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# The Effect of Child Care Costs on the Employment and Welfare Recipiency of Single Mothers

Rachel Connelly* and Jean Kimmel†

This paper considers the effect of child care costs on two labor market outcomes for single mothers—whether to work for pay and whether to receive welfare. Hourly child care expenditures are estimated using data drawn from the 1992 and 1993 panels of the Survey of Income and Program Participation (SIPP). These expenditures are then used to predict the probability of welfare recipiency and employment. While the direction and significance of key variables are robust to changes in specification, the quantitative results are found to be sensitive to identification restrictions. All results show a substantial positive effect of child care costs on welfare recipiency, with the child care price elasticity of welfare recipiency varying from 1.0 to 1.9. Similarly, we find a significant negative effect of child care price on employment with elasticity estimates from $-.3$ to $-1.1$, showing that controlling for the welfare choice does not reduce the price elasticity of employment found in other studies.

## 1. Introduction

For all mothers of young children, entering the labor market is strongly linked with the need for child care. Opportunities for caring for children while in the labor market are few in a developed economy. In many cases, the husband or another family member serves as caregiver, but approximately 50% of preschoolers with a working mother are cared for by nonrelatives (Casper 1997). Some of these arrangements involve a substantial amount of money. In 1993, the average weekly cost of care was $59 for home-based care, $68 for center-based care, and $48 for care provided by a relative. This can represent one-fourth of earnings for single mothers working full time at the minimum wage (Kimmel 1994). Such substantial money expenditures, coupled with transportation needs both to work and to day care, as well as the uncertainty of many child care arrangements, are expected to keep many mothers of young children out of the labor market. Thus, the relationship between employment and child care for these mothers is thought to play a strong role in the link between welfare recipiency and child care.

Welfare programs before and after welfare reform have targeted child care as a barrier to employment.[1] Before welfare reform, child care subsidies were available to some recipients through

* Department of Economics, Bowdoin College, Brunswick, ME 04011, USA; E-mail connelly@bowdoin.edu.

† Department of Economics, Western Michigan University, Kalamazoo, MI 49008, USA; E-mail jean.kimmel@wmich.edu; corresponding author.

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[1] See Blau (2000) for a comprehensive discussion of child care subsidy programs.

federal Title IV-A funding sources for child care (AFDC/JOBS, At-Risk, Transitional Child Care) and through the Child Care Development Block Grant. These funds often came with matching requirements from the states. The Personal Responsibility and Work Opportunity Reconciliation Act of 1996 (PRWORA) consolidated all these funds into state block grants, thereby permitting the states to design their own child care assistance schemes. States may supplement federal child care block grants with state dollars, but there is no longer a required state match. Thus, while the total federal dollar amount allocated to child care in Temporary Assistance for Needy Families (TANF) exceeds former federal Aid to Families with Dependent Children (AFDC) child care commitments, because TANF requires less in state matching expenditures, it is unclear what will happen to total child care expenditures as welfare reform evolves. Early postreform evidence suggests that while overall child care spending at the state level has increased, the increase is less than would have occurred had the matching requirements been retained. A recent study of welfare leavers reports that few are receiving subsidies (Schumacher and Greenberg 1999), and only 1.24 million of the approximately 10 million children eligible for federally funded support received assistance in 1997 (U.S. Department of Health and Human Services 1999).

Underlying states' expenditures on child care subsidies are their subsidy eligibility guidelines, participation in such subsidy programs by the eligible population, and availability of subsidized slots or funds for those families applying for such funds. Only a small percentage of families eligible for subsidies based on the federal maximum income limits receive such support. Federal guidelines as outlined in PRWORA stipulate that federally financed child care subsidies can be made available to families with incomes up to 85% of the state's median income. However, as of July 1999, only five states had set their eligibility guidelines at the federal maximum. In addition, participation by the state-defined eligible group is quite low, partially because of a lack of information. City officials in San Francisco have used an innovative peer outreach program to increase participation by the eligible population, and by the start of 2000, the city was enrolling 50% of the estimated eligible population, an enrollment rate twice the statewide average (Heymann 2000b).

Extensive data on post-TANF behavior are not yet available, nor will they be for some time. However, there is some evidence that workers continue to report that availability and cost of child care are barriers to self-sufficiency. For example, the McKnight Foundation's recent survey found that 18% of employers report that their welfare-to-work workers face child care problems (Heymann 2000a).

This paper looks back to the relationship between AFDC recipiency and child care costs using data from the second half of 1994. It is offered not as a historical footnote but rather because child care costs will continue to be an important factor determining welfare participation in the post-welfare reform environment because of the low expected earnings of low-skilled workers and the high percentage of earned income that must be devoted to purchase reliable quality care. In addition to facilitating mothers' employment and thus reducing poverty and the need for income supplements, quality child care is also an important social concern in and of itself, given the strong link between quality child care and positive child outcomes, particularly for at-risk children. Finally, these data come from early in the 1990s' economic expansion and thus represent a more diverse population of welfare recipients than more recent data would contain. Later in the 1990s, after the economic expansion broke historical records, state welfare caseloads had fallen so substantially (because of both welfare reform and the unusually strong economy) that the remaining caseload is overrepresented by hard-to-place individuals with multiple (hard-to-quantify) barriers to employment (see, e.g., Council of Economic Advisers 1997; Ziliak et al. 2000). The earlier data permit the estimation of a link between child care costs and welfare recipiency that is likely to be observed in future periods of more typical moderate economic expansion or contraction.

In this paper, we measure the effectiveness of child care assistance policies indirectly by considering explicitly the effect of the cost of child care on welfare recipiency. We find that, over a set of alternative specifications, AFDC recipiency and employment of single mothers are sensitive to the predicted hourly price of child care. The elasticity of recipiency with respect to the predicted price of child care is sensitive to the specification of the final model ranging in value from 1.01 to 1.94 once the jointness of AFDC recipiency and employment are considered. The elasticity of employment with respect to the predicted price of child care is less sensitive to the specification and estimated to be between −0.32 and −0.42, which is similar to what other studies of single mothers have found. Finally, simulations of child care subsidies show that substantial declines in AFDC recipiency and increases in employment could be achieved with modest means-tested child care subsidies available to all single mothers.

We begin with a summary of evidence concerning the importance of child care costs in the determination of welfare recipiency available from welfare-to-work programs as well as a summary of the existing econometric evidence on this issue. Then we summarize a theoretical model of employment and welfare recipiency and estimate the model using data from 1994 obtained by merging overlapping interviews from the 1992 and 1993 panels of the Survey of Income and Program Participation (SIPP). Finally, we discuss policy simulations designed to enumerate more clearly the importance of child care costs to the welfare population.

## 2. Review of Existing Evidence

There are three main sources of information related to our research question on the effect of the price of child care on employment and welfare recipiency. The first source is a large body of econometric work on the effect of child care costs on employment. Much of that literature focused on married women, but a few more recent papers have highlighted differences between married and single mothers. Second is a much smaller set of papers focused on the welfare side of the coin. Finally, there is some evidence from evaluations of welfare-to-work demonstration projects of the importance of child care costs to employment and welfare recipiency.

In terms of the econometric work on the effect of child care costs on employment, that body of work has been well summarized elsewhere (see, e.g., Anderson and Levine 1999; Blau 2000). This collection of research includes the early work by Heckman (1974) and the economics of child care revival of the late 1980s and early 1990s, which includes, for example, Ribar (1992). Almost all the studies on employment find a significant negative effect of child care costs on women's employment, although the estimated child care price elasticity with respect to employment varies widely across studies. Most relevant to our current topic are three papers—Han and Waldfogel (1998), Anderson and Levine (1999), and Connelly and Kimmel (in press)—each of which uses SIPP data from the early 1990 panels to look at differences across marital status. Each of these papers finds evidence that the elasticity of single mother's employment with respect to child care costs is greater in absolute value than married mother's employment elasticity.

