Research Article

CHILD-CARE STRUCTURE → PROCESS → OUTCOME: Direct and Indirect Effects of Child-Care Quality on Young Children’s Development

NICHD Early Child Care Research Network

Abstract—With data from the NICHD Study of Early Child Care, we used structural equation modeling to test paths from structural indicators of child-care quality, specifically caregiver training and child-staff ratio, through a process indicator to child outcomes. There were three main findings: (a) Quality of maternal caregiving was the strongest predictor of cognitive competence, as well as caregivers’ ratings of social competence; (b) quality of nonmaternal caregiving was associated with cognitive competence and caregivers’ ratings of social competence; and (c) there was a mediated path from both caregiver training and child-staff ratio through quality of nonmaternal caregiving to cognitive competence, as well as to caregivers’ ratings of social competence, that was not accounted for entirely by family variables. These findings provide empirical support for policies that improve state regulations for caregiver training and child-staff ratios.

One of the most robust findings in the early-childhood literature is that good child-care quality is associated with a variety of positive outcomes for young children. Specifically, children in higher-quality child-care programs perform better on measures of social, language, and cognitive development when compared with other children (Clarke-Stewart & Allhusen, in press; Lamb, 1998). Quality is typically measured by both process features, such as caregiving quality, and structural features, such as child-staff ratio. By definition, process variables are assumed to have a direct impact on children’s development, whereas structural variables are assumed to have an indirect impact via process quality (see Friedman & Amadeo, 1999).

Researchers who ask policy questions have focused on structural variables, which are regulable, whereas researchers more interested in early experience have focused on process variables. Findings from the Staffing Study (Howes, Phillips, & Whitebook, 1992) and the NICHD Study of Early Child Care (NICHD Early Child Care Research Network, NICHD ECCRN, 1999) show that structural variables matter, such that children attending programs where caregivers have more training and where child-staff ratios are smaller perform better across a range of measures. There is also evidence that process variables matter, especially in the case of sensitive and responsive caregiving, which has been associated with better developmental outcomes in most child-care research projects, including the Bermuda Study (McCartney, 1984; Phillips, McCartney, & Scarr, 1987), the Chicago Study (Clarke-Stewart, Gruber, & Fitzgerald, 1994), the Child Care and Family Study (Kontos, Howes, Shinn, & Galinsky, 1995), the Cost, Quality, and Outcomes Study (Peisner-Feinberg & Burchinal, 1997), and the NICHD Study of Early Child Care (NICHD ECCRN, 1998, 2000b).

Not surprisingly, there is some evidence that structural measures of child-care quality are associated with process measures. Across different types of infant, toddler, and preschool programs, lower child-staff ratios, smaller group sizes, and more caregiver education and training each predicted better child-caregiver interactions (Fischer & Eheart, 1991; Galinsky, Howes, & Kontos, 1995; NICHD ECCRN, 1996, 2000a; Phillips, Mekos, Scarr, McCartney, & Abbott-Shim, 2001; Stith & Davis, 1984). The best evidence of a link between structural and process measures of child-care quality comes from Howes, Smith, and Galinsky (1995), who found that the introduction of stricter standards in Florida led to improved quality of caregiving, as well as better child outcomes.

To summarize, studies have documented three types of associations: those between structural and process features of child-care quality, those between structural features of child-care quality and child outcomes, and those between process features of child-care quality and child outcomes. In no investigation to date, however, have these three associations been brought together in a single analytic model. Of particular interest is the often assumed but never tested mediated path from structural features of child-care quality through process features to child outcomes. The purpose of the present study was to test this mediated path using structural equation modeling (SEM).

USING SEM TO TEST MEDIATION

Baron and Kenny (1986) provided the seminal discussion of mediation with respect to regression. They argued that mediation occurs in a model when including a hypothesized mediator reduces the association between a predictor and an outcome. SEM has become a preferred analytic approach for testing mediation; for example, it provides an ecological description of associations among family factors, child-care factors, and child-outcome factors. With SEM, interrelations among factors, both direct and indirect, can be tested. Note that neither regression analysis nor SEM can distinguish between a “true” indirect path (e.g., the effect on child outcomes due to child-care experiences per se but associated with the family placing their child in that setting) and a “spurious” indirect path (e.g., the effect on child outcomes solely due to more advantaged families selecting better quality care).

