The Effects of Child Care Costs and Taxes on the Employment of Single Mothers: Evidence from a SIPP-CPS Matching Procedure [1]

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Abstract

Child care subsidies and the Earned Income Tax Credit are among the most prominent government tools for increasing the work incentives of single mothers. Although a substantial literature examines the labor supply effects of child care costs and taxes separately, no study has done so within a modeling framework that accounts for both factors simultaneously. Given that previous research consistently demonstrates the importance of prices and taxes for single mothers, excluding one of these might lead to an omitted variables problem. Therefore, a contribution of this paper is to merge empirical techniques from previous child care and EITC studies to estimate the effect of prices and taxes on the employment of single mothers. A dataset is created by merging child care expenditure data from the SIPP with labor market information from CPS samples over the period 1990-2004. Main results suggest that single mothers' work decisions are sensitive to child care costs and taxes, although the effects are economically small. Estimated price effects are considerably smaller than those reported in previous child care studies, and I explore reasons for these differing results.

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[1] SIPP: Survey of Income and Program Participation; CPS: Current Population Survey.
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I. Introduction

The 1990s marked a watershed period in the evolution of U.S. social policy. Indeed, significant changes were introduced across a number of policy domains, each with the goal of increasing the incentive for single mothers to reduce welfare dependency and enter the labor force. Additional funding for child care subsidies and the Earned Income Tax Credit (EITC) are among the most prominent vehicles through which the federal and state governments have eased the transition from welfare to work. Expenditures on child care subsidies increased from $168 million in 1990 to $9.4 billion in 2004, owing in large part to the 1996 passage of the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) and the creation of the Child Care and Development Fund (CCDF) (Besharov & Higney, 2006). Similarly, dramatic expansions of the EITC in 1990 and 1993 increased funding from $10.5 billion to $33.8 billion over the same period (Green Book, 2004). With annual expenditures exceeding those of the Temporary Assistance to Needy Families (TANF) program, the EITC is now the single largest antipoverty program in the U.S. and the fastest growing item in the federal budget.

Concurrent with these policy changes has been the explosion in employment among single mothers and a rapid decline in the welfare rolls. Specifically, between 1990 and 2004, the employment rate for single women with children (ages 0-12) increased from 68.9 percent to 76.6 percent, peaking at 81.7 percent in 2000. Conversely, after reaching 5.0 million families in 1994, welfare caseloads declined to approximately 2.2 million, its lowest level in 30 years.

Given the recent changes to child care and tax policy, a growing empirical literature has attempted to estimate their causal effects on the employment of single mothers. Isolating these policies is a difficult task for several reasons. First, at the same time that reforms to child care and tax policies were implemented, other changes occurred that created similar incentives for single mothers to work. States began experimenting with welfare reforms in the early 1990s that ultimately became the basis for the 1996 PRWORA. Distilling the effects of welfare reform has been challenging in and of itself because states often made many changes simultaneously. In addition, the strong economy throughout the 1990s increased real earnings for low-skilled workers for the first time in nearly two decades, providing an unambiguous incentive to leave welfare for work. Second, although a substantial literature examines the labor supply effects of child care prices and taxes separately, to date no study has done so within a modeling framework that accounts for both factors simultaneously. Given that previous research demonstrates the importance of prices and taxes for single mothers, excluding one of these factors might lead to an omitted variables problem.[3]

However, estimating child care prices and taxes is complicated because these variables are endogenous to the work decision and are observed only among employed single mothers. Therefore, a large number of supporting equations must be specified in order to handle these issues. A further complication arises from the fact that nationally representative surveys either do not collect data on child care expenditures (Current Population Survey, CPS) or do not contain large samples of single mothers (Survey of Income and Program Participation, SIPP). Previous child care studies therefore suffer from a number of shortcomings: most research focuses on the effects of pre-welfare reform prices; variation in prices comes from only a single cross-section of data; no study in the literature controls for other changes in the policy environment; and many do not control for unobserved heterogeneity.

Accordingly, this paper makes a number of contributions. First, I join together empirical techniques from previous child care and EITC studies to simultaneously estimate the effects of prices and taxes on the labor supply of single mothers. A unique dataset is created by merging child care expenditure data from multiple panels of the SIPP with demographic and employment

[3] It is common for child care studies to include pre-tax hourly wages in the employment equation. However, no published child care or tax study incorporates child care prices and net-wages in its models. Some recent work attempts to adjust child care prices for subsidies [Averett, Peters, and Waldman (1997); Tekin (2007)]. Both studies are based on cross-sectional data, and given the longitudinal nature of this study and the poor availability of child care subsidy data over time, I choose not adjust prices in the present study.

data from the 1990 to 2004 March CPS. This is accomplished by first constructing SIPP and CPS samples in an identical manner and then creating an imputation procedure that assigns a potential child care expenditure to single mothers in the CPS based on shared characteristics with mothers in the SIPP. To account for EITC expansions, NBER's TAXSIM model is used to create a net-of-taxes wage variable for CPS single mothers. These data are supplemented by state-level information on welfare policies and economic conditions. The construction of a rich dataset over a significant time period allows for a more rigorous test of the effect of child care prices and taxes than previous work.

This study also employs a new methodology to account for the endogeneity of child care expenditures by using a tri-variate sample selection framework estimated via simulated maximum likelihood. A three-equation sample selection procedure accounts for censoring in the SIPP survey design as well as single mothers' underlying decisions regarding employment and child care use that lead researchers to observe child care costs for some women. Furthermore, given that previous child care studies neglect careful specification checks, I assess the robustness of price-effects by comparing estimates from tri-variate and bi-variate sample selection models with those that do not assume selection bias, and by using data on states' child care regulatory and workforce environments to examine the impact of alternative exclusion restrictions.

Estimates from the main employment model yield elasticities of employment (last year) with respect to child care expenditures and net-of-taxes wages of -0.08 and 0.12. One of the central implications of this finding is that child care price-effects are considerably smaller than what is commonly found in the literature. Most child care studies find price elasticities in the range -0.45 to -0.75. In addition, contrary to previous studies, I find low-skilled single mothers and those with young children are not more responsive to child care prices. Specification tests indicate that child care price-effects are not sensitive to tri-variate versus bi-variate sample selection procedures, nor are the estimates particularly sensitive to the identifying instruments in the child care expenditure equation. Instead, what appears to distinguish my results from those of previous child care studies is the use of detailed controls for the social policy environment, macro-economic conditions, and unobserved heterogeneity.

The remainder of this paper is organized as follows. Section II provides an overview of the child care and EITC policy changes between 1990 and 2004. It also reviews the relevant empirical literature in each policy domain. Section III introduces the data and empirical strategy and discusses the construction of key policy variables Section IV presents results from several employment models and conducts a number of policy simulations. Finally, conclusions and policy implications are discussed in Section V.

II. Review of Policy Changes and Relevant Literature

Child Care Subsidy Policy

Throughout the early-1990s, the federal government operated four major child care assistance programs that eventually became the Child Care and Development Fund (CCDF) after welfare reform in 1996. One of the key features of the CCDF is that subsidy recipients must be engaged in a state-defined "acceptable" work activity. In addition, states can serve a broad population of non-TANF families, and are given significant latitude in the design of their subsidy regimes. Overall, PRWORA allocated $21 billion for child care assistance over a seven year period, 70 percent of which must be used to subsidized costs for families receiving TANF or transitioning from welfare to work (Greenberg, Lombardi, & Schumacher, 2000).[4]

Table 1 highlights changes to child care subsidy policy throughout the 1990s. Expenditures on the programs that eventually became the CCDF grew steadily in the period 1990 to 1996, but exploded after the passage of welfare reform. By 2004, approximately $9.4 billion

[4] See Blau (2000) for a detailed review of child care subsidy policy. Eligibility for CCDF subsidies is set at 85 percent of a state's median income (SMI), although states are able to establish a lower ceiling. States are given substantial flexibility in designing their subsidy systems, including being able to transfer up to 30 percent of their TANF block grant to the CCDF, setting reimbursement and co-payment rates, and defining acceptable work activities.

was spent on child care subsidies through the CCDF, compared to $168 million in 1990. This led to a steep rise in the number of subsidy dollars available per child under age 5, the age group targeted by child care assistance programs. Consistent enrollment data prior to 1995 cannot be computed, but the available evidence shows a moderate increase in the (monthly) number of children served by CCDF subsidies: from 1.4 million in 1995 to 1.7 million in 2004. The 1990s, moreover, witnessed equally important changes in the broader child care market. Demand for nonparental child care services rose during this period, while private child care wages, which are a good proxy for prices, remained basically unchanged.[5] As shown in Table 1, the private child care workforce grew 83 percent between 1990 and 2004, but private child care earnings increased only 25 percent.

A large body of research examines the relationship between child care costs and women's work decisions. Non-experimental evidence comes from two primary sources: studies of price effects and studies of actual subsidy programs. The review herein focuses on the former type, given its relevance to this study.[6] The most common methodological approach to examining price effects includes a discrete choice participation probit with predicted child care costs and wages as the key right-hand-side variables. Both measures are derived from OLS models that control for selectivity on employment and the decision to pay for child care (expenditures only). This basic approach is quite common in the literature, and the results are surprisingly uniform (Baum, 2002; Blau & Robbins, 1991; Ribar, 1992, Connelly & Kimmel, 2001; 2003a; 2003b; Kimmel, 1995, U.S. GAO, 1994; Connelly, 1992; Han & Waldfogel, 2001; Anderson & Levine, 2000). Although nearly every study finds a negative relationship between child care costs and mothers' labor

[5] The production of child care is a very labor intensive process, accounting for nearly 70 percent of the price of child care (Helburn, 1995).

[6] Labor supply studies of subsidy programs include Berger and Black (1992), Gelbach (2002), Meyers Heintz, and Wolf (2002), Blau and Tekin (2007), and Tekin (2005; 2007a, 2007b). Every study finds that receipt of a child care subsidy increases substantially the probability of employment. Two studies investigate the labor supply effects of the Dependent Care Tax Credit (DCTC) (Averett, Peters, & Waldman, 1997; Michalopoulos, Robins, & Garfinkel, 1992). The former study finds an elasticity of hours worked of –0.78, while the latter estimates elasticities of essentially zero.

supply, the range of elasticities is large (from 0.06 to –1.36). However, there appears to be a recent convergence of estimates in the range –0.45 to –0.75.

Federal and State Tax Policy

Arguably the most important change to work incentives faced by single mothers comes from the EITC.[7] Enacted in 1975 as part of the Tax Reduction Act (TRA), expenditures on the EITC increased dramatically throughout the 1990s. By 2003, foregone revenue due to the credit totaled $33.8 billion, up from $10.5 billion in 1990. Claimant families also grew steadily during this period, from 12.5 million to 19.3 million. Single-parent families comprise 48 percent of all claimants, and 76 percent of EITC dollars go to these families (Liebman, 1999; Green Book, 2004).

The EITC received three major expansions, but for the purposes of this study, the key reforms came through the 1990 Omnibus Budget Reconciliation Act and the 1993 Omnibus Budget Reconciliation Act.[8] By 1996, when the changes were fully phased in, families with one child received a wage subsidy of 34 percent, while families with two or more children received a subsidy of 40 percent. The maximum credit available to both families was, respectively, $2,152 and $3,556. Moreover, the introduction and proliferation of state EITCs further eased the tax burden for single mothers. By 2004, 16 states operated their own EITC—compared to five in 1990—12 of which made it refundable like the federal credit. These credits simply "piggyback" onto the federal EITC by using its eligibility rules and credit rates.[9] Annual foregone revenue

[7] See Hotz and Scholz (2001) for a detailed description of the EITC. The EITC introduces a complicated set of labor supply incentives for low-income workers. Because the program comprises three credit regions, it is useful to think of it as three separate programs. The first region is called the phase-in range, which, due to its negative marginal tax rate, operates like a wage subsidy by increasing workers' net-of-taxes wages. The plateau range, where the credit rate is zero for each additional dollar earned, acts like a lump sum transfer. Finally, the phase-out range is essentially a negative income tax because of the way it gradually phases out benefits as earnings rise.
[8] The first expansion came with the passage of the 1986 Tax Reform Act (TRA86). This legislation indexed the EITC for inflation, increased the phase-in rate, and decreased the phase-out rate.
[9] Wisconsin's EITC provides a striking example. Introduced in 1995, the Wisconsin EITC supplements the federal credit by 4 percent for families with one child, 14 percent for families with two children, and 43 percent for families with three or more children. This translates to a maximum credit that is $1,107 larger for Wisconsin's two-child families and $1,641 larger for three-child families (Cancian & Levinson, 2006).

from state EITCs ranges from $17 million in Vermont to $591 million in New York (Nagle & Johnson, 2006).

