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Close to Home: A Simultaneous Equations Model of the Relationship Between Child Care Accessibility and Female Labor Force Participation

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Abstract Rising rates of maternal employment among current and former welfare recipients have increased the use of non-parental child care. Little empirical work examines the relationship between women's labor supply and the geographic supply of child care. We combine census data with child care provider information for the state of Maryland to explore the relationship between female labor supply and the geographic supply of child care. OLS and 3-SLS equations are estimated, and the findings are consistent across each estimator: Women's labor supply is sensitive to the geographic supply of child care and vice versa. These results are important because states now spend significant money on quality improvement initiatives, many of which increase child care supply in low-income neighborhoods.

Keywords Child care · Female labor supply · Geography · Simultaneous equations

Introduction

Economic trends since the early 1970s coupled with the recent restructuring of the U.S. welfare system have induced women to increase their participation in the paid labor force. Indeed, 71% of women with children are currently in the labor force, up from 42% in 1970 (Cohen 2001; U.S. Bureau of Labor Statistics 2004). The enactment of the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) in 1996 has led to even more dramatic increases in employment among current and former welfare recipients,

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most of whom are women. Recent evidence suggests that two-thirds of welfare leavers are working, while fully 71% of single-parent leavers are currently employed (Loprest 2001).

These factors have taken the issue of child care into the national spotlight. Substitute child care is an important component of current federal and state efforts to move low-income parents into employment.[1] Although increased federal assistance has ostensibly made the purchase of child care more feasible, a lack of high-quality, affordable, and accessible child care still represents a critical barrier for low-income families, many headed by single women, to balance responsibilities at home and work (Baum 2002; Urban and Olson 2005). With child care costs sometimes consuming over one third of poor families' monthly income, there is reason to believe that these services pose an important financial and logistical burden that may not be offset by the gains to employment (Smith 2000).

It is, therefore, essential for scholars and policymakers to focus on the ways in which the price of child care affects maternal employment. There is a substantial literature that estimates the sensitivity of women's employment to changes in the price of child care. The evidence, whether from studies on the price of purchased care or actual subsidy programs, supports the hypothesis that higher prices for child care lead to lower maternal employment rates. Although the range of findings is quite large, there appears to be a convergence of recent price elasticities centering on −0.40.

Yet, little empirical work has focused on the relationship between non-monetary aspects of the cost of child care and women's employment. The geographic supply of child care represents one area that has not been explored. An underprovision of child care services within a reasonable distance from home indicates real costs in terms of travel time and other expenditures. Therefore, prospective workers in communities with less spatial accessibility face higher child care costs, and thus greater constraints on employment.

This research begins to address this gap by focusing on the supply of child care services in neighborhoods with varying socioeconomic characteristics. The analyses draw upon a unique dataset for the state of Maryland, employing the 2000 decennial census and information from the state's child care resource and referral (CCR&R) network. Specifically, it explores several hypotheses about the reciprocal relationship between the geographic supply of child care and female labor force participation rates. However, this study must surmount the classic identification problem: The theoretical framework suggests the presence of two dependent variables, both of which are likely to be endogenous. Drawing from economic models of labor supply, we exploit several theoretically valid sources of exogenous variation in female labor supply and child care supply to identify the system of equations. We then account for a number of demographic and socioeconomic attributes to estimate participation and child care supply equations with both ordinary least squares (OLS) and simultaneous equation techniques.

The next section reviews the literature on the price of child care. We then develop a theoretical framework based on economic models of female labor supply and child care service supply to test several hypotheses about their simultaneity. Following that, we introduce the data and empirical methods, present results, and conclude with a discussion of findings.

[1] The term "substitute" child care is used in this paper to denote sources of formal and informal child care that are not provided by the parent. Formal modes of child care include center-based or family-based providers, for example, while informal modes are typically defined as relatives, neighbors, and babysitters.

Review of Literature

As previously mentioned, much scholarly work focuses on the effect of monetary components of child care costs on women's work decisions. Non-experimental evidence on this relationship comes from two primary sources: studies on price effects and studies of actual child care subsidy programs. We consider each in turn.

A substantial literature demonstrates the effect of child care costs on the decision to enter the labor market. 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 sample selection bias 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 (Anderson and Levine 2000; Blau and Hagy 1998; Blau and Robins 1988, 1989, 1991; Baum 2002; Connelly 1992; Connelly and Kimmel 2001; Han and Waldfogel 1998; Jenkins 1992; Kimmel 1998; Liebowitz et al. 1992, 1988; Ribar 1992, 1995; U.S. General Accounting Office 1994). Although nearly every study found a negative relationship between child care costs and mothers' labor supply, the range of elasticities is large (from 0.06 to −1.36). However, there appears to be a recent convergence of estimates centering on −0.40.

A smaller body of research examined the effect of child care subsidies on labor supply. Using data from a day care subsidy program in Kentucky, Berger and Black (1992), compared the employment probabilities of individuals on the waiting list with those actually receiving a subsidy. Estimates from a probit model imply that the predicted employment rate for those on the waiting list was 85.5%, compared to 97.5% for those receiving child care assistance. Meyers et al. (2002) used data on a sample of California AFDC recipients and estimated subsidy effects through a two-stage model. The first stage models the probability of subsidy receipt, conditional on using market child care. Using the predicted probabilities from the first stage, the authors then estimated a discrete choice employment equation. Policy simulations imply that as the probability of subsidy receipt goes from 0.10 to 0.60, the employment probability increases from 0.30 to 0.81. A final piece of evidence comes from a recent study by Blau and Tekin (2007), who used data on a nationally representative sample of unmarried women with at least one child under age 13. Results suggest that child care subsidies are associated with a 0.05–0.11% point increase in the probability of employment.

Very little is known about the effect of spatial accessibility—or the physical presence of child care providers within communities. To our knowledge, there are only two studies conducted in the United States that apply multivariate techniques to analyze the effect of child care accessibility.[2] The first, by Presser and Baldwin (1980), used the 1977 Current Population Survey (CPS) to determine whether child care availability constrains the labor supply of women with young children. They found that women who would benefit most from employment (young, unmarried mothers with limited education) are also more likely to report that inaccessible child care restricts their labor supply. One drawback of this research is that it did not explicitly model female labor supply as a function of service availability.[3] The other study, by Stolzenberg and Waite (1984), found that women's labor

[2] A study conducted in Wales using a reduced form specification finds that an increase of 10 child care providers within a community leads to a 3% increase in the female labor supply (Webster and White 1997b). Another analysis from the Netherlands shows that female labor supply increases significantly when the supply of accessible subsidized child care is raised (Van Dijk and Siegers 1996).

[3] This study models the social and economic correlates of "feeling" constrained, rather than a quantitative measure of service accessibility.

supply, conditional on the presence and number of young children, is very sensitive to the availability of substitute child care. However, the authors' definition of service accessibility assumed that the number of paid child care workers in a geographic area is a valid measure of service capacity and that these individuals live and work in the same area.

Theoretical Model

This research tests two principal hypotheses regarding aggregate household decision-making and child care provider behavior. Specifically, we estimate the response of women's labor supply to changes in the level of child care accessibility. It is hypothesized that increases in the geographic supply of child care will be associated with higher levels of female labor force participation. We also estimate the effect of changes in aggregate female labor supply on child care accessibility. It is postulated that communities with higher female labor supply rates will have a greater number of accessible child care slots. The theoretical framework used to test these hypotheses is based on economic models of female labor supply (Becker 1981; Blau and Hagy 1998; Blau and Robbins 1988; Brandon 2000) and child care firm behavior (Edwards et al. 1996; Queralt and Witte 1998; Webster and White 1997a).

Female Labor Supply

The simple model described here is intended to highlight issues for estimating single-period models of female labor force participation using cross-sectional data. Although the model is estimated on aggregate data, it focuses on households with children who require continuous care throughout the day, either from the mother, a spouse, or other child care provider. Child care providers are assumed to be a constellation of formal (paid) and informal (unpaid) sources. Standard models imply that not all households have access to paid providers, but they do have access to potential unpaid sources of care. Households have preferences over child care quality (Q), leisure (L), other consumption goods (G), and the employment state (E):

$$U = U(Q, L, G, E, X, \varepsilon_1), \tag{1}$$

where $\mathbf{X}$ is a matrix of measured exogenous variables that influence utility, and ε_1 represents unobserved components.[4] The observed labor state, E, represents various employment—child care mode—payment type combinations (Blau and Hagy 1998).