The econometrics literature that focus on child care costs and welfare recipiency is more limited. Four papers using national databases are Connelly (1990), Kimmel (1995), Houser and Dickert-Conlin (1998), and Crecelius and Lin (2000). The first three use SIPP data similar to those in our analysis here. Crecelius and Lin use Panel Study of Income Dynamics (PSID) data. Connelly (1990) used the 1984 panel of SIPP and found a small effect of child care costs on welfare recipiency. Kimmel (1995) used a low-income subsample of a merged file from the 1987 and 1988 SIPP panels and found a nearly zero elasticity. Houser and Dickert-Conlin (1998) used 1993 SIPP data in a complex microsimulation model of labor market and transfer program participation, incorporating after-tax wages, transfer payments, and child care payments and examining married and single mothers separately (the former in order to discern secondary worker effects). Their simulations suggest that a 50% child care subsidy would increase the labor force participation of single parents by 2.9 percentage points and that a 20% reduction in the AFDC guaranteed payment would increase the labor force participation of single parents by 1.6% and reduce their welfare transfer program participation by 1.2 percentage points. These results, although in the same direction as our findings, are much smaller.

Crecelius and Lin's (2000) model also differs from ours in several ways. First, they estimate a joint model of employment/welfare participation that includes hours worked truncated at zero rather than an employment probit as we do. Previous child care studies have shown that the bulk of the behavioral "action" is in the discrete employment outcome rather than the continuous hours outcome. They find that for each 10-cent reduction in child care costs, there are 0.154 to 0.212 more hours worked per week.

Evidence of a positive relationship between child care costs and welfare recipiency can also be found in a number of evaluation studies of welfare-to-work demonstration projects, though the results are not uniform. Anderson and Levine (1999) reviewed evidence from several major welfare-to-work demonstration projects from the late 1980s and early 1990s that included child care components.[2] They wrote, "Although the confluence of services, mandates, and incentives in these demonstrations suggests caution is required in interpreting their results, based on this evidence it seems reasonable to conclude that subsidized child care may have a modest effect, at best, in increasing employment levels of very low-skilled, single mothers with small children" (p. 12). However, as the authors point out, none of these demonstrations explicitly examined the importance of child care costs within an experimental framework, so any conclusions relating to the importance of child care costs are tentative at best.

The Minnesota Family Investment Program (MFIP), which was included in Anderson and Levine's review, deserves extra scrutiny because new findings from the three-year follow-up study (conducted with a desirable experimental design based on random assignment into MFIP or AFDC) have now been released. This program was an innovative program based on the dual (and often competing) goals of encouraging work and making work pay. It contained two key work incentive provisions, the second of which related to child care. The MFIP paid child care costs directly to providers for all parents working or participating in employment-related activities. The AFDC reimbursement scheme differed because the parents paid the providers directly and were reimbursed later. According to the MFIP report summary (2000), the practice of reimbursing the mother after the expenditure occurred may have hindered the mother's efforts to get and stay employed. Also, the AFDC reimbursement rules tend to discourage providers from accepting such subsidized clients because of the uncertainty of receiving payment. The third-year follow-up report finds significant impacts in numerous areas, including employment rates and earnings of the MFIP approach.

Finally, Lemke et al. (2000) analyzed Massachusetts state data on current and former TANF recipients who also receive child care vouchers. They find that increased funding for child care subsidies and availability of full-day kindergarten are associated with increased probabilities that current and former welfare recipients will work.[3]

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[2] See also papers by Robins (1988), Joesch (1991), Berger and Black (1992), and Bowen and Neenan (1993). These papers are summarized in relation to the question posed here in Connelly and Kimmel (2001).

[3] This study has two serious limitations. First, only those currently receiving child care vouchers are included, making it difficult to draw conclusions about the importance of the availability of such vouchers in employment and training decisions. Second, the probit model of employment has, as its alternative to employment, participation in formal training or education programs rather than the broader category of nonemployment.

In sum, a thorough review of the broad literature relevant for this paper reveals a uniformity in the direction and significance of the child care price effect but a rather broad range of empirical estimates concerning the importance of child care costs on employment probabilities of single mothers. Less has been done in reference to welfare recipiency, but there, too, findings are consistent in the direction of the effect and differ substantially in terms of the magnitude. What are the likely sources of these disparate findings? First, equation specification matters (for an explicit focus on the importance of equation specification, see, e.g., Kimmel 1998). Without careful justification of equation specification and robustness checks, results could be unstable. Second, studies that rely on regional child care price data or complicated across-equation error structures (e.g., Blau and Hagy 1998; Tekin 2000) tend to produce smaller elasticities. On the other hand, studies (such as this one) that rely on predicting child care prices from individual characteristics tend to get larger elasticities. Since the intracity variation in child care expenditures are substantial and SIPP data constitute the only continuing national data set with child care price information, we believe that studies such as ours using individually generated child care prices should not be dismissed or their findings discounted. One of the most important aspects of the market for child care is that individuals face widely different costs for similar services depending on the availability of low- or no-cost child care options. Only individual based models take this variation into account systematically.

## 3. Underlying Theoretical and Econometric Models

We begin with a simple model of individual decision making from which equations can be derived that represent the discrete choices about welfare recipiency and employment of mothers with young children. In our model, we assume that mothers of young children seek to maximize their utility over goods and child services, subject to four constraints: a money budget constraint combining the mother's labor income and nonlabor income, a production function for child services, a mother's time constraint, and a child's time constraint. Child services are the commodity parents are consuming from their children; it could be companionship or love or pride in one's progeny. They are produced with a combination of the mother's time at home, the child's time with other caregivers, and money inputs. Total nonlabor income is the sum of family income from sources other than the mother's labor market participation and means-tied transfer income, such as welfare payments. Mothers have three uses of their time: work in the labor market, time spent with children, and leisure. The child has two types of time: time with the mother and time with a nonmaternal caregiver.

From this theoretical model, we derive the individual's indirect utility function that takes on two or four different values corresponding to the different possible work and welfare outcomes.[4] Based on the indirect utility function, we derive estimating equations for AFDC participation and employment in which both discrete dependent variables represent underlying continuous latent indices reflecting preferences for welfare recipiency and market work. Estimation of these equations using variants of the probit model produce estimates of the probabilities associated with employment and welfare recipiency.

Included among the factors affecting welfare recipiency and employment will be predicted child care expenditures, which are expected to be positively related to the probability of welfare receipt and negatively related to the probability of employment. Increased expenditures on child care lower

[4] See, for example, Blank (1985, 1989) and Crecelius and Lin (2000) for models employing this indirect utility approach to AFDC recipiency.

a woman's effective wage in the labor market when she is not receiving AFDC. Also included among these variables will be her predicted wage (proxying potential earned income), nonlabor family income, dichotomous variables indicating that the mother is nonwhite or unhealthy or lives in an urban area or in the South, factors affecting the value of a woman's time at home (specifically, two dichotomous variables indicating whether the youngest child is age zero to two years and whether there are two or more preschoolers in the family), the state's average Medicaid expenditures per enrollee, the state's average monthly AFDC payment, and the state's unemployment rate. We expect that the woman's wage will be negatively correlated with welfare receipt but positively associated with employment, while those variables that are positively correlated with the value of a mother's time at home, particularly the number of young children in the family, will have the opposite effects on both outcomes.