Figure 1 presents a general child-care model that has guided researchers. Note that, as with child care, there are two kinds of family variables: structural (e.g., mothers’ education) and process (e.g., maternal caregiving). These two family variables are each linked to both structural and process child-care quality, as well as to child outcomes; the link is direct for family process and indirect for family structure. Structural child-care quality is directly linked to process child-care quality. Finally, process child-care quality is linked to child outcomes. We predicted that the mediated path (in boldface), from structure to process to outcome, would be significant; further, we predicted that this mediated path would not be accounted for entirely by family influences. In the study we report here, we focused on child care at age 54 months, a time when the greatest proportion of children are in care, and we included all types of child care, from center-based care to home care to relative care.

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METHOD

Overview of the Study Design

Data from the NICHD Study of Early Child Care were used in this investigation. A detailed description of the study design can be found in NICHD ECCRN (1994). In brief, children were followed from birth through 54 months. Mothers were interviewed in person when their infants were 1 month old. Measures of the family environment and measures of primary child-care settings were obtained when the children were 6, 15, 24, 36, and 54 months old. The 54-month child-outcome data used in this report were obtained from parents' ratings, caregivers' ratings, and laboratory assessments. All of the data collectors were trained and certified on data-collection procedures. Their performance was monitored centrally to ensure uniform, high-quality data across sites.

Participants

Families were recruited through hospital visits to all 8,986 women giving birth during selected 24-hr intervals at 10 sites. Approximately 60% met eligibility requirements and agreed to consider participating in the study. A stratified random sample of those families was selected, and 1,364 participated in the 1-month home visit. The families in this study are not a nationally representative sample. Nevertheless, the resulting sample was diverse; 24% of the children were ethnic-minority children, 11% of the mothers had not completed high school, and 14% of the mothers were single mothers. When the children were 54 months old, 1,083 were still enrolled in the study. The 1,083 participants differed from the 281 children who were recruited but lost to follow-up: Mothers of participants had significantly more education ($M_s = 14.4$ years vs. 13.6 years), had higher family incomes (income/poverty ratio: $M_s = 3.6$ vs. 3.2), and were more likely to have a husband or partner in the household (85% vs. 76%); their children were less likely to be African American (11% vs. 19%). The 813 children who were in 10 or more hours per week of observable child care at age 54 months and had been in that setting for at least 6 months were included in this report. The sample sizes in the structural equation models ranged from 656 to 789 because incomplete data were addressed with pair-wise deletion.

Measures

Process measures of child-care quality

The Observational Record of the Caregiving Environment (ORCE) was developed to assess characteristics of child-care quality generally and nonmaternal caregiving specifically (NICHD ECCRN, 1996). Behaviors were coded during observations of primary child-care settings for two 44-min cycles when the children were 54 months old. Then eight qualitative ratings were made. Four assessed the caregivers' relationship with the children (sensitivity to nondistress, detachment, stimulation of cognitive development, and intrusiveness), and four assessed the classroom setting (chaos, overcontrol, positive emotional climate, and negative emotional climate).

Structural measures of child care

We chose to focus on the two most policy-relevant indicators: caregivers’ training and child-staff ratio. Information on caregivers’ training in child development or early-childhood education was obtained from interviews with the caregivers, and each caregiver’s training was scored on an ordinal scale as follows: none (0), high school (1), vocational or technical school course (2), college course (3), bachelor’s degree (4), master’s degree (5), or Ph.D. (6). Child-staff ratios were recorded by child-care observers at the beginning and end of each ORCE cycle. An average across the four recordings was then computed.

Family background

Two well-established measures of family background were selected: mothers’ education in years and an income-to-needs ratio (i.e., total family income, including government payments, divided by the appropriate poverty threshold for the household, as determined by the U.S. Department of Labor).