The cumulative effects of changes to federal and state tax policy over the 1990s are summarized in Table 1. Two of the most important developments are the decrease in the bottom income MTR and the increase in the federal EITC's phase-in rate and maximum credit. The income tax liability for the average single mother with one child fell $928 between 1990 and 2004. A single mother with two or more children experienced a decline of $2,034. Overall the amount of income taxes paid by the average single mother in 2004 was 160 percent less than the amount paid in 1990. Most of the decline can be attributed to the decreased federal MTR, the 1993 federal EITC expansion, and the proliferation of state EITCs.

A growing body of research evaluates the labor supply effects of the EITC. A majority of this work focuses on analyzing major changes to the EITC embedded in tax laws (Ellwood, 2000; Hotz, Mullin & Scholz, 2005; Eissa & Liebman, 1996; Meyer & Rosenbaum, 1999; 2000; 2001) or geographic disparities in the generosity of state EITC's (Cancian & Levinson, 2005). The basis for this approach is to observe participation rates for a sample of individuals most likely affected by an EITC expansion before and after passage of the law, relative to changes in a comparison group. Results from these studies as a whole find strong, positive effects of EITC expansions on the labor supply of single mothers. Another set of studies uses a structural approach, drawing on economic theory to suggest parameterizations of policy and budget constraint variables that enter the work decision. Most of this research focuses on estimating employment models at the extensive margin (Fang & Keane, 2004; Looney, 2005; Meyer & Rosenbaum, 1999; 2001; Grogger, 2003; Neumark & Wascher, 2000), while others concentrate on the intensive margin (Dickert, Houser, & Scholz, 1995; Hoffman & Seidman, 1990; Keane, 1995; Keane & Moffitt, 1998). Results from these studies find elasticities of employment with respect to the return to work in the range 0.59 to 1.16.

III. Empirical Implementation

The forthcoming discussion introduces the primary data sources used in the analysis, and describes the process by which child care expenditure data from the SIPP are merged with demographic and labor market information from the CPS. I also discuss the main modeling strategy, which examines the effects of child care expenditures, net-of-taxes wages, welfare policies, and macro-economic conditions on the employment of single mothers over the period 1990 to 2004. I then describe the construction of the child care expenditure and net-of-taxes wage variables.

Data Sources and SIPP-CPS Matching Procedure

Data for this research are drawn from multiple sources, principally the March Current Population Survey (CPS) and the Survey of Income and Program Participation (SIPP). The CPS is a nationally representative survey of approximately 60,000 households, providing detailed data on labor market behavior, income, and demographic characteristics for individuals ages 15 and over. March CPS surveys for years 1991 to 2005 are used, yielding information on employment and income for the years 1990 to 2004. I include in the sample single women (widowed, separated, divorced, and never married) ages 21 to 64, who have at least one child ages 0 to 12.[10] Single mothers from census-defined families comprise the unit of analysis. I include not only independent female-headed families (primary families), but also female heads of related sub-families and (unrelated) secondary families.[11] After applying a number of standard exclusions on the sample composition, the final analysis sample consists of 74,042 single mothers with at least one child ages 0 to 12.[12]

[10] The sample is limited to children in this age range because it is the one most relevant to simultaneous eligibility for child care subsidies, the EITC, and welfare.
[11] Defining families in this manner provides the closest match to a tax-filing unit, which is crucial for determining eligibility for the EITC and other means-tested programs.
[12] Exclusions to the sample include women in the armed services; women with negative earnings, negative non-labor income, positive earnings but zero hours of work, or positive hours of work but zero earnings; and women with hourly wages over

Table 2 presents summary statistics for selected years of the CPS sample. The human capital and demographic variables include age, educational attainment, race, marital status, non-wage income, and the presence and number of children in various age groups. A few observations about the data are worth making. First, the average skill level of single mothers increased during the observation period, as evidenced by the upward shift in educational attainment. Specifically, the fraction of single mothers with at least some college education increased 18 percentage points between 1990 and 2004. Second, marital behavior within the population of single mothers changed dramatically. Never married mothers comprised about 41 percent of all single mothers in the early-1990s, but their representation grew to 51 percent by 2004.[13]

A major drawback of the CPS is that data on child care costs are not collected. Therefore, I must draw from various panels of the SIPP to impute child care expenditures for single mothers in the CPS. The SIPP comprises a series of national panels, with sample sizes ranging from approximately 14,000 to 37,000 households.[14] Although the majority of SIPP survey content focuses on a "core" of labor force, program participation, and income questions, the survey is supplemented by several "topical" modules, one of which covers child care. A typical child care module collects data on all child care arrangements for children under age 15. Detailed information is ascertained on the type of child care used, the number of hours per week a child spends in care, and the cost associated with purchasing it.[15]

$150. Also, approximately one-fourth of single mothers appear in the sample for two consecutive years, given the CPS structure.

[13] The changes in composition might be indicative of a larger issue. Grogger (2003) notes that constraining a sample to only single mothers in the context of studying the effects of welfare and tax policies could lead to a type of sample selection bias. Such policies have altered marriage and fertility incentives, leaving the population of female heads after the policy changes to appear significantly less employable. Of course, this assumes that success in the marriage and labor markets are positively correlated, but if they are, it could lead to conservative estimates of the effects of EITC and welfare policies on labor supply.

[14] The duration of each panel ranges from 2.5 to four years. Households included in a given panel are divided into four rotation groups, each of which is interviewed in successive months. The four-month period required to interview each rotation group is called a wave.

[15] There are well-known criticisms of the SIPP child care module, many of which I attempt to handle in this paper. For a review of these criticisms, see Besharov, Morrow, and Shi (2006). Many of these issues focus on changes to the survey design throughout the 1990s. For example, it was fundamentally altered three times, leading to changes in the wording of the child care questions and the timing of the module itself. During the early-1990s, the child care module was conducted throughout the fall but was changed to the spring during the late-1990s. The coverage of child care questions was also

Since the SIPP collects much of the same information as the CPS, it is possible to define both samples in exactly the same manner.[16] A critical step in this process is to achieve a close temporal match between the collection of SIPP child care data and CPS labor market and earnings data.[17] Fortunately, the SIPP introduces a child care module at several points throughout the sampling period. Specifically, I draw from the 1990, 1991, 1992, 1993, 1996 (Waves 4 and 10), and 2001 panels to conduct the match in the following manner:

SIPP Panel/Wave for the Child Care Module	**Calendar Months to Which the SIPP Child Care Data Apply**	**CPS "Data" Year(s) Matched to the SIPP Child Care Expenditure Data**
1990 Panel, Wave 3	9/90 – 12/90	1990
1991 Panel, Wave 3	9/91 – 12/91	1991, 1992
Overlapping 1992 Panel, Wave 6 and 1993 Panel, Wave 3	9/93 – 12/93	1993, 1994
1993 Panel, Wave 9	9/95 – 12/95	1995, 1996
1996 Panel, Wave 4	3/97 – 6/97	1997, 1998
1996 Panel, Wave 10	3/99 – 6/99	1999, 2000, 2001
2001 Panel, Wave 4	1/02 – 4/02	2002, 2003, 2004

For example, child care and demographic data in SIPP's 1990 panel (wave 3), which were fielded in the fall of 1990, are used to assign child care expenditures to a similarly constructed sample of single mothers in the 1990 CPS. Since the child care module is not implemented every year, there are several years during the study period that a single wave of child care data is applied to multiple years of CPS data. After both samples are created, I estimate a separate OLS child care expenditure equation for each SIPP child care module, yielding a total of seven equations. I do so to allow for shifts in the price structure over the study period. Parameter estimates from

dramatically expanded to include a larger number of child care arrangements per child, a larger number of children per family, and non-working (in addition to working) mothers. Finally, the list of available arrangements increased, and the SIPP tailored many of these arrangements to specific age ranges.

[16] Appendix A presents summary statistics for the employment, demographic, and child care characteristics of the SIPP sample of single mothers. The characteristics of SIPP single mothers match closely their counterparts in the CPS.

[17] Obtaining a close temporal match between the datasets is justified because the structure of child care prices likely changed in important ways over the sampling period. First, employment growth among single mothers lead to an increase in the demand for and supply of child care. A by-product of increased demand for child care services is the growing demand for child care labor, which accounts for 70 percent of child care prices (Helburn, 1995). Finally, public policies aimed at lowering costs and increasing quality have also contributed to a changing price structure.

these equations are then applied to the corresponding characteristics of single mothers in the CPS. It is important to note that SIPP child care data are observed only if a single mother is employed, using a SIPP-defined mode of paid child care, and are observed paying for that care. To predict child care expenditures for all CPS single mothers, several supporting equations must be estimated prior to the expenditure model. I describe this process in a forthcoming section.

Estimating the Employment Model

Using CPS data over the period 1990-2004, I examine the effects of child care prices and taxes, along with changes to welfare policy and the economy, on the employment decisions of single mothers. Therefore, I estimate a discrete choice participation equation that uses parameterizations of budget constraint, policy, and economic variables thought to influence the relative utility from employment. Stated formally, the estimated employment probit is:

[1] $\Pr[emp_{ist} = 1 \mid \mathbf{x}] = \Phi\{\alpha + \beta_1 \mathrm{E}[\ln E_{ist}] + \beta_2 \mathrm{E}[\ln w_{ist}(1-\tau)] + \mathbf{P}_{ist}'\gamma + \mathbf{X}_{ist}'\theta + \mu_s + \nu_t + \varepsilon_{ist}\}$

for $i = 1, \ldots, N_{si}$; $s = 1, \ldots, S$; $t = 1, \ldots, N$, where $\varepsilon \sim$ i.i.d. $N(0,1)$ and emp is the binary employment status for the i^{th} mother in state s at year t. The variables $\ln E$ and $\ln w(1-\tau)$ are, respectively, the natural logarithms of predicted hourly child care expenditures and net-of-taxes wages. The $\mathbf{P}'$ is a vector of policy and economic controls. Included here are states' maximum AFDC/TANF benefits available to a family of three; a dummy variable that equals 1 for all state-years after the initial implementation of any statewide waiver or welfare reform; a dummy variable that equals 1 for all state-years after the implementation of a time limit[18]; the predicted amount of disregarded earnings when calculating welfare benefits for employed single mothers; the AFDC/TANF participation rate for female-headed families; and the average, annual state

[18] As expected, the welfare reform dummy variables are highly correlated because states often made many changes simultaneously. But a multicolinearity problem should be mitigated by the fact that I interact the time limit variable with the age of the mother.

unemployment rate.[19] Child care expenditures, net-wages, and disregarded earnings are allowed to vary across women, state of residence, and year, while the remaining policies vary across state-year cells. The $\mathbf{X'}$ is a vector of human capital and demographic controls, including age (and age-squared), education, marital status, race, non-wage income, and the presence and number of children in various age groups. These variables capture underlying preferences for work and leisure as well as the value of opportunity costs associated with remaining out of work. I also include a number of controls for unobserved heterogeneity.[20] The parameters μ and ν denote state and year effects: state fixed-effects capture state-specific, time-invariant determinants of child care costs and taxes (or net-wages) that are related to the work decision, and the year dummies net out time-varying factors affecting all states. In other models, I include state-specific linear time trends, which account for unobservable factors that change linearly over time and influence employment rates. The coefficients of interest are β_1 and β_2, which measure the average change in employment associated with a one percent increase in predicted hourly child care expenditures and net-of-taxes wages.

