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Early Child-Care Selection: Variation by Geographic Location, Maternal Characteristics, and Family Structure

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More than half of all U.S. infants and toddlers spend at least 20 hr per week in the care of a nonparent adult. This article uses survival analysis to identify which families are most likely to place their child in care and the ages when these choices are made, using data from a national probability sample of 2,614 households. Median age at first placement is 33 months, but age varies by geographic region, mother's employment status during pregnancy, mother's education level, and family structure (1 vs. 2 parents, mother's age at 1st birth, and number of siblings). Controlling for these effects, differences by race and ethnicity are small. Implications for studies of child-care selection and evaluations of early childhood programs are discussed.

The U.S. family's burgeoning demand for child care and preschooling has gained broader public attention in recent years. Between 1950 and 1990, the percentage of mothers with preschool-age children who were employed quadrupled, from 14% to 58% (Hofferth, 1989). More than two thirds of all children, ages 3–5, now spend an average of 19 hr per week in the care of a nonparent adult. More than 60% of all young children attend a formal center or preschool prior to enrolling in kindergarten (West, Hausken, & Collins, 1993).

Less well-known is the fact that more than one half of all infants and toddlers spend some time in the care of a nonparent provider prior to reaching age 12 months (Caspar, 1996). Surprisingly little knowledge is available on which families are most likely to select nonparental care and at what age in the child's life. The present article details how most families make initial nonparental care choices during the child's infancy or toddler years, despite all of the attention that center-based programs, primarily serving 3- and 4-year-olds, have received in family and early childhood policy circles.

This paucity of evidence makes it difficult to determine (a) when young children from different types of families enter nonparental care, (b) whether access to early child care is equally distributed across diverse U.S. families, (c) whether early child-care selection puts only certain youngsters on a developmental trajectory that eventually includes exposure to center-based preschool, and (d) whether the apparent effects of formal programs, such as Head Start, stem from the preschool "treatment" per se or from prior selection effects rooted in the family's attributes and practices. The length of time that children spend in nonparental care will likely affect the magnitude of such treatment effects, but we know very little about when—at what age—young children enter nonparental settings (Fuller, Holloway, & Liang, 1996).

One recent analysis of data collected as part of the Children of the National Longitudinal Survey of Youth (NLSY), found that the number of years a child spent in nonparental care and how early it began was associated with his or her reading-related skills at ages 5 and 6 (Caughy, DiPietro, & Strobino, 1994). To control for potential confounding factors, these researchers used global socioeconomic indicators and home environment scores as statistical controls before assessing the effects of child care of variable length. But a wider set of parent attributes and practices is likely associated with the selection of child care outside the home, indirectly shaping early cognitive development. This selection process may vary over time, from ages 0–5, for different types of families. Designs to date have assumed that home effects are direct, rather than having recognized that parents' management of their child's time outside the home may yield indirect effects. Also the child-care selection process over time, commencing with the child's birth, has not been adequately modeled.

In this article we provide initial evidence as to which families are more likely to use nonparental forms of child care (babysitters, family day-care homes, kin members, and centers) and at what age initial entry into these settings occurs. We begin by reviewing what is known about the child-care selection process. We show that this young body of evidence continues to focus on later selection of center-based programs for children ages 3–4 years. Researchers also continue to rely on traditional statistical methods—linear and logistic regression analysis—ill-suited for analyzing the unfolding pattern of child-care selection over

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time. We then turn to the present study, which address both of these limitations by using discrete-time survival analysis to model the event histories of a national probability sample of 2,614 children under age 6 to determine whether, and if so when, they first entered nonparental care.

Family Selection of Nonparental Child Care

Three theoretical explanations have been advanced for why families differ in their propensity to select nonparental or center-based child care. The first two explanatory processes operate at the family level: the household's economic situation and its demographic structure. The third relates to the organized supply of nonparental child-care options. Operating above the family, recent research has begun to document how supply conditions of, for example, family day care or center-based programs, vary sharply across geographic regions and local communities. These organization-level dynamics likely condition family-level selection effects.

Household Economy and Social Class

Maternal employment is predictably related to the propensity to use nonparental care, with nearly three fourths of families with working mothers using child care for youngsters under 5 (West et al., 1993). Less well-known is the fact that about one third of children with nonemployed mothers are also placed in nonparental care. Family income appears unrelated to this decision. Low-income families (with incomes under $15,000 annually) are only somewhat more likely to use kin members and less likely to select centers, compared with affluent households earning more than $50,000 (Hofferth, Brayfield, Deich, & Holcomb, 1991). The relative prices of different forms of child care appear associated with selection decisions, but the effects of this factor are relatively small because many low-income and working-poor families receive subsidies that largely offset private costs (Hofferth & Wissoker, 1992). The scarcity and high cost of infant and toddler care does point to family income as a possible determinant of selection decisions (Culkin, Morris, & Helburn, 1991). Yet, one initial article on selection of infant care shows that high-income and impoverished families display the greatest (and similar) probabilities of selecting nonparental providers (National Institute on Child Health and Human Development [NICHD], Early Child Care Research Network, 1997).

Research in a limited number of geographic areas suggests that social class may be related to the selection of nonparental care. Phillips, Voran, Kisker, Howes, and Whitebook (1994) argued that working-class households are least able to use centers, inasmuch as they are ineligible for subsidies and cannot afford private fees. But this dip in center selection may also be linked to the higher rate of part-time employment among these mothers, many of whom work night or swing shifts, when fathers or kin members can provide child care (Presser, 1986).[1] Moreover, social class is a complex construct. Operationalizing it using variables such as household income raises questions about which of the two changes occurs first. Income may rise after selecting nonparental care because a mother can return to the workforce (the problem of endogeneity). Research is needed that identifies elements of the family's economic and class position that are closer to the proximate causes and processes driving early child-care selection.

Family Demographics and Social Structure

Better educated parents are consistently more likely to place their children in nonparental care. Among families nationwide with children ages 3–5, 48% were in nonparental care when one parent's highest schooling level was less than a secondary school diploma, versus 72% when one parent had attended some college (West et al., 1993). We know very little, however, about whether parents' schooling levels also influence the selection of early nonparental care and whether these effects persist after controlling for other factors (e.g., maternal employment or single-parent family status) known to be correlated with education. And from a psychological viewpoint, researchers are just beginning to learn how maternal education levels may shape proximal determinants of selection.

Many studies have found racial and ethnic differences in the use of child care, although the reasons for these differentials remain unclear. Nationally, only 46% of Latino families use nonparental care for children under age 5, in comparison to 64% for White families and 75% for Black families (Hofferth, West, Henke, & Kaufman, 1994). Liang (1996) found that the differential between Latino and non-Latino families is most pronounced among those families who speak Spanish at home. One explanation for this differential is that Latino women are less likely to be employed, either part time or full time (Folk & Beller, 1993). But even in households where the mother is working, the percentage of Latino families who enroll children ages 3–5 in center-based programs is 23 points below the rate of Black households (Fuller et al., 1996). Variability across ethnic communities in the organized supply of centers or family day-care homes may further contribute to differences in the rate of selecting nonparental care.[2]

The balance between the family's level of need for child care (e.g., represented by the number of young children in the household) and the availability of supportive adults in the household appears associated with the selection of nonparental care. Parents, as they have more children, are less likely to select nonparental providers (Leibowitz, Waite, & Witsberger, 1988; NICHD, 1995) or they shift to less expensive forms of care (e.g., shifting from centers to family day care homes, Lehrer, 1983). The rate at which infants enter nonparental care also appears to be higher for smaller families (NICHD, 1997). But the rate at which this family-size effect occurs (is it linear with each additional child or is there a ceiling effect?) is not well understood. Cost considerations become more potent as family

[1] Almost 38% of all employed women with at least one child under age 5 work part time. These families tend to rely just on the parents, or proximate kin, to care for their young children. Families with mothers working full time, who tend to be lower income or affluent and well educated, tend to enter the child-care market and select family day care or a center-based program. But again, we don't know when these effects occur in the child's life (Folk & Beller, 1993; Michalopoulos, Robins, & Garfinkel, 1992).

[2] A recent study of availability of center-based programs in Massachusetts found that supply was significantly lower in local communities with high concentrations of Spanish-speaking parents (Fuller & Liang, 1996).

size grows, causing more and more mothers, not eligible for subsidies, to stay at home. At the same time, the presence of a kin member or other nonparent adult in the household, or living close by, makes selection of nonparental care much easier. Heckman (1974), for example, found that families with proximate kin members relied on them more and on center-based programs less (see also, Hofferth & Wissoker, 1992).[3]

Most of this research has focused on entry into center-based programs. Focusing on infants and toddlers less than 15 months of age, the NICHD Early Child Care Research Network (1995) found that children were more likely to be placed into nonparental care earlier when the mother was employed, single, had less formal education, and was of African American ethnicity. Firstborns were placed at a younger age than were children further down the birth order. The children of mothers who reported being worried about possible negative effects from working outside the home on their infant or toddler were placed less frequently or at an older age. This moves theory closer to the mother's own reasoning about child-care settings. But these conclusions are based on multiple regression analyses using age of placement as the outcome, excluding parents who did not use nonparent infant or toddler care prior to 15 months. This, of course, sets aside families who select nonparental care at a later age. These researchers also were faced with difficult model specification issues surrounding the endogeneity of their predictors. For instance, the use of current maternal employment and mothers' beliefs about the risks or benefits of employment as predictors implies that these factors must be causes, not consequences, of child-care selection.