Maternal caregiving

Three measures of the quality of maternal caregiving were included. First, a composite measure of maternal sensitivity was created.
from observers’ ratings of structured play sessions (NICHD ECCRN, 1997). When the children were 6, 15, and 24 months old, the score reflected the sum of the average ratings of sensitivity to nondistress, positive regard, and intrusiveness (reversed). When the children were 36 and 54 months old, the score reflected the average sum of ratings of supportive presence, respect for autonomy, and hostility (reversed). Second, overall quality of the physical and social resources available to the child in the family context was assessed through the Home Observation for Measurement of the Environment (HOME; Caldwell & Bradley, 1984). The infant/toddler version was used at ages 6 and 15 months, and the Early Childhood version was used at ages 36 and 54 months. Third, when the infants were 1 month old, mothers completed a questionnaire that assessed nonauthoritarian child-rearing attitudes and values (Schaefer & Edgerton, 1985).

**Cognitive competence**

Seven measures of competence were used to form a latent variable. Two were from the Woodcock-Johnson Tests of Cognitive Ability (Woodcock & Johnson, 1989, 1990): Incomplete Words and Memory for Sentences; two were from the Woodcock-Johnson Tests of Achievement (Woodcock & Johnson, 1989, 1990): Letter Word Identification and Applied Problems; and two were from the Preschool Language Scale (Zimmerman, Steiner, & Pond, 1992): Auditory Competence and Expressive Language. The seventh measure was the number of omission errors (i.e., missed target stimuli) derived from the Continuous Performance Task (Rosvold, Mirsky, Sarason, Bransome, & Beck, 1956), a measure of sustained attention.

**Caregivers’ and mothers’ ratings of social competence**

Caregivers and mothers each completed the Child Behavior Checklist (Achenbach, 1991), a measure of internalizing and externalizing behavior problems. In addition, caregivers completed the California Preschool Social Competence Scale (Levine, Elzey, & Lewis, 1969), and mothers completed the Social Skills Rating System (Gresham & Elliott, 1990), each yielding a measure of social skills.

**RESULTS**

**Data Analysis Plan**

Specified models were tested using the Amos Version 3.6 program (Arbuckle, 1997). The models were fitted in two stages. First, measurement models were fitted, yielding latent variables, and then hypothesized models were fitted, including both latent and manifest variables. Indirect paths were tested using a $t$ statistic developed by Sobel (1982).

Multiple indices of fit were examined because the chi-square overall goodness-of-fit test statistic is adversely affected by a large sample size, model misspecification, or violation of distribution assumptions (Bollen, 1990). Following Hu and Bentler (1999), we selected three widely used measures to supplement the chi-square statistic. The standardized root mean squared residual (SRMR) is an absolute fit index that measures how well the sample covariance structure fits the proposed model’s covariance structure and is less sensitive to latent-structure misspecification than chi-square. Hu and Bentler suggested that a value less than or equal to .08 represents good fit. The root mean square error of approximation (RMSEA; see Browne & Cudek, 1993; Steiger & Lind, 1980) and Bollen’s goodness-of-fit index (Bollen GFI; see Bollen, 1990) each determine factor-loading misspecification. Both are less sensitive to measurement-model misspecification than chi-square. The RMSEA, like the SRMR, is an absolute fit index. Hu and Bentler suggested that a value less than or equal to .06 represents good fit. The Bollen GFI is an incremental fit index that measures the proportion of improvement in fit between the model and a null model. Hu and Bentler suggested that a value greater than or equal to .95 represents good fit.

**Measurement Models**

The two demographic characteristics, maternal education and family income-to-needs ratio, were regarded as manifest exogenous variables. The three measures of maternal caregiving (i.e., maternal sensitivity with the child, overall level of stimulation and responsiveness on the HOME, and traditional beliefs of the mother) were combined into a single manifest exogenous variable using principal-components analysis; loadings were $-.81$ for traditional child-rearing attitudes, $.85$ for the across-time mean HOME total score, and $.86$ for the across-time mean maternal-sensitivity score.

