Popul Res Policy Rev (2010) 29:745–774
DOI 10.1007/s11113-009-9171-4

Center-Based Care in the Context of One-Child Policy in China: Do Child Gender and Siblings Matter?

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Received: 22 February 2009 / Accepted: 25 November 2009 / Published online: 18 December 2009

Abstract We examined the effects of child gender and siblings on center-based care enrollment in the context of China's one-child policy and its tradition of preference to have many children, especially sons. Using data from the China Health and Nutrition Survey (CHNS) 2000 wave and multilevel logistic regression models, we found that children without siblings consistently had higher odds of receiving center-based care than those with siblings, while there was no evidence that child gender mattered. Further analyses did not show evidence that the effects of child gender and siblings were moderated by household and community resources or local one-child policy. However, we did find that the presence of male, older, or school-age siblings (as compared to female, younger, or preschool-age siblings) reduced preschoolers' odds of receiving center-based care. This was possibly because parents valued formal education much more than preschools and thus focused more on boys when they entered elementary schools than on their sisters or younger brothers. These findings suggest that more attention needs to be given to the equal education opportunities for boys and girls as well as for children with and without siblings.

Keywords Center-based care · One-child policy · China · Gender · Sibling

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Introduction

The one-child policy in China has been implemented for three decades. Under this policy, most couples are allowed to have only one child, while ethnic minorities and many rural residents have been granted the freedom to have more than one child. Meanwhile, the demand for center-based care has continuously increased, though parental care and grandparent care are still popular in most Chinese families. Given that traditionally many families prefer to have many children, especially boys, it is of interest to examine the effects of child gender and siblings, including siblings' gender- and age-related characteristics (i.e., male/female, younger/older than the focal child, and preschool/school ages), on children's center-based care enrollment, which was the main goal of this study.

One-Child Policy and the Tradition of Son Preference

China launched a series of family planning policies after the establishment of the socialist regime in 1949, which eventually evolved into the widely known one-child policy in 1979 (Attane 2002; Fong 2002; Riley 2004; Zhai and Gao 2008). In 1982, fertility control and family planning became constitutional duties for couples (Attane 2002). In 2002, the first state legislation of the one-child policy, the Law on Population and Birth Planning, claimed birth planning as the basic national policy and a duty for all citizens (Leung 2003; Winckler 2002). Nevertheless, the one-child policy does not always prevent parents from having more than one child. The 55 ethnic minorities, which account for about 8.4% of the total population, are exempted from the one-child requirement. In many rural areas, couples are allowed to have a second child if their first one is a girl, while some other local officials allow two children in all cases (Attane 2002; Bernman 1999; Hutzler and Chang 2004; Short et al. 2001).

Meanwhile, derived from an agriculture-based economy and influenced by Confucianism, the patriarchal family system in China traditionally emphasized having many children with a preference for sons to perpetuate the surname and lineage of the family (Chow and Zhao 1996; Haughton and Haughton 1995; Li 2004). Sons—especially the eldest son—have traditionally been more highly valued than daughters, but also expected to fulfill responsibilities such as providing elderly care for their parents and maintaining financial and social ties to the family throughout their lives (Short et al. 2001). In many families, children still serve as the primary caregivers for their aged parents and even grandparents. Therefore, parents with only one child may worry more about their elderly care. Those couples whose only child is a daughter see themselves particularly at risk because, once married, the daughter usually moves away to join her husband's family and may be much less likely to return to provide elderly care to her own parents (Chow and Zhao 1996; Jowett 1991; Meulenberg 2004). Moreover, despite recent rapid socioeconomic changes and urbanization, agriculture still dominates the economy in many rural areas, which demands children, particularly sons, as an important source of labor.

Child Care Arrangements in the Context of One-Child Policy

The types of child care arrangements have been closely related to children's short- and long-term developmental outcomes. Prior studies have shown that children who received center-based care tend to have higher cognitive and social skills (e.g., more knowledgeable about the world, self-confident, and engaging interaction with peers and adults) but more problem behaviors (e.g., less polite, agreeable, or compliant, and more aggressive) than children who were cared for by parents or grandparents (Clarke-Stewart and Allhusen 2005; Magnuson and Waldfogel 2005; Smolensky and Gootman 2003). Cross-national studies have also reported long-term positive effects of high-quality center-based care programs, such as improvement in school achievement, social skills, college attendance, health, and future earnings; and reduction in grade retention, high-school dropout rates, teen pregnancies, delinquency, and crime (see reviews by Blau 2001; Currie 2001; Organization for Economic Co-operation and Development [OECD] 2001, 2006; Smolensky and Gootman 2003; United Nations Educational, Scientific and Cultural Organization [UNESCO] and OECD 2002).