Estimating the welfare recipiency equation by itself will provide an initial look at the effect of child care costs on AFDC recipiency. However, estimating this equation alone ignores the interaction between AFDC recipiency and employment. Because of kinks in the budget line caused by AFDC regulations, as well as possible discontinuities in hours of employment and child care availability, it is reasonable to suspect that decisions about AFDC recipiency are made jointly with decisions to work for pay. In other words, the error terms in the two equations are correlated. Jointly estimating these two equations is accomplished by estimating a bivariate probit with four possibilities corresponding to the joint outcomes of AFDC recipiency, yes or no, and employed, yes or no. Estimates of the bivariate probit model refine our understanding of the effect of child care expenditures on both AFDC recipiency and employment of single mothers. In addition, use of the bivariate probit model produces more efficient estimates of the parameters and the standard errors.

## 4. Description of the Data

The sample of single mothers with children age five or younger used in this paper was drawn from a merged file from the 1992 and 1993 SIPP panels. The SIPP, which is conducted by the U.S. Bureau of the Census, is a large, nationally representative sample of households in the United States.[5] In these two panels, SIPP respondents are interviewed every four months for nine interviews, and a special set of child care questions are asked at the sixth interview of the 1992 panel, which overlaps the same calendar time period as the third interview of the 1993 panel. In these overlapping child care interviews, which took place in the second half of 1994, currently employed respondents with children younger than six were asked a number of detailed questions regarding their child care utilization patterns and expenditures. Mothers of such young children are subject to strongly binding child time constraint; that is, these children must be cared for 24 hours of the day by either a parent or a nonparental child care provider. Thus, while some child care costs are also associated with older children, the labor market decisions of mothers with young children are the mostly likely to be affected by the costs of child care.

Using the detailed labor force information from the fourth month of the wave, each mother is defined as employed if she reports positive earnings, hours, and weeks worked. The hourly wage is defined as monthly earnings divided by monthly hours worked. Finally, welfare recipiency equals one if the mother reports any AFDC recipiency during the fourth month of the wave.

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[5] The SIPP survey was designed to represent the noninstitutional population of the United States. There was no oversampling in SIPP panels 1984 through 1993 except for the 1990 panel (see Nelson, McMillen, and Kasprzyk 1984; Kalton et al. 1999; and communication with Smanchai Sae Ung of the U.S. Bureau of the Census).

We added a set of state-based variables to the SIPP's individual-based information. These variables include the constructed dummy variables for urban residence (equals one if the mother lives in a standard metropolitan statistical area [SMSA]), and southern residence (equals one if the mother lives in the South). An additional set of state-based variables was added that includes information drawn from a variety of sources. These variables include the state's average Medicaid payment per enrollee, the state's average monthly AFDC payment, the state's unemployment rate, the state's regulated child:staff ratio of less than 10:1, the state regulated center teachers' education, state per capita income, and, finally, the employers' estimated workers' compensation payment by state.[6,7]

Table 1 presents the mean values of the variables included in the analysis for five categories of single mothers: all single mothers, those employed, those employed and paying for child care, single mothers receiving welfare payments, and single mothers not receiving welfare payments. Table 2 provides a more detailed breakdown of variable means using subgroups stratified by both welfare and employment status, which is the specific focus of this paper. First looking at Table 1, we see that 43% of the 1523 women in our full sample are welfare recipients. Thirteen percent of the welfare recipients are employed in the labor market, while 73% of the nonrecipients are employed. In addition, AFDC recipients are slightly younger than nonrecipients (27.7 vs. 28.2 years old) and have, on average, 11.2 years of education—more than one year fewer than the nonrecipients. The AFDC recipients have more children aged zero to two and three to five, are more likely than nonrecipients to be nonwhite, and are considerably more likely to live in poverty.

Employed single mothers are 28.5 years of age, on average, and have 12.5 years of education. Only 26% live in poverty, but two-thirds have income less than twice the poverty threshold. Approximately one-fourth work part time, and 53% report paying for child care. The oldest single mothers are those who are employed and paying for child care, and this subgroup also reports the highest education levels, with 12.6 years of education. Focusing further on the issue of paying for child care, those single mothers employed and paying for care are a bit less likely to be nonwhite and less likely to live in poverty or receive welfare than all employed single mothers. Additionally, they are less likely to work part time, and they earn higher average hourly wages ($8.96 vs. $8.25 an hour).

Turning to Table 2, the working single mothers not reporting welfare recipiency are the oldest and have the most education and the lowest poverty rates. Their higher nonlabor income may indicate that they are more likely to be receiving child support payments. The other group with relatively higher nonlabor income is the group not employed and not on welfare. Some of these women are also receiving child support, but there is substantial variation among themselves, as the high poverty rate indicates. Others may be queued for welfare, waiting for their savings to be depleted.

Looking now at the two employed subgroups in Table 2, note that the nonwelfare group is far less likely to be employed part time and receives a considerably higher average hourly wage ($8.61 vs. $5.41 an hour). In addition, note that while the welfare recipient group is less likely to pay for care

[6] The origin of these added state-level variables are listed here: average Medicaid payment per enrollee (Table D5, *State-Level Databook on Health Care Access and Financing*, by David W. Liska, Niall J. Brennan, and Brian K. Bruen), average monthly AFDC payment (Table 605, *Statistical Abstract of the United States*), average unemployment rate (BLS data downloaded from the BLS Web site), regulated child:staff ratio (data compiled by the Center for Career Development in Early Care and Education at Wheelock College, based on data provided by Work/Family Directions, Inc.), center teachers' education regulated (data compiled by the Center for Career Development in Early Care and Education at Wheelock College, based on data obtained in their review of state licensing regulations conducted in 1994), state per capita income (Table 1, *Survey of Current Business*, 1999, 79, p. 35), and employers' estimated workers' compensation (data compiled by Ed Welch, editor of *Worker's Compensation*).

[7] Seven states are not identified uniquely. Iowa, North Dakota, and South Dakota are in a first group, and Alaska, Idaho, Montana, and Wyoming are in a second group. For these two groups of states, the state-level variables are state group averages.

**Table 1.** Means and Standard Deviations for Demographics, Employment, and Child Care Variables[a]

| | Single Mothers | | | | |
|---|---|---|---|---|---|
| Variables | All | Not on Welfare | On Welfare | Employed | Employed and Pays for Care |
| **Demographics** | | | | | |
| Age | 28.01 | 28.24 | 27.70 | 28.48 | 28.56 |
| | (6.82) | (6.77) | (6.88) | (6.65) | (6.22) |
| Education | 11.82 | 12.31 | 11.15 | 12.50 | 12.55 |
| | (2.12) | (2.04) | (2.04) | (1.96) | (2.11) |
| Nonlabor income | 849.96 | 1016.12 | 625.41 | 919.65 | 849.56 |
| | (1536.21) | (1683.57) | (1277.11) | (1665.34) | (1577.61) |
| Number of children age 0 to 2 | 0.59 | 0.55 | 0.65 | 0.50 | 0.52 |
| | (0.59) | (0.55) | (0.65) | (0.54) | (0.54) |
| Number of children age 3 to 5 | 0.72 | 0.64 | 0.83 | 0.65 | 0.65 |
| | (0.63) | (0.58) | (0.68) | (0.56) | (0.57) |
| Nonwhite | 0.39 | 0.33 | 0.48 | 0.35 | 0.32 |
| | (0.49) | (0.47) | (0.50) | (0.48) | (0.47) |
| Poverty | 0.55 | 0.36 | 0.80 | 0.26 | 0.23 |
| | (0.50) | (0.48) | (0.40) | (0.44) | (0.42) |
| Poverty$^2$ | 0.80 | 0.71 | 0.93 | 0.67 | 0.62 |
| | (0.40) | (0.45) | (0.26) | (0.47) | (0.49) |
| Welfare | 0.43 | — | — | 0.11 | 0.08 |
| | (0.49) | | | (0.32) | (0.27) |
| **Employment** | | | | | |
| Proportion in labor force | 0.47 | 0.73 | 0.13 | — | — |
| | (0.50) | (0.45) | (0.33) | | |
| Part time | — | — | — | 0.27 | 0.20 |
| | | | | (0.45) | (0.40) |
| Weekly work hours | — | — | — | 35.60 | 37.16 |
| | | | | (10.06) | (9.10) |
| Hourly wage | — | — | — | 8.25 | 8.96 |
| | | | | (5.43) | (6.11) |
| **Child care** | | | | | |
| Proportion paying for care | — | — | — | 0.53 | 1.00 |
| | | | | (0.50) | |
| Weekly child care for youngest child ($) | — | — | — | — | 57.58 |
| | | | | | (33.70) |
| Hourly child care for youngest child ($) | — | — | — | — | 1.65 |
| | | | | | (1.20) |
| Number of observations | 1523 | 912 | 611 | 738 | 395 |