Organization-Level Factors

Variation in the supply of nonparental care across states and localities may also influence the individual family's propensity to use care. Early studies used family-level data to draw inferences about supply, access, and the distribution of quality of nonparental care providers across local communities. Drawing from data collected in five cities, for example, Whitebook, Howes, and Phillips (1989) argued that parents of children from working-class and middle-income families selected centers at a lower rate than parents of youngsters from either low- or high-income households. More recent work has directly observed the per capita availability of center-based programs, attempting to understand how supply may constrain the family-level selection process. For example, one national survey of nonparental care providers found that the supply of family day care was highest in the West, relative to these states' share of preschool-age children (Kisker, Hofferth, Phillips, & Farquhar, 1991). In contrast, the supply of center-based programs was highest in the South, relative to this region's child population.[4] The greater availability of centers in the South could be due to the higher proportion of Black mothers, a group that works full time at a rate higher than other ethnic groups, or could be due to early supply gains, such as through Head Start, since the 1960s (Folk & Beller, 1993).

Aims of the Present Study

Research on the family's child-care selection process has been hampered by four limitations. First, it has emphasized the selection of center-based programs for 3- and 4-year-olds. Although most parents initially place their child into nonparental care much earlier, we know little about who does this and how young the children are when they are first placed. Second, researchers are just beginning to disentangle how organizational supply, varying across states and locales, constrains the process of finding and selecting nonparental care. Few studies look across levels at both family and community factors. Third, endogeneity problems limit the kinds of data that can be analyzed and the inferences that can be drawn. Current maternal employment and current family income are commonly used as predictors of child-care selection, for example, without asking whether such variables should really be treated as the results of these choices, not potential causes. Fourth, most studies have used either linear regression to predict age at entry (forcing researchers to discard data for children who have not yet entered care or to impute incorrect placement times) or logistic regression (forcing researchers to discard potentially meaningful information about when entry occurs). These techniques are inadequate for analyzing age-heterogeneous data sets in which some children have not yet entered care.

Our project was designed to address each of these limitations. First, rather than focus on a single age period, such as infancy or the preschool years, we examined the entire child-care history of the youngest child (age 6 or under) in the nationally representative sample of 2,614 families participating in the 1990 National Child Care Survey (NCCS; detailed in Hofferth et al., 1991). This allowed us to determine whether these children had ever been placed in nonparental care and if so, when—at what age—they were first placed. Second, we specify more precisely the theoretical processes that may drive early selection of nonparental child care and examine not only family factors, but also two community-level factors—urbanicity and region of residence. We find, in fact, that the entire profile of probability of entry into care varies substantially across the country. Third, rather than use current data on maternal employment to predict the earlier event of whether the child had been placed in care, we reconstructed the mother's employment history to determine whether she was known to be working during pregnancy. As we show, this measure of maternal employment is a powerful predictor of child-care decision making, especially during the child's first few years of life. Fourth, rather than traditional regression methods, we use survival analysis to simultaneously analyze the timing of placement among the full sample of chil-

[3] Resident kin members may allow mothers with young children to enter the labor force by assisting with child care. For example, Figueroa and Melendez (1993) found that Puerto Rican mothers in New York, with a child under 5, are more likely to be employed when a nonparent kin member was resident in the home.

[4] Analyzing data on all preschools and centers operating in 100 counties nationwide, Fuller and Liang (1996) found inequalities in availability per capita across regions of the country and between affluent versus low-wealth counties. Among the 25 most affluent counties, one center-based classroom was available for every 45 children, ages 3–5. In contrast, one classroom was available for every 77 children in the lowest wealth counties. These county-level variations in supply were most strongly related to mean family income observed in the local areas, presence of single-parent households, and population growth, after accounting for the suppressing effect of prices or fee levels.

dren, those who have already entered care and those who have yet to do so.

Method

Sample

In early 1990, 4,392 parents with at least one child under age 13 participated in the NCCS, a telephone survey focusing on child-care issues. Like most national studies, the NCCS used a multistage cluster sample design. In the first stage, a stratified random sample of 100 counties was selected from the 2,683 counties in the United States with 5,000 or more residents. The largest 20 counties (e.g., Los Angeles and Cook counties) were selected with certainty. That is, a design decision was made to include all 20 in the sample. A total of 80 additional counties were then chosen by stratifying the remaining counties by region, urbanicity, and poverty level, and selecting pairs of counties with probability proportional to the estimated number of children in the county who were under age 5. In the second stage, banks of contiguous telephone numbers were selected from the three-digit telephone exchanges used in the 100 selected counties. In the third stage, telephone numbers were drawn from the selected banks, and after a screening interview, parent interviews were successfully completed in 69% of the eligible households.

The analyses reported in this article focus on the youngest child in the subsample of 2,614 households with a child age 6 or under who had not yet entered kindergarten, in which a mother was present and for whom the parent provided information on whether and, if so, when the child first entered nonparental care. To arrive at this subsample, we set aside the 1,558 households in which the youngest child had already entered school, because the interview was designed to have these families skip the child-care questions. We also set aside the 63 households with no mother present, because the factors driving their selection process are likely to differ from those in households with a mother present. Finally, we set aside two small groups of households that did not provide the information necessary for constructing outcome measures, 77 in which the parent was inadvertently not asked the appropriate questions about child care and 80 in which the parent did not indicate whether or when the target child was first placed in care.

Determining Whether and When Children Enter Regular Nonparental Care

A total of 1,881 parents reported that they had placed their youngest child in some type of care prior to the interview; 733 reported that they had not.[5] Because of our interest in identifying the child's first regular arrangement—not one that was simply for a few hours a week—we further examined the intensity (number of hours per week) of care. Among those children who had ever been in care, more than one half (1,045) were placed in their initial arrangement for at least 20 hr per week. Among the remaining 836, most (n = 700, 84%) were placed for 10 hr or less per week. Adopting a cutoff of 20 hr a week, we divided the sample into three groups: 1,045 for whom we know when they were first placed in care; 733 who had not yet entered regular care by the date of the interview; and 836 who may have entered more intensive care after the date provided, but for whom the initial arrangement identified is of low intensity. We used this information to create two measures: one that indicated whether the child was known to be placed in regular nonparental care and a second that indicated when (at what age) that had occurred.

Why and How We Used Discrete-Time Survival Analysis

The problem of censoring. Most previous research on whether and when children enter child care has suffered from serious methodological limitations (Singer & Willett, 1991). The core dilemma has been how to include the children who had been censored by the end of data collection (in our sample, the 733 children who had not yet entered care by the time of the interview) and for whom the outcome—age at first entry—is unknown. For these children (28% of the sample), we do not know when they will enter care, or even if they will enter care. All we know is that they had not yet entered care by the time of the interview. Although some children may begin soon thereafter and some will begin before kindergarten, others will never enter care (prior to the start of elementary school). For these children, we are missing the very item of interest, data about whether and when they enter care. Yet they do provide much information, especially about the probability that parents do not place their children in care (or delay placement). Our analyses were further complicated by a second type of censoring created by the 836 parents who named an initial arrangement that was of too low intensity to meet our criterion. For these children (another 32% of the sample), all we know is that they had not yet been placed in regular child care by a certain date. These children also have censored child-care histories, but the censoring occurs at the date of first placement given by the parents.

Taken together, 60% of the children in the NCCS have censored event histories. To analyze the data for these children simultaneously with that of the 40% of children with known event times, we used survival analysis (Singer & Willett, 1993; Willett & Singer, 1993). Commonly used by biostatisticians studying human lifetimes (in which the event of interest is death), survival analysis can be used to study how long it takes for any event to occur, even when the event is within an individual's (or his or her parent's) control. Although it might appear that the level of censoring here is so high as to curtail statistical power, the large sample size, coupled with variable censoring times, ensures adequate statistical power to detect even relatively small effects (Singer & Willett, 1991).

