employee share of payroll taxes as falling completely on workers. (Keane and Moffitt include only the employee share.)

3.4 Labor Force Participation and Hours of Work

Labor force participation rates in our sample vary from 56.4 percent among single parents to 91.6 percent for primary wage earners in two-parent families. In two-parent families in our sample, 61.0 percent of spouses work outside the home. In Section 4 we examine factors correlated with labor market and transfer program participation.

Figure 3 shows the distribution of the usual hours worked per week for single-parent families, primary earners in two-parent families, and secondary earners in two-parent families. Part-time work is uncommon

[36] Table B-1 in Keane and Moffitt (1991) shows implicit and explicit tax rates faced by representative single-parent families; if one assumes wages are $5.20 an hour, nonlabor income is $4 per month, child care expenses are zero, and all families participate in programs for which they are eligible. In contrast, we use data for each family in our sample to make our calculations. We calculate tax rates for all families in the data using their reported characteristics and present rates applying to the median and 90th-percentile families in a low- and high-benefit state.

[37] Housing benefits are clearly important for those who receive them; however, housing programs are not operated as entitlements (i.e., not everyone who meets eligibility criteria is entitled to benefits), and only around 30 percent of eligible families receive housing benefits (see, e.g., Casey, 1992). Thus, the Keane and Moffitt convention for housing overstates tax rates for the typical family.

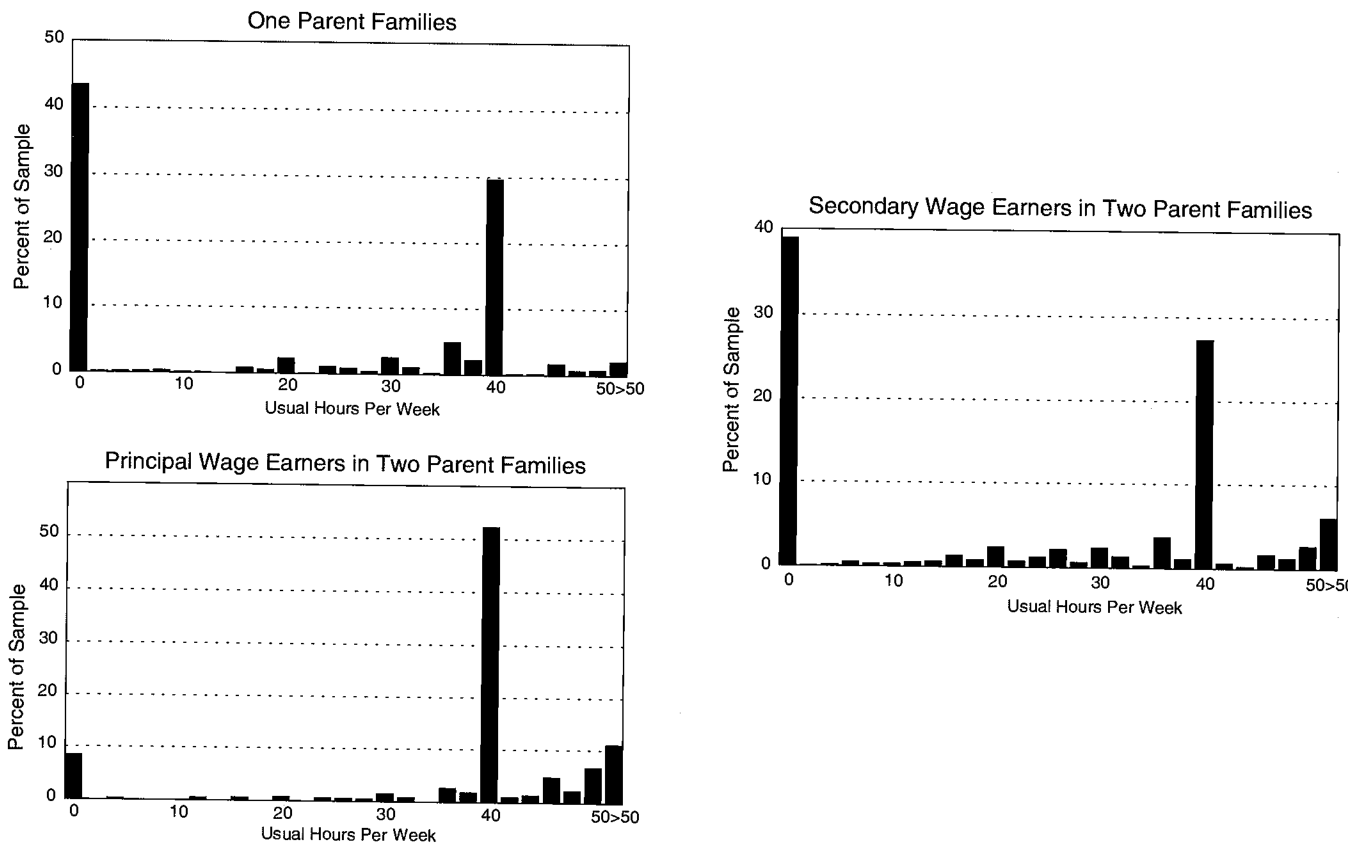


FIGURE 3. ***Distribution of Usual Hours Worked per Week by Person Type.***

for families in our sample. Among single parents who work in the paid labor market, 76.1 percent work 35 or more hours in the typical week. Similarly, of the secondary earners in two-parent families who work in the paid labor market, 74.5 percent work 35 or more hours in the typical week. Part-time work is even less common for employed primary earners, 92.6 percent of whom work 35 or more hours a week. As is also clear from the figures, hours in part-time jobs are distributed fairly uniformly between 20 and 35 hours.