Identification of policy-effects is achieved through a number of channels. Single mothers face different child care price structures and tax rates depending on the state of residence and sample year. These *individual* sources of variation are further exploited by the timing and intensity of federal and state changes to child care and tax policy. Recall that the sampling period covers the creation of two large child care subsidy programs in the early-1990s and a major shift in child care policy in 1996. These policy changes, coupled with dramatic changes in the broader

[19] Appendix B provides a detailed description of how the key policy variables are constructed and the theoretical effect of each on the employment of single mothers.

[20] Policy endogeneity (a manifestation of unobserved heterogeneity) is a concern for all evaluations of public policies. Grogger's (2003) explicit discussion is quite helpful in highlighting the problem. Simply stated, policy endogeneity occurs when unobserved attributes of states' single mothers are correlated with both the timing (and intensity) of policy changes *and* the propensity for employment. Another manifestation occurs when the effects of policies cannot be distinguished from other factors, such as the unemployment rate, because the onset of policy changes co-occurs with trends in economic conditions. Nearly every study in the literature deals with this issue through use of state fixed-effects, year dummies, and state-specific time trends. I do the same in this paper.

child care market, have produced significant interstate and temporal variation with which to identify child care costs. Federal and state tax structures have undergone similar dramatic changes over the sampling period. The creation and expansion of a separate EITC schedule for families with two (or more) children allows me to compare families of different sizes over time. Proliferation of state EITCs allows me to compare the tax treatment of single mothers across states and over time. Finally, for the welfare variables, I rely on the gradual phasing-in and increased intensity of states' welfare reform efforts throughout the early-1990s, culminating in a major shift in federal welfare policy through the 1996 PRWORA. As others have noted, the variation in welfare reform during the waiver period is quite rich because states phased in policy changes over a five-year period (1992-1996). However, the identification of TANF is weaker: all states implemented TANF within a 16-month period, with the majority of states doing so in 1997.

Construction of Key Policy Variables

Procedure for Adjusting Child Care Expenditures

As previously mentioned, a number of adjustments are made to child care expenditures (ln*E*) before moving these data from the SIPP to the CPS. These adjustments are required for several reasons. First, idiosyncrasies in the SIPP survey design coupled with the underlying decision-making process of single mothers leads researchers to observe child care expenditures if the following conditions are met: the mother is employed, using a SIPP-defined source of paid child care, and paying for that care. It is therefore necessary to assign a potential child care expenditure to mothers with missing data, because the cost structure faced by working mothers may not reflect that of non-working mothers had they been employed. In other words, the single mothers for whom these data are non-missing are likely a self-selected group. Second, child care expenditures are endogenous in the presence of unobserved factors related to the work decision. Unobserved child care quality is a common example: quality is related to the decision to use paid care, which in turn affects how much is spent and ultimately the employment decision.

To deal with these issues, I estimate four supporting equations: first-stage employment, use-of-paid child care, and pay-for-care equations, followed by a predicted hourly child care expenditure equation. The three first-stage equations are used to construct sample selection terms for the expenditure model; that is, child care expenditures are corrected for selectivity on employment and the joint decision to use a paid source of child care and pay for that care. It is important to note that I estimate these models separately for each SIPP child care module implemented during the observation period, yielding a total of seven sets of supporting equations.

The introduction of a tri-variate sample selection framework represents an important innovation in the child care literature. Previous research uses a bi-variate selection correction, controlling only for selection into employment and paying for child care. However, this framework ignores a common skip pattern in SIPP's questionnaire as a potential third source of selection bias. Specifically, the SIPP child care module is designed such that child care payment questions cover only a subset of arrangements, making the child's participation in one of these arrangements a requirement before expenditure data are ascertained.[21] Accounting for all three factors leads to a modeling strategy that reflects both the mechanical and behavioral processes that lead researchers to observe child care expenditures for a subset of women. Stated formally, joint estimation of the tri-variate sample selection framework is given by the following multivariate probit model:

[4] $\Pr[emp_{ist}] = \Pr[z_1 = 1 \mid \mathbf{x}) = \mathbf{X}_{ist}'\pi_1$

$\Pr[paidcare_{ist}] = \Pr[z_2 = 1 \mid z_1 = 1, \mathbf{x}) = \mathbf{X}_{ist}'\pi_2$

$\Pr[pay_{ist}] = \Pr[z_3 = 1 \mid z_1 = 1, z_2 = 1, \mathbf{x}) = \mathbf{X}_{ist}'\pi_3$,

where *emp* in represents the dichotomous employment decision; *paidcare* is the decision to use a SIPP-defined source of paid child care; *pay* represents whether the mother pays for care; and μ_{ist1},

[21] The SIPP-defined modes of paid child care include relatives, non-relatives, center-based care (including pre-school), and other school-based programs.

μ_{ist2}, and μ_{ist3} are disturbance terms distributed multivariate normal with a mean of zero and a standard deviation of unity. The $\mathbf{X}'$ is a vector of demographic and human capital characteristics of the mother and one-year lags of the state unemployment rate and maximum welfare benefit. Variables included in these models reflect single mothers' underlying preferences for work and non-maternal child care, including martial status, the availability of informal arrangements, ages of children, and disposable income. Since algorithms to evaluate multivariate normal integrals are not readily available, I rely on simulated maximum likelihood methods to jointly estimate [4]. Specifically, I use the Geweke-Hajivassilioiu-Keane (GHK) smooth recursive simulator.[22] The GHK simulator exploits the computational tractability and accuracy of the univariate normal by approximating the multivariate normal as the product of sequential univariate normal distribution functions (Cappellari & Jenkins, 2003).[23]

I then estimate the following OLS model on the sub-sample of single mothers with non-zero child care expenditures, the coefficients from which will be applied to single mothers in the CPS:

[5] $\ln E_{ist} = \mathbf{X}_{ist}'\theta + \lambda_{1\text{-}3} + \nu_{ist},$

where ln$E$ is the natural logarithm of child expenditures per hour of employment. To construct this variable, I sum expenditures across all child care arrangements for the three youngest children in a family, and then divide this amount by total hours worked during the reference week.[24] The $\mathbf{X}'$ is a vector of exogenous determinants of child care costs, including age, education, race, non-

[22] For reviews of the GHK simulator, see Greene (2003), Keane (1994), and Stern (1997). The GHK simulator is widely acknowledged as one of the fastest and most accurate simulators available. It also has a number of desirable properties: the simulated probabilities are unbiased, they are bounded within 0, 1, and the simulator is a continuous function of the model's parameters. As with all simulation methods, the GHK estimator is consistent and unbiased as the number of replications and observations increases. However, it has been shown that for the GHK, simulation bias is reduced substantially even for a moderate number of replications (Cappellari & Jenkins, 2003).
[23] It derives values for each error term by randomly drawing values from truncated normal distributions and then recursively computing simulated multivariate probability values. The process is repeated R times (as set by the analyst), with each iteration producing a value for the contribution to the simulated log-likelihood function.
[24] This definition deviates from others in the literature, which includes expenditures covering only the primary arrangement of the youngest child. However, the approach taken in this paper is preferable because it exploits all available information on mothers' child care use, and it assumes that employment decisions depend on total expenditures (and not just those of a single child).

wage income, the number of children in various age groups, metropolitan residence, and region. Variables such as age, education, and race control for individual preferences in the choice of child care services, while children's age groupings account for the fact that market prices vary according to the age of the child. Finally, selection bias is accounted for through the inclusion of three inverse Mill's ratio terms ($\lambda_{1\text{-}3}$) derived from the set of first-stage equations.

A key estimation issue is the identification of child care expenditures in the main employment model ([1]). To do so, at least one statistically significant variable appearing in [5] must be omitted from the main equation. I draw upon the most recent work by Anderson and Levine (2000) and Connelly and Kimmel (2003b) for guidance on an appropriate set of exclusion restrictions. In addition, I experiment with several alternative sources of identifying variation that are often neglected in the literature. Specifically, detailed data on state-level child care regulations, private child care establishments, and private child care workers' wages provide a rich set of instruments to identify [5]. One must assume that these variables reflect structural attributes of states' child care markets, and are therefore associated with market prices but have no direct effect on employment.[25] To assess the sensitivity of price-effects in [1], I estimate several permutations of the expenditure equation, each one varying the exclusion restrictions and the assumptions regarding selection bias.

To illustrate the results from [4] and [5], Table 3 presents estimates using the 1990 SIPP. While space limitations preclude a full discussion of the results, they are consistent with those found in the literature. In addition to the *baseline* child care expenditure equation in Table 3, I estimate a number of models that include the richer set of instruments and alter the sample selection framework, the results of which are shown in Table 4. The effects of minimum

[25] Several studies find that more stringent regulations lead to higher prices for child care (Blau, 2002; Heeb & Kilburn, 2004; Hotz & Kilburn, 1995), with either a small or statistically insignificant effect on employment (Blau, 2003; Ribar, 1992; Heeb & Kilburn, 2004; Hotz & Kilburn, 1995). To date, only a handful of studies use child care regulations as instruments in the expenditure equation, and in each case, regulations are strongly related to prices. These results suggest that child care regulations influence labor supply indirectly and only through their influence on child care prices.

standards are consistent with theoretical predictions and recent empirical work. Higher child-staff ratios and educational requirements are associated with lower child care expenditures. Both results accord with the findings in Hotz and Xiao (2005), whose estimates indicate that private child care firms gain when state regulations mandate lower child-staff ratios but lose from increased educational requirements. Results in Table 4 also suggest that raising salaries for and the supply of private child care workers are associated with greater expenditures among single mothers, confirming theoretical predictions.

Procedure for Adjusting Wages

Non-linearities arising from federal and state income taxes and transfer programs plague empirical work on labor supply, because the average low-income worker faces a large number of plausible tax rates, all of which are endogenous to the amount of labor supplied. As Moffitt (1986; 1990) points out, it is difficult to discern how changes to the budget set affect the movement of individuals from one segment to another or cause individuals to bunch at kink points.[26] Moreover, it appears that workers facing an identical set of tax rates choose very different locations on the budget constraint. This suggests a large role for unobserved work preferences in explaining the labor supply decision.

A number of methods have been used to attempt to surmount this issue (Meyer & Rosenbaum, 1999; 2001; Eissa & Hoynes, 2004; Dicket-Conlin, Houser, & Scholz, 1995). In this paper, I develop a simple approach that closely resembles those proposed by Hausman, Kinnucan, and McFadden (1979) and Heckman and MaCurdy (1981). Specifically, I rely on a set of instrumental variables to net out selection bias arising from the decision to choose a given location on the budget constraint, and I rely on a set of exogenous characteristics to predict a net-

[26] Recent work by Saez (2002) suggests that there is little evidence of bunching at kink points. Even the large, discontinuous jumps in MTR's caused by the ETIC do not appear to be associated with bunching.

of-taxes hourly wage for CPS single mothers. This approach exploits the general sample selection framework developed by Heckman (1979).

I begin by adjusting observed annual earnings in the CPS for federal (and state) income and payroll taxes using the NBER's TAXSIM calculator. TAXSIM is a micro-simulation model capable of generating income tax rates and liabilities to a fairly high degree of accuracy. I then create a net-of-taxes hourly wage variable by dividing net-earnings by annual hours worked. Next, the goal is to predict an exogenous net-wage for workers and the sub-sample of single mothers for whom wage data are not observed. This is accomplished through a two-step Heckman wage procedure, with the first step modeling the participation decision and the second step estimating an OLS net-of-taxes wage equation of the form:

[5] $\ln w_{ist}(1-\tau) = \mathbf{X}_{ist}'\psi + \lambda + \zeta_{ist}$,

where $\mathbf{X}'$ is a vector of exogenous characteristics of the mother (age, race, and education), controls for state-level economic conditions, and state fixed-effects. The λ is a sample selection term constructed from a first-stage participation probit. Variables used to identify the participation equation include dummies for the youngest child in the family, the number of children ages 0 to 18, and non-wage income. The sample selection term controls for differential employment tastes across mothers, and by extension, the propensity to choose a segment on the budget constraint. A separate wage equation was estimated for every year in the sampling period to allow for shifts in the wage structure. Table 5 presents an example of the results from this procedure using the 1991 CPS. These results accord with those from previous studies.