[a] Means and standard deviations are weighted to obtain population averages using the "topical module" weights supplied by SIPP. Standard deviations are shown in parentheses.

(36% vs. 56%), the recipient group pays a higher hourly price for child care. This may reflect the higher cost of part-time child care (see, e.g., Connelly and Kimmel in press) or the receipt of child care subsidies.

Table 3 provides additional detail concerning child care expenditures by particular mode for all single mothers, then the single mother group is broken down by recipiency status. Single mothers receiving welfare are more likely to rely on relative care and less likely to rely on center-based care. But recall that they are also more likely to work part time, an employment state more often associated with this pattern of modal choice. In addition, the welfare recipients are less likely to pay for relative

**Table 2.** Means and Standard Deviations for Demographics, Employment, and Child Care Variables by Employment and Welfare Status[a]

| | Employed | | Not Employed | |
|---|---|---|---|---|
| Variables | On Welfare | Not on Welfare | On Welfare | Not on Welfare |
| Demographics | | | | |
| Age | 28.12 | 28.53 | 27.64 | 27.47 |
| | (7.51) | (6.52) | (6.78) | (7.33) |
| Education | 11.77 | 12.59 | 11.06 | 11.57 |
| | (1.70) | (1.97) | (2.07) | (2.04) |
| Nonlabor Income | 659.35 | 953.42 | 620.44 | 1183.69 |
| | (1378.94) | (1696.05) | (1261.45) | (1638.04) |
| Number of children age 0 to 2 | 0.52 | 0.50 | 0.67 | 0.69 |
| | (0.56) | (0.54) | (0.65) | (0.55) |
| Number of children age 3 to 5 | 0.60 | 0.66 | 0.86 | 0.59 |
| | (0.53) | (0.56) | (0.69) | (0.62) |
| Nonwhite | 0.43 | 0.34 | 0.48 | 0.29 |
| | (0.49) | (0.47) | (0.50) | (0.45) |
| Poverty | 0.57 | 0.22 | 0.83 | 0.74 |
| | (0.50) | (0.41) | (0.37) | (0.44) |
| 2 × poverty | 0.85 | 0.65 | 0.94 | 0.88 |
| | (0.36) | (0.48) | (0.24) | (0.32) |
| Employment | | | | |
| Part time | 0.58 | 0.23 | — | — |
| | (0.49) | (0.42) | | |
| Weekly work hours | 28.28 | 36.55 | — | — |
| | (13.06) | (9.18) | | |
| Hourly wage | 5.41 | 8.61 | — | — |
| | (2.45) | (5.60) | | |
| Child care | | | | |
| Proportion paying for care | 0.36 | 0.56 | — | — |
| | (0.48) | (0.50) | | |
| Weekly child care for youngest child ($) | 61.91 | 57.22 | — | — |
| | (39.37) | (35.35) | | |
| Hourly child care for youngest child ($) | 2.46 | 1.59 | — | — |
| | (2.08) | (1.06) | | |
| Number of observations | 79 | 659 | 532 | 253 |

[a] Means and standard deviations are weighted to obtain population averages using the "topical module" weights supplied by SIPP. Standard deviations are shown in parentheses.

care and less likely to pay for center-based care. Neither subgroups are very likely to pay for relative care. The welfare recipient subgroup's average weekly payment for center-based care is considerably higher than for those not receiving welfare, but note that only nine single mothers fit this category, a sample of insufficient size for a meaningful statistical comparison. For all single mothers, center-based care is the most expensive, followed by home-based care and relative care, respectively.

## 5. Measuring Child Care Costs and the Problem with Censored Data

Child care costs present a problem for the empirical researcher in that they are often unknown unless the mother is engaged in market work. This is the case with the SIPP data. This situation is similar to the problem of wages that are unobserved if the person is not employed. In addition to the

**Table 3.** Child Care Mode Choice and Weekly Expenditures by Mode of Care for Employed Single Mothers[a]

| | All | On Welfare | Not on Welfare |
|---|---|---|---|
| Weekly expenditure on child care for each mode for those who pay for care ($) | | | |
| Relative care | 48.06 | 58.62 | 47.21 |
| Home-based care | 59.27 | 49.98 | 60.41 |
| Center-based care | 68.38 | 97.32 | 66.59 |
| Percentage using each child care mode | | | |
| Relative care | 44.78 | 54.73 | 43.49 |
| (No. of observations) | (325) | (42) | (283) |
| Home-based care | 17.40 | 17.65 | 17.37 |
| (No. of observations) | (133) | (16) | (117) |
| Center-based care | 37.82 | 27.62 | 39.14 |
| (No. of observations) | (280) | (21) | (259) |
| Of those who use each mode, percentage who pay for it | | | |
| Relative care | 27.65 | 14.67 | 29.77 |
| (No. of observations) | (88) | (6) | (82) |
| Home-based care | 90.51 | 85.04 | 91.23 |
| (No. of observations) | (121) | (14) | (107) |
| Center-based care | 66.48 | 46.19 | 68.33 |
| (No. of observations) | (186) | (9) | (177) |

[a] Means are weighted to obtain population averges using the "topical module" weights supplied by SIPP. All numbers relate to care arrangements for each employed mother's youngest child except for weekly expenditure figures or where indicated otherwise.

problem of limited observation of the relevant variable, child care is complicated by the fact that many families do not pay the "market price" for child care. Nonprofit centers are often subsidized in the form of free rent and require no return on investment capital. Relatives and friends may be willing to provide child care at a reduced price or at no charge either because they receive in-kind payments or because they enjoy caring for the child. In addition, some families in our sample may already receive a subsidy for their child care costs.

How one approaches this problem depends in part on the information available and in part on the question one is trying to answer. Because the focus here is on the mother's decision, only the portion of the cost she pays is relevant. Since we are interested in the effect of child care costs on welfare recipiency and employment, we use the cost of child care per hour of employment, not the cost per hour of child care used. This is the relevant decision variable for mothers of young children who are evaluating the costs and benefits of entering the labor market, with one alternative being receiving welfare.

As we argued previously, differences among families in their access to low- or no-cost care is a very pertinent issue for our problem. Using the average local market price of child care alone ignores substantial differences among families in access to below-market child care. The problem is that there is not really an exogenously given price of child care that is relevant to all consumers in the marketplace. Instead, because of differences in family circumstances and location of residence (which are assumed to be exogenous to current decision making), each individual faces her own (exogenously given) price per hour of child care. The approach we use follows from Heckman (1974), who estimated a price of child care for each woman given information about the availability of other potential caregivers.