Child care quality is represented through a production function that includes maternal child care time (M), non-parental child care time (N), and the number and age structure of children in the household (C). Recall that non-parental care can be specified as formal

[4] To make the framework tractable, we allow a few simplifying assumptions: (a) the decisions of other household members, including the spouse, are exogenous to the work decision of the mother; (b) that all of the mother's non-work time is spent caring for her children, which removes the complication of having to distinguish between leisure and home production of child care; (c) that one type of non-mother child care is used; and (d) that the price of child care reflects its quality.

(paid) and informal (unpaid) services, all of whose quality depends on a bundle of characteristics, *B*, chosen by the household. The quality production function, then, is written as:

$$Q = Q(M, N, C, B, X, \varepsilon_2), \tag{2}$$

where ε_2 is a vector of unobserved components that produce child care quality.

Households maximize the utility function specified in (1) subject to the following cost constraint, assuming no savings or borrowing:

$$\alpha + \pi + \mathrm{wh} - \tau - \mathrm{p_{AQ}}AQ - \mathrm{p_G}G - \ \mathrm{f_E}T = 0, \tag{1}$$

where α denotes all transfer payments to a household; π represents sources of non-labor income (including income from the spouse); w is the mother's wage rate; h is her hours of work; τ denotes tax payments by the household; $\mathrm{p_{AQ}}$ is the price of child care for *A* number of slots at *Q* level of quality; $\mathrm{p_G}$ is the price of all other consumptions goods *G*; and $\mathrm{f_E}$ represents fixed employment costs (e.g., travel, time, and other expenditures) from *T* trips to consume child care services. The level of child care supply in a given neighborhood—also referred to as child care accessibility—can be rewritten as A_{ki} to denote the number of slots, *k*, provided by firm *i*:

$$Ak_i = A_{11},\ A_{21}, \ldots,\ A_{n1};\ A_{12},\ A_{22}, \ldots, A_{n2};\ A_{1m},\ A_{2m}, \ldots, A_{nm}. \tag{4}$$

This framework suggests that the distance a mother travels to a given child care provider represents real costs in terms of travel time and other expenditures. Therefore, individuals in communities with less spatial accessibility face higher child care costs and thus greater constraints on employment. In some communities, price (distance) may be so great that households choose not to consume child care services. Other communities face not only large variations in the number of providers close to home, but also in the price (monetary and non-monetary) at which they can purchase.

It should also be noted that the geographic boundaries of local child care markets are unknown, but they most likely vary according to population density, income level, and place of employment (Gordon and Chase-Lansdale 2001). Prior studies emphasize that parents, especially low-income parents, rate proximity to home as an important criterion in selecting a child care arrangement (U.S. Department of Education 1995). Although home-child care journeys are short (Meyers-Jones and Brooker-Gross 1996), they are made overwhelmingly by women (Blumen 1994; Rosenbloom 1988, 1993). Furthermore, one study found the mother's trip to child care is shorter than her trip from child care to work, but journeys to child care add 28% more time to the total commute (Michelson 1985). This suggests that a model based on familial preferences assumes that all else equal, mothers will consume child care services close to their residential area.

Child Care Firm Behavior

Economic models of child care supply assume that providers have two general motivations for entering the market: making profits and deriving satisfaction from working with young children (Blau 2001). All providers have both motivations, but each places its own value on the pecuniary versus non-pecuniary benefits. The framework assumes that providers choose the quality of care by selecting the amount of labor, including its skill level, and other operational inputs.[5] As a result, it is assumed that the price per hour of care is a function of its quality, which determines the provider's revenue and costs. Providers choose the level of quality that maximizes their satisfaction under competitive market conditions.

The model of child care supply developed here assumes that child care providers are profit-maximizing firms operating in a competitive market. For-profit firms can have non-pecuniary motives, but it certainly is the case that to survive in the market, child care providers must focus heavily on profit. Research on the locational preferences of child care centers highlights this point. For example, Kahn and Kamerman (1987) found that for-profit child care providers identify profitable markets as those "near a major highway, a location between a middle-class residential area and a commercial area, a community with high female labor force participation rates and husband/wife families with two earners and income more than 50% above the median income" (p. 105).

Recent empirical work supports this assertion, and provides the groundwork for an empirical model based on theoretical predictions. It is well documented that substantial variation exists in the geographic supply of child care (Kreader et al. 2000; Queralt and Witte 1998; Truelove 1996). Many of the differences in child care supply can be explained by community-level differences in demographic attributes, social and economic indicators, and urbanicity.

In formal terms, consider the simple model in which the primary inputs to child care production are labor and capital. Firms, whether they be family- or center-based, are motivated by profits, π, denoted by:

$$\pi = pA - wL - rK, \tag{5}$$

where A is a matrix of child care services offered by each firm; L is a matrix of child care workers with varying skill sets; K represents the raw materials (floor space, etc.) that go into producing child care services; and r is the cost of rent, overhead, and other raw materials. Assuming that firms have child care service production functions of the form $A(L, K)$ for each of its a services, then providers demand child care labor according to

$$L = L(p,\ w,\ r), \tag{6}$$

which yields the firm's profit function:

$$\pi = \pi(p,\ w,\ r). \tag{7}$$

The supply function, therefore, is the partial derivative of the profit function with respect to price, written as:

$$\frac{\partial \pi}{\partial p} = a(p, w, r). \tag{8}$$

The above model suggests that child care providers, like most in the service industry, do not choose the quantity of the service. They instead choose the quality of the care, and the price is determined within the market. The supply of child care slots is determined by the number and characteristics of consumers who prefer the bundle of services offered by the provider compared to others available in the market (Blau 2001). A number of straightforward predictions are derived from the model: It suggests that child care firms

[5] Highly trained workers who possess advanced education in child development are more skilled and provide higher-quality care, but they also cost more to employ. Child care services provided to smaller groups or special populations are deemed higher quality, but cost more to provide.

decrease supply when adverse demand conditions are present (e.g., depleted financial resources in a neighborhood and barriers to entry from competitors or regulations), and supply increases when positive demand conditions are present (e.g., higher female employment rates, households with young children, and the use of child care subsidies).

Empirical Approach

Data Sources

Data for this research come from two sources: LOCATE: Child Care database from the Maryland Committee for Children and the 2000 U.S. Decennial Census. LOCATE: Child Care is a compilation child care resource referral (CCR&R) database system owned and managed by the Maryland Committee for Children. It includes information on all licensed services in child care centers, infant programs, Head Start programs, nursery schools, pre-kindergarten programs, and summer programs. The database also provides information on regulated family care providers, which are defined as services for up to eight children under age 13 in place of parental care. To be classified as such, family day care provides fee-based services for less than 24 h in a residence other than the child's home. This study uses the December 2000 wave of data on the number of center-based, family-based, and Head Start child care slots in Maryland.[6] As of the December 2000 survey, there were 10,808 and 75,191 family-based providers and slots, respectively. There were 1,334 center-based child care services, providing 80,419 slots. Finally, the survey records 235 Head Start providers with 9,542 slots.[7]

Demographic and socioeconomic data for Maryland are taken from the 2000 U.S. Decennial Census' Summary Tape File 3A (STF3A). This file contains sample data weighted to represent the total population, as well as 100% counts for total persons and total housing units. The geographic unit of analysis in this study is the census tract. Census tracts contain 2,500–8,000 relatively homogenous residents with respect to demographic, social, and economic attributes (U.S. Bureau of the Census 2000).

Geographic information systems (GIS) technology allowed us to geocode the location of center-based, family-based, and Head Start providers to specific street addresses and census tracts based on 2000 U.S. Decennial Census definitions. After each child care provider was assigned a census tract identification number, the data were merged with tract-level information from the decennial census. Thus, the analytic sample contains detailed information on child care providers and the demographic, social, and economic characteristics for all 1,218 census tracts in Maryland.[8]

[6] The dependent variable in our equations, however, only uses slots from center-based and family-based providers. Head Start is used as a control variable in the child care supply equation.