IV. Estimation Results from the Employment Model

Results from the main employment equation are presented in Table 6. I begin by estimating a model that mirrors a typical one found in the child care literature. I then present a number of estimates based on the full specification, followed by sub-samples of low-skilled mothers and those with children ages 0 to 5. For each model, two sets of results are shown based on the exclusion restrictions in the child care expenditure equation. The first column shows price-effects derived from the *baseline* expenditure model, while the second column shows price-effects derived from the richer set of child care instruments.[27]

Results presented in columns (1a) and (1b) come from the 1994 March CPS, and are meant to recreate a typical model in the child care literature. Such models are estimated on a single cross-section of data from the pre-PRWORA period, and include only the policy variables listed in each column and the basic demographic controls.[28] These models also omit controls for unobserved heterogeneity. A comparison of columns (1a) and (1b) shows that the coefficients on hourly child care expenditures are very similar, indicating that price-effects are not sensitive to the choice of exclusion restrictions in the expenditure equation.[29] The estimates in column (1b) imply that a unit increase in log(prices) is associated with a 36 percentage point decrease in the probability of employment (last year), while a similar increase in log(net-of-taxes wages) increases the probability of employment by 40 percentage points.

Columns (2a) and (2b) provide estimates from the full specification and sampling period. Added to these models are the complete set of policy variables, controls for the macro-economy, and a number of controls for unobserved heterogeneity. Comparing the model in (2b) with (1b) it is immediately clear that the estimates of child care expenditures and net-of-taxes wages experience a significant reduction. Marginal effects imply that increases in log hourly child care costs decreases employment by 6.1 percentage points, and a similar increase in net-wages increases employment by 8.6 percentage points. These estimates represent a reduction of 83

[27] In addition to the variables shown in Table 6, all models control for age, race, marital status, education, age ranges for the youngest child in the family, presence of a child ages 13-17, and number of children ages 0 to 5. All variables are correctly signed and statistically significant. In addition, recall that the richer set of child care instruments includes measures of minimum quality standards, the supply of private child care establishments, private child care earnings.

[28] There are exceptions to this, however. Anderson and Levine (2000) use three years of SIPP data, Connelly and Kimmel (2003) use two years of SIPP data, Han and Waldfogel (2001) use fours years of CPS data, and Tekin (2006) uses a single cross-section of NSAF data.

[29] The interpretations presented hereafter will focus on the second set of marginal effects [columns (2b), (3b), (4b)].

percent and 79 percent, respectively, in the magnitude of the coefficients from the previous model. It is interesting to note that a comparison of columns (2a) and (2b) shows once again that price-effects are not sensitive to additional instruments in the expenditure equation.

The next several variables summarize the effects of state and federal welfare reform initiatives. As expected, the monthly maximum welfare benefit shows a negative association with employment, and the amount of disregarded earnings shows a positive association. Interestingly, the size of states' maximum welfare grant appears to be a greater negative work incentive than the positive incentive introduced by increased welfare benefits for employed recipients. However, it does appear that increasing the initial earnings disregard and lowering the benefit reduction rate generates a sizable work incentive: a one percent in disregarded earnings is associated with a 1.4 percentage point increase in the probability of employment. Only one other observational study has attempted to capture the generosity of states' earnings disregards for employed welfare recipients, and it also found significant positive effects (Meyer & Rosenbaum, 1999; 2001).

Turning to the welfare reform dummy variables, I find that the onset of welfare waivers and TANF increased employment by 1.8 percentage points. Some caution must be used when interpreting this coefficient, because states often began several reforms simultaneously. This makes it impossible to disentangle the effects of individual reforms, but the positive coefficient accords with the theoretical prediction for most of the individual changes and reflects the overarching goal of welfare reform to increase employment. Building on research by Grogger (2003) and Grogger and Michalopoulos (2003), I capture the effects of time limits through a dummy variable and its interaction with the age of the mother. Allowing the effect of time limits to vary by age accounts for the possibility that mothers save their welfare benefits until an employment shock occurs. Indeed, the theoretical model developed by Grogger and Karoly (2005) suggests that forward-looking mothers will not draw upon their benefits today, opting instead to save them for future use.[30] The results in Table 6 corroborate this previous work: the coefficient on the time limit-age interaction suggests that this policy leads to smaller increases in employment as the mother ages.

Finally, I capture stigma costs and changes to the culture of states' welfare offices through the AFDC/TANF participation rate. This variable is the fraction of single mothers in a given state-year who receive welfare benefits. Recall (from Table 1) that this measure reached a peak of around 58 percent in 1992 and 1993, but declined dramatically thereafter. Therefore, one would expect much higher stigma costs associated with welfare participation in recent years. This decline also captures important changes to the culture of states' welfare offices. Indeed, it is argued that welfare offices transformed from "check writing" to "people changing" entities throughout the 1990s. This changing philosophy is borne out by the fact that several states now operate formal diversion programs that provide welfare applicants with small cash grants if they agree to stay off welfare. My findings suggest that higher psychic or transaction costs are associated with increased employment propensities. In fact, the coefficient on the AFDC/TANF participation rate suggests that a one percentage point decrease in the participation rate is expected to increase employment among single mothers by 15.0 percentage points.[31]

Table 6 also presents separate estimates for two policy-relevant sub-groups: low-skilled single mothers, defined as those with a high school degree or less, and single mothers with young children. Generally speaking, one would expect the effect of the policy variables to be greater among these sub-samples. The evidence somewhat supports these predictions, especially for mothers with low educational attainment. Low-skilled single mothers and those with young

[30] The precise relationship between time limits and employment depends on the age of the mother's youngest child. Beginning with the observation that AFDC/TANF eligibility ends when her youngest child reaches age 18, a five-year time limit does not influence work decisions when the youngest is between ages 13 and 17. However, the younger the younger the youngest child is below age 13, the stronger the incentive to "bank" welfare benefits for future use.

[31] Relative to the other welfare variables (especially the maximum credit and disregarded earnings), this variable is likely measured with substantial error, and as a result, the coefficient is biased. Meyer and Rosenbaum (2001) make a similar point, and to this I would add that the bias probably explains why the magnitude of the coefficient is so large.

children are moderately more responsive to child care expenditures and net-wages. In addition, both groups appear to be more sensitive to states' maximum welfare benefits and earnings disregards. Interestingly, the marginal effect associated with "any statewide welfare reform" is small in magnitude and statistically insignificant for low-skilled mothers. In other analyses not shown, I find that welfare reform is associated with a significant 3.3 percentage point increase in employment among mothers with some college or above. This suggests that states' welfare offices are "creaming" or that policy reforms are more successful among women already likely to be employed. However, this finding contradicts the work of others, which typically find a larger role for welfare reform among less-skilled women (Moffitt, 1999; Meyer & Rosenbaum, 1999; 2001).

Sensitivity Tests and Alternative Specifications

Estimates from the main employment model are subjected to extensive sensitivity tests to determine whether price- and wage-effects are robust to changes in the specification. Results from this exercise are shown in Table 7. Generally speaking, the estimates are robust to many alternative specifications. Given space limitations, I consider only a few below.

Rows (1) and (2) alter the assumptions regarding the selectivity of mothers for whom child care expenditures are observed. In the first case, I follow previous child care studies and estimate a bi-variate sample selection framework, controlling for selectivity on employment and paying for care, and in the second I assume the absence of selection bias. Both estimates remain negative and statistically significant at conventional levels, although the latter estimate is substantially smaller in magnitude. These results, coupled with the fact that price-effects are robust to changes in the exclusion restrictions, are contrary to Kimmel's (1998) work. Analyses by this author reveal substantial sensitivity of price elasticities to changes in the model specification.

The next two rows (3, 4) add state-specific time trends, interactions between the education dummies and year effects, and interactions between the education dummies and the unemployment rate. The interactions control for differential work propensities by education level that vary over the study period and the differential effects of educational attainment across varying economic conditions. Controls for these additional sources of unobserved heterogeneity do not alter the effects of child care expenditures and net-wages, and both coefficients remain precisely estimated.

Given the criticisms of SIPP child care data, it is instructive to examine the sensitivity of price-effects to changes in the measure of child care expenditures. One of the primary drawbacks of the SIPP is that analysts cannot uniquely identify nine states in the early panels. More recent data reduce this number to five states. This is a potentially serious omission that precludes even multiple cross-sections of data from taking full advantage of geographic variation in child care prices. To assess the influence of these states, the estimates in row (7) are derived from a model that deletes from the CPS the nine states that cannot be uniquely identified in the SIPP. The results remain unchanged. In addition, Besharov, Morrow, and Shi (2006) criticize the quality of SIPP's child care expenditure data, arguing that such information is "largely unusable for most analyses." I deal with this issue by substituting a proxy for hourly expenditures: states' weekly wage for private child care workers. Although this reduces substantially the available variation to identify price-effects, the results are once again unchanged. The coefficient on the wage variable implies that a $100 increase in average weekly wages for private child care workers reduces single mothers' employment by 5.2 percentage points.

The analysis sample for this research comprises all single mothers, regardless of marital status. In addition to never-married mothers, I include those who are divorced, separated, and widowed at the time of the survey. However, this group is heterogeneous with respect to a number of labor market, human capital, and demographic characteristics. Never-married mothers, in particular, tend to be younger (and have younger children), are lower skilled, have less education and work experience, and are more welfare prone than their previously married counterparts. From a policy perspective, never-married mothers appear to be a more relevant group for the analysis, and so I estimate the employment model with only these mothers included. As expected, never-married mothers are more responsive to child care expenditures and net-of-taxes wages. Specifically, the estimates in Table 7 suggest that a one-unit increase in the log hourly child care expenditures (and net-wages) is associated with a reduction (an increase) in the employment probability of 7.4 (11.1) percentage points.

One final set of alternative specifications is discussed. Most previous research focuses exclusively on employment at the extensive margin (the work decision), neglecting potentially heterogeneous effects of policies across infra-marginal work outcomes. Therefore, I extend the literature by examining two additional employment margins: whether a mother was employed and not receiving welfare in the previous year and (conditional on working) whether a mother was employed full-time (35+ hours), full-year (48+ weeks). Participation rates at these margins increased dramatically throughout the 1990s. By 2004, 71.5 percent of mothers were working and not receiving welfare, compared to 56.3 percent in 1990. The comparable numbers at the full-time, full-year margin are 61.5 percent and 54.3 percent. Results in Table 7 suggest that single mothers remain sensitive to child care expenditures and net-wages across these infra-marginal work outcomes. In fact, these results imply a strong role for child care subsidies and the EITC in moving single mothers from welfare to work.

Estimated Price and Wage Elasticities

To assess the sensitivity of single mothers to child care expenditures and net-wages, it is instructive to translate the marginal effects to elasticities. Table 8 presents these estimates, along with a comparison to previous child care and EITC studies. All elasticities are calculated using the coefficients in Table 6 [specifically, columns (1b), (2b), (3b), and (4b)].

My replication of previous child care studies yields a price elasticity (-0.48) within the range of previous estimates. Earlier research on child care costs produces elasticities between -0.45 and -0.75. When the model is extended to account for the full observation period and the set of controls, I find an elasticity of employment (last year) with respect to child care expenditures of -0.08 and an elasticity of employment with respect to net-of-taxes wages of 0.12. This pattern of results is more pronounced among less-skilled mothers and those with young children. I find that these mothers are no more responsive to changes in child care prices and wages, whereas previous child care studies, in particular, find such mothers are significantly more responsive.