Because child care costs differ on the basis of the number and ages of young children in the family, we include variables measuring the number of children in fairly specific age categories that relate directly to child care options available to children of various ages. Our measure of child care costs is the predicted cost per hour of employment of child care for the youngest child in the family controlling for the number of other young children in the household.[8]

The problem of censored data is handled using the methodology described by Tunali (1986) and first applied to the problem of child care by Connelly (1992). This is a bivariate sample selection correction akin to the well-known Heckman correction to the wage equation (Heckman 1976). This method has since been used by a number of researchers interested in estimating child care costs, including the U.S. General Accounting Office (1994), Kimmel (1995), Powell (1997, 1998), Han and Waldfogel (1998), Kimmel (1998), and Anderson and Levine (1999), among others. Hourly child care costs are estimated using information from all women who are currently employed, taking into account both the selection in the employment decision and the large number of women who are employed but whose money costs of child care are zero. Child care expenditures (measured in natural logarithm form) are assumed to be a linear function of a set of individual and family and locational variables, which includes the number of children of various ages, the presence of other potential caregivers in the family, age, race, nonlabor income, region, and state child care regulations.

The statistical technique used involves estimating a bivariate probit model predicting employment and nonzero expenditure for child care. The results of this bivariate probit are used to create the selection terms that are used in the second-stage linear estimation of hourly expenditures. The results of the bivariate probit and other supporting estimations are presented in appendix tables. The coefficients estimated in this two-stage procedure are then used with the individual woman's characteristics to predict an hourly price of child care for each mother in the sample. This prediction is for care as well as the expected cost of paid care; that is, we estimate the unconditional expected price of child care (which accounts for the expected probability of paying), and use the resulting coefficients and individual characteristics of the women to estimate $E[P_{cc}] = E[P_{cc} \mid \text{Paying}$ Paying $= \text{yes}] * \text{Prob}[\text{Paying}]$.[9]

One should note that while we think this method of estimating child care costs has substantial benefits over alternatives such as average child care costs in the location of residence (which is not available with SIPP data), because of its acknowledgment of differences in the probability of paying for care, the disadvantage is that bivariate probits are in general quite sensitive to sample size. In this research context, we found that we could not get robust estimates of the price of child care using the single mothers sample only. So to increase our sample size, we included in our preliminary regressions all women with young children, both married and unmarried women, who are employed and paying for care. With married women included in the sample used for estimating the price of child care (and wage rates), the estimated price of child care is robust to other issues of model specification (Anderson and Levine 1999 also use this technique to resolve robustness problems arising from small subsamples). As long as married and unmarried women do not differ in the structure that converts individual and family characteristics into the probability of paying for child care and the amount paid if the cost is greater than zero other than a shift in the intercept (which we do allow), then our strategy is an appropriate one. If differences between single and married women cannot simply be captured by a single dummy variable, then our estimated price of child care may not fully capture the experience of single mothers' decision making.

With predicted child care expenditures for the youngest child of each single mother, we can analyze how changes in the price of child care might affect the probability of employment and the

[8] See Gelbach (1999) for a model of the natural experiment of having a child turn eligible for public school on employment of mothers.

[9] See Connelly (1992) for the explicit derivation of the unconditional expected price.

probability of AFDC receipt. We can also simulate "tied" programs, such as increased child care subsidies enacted in conjunction with lowered AFDC benefits. A set of policy simulations are discussed after our analysis of the main results.

## 6. Summary of Estimation and Identification

Our full estimation involves several steps that we summarize here. First, as discussed previously, we must create the two predicted regressors (predicted child care prices and predicted wages). These are constructed with two different sets of preliminary regressions. To construct predicted wages, we use the full sample of married and single mothers to run a reduced-form employment probit equation. This is used to construct the single Heckman correction term for inclusion in the wage equation. The Heckman correction addresses the econometric problem of sample selection resulting from estimating the wage equation only for those individuals with positive wages. Still using the full sample, we then estimate the wage equation including this Heckman selection as one of the included variables. The resulting coefficients from that model are used to construct predicted wages for each individual in the single mothers' sample. The coefficient on the Heckman correction term is not used in the construction of the predicted wage, thus giving us the E[W], not the E[W | Employment = yes].

To construct predicted child care price for the youngest child, we first run a reduced-form bivariate probit model that includes both a reduced-form employment equation and a reduced-form probability of paying for care equation, again using the full sample of married and single mothers of children under age six. These results are used to construct the two correction terms needed for inclusion in the price of the child care equation. The price per hour worked of child care for the youngest child is estimated using the sample of married and single mothers who are both employed and pay for care. The resulting coefficients of this price of child care equation are then used to construct predicted unconditional hourly price of child care for the youngest child for each single mother in the sample, $E[P_{cc}]$.

The strategy used requires that the selection terms that are constructed from a nonlinear combination of reduced-form variables be identified in the second-stage equation.[10] For the wage equation, nonlabor income, the set of household composition variables, and the state-level variables related to the price of child care and the generosity of the state's welfare system, such as the state's regulated child:staff ratio for four year olds and the state's average monthly AFDC payment, serve as identifiers of the inverse Mills ratio.[11] For the price-of-child-care equation, we have only one identifier other than the functional form that is our measure of the health status of the mother. However, this variable seems to satisfy both criteria of an adequate identifier. It is a significant predictor of both employment and the probability of paying but should not be expected *a priori* to affect the amount paid for care once one does decide to pay for care.

Once we have the two predicted values in hand, we run two versions of the full model. First, we estimate the final AFDC and employment probits separately. Second, we implement a full bivariate probit model that takes into account the error structure relationship between employment and recipiency. Our policy simulations and cost estimates are constructed from these final bivariate probit results.

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[10] Technically, one can identify off of the nonlinearity itself, but one prefers not to.

[11] The full set of identifiers of the inverse Mills ratio of the wage equation includes nonlabor income, number of other preschoolers, youngest child is an infant, number of children age 3 to 5, number of children age 6 to 12, number of children age 13 to 17, presence of other adults, state's regulated child:staff ratio less than 10:1, state's regulated center teachers' education, state's average Medicaid per enrollee, and state's average monthly AFDC payment.

Here, too, issues of identification arise. What is needed to identify the price of child care and wage variables are variables included in those estimating equations that are excluded from the final probits. Again we look for exclusion restrictions that can both be justified theoretically and have empirical significance in the first-stage equation. The full set of identifiers are years of education, age, age squared, number of children aged 6 to 12, number of children aged 13 to 17, presence of other adults, the state's regulated child:staff ratio of less than 10:1, the state regulated center teachers' education, employers' estimated workers' compensation payment by state, and state per capita income. These restrictions are similar to those made by a number of other authors (Anderson and Levine 1999; Crecelius and Lin 2000; Michalopoulos and Robins 2002) and ourselves in previous work (Connelly 1992; Kimmel 1998). Several of these variables satisfy the criteria of empirical significance in the first-stage equation. These include years of education, age, age squared, number of children aged 6 to 12, number of children aged 13 to 17, and state per capita income. The theoretical justification for exclusion is that the number of children age 6 to 12 and children 13 to 17 reflect the probability of paying for care but do not directly affect employment and welfare recipiency. Similarly, the argument is that education, age, and age squared are strongly associated with the wage and price paid for child care but do not directly affect employment and recipiency probabilities. State per capita income is expected to be correlated with price levels in the state but not directly with employment and recipiency probabilities. Of these restrictions, probably the most controversial are the exclusion of education, age, and age squared from the final equation. We estimated the final equation with and without these variables. Our findings are qualitatively robust to the change in specification, though the elasticities of AFDC recipiency with respect to the price of child care and wages are increased when education, age, and age squared are included in the final probit. The elasticity of employment with respect to the price of child care and wages are largely unchanged. We return to this comparison in Table 5.