Child-care quality was represented by both structural and process measures. The structural variables (training and child-staff ratio) were not used as indicators of a latent variable because each has independent policy relevance; note that the correlation between the two was only $.16$ ($p < .001$). Thus, the two structural variables were investigated in separate models. A latent nonmaternal-caregiving variable was indicated by three child-level process measures (caregivers’ sensitivity, caregivers’ detachment, and caregivers’ cognitive stimulation) and a classroom-level process measure (positive emotional climate). The goodness-of-fit statistics for the nonmaternal-caregiving model all indicated good fit, $\chi^2(2) = 5.27, p = .072; SRMR = .009; RMSEA = .047$; and Bollen $GFI = .99$. The remaining four child- and classroom-level variables were excluded from the measurement of nonmaternal caregiving because their exclusion improved the fit of the measurement model substantially, $\chi^2(7) = 154.13, p < .001$.

Each structural equation model included one of three child-outcome latent variables. Outcome latent variables were estimated to describe cognitive competence, the caregivers’ perceptions of social competence, and the mothers’ perceptions of social competence. The cognitive-competence latent variable was indicated by measures of the child’s Incomplete Words, Letter Word Identification, Applied Problems, Memory for Sentences, Auditory Comprehension, Expressive Language, and omission errors. Modification indices indicated that correlating the errors among the measures of cognitive competence, the measures of processing and memory, and the measures of language and attention would greatly improve the overall goodness of fit. Three of the four goodness-of-fit statistics for the cognitive model indicated good fit, $\chi^2(11) = 25.42, p = .008; SRMR = .02; RMSEA = .04; and Bollen GFI = .99$.

Two social-competence latent variables were constructed, one for caregivers’ ratings of children and the other for mothers’ ratings. Three ratings defined each of these latent variables: internalizing behavior problems, externalizing behavior problems, and social skills. The fit of these models could not be tested because exactly three exogenous variables defined each of the social-competence latent variables, and thus, these measurement models were saturated. Nevertheless, the loadings indicate that a single underlying latent construct was reasonable for both ratings of social competence (i.e., loadings of $-.64$,
Six models were fitted to the data. Each model included one of the two structural measures of child-care quality (training and child-staff ratio) and one of the three child outcomes (cognitive competence and caregivers’ and mothers’ ratings of social competence). The models also included mothers’ education and family income-to-needs ratio, which were treated as exogenous variables, as well as maternal caregiving, a factor, and nonmaternal caregiving, a latent variable. For each model, the significance of 11 direct paths, 4 indirect paths with three variables, and 3 indirect paths with four variables was tested (outlined in Table 1). Six of the 7 indirect paths concerned family selection of child-care structure and process. Family selection was also assessed by contrasting the fit of each of the six models with the fit of a corresponding model that excluded these 6 indirect paths. The 7th indirect path concerned the hypothesized mediated path from child-care structure to child-care process to child outcome. Hypotheses regarding each of these indirect paths were directional and therefore one-tailed t tests were used to evaluate their fit.

### Structural Equation Models

The two cognitive-competence models were fitted with data from 738 children. Figures 2a and 2b show the results of the analyses involving caregiver training and child-staff ratio, respectively, by displaying the estimated standardized path coefficients and loadings. Table 1 shows the unstandardized coefficients, which were used in the tests for indirect effects. The chi-square tests for training, $\chi^2(77) = 143.2$, $p < .001$, and for ratio, $\chi^2(77) = 155.6$, $p < .001$, were both significant, perhaps as a result of the large sample size. Because the other three indices indicated good fit for caregiver training ($SRMR = .03$, $RMSEA = .03$, and $Bollen GFI = .98$), as well as for child-staff ratio ($SRMR = .03$, $RMSEA = .04$, and $Bollen GFI = .98$), we argue that interpretation of the estimated paths is reasonable.