Why do the price elasticities in this study differ dramatically from previous child care research? This study follows closely the methodological developments in recent EITC studies by using repeated cross-sectional data and including a large number of controls for social policy reforms, macro-economic conditions, and unobserved heterogeneity. However, with a few exceptions, previous child care studies omit all of these components in the data collection and estimation. For example, Connelly and Kimmel (2003a; 2003b) use multiple years of SIPP data, but do not control for other policy reforms, the unemployment rate, and state/year effects in the main employment equation. The authors also exclude critical human capital controls such as educational attainment. Anderson and Levine (2000) use three years of SIPP data, and although they omit state effects and several basic demographic controls the authors do include state effects their specifications. It appears that these omissions are the primary drivers in explaining the divergent child care results.[32] Assumptions about the sample selection framework and the mix of variables in the child care expenditure equation do not explain much of the difference. Given the large cross-state variation in child care and labor markets, state and year fixed effects are

[32] It is important to note that my price elasticities are very close to those that use more plausible sources of exogenous variation in child care costs, such as waiting lists (Berger and Black, 1992), quarter-of-birth dummies (Gelbach, 2002); geographic-based policy shifts (Baker, Gruber, and Milligan, 2005), and state fixed effects (Tekin, 2006). These studies estimate elasticities at the low end of the range: -0.12 to -0.36.

particularly important to include in the employment model. Failure to do so explains why reported elasticities in previous child care studies are severely biased upward.

Policy Simulations

To summarize the findings in the previous sections, I use estimates from the full employment model [column (2b) in Table 6] to conduct a number of policy simulations. As shown in Table 9, the Panel A presents child care subsidy simulations, Panel B simulates changes to net-wages, and Panel C explores joint policy simulations. Most of the simulation results are compared to the baseline predicted probability of employment for the full sample of single mothers (0.779).

The first two subsidy simulations provide estimates of the anticipated employment-effect from subsidy programs that do not take income into account. That is, they simulate the effects of universal child care subsidy programs that reduce expenditures by 25 percent and 50 percent. Estimates suggest that a government subsidy covering one-half of child care costs is expected to raise the employment rate to 0.825, an increase of 4.6 percentage points above the baseline.

The next policy experiment conceives child care subsidy benefits in relation to earnings. Previous studies find that low-income families pay a much larger share of their earnings on child care services than their high-income counterparts (Giannarelli, Adelman, & Schmidt, 2003). Moreover, CCDF language states that low-income families have "equal access" to high-quality providers if they do not spend more than 10 percent of their income on child care. It is therefore instructive to examine the employment-effects of a subsidy regime that limits payments to this rule-of-thumb percentage. As shown in Table 9, a subsidy system that covers child care costs after they exceed 10 percent of net-wages is predicted to increase employment to 0.848, an increase of 6.9 percentage points. Interestingly, such results imply that the average singe mother would need to receive a subsidy exceeding 50 percent of hourly costs in order to reduce them to 10 percent of net-wages.

There are a number of flaws, however, with the above simulations. First, employment-effects are derived from a universal subsidy program, whereas CCDF subsidies are means-tested, restricting eligibility to 85 percent of SMI. Second, these simulations are based on linear benefit schedules, whereas CCDF subsidies are non-linear. Therefore, the final policy experiment attempts to mimic a targeted, non-linear subsidy system comprised of six benefit segments.[33] It provides free child care to families with incomes no greater than $5,000, and then reduces the subsidy in increments of 15 percentage points until income exceeds $30,000. At this point, the family is no longer eligible for subsidies.[34] Results from this policy experiment imply a dramatic increase in employment among single mothers. Approximately 87 percent of mothers are predicted to be employed under this non-linear subsidy system, an increase of nine percentage points.

The simulations presented in Panel B begin by showing the effects of similar percent increases in the net returns to work. The predicted employment response is generally similar in magnitude to child care subsidies, which is surprising in light of the fact that wage elasticities are larger than price elasticities. A partial explanation for these results—borne out by several studies—is that low-income women are more sensitive to taxes than high-income women. I assess this explanation in the last wage experiment by comparing the employment response to a 50 percent increase in net-earnings across the bottom and top predicted wage quartiles. These results show substantial heterogeneity across the wage distribution. Employment is predicted to increase 4.5 percentage points across the bottom wage quartile, compared to 2.2 percentage points across the top quartile.

[33] It is common for states' subsidy systems to have several piecewise linear segments. Maryland's subsidy system, for example, has 10 segments that increase in approximately $2,000 to $3,000 increments. The system is furthered complicated by separate benefit and co-payment schedules for families of different sizes.

[34] This reflects quite well the chosen the break-even point among states. In 2004, the average state set its eligibility limit (for a family of three) at $29,184.

The final set of simulations, presented in Panel C, shows the predicted employment generated from simultaneous increases in child care subsidies and net-wages. Such an exercise is important given that a non-trivial fraction of low-income families are observed participating in multiple means-tested programs. For example, a study using NSAF data found that 25 percent of working poor families with a recent connection to the welfare system received child care subsidies and the EITC (in addition to Medicaid and food stamps) (Zedlewski, Adams, Dubay, & Kenney, 2006). Results from these simulations imply large employment effects. For example, providing single mothers with means-tested child care subsidies and raising the wage subsidy by 50 percent is predicted to increase employment from 0.779 to 0.891, an increase of 11.2 percentage points.

V. Conclusions and Policy Implications

Throughout the 1990s, significant changes were enacted across a number of social policy domains that increased the incentive for single mothers to work. Two of the most significant policy shifts were expansions to child care subsidy programs and the EITC. Each of these changes was implemented against a backdrop of federal and state welfare reform initiatives and unprecedented economic growth. The purpose of this paper, therefore, is to examine the effects of child care prices and taxes, controlling for welfare reform and macro-economic conditions, on the employment of single mothers between 1990 and 2004. Main results suggest that the employment decisions of single mothers are sensitive to child care expenditures and taxes, although the effects are economically small.

One of the primary implications of my findings is that child care price-effects are considerably smaller than what is commonly found in the literature. This study makes several data and methodological improvements over previous research that likely account for the differences in estimated price-effects. First, I merge empirical techniques from previous child care, EITC, and welfare reform studies to jointly estimate multiple policies alongside controls for macro-economic conditions and a rich set of demographic characteristics. Given that previous employment research demonstrates the importance of child care prices, taxes, and welfare reform for single mothers, excluding one of these factors might lead to an omitted variables problem. Second, whereas as previous child care studies use a single section cross-section of data, this study takes full advantage of cross-state and temporal variation in child care policies and markets over a significant time period. As a result, I rely on substantially more exogenous variation to identify price-effects. Third, I develop a methodological approach that characterizes accurately the self-selection of single mothers into employment and the use of paid child care.

Policy implications of this research are borne out by the simulation results. I compare the employment response of universal increases in child care subsidies and decreases in taxes to a system that provides generous, targeted assistance. Specifically, I examine the amount of employment generated by a non-linear child care subsidy system (that includes means-testing) and increases in the EITC over the program segments. Results in each exercise suggest that a system of generous, targeted work supports generates more employment than one that provides limited, universal assistance. These findings are important in light of the reauthorization of TANF and the CCDF through the 2005 Deficit Reduction Act. This legislation introduces several punitive measures for welfare recipients and states, including greatly accelerated work participation rates, a narrowing of acceptable work activities, and the imposition of financial penalties on states that fail to comply with federal guidelines. The new work requirements are, furthermore, matched with small increases in funding for child care subsidies, a TANF block grant that is not adjusted for inflation, and an economic climate less favorable than the one throughout the late-1990s. The cumulative effects of these policy changes imply that the federal government endorses the "work first" over the "make work pay" philosophy. However, results of my policy simulations suggest that the latter might actually be a more effective vehicle for increasing the employment of welfare recipients.

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TABLE 1:
Summary of Characteristics and Policy Changes Affecting Single Mothers, 1990-2004

Policy / Year	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004
Panel A: Employment (March CPS)															
Employment (%)															
All Mothers	0.689	0.676	0.662	0.672	0.703	0.713	0.736	0.755	0.797	0.817	0.817	0.796	0.788	0.770	0.766
With High School/Less	0.624	0.597	0.568	0.588	0.612	0.625	0.652	0.681	0.727	0.768	0.767	0.729	0.725	0.709	0.695
With Children Ages 0-5	0.629	0.595	0.595	0.612	0.646	0.651	0.691	0.716	0.761	0.785	0.784	0.768	0.751	0.731	0.729
Panel B: States' Child Care Characteristics															
CCDF Spending ($ in millions)	168	2154	2355	2912	3426	3740	3762	4601	6105	7124	7922	8479	9018	9720	9380
$ Per Child Ages 0-4	9	107	115	141	166	181	186	229	304	355	393	417	439	468	467
Private Child Care Earnings ($)	12195	12522	12828	12856	12976	13038	13158	13468	13824	14157	14350	14639	14976	15174	15249
Private Child Care Workers	7670	8146	8685	9404	10065	10611	10976	11481	12119	12638	13167	13695	13954	14004	14061
Panel C: Federal and Sate Tax Policies															
Bottom Income MTR	0.15	0.15	0.15	0.15	0.15	0.15	0.15	0.15	0.15	0.15	0.15	0.15	0.10	0.10	0.10
Federal EITC (%, $)															
Phase-in Rate															
1 Child	0.14	0.167	0.176	0.185	0.236	0.34	0.34	0.34	0.34	0.34	0.34	0.34	0.34	0.34	0.34
2+ Children	0.14	0.173	0.184	0.195	0.30	0.36	0.40	0.40	0.40	0.40	0.40	0.40	0.40	0.40	0.40
Maximum Credit															
1 Child	1377	1653	1783	1875	2595	2596	2591	2601	2632	2621	2581	2590	2631	2615	2604
2+ Children	1377	1713	1863	1975	3223	3855	4281	4303	4353	4327	4265	4275	4347	4316	4300
No. States with an EITC	5	6	6	6	7	7	7	9	10	11	15	15	16	16	16
Income Tax Liability															
All Mothers	966	806	739	682	338	229	94	9	-299	-226	49	-16	-303	-758	-604
With 1 Child	1465	1175	1054	971	926	1229	1127	838	561	749	1011	1337	657	354	537
With 2+ Children	542	501	450	403	-149	-582	-694	-674	-947	-1018	-716	-1092	-1111	-1625	-1492
Panel D: States' AFDC/TANF Programs															
Maximum Benefit ($)	557	545	532	515	505	494	481	467	456	451	441	434	438	430	419
Welfare Reform															
Any Reform (%)	0	0	0.023	0.218	0.325	0.363	0.606	0.907	1.0	1.0	1.0	1.0	1.0	1.0	1.0
Time Limits (%)	0	0	0	0.001	0.005	0.015	0.247	0.679	0.958	0.962	0.964	0.961	0.958	0.958	0.960
Disregarded Earnings ($)	6254	6377	6323	6587	6807	7118	7440	8754	10254	10635	13153	13978	13420	12930	13129
Participation Rate (%)	0.511	0.546	0.578	0.575	0.549	0.500	0.439	0.371	0.273	0.209	0.185	0.168	0.160	0.157	0.152
Panel E: Economic Environment															
Unemployment Rate	0.056	0.068	0.075	0.069	0.061	0.056	0.054	0.049	0.045	0.042	0.040	0.047	0.058	0.060	0.055

Notes: Dollars are adjusted for inflation to reflect 2004 prices. All means, except for the employment, welfare reform, and income tax liability variables, represent state-level averages. Welfare reform parameters indicate the proportion of single mothers in the CPS who are affected by a given reform. Time limits never reach full coverage because Michigan and Vermont do not have them. The labor supply, earnings, income tax liability, and earnings disregard variables are calculated only among those who are employed. These are calculated from the CPS sample for each year. The CCDF spending figures are from Besharov & Higney (2006).