Our data showing few families working part-time are consistent with other data. Commonly cited explanations for the paucity of part-time jobs focus on fixed labor-hiring costs to the employer. Because of these fixed costs, empirical models that treat the wage rate as being exogenous generally overpredict the number of people working part-time. One approach to addressing this problem is to model wages as increasing with hours of work (Moffitt, 1984; Rosen, 1976) or to model some process of rationing part-time jobs (Dickens and Lundberg, 1993). As a practical matter, the empirical distribution of hours leads us to follow most others in the literature (Moffitt, 1984; Fraker and Moffitt, 1988; Keane and Moffitt, 1991; Moffitt and Wolfe, 1992; Hoynes, 1993) who treat families as having three choices in the number of hours they can work: 0, 20, and 40. Thus, in our empirical work, the taxes and benefits families pay and receive generally will be based on the difference between working 0 and 20 hours.

4. THE EFFECTS OF WAGES, TAXES, AND TRANSFERS ON LABOR MARKET AND PROGRAM PARTICIPATION

The literature on taxes and labor supply focuses primarily on the intensive margin—the effect of taxes on hours of work for those in the labor market. In Table 2 we showed that estimates from this literature imply that the 1993 EITC expansion, when fully phased in, will have a modest negative overall effect on the hours of work among families already working, relative to the EITC in 1993. The negative effect occurs because nearly 80 percent of working families that receive the credit are in its flat or phaseout range, where the incentives are to reduce hours. The EITC is expected to have a beneficial effect on the extensive margin—the decision to work for those not in the labor market. Existing studies of the EITC's effect on labor market behavior ignore labor market participation. The broader literature on the effects of taxes and transfers on labor market behavior also provides little guidance for understanding the EITC's likely effects on labor market participation.

In this section we examine the responsiveness of labor market and transfer program participation to changes in after-tax wages, controlling for transfer benefits, demographic characteristics, and other factors likely to affect labor market and transfer program participation.

4.1 Empirical Model

To study the effect of wages, taxes, and program benefits on labor market and program participation, we follow Moffitt and Wolfe (1992) and estimate bivariate probit models of labor market and transfer program participation.[38] We include variables for incomes, transfers, and demographic characteristics in both participation equations. The models are estimated separately for one-parent families and primary earners in two-parent families. We also estimate a single-equation probit model for the labor market participation decisions of secondary earners in two-parent families.

Income, tax rates, and benefits depend on labor market and transfer program participation decisions, which leads to a potential endogeneity problem with these policy variables. To circumvent this problem we simulate families' taxes and benefits if they work 0 hours and if they work 20 hours at their market wage rates. Explicit tax rates are then defined as one minus the change in after-tax income divided by gross earnings.

The main effect of the EITC on participation will be through its influence on net wages. We include net wages in the empirical model, defined as gross wages (predicted wages for those not in the labor market) multiplied by one minus the explicit tax rate. We expect that net wages are positively correlated with labor market participation and negatively correlated with transfer program participation. In sensitivity analyses we examine the robustness of our results to (1) using the tax rate that would arise from moving from 0 to 40 hours of work, and (2) using predicted wages for all observations in the sample rather than for only those people not in the labor force.

We examine the effects of program benefits by including the AFDC and food stamp benefits available to a family if members work 0 hours. We expect benefit guarantees to show a negative relationship to labor market participation and a positive relationship to program participa-

[38] To make inferences about behavioral effects from cross-state variation in benefit and tax rules, it is important that households do not make location decisions based on these benefits and taxes. If households with unobserved "tastes" for work systematically locate in low-benefit, low-tax states, we would observe a spurious negative correlation between taxes and benefits on one hand and hours (or labor market participation) on the other. In a careful study using the 1980 Census, Walker (1994) finds little evidence of welfare-induced migration.

tion. We include family income at 0 hours of work, which is primarily capital income. Income at 0 hours is expected to have a negative effect on both labor market and transfer program participation.

We also include a number of other variables that are common to studies of labor market and transfer program participation. Age and age squared are included to capture life-cycle effects that might affect labor market or program participation. We include years of education, family size, and dummy variables for nonwhite, female, a self-reported measure of being in poor or fair health, presence of children under 6 years old, presence of children aged 6–12, and region of the country. The dependent variable in the labor force participation model takes the value 1 if the person is usually employed during the sample month. The dependent variable in the transfer program model takes the value 1 if the family reports participating in either the AFDC or the food stamps program during the sample month.