[7] The extent to which our data capture the total paid child care market in Maryland is unknown. However, estimates from recent national studies suggest that the two modes of child care observed in this study capture approximately 50% of the total (paid and unpaid) child care market (Capizzano et al. 2000; Ehrle et al. 2001).

[8] The final analysis sample contains full information on 1,128 census tracts. Deletions were made for the following purposes: five census tracts were deleted for having zero population or zero working age females; three census tracts were deleted for having zero for the female labor force participation rate; 43 and 35 mutually exclusive census tracts, respectively, were deleted because they were not able to be geocoded; and an additional four observations were deleted due to missing data.

Recent evidence suggests that child care markets may be smaller than originally conceptualized.[9] Low-income working mothers have stressed the importance of finding child care conveniently located near school, work, or home (Henly and Lyons 2000). This is corroborated by findings from the National Household Education Survey (NHES): fully 57% of all parents with young children indicated that services close to home was highly influential in choosing a child care arrangement (U.S. Department of Education 1995). Conveniently located services were very important to nearly 70% of low-income parents, compared to 50% among wealthy parents. Finally, a study in Cuyahoga County, Ohio found that child care subsidy users travel approximately 2 miles to center-based providers and 1.5 miles to family child care homes (Bania et al. 2000).

Estimation Approach

The econometric task in this paper is to estimate the simultaneous relationship between female labor supply and the geographic supply of child care, measured at the census tract level. Therefore, the analytic framework contains two dependent variables. First, a measure of women's labor supply is calculated for each census tract by dividing the number of females ages 16 and over in the labor force by the total number females ages 16 and over.[10] Second, total child care supply is measured as a simple continuous variable of the number of center-based and family-based child care slots in a given census tract.

Recall the first hypothesis states that greater levels of child care supply are associated with increased female labor force participation rates, *ceteris paribus*. Drawing upon the theoretical model and previous empirical work on labor supply for guidance, we postulate that women's aggregate labor supply is a function of child care supply, the cost of such services, local labor market conditions, and other economic and demographic characteristics of neighborhoods.[11] This suggests estimating the following structural labor supply equation by ordinary least squares (OLS):

[9] The census tract, rather than zip codes or counties, is used as the unit of analysis for theoretical and empirical reasons. First, recall that the overarching hypothesis of this research is that there is a reciprocal relationship between the characteristics of families and the social, economic, and institutional environment in which they reside. This provides a framework for examining community-level factors related to the level of formal child care availability, as well as the extent to which service accessibility has an effect on the labor supply of working-age women. It is expected that the precision with which this relationship can be observed and modeled increases as the geographic area becomes smaller and thus more homogenously defined. Although there is no consensus on the size of local child care markets, most studies opt to define larger rather than smaller geographic boundaries. Gordon and Chase-Lansdale (2001), for example, argue that a 30-mile radius around a particular zip code might closely approximate center-based markets. However, Queralt and Witte (1998) state that their decision to focus on census tracts was made in consultation with workers at the Child Care Search Resource and Referral Agency. This agency apparently also provided evidence that census tracts best proxy what is commonly known as the boundaries of child care markets.

[10] Note that this participation measure is an aggregate for an entire census tract. As a result, we are unable to estimate models for part-time and full-time work status across individuals, nor can we examine participation at the extensive (decision to work) and intensive (hours) margins.

[11] Variables in the labor supply equation are common in the literature. To guide the selection of our variables, we carefully reviewed the models in Blau and Hagy (1998), Anderson and Levine (2000), and Meyers et al. (2002). These studies used individual-level data, and so our task was to create a set of control variables using aggregate data that mirror as closely as possible the controls in the above studies. Specifically, we create variables that reflect women's underlying preferences for work (age, disability rate and race), their human capita (educational attainment), characteristics of the local child care market (average price and subsidy-acceptance rate), and potential barriers to employment (non-English speaking families and those that do not own a car).

$$\begin{aligned} WORK_{\mathrm{i}} = \alpha_{\mathrm{i}} &+ \beta_1 SUPPLY_{\mathrm{i}} + \beta_2 PRICE_{\mathrm{i}} + \beta_3 SUBSIDY_{\mathrm{i}} + \beta_4 UNEM_{\mathrm{i}} + \beta_5 BLACK_{\mathrm{i}} \\ &+ \beta_6 AGE_{\mathrm{i}} + \beta_7 FEMEDU_{\mathrm{i}} + \beta_8 MALEEDU_{\mathrm{i}} + \beta_9 INCOME_{\mathrm{i}} + \beta_{10}(INCOME)^2_{\mathrm{i}} \\ &+ \beta_{11} DISABILITY_{\mathrm{i}} + \beta_{12} ENGLISH_{\mathrm{i}} + \beta_{13} NOCAR_{\mathrm{i}} + \pi_{\mathrm{i}} + \varepsilon_{\mathrm{i}} \end{aligned} \tag{9}$$

where, *WORK* is the female labor force participation rate; *SUPPLY* is the total geographic supply of child care slots; *PRICE* is average weekly price of child care, across all providers and age groups; *SUBSIDY* is the percentage of child care providers in each census tract that accept CCDF subsidies; *UNEM* is the census tract unemployment rate; *BLACK* is the proportion African American; *AGE* is the ratio of the number of women ages 40–64 to women ages 16–39; *FEMEDU* is the proportion of females ages 25 and over with at least a BA degree; *MALEEDU* is the proportion of males ages 25 and over with at least a BA degree; *INCOME* is the median family income; $(INCOME)^2$ is a quadratic expression of the income variable; *DISABILITY* is proportion of individuals ages 16–64 with a disability; *ENGLISH* is the proportion of individuals ages 18–64 who speak English *not well* or *not at all*; and *NOCAR* is the proportion of households without a vehicle.

Note that α and $\beta_1,\ldots,\beta_{13}$ are the parameter estimates, π represents a matrix of county fixed effects, and ε is an iid $\sim N(0,\sigma^2)$ error term. County fixed effects capture unobserved county-specific attributes that affect the supply of child care and women's employment.[12] The effect of child care supply—the regressor of primary interest—which is denoted by β_1, represents the average participation effect for each additional child care slot, holding all other variables constant. The framework suggests that the coefficient on child care supply should be positive.

The second hypothesis implies that the labor supply of working-age women in a neighborhood influences the supply of child care services, ceteris paribus. We posit that child care accessibility will be a function of female labor supply, the cost of child care services, demand indicators such as the number of young children, sources of competition from Head Start and informal providers, and other economic and demographic attributes of neighborhoods.[13] Therefore, a structural child care supply equation can be estimated by OLS, taking the form:

$$\begin{aligned} SUPPLY_{\mathrm{i}} = \mu_{\mathrm{i}} &+ \gamma_1 WORK_{\mathrm{i}} + \gamma_2 PRICE_{\mathrm{i}} + \gamma_3 KIDS_{\mathrm{i}} + \gamma_4 ONEPAR_{\mathrm{i}} + \gamma_5 TWOPAR_{\mathrm{i}} \\ &+ \gamma_6 SAMEHOUSE_{\mathrm{i}} + \gamma_7 INCOME_{\mathrm{i}} + \gamma_8 (INCOME)2_{\mathrm{i}} + \gamma_9 HSTART_{\mathrm{i}} \\ &+ \gamma_{10}(HSTART * INCOME)_{\mathrm{i}} + \gamma_{11} INFORMAL_{\mathrm{i}} + \gamma_{12} WORKHOME_{\mathrm{i}} \\ &+ \pi_{\mathrm{i}} + \nu_{\mathrm{i}} \end{aligned} \tag{10}$$

where, *SUPPLY* is the total geographic supply of child care slots; *WORK* is the female labor force participation rate; *PRICE* is average weekly price of child care, across all

[12] One such unobserved factor is child care quality. Quality likely influences child care prices, which in turn influences the employment decision. Fixed effects control for many of these unobserved factors, at least as they exist at the county-level. Ideally, we would like to use fixed effects at a lower level of aggregation. Census-tract-level fixed effects are not possible give that this is the unit of analysis. However, we did estimate models that control for census tract that reside in specific cities or major towns, as well as others that code census tracts along an urban-rural dimension. None of these alterations substantially changed our results.