## 7. Estimation and Simulation Results

Table 4 presents the results from a bivariate probit estimation model in which the dependent variables are AFDC recipiency and employment.[12] For AFDC recipiency, very similar results have been obtained from other data sets.[13] Nonwhite mothers, mothers who reside in urban areas, and mothers reporting poor health are more likely to receive AFDC. The state's average AFDC payment per enrollee is related positively to AFDC recipiency, but the average Medicaid expenditure per enrollee is related negatively.

The newer finding of Table 4 is the effect of predicted child care expenditures on the probability of AFDC recipiency. As the theoretical model predicts, that effect is positive and significant, with an estimated price elasticity of AFDC recipiency equal to 1.0. Controlling for the price of care, the predicted wage (a proxy for earned income in this equation) is related negatively to the probability of welfare recipiency, with the wage elasticity equal to −0.8. Those with higher nonlabor incomes are also less likely to receive welfare, while families in which the youngest child has one or more siblings under the age of six are more likely to receive welfare.

Results for the employment equation are also consistent with *a priori* expectations. The child care price elasticity of employment equals −0.4, which falls well within the broad range of estimates

[12] We report marginal effects in Table 4. These unconditional marginal effects were evaluated at the means of the data.

[13] Graham and Beller (1989) used the 1979 and 1982 March CPS, Blank (1989) used the National Medical Care Utilization and Expenditure Survey, and Crecelius and Lin (2000) used the 1988 PSID.

**Table 4.** Marginal Effects from the Bivariate Probit Model of Employment and Welfare Recipiency

| | Welfare | Employment |
|---|---|---|
| Predicted child care price | 0.329*** | −0.143*** |
| | (3.19) | (−2.44) |
| | [1.013] | [−0.422] |
| Predicted wage | −0.269*** | 0.273*** |
| | (−8.29) | (8.23) |
| | [−0.828] | [0.808] |
| Nonlabor income | −0.434E-04*** | −0.641E-05*** |
| | (−5.94) | (−2.40) |
| Nonwhite | 0.137*** | −0.020*** |
| | (6.32) | (−3.85) |
| Unhealthy | 0.012 | −0.047* |
| | (1.22) | (−1.65) |
| Youngest child is an infant | −0.078*** | −0.033 |
| | (−2.68) | (0.40) |
| Number of other preschoolers | −0.026 | −0.060 |
| | (0.16) | (−1.27) |
| Urban residence | −0.007 | −0.028 |
| | (0.33) | (−1.02) |
| Southern residence | 0.058 | 0.056 |
| | (0.94) | (0.80) |
| State's average Medicaid per enrollee | −0.209E-04 | −0.265E-04 |
| | (−0.79) | (−1.64) |
| State's average monthly AFDC payment | 0.526E-03*** | 0.186E-03 |
| | (3.39) | (−0.78) |
| State's unemployment rate | −0.012 | −0.016 |
| | (−0.38) | (−1.03) |
| Constant | 0.227*** | −0.308*** |
| | (3.92) | (−4.11) |
| Rho | −0.759 | |
| | (−30.29) | |

Note: *T*-statistics relating to the estimated coefficient are in parentheses, and elasticities are in brackets. Significance level: * = 10%, ** = 5%, *** = 1%.

found in the current literature. The employment elasticity with respect to wage changes equals 0.8, which is also consistent with previous findings of employment elasticities for single mothers. For example, our employment elasticities are very similar to those reported by Anderson and Levine for unmarried mothers with children under six. Their employment elasticity with respect to the wage is 0.6, and their employment elasticity with respect to price of child care of is −0.6.

The bivariate probit used to estimate the model reported in Table 4 accounts for the correlation between employment and welfare recipiency. Accounting for the correlation in this case is important because unobserved variables relevant to the AFDC outcome are also likely to be relevant to the employment outcome. As expected, the estimated correlation coefficient between the two equations' error terms is negative, significant, and quantitatively large. This suggests that unobserved factors that increase the probability of employment decrease the probability of receiving AFDC.

One concern of models of this type is the robustness of the findings in terms of specification. We discussed the identifying restrictions in the previous section. We experimented with many different specifications of the early stage equations, and as long as we included married women in our sample, our results were robust to these changes. We also experimented with adding some of the overidentifying variables back into the final probit and were encouraged by the retention of significance of both of the generated regressors regardless of the specification. Of particular interest was a final model that included age and education in addition to the predicted wage and predicted price of child care. The elasticities that result from that specification are almost the same in terms of the employment elasticities but are much larger in terms of the welfare recipiency elasticities. The comparison is shown in Table 5. Since age and education figure so prominently in the value of the wage variable, it would be "pushing" our 1523 observations too hard to expect enough variation to keep education, age, wage, and the price of child care all in the final stage equation. Thus, we prefer our specification over the expanded version but caution that the reported elasticities are sensitive to this specification choice.

The quantitative results are also sensitive to the estimation strategy used. We experimented with several alternatives, including univariate probits of employment and recipiency separately and a multinomial logit model that treats the four cells of our bivariate probit as four separate states of the world. The univariate probit might be preferred for ease of calculation. However, the bivariate probit model of Table 4 allows the error terms of the two equations to be correlated, improving the efficiency of the estimation process and producing more accurate standard errors. A weakness of the bivariate model is that it constrains the model to a single coefficient vector for employment and one for recipiency, allowing only for interactions in the error terms. The third alternative, the multinomial logit model, allows the effect of price of child care, for example, to differ between the state of employed/not receiving AFDC and employed/receiving AFDC. While more freedom for the coefficients is usually preferred in econometric models, the multinomial logit requires the assumption of independence of irrelevant alternatives (see, e.g., Greene 2000). In our model, this requires the assumption that if we were to remove one of the four possible cells (corresponding to the $2 \times 2$ matrix for labor force participation and welfare recipiency), the estimated coefficients corresponding to the other three cells would not be affected. In other words, removing the option of not working and not receiving welfare would not affect the coefficients corresponding to the option of not working and receiving welfare. This seems to us to be a serious failing of this model, as one expects that the decision to receive AFDC and work is closely linked with the decision to receive AFDC and not work. Michalopoulos and Robins (2000) discuss this shortcoming in their paper and explain that they rely on the multinomial logit for their model only because of the lack of a better option in light of their 12-choice model. Because our model has only four choices (or cells), we do have another option.

The most common alternative to the multinomial logit model is a nested logit model, but this model is basically equivalent to the bivariate probit in the $2 \times 2$ case.[14] Table 5 presents the elasticities of changes in employment and welfare recipiency due to changes in the price of child care and wages for three models: the univariate probit, the bivariate probit, and the multinomial logit for the same specification of the final model. The reader will note that the elasticities are sensitive to the change in estimation strategy with our preferred bivariate probit providing, in general, the smallest elasticities.

Table 6 presents a set of simulations designed to assess the impact of child care subsidies on the probability of AFDC recipiency and on the probability of being employed. While these simulations do not address specific welfare reform proposals, the simulations help illustrate the study's estimates of price effects. The simulations were done using the coefficient estimates of Table 5 and the actual

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[14] The difference, of course, is the assumption of the distribution of the errors are extreme value in the case of the logit and normal in the case of the probit.