Parental care and grandparent care have been the most traditional and economical options and currently still are the most popular child care arrangements in China, especially in rural areas. Meanwhile, under the influence of Confucianism, education has been highly valued in Chinese culture and has played an important role in personal development and social mobility (Zhai and Gao 2008). The benefits of center-based care, especially in promoting children's academic achievement, have been recognized by more and more parents, particularly young and educated parents. Many parents and grandparents, including those who had no or limited educational opportunities during the Cultural Revolution (1966–1976), see the future success of their children or grandchildren as the fulfillment of their own lives and thus are willing to invest heavily in them from an early age (Chow and Zhao 1996; Li et al. 2001). As a result, the demand for center-based care has continuously increased in recent decades (Zhai and Gao 2008).

Nevertheless, the potential conflicts between the one-child policy and the traditional preference to have many children, especially sons, may affect parents' investment in their children, including child care arrangements. Evidence from many Asian and African societies characterized by son preference, such as China, Korea, Vietnam, India, Bangladesh, Pakistan, Nepal, Egypt, Morocco, and Tunisia, shows that boys are more likely to receive breast-feeding, quality food, education, child care, immunizations, and medical treatment than girls (Haughton and Haughton 1995; Hossain and Glass 1988; Li 2004; Mishra et al. 2004; Obermeyer and Cardenas 1997; Park and Cho 1995; Short et al. 2001). In previous generations in China, many parents avoided investing family resources in daughters and sometimes even forced daughters to drop out of school and work to financially support their brothers' education (Fong 2002). Furthermore, if family resources were scarce, parental investments often heavily focused on the eldest son even if he had younger brothers. Without competition from siblings, children without siblings (i.e., only children), including only girls, may be more likely to receive better child care and education compared to those with siblings (Hesketh and Zhu 1997).

Nevertheless, investing a great deal of family resources in their only children, particularly only girls, may still be perceived as risky by some parents, who might worry about having nothing to count on if their only children are not able or willing to provide care for them in old age. As a result, the one-child policy might exacerbate discrimination against girls, especially in rural areas (Wang et al. 2000).

Few empirical studies have directly examined the effects of child gender and siblings on child care in China. The limited number of existing studies found that compared to families with multiple children, one-child families invested substantially more in child care services, academic tutoring, toys, and family activities (Chow and Zhao 1996) and that the only children tended to receive more high-quality care from both parents (Short et al. 2001) and center-based care than their peers with siblings (Kilburn and Datar 2002). The gender of only children did not have significant effects on parents' attitudes regarding devoting family resources to children (Chow and Zhao 1996). Nevertheless, the effects of child gender and siblings' characteristics, such as male/female siblings, younger/older siblings, and preschool-/school-age siblings, on child care arrangements have not been investigated directly in previous studies.

The present study aimed to fill in gaps in the literature by examining the effects of child gender and siblings on center-based care enrollment in the context of one-child policy and the traditional preference to have many children, especially sons, in China. In particular, this study focused on three research questions: (1) whether gender had significant effects on the chance of child's enrollment in center-based care; (2) whether the presence of siblings in the household had significant effects on the focal child's chance of being enrolled in center-based care; and (3) whether siblings' gender- and age-related characteristics (i.e., male/female, younger/older than the focal child, and preschool/school ages) mattered in affecting the focal child's chance of being enrolled in center-based care. To further understand the causal mechanisms, we also examined the moderating roles of household and community resources and the one-child policy in the relationships between center-based care enrollment and child gender and sibling variables.

Methods

Data and Measures

This study used data from the China Health and Nutrition Survey (CHNS). The CHNS was conducted by the Carolina Population Center at the University of North Carolina at Chapel Hill and the National Institute of Nutrition and Food Safety at the Chinese Center for Disease Control and Prevention. It was designed to examine the effects of health, nutrition, and family planning policies and programs on health and nutrition. The CHNS used a multistage, random cluster process to draw the sample in each of the nine provinces (i.e., Guangxi, Guizhou, Heilongjiang, Henan, Hubei, Hunan, Jiangsu, Liaoning, and Shandong). Counties were stratified by income (low, middle, and high) and a weighted sampling scheme was used to randomly select four counties in each province. Villages and townships within the counties and urban and suburban neighborhoods within the cities were selected randomly. Data have been collected every two to four years since 1989. More details about the design of the CHNS can be found on its website (http://www.cpc.unc.edu/projects/china).