As expected, indicators of structural child-care quality were related to nonmaternal caregiving. In addition, family variables as well as nonmaternal caregiving were related to child outcomes. There were significant paths from the model from training to nonmaternal caregiving ($\beta = .17$) and from child-staff ratio to nonmaternal caregiving ($\beta = -.10$). There were also significant paths from nonmaternal caregiving to cognitive competence ($\beta = .10$ in the training and ratio models). The betas were relatively small. In contrast, there were larger direct paths between maternal caregiving and cognitive competence ($\beta = .46$).

### Table 1. Unstandardized coefficients representing direct and indirect paths for the models

<table>
<thead>
<tr>
<th>Path</th>
<th>Cognitive competence</th>
<th>Caregiver ratings of social competence</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Training</td>
<td>Child-staff ratio</td>
</tr>
<tr>
<td>Direct effects</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Maternal caregiving → O</td>
<td>2.407 (0.28)**</td>
<td>2.409 (0.28)**</td>
</tr>
<tr>
<td>Mothers’ education → O</td>
<td>0.405 (0.126)**</td>
<td>0.406 (0.126)**</td>
</tr>
<tr>
<td>Income → O</td>
<td>0.219 (0.089)**</td>
<td>0.219 (0.089)**</td>
</tr>
<tr>
<td>Nonmaternal caregiving → O</td>
<td>0.837 (0.306)**</td>
<td>0.818 (0.305)**</td>
</tr>
<tr>
<td>S → nonmaternal caregiving</td>
<td>0.085 (0.019)**</td>
<td>-0.021 (0.008)**</td>
</tr>
<tr>
<td>Maternal caregiving → nonmaternal caregiving</td>
<td>0.088 (0.029)**</td>
<td>0.096 (0.029)**</td>
</tr>
<tr>
<td>Income → nonmaternal caregiving</td>
<td>0.001 (0.016)</td>
<td>0.005 (0.016)</td>
</tr>
<tr>
<td>Maternal caregiving → S</td>
<td>0.017 (0.056)</td>
<td>0.327 (0.129)</td>
</tr>
<tr>
<td>Mothers’ education → S</td>
<td>0.069 (0.031)*</td>
<td>-0.091 (0.073)</td>
</tr>
<tr>
<td>Income → S</td>
<td>0.062 (0.022)**</td>
<td>0.131 (0.052)</td>
</tr>
<tr>
<td>Indirect effects* with 3 variables</td>
<td></td>
<td></td>
</tr>
<tr>
<td>S → nonmaternal caregiving → O</td>
<td>0.071 (0.031)*</td>
<td>-0.017 (0.009)*</td>
</tr>
<tr>
<td>Maternal caregiving → nonmaternal caregiving → O</td>
<td>0.074 (0.04)*</td>
<td>0.079 (0.03)*</td>
</tr>
<tr>
<td>Mothers’ education → nonmaternal caregiving → O</td>
<td>0.001 (0.014)</td>
<td>0.004 (0.014)</td>
</tr>
<tr>
<td>Income → nonmaternal caregiving → O</td>
<td>0.003 (0.011)</td>
<td>0.01 (0.011)</td>
</tr>
<tr>
<td>Indirect effects* with 4 variables</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Maternal caregiving → S → nonmaternal caregiving → O</td>
<td>0.001 (0.004)</td>
<td>-0.006 (0.04)</td>
</tr>
<tr>
<td>Mothers’ education → S → nonmaternal caregiving → O</td>
<td>0.005 (0.003)</td>
<td>0.002 (0.002)</td>
</tr>
<tr>
<td>Income → S → nonmaternal caregiving → O</td>
<td>0.004 (0.002)*</td>
<td>-0.002 (0.002)</td>
</tr>
</tbody>
</table>

Note. Standard errors are given in parentheses. $S =$ structural child-care quality, in this case caregiver training (columns 1 and 3) and child-staff ratio (columns 2 and 4); $O =$ outcome.

*Indirect paths are computed as the product of the included direct paths (e.g., for two paths labeled b, c). The estimated unstandardized coefficient for the indirect path is $b \cdot c$. The estimated standard error is computed as the square root of $(c_1^2(SE^2_b) + (b_2^2)(SE^2_c) + (c_1b_2)(SE^2_{bc}))$. The $t$ ratio is computed as the estimated coefficient divided by the estimated standard error.