**Table 5.** Comparison of Estimated Elasticities across Specifications

| | Bivariate Probit as Shown in Table 4 | Bivariate Probit with Education, Age and Age Square Included | Univariate Probit | Bivariate Probit as Shown in Table 4 | Multinomial logit |
|---|---|---|---|---|---|
| Elasticity of employment with respect to price of child care | −0.42 | −0.32 | −1.18 | −0.42 | −1.07 |
| Elasticity of employment with respect to wage | 0.81 | 0.92 | 1.58 | 0.81 | 1.33 |
| Elasticity of receipency with respect to price of child care | 1.01 | 1.94 | 1.50 | 1.01 | 1.22 |
| Elasticity of receipency with respect to wage | −0.83 | −2.25 | −1.58 | −0.83 | −1.36 |

characteristics of the 1523 women in the sample. Row 2 shows that using the predicted child care expenses and the other actual characteristics of women in our sample, 40.2% of single mothers are predicted to receive AFDC and 48.5% to be employed. These baseline probabilities compare with the actual proportions in the data of 40.1% for AFDC recipiency and 48.5% for employment. If child care expenditures were subsidized 10% for all single mothers, the predicted level of AFDC recipiency falls to 34.9%, and employment rises to 52.8%. A means-tested subsidy of 10% for all women below median annual income of $24,600 has little impact on the probability of receiving AFDC or being employed compared to the non-means-tested subsidy but would cost considerably less. Tying a means-tested 10% child care subsidy with a reduction in average AFDC receipts is successful in reducing AFDC recipiency from 36.0% to 32.2% but has almost no impact on employment.

With child care expenditures reduced to one-half for all single mothers, AFDC recipiency would fall further to 12.5%, while employment is predicted to rise to 74.7% (row 6). Making the child care subsidy means tested moves the AFDC recipiency rate up to 17.6% (row 7), still a substantial reduction from the baseline 40.2% with a large cost savings. Tying the child care subsidy to a reduction in average state benefits (row 8) reduces the receipency rate still further to 15.1% and increases the employment rate to 69.5% with further cost saving in AFDC expenditures. Taken as a whole, these results of our simulations indicate that subsidizing child care costs for all single mothers may be an important policy tool leading to lower AFDC recipiency rates. These subsidies could be packaged with existing federal TANF program restrictions on length of total, lifetime welfare recipiency, and work requirements to improve living standards for ex-recipients by helping to "make work pay."

Table 7 makes explicit the cost versus saving trade implicated to our discussion of Table 6. Table 7, column 1, shows the estimated annual savings in the total AFDC expenditures that would result from the lower AFDC recipiency rates alongside estimated annual costs of the subsidy. These are "back-of-envelope" calculations using each woman's predicted wage assuming full-time employment and full-time use of child care and predicted price of child care for the youngest child. Savings are accrued if the woman was predicted to be receiving AFDC in the baseline calculation and predicted to be not receiving AFDC in the simulation. Child care subsidy costs were accrued if the woman was predicted to be employed in the simulated scenario. The savings ignore potential savings from Medicaid, food stamps, and other means-tested programs, such as housing and potential gains of income tax dollars. The costs columns ignore the child care costs of a second or third child in the same family. Column 2 assumes that only single mothers' child care costs are subsidized and ignores increased government obligations from the earned income tax credit. Column 3 again assumes that only single mothers' child care costs are subsidized but

**Table 6.** Simulation Results

| Row | | Predicted Probability of Receiving AFDC (%) | Predicted Probability of Being Employed (%) |
|---|---|---|---|
| 1 | Actual data means | 40.1 | 48.5 |
| 2 | Baseline predictions from bivariate probit model (Table 5) | 40.2 | 48.5 |
| 3 | 10% subsidy of predicted hourly child care cost ($P_{cc}$) | 34.9 | 52.8 |
| 4 | 10% subsidy of $P_{cc}$ for those below median predicted annual income | 36.0 | 51.8 |
| 5 | 10% subsidy of $P_{cc}$ for those below median predicted annual income and 20% reduction in average AFDC benefits in state of residence | 32.2 | 52.7 |
| 6 | 50% subsidy of $P_{cc}$ | 12.5 | 74.7 |
| 7 | 50% subsidy of $P_{cc}$ for those below median predicted annual income | 17.6 | 68.7 |
| 8 | 50% subsidy of $P_{cc}$ for those below median predicted annual income and 20% reduction in average AFDC benefits in state of residence | 15.1 | 69.5 |

Note: Simulations were done using actual characteristics of the 1523 single mothers except for the predicted price of child care. The predicted price of child care was reduced for the given percentage for each woman in the sample in lines 3 and 6. In simulations 4 and 7, a predicted income is calculated using the predicted wage and assuming 2000 hours of employment. The predicted price of child care was reduced for any woman in the sample with a predicted income less than $24,800 per year. Simulations 5 and 8 couple the means-tested subsidy of $P_{cc}$ with a simulated 20% reduction in average AFDC benefits in one's state of residence.

included an estimated earned income tax credit for newly employed single mothers. Column 4 estimates the costs of a child care subsidy that would apply to all employed mothers of young children and included the earned income tax credit (EITC) costs for both single and married EITC eligible mothers. The number in column 5 represents the net cost of the subsidy comparing the cost calculations of column 4 with the AFDC-derived savings of column 1. The results of column 5 compared with column 4 show that the net cost of a child care subsidy program is reduced by the savings from lower recipiency rates. Even without a reduction in the amount of AFDC benefits, the cost of subsidizing child care for low-income mothers appears to be low because of substantial savings from lower recipiency rates.

## 8. Conclusions

Many papers have examined the effect of child care costs on the labor market decisions of mothers of young children. This paper is one of only a few that looks specifically at the effect of child care costs on the decisions of single mothers concerning employment and AFDC recipiency. In doing so, it seeks to answer the questions made so relevant first by the Family Support Act of 1988 and more recently by the Personal Responsibility and Work Opportunity Reconciliation Act of 1996: Can subsidizing child care reduce the welfare dependency of single mothers?

The answer seems to be an unequivocal yes, though the size of the estimated effect is found to be sensitive to the specification of the model and the estimation strategy used. Simulations using our preferred specification, which has much smaller elasticities with respect to recipiency, show that

**Table 7.** Cost Simulation Results

| | | 1 | 2 | 3 | 4 | 5 |
|---|---|---|---|---|---|---|
| | | Predicted Annual Savings from Reduction of AFDC Recipiency and/or Reduction in Recipient Amounts (in Millions) | Predicted Annual Cost of the Subsidy for Single Women Only (in Millions) | Predicted Annual Cost of the Subsidy for Single Women Only Plus Extra EITC | Predicted Annual Cost of the Subsidy for All Women Plus Extra EITC | Net Cost of the Child Care Subsidy Cost Savings (in Millions), Column 1 Minus Column 4 |
| 1 | 10% subsidy of predicted hourly child care cost ($P_{cc}$) | 1803.5 | 604.1 | 1159.9 | 3738.8 | 1935.3 |
| 2 | 10% subsidy of $P_{cc}$ for those below median predicted annual income | 1588.8 | 436.4 | 992.2 | 1279.9 | −308.9 |
| 3 | 10% subsidy of $P_{cc}$ for those below median predicted annual income and 20% reduction in average AFDC benefits in state of residence | 2764.8 | 447.0 | 1090.5 | 1338.6 | −1426.2 |
| 4 | 50% subsidy of $P_{cc}$ | 6237.0 | 4658.0 | 7323.3 | 22821.9 | 16584.9 |
| 5 | 50% subsidy of $P_{cc}$ for those below median predicted annual income | 5687.7 | 3464.3 | 6129.0 | 7978.7 | 2291.0 |
| 6 | 50% subsidy of $P_{cc}$ for those below median predicted annual income and 20% reduction in average AFDC benefits in state of residence | 6105.4 | 3513.2 | 6258.2 | 8065.8 | 1960.4 |

Note: Simulated costs of columns 1, 2, and 3 are based on actual characteristics of 1523 single mothers weighted with the wave weights and the estimated coefficients of Table 5. Costs are added in terms of subsidized child care if the woman was predicted to be employed $Y^* > 0.5$. Savings were added in terms of AFDC savings if the predicted probability of receiving AFDC is $>0.5$ in the baseline prediction and $<0.5$ with the simulated values. Column 4 added the simulated costs of the child care subsidy for married women using our married women sample and coefficients for the probability of employment. Columns 3 and 4 also estimate the increase in earned income tax credits (EITC) due to increased employment probability of low-income (EITC-eligible) families, assuming our predicted wage if employed and 2000 hours of employment.