4.2 Empirical Results

Empirical results for our primary specifications are given in Table 4. The primary variable of interest for this study is the effect of the net wage rate on labor market and program participation. For single parents (the first two columns of Table 4), we find that net wages positively affect labor market participation and negatively affect transfer program participation. Both coefficients are significant at typical levels of confidence. As shown in the following section, where we summarize the results through policy simulations, the economic significance of both estimates is fairly large. In particular, a 10-percent increase in the after-tax wage (a \$0.61 increase) raises the single parent's probability of working by two percentage points and lowers the probability of transfer program participation by more than four percentage points, holding other variables in Table 3 at their means (and dummy variables at 0, except for female). Net wage takes the expected signs in the specifications for two-parent families, but it is not significant in the labor market decisions of primary earners. It is strongly significant in the transfer program participation regressions where a 10-percent increase in the net wage implies a reduction of 0.7 percentage points in the probability of participating (recall that relatively few two-parent families receive transfers). These results imply that policies, like the EITC, that alter the after-tax wage rate can substantially increase labor market participation and reduce transfer program participation.

The regression evidence showing strong effects of net wages is not an artifact of our particular empirical specification, but it clearly emerges in the underlying data even when we do not condition on other factors. In

TABLE 4.
Bivariate Probit Estimates of Labor Market and Program Participation

Variable	Single parents		Principal wage earners		Spouses
	Labor market	*Transfers*	*Labor market*	*Transfers*	*Labor market*
Constant	−0.833	−0.646	3.941	0.658	−1.871
	(0.863)	(0.937)	(1.805)	(1.461)	(0.747)
Net income at 0 hours (1,000s)	−0.982	−0.241	−3.232	0.102	−1.042
	(0.497)	(0.443)	(0.751)	(0.535)	(0.385)
Net wages (in 1,000s)	89.95	−195.3	78.23	−219.49	57.83
	(34.56)	(31.47)	(57.46)	(43.74)	(23.61)
AFDC at 0 hours (in 1,000s)	−1.986	1.739	−1.812	1.086	−1.394
	(0.435)	(0.406)	(0.675)	(0.409)	(0.538)
Food stamps at 0 hours (1,000s)	0.423	−1.064	−2.904	0.895	−0.206
	(1.326)	(1.291)	(1.936)	(1.06)	(0.637)
Female	−0.600	1.039	−1.551	0.159	−0.757
	(0.258)	(0.239)	(0.238)	(0.189)	(0.143)
Age	0.055	0.029	−0.093	−0.078	0.130
	(0.047)	(0.048)	(0.094)	(0.075)	(0.040)
Age squared	−0.00076	−0.00026	0.00076	0.00106	−0.00182
	(0.00065)	(0.00066)	(0.00119)	(0.00099)	(0.00053)

Education	0.061	−0.054	−0.020	−0.052	0.057
	(0.025)	(0.028)	(0.039)	(0.030)	(0.019)
Nonwhite	−0.066	0.229	−0.175	0.297	−0.124
	(0.116)	(0.125)	(0.23)	(0.175)	(0.115)
Bad or poor health	−0.592	0.572	−0.246	0.453	−0.533
	(0.141)	(0.154)	(0.290)	(0.228)	(0.159)
Kids under 6	−0.413	0.598	−0.255	0.247	−0.321
	(0.096)	(0.114)	(0.149)	(0.109)	(0.079)
Kids from 6 to 12	−0.122	0.258	−0.313	0.076	−0.035
	(0.088)	(0.097)	(0.136)	(0.106)	(0.072)
Family size	0.050	0.059	0.339	0.141	−0.049
	(0.105)	(0.102)	(0.175)	(0.091)	(0.061)
South	0.252	−0.227	0.702	−0.059	0.204
	(0.171)	(0.170)	(0.240)	(0.247)	(0.135)
Midwest	0.217	−0.382	0.183	0.138	0.144
	(0.177)	(0.185)	(0.283)	(0.260)	(0.159)
West	0.423	−0.511	0.613	−0.695	0.201
	(0.191)	(0.205)	(0.283)	(0.286)	(0.167)
Correlation of errors	−0.324		−0.083		
	(0.072)		(0.125)		

Note: The spouse's equation includes primary earner's income, which equals 0.101 (standard error of 0.120).

the top panel of Figure 4, we classify one-parent families by their predicted wage rate and then plot their probability of working. Within each wage group, we allocate families based on their quartile of the cumulative tax rate distribution when increasing hours from 0 to 20. If one looks across wage groups, it appears that the probability of working generally rises with wage rates, though the relationship is not monotonic. Within wage groups, the probability of working falls sharply with our exogenous measure of tax rates. Even after crudely controlling for human capital by looking within predicted wage categories, tax rates appear to exert a strong negative effect on the probability of labor force participation.

The labor force participation rate of primary earners in two-parent families is very high. In the bottom panel of Figure 4, therefore, we graph the effects of wages and taxes on the labor force participation of secondary earners in two-parent families. As with single parents, the probability of working generally increases with predicted wage rates. There is a weaker negative relationship between tax rates and labor market participation for secondary earners than there is for single parents. This result is somewhat surprising given that the literature suggests two-parent families are more responsive to economic variables than are single-parent families. However, the figures do not take into account many factors that are likely to affect labor market behavior.