TABLE 2:
Variable Means for the CPS Sample of Single Mothers: 1992, 1995, 1998, 2001, 2004

CPS "Data" Year	**1992**	**1995**	**1998**	**2001**	**2004**
Number of Observations	N = 4,397	N = 3,849	N = 3,773	N = 6,764	N = 6,572
Age	32.56 (8.18)	33.09 (8.69)	33.44 (8.70)	33.76 (9.04)	33.86 (9.41)
Less than High School (%)	0.236 (0.424)	0.220 (0.414)	0.201 (0.400)	0.187 (0.390)	0.198 (0.399)
High School/GED (%)	0.399 (0.489)	0.360 (0.480)	0.354 (0.478)	0.361 (0.480)	0.336 (0.472)
Some College (%)	0.275 (0.447)	0.325 (0.468)	0.332 (0.471)	0.325 (0.468)	0.346 (0.475)
BA+ (%)	0.087 (0.283)	0.092 (0.289)	0.111 (0.315)	0.125 (0.331)	0.118 (0.322)
Widowed (%)	0.044 (0.205)	0.042 (0.202)	0.046 (0.210)	0.046 (0.210)	0.038 (0.191)
Separated (%)	0.186 (0.389)	0.189 (0.391)	0.154 (0.361)	0.138 (0.345)	0.136 (0.343)
Divorced (%)	0.359 (0.479)	0.343 (0.475)	0.330 (0.470)	0.322 (0.467)	0.317 (0.465)
Never Married (%)	0.409 (0.491)	0.423 (0.494)	0.468 (0.499)	0.492 (0.499)	0.507 (0.499)
Non-white (%)	0.368 (0.482)	0.370 (0.483)	0.359 (0.479)	0.354 (0.478)	0.356 (0.479)
Non-wage Income ($)	466.99 (761.58)	508.65 (936.66)	408.60 (718.44)	428.91 (815.10)	437.06 (833.53)
Child Ages 0-2 (%)	0.283 (0.450)	0.266 (0.442)	0.239 (0.426)	0.248 (0.432)	0.258 (0.437)
Child Ages 3-5 (%)	0.351 (0.477)	0.355 (0.478)	0.343 (0.474)	0.317 (0.465)	0.335 (0.472)
Child Ages 6-12 (%)	0.674 (0.468)	0.678 (0.466)	0.694 (0.460)	0.691 (0.461)	0.682 (0.465)
Child Ages 13-17 (%)	0.211 (0.408)	0.224 (0.417)	0.230 (0.421)	0.233 (0.423)	0.232 (0.422)
Youngest Child: 0-2 (%)	0.283 (0.450)	0.266 (0.442)	0.239 (0.426)	0.248 (0.432)	0.258 (0.437)
Youngest Child: 3-5 (%)	0.256 (0.436)	0.265 (0.441)	0.269 (0.443)	0.247 (0.431)	0.253 (0.434)
Youngest Child: 6-8 (%)	0.204 (0.403)	0.222 (0.415)	0.239 (0.426)	0.218 (0.413)	0.214 (0.410)
Youngest Child: 9-12 (%)	0.255 (0.436)	0.245 (0.430)	0.251 (0.434)	0.285 (0.451)	0.274 (0.446)
No. of Children Ages 0-2	0.328 (0.562)	0.299 (0.531)	0.265 (0.498)	0.275 (0.503)	0.289 (0.523)
No. of Children Ages 3-5	0.410 (0.606)	0.411 (0.598)	0.385 (0.569)	0.353 (0.552)	0.375 (0.564)
No. of Children Ages 6-12	0.916 (0.819)	0.936 (0.842)	0.963 (0.837)	0.946 (0.820)	0.924 (0.813)
No. of Children Ages 0-17	1.916 (1.088)	1.930 (1.083)	1.900 (1.035)	1.865 (0.991)	1.885 (1.007)
Urban Residence (%)	0.812 (0.390)	0.818 (0.385)	0.834 (0.371)	0.829 (0.376)	0.841 (0.364)
Southern Residence (%)	0.368 (0.482)	0.378 (0.484)	0.364 (0.481)	0.371 (0.483)	0.391 (0.488)

Source: Author's calculations from the 1993, 1996, 1999, 2002, and 2005 March Current Population Survey (CPS).
Notes: Standard deviations are in parentheses. Data are weighed using the March Supplemental Person Weight. Dollars are adjusted for inflation to reflect 2004 prices.

TABLE 3:
Simulated Maximum Likelihood Estimates of the Tri-variate Sample Selection and Child Care Expenditure Equations, SIPP 1990 (3)

Variable	Multivariate Probit Sample Selection Equations			OLS
	Participation	Use of Paid Care	Paying for Care	ln(child care expenditures per hour of work)
Age	0.072 (0.036)**	0.133 (0.031)***	0.106 (0.033)***	0.005 (0.010)
Age^2	-0.001 (0.000)**	-0.002 (0.000)***	-0.001 (0.000)***	--
High School/GED	0.448 (0.086)***	0.460 (0.084)***	0.446 (0.089)***	-0.235 (0.160)
Some College	0.578 (0.101)***	0.572 (0.103)***	0.469 (0.108)***	-0.095 (0.190)
BA+	1.287 (0.161)***	1.067 (0.126)***	1.027 (0.130)***	-0.521 (0.262)**
Widowed	0.458 (0.190)**	0.362 (0.193)*	0.279 (0.227)	--
Divorced	0.536 (0.094)***	0.326 (0.086)***	0.382 (0.092)***	--
Separated	0.328 (0.102)***	0.290 (0.094)***	0.390 (0.096)***	--
Non-white	-0.062 (0.079)	-0.092 (0.071)	-0.095 (0.076)	0.060 (0.095)
ln(non-wage income)	-0.109 (0.009)***	-0.073 (0.007)***	-0.051 (0.007)***	0.035 (0.017)**
Child Ages 0-2	-0.264 (0.106)**	0.119 (0.113)	0.185 (0.181)	--
Child Ages 3-5	0.031 (0.102)	0.208 (0.100)**	0.225 (0.146)	--
Child Ages 6-12	0.002 (0.121)	-0.220 (0.120)*	-0.196 (0.134)	--
Child Ages 13-17	0.313 (0.110)***	-0.101 (0.106)	0.216 (0.230)	--
No. of Children Ages 0-2	--	--	0.332 (0.207)*	0.296 (0.115)**
No. of Children Ages 3-5	--	--	0.383 (0.205)*	0.201 (0.097)**
No. of Children Ages 6-12	--	--	0.272 (0.188)	0.114 (0.068)*
No. of Children Ages 0-17	-0.187 (0.051)***	-0.116 (0.051)**	-0.401 (0.185)**	--
Unemployed Adult	--	-0.351 (0.298)	-3.197 (0.252)***	--
Urban Residence	-0.018 (0.087)	0.120 (0.082)	0.070 (0.084)	0.321 (0.101)***
Southern Residence	0.115 (0.095)	0.072 (0.067)	0.063 (0.068)	-0.020 (0.085)
Unemployment $Rate_{t-1}$	-0.051 (0.025)**	--	--	--
Maximum AFDC $Benefit_{t-1}$	-0.000 (0.000)	--	--	--
Intercept	-0.224 (0.641)	-2.241 (0.563)***	-2.560 (0.599)***	0.784 (0.730)
$\rho_{1,2}$		0.982 (0.007)		--
$\rho_{1,3}$		0.968 (0.013)		--
$\rho_{2,3}$		0.967 (0.011)		--
$\lambda_{participation}$	--	--	--	0.955 (1.189)
$\lambda_{use\ of\ paid\ care}$	--	--	--	-0.444 (2.398)
$\lambda_{paying\ for\ care}$	--	--	--	-2.536 (2.166)
R^2	--	--	--	0.138
Log Pseudolikelihood		-1,886.420		--
Number of Observations		1,664		414

Source: Author's calculations from SIPP's 1990 Panel, Wave 3.
Notes: Marginal effects are presented. Robust standard errors are in parentheses. *, **, *** indicate that the coefficient is statistically significant at the 10%, 5%, and 1% levels, respectively. $\rho_{1,2}$, $\rho_{1,3}$, and $\rho_{2,3}$ is the correlation between the errors in models 1 and 2, 1 and 3, and 2 and 3, respectively. The lambdas are the sample selection parameters derived from the joint participation, use of paid care, and pay for care equations. Estimates from the other SIPP panels (waves) are available from the author upon request.

TABLE 4:
Alternative Specifications of the OLS Child Care Expenditure Equation: All SIPP Panels (Waves)

Variable	1990 (3)	1991 (3)	1992 (6)/ 1993 (3)	1993 (9)	1996 (4)	1996 (10)	2001 (4)
Panel A: Tri-variate Sample Selection Model and the Inclusion of State-level Child Care Instruments							
Average Maximum Child-Staff Ratio (centers)	-0.050 (0.023)**	-0.010 (0.030)	0.028 (0.020)	-0.023 (0.034)	0.009 (0.025)	-0.003 (0.031)	--
Educational Requirement: All Staff (centers)	-0.016 (0.009)*	-0.000 (0.015)	-0.015 (0.007)**	0.024 (0.015)*	--	--	--
Educational Requirement: Director (centers)	--	--	--	--	-0.024 (0.011)**	-0.099 (0.055)*	--
ln(annual wages for private child care staff)	0.068 (0.033)**	0.016 (0.045)	0.541 (0.247)**	-0.073 (0.045)*	0.715 (0.411)*	0.156 (0.084)*	-0.022 (0.034)
No. of Private Child Care Establishments (/ 1,000)	0.078 (0.043)*	-0.009 (0.040)	0.006 (0.028)	0.057 (0.030)*	0.008 (0.032)	-0.003 (0.027)	0.038 (0.026)
$\lambda_{\text{participation}}$	0.539 (1.230)	-0.591 (1.491)	-1.777 (1.361)	5.171 (2.213)*	5.652 (2.066)***	-1.598 (2.191)	3.693 (2.801)
$\lambda_{\text{use of paid care}}$	0.444 (2.422)	-0.576 (3.061)	0.012 (2.549)	-11.597 (3.634)***	-2.203 (3.632)	2.933 (3.495)	5.440 (3.965)
$\lambda_{\text{paying for care}}$	-3.005 (2.163)	-1.693 (2.478)	-1.044 (2.067)	2.830 (2.767)	-1.298 (2.850)	-2.895 (2.828)	-5.958 (2.863)**
Panel B: Bi-variate Sample Selection Model and the Inclusion of State-level Child Care Instruments							
Average Maximum Child-Staff Ratio (centers)	-0.051 (0.022)**	-0.008 (0.030)	0.027 (0.022)	-0.016 (0.031)	0.007 (0.025)	-0.001 (0.031)	--
Educational Requirement: All Staff (centers)	-0.017 (0.009)*	-0.001 (0.014)	-0.015 (0.009)*	0.026 (0.014)*	--	--	--
Educational Requirement: Director (centers)	--	--	--	--	-0.025 (0.011)**	-0.097 (0.055)*	--
ln(annual wages for private child care staff)	0.070 (0.033)**	0.013 (0.045)	0.539 (0.270)**	-0.073 (0.045)*	0.696 (0.411)*	0.153 (0.084)*	-0.034 (0.033)
No. of Private Child Care Establishments (/ 1,000)	0.079 (0.042)*	-0.009 (0.040)	0.006 (0.027)	0.064 (0.033)*	0.007 (0.032)	-0.002 (0.027)	0.043 (0.027)
$\lambda_{\text{participation}}$	0.565 (1.037)	-0.351 (1.181)	-1.365 (1.065)	1.711 (1.524)	4.993 (1.741)***	-1.742 (1.832)	5.768 (2.306)**
$\lambda_{\text{paying for care}}$	-2.640 (1.050)**	-1.717 (1.186)	-1.079 (0.798)	-3.414 (1.385)**	-2.507 (1.503)*	-0.330 (1.933)	-2.878 (1.483)*
Panel C: No Correction for Selection Bias and the Inclusion of State-level Child Care Instruments							
Average Maximum Child-Staff Ratio (centers)	-0.053 (0.022)**	-0.008 (0.030)	0.028 (0.020)	-0.015 (0.035)	0.018 (0.025)	-0.000 (0.031)	--
Educational Requirement: All Staff (centers)	-0.017 (0.009)*	-0.001 (0.014)	-0.015 (0.007)**	0.024 (0.015)	--	--	--
Educational Requirement: Director (centers)	--	--	--	--	-0.024 (0.010)**	-0.101 (0.054)*	--
ln(annual wages for private child care staff)	0.067 (0.033)**	0.011 (0.044)	0.442 (0.237)*	-0.075 (0.044)*	0.791 (0.412)*	0.155 (0.084)*	-0.017 (0.032)
No. of Private Child Care Establishments (/ 1,000)	0.084 (0.043)**	-0.007 (0.039)	0.004 (0.028)	0.077 (0.030)**	0.014 (0.032)	-0.007 (0.027)	0.037 (0.027)