Why the analyses were conducted in discrete time. We initially planned on analyzing age at entry (expressed in months) as a continuous variable. Examination of its distribution, however, revealed that the event times were both highly skewed and highly discretized (concentrated in a small number of specific values). Although many continuous-time survival methods (e.g., proportional hazards models) require no distributional assumptions about event times, all are sensitive to the "ties" that arise with highly discretized event times (Cox & Oakes, 1984). For example, among the 1,045 children whose parents provided a date of entry, 621 were reported to have entered care before age 6 months and 331 were reported to have entered care either during their birth month (resulting in reported placement ages of 12, 24, 36, or 48 months precisely) or in their half birth month (resulting in reported placement ages of 6, 18, 30, 42, or 54 months precisely). In only 86 cases (8%) did parents report that their child entered care in any other month.

Is this pattern a result of overt parental decision making to place children in care when they reach their birthday (or half birthday)? Or, does it stem from the rounding that clouds respondents' memories during retrospection (Bradburn, Rips, & Shevell, 1987)? With the NCCS data, we have no way of knowing. Regardless of its cause, the behavior of this variable suggested the need for discrete-time methods, the major decision then being the selection of appropriate time intervals. After

[5] We used several lines of questions to determine whether and when the youngest child in the household first entered care. Parents of children currently in care or who had been in care during the past year were asked, "Thinking back on all the child care and preschool programs you have ever used for [target child], how old was [s/he] when you first started leaving [him/her] with someone other than you [or child's other parent] on a regular basis?" All other parents were asked, "Have you ever used any form of child care on a regular basis for [target child] other than you [or child's other parent]?" Parents who responded yes to this question were then asked the age when the child was first left in care.

comparing parameter estimates and standard errors from models based on quarters (3-month groups), half years (6-month groups), and years (12-month groups), as well as those from models based on unequal interval lengths (e.g., monthly for the first few months, half years thereafter), we selected a half-year approach. Children's entry times are collapsed into 10 intervals, from 0–5 months, 6–11 months, and 12–17 months up to 54–59 months. Resonant with the work of Efron (1988), who compared levels of bias arising from alternative degrees of discretization, our substantive findings remained virtually identical, regardless of the time intervals chosen.

Building discrete-time hazard models. Survival analysis models focus not on age of entry directly but rather the *hazard rate,* a transformation of age that remains meaningful in the presence of censoring. In this study, hazard measures the probability that a child will be placed in nonparental care in any particular half-year period, given that he or she had not been placed in care in any previous time period. The conditionality inherent in the hazard rate ensures that it measures risk (the probability of event occurrence) in each time period among a key group of children, those who have not yet entered care. Just like any statistical quantity, hazard rates can be estimated from sample data. If many children enter care between birth and 5 months, for example, the first time period's hazard rate is high. If few children who had not yet entered care do so between 36–41 months, the risk in this later time period is low. Plots of hazard versus age describe the period-by-period risks of initially entering care across the childhood years. Because of the negative connotations associated with the technical term *risk* when referring to first entry into nonparental care, we use the more neutral term *probability.*[6]

We explored the relationship between the hazard rate and predictors by fitting a sequence of discrete-time hazard models that linked the probability of entering care, on one hand, to family and community characteristics on the other. We began by exploring the relationship between hazard and where the child lived (the effects of urbanicity and region), because these two variables were features of the original study design. Subsequent models explored the effects of maternal characteristics and family structure. At every stage, we explored the main effects of each predictor and all possible interactions between predictors. This enabled us to discover, for example, that not only is the number of siblings in the home associated with hazard (the main effect), but also that the effect of family size differs depending on whether the mother was working during her pregnancy (the interaction effect). We note, however, that this was the only statistically significant interaction between predictors. Unless otherwise noted, all comparisons cited in the text are significant at the .05 level (two-tailed tests).

We summarize the results of fitting the series of discrete-time hazard models numerically and graphically. Numerical summaries include (a) Table 2, which presents parameter estimates, standard errors, and the goodness-of-fit statistics for five selected discrete-time hazard models; (b) textual references to antilogged parameter estimates, which can be interpreted as odds ratios; and (c) Table 3, which presents the estimated median age at first placement—the number of months by which one half of a given group of children have entered care—for 96 distinct demographic groups. Because these models describe event occurrence over time, we also present four graphical displays (Figures 1–4) that we believe are even more informative—fitted hazard and survivor functions constructed for prototypical children. The top panel of each figure presents the fitted hazard function, which displays the conditional risk of event occurrence in each time period. The bottom panel presents the fitted survivor function, which cumulates these period-by-period risks to display the estimated proportion of the population that survive through each time period, that is, the proportion who do not enter child care. Each graphical summary conveys important and distinct information. The hazard function describes the risk associated with each individual time period, whereas the survivor function aggregates these risks to describe event occurrence over a much broader period of time.

Testing model assumptions. Like all statistical models, discrete-time hazard models have assumptions—most prominently, linearity and proportionality—that may not be met in practice (Singer & Willett, 1993). To ensure the appropriateness of our models, we examined the tenability of these assumptions for every predictor, and we relaxed them when necessary. We examined the linearity assumption by comparing models that used each predictor expressed in linear and quadratic forms (for continuous variables) with models using the predictor expressed as a set of indicators. In the one instance in which the linearity assumption was violated (when examining the effect of number of siblings), we addressed the problem by categorizing the predictor into three levels (no siblings, one sibling, and two or more siblings) and using two dichotomous indicator variables. As no other violations of the linearity assumption were found, all other predictors were analyzed in their continuous form (although we categorize several for descriptive purposes in Table 1).

Violations of the proportionality assumption were far more common. In the three instances when this assumption was violated (when investigating the effects of region, mother's age at first birth, and maternal employment), we addressed the problem by including interactions with time (Willett & Singer, 1993). Although an interaction with time may seem like nothing more than a methodological nuisance, it has two equivalent interpretations, each of substantive interest. One interpretation is that the shape of the hazard profile differs across groups of children. In Figure 2, for example, we show that the shape of hazard function for children in the South differs from that for children from other regions. The other interpretation focuses on the changing effects of the predictor depending upon the child's age. When exploring the effect of maternal employment during pregnancy, for example, we find that it is large when the child is young but that it dissipates as the child gets older.

Accounting for the Stratified Multistage Cluster Sample Design

The households in the NCCS are not a simple random sample of families with children under age 13. Complex multistage samples, whereas economical for data collection, must be analyzed carefully to account for both the stratification (with unequal probabilities of selection) and the clustering (pairs of counties within strata and banks of telephone numbers within exchanges). Parameter estimates that do not account for the unequal probabilities of selection will be biased toward the results for the oversampled groups. Standard errors that do not account for the clustering will generally underestimate the true level of variability in the parameter estimates. The data analysis routines available in most statistical packages (e.g., SAS, BMDP) assume simple random sampling. When this assumption is patently false, as it is here, alternative approaches are necessary (Lee, Forthofer, & Lorimor, 1989). All of the estimates presented in this article were computed using SUDAAN (Shah, Barnwell, & Bieler, 1995), a statistical package specially designed for the analysis of complex cluster sample data. SUDAAN uses a Taylor series linearization routine to adjust parameter estimates, standard errors, and test statistics for the sample design.[7]

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[6] This semantic substitution is possible because event occurrence is measured in discrete periods. Were we modeling continuous-time data, this substitution would be inappropriate (Allison, 1984).

[7] Because of these adjustments, readers are cautioned that seemingly straightforward calculations from our tables (such as computing percentages using sample sizes) may yield incorrect nation-level estimates of population parameters. Table 1, for example, presents the distribution of the sample according to values of the predictors; the percentages in the third column of the table differ from those that a reader might compute using the sample sizes presented in the second column.

Results

Which Children Have Ever Been Placed in Regular Nonparental Care?

Table 1 presents national estimates—pertaining to the youngest child in the 15.6 million American households in which there is at least one child, 6 years old or younger, and a mother present—of the percentage who entered regular nonparental care at some time before age 5. Table 1 shows that the percentage of children who have ever been in care is strongly associated with age. Older children, having lived longer, have had more opportunity to enter care. More than twice as many 2-year-olds have entered care, for example, in comparison to infants under age 1. It is for this reason that survival analysis is necessary. Were we to model placement in this age-heterogeneous sample (using logistic regression), children would not be on a level playing field, in the sense that they would not have been at risk of entry for equal lengths of time. Nevertheless, we review the findings in Table 1 for two reasons. First, in some cases, they presage those of the survival analyses, helping to identify predictors associated with the use of child care. In other cases, they suggest findings that are refuted by the survival analyses, thus demonstrating the superiority of this analytic approach.