The AFDC benefit guarantee appears to exert an economical and statistically significant effect in our regression specification of labor market and transfer program participation decisions. The influence of the AFDC variable presumably arises both from its size and from its influence on the benefit reduction rate. As is clear when one examines tax rates in Texas (a low-benefit state) and New York (a high-benefit state), there is a close positive relationship between the guarantee and the benefit reduction rate. The empirical estimates show a consistent negative relationship between the AFDC benefit guarantee and labor force participation, and a consistent positive relationship between the guarantee and program participation. For single-parent families, a 10-percent increase in the benefit guarantee implies a 1.65-percentage-point reduction in the probability of working and a 1.45-percentage-point increase in the probability of receiving program benefits. The economic significance of the effects is much smaller for two-parent families. As with wage rates, the estimated effects of benefits from the empirical model are also clearly present when we graph program or labor force participation against benefits, holding wages constant. (The figure is not shown.)

Several other coefficients generally have consistent patterns across family types and are economically and statistically significant influences

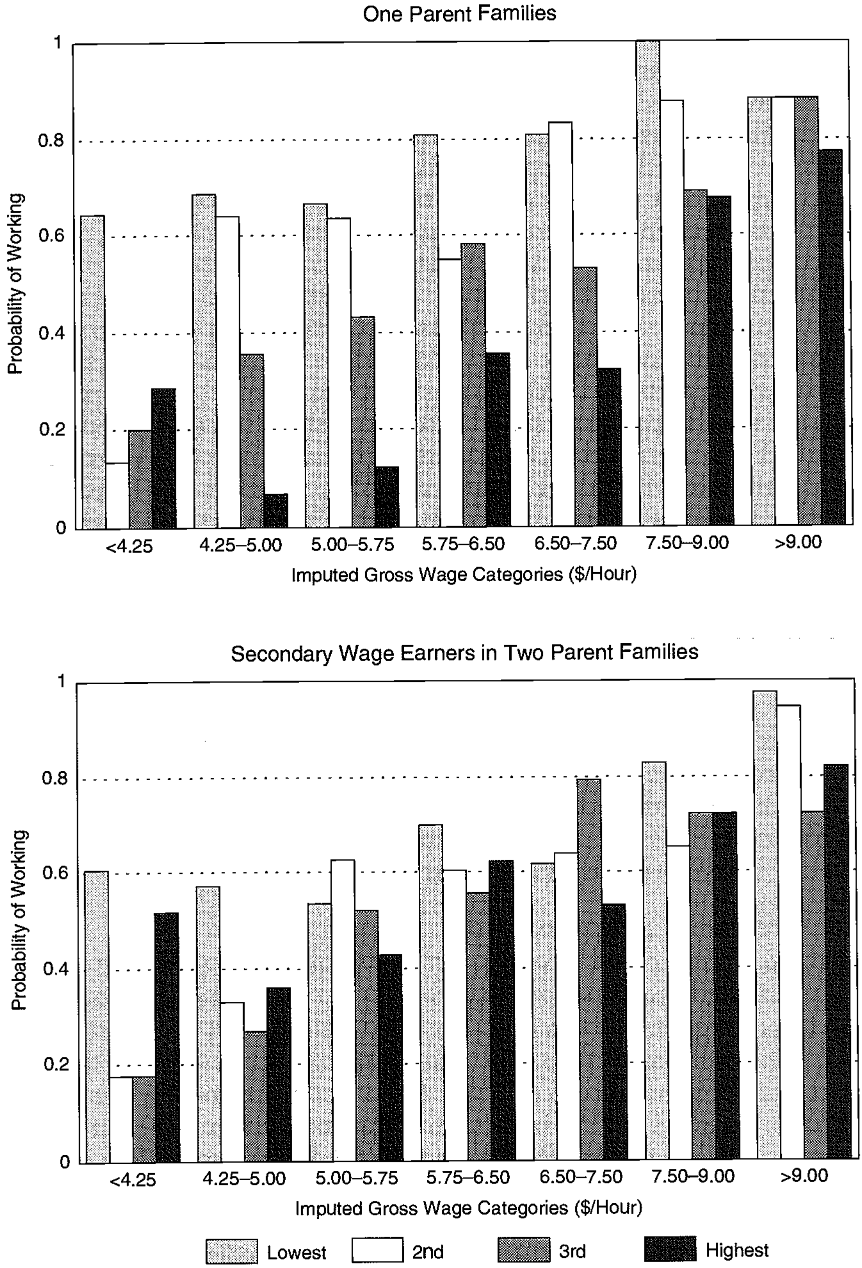


FIGURE 4. ***Labor Force Participation by Wages and Tax Rates.***

on labor force and transfer program participation. The coefficient on net private income at 0 hours is negatively related to labor market and transfer program participation for single parents, but it is statistically significant only in the labor market participation equations. The economic significance of the other income variable is small. A 10-percent increase in other income, for example, reduces the probability of work by 0.13 percentage point for secondary earners in two-parent families. Labor force participation is generally negatively correlated with being female, in poor or fair health, and with the number of children in the family under six years of age, while these characteristics are positively correlated with transfer program participation (conditioning on other characteristics). Being in poor or fair health has an economically large effect, lowering the probability of working by 23 percentage points for single parents and 21 percentage points for secondary earners. It raises the probability of participation in transfer programs by 22 percentage points for single parents holding other variables at their means (and other dummy variables at 0 except for female). Whites, more highly educated persons, and individuals living in the South and West tend to have higher labor force participation rates and have lower rates of transfer program participation than others.