[13] As in the labor supply model, we consult recent work on child care supply to select and justify the inclusion of control variables in (10). Specifically, we pay particular attention to Webster and White (1997a; b) and Queralt and Witte (1998).

providers and age groups; *KIDS* is the total number of children ages 0–5; *ONEPAR* is proportion of children ages 0–5 living with a single mother, in the labor force; *TWOPAR* is the proportion of kids ages 0–5 living with two parents, both in the labor force; *SAMEHOUSE* is the proportion of individuals ages 5 and over living in the same house for the past 5 years; *INCOME* is the median family income;[14] $(INCOME)^2$ is a quadratic expression of the income variable; *HSTART* is a dummy variable for whether or not a census tract contains a Head Start provider; (*HSTART* * *INCOME*) is an interaction term between the Head Start dummy and median family income, which should account for the large number of Head Start providers located in low-income communities; *INFORMAL* is a proxy for neighborhood informal child care providers, measured as a ratio of the number of children ages 0 to 5 to the number of adults ages 55–74; and *WORKHOME* is the proportion of employed individuals ages 16 and over who work from home.

Note that μ and $\gamma_1,\ldots,\gamma_{12}$ are the parameter estimates, π represents a matrix of county fixed effects, and v is an iid $\sim N(0,\sigma^2)$ error term. The coefficient on female labor force participation—the regressor of primary interest—which is denoted by γ_1, represents the average effect of a unit increase in aggregate labor force participation on child care supply, holding all other variables constant. The economic framework suggests that the coefficient on this variable should be positive.

The separate and joint interpretation of the structural parameters in Eqs. 9 and 10 are hampered because it is unclear whether child care supply causes increased female labor supply or whether higher levels of female labor supply lead to more child care slots. The theoretical model suggests these two are most likely determined simultaneously. Utilizing OLS to model systems of equations in which there are two or more endogenous variables yields biased estimates of the parameters and causes the standard errors to become unreliable. The problem of simultaneity bias is observed through the dependence of the total geographic supply of child care slots *(SUPPLY)* on v, the error term from the child care supply equation, and the female labor force participation rate *(WORK)* on ε, the error term from the labor supply equation.

The appropriate method for estimating systems of simultaneous equations in which each function contains an endogenous explanatory variable among the dependent variables is three-stage least squares (3-SLS; Greene 2003). This method defines both the labor supply function and the child care service supply function as structural equations. Each dependent variable, in this case the female labor force participation rate and the total supply of child care slots in the area, is assumed to be endogenous to the system, and as a result, treated as correlated with the error terms. Unless otherwise specified, all other explanatory variables in the system are considered exogenous and uncorrelated with the error terms. This study exploits several theoretically valid sources of exogenous variation in female labor supply and child care supply to identify the system of equations. Although most of the variables in each equation are excluded from the other—and therefore are considered exogenous—we would like to focus on a few variables in each equation as being particularly effective instruments.

As for the labor supply equation, the subset of instrumental variables includes the proportion of individuals ages 16–64 with a disability *(DISABILITY)*, the proportion of individuals ages 18–64 who speak English *not well* or *not at all (ENGLISH)*, and the proportion of individuals without a vehicle (*NOCAR)*. That is, each variable is expected to affect female labor force participation but not child care supply. Economic models of labor

[14] Ideally, we would prefer to use income net of the woman's earnings, but that is not possible for our data source because family income depends in part on whether the woman works. This problem would likely be more severe if we were using individual data rather than aggregate data.

supply (and much empirical evidence) suggest that a disability (Anderson and Burkhauser 1985; Danziger et al. 2000; Gordon and Blinder 1980; Hanoch and Honig 1983; Sickles and Taubman 1986; Parsons 1980; Stern 1989; Wolfe and Hill 1995) a lack of English proficiency (Carliner 1996; Grenier 1984; McManus et al. 1983; Mora 1998; Trejo 1997), and a lack of reliable transportation (Blumenberg 2002; Ong 1996, 2002; Stoll and Raphael 2000) decrease the probability of employment, and contingent on being employed, decrease the number of hours of work. In other words, these variables pose significant *barriers to employment*: The greater their prevalence in a given neighborhood, the lower the expected level of labor force participation. Thus, the coefficient on each variable is hypothesized to be negative in the female labor supply equation.

Turning to the child care supply equation, the four instrumental variables are a dummy variable for whether or not a census tract contains a Head Start provider (*HSTART*), an interaction term formed by multiplying Head Start presence by median income (*HSTART* * *INCOME*), the ratio of the number of children ages 0–5 to the number of adults ages 55–74 (*INFORMAL)*, and the proportion of employed individuals ages 16 and over who work from home *(WORKHOME).* That is, each variable should affect child care supply (in this case, the number of center- and family-based slots) but should not have a direct effect on female labor supply. Recall that the economic model of child care supply developed here assumes that child care providers are profit-maximizing firms operating in a competitive market. For-profit firms can have non-pecuniary motives, but the model implies that to survive in the market, child care providers must focus heavily on profit. Neighborhood factors that signal lower profits will dissuade for-profit providers from locating in such neighborhoods, leading ultimately to decreased child care supply. The three instruments used in this study—the presence of Head Start providers, the extent of informal caregivers, and individuals working from home—represent child care options from other sources, and therefore signal potential *barriers to entry* into the formal child care market. In other words, increased supply from these sources is hypothesized to reduce the supply of for-profit providers. Indeed, the empirical evidence suggests that Head Start providers (Edwards et al. 1996), informal child care providers (Queralt and Witte 1998), and parental caregivers (Blau and Robins 1991; Connelly and Kimmel 2001) crowd out for-profit child care providers. Thus, the coefficient on each variable is expected to be negative in the child care supply equation.

Empirical Results

Summary Statistics

Table 1 presents a descriptive look at the tract-level child care and census data for the state of Maryland. The average neighborhood contains a total supply of 135 child care slots, which includes family- and center-based slots. There are approximately equal numbers of family- and center-based child care slots, 66 and 69 on average, respectively. Child care providers charge $102 per week on average, and there is, likewise, a great deal of variation across neighborhoods: a minimum price of $55 per week and a maximum price of $285.[15]

[15] It is important to note that these figures reflect the weekly price of child care, averaged across all family- and center-based providers in a given census tract. The variable also reflects differential pricing across different age groups. In other words, the price variable reflects the average aggregate cost of child care services across all providers and age groups served in a given neighborhood.

Table 1 Descriptive statistics for analysis variables (N = 1,128 census tracts)

Variable	Mean	*SD*	Min	Max
Panel A: Female labor force participation				
Total child care supply (slots)	135.43	115.05	4	831
Family-based child care supply	66.25	55.63	0	354
Center-based child care supply	69.18	90.40	0	648
Child care cost (average weekly price, $)	102.39	24.86	55.0	285
Providers accepting subsidies (%)	69.80	22.96	0	100
Unemployment rate (%)	5.47	5.15	0	34.21
African American (%)	30.14	32.51	0	99.77
Ratio of women ages 40–64 to women ages 16–39	1.06	0.41	0.02	3.91
Women ages 25+ with at least a BA degree (%)	27.08	17.35	0	85.97
Males ages 25+ with at least a BA degree (%)	30.34	21.0	0	93.43
Median family income ($)	54,726	23,728	7,944	200,001
Disability rate (ages 16–64, %)	18.07	7.68	0	63.07
Individuals ages 18–64 who speak English *not well* or *not at all* (%)	2.22	3.99	0	47.53
Households without a vehicle (%)	12.54	15.29	0	84.98
Panel B: Child care supply				
Female labor force participation (%)	62.04	8.96	19.13	84.96
Child care cost (average weekly price, $)	See above	See above	See above	See above
Number of children ages 0–5	363.71	225.24	12.0	1734.0
Median family income ($)	See above	See above	See above	See above
Children ages 0–5 living with an employed single mother (%)	18.01	15.07	0	88.46
Children ages 0–5 living with two employed parents (%)	41.26	17.58	0	100
Individuals ages 5+ living in the same house for at least 5 years (%)	57.03	11.92	7.54	80.45
Head Start provider (dummy variable)	0.17	0.38	0	1
Ratio of children ages 0–5 to adults ages 55–74	0.71	3.44	0.06	115.0
Employed individuals working from home (%)	3.15	2.23	0	13.14

Table 1 also shows that 70% of providers accept subsidies, and nearly one fifth of neighborhoods contain at least one Head Start provider. Turning to the demographic and economic variables, we find that nearly two thirds of female ages 16 and over are in the labor force. The average neighborhood has an unemployment rate of 5.5% and a median family income of approximately $55,000. African Americans comprise 30% of the population in Maryland, and nearly equal proportions of women and men ages 25 and over have at least a college degree, 27% and 30%, respectively.