Turning to geographic location, which represents the basic contexts that may condition family-level processes, we find few differences with respect to urbanicity but large differences with respect to region. Children in the South are most likely to enter care; those in the Northeast are least likely. As we will soon

Table 1
Demographic Characteristics of the Sample and the Estimated Percentage of Children Ever in Regular Nonparental Care by Demographic Groups

	Total sample		Ever in regular nonparental care?		
Characteristic	*n*	%	%	χ^2 (*df*)	*p*
Child's age at interview (in years)				107.72 (6)	<.0001
<1	596	22.3	22.8		
1–1.99	500	19.8	34.9		
2–2.99	449	16.1	44.5		
3–3.99	362	13.9	47.2		
4–4.99	335	13.2	49.8		
5–5.99	372	10.9	48.2		
6–6.99	87	3.8	39.9		
Urbanicity				0.28 (2)	.87
Urban	1,042	41.3	39.9		
Suburban	890	33.9	38.9		
Rural	682	24.8	38.0		
Region				19.66 (3)	<.0001
South	855	33.3	47.2		
Midwest	729	24.6	39.0		
West	521	22.5	34.3		
Northeast	509	19.5	30.8		
Mother's education				22.37 (4)	<.0001
Nonhigh school graduate	251	10.7	29.0		
High school graduate	1,058	40.6	37.5		
Some college	625	23.9	40.8		
College graduate	461	17.0	41.8		
Postgraduate education	207	7.8	51.9		
Mother's age at first birth (in years)				5.07 (4)	.28
<18	113	4.2	44.3		
18–20	389	16.2	38.9		
21–24	775	29.4	38.1		
25–29	819	30.8	37.2		
30 or older	507	19.4	43.4		
Race–ethnicity				17.61 (3)	<.001
Latino	238	10.6	34.9		
White	2,060	75.8	37.5		
Black	255	12.7	51.3		
All others	39	0.9	54.0		
Single-parent family?				29.31 (1)	<.0001
Two parents	2,201	83.0	35.8		
Single mother	413	17.0	54.9		
Mother working before birth				127.98 (1)	<.0001
Working	920	33.8	56.1		
Not working	1,694	66.2	30.4		
Number of siblings				100.58 (2)	<.0001
None	1,040	39.1	48.1		
One	1,044	40.7	38.6		
Two or more	505	20.2	23.5		

show, these regional differences are complex, varying not only in level but in the entire shape of the hazard profile: there are times in children's lives when those living in the Northeast are most likely to enter care.

Next, consider maternal and family-level factors. Although mother's education is positively associated with child-care use, there appears to be no effect of maternal age at first birth (a conclusion refuted by the survival analyses). As others have found, children who are either White or Latino are less likely to be placed in care than African American children. Mothers who are single, who worked during pregnancy, and who have no other children are far more likely to use nonparental care.

When Do Children Enter Their First Regular Nonparental Care Arrangement?

Nearly one fourth of all children are placed into their first nonparental child-care setting in their first 5 months of life; as children grow older, the probability of initial placement declines (Figure 1). Among children not placed between birth and 5 months, an estimated 10% enter during the next 6 months (between ages 6 and 11 months). Among those not placed in either of these periods, an estimated 7% start between 12 and 17 months and an estimated 6% of those not placed before turning 1½ start between 18 and 23 months. Although the probability of initial placement in all subsequent 6-month periods never exceeds 11%, the zig-zag pattern in the hazard profile reflects the pattern alluded to in the Method section. Parents are either far more likely to initiate child care when the child has a birthday (the time periods beginning at 24, 36, and 48 months) or they are more likely to report initiating placement during these periods. In the off periods after age 2, the risk of placement never exceeds 3%.[8]

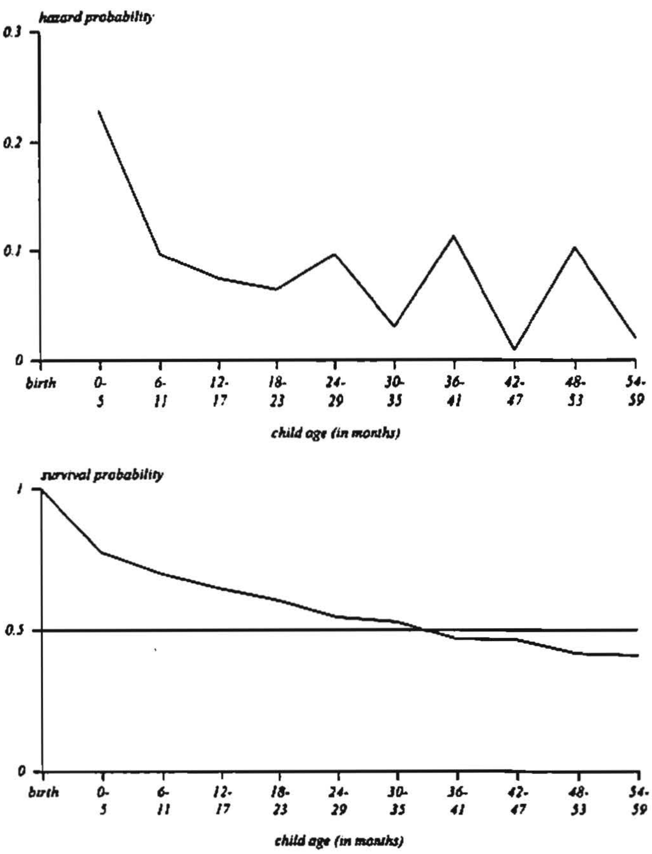

Figure 1. Sample hazard function (top panel) and survivor function (bottom panel) describing the risk of initial entry into regular nonparental care by child age (in months) for the youngest child in a probability sample of U.S. households.

The cumulative impact of these period-by-period probabilities is displayed in the sample survivor function in the bottom panel of Figure 1. Recall that the survivor function represents the percentage of children not yet placed in regular care. Thus, an estimated 60% of the children had not yet entered care before 24 months, and an estimated 41% had not yet entered care before 60 months. To facilitate interpretation, subtract each of these percentages from 100 to find that by age 2 an estimated 40% of all children have begun their first regular child-care arrangement and that by age 5, 59% have done so. Using the horizontal line drawn at a survival probability of .50 to estimate the median lifetime, we estimate that one half of the nation's children begin their first regular care arrangement before they turn 3 (by age 33 months). Before age 5 (by 59 months), approximately 60% have done so.

Differences in Timing by Geographic Region and Urbanicity

The descriptive results presented in Table 1 suggest that although children are equally likely to have entered care regardless of whether they live in urban, suburban, or rural communities, there are major differences by geographic region. To identify when these differences occur, we fit a discrete-time hazard model predicting probability of placement using both predictors. We continued to find no effect of urbanicity (although we retain this variable in all models because of its role in the design of the NCCS). Detailed investigation of the effects of region revealed that not only does the child's probability of placement vary by geographic region, the entire shape of the hazard profile differs as well (see Model 1 of Table 2). These effects are portrayed in Figure 2, which displays fitted hazard and survivor functions for children from each of the four regions.[9]

The easiest way to interpret the complex regional effects is to focus on how they change as children grow up. During infancy, a child's probability of initial placement differs dramatically across regions. In the first 6 months after birth, for example, an estimated 29% of children in the South and 25% of children in the Midwest enter care versus 16% of children in the Northeast and 18% of children in the West. We can use the parameter estimates in Table 2 to translate these effects into another metric: the comparative odds of initial placement for children by region. In comparison to a child born in the Northeast, for example, the odds that a child born in the South will enter care during the first 6 months are 2.1 times higher and the odds that a child born in the Midwest will enter care are 1.8 times higher. As children age, regional differences

[8] One reviewer suggested that 6- or 12-month leave policies, offered by some employers, may contribute to this zig-zag pattern as well.

[9] These fitted hazard and survivor functions were computed from Model 1 of Table 2 by setting the urbanicity dummy variables at their mean.

Table 2

Results of Fitting a Series of Discrete-Time Hazard Models Predicting First Entry into Regular Nonparental Care