Source: Author's calculations from SIPP's Core File and Child Care Topical Module in the 1990, 1991, 1992, 1993, 1996, and 2001 Panels.
Notes: Robust standard errors are in parentheses. *, **, *** indicate that the coefficient is statistically significant at the 10%, 5%, and 1% levels, respectively. The lambdas are the sample selection parameters derived from the joint participation, use of paid care, and pay for care equations. All models include controls for age, education, race, non-wage income, the number of children in various age groups, metropolitan residence, and region

TABLE 5:
Estimates from the Heckman Selection-Corrected Wage Equation, 1991 CPS

Variable	Probit Participation Equation		OLS Net-of-Taxes Wage Equation	
	Coefficient	SE	Coefficient	SE
Youngest Child Ages 3-5	0.228	(0.058)***	--	--
Youngest Child Ages 6-8	0.331	(0.065)***	--	--
Youngest Child Ages 9-12	0.476	(0.067)***	--	--
No. of Children Ages 0-18	-0.080	(0.021)***	--	--
ln(non-wage income)	-0.098	(0.005)***	--	--
Age	0.027	(0.016)	0.050	(0.009)***
Age-squared	-0.0004	(0.0002)**	-0.0005	(0.0001)***
High School/GED	0.643	(0.051)***	0.219	(0.035)***
Some College	1.006	(0.065)***	0.315	(0.041)***
BA+	1.294	(0.103)***	0.581	(0.050)***
Widowed	0.171	(0.104)	0.017	(0.058)
Divorced	0.436	(0.059)***	0.096	(0.030)***
Separated	0.078	(0.062)	0.027	(0.033)
Non-white	-0.193	(0.049)***	0.047	(0.026)*
Urban Residence	-0.076	(0.053)	0.089	(0.026)***
Southern Residence	0.203	(0.049)***	-0.094	(0.024)***
Unemployment Rate	-0.040	(0.023)*	-0.007	(0.011)
Intercept	-0.008	(0.315)	0.392	(0.182)**
$\lambda_{participation}$	--	--	-0.006	(0.054)
R^2	--		0.122	
Number of Observations	4,352		2,959	

Source: Author's calculations from the 1991 CPS.
Notes: Marginal effects are presented. Standard errors are in parentheses. *, **, *** indicate that the coefficient is statistically significant at the 10%, 5%, and 1% levels, respectively. Estimates from the other CPS years are available upon request from the author.

TABLE 6:
Parameter Estimates from the Employment Probit Equation

Variable	All Single Mothers				Single Mothers With A High School Degree or Less		Single Mothers With Children Under Age 6	
	(1a)	(1b)	(2a)	(2b)	(3a)	(3b)	(4a)	(4b)
ln(predicted hourly child care expenditure)	-0.392 (0.049)***	-0.360 (0.042)***	-0.063 (0.006)***	-0.061 (0.006)***	-0.077 (0.009)***	-0.075 (0.009)***	-0.078 (0.009)***	-0.074 (0.009)***
ln(predicted net-of-taxes hourly wage)	0.420 (0.091)***	0.404 (0.091)***	0.092 (0.025)***	0.086 (0.025)***	0.098 (0.045)**	0.089 (0.044)**	0.057 (0.040)	0.048 (0.040)
ln(monthly maximum welfare benefit)	-0.079 (0.024)***	-0.057 (0.025)**	-0.076 (0.031)**	-0.078 (0.031)**	-0.097 (0.045)**	-0.099 (0.045)**	-0.079 (0.047)*	-0.082 (0.048)*
ln(predicted annual disregarded earnings)	--	--	0.014 (0.005)***	0.014 (0.005)***	0.019 (0.007)***	0.019 (0.007)***	0.027 (0.007)***	0.026 (0.007)***
Any Statewide Welfare Reform	--	--	0.018 (0.009)*	0.018 (0.009)*	0.001 (0.014)	0.001 (0.014)	0.030 (0.014)**	0.030 (0.014)**
Time Limit	--	--	0.054 (0.019)***	0.049 (0.019)**	0.077 (0.028)***	0.071 (0.028)**	0.080 (0.029)***	0.074 (0.029)**
Time Limit x Age	--	--	-0.002 (0.0004)***	-0.002 (0.0004)***	-0.002 (0.0006)***	-0.002 (0.0006)***	-0.003 (0.0006)***	-0.003 (0.0006)***
AFDC/TANF Participation Rate	--	--	-0.159 (0.030)***	-0.150 (0.030)***	-0.221 (0.044)***	-0.209 (0.044)***	-0.195 (0.046)***	-0.184 (0.046)***
Unemployment Rate	--	--	-0.008 (0.003)***	-0.008 (0.003)***	-0.009 (0.004)**	-0.008 (0.004)**	-0.009 (0.004)**	-0.009 (0.004)**
ln(non-wage income)	-0.020 (0.003)***	-0.020 (0.003)***	-0.016 (0.0004)***	-0.016 (0.0004)***	-0.021 (0.0006)***	-0.021 (0.0006)***	-0.015 (0.0006)***	-0.015 (0.0006)***
State-level Child Care Instruments	No	Yes	No	Yes	No	Yes	No	Yes
State Fixed-effects	No	No	Yes	Yes	Yes	Yes	Yes	Yes
Year Dummies	No	No	Yes	Yes	Yes	Yes	Yes	Yes
Month-in-sample Dummies	No	No	Yes	Yes	Yes	Yes	Yes	Yes
McFadden's R^2	0.192	0.194	0.160	0.160	0.149	0.149	0.151	0.151
Log-pseudolikelihood	-2,215.25	-2,211.63	-35,235.54	-35,235.76	-23,282.26	-23,281.67	-19,645.37	-19,647.67
Number of Observations	4,342	4,342	74,042	74,042	43,156	43,156	37,722	37,722

Source: Author's calculations from the 1991-2005 March Current Population Survey (CPS).

Notes: Marginal effects are shown, along with robust standard errors (in parentheses). *, **, *** indicate that the coefficient is statistically significant at the 10%, 5%, and 1% levels, respectively. All models include controls for age; age-squared; marital status; non-white; educational attainment; whether the youngest child in the family is ages 3-5, ages 6-8, and ages 9-12; the presence of a child ages 13-17; and the number of children ages 0-5. Estimates for these variables are available from the author upon request. The estimates in models (1a) and (1b) come from the 1994 March CPS, while those in (2a) and (2b) are derived from the full observation period. Both sets of models include all single mothers with at least one child ages 0-12. Models (3a) and (3b) include only single mothers with no more than a high school education, while models (4a) and (4b) reduce the sample to families with at least one child ages 0-5. The models in columns (a) and (b) differ only in the way the hourly child care expenditure variable is identified in the OLS prediction equation. The variable in column (a) uses the standard demographic instruments, while the variable in column (b) uses a richer set of state-level child care characteristics. See text for details. All analyses are weighted using the March Supplemental Person Weight.

TABLE 7:
Tests of Robustness and Alternative Specifications

Specification	ln(predicted hourly child care expenditure)	ln(predicted net-of-taxes hourly wage)
(1) Hourly child care expenditures are corrected for selectivity using a bi-variate sample selection procedure	-0.046 (0.007)***	0.076 (0.025)***
(2) Hourly child care expenditures are not corrected for selectivity	-0.016 (0.009)*	0.043 (0.025)*
(3) Specification adds state-specific time trends	-0.059 (0.006)***	0.075 (0.026)***
(4) Specification adds state-specific time trends, education dummies interacted with year dummies, and education dummies interacted with the unemployment rate	-0.067 (0.006)***	0.121 (0.028)***
(5) Omit years 1990-1993	-0.034 (0.006)***	0.102 (0.027)***
(6) Omit years 2001-2004	-0.061 (0.008)***	0.051 (0.031)
(7) Omit the nine states that are not uniquely identified in the SIPP (Maine, Vermont, Iowa, North Dakota, South Dakota, Alaska, Idaho, Montana, and Wyoming)	-0.063 (0.006)***	0.086 (0.026)***
(8) Substitute states' weekly, private child care worker wages (/100) for the measure of hourly child care expenditures	-0.052 (0.020)***	0.028 (0.024)
(9) Limit the sample to never married mothers	-0.074 (0.009)***	0.111 (0.047)***
(10) Limit the sample to non-white mothers	-0.059 (0.011)***	0.066 (0.053)
(11) Limit the sample to mothers with at least one child ages 6-12	-0.059 (0.007)***	0.071 (0.029)**
(12) Use as the outcome variable an indicator for whether the mother worked and did not receive welfare in the previous year	-0.090 (0.007)***	0.212 (0.030)***
(13) Use as the outcome variable an indicator for whether the mother was employed full-time (35+ hours), full-year (48+ weeks), conditional on being employed	-0.023 (0.008)***	0.053 (0.031)*
(14) Estimate the model using OLS regression (linear probability model)	-0.067 (0.005)***	0.054 (0.019)***

Source: Author's calculations from the 1991-2005 March Current Population Survey (CPS).
Notes: Marginal effects are shown, along with robust standard errors (in parentheses). *, **, *** indicate that the coefficient is statistically significant at the 10%, 5%, and 1% levels, respectively. All models include controls for age; age-squared; marital status; non-white; educational attainment; whether the youngest child in the family is ages 3-5, ages 6-8, and ages 9-12; the presence of a child ages 13-17; the number of children ages 0-5; and all policy variables listed in Table 6. Estimates for these variables are available from the author upon request. Unless otherwise specified, all models include state fixed effects, year dummies, and month-in-sample dummies. All analyses are weighted using the March Supplemental Person Weight.

TABLE 8:
Comparison of Price and Wage Elasticities to Previous Studies

Study	Estimated Elasticity
Panel A: Elasticity of Employment with Respect to Child Care Expenditures	
All mothers:	
Current Study	
Replication model	-0.48
Full Model	-0.08
Anderson and Levine (2000)	-0.47
Mothers with a high school degree or less:	
Current study	-0.11
Anderson and Levine (2000)	-0.69/-0.41[1]
Mothers with at least one child ages 0-5:	
Current Study	-0.11
Connelly and Kimmel (2003a)	-0.76
Connelly and Kimmel (2003b)	-0.98
Han and Waldfogel (2001)	-0.50 to -0.73[2]
Anderson and Levine (2000)	-0.59
Panel B: Elasticity of Employment with Respect to Wages/Effective Taxes	
All mothers:	
Current Study	0.12
Hotz, Mullin, and Scholz (2005)	1.30
Looney (2005)	0.59
Meyer and Rosenbaum (2001)	1.07
Mothers with a high school degree or less:	0.13
Mothers with at least one child ages 0-5:	0.07

Notes: [1] This study reports price elasticities separately for mothers with less than a high school degree and those with exactly a high school degree. Therefore, I report both estimates. [2] These authors only provide a range of elasticity estimates.

TABLE 9:
Policy Simulations

Policy Scenario	Pr(employment)	Percentage Point Change
Baseline Predicted Probability	0.779	--
Panel A: Child Care Subsidy Simulations		
25% Subsidy	0.802	2.3
50% Subsidy	0.825	4.6
Subsidy Reducing Child Care Costs to 10% of Net-of-Taxes Hourly Wages	0.848	6.9
Targeted Nonlinear Subsidy a) \$0 - \$5,000: 100% subsidy b) \$5,001 - \$10,000: 85% subsidy c) \$10,001 - \$15,000: 70% subsidy e) \$15,001 - \$20,000: 55% subsidy f) \$20,001 - \$25,000: 40% subsidy g) \$25,001 - \$30,000: 25% subsidy h) Eligibility ends for families above \$30,000	0.868	8.9
Panel B: Wage and EITC Simulations		
25% Increase in Net-of-Taxes Hourly Wages	0.803	2.4
50% Increase in Net-of-Taxes Hourly Wages	0.818	3.9
Employment Change from a 50% Increase in Net-of-Taxes Hourly Wages for Mothers at the Following Predicted Wage Quartiles		
At or below the 25th percentile	0.615 to 0.660	4.5
At or above the 75th percentile	0.884 to 0.906	2.2
Panel C: Joint Subsidy and EITC Simulations		
50% Subsidy and Increase in Net-of-Taxes Hourly Wages	0.853	7.4
Targeted Nonlinear Subsidy and 50% Increase in Net-of-Taxes Hourly Wages	0.891	11.2

Notes: All simulations use estimates from the model in column (2b) in Table 6. The baseline predicted probability is the probability of employment based on the estimates from that model. With the exception of the predictions from the wage decile and EITC schedule exercises, percentage point change is in relation to the baseline probability of employment. Wage deciles and EITC regions are expressed in 2004 dollars.