In the bivariate probit model for single-parent families, we estimate a highly significant negative correlation between the labor market and transfer program participation equations. The negative correlation implies that unobserved factors, such as parental background variables, that affect the probability of working are negatively correlated with the probability of participating in transfer programs. The correlation of the error terms in the two-parent primary earner specification is also negative but is much smaller and not significantly different from zero.

4.3 Alternative Specifications

In all the specifications reported in Table 4, we use the observed market wage rate for people in the labor force. A common alternative is to use predicted wages for all people in the sample (Blank, 1985; Hoynes, 1993). Using the predicted rather than observed wage nearly doubles the estimated effect of net wages on labor market participation for one-parent families, and the coefficient remains statistically significant at the 5-percent level. The wage effect more than doubles for secondary earners in two-parent families, and it is also significant at usual levels of confidence. The wage effect flips sign for primary earners, but the coefficient is imprecisely estimated. We conclude that using predicted wages would make the effects of the EITC on labor market participation considerably larger than those based on estimates from Table 4.

The net wage effect is again larger when we calculate the exogenous tax rate measures assuming people increase hours to 40 from 0, in this case by roughly 35 percent. The 40-hour tax rates are higher than the 20-hour rates for 60 percent of the sample, who generally live in high-benefit states. It is not clear which measure of tax burdens is preferable, but, like our treatment of wages, the net wage effects in Table 3 are lower than they would be in alternative, plausible specifications.

We also estimated the bivariate probit equations including benefit reduction rates in the specification, along with all other covariates. The benefit reduction rates are defined as the ratio of changes in AFDC or food stamps benefits to the change in earnings between 0 and 20 hours of work. Since families living in high-benefit states have more AFDC at 0 hours and, thus, a longer clawback range, our definition of benefit reduction rates is highly correlated with the benefit at 0 hours of work. The addition of the benefit reduction rate causes the coefficient on AFDC benefits at 0 hours to flip signs and lose significance. In this case, the net wage effects are lower than those in Table 4 for single parents but higher for both principal and secondary earners in two-parent families. We have little confidence that this specification adequately disentangles the independent effects of benefit guarantees and benefit reduction rates.

The labor market effects of the EITC depend on changes at the intensive and extensive margin. When we estimate hours equations like equation (1), adjusting for selection, we get small wage and income elasticities, though the parameters are estimated imprecisely. Our hours estimates are similar to others in the existing literature, which suggests that our data are consistent with data used in other studies of labor market behavior.

5. POLICY IMPLICATIONS

To help interpret the economic significance of the coefficient estimates, we use our regression results and labor market simulations to estimate the effects of the OBRA93 EITC expansion on labor force participation. As with Table 2, we simulate the effect of the 1996 EITC, when the increase is fully phased in, relative to the law that applied in 1993. We first model the effect of the expanded EITC on net-of-tax wages and calculate the implied change in the probability that individuals work. For people not in the labor force, we calculate the effect of the EITC on their after-tax and after-transfer wage in both 1993 and 1996, assuming that if they enter the labor force, they would work 20 hours per week.[39]

[39] As with our earlier simulations, we calculate the changes in net wages and virtual incomes that arise from the EITC for secondary wage earners, holding the hours of the primary earner fixed at their observed value.

The simulations show that the EITC increases the net wage of single parents by 15 percent. The higher net wage increases their probability of working by 3.3 percentage points.[40] If each of these single parents works an average of 20 hours per week for 20 weeks per year, our simulation implies that the hours of single-parent families would increase by roughly 72.8 million hours per year. The EITC expansion increases the mean net-of-tax wage of primary earners in two-parent families by 19.6 percent. This leads to a much smaller 0.7-percentage-point increase in labor force participation because most primary earners in two-parent families are already working. Our simulation implies that primary wage earners entering the labor force because of the EITC expansion will work about 12.1 million hours, again at an average of 20 hours per week for 20 weeks per year.

A feature of the EITC that has received little attention is that the average net wage of secondary wage earners decreases by 5.0 percent because the earnings of the second worker frequently either moves the family into the clawback range of the credit or makes the family ineligible for the EITC. Therefore, we expect secondary workers to reduce their hours of work by roughly 10.4 million because of their lower mean net wages. Overall, our simulation results imply that greater labor market participation will lead to an increase of 74.4 million hours, given our assumptions that new labor market participants will work 400 hours per year.

The increased hours resulting from higher rates of labor force participation can be compared with the reduction in hours caused by the credit shown in Table 2. We use Triest's (1990) labor supply parameters for the simulation. As is evident from Table 2, most EITC recipients in our sample are concentrated in the flat and clawback ranges of the credit and their reduction in hours is larger than the increase for those in the subsidy range. The results from Table 2 imply that single parents will reduce their hours of work by 26.4 million, primary earners in two-parent families by 13.6 million, and secondary earners by 14.5 million.