Predictor	Model 0		Model 1		Model 2		Model 3		Model 4	
	Estimate	*SE* (est)	Estimate	*SE* (est)	Estimate	*SE* (est)	Estimate	*SE* (est)	Estimate	*SE* (est)
Time (in months)										
0–5	−1.2221	(0.06)	−1.0248	(0.15)	−3.0197	(0.29)	−3.037	(0.30)	−1.936	(0.35)
6–11	−2.2380	(0.10)	−2.0513	(0.19)	−3.5776	(0.30)	−3.623	(0.31)	−2.329	(0.34)
12–17	−2.5221	(0.12)	−2.3444	(0.21)	−3.4100	(0.36)	−3.451	(0.36)	−2.119	(0.54)
18–23	−2.6841	(0.17)	−2.5267	(0.31)	−3.1593	(0.53)	−3.228	(0.53)	−1.960	(0.54)
24–29	−2.2377	(0.16)	−2.1019	(0.30)	−2.2907	(0.57)	−2.369	(0.54)	−1.182	(0.53)
30–35	−3.4651	(0.26)	−3.3536	(0.34)	−3.1039	(0.70)	−3.190	(0.67)	−2.119	(0.65)
36–41	−2.0658	(0.20)	−1.9713	(0.33)	−1.2755	(0.74)	−1.360	(0.71)	−0.400	(0.70)
42–47	−4.6995	(0.69)	−4.6274	(0.72)	−3.5254	(1.11)	−3.628	(1.06)	−2.823	(1.06)
48–53	−2.1688	(0.29)	−2.1342	(0.43)	−0.6043	(0.91)	−0.704	(0.86)	0.008	(0.84)
54–59	−3.8999	(0.73)	−3.9202	(0.84)	−2.0280	(1.08)	−2.053	(1.02)	−1.457	(1.04)
Urbanicity										
Urban			0.2194	(0.15)	0.1414	(0.14)	0.093	(0.13)	0.165	(0.13)
Suburban			0.0656	(0.15)	0.0186	(0.14)	0.038	(0.14)	0.134	(0.14)
Region										
Northeast			−0.7605	(0.20)	−0.7903	(0.20)	−0.740	(0.21)	−0.759	(0.20)
West			−0.6126	(0.13)	−0.5479	(0.13)	−0.469	(0.14)	−0.455	(0.14)
Midwest			−0.1994	(0.11)	−0.1930	(0.11)	−0.107	(0.12)	−0.076	(0.12)
Region by time										
Northeast by time[a]			0.0209	(0.01)	0.0200	(0.01)	0.021	(0.01)	0.020	(0.01)
West by time[a]			0.0129	(0.01)	0.0158	(0.01)	0.018	(0.01)	0.017	(0.01)
Midwest by time[a]			−0.0136	(0.01)	−0.0151	(0.01)	−0.013	(0.01)	−0.010	(0.01)
Maternal characteristics										
Education					0.1218	(0.02)	0.140	(0.02)	0.100	(0.02)
Age at first birth					0.0161	(0.01)	0.025	(0.01)	−0.013	(0.01)
Age by time[a]					−0.0030	(0.00)	−0.002	(0.00)	−0.001	(0.00)
Race–ethnicity										
Latino							−0.015	(0.13)	0.002	(0.14)
Black							0.250	(0.14)	0.199	(0.14)
Other non-White							0.388	(0.27)	0.300	(0.31)
Single parent family							0.645	(0.11)	0.635	(0.11)
Work before birth?									1.389	(0.15)
Work by time[a]									−0.040	(0.01)
No. of siblings										
One									−0.592	(0.12)
Two or more									-1.016	(0.15)
No. of siblings by work										
One by work[a]									0.737	(0.18)
Two or more by work[a]									0.1239	(0.28)
Deviance (*df*)	5,662.48(10)		5,598.21(18)		5,494.21(21)		5,390.55(25)		4,934.49(31)	
Δ Deviance (*df*)			64.27(8)****		104.00(3)****		103.66(4)****		456.06(6)****	

Note. Est = estimate.
[a] Interaction term.
**** $p < .0001$.

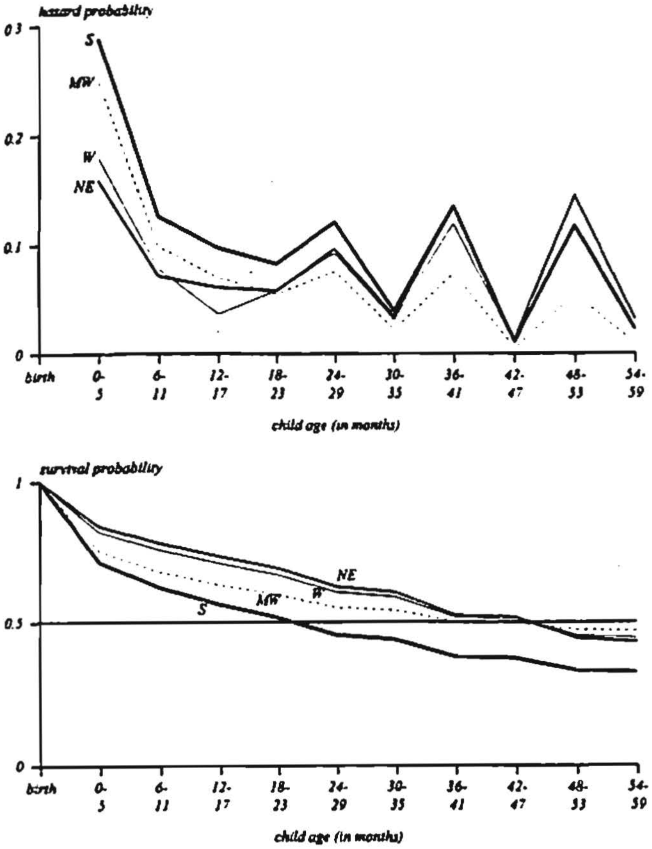

Figure 2. Fitted hazard functions (top panel) and survivor functions (bottom panel) by geographic region. S = South; W = West; MW = Midwest; NE = Northeast.

between the South and both the Northeast and West diminish, and the Midwest begins to stand out with especially low probabilities of placement. Among children between 36 to 41 months who had not yet been placed in care, for example, the probability of initial placement in the Northeast, South, and West hovers between 12% and 13%, in comparison to 7% in the Midwest. Eventually, the regional rank ordering reverses itself almost entirely. Among children 48 to 53 months who had not yet been placed into care, those in the Northeast and West (and to a lesser extent the South) display the greatest probabilities of entry; those in the Midwest display the lowest.

Because so many children born in the South enter care during the first year of life, and because this Southern propensity to use nonparental care is maintained as children age, the estimated median age at placement in this region is only 20 months (see the bottom panel of Figure 2). The average child in the other three regions begins his or her initial arrangement much later: at age 40 months in the Midwest and age 43 months in the Northeast and West (by which time nearly two thirds of children in the South have entered care).

Differences in Timing by Maternal Education and Age at First Birth

Next, we turn to the added influence of maternal and family-level attributes on whether and when children enter care. Consonant with the work of others, we find a positive relationship between mother's education and child-care use. Because we used survival methods, however, we were also able to detect the effect of another maternal characteristic—age at first birth—which has eluded detection in previous studies. The inability to detect this effect has been due, we believe, to two factors: (a) the reliance on linear or logistic regression and (b) the fact that the effect of mother's age at first birth is not constant throughout childhood, but increases in size as children grow older.

One way to understand the effects of maternal education is through closer examination of its parameter estimate in Model 2, Table 2 (.1218). Multiplying by 8 (to represent the 8-year difference between Grades 8 and 16), then antilogging the result, yields an estimated odds ratio of 2.6. This indicates that during every 6-month period studied, the odds that a mother who completed college will place her youngest child into care are 2.6 times higher than are those for a mother who completed only 8th grade. The effects of education can also be seen in the prototypical hazard and survivor functions presented in Figure 3 for mothers who completed 8th grade (left pair of panels), high school (middle pair), and college (right pair).[10]

Because the effects of education persist throughout the child's first 5 years of life, the large period-by-period probabilities cumulate to produce dramatic differences in survivorship (as seen in the three bottom panels). Mothers with only an 8th-grade education are so unlikely to use child care (regardless of when they had their first child) that we cannot even estimate a median age at first placement. Less than one half of these children (and less than one third of those whose mothers first gave birth after age 30) have been in care before turning 5. College-educated mothers, in contrast, are so likely to use child care that the estimated median age at placement for their children (once again, regardless of when they had their first child) is approximately 19 months. By the time these children turn 5, more than two thirds have been in care.

The effects of maternal age at first birth are more complex to interpret because they vary with child age. Moreover, unlike the time-varying effects of geographic region (which diminish over time), the effects of maternal age at first birth increase over time. This pattern can be seen in the increasing size of the gap between the prototypical hazard functions in the three top panels of Figure 3, computed for women who had their first child when they were 20 (the higher function) and 30 (the lower function). During infancy, all mothers are about equally likely to use child care, regardless of how old they were when they started their family. Among toddlers, however, differences begin to emerge, and among preschoolers these differences become pronounced. Given the positive correlation between mother's age at first birth and mother's education ($r = .39$, $p < .0001$ in this sample) and the finding that better educated women are more likely to use child care, we might expect that women who began their families later are also more likely to use child care. But just the opposite it true. The older the woman was when she began her family, the less likely she is to place her children in care.

[10] We emphasize that these are fitted functions estimated from the parameter estimates in Table 2, computed across the full sample; they are not sample functions computed for subgroups of respondents.