*p < .05. **p < .01. ***p < .001.
Fig. 2. Structural equation models predicting cognitive competence from family variables, caregivers’ training (a) or child-staff ratio (b), and nonmaternal caregiving. Significant paths are indicated by asterisks: **p < .01; ***p < .001.
Next, we tested the indirect effects (see Table 1). As hypothesized, significant indirect paths were observed from training, $\beta = .071$, $t(75) = 2.29$, $p < .05$, and from child-staff ratio, $\beta = -.017$, $t(75) = -1.81$, $p < .05$, to cognitive competence as mediated by nonmaternal caregiving. Of the six indirect family paths tested for training and the six indirect family paths tested for child-staff ratio, three paths reached statistical significance: the path from maternal caregiving through nonmaternal caregiving to cognitive competence for training, $\beta = .074$, $t(75) = 1.97$, $p < .05$, and for ratio, $\beta = .079$, $t(75) = 2.04$, $p < .05$, and the path from income through training and nonmaternal caregiving to cognitive competence, $\beta = .004$, $t(75) = 1.71$, $p < .05$. Comparisons of models with and without these six indirect paths from family through nonmaternal caregiving to cognitive competence suggested that these paths taken together added significantly to the model including training, $\chi^2(6) = 50.2$, $p < .001$, and the model including child-staff ratio, $\chi^2(6) = 45.6$, $p < .001$.

**Caregiver report of social competence**

A similar pattern of results emerged for the models predicting caregivers’ ratings of social competence. These models were fitted with data from 656 children (see Figs. 3a and 3b). Again, both chi-square tests were significant, $\chi^2(34) = 72.2$, $p < .001$, for caregiver training and $\chi^2(34) = 101.0$, $p < .001$, for child-staff ratio, and again the other three indices indicated good fit for training ($SRMR = .027$, $RMSEA = .041$, and Bollen $GFI = .98$), as well as for ratio ($SRMR = .032$, $RMSEA = .055$, and Bollen $GFI = .97$). Both caregiver training ($\beta = .19$) and child-staff ratio ($\beta = -.10$) were significantly associated with nonmaternal caregiving, which was associated with caregiver report of social competence ($\beta = .15$). The direct path from maternal caregiving to social competence was larger and significant in both models ($\beta = .20$).

Significant indirect paths from training, $\beta = .017$, $t(34) = 2.73$, $p < .01$, and child-staff ratio, $\beta = -.004$, $t(34) = 1.86$, $p < .05$, to social competence as mediated by nonmaternal caregiving were tested (see Table 1). For the training model, one indirect path from the family measures was statistically significant, namely the path from income level through training and nonmaternal caregiving to social outcomes, $\beta = .001$, $t(75) = 1.62$, $p < .05$. None of the other indirect paths from the family measures through child-care quality were significant in either the training or the ratio model. Comparisons of models with and without those six indirect paths from family to child outcomes showed that including family paths added to the overall fit of the model including training, $\chi^2(6) = 27.0$, $p < .001$, and of the model including child-staff ratio, $\chi^2(6) = 28.9$, $p < .001$.

**Mother report of social competence**

The models predicting mothers’ reports of social competence were fitted for 789 children. As before, both chi-square tests were significant, $\chi^2(34) = 120.2$, $p < .001$, for caregiver training and $\chi^2(34) = 143.5$, $p < .001$, for child-staff ratio. The other three indices indicated an adequate fit for the model that included caregiver training ($SRMR = .048$, $RMSEA = .057$, and Bollen $GFI = .97$) and for the model that included child-staff ratio ($SRMR = .05$, $RMSEA = .064$, and Bollen $GFI = .96$). These models did not provide evidence for a link from nonmaternal caregiving to this child outcome ($\beta = .01$, $p > .05$). Thus, no tests of indirect effects were conducted.