Estimation Results

Table 2 presents the OLS estimation results from the structural female labor supply and child care supply equations. The first two models (Models 1 and 2) show the parameter

Table 2 OLS estimates of the female labor force participation and child care supply equations

Equation	Model 1	Model 2	Model 3	Model 4
Female labor force participation				
Total child care supply	0.005 (2.92)***	0.003 (2.15)**	–	–
Child care cost	−0.021 (1.63)	−0.037 (2.72)***	–	–
Providers accepting subsidies	−0.011 (1.20)	−0.002 (0.26)	–	–
Unemployment rate	−0.293 (3.70)***	−0.255 (3.00)***	–	–
African American	0.110 (13.14)***	0.110 (10.34)***	–	–
Ratio of women ages 40–64 to women ages 16–39	−9.20 (11.14)***	−8.311 (9.39)***	–	–
Women ages 25+ with at least a BA degree	0.208 (4.80)***	0.236 (5.31)***	–	–
Males ages 25+ with at least a BA degree	−0.144 (3.55)***	−0.149 (3.57)***	–	–
Median family income	0.0003 (6.70)***	0.0002 (3.42)***	–	–
(Median family income)^2	−1.300e-09 (5.59)***	−8.870e-10 (3.71)***	–	–
Disability rate	−0.109 (1.92)*	−0.109 (1.91)*	–	–
Individuals ages 18–64 who speak English *not well* or *not at all*	−0.027 (0.55)	−0.099 (1.82)*	–	–
Households without a vehicle	−0.192 (5.28)***	−0.212 (4.80)***	–	–
Constant	63.633 (27.71)***	61.795 (25.80)***	–	–
Child care supply				
Female labor force participation	–	–	0.860 (1.94)*	0.851 (1.91)*
Child care cost	–	–	0.161 (1.01)	0.556 (2.64)***
Number of children ages 0–5	–	–	0.272 (14.48)***	0.260 (13.53)***
Children ages 0–5 living with an employed single mother	–	–	0.059 (0.22)	0.141 (0.49)
Children ages 0–5 living with two employed parents	–	–	−0.021 (0.10)	−0.115 (0.50)
Individuals ages 5+ living in the same house for at least 5 years	–	–	0.752 (2.29)**	0.790 (2.20)**
Median family income	–	–	0.001 (1.95)*	0.001 (1.67)*

Table 2 continued

Equation	Model 1	Model 2	Model 3	Model 4
(Median family income)^2	–	–	−8.850e-09 (2.41)**	−8.320e-09 (1.98)**
Head Start provider	–	–	−39.321 (2.49)**	−46.232 (2.87)***
(Head Start * Median family income)	–	–	0.001 (2.47)**	0.001 (2.86)***
Ratio of children ages 0–5 to adults ages 55–74	–	–	−1.259 (4.05)***	−1.109 (3.36)***
Employed individuals working from home	–	–	−0.102 (0.08)	−2.492 (1.57)
Constant	–	–	−107.400 (2.74)***	139.166 (3.54)***
County fixed effects (No/Yes)	No	Yes	No	Yes
R^2	0.589	0.607	0.334	0.359
F-statistic for subset of IV's/All exclusions	17.62/59.69	15.39/46.35	5.61/35.85	6.14/34.88

Notes: Absolute value of *t*-statistics is in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

estimates and *t*-statistics for the labor supply equation, and the last two models (Models 3 and 4) show the same results for the child care supply equation. Both equations are first estimated without county fixed-effects, and then with them in the second equation. Significance levels for the *t*-statistics have been adjusted for heteroskedasticity using robust standard errors.

Beginning with the labor supply equation, we find few differences between Models 1 and 2, and so we focus on the coefficients in the latter model (with county fixed-effects). Most of the coefficients have the theoretically correct sign, and all are statistically significant. We are concerned primarily with testing the null hypothesis that the geographic supply of child care has no association with women's labor force participation. The set of regressions in Table 2 suggests that we can reject the null in both models. In fact, the coefficient on total child care supply implies that a 10-slot increase is associated with a 0.03% point increase in the female labor force participation rate, ceterus paribus, while an increase of 100 slots should raise the participation rate by 0.30% points.

As expected, female labor supply is fairly sensitive to the neighborhood cost of purchasing child care services, with higher prices leading to reductions in labor supply. Specifically, we find that a $100 increase in the price of child care is associated with a 3.7% point decrease in a neighborhood's female labor force participation rate. It is interesting to note the change in statistical significance for the cost-of-services variable between Models 1 and 2. Previous research treats the price of child care as endogenous in labor supply equations, and this is likely the case in the present study. As previously stated, we deal with the endogeneity by including county-level fixed effects in the models, which likely mitigates the effect of unobserved factors that influence local child care prices and female employment.[16]

With the exception of the proportion African American, the remaining variables in the labor supply equation have the theoretically correct sign for their coefficients. Higher unemployment rates, older female age structures, and greater proportions of highly educated males are associated with lower female labor force participation rates. Conversely, neighborhoods with greater numbers of highly educated females and families with larger incomes are associated with increased female labor supply. The coefficient on the quadratic income variable is negative, suggesting that female labor rates actually begin to decline in neighborhoods with higher earning families.[17] Finally, estimates on the subset of instrumental variables for female labor supply—the proportion disabled, English non-proficient, and households without a vehicle—are all negative, as predicted by the theoretical model, and are statistically significant. In fact, a joint F-test on the null hypothesis that all three variables are simultaneously equal to zero is rejected unequivocally ($F = 15.39$; $p < 0.01$). We conclude that these variables form an important constellation of *barriers to employment* and, therefore, are robust sources of identifying variation in female labor supply.

Turning our attention to the child care supply equation in Table 2 (Models 3 and 4), we again observe few differences between the models with and without county fixed-effects, and so we will concentrate the discussion on the latter model. Most of the coefficients have the theoretically correct sign and most are statistically significant. We are concerned primarily with testing the null hypothesis that female labor supply is not associated with the geographic supply of child care. As Model 4 shows, we are able to reject the null, and

[16] This perhaps explains why the estimate of child care price becomes statistically significant when accounting for such unmeasured factors.

[17] This is most likely due to the presence of husbands in certain neighborhoods who earn enough to allow the female to stay home and specialize in child care production.

the coefficient on female labor supply implies that a 1% point increase in a neighborhood's female labor force participation rate is associated with approximately one (0.85) additional child care slot, ceterus paribus.

Also noteworthy is the finding that a neighborhood's supply of child care slots is quite sensitive to the price for such services. As expected, higher average prices for child care services are associated with greater geographic supply: The coefficient on price suggests that a $10 increase in the average cost of child care within a given neighborhood is expected to increase supply by nearly six slots. Neighborhoods with greater numbers of children ages 0–5 also have a larger stock of child care slots, and the proxy for residential stability—individuals living in the same house for the past 5 years—is positively correlated with child care supply.

Estimates on the subset of instrumental variables for child care supply show that all coefficients take on the theoretically correct sign, and all but one (the proportion working from home) are statistically significant. Looking first at the coefficients on the Head Start dummy and the interaction between it and family income, we must proceed with interpretations very carefully. To estimate the partial effect of Head Start providers on for-profit supply, we cannot simply look at the coefficient on Head Start, as this is interpreted as the Head Start-effect when median family income is zero. We must, therefore, plug in interesting values of family income to arrive at the partial effect. Since Head Start providers are located predominately in lower-income neighborhoods, it would be useful to look at neighborhoods at the lower end of the income distribution. For example, in a neighborhood at the first percentile of family income ($14,286), there are 32 fewer child care slots when a Head Start provider is located in the neighborhood [−46.232 + 0.001(14,286)]. Similarly, neighborhoods at the 10th percentile of family income ($28,333) have, on average, 18 fewer child care slots when a Head Start provider is located there [−46.232 + 0.001(28,333)].[18] It therefore appears that Head Start providers crowd-out for-profit services only in low-income neighborhoods.