APPENDIX A:
Variable Means for the SIPP Sample of Single Mothers: 199-1993 Panels

SIPP Panel (Wave)	**1990 (3)**	**1991 (3)**	**1992 (6), 1993 (3)**	**1993 (9)**
Dates of Data Collection	9/90 – 12/90	9/91 – 12/91	9/93 – 12/93	9/95 – 12/95
CPS "Data" Year(s)	1990	1991, 1992	1993, 1994	1995, 1996
Number of Observations	N = 1,664	N = 875	N = 2,677	N = 1,201
Demographics				
Age	32.23 (7.48)	32.48 (7.65)	32.29 (7.63)	32.95 (7.78)
Less than High School (%)	0.272 (0.445)	0.278 (0.448)	0.249 (0.432)	0.253 (0.434)
High School/GED (%)	0.439 (0.496)	0.410 (0.492)	0.424 (0.494)	0.419 (0.493)
Some College (%)	0.205 (0.403)	0.226 (0.418)	0.244 (0.429)	0.234 (0.423)
BA+ (%)	0.082 (0.274)	0.084 (0.278)	0.082 (0.274)	0.093 (0.290)
Widowed (%)	0.045 (0.209)	0.055 (0.228)	0.037 (0.189)	0.039 (0.194)
Separated (%)	0.202 (0.402)	0.180 (0.385)	0.191 (0.393)	0.174 (0.379)
Divorced (%)	0.378 (0.485)	0.389 (0.487)	0.367 (0.482)	0.373 (0.483)
Never Married (%)	0.372 (0.483)	0.374 (0.484)	0.403 (0.490)	0.412 (0.4920
Non-white (%)	0.394 (0.488)	0.363 (0.481)	0.383 (0.486)	0.377 (0.485)
Non-wage Income ($)	512.87 (669.81)	539.25 (700.39)	521.05 (634.66)	500.99 (809.44)
Child Ages 0-2 (%)	0.292 (0.455)	0.278 (0.448)	0.282 (0.450)	0.240 (0.427)
Child Ages 3-5 (%)	0.330 (0.470)	0.339 (0.473)	0.357 (0.479)	0.372 (0.483)
Child Ages 6-12 (%)	0.665 (0.472)	0.698 (0.459)	0.656 (0.475)	0.688 (0.463)
Child Ages 13-17 (%)	0.196 (0.397)	0.195 (0.397)	0.214 (0.410)	0.232 (0.422)
No. of Children Ages 0-2	0.343 (0.587)	0.316 (0.545)	0.318 (0.542)	0.259 (0.480)
No. of Children Ages 3-5	0.378 (0.580)	0.378 (0.563)	0.414 (0.602)	0.421 (0.589)
No. of Children Ages 6-12	0.899 (0.811)	0.995 (0.869)	0.914 (0.851)	0.991 (0.882)
No. of Children Ages 0-17	1.86 (1.01)	1.93 (1.07)	1.92 (1.05)	1.95 (1.02)
Unemployed Adult (%)	0.010 (0.100)	0.010 (0.101)	0.008 (0.093)	0.009 (0.099)
Urban Residence (%)	0.757 (0.428)	0.720 (0.449)	0.772 (0.419)	0.780 (0.413)
Southern Residence (%)	0.385 (0.486)	0.355 (0.478)	0.356 (0.479)	0.347 (0.476)
Employment/Child Care				
Labor Force Participation (%)	0.641 (0.479)	0.643 (0.479)	0.637 (0.480)	0.662 (0.473)
Uses Paid Child Care (%)	0.675 (0.468)	0.621 (0.485)	0.651 (0.476)	0.735 (0.441)
Pays for Child Care (%)	0.612 (0.487)	0.604 (0.489)	0.570 (.495)	0.568 (0.495)
Weekly Child Care Costs ($)	78.87 (53.47)	77.50 (47.08)	75.21 (51.75)	79.34 (76.49)
Cost per Hour of Work ($)	2.08 (1.51)	2.01 (1.35)	2.06 (1.66)	2.07 (2.02)
Share of Income Paid (%)	0.159 (0.240)	0.168 (0.167)	0.193 (0.378)	0.165 (0.201)

Source: Author's calculations from the SIPP Core File and Child Care Topical Module.
Notes: Standard deviations are in parentheses. All means are weighted using the final person weight from the fourth month of a given wave of data collection. Dollars are adjusted for inflation to reflect 2005 prices.

APPENDIX A (Continued):
Variable Means for the SIPP Sample of Single Mothers: 1996 and 2001 Panels

SIPP Panel (Wave)	**1996 (4)**	**1996 (10)**	**2001 (4)**
Dates of Data Collection	3/97 – 6/97	3/99 – 6/99	1/02 – 4/02
CPS "Data" Years	1997, 1998	1999, 2000, 2001	2002, 2003, 2004
Number of Observations	N = 2,605	N = 2,015	N = 1,985
Demographics			
Age	33.16 (8.22)	33.20 (8.33)	33.25 (8.12)
Less than High School (%)	0.216 (0.411)	0.192 (0.394)	0.196 (0.397)
High School/GED (%)	0.360 (0.480)	0.384 (0.486)	0.328 (0.469)
Some College (%)	0.339 (0.473)	0.328 (0.469)	0.364 (0.481)
BA+ (%)	0.083 (0.277)	0.094 (0.292)	0.110 (0.313)
Widowed (%)	0.042 (0.202)	0.035 (0.185)	0.034 (0.183)
Separated (%)	0.171 (0.377)	0.144 (0.351)	0.147 (0.354)
Divorced (%)	0.340 (0.473)	0.344 (0.475)	0.337 (0.472)
Never Married (%)	0.444 (0.497)	0.476 (0.499)	0.480 (0.499)
Non-white (%)	0.388 (0.487)	0.403 (0.490)	0.370 (0.483)
Non-wage Income ($)	448.73 (612.09)	407.85 (617.21)	406.19 (607.70)
Child Ages 0-2 (%)	0.244 (0.429)	0.230 (0.421)	0.259 (0.438)
Child Ages 3-5 (%)	0.356 (0.479)	0.335 (0.472)	0.329 (0.4700
Child Ages 6-12 (%)	0.680 (0.466)	0.692 (0.461)	0.685 (0.464)
Child Ages 13-17 (%)	0.217 (0.412)	0.208 (0.406)	0.205 (0.404)
No. of Children Ages 0-2	0.266 (0.492)	0.252 (0.484)	0.286 (0.509)
No. of Children Ages 3-5	0.406 (0.588)	0.373 (0.555)	0.368 (0.559)
No. of Children Ages 6-12	0.941 (0.841)	0.973 (0.865)	0.953 (0.827)
No. of Children Ages 0-17	1.88 (1.06)	1.86 (1.04)	1.87 (1.00)
Unemployed Adult (%)	0.013 (0.116)	0.013 (0.117)	0.008 (0.093)
Urban Residence (%)	0.810 (0.391)	0.848 (0.358)	0.771 (0.419)
Southern Residence (%)	0.385 (0.486)	0.384 (0.486)	0.372 (0.483)
Employment/Child Care			
Labor Force Participation (%)	0.743 (0.436)	0.778 (0.415)	0.777 (0.416)
Uses Paid Child Care (%)	0.762 (0.425)	0.775 (0.417)	0.756 (0.429)
Pays for Child Care (%)	0.573 (0.494)	0.534 (0.499)	0.541 (0.498)
Weekly Child Care Costs ($)	81.33 (71.43)	81.90 (73.43)	85.86 (80.07)
Cost per Hour of Work ($)	2.29 (3.28)	2.22 (2.77)	2.32 (2.99)
Share of Income Paid (%)	0.171 (0.232)	0.207 (0.547)	0.156 (0.187)

Source: Author's calculations from the SIPP Core File and Child Care Topical Module.
Notes: Standard deviations are in parentheses. All means are weighted using the final person weight from the fourth month of a given wave of data collection. Dollars are adjusted for inflation to reflect 2005 prices.

APPENDIX B:
Construction of Other Key Policy Variables and Theoretical Effects

The following describes the construction of other key policy variables that appear in the employment models, and discusses the theoretical effect of each on the work decision of single mothers.

First, to capture the effect of states' *earnings disregard policies*, I assigned to each single mother a predicted amount of annual disregarded earnings. This is accomplished by coding both the initial (fixed) component and the variable component of each state's disregard policy over the study period and then applying these rules to the earnings of employed single mothers. The fixed component refers to the first $30 of earnings, for example, while the variable component is 33% of the remainder. I code only the initial earnings disregard, omitting both the work expense and child care expense components. This process assumes that women are in the first four months of welfare receipt. After four consecutive months, states continued only the initial $30 disregard; after one year on welfare, individuals faced a 100 percent implicit tax on earnings. To predict disregarded earnings for non-working mothers, I estimated for each CPS survey a simple OLS regression of annual disregarded earnings on several exogenous demographic and human capital characteristics plus a vector of state fixed effects. Insertion of fixed effects controls for variation in states' disregard policies, especially in the period after welfare reform. Second, I capture the effect of *states' welfare reform efforts* through two dummy variables: enactment of any *statewide welfare reform* and *time limits*. The former is defined as the first reform policy a state passed under its waiver authority, including expanded earnings disregards, family caps, and work requirement sanctions. If a state did not implement any changes under a waiver, I take the implementation of its TANF program to be the first welfare policy considered. Both variables equal one starting in the year after the policies are executed, and they equal zero in the years prior to their implementation. In the year of implementation, I follow the standard practice of coding both variables as the faction of the year during which these policies were in effect. Finally, the *AFDC/TANF participation rate* was constructed by dividing each state's (adult) female caseload by the total number of female-headed households (with children under age 18).

Economic models suggest that every hour of work by women with children requires the use of substitute child care. *Child care expenditures* are viewed as a fixed cost of employment, such that each hour of care purchased in the market reduces the returns to work. A testable hypothesis is that an increase in hourly child care expenditures reduces the incentive to work. This is particularly applicable in the case of single mothers because child care costs often represent a significant fraction of their earnings. The theoretical effect of introducing an *EITC* is unambiguously positive. Eligibility for the program is confined to those with positive earnings, and EITC recipients experience an expanded budget set that makes work look more attractive at every wage level. In other words, the increased net-of-taxes wage rate for recipients previously not working leads to only a positive substitution effect. States' *maximum AFDC/TANF benefit* is predicted to decrease the incentive to work because the income effect from guaranteed benefits to those not working reduces the attractiveness of entering the labor force. This is particularly relevant for low-skilled workers in high benefit states, whose reservation wages are below the maximum welfare benefit. However, increasing the initial *earnings disregard* and lowering the *welfare phase-out rate* encourages the combination of welfare and work over pure welfare receipt (among those previously not working). This is because employed recipients can keep more of their earnings until they reach the break-even point. The theoretical effect of states' bundled *welfare reform efforts* is ambiguous, since states often implemented several policies are the same time. However, most of the individual policies—such as work requirements and benefit sanctions—are expected to increase the incentive to work. Finally, *time limits* are hypothesized to increase employment through two channels. One is purely the mechanical effect experienced by those who hit a state's time limit. The other is behavioral, and incorporates the assumption that forward-looking women will save their welfare benefits until they experience an employment shock.