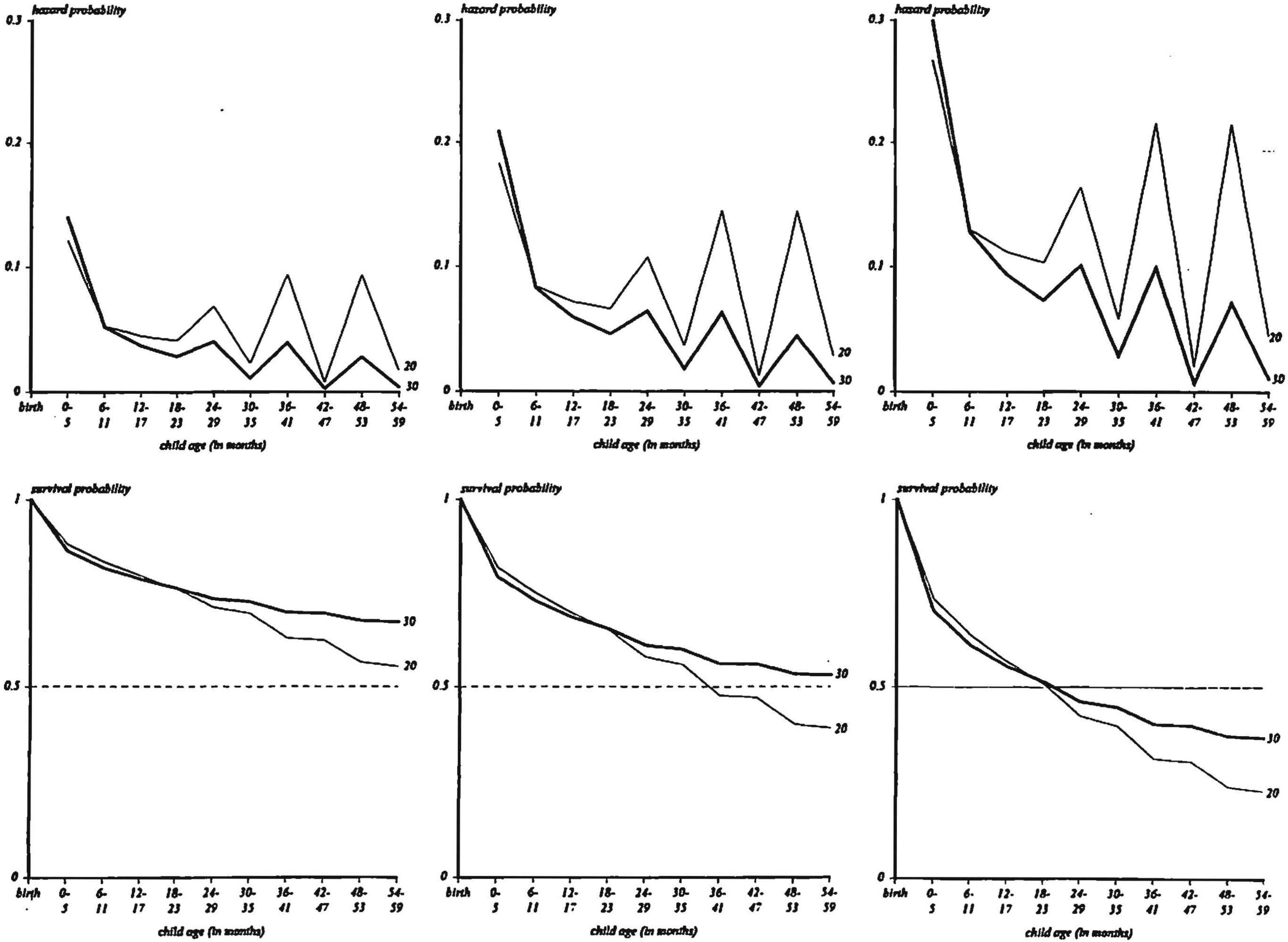

Figure 3. Fitted hazard functions (top panels) and survivor functions (bottom panels) by maternal education and age at first birth. The left pair of panels is for children whose mothers completed 8th grade, the middle pair is for children whose mothers completed high school, and the right pair is for children whose mothers completed college. Within each graph, separate functions are presented by the age when the child's mother had her first child (age 20 and age 30).

We can summarize these time-varying risks by expressing the parameter estimates in Model 2 as odds ratios. Controlling for the effects of maternal education (and all other variables in the model), the odds ratio comparing a woman who had her first baby at age 30 to one who had her first baby at age 20 is 0.6 among children 24–29 months, 0.4 among children 36–41 months, and 0.3 among children 48–53 months. Although these differences appear small, recall that odds ratios are symmetric about 1.0. Thus, this last odds ratio of 0.3 is equivalent to an odds ratio of 3.6. In other words, among 4-year-olds who had not yet entered care, the odds that a child would do so during the next 6 months are 3.6 times greater if the mother began her family at age 20 instead of 30.[11]

Differences in Timing by Single-Parent Status and Ethnicity

At the individual level, the effect of single parenthood on a mother's reasoning about child care is more straightforward than the effect of ethnicity. We examined the effects of single-parent status and ethnicity together, however, not because these variables are related substantively but because they are correlated statistically. In comparison to White and Latino families, Black families are more likely to be headed by a single woman. To ensure that ethnic differentials could not be just as easily attributed to the effects of single-parent status (or vice versa), we modeled both variables simultaneously. We discuss these effects by focusing on the parameter estimates presented in Model 3, Table 2.

Of the two variables, single-parent family status has the larger effect. Controlling for the effects of geographic location, maternal demographics, and ethnicity, the odds of initial entry into care for children from single-parent families are nearly twice as high (1.9) as those for children from two-parent families. Because this effect persists throughout childhood, it cumulates into major differences in the total use of care. Among children who are Black, we estimate that before turning 5, 82% of those from single-parent homes have been in care, in comparison with only 61% of those from two-parent homes. Similar differentials are found for children from other ethnic groups.

The effects of ethnicity are much smaller by comparison and focus on one particular contrast: Black families versus White and Latino families. In each time period under study, the odds that a child who is Black will be placed into care are 1.3 times higher than the odds for a peer who shares all other demographic characteristics, but who happens to be White or Latino. We urge caution, however, when interpreting these small differentials. Upon further control for additional demographic factors (as shown below), they disappear entirely.

Differentials in Timing by Maternal Employment and Number of Children at Home

Some women use child care because they find a job. Some women find a job because they have child care. To ensure that we examined a maternal employment variable that could not be an outcome of the selection process, we assessed whether the mother was known to be working during pregnancy. We hypothesized that women known to be working during pregnancy would be more likely to return to work (and thus seek child care). What we found is that the effect of previous employment varies both by the child's age and by the number of other children in the home. This is an important case of where the theoretical process by which an economic factor (maternal employment) affects child-care selection must be understood in the context of the family's social structure, not based solely on parents' attributes. Our research team has completed some qualitative work to illuminate how these family structure factors may be related to the mothers' reasoning process (Holloway, Fuller, Rambaud, & Eggers-Pierola, 1998).

These relationships are shown in Model 4 of Table 2 and Figure 4. We focus first on the effects of maternal employment for only children—those with no siblings under age 16 at home—by comparing the hazard functions in the top left-hand panel. In the first 6 months after birth, the odds of placing an only child into nonparental care are approximately 4 times higher if the mother was known to be working during pregnancy. As children grow up, the effect of previous employment diminishes among families that have not already placed their youngest child in care. For example, by the time the youngest child is 18 months, the odds ratio drops to near 2, and by the time he or she is 36 months, it has disappeared entirely, falling almost to 1. In other words, immediately after having a first child, previously working mothers are much more likely to place that child in care (and presumably reenter the workforce). Among mothers who choose not to place that first-born child in care during this time period (and who have no other children), the effect of previous employment diminishes, disappearing entirely by the time the child turns 3.

A virtually identical pattern is found for second-born children of mothers who were known to be working before the new child was born (middle panel). The maternal employment effect is large when the child is young and dissipates as the child grows up. In fact, the fitted hazard profiles for children whose mothers were known to be working at birth are statistically indistinguishable, regardless of whether the target child (the youngest child in the family) is the first or second born. The children of women who already had one child and went back to work follow almost the same hazard profile as their peers whose mothers worked while they were pregnant with their first child.

But if the target child is third born or more, the probability of placement is much lower, even if the mother was working before the child was born. We examined where the family size tipping point occurred (i.e., whether increased hazard accrued for each additional child beyond two). We discovered that once there were already two children in the household, we could discern no additional differences. For these families—even ones in which the mother worked before the child was born—the risk of initial placement into child care was quite low. We estimate that only 36% of these youngest children entered care before turning 5.

Child care is not just for working mothers, however. As shown in the bottom hazard functions in each of the top panels of Figure 4, many women not known to be working during pregnancy place their children into nonparental care. Although some

[11] Because the effects of age at first birth begin rather small and escalate over time, the survivor functions in the three bottom panels of Figure 3 are not as informative for interpreting the effects of this variable.