This study used data from the CHNS 2000 survey, which included 4,064 households and 15,648 individuals. Center-based care programs in China include nurseries for children under the age of three and kindergartens serving 3–6 years olds (Lee 1992; Wong and Pang 2002; Zhai and Gao 2008). Thus we focused on children who were 6 years old and younger in the CHNS 2000 survey. Since our goal was to examine the effects of child gender and siblings on center-based care enrollment, children with missing information on key variables of gender, age, or siblings (less than 1% in total) were excluded. As a result, the full analytical sample included 784 children from 704 households in 190 communities (i.e., neighborhoods in urban areas and villages in rural areas).

The outcome variable of this study was center-based care, which referred to child care provided by daycare centers, nursery schools, or kindergartens run by the government, communities, private institutions, or work units. Appropriately 16% of children in the full sample received center-based care. Other children received care from parents, grandparents, other relatives, neighbors, and others.

Based on previous research, the covariates in the analyses included child, household, and community characteristics. Child characteristics included gender, age, member of ethnic minority, and the presence of siblings. Birth order and siblings' age and gender have been reported as important factors that influence intra-household educational investments among children (Ejrnaes and Portner 2004; Garg and Morduch 1998; Hatzitheologou 1997). To capture the roles of siblings' gender- and age-related characteristics in affecting the center-based care enrollment of the focal child, we further divided siblings into subgroups, including male or female, younger or older (compared to the focal child), and preschool (0–6 years old) or school (7–18 years old) ages. The presence of preschool-age siblings may indicate competition for child care with the focal children, whereas school-age siblings may dominate the family resources invested in education.

The household characteristics included parents' marital status, ages, education (i.e., primary school or lower, middle school, and high school or higher), and employment, household monthly income, and the presence of grandparents and other adults. The total household monthly income was calculated by adding all of the cash and in-kind incomes from work (e.g., wages, subsidies, and bonuses), farmland and related incomes (e.g., from gardening, farming, raising livestock, and fishing), and other sources (e.g., small business earnings, subsidies, welfare, and in-kind benefits). The total monthly household income was grouped into five categories: less than 500 yuan, 500–1,000 yuan, 1,000–2,000 yuan, 2,000–3,500 yuan, and 3,500 yuan and higher.

The characteristics of communities included urban/rural residence, availability of child care facilities in the community, local one-child policy (i.e., one child only, allowing for a second child if the first one was a girl, or allowing for a second child in all cases), whether local cadres' responsibility for family planning was established, and whether one-child subsidy was provided for one-child families. The local one-child policy variable indicated the local family planning policy, while the other two variables reflected the pressures for local officials and the motivations for families to implement the family planning policy.

A small portion of children in the sample had missing values on some covariates. In particular, about 18% of children had missing data on mother's age and education, 13% on membership of ethnic minority, 11% on father's age and education, 5% on household income, and 2% on whether local cadre responsibility of one-child policy was established and whether one-child subsidy was provided. As a result, complete case analyses on all valid covariates would reduce the sample to 566 children, a reduction of 28% of the original sample ($n = 784$). Such a substantial reduction in sample size would limit the statistical power of the analyses and introduce potential biases (Hill et al. 2004; Little and Rubin 1987). Meanwhile, the method of dummy variable adjustment for missing data, advocated by Cohen and Cohen (1983), has been shown to produce biased estimates (Allison 2001; Jones 1996).

To address the issue of missing data in the sample, we adopted a multiple imputation method. Multiple imputation uses multiple predictions for each missing value of variables based on other observed variables and the assumption that data are missing at random (Hill et al. 2004; Little and Rubin 1987; Rubin 1987, 1996; Schafer 1997; van Buuren et al. 1999). We conducted two-tailed *t*-tests and did not find significant differences between children who had valid values on the observed variables and those who had missing data on these variables. Nevertheless, since whether the data are missing at random can never be thoroughly tested, it is possible that the missing data depended on some unobserved variables or the unseen observations themselves, which could introduce potential biases when multiple imputation methods were used. Our supplemental analyses using complete cases found results that were very similar to the models using data from multiple imputation, which could relieve the concerns regarding the assumption of missing at random in the data.