We also find evidence that communities with many older adults (retirees) may crowd-out formal child care providers. This occurs when households choose to utilize retirees as resources in caring for their children, thus lowering aggregate demand for formal child care services. Finally, increased numbers of individuals who work from home appears to be negatively correlated with neighborhood child care supply; however it just misses statistical significance at the 10% level. But a joint F-test on the null hypothesis that all four variables are simultaneously equal to zero is rejected ($F = 6.14$; $p < 0.01$). We conclude that these variables form an important constellation of *barriers to entry* into the formal child care market, and, therefore, are robust sources of identifying variation in child care supply.

Recall that the theoretical and empirical models described earlier assume that both dependent variables are endogenous, with the endogeneity originating from the simultaneous causation between them. Table 3 presents results from an explicit test of the endogeneity of female labor supply and child care supply, as described in Wooldridge (2001).[19] We conduct the empirical test using various combinations of the instrumental

[18] However, neighborhoods at the mean of family income ($54,726) actually have a greater supply of child care when a Head Start provider is present.

[19] The procedure for conducting the endogeneity test is as follows: First, we estimate a separate reduced form equation for child care supply and female labor force participation, controlling for all exogenous variables and the additional instrumental variables. Second, we calculate the reduced form residuals for each equation, or τ. The final step is to include τ as a regressor in the structural participation and child care supply equations, which include the endogenous independent variables. Evidence of endogeneity is found when τ is statistically significant, as shown in Table 3.

Table 3 Endogeneity tests of child care supply and female labor supply

	Structural female labor supply model		Structural child care supply model	
	τ (*t*-Statistic)	τ (*t*-Statistic)	τ (*t*-Statistic)	τ (*t*-Statistic)
Instruments for child care supply				
A and B	−0.013 (2.43)**	−0.002 (0.17)	–	–
A, B, and C	−0.011 (2.11)**	0.006 (0.39)	–	–
A, B, C, and D	−0.014 (2.63)**	−0.011 (0.70)	–	–
*Instruments for female labor force participatio*n				
E	–	–	0.288 (0.14)	2.134 (0.51)
E and F	–	–	−2.076 (1.11)	−6.014 (1.86)*
E, F, and G	–	–	−2.724 (2.25)**	−3.702 (2.69)***
County fixed effects (No/Yes)	No	Yes	No	Yes

Notes: The child care supply instruments are as follows: A is the Head Start dummy; B is the Head Start/median family income interaction; C is the ratio of children ages 0–5 to adults ages 55–74; and D is the proportion of employed individuals working from home. The female labor supply instruments are as follows: E is the disability rate; F is the proportion of individuals ages 18–64 who speak English *not well* or *not at all*; and G is the proportion of households without a vehicle. All reduced form equations are estimated with county fixed effects

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

variables for both dependent variables and also report an estimate of τ (the reduced form residual) in structural models with and without county fixed-effects. Beginning with the test for child care supply (upper left-hand corner), we find moderate evidence of a negative correlation between the error terms (ε and τ), that is, the nature of the endogeneity is to place a downward bias on the effect of child care supply. The results do not appear to be sensitive to the number of instrumental variables included in the reduced form equation, but they are quite sensitive to whether the fixed-effects appear in the structural labor supply equation. Of interest, τ becomes insignificant when fixed-effects are included. This is most likely due to the effectiveness of county fixed-effects in accounting for unobserved heterogeneity correlated with child care supply and labor supply. Looking at the test for female labor supply (lower right-hand corner), we also find moderate evidence of a negative correlation between the error terms. However, in this case, the results appear to be more sensitive to the inclusion of additional instrumental variables. From this we conclude that county fixed-effects will not completely mitigate the endogeneity of female labor supply, and we must rely on the instrumental variables to do so. Nevertheless, given the evidence of endogeneity in both dependent variables, it is prudent to re-estimate the equations using 3-SLS.

Table 4 presents the 3-SLS estimates of the joint labor supply and child care supply equations. Models 1 and 2 estimate the equation using a linear functional form, and Models 3 and 4 provide results from a log-linear specification. Both models are estimated with and without county fixed effects, but the discussion focuses on models that include the fixed effects (Models 2 and 4).

Looking first at the labor supply equation, we observe few differences between the 3-SLS and OLS coefficients. The effect of child care supply increases fourfold, with a coefficient of 0.013 compared to 0.003 when estimated using OLS. This result accords with the endogeneity tests, which found that the OLS estimate of child care supply is biased

Table 4 3-SLS estimates of the female labor force participation and child care supply equations

Equation	Model 1: Linear	Model 2: Log-linear	Model 3: Linear	Model 4: Log-linear
Female labor force participation				
Total child care supply	0.014 (5.07)***	0.013 (4.15)***	0.035 (4.99)***	0.036 (4.53)***
Child care cost	−0.017 (1.54)	−0.042 (3.37)***	−0.117 (5.72)***	−0.149 (5.24)***
Providers accepting subsidies	−0.001 (0.12)	0.009 (1.08)	0.016 (3.50)***	0.017 (3.64)***
Unemployment rate	−0.260 (4.49)***	−0.215 (3.80)***	−0.033 (5.83)***	−0.024 (4.39)***
African American	0.105 (13.16)***	0.108 (10.06)***	0.021 (9.70)***	0.022 (8.04)***
Ratio of women ages 40–64 to women ages 16–39	−9.375 (15.29)***	−8.472 (12.85)***	−0.147 (16.09)***	−0.127 (13.23)***
Women ages 25+ with at least a BA degree	0.189 (5.19)***	0.207 (5.66)***	0.031 (3.02)***	0.038 (3.70)***
Males ages 25+ with at least a BA degree	−0.125 (3.90)***	−0.118 (3.77)***	−0.012 (1.25)	−0.010 (1.06)
Median family income	0.0003 (6.39)***	0.0002 (3.27)***	0.189 (10.47)***	0.150 (6.95)***
(Median family income)^2	−1.090e-09 (4.91)***	−6.7500e-10 (2.80)***	–	–
Disability rate	−0.075 (1.66)*	−0.075 (1.69)*	0.036 (2.88)***	0.032 (2.68)***
Individuals ages 18–64 who speak English *not well* or *not at all*	−0.080 (1.96)**	−0.135 (2.97)***	−0.002 (0.70)	−0.003 (1.49)
Households without a vehicle	−0.238 (8.66)***	−0.256 (8.18)***	−0.019 (4.44)***	−0.016 (3.86)***
Constant	62.760 (30.52)***	60.687 (24.47)***	2.243 (11.49)***	2.682 (10.79)***
Child care supply				
Female labor force participation	1.682 (1.68)*	2.702 (2.70)***	2.375 (3.43)***	2.964 (4.32)***
Child care cost	0.195 (1.14)	0.820 (3.69)***	0.008 (0.05)	0.509 (2.27)**
Number of children ages 0–5	0.263 (18.12)***	0.231 (15.41)***	0.775 (12.91)***	0.693 (11.22)***
Children ages 0–5 living with an employed single mother	0.862 (2.27)***	1.023 (2.71)***	0.032 (1.00)	0.029 (0.93)
Children ages 0–5 living with two employed parents	0.499 (1.71)*	0.377 (1.33)	0.053 (1.03)	0.004 (0.08)
Individuals ages 5+ living in the same house for at least 5 years	0.739 (1.95)*	0.883 (2.36)**	0.751 (6.75)***	0.805 (7.08)***
Median family income	0.001 (1.49)	0.001 (1.40)	-0.327 (1.88)*	−0.340 (2.00)**
(Median family income)^2	−1.030e-08 (2.17)**	−8.270e-09 (1.74)*	–	–

Table 4 continued

Equation	Model 1: Linear	Model 2: Log-linear	Model 3: Linear	Model 4: Log-linear
Head Start provider	−39.801 (1.96)**	−48.475 (2.43)**	−0.329 (2.25)**	−0.332 (2.27)**
(Head Start * Median family income)	0.001 (2.32)**	0.001 (2.94)***	7.260e-06 (2.80)***	7.840e-06 (2.98)***
Ratio of children ages 0–5 to adults ages 55–74	−1.964 (2.98)***	−1.487 (2.33)**	−0.015 (0.17)	−0.016 (0.18)
Employed individuals working from home	−1.081 (0.58)	−5.819 (2.96)***	−0.008 (0.28)	−0.040 (1.22)
Constant	−190.924 (3.20)***	−279.441 (4.51)***	−9.457 (4.79)***	−13.247 (5.89)
County fixed effects (No/Yes)	No	Yes	No	Yes

Notes: Absolute value of *t*-statistics is in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

downward. With the exception of the proportion of providers accepting subsidies, all other covariates assume the theoretically correct sign and are statistically significant. In fact, the subset of instruments for female labor supply—the three *barriers to employment*—are all negatively related to labor supply, as predicted by the theoretical model.