Together, the simulations suggest that the aggregate reduction in hours supplied by working households, 54.5 million, would be more than offset by the hours of new entrants, 74.4 million, if new labor force participants work an average of 20 hours per week for 20 weeks per year. If new labor market entrants work less, the "participation effect" will be smaller. If they work more, the participation effect will be larger. If

[40] We have not explicitly modeled Medicaid. Over time, a new labor force participant could, depending on income and family characteristics, lose Medicaid benefits, which presumably will inhibit labor market participation. Further work examining the degree to which incorporating Medicaid would alter our results would be worthwhile.

estimates of participation elasticities with respect to wages from the alternative specifications described in Section 4.3 were used, our estimated offset would also be somewhat higher. For example, using imputed wages for all persons implies that new labor market participants will increase hours of work by 108.4 million. Replacing the 20-hour average tax rate with the tax rate from 0 to 40 hours implies that the new workers will work 91.8 million hours.

We also use our empirical model to simulate the effects of the 1993 EITC expansion on participation in transfer programs. The results are summarized in the lower panel of Table 5. The 15-percent increase in net-of-tax wages for single parents implies that transfer program participation among this group will decrease by 7.2 percentage points or that almost 400,000 families will no longer participate either in AFDC or in the food stamps program. The increased net wages of primary wage

TABLE 5.
Labor Market and Transfer Program Effects of the OBRA93 EITC Expansion, 1993 to 1996

	Labor market effects			
	New labor force participants		Families in the labor market	
	Percent change in net wage	*Annual hours change due to labor force participation (million)*[a]	*Annual hours reductions of workers (million)*[b]	*Average annual reduction in hours*
Single parents	15.0	72.8	26.4	10.1
Primary wage earners	19.6	12.1	13.6	7.7
Secondary wage earners	−5.0	−10.4	14.5	30.3
Total		74.4	54.5	11.2

	Transfer program participation effects		
	Number leaving program	*Mean annual benefit*	*Mean EITC payment*
Single-parent families	398,384	$6,844	$2,040
Two-parent families	117,757	$4,702	$2,842

[a] The estimation of the change in hours from new labor force participation assumes that, on average, these persons work 20 hours per week for 20 weeks per year.

[b] The reductions in hours for workers are simulated using the kinked budget set approach described in Section 2.1 and in Table 2. We use the estimated net wage and virtual income elasticities from Triest (1990).

earners imply a reduction in transfer program participation of roughly 117,000 families. The mean annual benefits (in 1994 dollars) for single-parent families are $6,844 and $4,702 for two-parent families, and the mean EITC payments for these family types are $2,040 and $2,842, respectively. Therefore, our simulations imply that a potentially substantial savings in transfer payments could result from the EITC expansion.

6. CONCLUSION

Over the past 20 years the EITC has been a favored policy tool for assisting low-income families with children. Between 1978 and 1996 the maximum credit available to families with children will have increased nearly 800 percent in nominal dollars. No other major program directed toward low-income families has grown at a comparable rate. The EITC is now the cornerstone of the Clinton administration's welfare reform agenda.

The effectiveness of the EITC will depend, in part, on its effect on labor market behavior. Most workers that will receive the credit have incomes in the flat or phaseout range of the credit, where the credit provides an unambiguous incentive for people to work fewer hours. Using recent estimates from the empirical literature on taxes and labor supply, we find that the change in incentives caused by the 1993 expansion of the credit is expected to lead to a modest reduction in hours of work by those in the phaseout range of the credit. When evaluated over all workers that could receive the credit, our central estimate predicts an overall reduction of 0.54 percent in hours of work.

No EITC labor market study examines the effect of the credit on labor market participation, though it is through this dimension—"making work pay"—that the EITC appears to attract its favored status. The effect of the credit on labor market participation may be large. As Heckman (1993, p. 118) writes, "A major lesson of the past 20 years is that the strongest empirical effects of wages and nonlabor income on labor supply are to be found at the extensive margin—the margin of entry and exit—where the elasticities are definitely not zero."

Before estimating participation rates, one needs to characterize the tax environment families face. If the cumulative tax burdens faced by the poor are very high, even the 40-percent wage subsidy that the EITC will offer to low-income taxpayers with two or more children may not be enough to make work an economically attractive option relative to welfare. We use a detailed microsimulation model to characterize the tax rates associated with labor market decisions for each family in our sample. Using data from the Survey of Income and Program Participation, we calculate tax rates faced by low-income families. We show that both marginal and average tax rates can be very high (over 90 percent) for high-tax households in both high- and low-benefit states. At the same time, because of earnings disregards and other features of the tax transfer system, tax rates on families in both high- and low-benefit states are often considerably lower than this high-tax case.