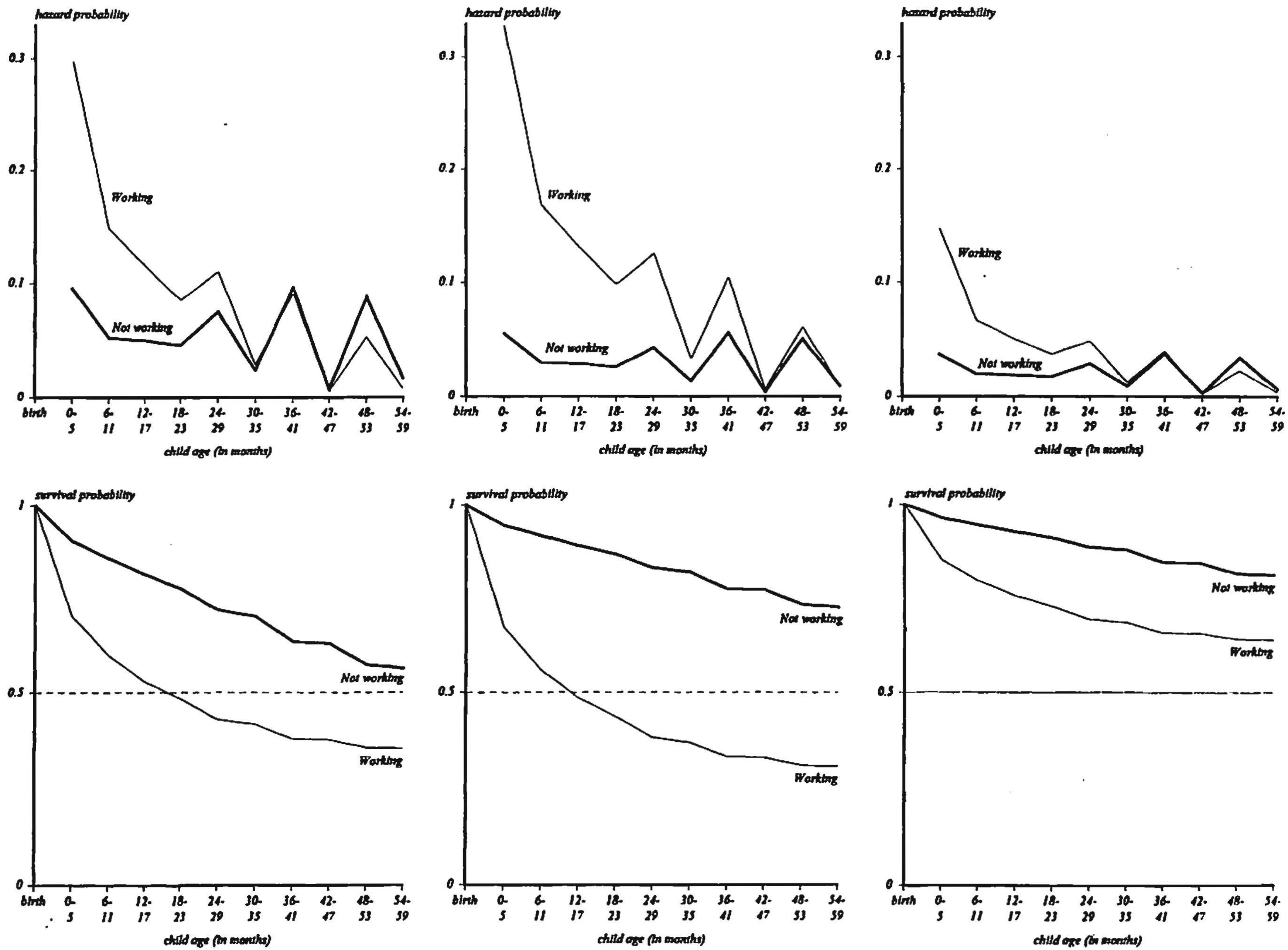

Figure 4. Fitted hazard and survivor functions by mother's employment status during pregnancy and number of siblings. The left pair of panels is for children with no siblings, the middle pair is for children with only one sibling, and the right pair is for children with two or more siblings. Within each graph, separate functions are presented for the children of women known to be working during pregnancy ("working") and not known to be working ("not working").

of these mothers were working during pregnancy (we could not discern this fact from the NCCS data), and others went back to work once the child was born, some use child care for reasons having nothing to do with employment. Yet even among this group, larger families are less likely to place their youngest in care. In comparison with only children, the odds of placement are 1.8 times lower if the youngest child has one sibling and 2.7 times lower if he or she has two or more siblings. Among children of mothers not known to be working during pregnancy, we estimate that 19% of those with two or more siblings, 27% of those with one sibling, and 44% of those with no siblings enter care before turning 5. (The comparable figures for children of mothers known to be working during pregnancy are 44%, 69%, and 64%, respectively.)

Magnitude of Effects From Alternative Selection Factors

Table 3 presents estimates of the median age at first child-care placement for 96 demographic groups (computed from Model 4 in Table 2). These estimates allow us to summarize our results and directly compare the relative magnitude of effects stemming from alternative selection factors. To facilitate interpretation, we have divided Table 3 by maternal employment status during pregnancy. We then sort the demographic groups according to their age at entry. Some groups are so unlikely to use nonparental care before the child reaches age 5 that we cannot estimate a median age at first placement: these groups are denoted by a dash in the table.

First, notice the profound effect of maternal employment during pregnancy. Independent of the child's or family's specific demographic characteristics (e.g., single-parent status, maternal education, race, or mother's age at first birth), if the mother was known to be working during her pregnancy and had no more than two children total (the first two columns of the table), the estimated median age at placement was consistently less than 12 months, usually less than 6 months. Differences by single-parent family status, maternal education, race, and mother's age at first birth pale by comparison. One reason the effect of employment appears so large stems from the time-dependent way it is associated with the hazard function: Its largest effect occurs immediately after birth, then diminishes over time. The

Table 3
Estimated Median Age at First Entry (in Months) into Regular Nonparental Child Care by Selected Maternal and Family Characteristics

	Mom known to be working			Mom not known to be working		
Demographic group	0 sibs	1 sib	2+ sibs	0 sibs	1 sib	2+ sibs
Single parent						
College grad						
Mom at 20						
Black	3.6	3.4	5.6	10.3	20.7	28.9
White	3.8	3.7	7.7	13.2	24.8	36.9
Mom at 30						
Black	3.8	3.6	7.2	13.2	27.1	48.9
White	4.0	3.8	9.8	17.0	36.5	—
High school grad						
Mom at 20						
Black	4.1	3.9	10.2	16.5	28.5	40.8
White	4.5	4.2	13.9	20.9	36.5	51.2
Mom at 30						
Black	4.4	4.1	13.8	22.7	48.1	—
White	4.8	4.5	20.4	27.4	—	—
Two parent						
College grad						
Mom at 20						
Black	4.6	4.3	14.6	21.7	37.2	52.0
White	5.0	4.7	20.3	25.6	41.2	—
Mom at 30						
Black	4.9	4.5	21.9	28.2	—	—
White	6.6	5.0	35.4	37.7	—	—
High school grad						
Mom at 20						
Black	7.1	5.6	27.1	30.1	51.5	—
White	9.5	7.8	39.1	37.3	—	—
Mom at 30						
Black	9.1	7.1	—	49.8	—	—
White	12.4	9.8	—	—	—	—

Note. Dashes indicate that we estimate that less than one half of these children enter into a regular nonparental child-care arrangement before age 5, therefore precluding the estimation of median age at first entry. Sib = sibling; grad = graduate.

immediate and sharp impact of employment on first placement occurs during the child's first few months of life, resulting in these extraordinarily young median lifetimes. This is why nationally we observe that many infants, especially those whose mothers worked during pregnancy, begin their first regular nonparental child-care arrangement before turning 1.

Table 3 also allows us to see that the effects of family size are not linear, but are observed most strongly after a third child is born. The family-size effect also varies according to the mother's employment status during pregnancy. Among working mothers, for instance, no difference in placement is observed between only children and second-born children in two-child families. Consider the child of a married Black high school graduate who began her family at age 20. If the mother was known to be working during her pregnancy, the median age at initial placement is 7.1 months for the first child and 5.6 months for the second; if the mother was not known to be working, the comparable medians are 30.1 and 51.5 months. For mothers who worked during pregnancy, the effect of family size kicks in only when there are three or more children in the home. It is only for this group that median placement ages routinely exceed one year (here, estimated at 27.1 months). It appears that once a mother places her first child in care (so she can go back to work we presume), the child-care decision for the next child is clear. If she stayed home with the first child, in contrast, the second child is either delayed from entering care or never enters.

Differences between other demographic groups—split by mother's age at first birth, mother's education, race, and single-parent status—may appear to be modest. But do not let the extraordinarily large effects of previous employment and family size diminish interpretation of the still large effects of these four family-level variables. When median lifetimes are low (below, say, 12 months), small differences in median lifetimes reflect fairly large effect sizes. For example, the greatest effect among these four variables stems from the mother's single-parent status. The median age at first placement for children from any two-parent family (regardless of race, education, or mother's age at first birth) exceeds the median for any type of single-parent family (with just one exception). To illustrate, consider a child of a Black high school graduate who began her family when she was 20 and who was known to be working when her youngest child was born. If this is her second child, the median age at first placement is 3.9 months if she is a single parent and 5.6 months if she has a partner. This may seem like a small difference. But examination of a comparable pair of median lifetimes for a different family size (or for nonworking mothers) reveals the large difference between two-parent and single-parent households. If this is her second child, for example, the comparable medians are 10.2 and 27.1 months.