Minor differences are also found in the child care supply equation. The one noteworthy difference is the coefficient on female labor supply, which becomes substantially larger in the 3-SLS equation: 2.702 compared to 0.851 when estimated using OLS. The coefficient on female labor supply implies that a 1% point increase in a neighborhood's female labor force participation rate is associated with approximately three additional child care slots, all else equal. Once again, the increased sensitivity of child care supply to changes in female labor supply is not surprising, given that the nature of the OLS bias is to place a downward bias on the effect of female labor supply. Most of the other variables in the model are statistically significant. Indeed, the subset of instruments for child care supply—the *barriers to entry* into the formal child care market—are all negatively related to child care supply. Thus, it appears that Head Start providers, informal caregivers, and parents working from home all crowd out for-profit family- and center-based child care services.

Conclusion and Policy Implications

This study represents the first to explore the simultaneous relationship between the geographic supply of child care and female labor supply. It adds to our knowledge about the geography of child care supply in several ways. First, we find Maryland to have a fairly high supply of child care services, with an average neighborhood stock of 135 family- and center-based slots. Furthermore, the average neighborhood has approximately five children ages 0 to 5 for every child care slot. And although high-income neighborhoods have, on average, a greater child care supply than low-income neighborhoods, 155 slots compared to 84 slots, there are very few differences when children are accounted for: The typical low-income neighborhood in Maryland has nearly seven young children per available child care slot, while high-income neighborhoods average around five young children per slot.[20]

The estimation results suggest that female labor supply is sensitive to the geographic supply of child care. Specifically, we find that increases in a neighborhood's stock of family- and center-based child care slots are associated with rising rates of female labor supply. Furthermore, these positive results hold for both OLS and 3-SLS specifications. The results, moreover, suggest that a neighborhood's supply of child care services is sensitive to the level of female labor supply. OLS and 3-SLS estimation consistently find that increases in female labor force participation rates leads to greater for-profit child care supply. We also find the supply of formal child care services is sensitive to several neighborhood attributes. Adverse demand conditions such as the presence of low-income households and increased competition from other services are associated with fewer for-profit child care slots. Indeed, we find that additional Head Start slots, particularly in

[20] High-income neighborhoods are define as census tracts with median family incomes of at least 400% the FPL, and low-income neighborhoods are those tracts with median family incomes of less than 200% the FPL.

low-income neighborhoods, and sources of informal child care tend to crowd out for-profit providers by posing *barriers to entry* into the market. On the other hand, neighborhoods with greater numbers of young children and increased residential stability appear to attract family- and center-based providers.

Our research has a number of policy implications, and we use developments in the state of Maryland to highlight these. States are required to spend no less than five percent of their annual Child Care and Development Fund allocation on quality activities. Many of these activities include incentives to increase the local supply of child care through grants that subsidize additional child care slots and initiatives to increase wages. For example, Maryland assists school districts in providing school-age services and contracts with community-based organizations to increase child care supply. The state also established a quality improvement grant program to award child care providers that increase the stock of accredited child care slots. Although many of these initiatives seek to raise average quality in existing child care environments, an ancillary effect of such endeavors is to raise child care supply and lower entry barriers for new providers. These programs are, therefore, likely to have a dramatic effect on the distribution of child care supply throughout Maryland, leading to increases particularly in low-income neighborhoods. Results in this study suggest that as the neighborhood supply of child care grows and becomes increasingly accessible to working families, employment is expected to increase as well. Therefore, one of unintended (positive) consequences of states' supply-side initiatives to increase quality is the creation of demand-effects through employment growth.

Several potential limitations of this research must be mentioned. First, it models the availability of only family- and center-based services while omitting other important sources, notably relative care. According to some estimates, the two modes of child care observed in this study capture approximately 50% of the child care market (Capizzano et al. 2000; Ehrle et al. 2001). To obtain accurate estimates of the parameters of interest, it is important to make corrections for these missing data. Second, the use of aggregate data, even at the census tract-level, is not optimal. Since many neighborhood characteristics are correlated with each other, it is significantly more difficult to find adequate instruments than in the case of individual-level data. Furthermore, the use of a single cross-section of data makes it problematic to locate exogenous sources of variation. The complexity arises from the fact that every neighborhood in the sample is *exposed* to the same child care policy regime and similar economic conditions. Although inclusion of county fixed effects should mitigate the bias from unobserved heterogeneity, we are forced to rely heavily on the economic models in order to justify the plausibility of the instruments. However, given that the OLS and 3-SLS estimates are reasonably close, we are fairly confident in the robustness of the results. Finally, the assumption that families prefer child care services close to home might be too restrictive. It is reasonable, for example, that some families prefer child care closer to the place of employment, but evidence cited in this paper show that this is not typically the case. Furthermore, data restrictions on commuting patterns and work locations preclude a full investigation of its effects, leaving this an important area for future research.

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References

Anderson, K., & Burkhauser, R. (1985). The importance of the measure of health in empirical estimates of the labor supply of older men. *Economics Letters, 16*, 375–380.

Anderson, P., & Levine, P. (2000). *Child care and employment decisions*. Paper presented at the conference on Labor Markets and Less-Skilled Workers (Sponsored by the Joint Center for Poverty Research). Chicago.

Bania, N., Coulton, C., & Leete, L. (2000). *Welfare reform and access to job opportunities in the Cleveland metropolitan area* (Working Paper No. 00-02). Cleveland, OH: Case Western Reserve University, Center on Urban Poverty and Social Change.

Baum, C. (2002). A dynamic analysis of the effect of child care costs on the work decisions of low-income mothers with infants. *Demography, 39*, 139–164.

Becker, G. (1981). *A treatise on the family*. Cambridge: Harvard University Press.

Berger, M., & Black, D. (1992). Child care subsidies, quality of care, and the labor supply of low-income, single mothers. *Review of Economics and Statistics, 74*, 635–642.

Blau, D. (2001). *The child care problem: An economic analysis*. New York: Russell Sage Foundation.

Blau, D., & Hagy, A. (1998). The demand for quality in child care. *Journal of Political Economy, 106*, 104–146.

Blau, D., & Robins, P. (1988). Child care costs and family labor supply. *Review of Economics and Statistics, 70*, 374–381.

Blau, D., & Robins, P. (1989). Fertility, employment, and child care costs. *Demography, 26*, 287–299.

Blau, D., & Robins, P. (1991). Child care and the labor supply of young mothers over time. *Demography, 28*, 333–351.

Blau, D., & Tekin, E. (2007). The determinants and consequences of child care subsidies for single mothers in the USA. *Journal of Population Economics, 20*, 719–741.

Blumen, O. (1994). Gender differences in the journey to work. *Urban Geography, 15*, 223–245.

Blumenberg, E. (2002). On the way to work: Welfare participation and barriers to employment. *Economic Development Quarterly, 16*, 314–323.

Brandon, P. (2000). Child care utilization among working mothers raising children with disabilities. *Journal of Family and Economic Issues, 21*, 343–363.