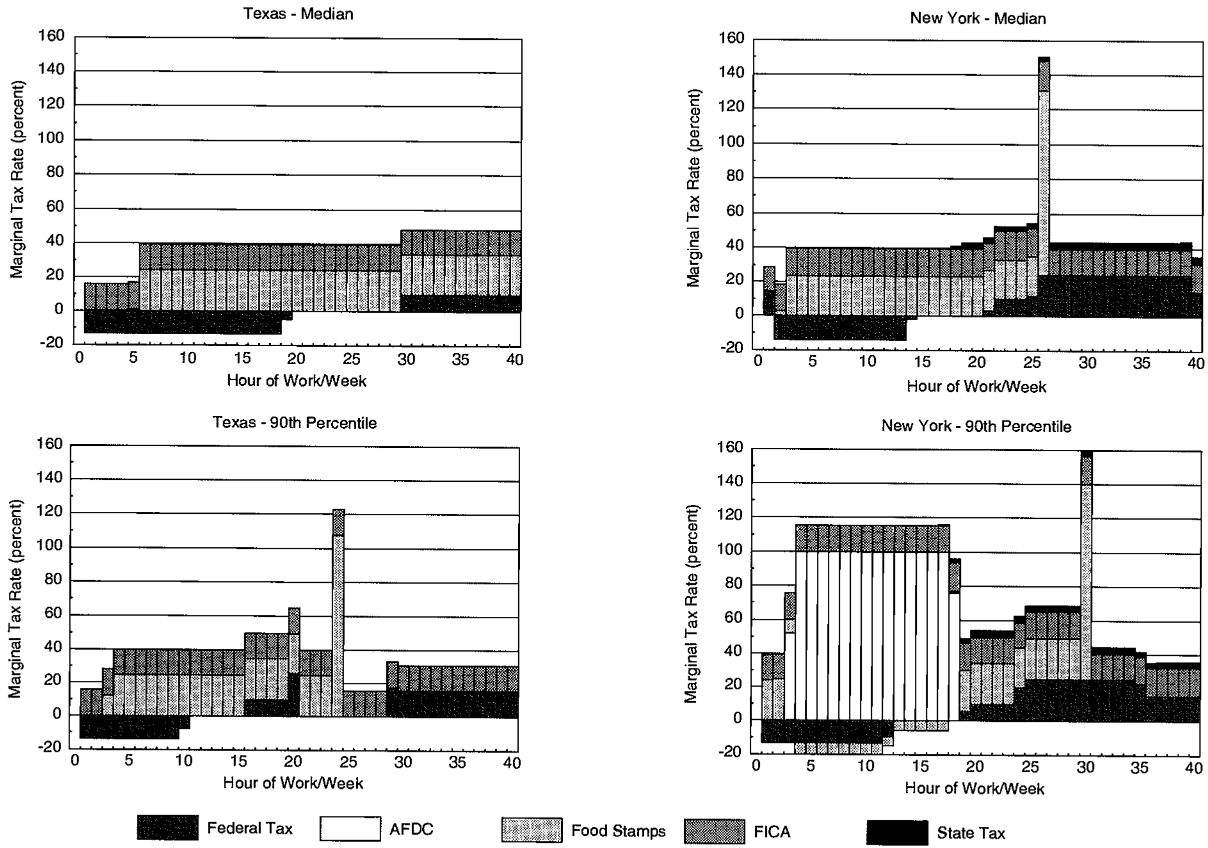
We find, both in descriptive and empirical models, that the after-tax wage has an economically and statistically positive effect on labor market participation and a negative effect on transfer program participation. Our results imply that when fully phased in, the 1993 EITC expansion will increase labor force participation rates by 3.3 percentage points for single parents in our sample, increase participation of primary earners in two-parent families by 0.7 percentage point, and decrease participation of secondary earners in two-parent families. When we simulate the overall effect on hours, assuming new participants work 20 hours per week for 20 weeks, the increase in labor force participation more than offsets the adverse effects of the EITC on hours supplied by low-income individuals. If new labor market participants work fewer total hours, the beneficial labor market effects of the credit will be smaller. Regardless of the precise estimates that are favored, it is clear that the participation effects of the credit can substantially offset, in aggregate, the negative effect of the credit on the labor supplied by people already in the labor market. Future work on the effect of the EITC on labor supply should incorporate the credit's effect on participation.