AFDC recipiency is reduced by 28 percentage points when child care expenditures are subsidized by 50% for women with annual incomes below the median and, equally important, that employment is increased by more than 25 percentage points. While that sounds like a large subsidy, recall that the average weekly expenditure on child care is about $58. However, any program that was designed to address the quality of child care would raise this average weekly cost. Availability would also be of concern, particularly for infants, and any solution to the availability problem could also increase overall subsidy costs.[15]

Finally, these simulations do not reflect a broad equilibrium system that would model reverberations of such a subsidy throughout the entire economy. Projection of the ultimate total impacts of such a policy is complicated and perhaps falls outside of what we can expect from data-based analysis. Yet the estimates presented in this paper do show the value of child care subsidies in encouraging self-sufficiency gained through market work.

[15] For example, see Mach and Reagan (2001).

**Appendix A**

Determinants of the Probability of Paying for the Primary Child Care Arrangement of the Youngest Child and the Amount Paid for That Care

| Variable | Pay for Care ($n$ = 5764) | Natural Logarithm of Hourly Price of Child Care ($n$ = 1677) |
|---|---|---|
| Years of education | 0.003 | 0.030*** |
| | (0.20) | (2.29) |
| Age | 0.005 | 0.014*** |
| | (1.17) | (4.37) |
| Nonwhite | −0.105* | −0.124*** |
| | (−1.70) | (−2.33) |
| Nonlabor income | 0.659E-04*** | 0.484E-04*** |
| | (5.57) | (2.84) |
| Youngest child is an infant | 0.174*** | 0.109** |
| | (3.70) | (2.10) |
| Number of other preschoolers | 0.091 | 0.260*** |
| | (1.56) | (5.54) |
| Number of children age 6 to 12 | −0.010 | −0.074*** |
| | (−0.24) | (−2.17) |
| Number of children age 13 to 17 | −0.135* | −0.166*** |
| | (−1.82) | (−2.63) |
| Presence of other adults | −0.339*** | −0.119 |
| | (−5.07) | (−1.25) |
| Unhealthy | 0.285*** | — |
| | (2.68) | |
| Urban residence | −0.122*** | 0.140*** |
| | (−2.33) | (3.02) |
| Southern residence | 0.158** | −0.011 |
| | (2.16) | (−0.18) |
| State's regulated child:staff ratio <10:1 | 0.025 | 0.066 |
| | (0.42) | (1.54) |
| State's regulated center teacher's education | −0.041 | 0.038 |
| | (−0.73) | (0.92) |
| State's average Medicaid per enrollee | −0.227E-04 | −0.883E-05 |
| | (−0.89) | (−0.44) |

**Appendix A**
Continued

| Variable | Pay for Care ($n$ = 5764) | Natural Logarithm of Hourly Price of Child Care ($n$ = 1677) |
|---|---|---|
| State's average monthly AFDC payment | 0.333E-03 | 0.253E-03 |
| | (1.21) | (1.19) |
| State per capita income | −0.120 | 0.238*** |
| | (−0.90) | (2.41) |
| Married | −0.339*** | 0.060 |
| | (5.50) | (0.66) |
| λ from YESPAY | — | −0.009 |
| | | (−0.03) |
| λ from employment | — | −0.010 |
| | | (−0.06) |
| Constant | 0.663** | −1.252*** |
| | (2.19) | (−4.32) |

Note: Table values are coefficients from bivariate probit for YESPAY and the OLS price equation. $T$ = statistics are in parentheses. Significance level: * = 10%, ** = 5%, *** = 1%. These results are used to construct the predicted price of child care for each mother in the sample, which is used in the models presented in Tables 4 and 5.

**Appendix B**
Determinants of the Probability of Being Employed and the Hourly Wages (Probit Model for Employment and OLS Selection Equation for Hourly Wages)

| Variable | Employment ($n$ = 5764) | Natural Logarithm of Hourly Wage ($n$ = 3088) |
|---|---|---|
| Years of education | 0.116*** | 0.106*** |
| | (14.27) | (16.37) |
| Age | 0.179*** | 0.126*** |
| | (7.70) | (6.80) |
| Age squared | −0.003*** | −0.002*** |
| | (−7.58) | (−5.48) |
| Nonwhite | −0.068 | −0.037 |
| | (−1.37) | (−1.13) |
| Total number of children | — | −0.110*** |
| | | (−6.13) |
| Nonlabor income | −0.899E-04*** | — |
| | (−9.87) | |
| Number of other preschoolers | −0.400*** | — |
| | (5.46) | |
| Youngest child is an infant | −0.150* | — |
| | (−1.65) | |
| Number of children age 3 to 5 | −0.055 | — |
| | (−0.70) | |
| Number of children age 6 to 12 | −0.263*** | — |
| | (−10.87) | |
| Number of children age 13 to 17 | 0.023 | — |
| | (0.37) | |
| Presence of other adults | 0.171*** | — |
| | (3.23) | |
| Unhealthy | −0.477*** | −0.226*** |
| | (−6.70) | (−3.72) |

**Appendix B**
Continued

| Variable | Employment ($n$ = 5764) | Natural Logarithm of Hourly Wage ($n$ = 3088) |
|---|---|---|
| Urban residence | 0.003 | 0.087*** |
| | (0.07) | (3.16) |
| Southern residence | −0.013 | −0.001 |
| | (−0.22) | (−0.04) |
| State's unemployment rate | −0.068*** | 0.016 |
| | (−3.39) | (1.41) |
| State's regulated child:staff ratio <10:1 | 0.021 | — |
| | (0.37) | |
| State's regulated center teachers' education | 0.124*** | — |
| | (2.63) | |
| State's average Medicaid per enrollee | −0.386E-04 | — |
| | (−1.61) | |
| Employers' estimated workers' compensation payment of state | −0.010 | −0.003 |
| | (−0.28) | (−0.18) |
| State's average monthly AFDC payment | 0.129E-03 | — |
| | (0.53) | |
| State per capita income | −0.142 | 0.207*** |
| | (−1.23) | (4.29) |
| Married | 0.195*** | 0.057** |
| | (3.84) | (1.98) |
| λ | — | 0.392*** |
| | | (4.89) |
| Constant | −2.938*** | −2.255*** |
| | (−7.17) | (−6.98) |

Note: Table values are coefficients from the employment probit equation and the OLS (ln)wage average equation. *T*-statistics are in parentheses. Significance level * = 10%; ** = 5%; *** = 1%. These results are used to construct the predicted wage for each mother in the sample, which is used in the models presented in Tables 4 and 5.

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