Capizzano, J., Tout, K., & Adams, G. (2000). *Child care patterns of school- age children with employed mothers* (Occasional Paper No. 41). Washington: The Urban Institute.

Carliner, G. (1996). *The wages and language skills of U.S. immigrants* (Working Paper No. 5763). Cambridge: National Bureau of Economic Research.

Cohen, S. (2001). *Championing child care*. New York: Columbia University Press.

Connelly, R. (1992). The effect of child care costs on married women's labor force participation. *Review of Economics and Statistics, 74*, 83–90.

Connelly, R., & Kimmel, J. (2001). *The effect of child care costs on the labor force participation and welfare recipiency of single mothers: Implications for welfare reform* (Working Paper 01–69). Kalamazoo: W.E. Upjohn Institute for Employment Research.

Danziger, S., Corcoran, M., Danziger, S., Heflin, C., Kalil, A., Levine, J., et al. (2000). *Barriers to the employment of welfare recipients* (Poverty Research and Training Center Working Paper) Ann Arbor: University of Michigan.

Edwards, J., Fuller, B., & Liang, X. (1996). The mixed preschool market: Explaining local variation in family demand and organized supply. *Economics of Education Review, 15*, 149–161.

Ehrle, J., Adams, G., & Tout, K. (2001). *Who's caring for our youngest children? Child care patterns of infants and toddlers* (Occasional Paper No. 42). Washington: The Urban Institute.

Gordon, R., & Blinder, A. (1980). Market wages, reservations wages, and retirement decisions. *Journal of Public Economics, 14*, 277–308.

Gordon, R., & Chase-Lansdale, L. (2001). Availability of child care in the United States: A description and analysis of data sources. *Demography, 38*, 299–316.

Greene, W. (2003). *Econometric analysis* (5th ed.). Upper Saddle River: Prentice Hall.

Grenier, G. (1984). The effects of language characteristics on the wages of Hispanic-American males. *Journal of Human Resources, 19*, 35–52.

Han, W., & Waldfogel, J. (1998, April). Child care and women's employment. Paper presented at the Annual Meeting of the Population Association of America, Chicago.

Hanoch, G., & Honig, M. (1983). Retirement, wages, and labor supply of the elderly. *Journal of Labor Economics, 1*, 131–151.

Henly, J., & Lyons, S. (2000). The negotiation of child care and employment demands among low-income parents. *Journal of Social Issues, 56*, 683–706.

Jenkins, S. (1992). Lone mothers' employment and full-time work probabilities. *The Economic Journal, 102*, 310–320.

Kahn, A., & Kamerman, S. (1987). *Child care: Facing the hard choices*. Dover: Auburn House Publishing.

Kimmel, J. (1998). Child care costs as a barrier to employment for single and married mothers. *Review of Economics and Statistics, 80*, 287–299.

Kreader, L., Brickman-Piecyk, J., & Collins, A. (2000). *Scant increases after welfare reform: Regulated child care supply in Illinois and Maryland, 1996–1998*. New York: Columbia University, National Center for Children in Poverty, Joseph L. Mailman School of Public Health.

Liebowitz, A., Klerman, J., & Waite, L. (1992). Employment of new mothers and child care choice: Differences by children's age. *Journal of Human Resources, 27*, 112–133.

Liebowitz, A., Waite, L., & Witsberger, C. (1988). Child care for preschoolers: Differences by child's age. *Demography, 25*, 205–221.

Loprest, P. (2001). *How are families that left welfare doing? A comparison of early and recent welfare leavers* (Policy Brief Series B, No. B-36). Washington: The Urban Institute.

McManus, W., Gould, W., & Welch, F. (1983). Earnings of Hispanic men: The role of English language proficiency. *Journal of Labor Economics, 1*, 101–130.

Meyers, M., Heintz, T., & Wolf, D. (2002). Child care subsidies and the employment of welfare recipients. *Demography, 39*, 165–179.

Meyers-Jones, H., & Brooker-Gross, S. (1996). The journey to child care in a rural American setting. In K. England (Ed.), *Who will mind the baby? Geographies of child care and working mothers*. New York: Routledge.

Michelson, W. (1985). *From sun to sun*. Toronto: Rowman & Allanheld.

Mora, M. (1998). Did the English deficiency earnings penalty change for Hispanic men between 1979 and 1989? *Social Science Quarterly, 79*, 581–594.

Ong, P. (1996). Work and car ownership among welfare recipients. *Social Work Research, 2*, 255–262.

Ong, P. (2002). Car access and welfare-to-work. *Journal of Policy Analysis and Management, 21*, 239–252.

Parsons, D. (1980). The decline in male labor force participation. *Journal of Political* Economy, 88, 117–134.

Presser, H., & Baldwin, W. (1980). Child care as a constraint on employment: Prevalence, correlates, and bearing on the work and fertility nexus. *American Journal of Sociology, 85*, 1202–1213.

Queralt, M., & Witte, A. (1998). Influences on neighborhood supply of child care in Massachusetts. *Social Service Review, 72*, 17–46.

Ribar, D. (1992). Child care and the labor supply of married women: Reduced form evidence. *Journal of Human Resources, 27*, 134–165.

Ribar, D. (1995). A structural model of child care and the labor supply of married women. *Journal of Labor Economics, 13*, 558–597.

Rosenbloom, S. (1988). Is there a women's transportation problem? *Women and Environments, 10*, 16–17.

Rosenboom, S. (1993). Women's travel patterns at various stages of their lives. In C. Katz & J. Monk (Eds.), *Full circles: Geographies of women over the life course*. NY: Routledge.

Sickles, R., & Taubman, P. (1986). An analysis of the health and retirement status of the elderly. *Econometrica, 54*, 1339–1356.

Smith, K. (2000). *Who's minding the kids? Child care arrangements: 1995* (Current Population Reports, No. P70–70). Washington: U.S. Census Bureau.

Stern, S. (1989). Measuring the effect of disability on labor force participation. *Journal of Human Resources, 24*, 361–395.

Stoll, M., & Raphael, S. (2000). Racial differences in spatial job search patterns: Exploring the causes and consequences. *Economic Geography, 76*, 201–223.

Stolzenberg, R., & Waite, L. (1984). Local labor markets, children and the labor force participation of wives. *Demography, 21*, 157–170.

Trejo, S. (1997). Why do Mexican Americans earn low wages? *Journal of Political Economy, 105*, 1235–1267.

Truelove, M. (1996). The locational context of child care center in metropolitan Toronto. In K. England (Ed.), *Who will mind the baby? Geographies of child care and working mothers*. New York: Routledge.

Urban, J., & Olson, P. (2005). A comprehensive employment model for low-income mothers. *Journal of Family and Economic Issues, 26*, 101–122.

U.S. Bureau of the Census (2000). Appendix 1: Census 2000 geographic terms and *concepts*. Retrieved September 2, 2002, from http://www.census.gov/geo/tiger/glossry2.pdf.

U.S. Bureau of Labor Statistics. (2004). Employment characteristics of families in 2003. Retrieved August 1, 2004, from ftp://ftp.bls.gov/pub/news.release/famee.txt.

U.S. Department of Education. (1995). Characteristics of children's early care and education programs: Data from the 1995 National Household Education Survey *(NHES)*. Washington: National Center for Education Statistics Department of Education.

U.S. General Accounting Office. (1994). *Child care: Child care subsidies increase likelihood that low-income mothers will work* (Report No. HEHS-95–20). Washington: U.S. General Accounting Office.

Van Dijk, L., & Siegers, J. (1996). The effect of the supply of subsidized day-care facilities on female labor supply. *Labour, 10*, 559–582.

Webster, C., & White, S. (1997a). Child-care services and the urban labour market. Part 1: Theurban child-care market. *Environment and Planning A, 29*, 1419–1431.

Webster, C., & White, S. (1997b). Child-care services and the urban labour market. Part 2: Modeling the relationships between child-care service accessibility and labour market participation. *Environment and Planning A, 29*, 1675–1695.

Wolfe, B., & Hill, S. (1995). The effect of health on the work effort of single mothers. *Journal of Human Resources, 30*, 42–62.

Wooldridge, J. (2001). *Introductory econometrics: A modern approach*. South-Western.