**DISCUSSION**

The model that has guided child-care researchers to date links structural indicators of quality through process indicators to child outcomes. Yet there have been no attempts to test this indirect or mediated path, until now. SEM offers a method for testing mediation within a single analytic model. Using the NICHD Study of Early Child Care data set, we computed six models, one for caregiver training and one for child-staff ratio for each of three outcomes. Note that these two structural variables are both potentially regulable by states.

There were three main findings. First, maternal caregiving was a strong predictor of cognitive competence and a moderate predictor of social competence as rated by caregivers. This finding is hardly surprising, given that contemporary studies on parenting, especially intervention studies, document the importance of sensitive parenting in young children’s lives (see the review by Collins, Maccoby, Steinberg, Hetherington, & Bornstein, 2000). The demographic indicators of family, specifically mothers’ education and family income-to-needs ratio, showed smaller effects for cognitive competence and nonsignificant effects for social competence. These distal measures of family functioning are well-established indicators of family process, but they are only indicators; research has consistently demonstrated that demographic effects are mediated by the home environment (Duncan & Brooks-Gunn, 2000).

The second finding from these models concerns the association between nonmaternal caregiving and both cognitive and social competence. These associations were revealed in our previous reports with 36-month outcomes (NICHD ECCRN, 1998, 2000b), as well as in other researchers’ work (see reviews by Clarke-Stewart & Allhusen, in press; Lamb, 1998). Note that the effect size for nonmaternal caregiving in this study is about 22% of the maternal caregiving effect for cognitive competence but 75% of the maternal caregiving effect for caregivers’ ratings of social competence. In comparing these two effects, differences in psychometric properties of the child-care and family variables should be evaluated (see McCartney & Rosenthal, 2000). Consider that nonmaternal caregiving was based on a single assessment at 54 months, whereas maternal caregiving was based on cumulative data, from 1 through 54 months. Regardless, some researchers would argue that even modest effects may aggregate when large numbers of children are involved, as is the case with child care (see Fabes, Martin, Hanish, & Updegraff, 2000). With respect to caregiver ratings, the relation between the predictor and the outcome also deserves some discussion. The quality of nonmaternal caregiving was negatively associated with the number of problems caregivers reported for children. Maybe caregivers from higher-quality programs are better class managers, maybe they are better interventionists, or maybe there is some bias.

The third finding concerns the heart of this investigation, specifically, the path from structure to process to outcome. This indirect or mediated path was tested for cognitive competence and caregivers’ ratings of social competence in separate models for caregiver training and child-staff ratio. In all four models the indirect effect was significant. These models provide the first empirical evidence of the mediated path: child-care structure $\rightarrow$ process $\rightarrow$ outcomes. Although the arrows suggest a causal path, we recognize that causality cannot be inferred from structural equation models, which rely on correlational data as input.

The child-care effects, both direct and indirect, do not appear to be due to family selection of child-care quality per se. Indirect paths from the family variables to nonmaternal caregiving to each of the outcomes were tested, and only 4 of the 24 such paths were significant. We know that family selection exists, however, because the models were better fitted when indirect paths from family variables to child-care variables
Fig. 3. Structural equation models predicting caregivers’ ratings of social competence from family variables, caregivers’ training (a) or child-staff ratio (b), and nonmaternal caregiving. Significant paths are indicated by asterisks: *p < .05; **p < .01; ***p < .001.
were included than when they were not. It is probably safest to say that family selection into child care occurs but does not account entirely for the path from child-care structure to process to outcome.

As Shonkoff (2000) warned, “The transmission of knowledge from the academy to the domains of social policy and practice is a formidable task” (p. 181). In part, this is because analyses do not always reflect the needs of policymakers. The present analyses, however, were designed to answer policy-relevant questions, and they provide empirical support for policies that improve state regulations for caregiver training and child-staff ratio. We suspect that more caregiver training may lead to better interactions between children and adults, while lower ratios may lead to more interactions. Because there is considerable variability across the United States in state-regulated child-care standards (Morgan et al., 1993), these data will be useful for policy analyses.

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