APPENDIX TABLE 1.
Means and Standard Deviations for Variables Used in Table 4

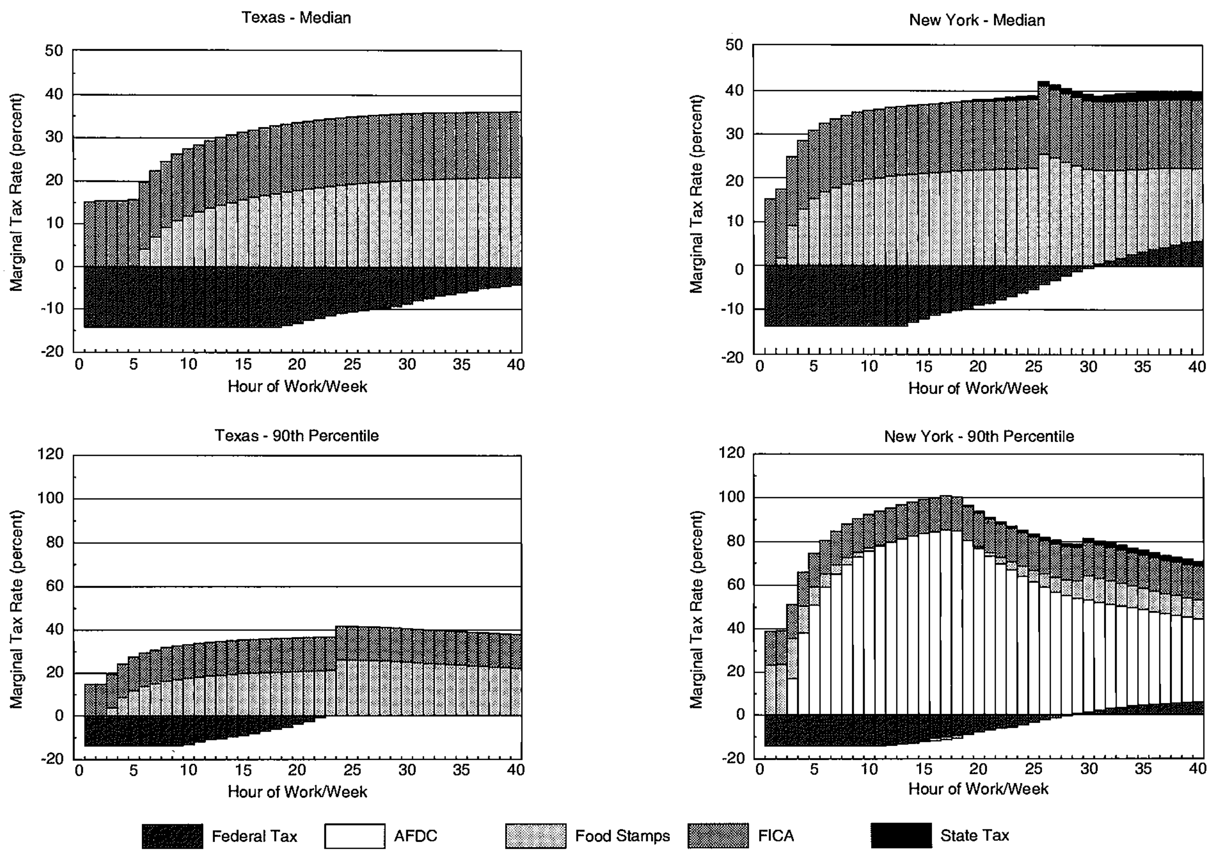
Variable	Single parents		Principal wage earners		Spouses
	Labor market	*Transfers*	*Labor market*	*Transfers*	*Labor market*
Dependent variable	.564	.505	.916	.094	.609
	(.496)	(.500)	(.277)	(.291)	(.488)
Net income at 0 hours	0.092	0.092	0.018	0.018	0.033
(1000s)	(0.189)	(0.189)	(0.103)	(0.103)	(0.124)
Net wages (in 1000s)	0.006	0.006	0.008	0.008	0.005
	(0.002)	(0.002)	(0.003)	(0.003)	(0.004)
AFDC at 0 hours (in 1000s)	0.228	0.228	0.088	0.088	0.014
	(0.240)	(0.240)	(0.237)	(0.237)	(0.096)
Food stamps at 0 hours	0.192	0.192	0.314	0.314	0.099
(1000s)	(0.084)	(0.084)	(0.102)	(0.102)	(0.135)
Female	0.945	0.945	0.260	0.260	0.740
	(0.229)	(0.229)	(0.439)	(0.439)	(0.439)
Age	32.832	32.832	34.654	34.654	33.541
	(8.081)	(8.081)	(7.444)	(7.444)	(7.628)
Age squared	1143.100	1143.100	1256.300	1256.300	1183.100
	(578.980)	(578.980)	(563.600)	(563.600)	(561.920)

Education	11.528	11.528	11.749	11.749	11.730
	(2.396)	(2.396)	(2.847)	(2.847)	(2.865)
Nonwhite	0.445	0.445	0.206	0.206	0.200
	(0.497)	(0.497)	(0.404)	(0.404)	(0.400)
Bad or poor health	0.174	0.174	0.071	0.071	0.089
	(0.380)	(0.380)	(0.257)	(0.257)	(0.285)
Kids under 6	0.635	0.635	0.775	0.775	0.775
	(0.813)	(0.813)	(0.830)	(0.830)	(0.830)
Kids from 6 to 12	0.709	0.709	0.814	0.814	0.814
	(0.800)	(0.800)	(0.846)	(0.846)	(0.846)
Family size	3.008	3.008	4.283	4.283	4.283
	(1.164)	(1.164)	(1.159)	(1.159)	(1.159)
South	0.398	0.398	0.495	0.495	0.495
	(0.490)	(0.490)	(0.500)	(0.500)	(0.500)
Midwest	0.231	0.231	0.194	0.194	0.194
	(0.422)	(0.422)	(0.395)	(0.395)	(0.395)
West	0.158	0.158	0.152	0.152	0.152
	(0.365)	(0.365)	(0.360)	(0.360)	(0.360)

Note: The spouse's equation includes primary earner's income divided by 1000. The mean is 1.114, and the standard deviation is 0.645.



APPENDIX FIGURE 1. ***Marginal Tax Rates for Low-(TX) and High-(NY) Benefit States—Principal Wage Earners in Two-Parent Families.***



APPENDIX FIGURE 2. *Average Tax Rates for Low-(TX) and High-(NY) Benefit States—Principal Wage Earner in Two-Parent Families.*

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