As in earlier research on selection processes, we find that maternal education is generally a strong predictor of entry into nonparental child care. But prior research finds that maternal education parallels the positive effects stemming other covariates with which it is positively correlated (such as two-parent status, later age at first birth, and being White). In contrast, we find that maternal education is associated with entry odds in the opposite direction from these related predictors. Mothers who are married, older when they begin their families, and White exhibit a lower probability of placing their youngest child in nonparental care at a young age; yet mothers who are better educated show a higher probability of placement. Moreover, this effect persists even after controlling for the positive effects of maternal employment. In contrast to the effects of maternal employment, family size, and single-parent status, in which differences in median placement ages often reach 3 to 4 years, the differential between mothers with high school diplomas and those with college degrees is often less than 6 months.

Women who delay childbearing are less likely to place their children into regular nonparental care. Although the effects of mother's age at first birth diminish after controlling for maternal employment and family size, they do remain statistically significant. Those who delay childbearing and then are not known to be working during pregnancy are especially likely to delay child care or not use it at all. For example, the first-born child of a married White college-educated woman who started her family at 30, and was not known to be working during pregnancy, entered care at an average child age of 37.7 months. A second- or third-born child in such a household was so unlikely to enter care that we cannot even estimate a median age at first placement.

Viewed against these large demographic differentials, the added effects of race are virtually nonexistent. After controlling statistically for the effects of all other variables, the White–Black differential dissipates to the point that we cannot determine whether it might just as easily be due to sampling variation. This inability to detect a White–Black differential is at odds with much prior research, although much of this work has focused on the placement of youngsters, age 3–5, in center-based programs (Fuller et al., 1996). Our inability to detect a direct ethnic effect may stem from insufficient statistical power; the sample has more than 2,060 White families but only 255 Black families. More likely, it reflects our model's focus on initial placement in any type of child care, as long as the intensity was at least 20 hr per week. In addition, the strong effect stemming from residence in the South (in which the proportion of families who are Black is significantly higher than for other regions) may be substituting for family-level ethnic effects. Children in the South show the highest probabilities of entering care, especially during infancy. This brings us back to the issue of how organization-level supply across regions or communities are conditioning the selection effects of family-level factors.

Discussion: Implications for Future Research and Program Evaluation

Research on children's social environment, and its subsequent effects on cognitive and social development, tends to focus either on home or nonparental child-care settings. For example, researchers focusing on the developmental effects of child care typically try to control for the prior effects of parental background or home factors to isolate the impact of nonparental care settings. The growing line of work on child-care selection, however, emphasizes one crucial point: Parents increasingly shape their young child's development not only within the home but also through their child-care placement decisions. Parental effects occur through direct interaction with the child and indirectly through the mix of nonparental settings in which they place their children (Holloway & Reichhart-Erickson, 1989).

We have shown how young children nationwide are placed in nonparental child-care settings (for at least 20 hr per week) for the first time at widely varying ages, depending on important features of the mother and the wider family structure. Sharp differences in child-care placement profiles are linked to the family's geographic location, an effect that may map against the unequal distribution of child-care supply observed across regions and communities (Fuller & Liang, 1996). Attributes of the mother hold telling effects on whether, and at what age, infants and toddlers are placed in child care, including maternal employment status, educational attainment, and the age at which she began her family. Earlier research also shows that parents' beliefs and early literacy practices further help to predict whether young children are enrolled in center-based programs and preschools (Fuller et al., 1996; Liang, 1996). In sum, these selection factors, whether operating within the family or at the organizational-supply level, move us far ahead in explaining why we observe differences in the length of young children's exposure to nonparental care and preschool programs.

This line of work holds implications for how we conceptualize home and child-care settings that jointly shape early development. First, evaluation research remains impoverished theoretically in that it often starts and ends with the question of how a discrete intervention (e.g., Head Start or child-care providers) directly influence developmental outcomes. Rarely taken into account is the stream of settings that children experienced earlier, including the home and a wide array of nonparental placements. Researchers earnestly attempt to control for such preexisting differences. Yet, eagerness to show the effects of child care should not inadvertently lead us to severely narrow how we define the variety of settings in which children are raised. So, too, interaction effects between home and child-care settings remain entirely underspecified.

Second, the field rarely articulates and models a set of social-environmental conditions under which early interventions are more, or less, likely to exert effects on early development. We sense that the conditions are important—such as variable home environments—but these a priori and concurrent settings are rarely observed and measured by those who focus on nonparental care settings. Our work suggests that an increasing number of infants and toddlers enter into these settings at younger and younger ages, long before organized or formal kinds of child care are encountered. The quality of these settings and the duration of exposure is important to understand, independent of the study of organized early interventions. For some populations, particularly low-income mothers losing welfare benefits, these early nonparental care settings may be changing substantially, in ways that place early development further at risk.

In this same light, we need to learn more about how and why many parents make positive selection decisions, placing their children in stimulating and warm nonparental care settings. Early Head Start, home visitor, resource and referral, and integrated-service programs all aim to provide richer consumer information to parents about their child-care options. As a result, public policy may indeed increase the influence of selection factors. Ironically, very little research is being conducted on how parents understand their options in this mixed market of informal and organized child care and preschooling, and whether intervention into the home alters selection behavior. If it does, then the child-development research community must become more serious in understanding selection processes before claiming effects of particular interventions.

Underlying these issues is the fact that we know very little about the psychological and cognitive reasoning processes used by parents that lead to their selection choices. Our findings help inform this question. For example, we found that better educated women who delay child bearing and have smaller families are less likely to place their child in nonparental care at a young age. Given their greater economic resources and ability to take a break from formal employment, these mothers appear to focus their energy on child rearing. Yet, we also found that a very different group of women—those from working-class backgrounds with more than two children—also stay at home, relying less on nonparental child care. Is it thus unclear just how mothers' variable education levels, child-bearing preferences, and gender roles enter into their reasoning about child care. With regard to low-income mothers, new evidence pertaining to their information processing and more tacit cultural scripts, both influencing child-care selection, is beginning to emerge (Holloway et al., 1998). Future research should be targeted on delineating whether and how proximal determinants of selection, manifest in parental practices and beliefs, can explain how the home environment directly affects early development.

From a methodological standpoint, those who heed this call to more carefully study full child-care histories will need to learn more about new longitudinal methods of analysis. Traditional methods do not allow researchers to look across the full age spectrum and uncover time-varying effects. For studying age at entry into child care (and the further study of the duration of each arrangement) survival analysis is the preferred analytic approach. Using survival methods, researchers can detect whether different predictors are more (or less) important at different points in a child's life. This is an important facility because the factors affecting child-care decisions for an infant will likely differ from those for a toddler or preschooler. Although these techniques are becoming popular in many fields, they are only beginning to be used in developmental psychology (for exceptions, see Capaldi, Crosby, & Stoolmiller, 1996; Eckenrode, 1993).

Our finding of geographic differences in child-care selection suggests the potential need for reinterpretation of previous studies; it also has implications for the design of future studies. Although many child-care researchers collect national data, others focus on a single community or handful of localities (for a review, see Fuller et al., 1996). In this article, we have shown that selection processes vary substantially across the country, both in level and in shape, suggesting that patterns uncovered in one area may not generalize to other localities. A study of child care in the Northeast, where many parents delay placement until the preschool years, could convey the mistaken impression that most parents wait until the preschool years (even though the estimated average age at placement in the South is 20 months). Future studies of child-care availability must be designed with this geographic variation in mind. This would help us better understand the conditions under which developmental effects may be observed, conditions characterized by family-level attributes and processes, as well as organization-level supply conditions within which parental selection occurs.

Finally, we must note that although our analytic methods are longitudinal in character, the data we analyzed were gathered cross-sectionally through retrospective reports of parents. Causal conclusions are not warranted. When we identify the effects of a predictor, we cannot definitively state that this particular variable drives any part of the child-care selection process. At best, we can comment on associations. Prospective data collection, in which researchers track families as they make child-care choices, is required to pin down the sequence of decisions that families actually make. Notwithstanding these limitations, our analysis documents how the majority of young children have already been exposed to nonparental child-care settings before turning 3, with enormous variation in the number of years that they experience before entering kindergarten. Much remains to be learned about how the quality of these early care settings influences early child development. But the first step is to recognize just how early infants and toddlers are being placed in social environments, which, after placement, are not controlled by parents. These selection decisions determine the length of exposure to child care and may shape the quality of care chosen as well.

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