To conduct multiple imputation, we used the command "ICE" in Stata 10, which implements multiple imputation by chained equations (MICE; Royston 2005; van Buuren et al. 1999). We included province-fixed effects in multiple imputation to control for the heterogeneity across provinces. A bootstrap method was also adopted, since it has the advantage of robustness as it estimates regression coefficients in a bootstrap sample of the non-missing observations (van Buuren et al. 1999). We generated five sets of imputations for missing data and performed analyses within each dataset separately (Hill et al. 2004; Little and Rubin 1987; Rubin 1987; Schafer 1997). Then we used the means of coefficients from the analyses of the five imputed datasets as the final estimates. Their standard errors were obtained using Rubin's (1987) combining rules for multiple imputation, as shown in Eq. 1:

$$\hat{s}_M = \left(M^{-1} \sum_m \hat{s}_m^2 + (1 + M^{-1})(M-1)^{-1} \sum_m (\hat{b}_m - \bar{b})^2 \right)^{1/2} \tag{1}$$

where $\hat{b}_m$ is the estimated coefficient, with a standard error of $\hat{s}_m$ in sample m of M imputed samples ($M = 5$ in our study); and $\bar{b}$ is the final estimate, calculated as the mean of $\hat{b}_m$'s.

Methods

To examine the effects of child gender and siblings on children's chance of receiving center-based care, we used hierarchical generalized linear models (HGLMs) to account for the hierarchical structure of our data in which children were nested within households and households, in turn, were nested within communities. Also known as generalized linear mixed models or generalized linear models with random effects, HGLMs provide coherent modeling framework for multilevel data with nonlinear structural models (Raudenbush and Bryk 2002). In this study, HGLMs also helped capture the multistage, random cluster sampling designs of the CHNS study. Using HLM6 (Raudenbush et al. 2008), we set a three-level Bernoulli model (i.e., a special case of binomial models) to estimate children's log odds of receiving center-based care with child data at Level 1, household data at Level 2, and community data at Level 3. The final estimates were obtained by the multilevel logistic regression models specified below.

Following the recommended practice by Raudenbush and Bryk (2002), the sample model at Level 1 is presented in Eq. 2:

$$Y_{ijk}|\varphi_{ijk} \sim B(1, \varphi_{ijk}) \tag{2}$$

where Y_{ijk} represents the binary outcome of receiving center-based care for child i from household j in community k (i.e., $Y_{ijk} = 1$ receiving center-based care and $Y_{ijk} = 0$ otherwise), and φ_{ijk} is the expected probability of receiving center-based care, which is normally distributed.

The nonlinear link function is shown in Eq. 3:

$$\eta_{ijk} = \log\left(\frac{\varphi_{ijk}}{1 - \varphi_{ijk}}\right) \tag{3}$$

where η_{ijk} is the log odds of receiving center-based care.

As shown in Eq. 4, the Level-1 structural model is:

$$\eta_{ijk} = \pi_{0jk} + \pi_{1jk}G_{ijk} + \pi_{2jk}S_{ijk} + \sum_{n} \pi_{njk}X_{nijk} \tag{4}$$

where G_{ijk} indicates whether the child was a boy or not; S_{ijk} stands for sibling covariates; and $\sum_n \pi_{njk}X_{nijk}$ represents the vector of n other child-level covariates (i.e., age and membership of ethnic minority).

The Level-2 structural model is presented in Eq. 5:

$$\pi_{njk} = \beta_{n0k} + \sum_{p} \beta_{npk}C_{pjk} + r_{njk} \tag{5}$$

where $\sum_{p} \beta_{npk}C_{pjk}$ denotes the sum of p household-level covariates, as detailed above, and r_{nij} is the random effect with the mean of zero.

Lastly the Level-3 model is shown in Eq. 6:

$$\beta_{pqk} = \gamma_{np0} + \sum_{q} \gamma_{npq} S_{qk} + u_{npk} \tag{6}$$

where $\sum_{q} \gamma_{npq} S_{qk}$ is the sum of q community-level covariate and u_{npk} is the random effect representing deviation of community k's coefficient (β_{pqk}) from its predicted value.

In the analyses, we first applied the models above among children with no siblings (i.e., only children; $S_{ijk} = 0$) to explore the effects of child gender on their enrollment in center-based care among only children, controlling for other covariates. Then we extended these models to children with siblings (i.e., $S_{ijk} = 1$) and included sibling covariates (S_{ijk}) in the models to investigate the effects of child gender and sibling on focal children's center-based care enrollment in the full sample. In doing so, first, we added a dummy variable of whether the child has one or more siblings in the model to examine the effects of having any siblings on the focal child's center-based care enrollment in the full sample. Second, to check whether the number of siblings mattered for children with siblings, we included the number of siblings as a covariate in the model among children with one or more siblings. Third, to further investigate the roles of siblings' gender- and age-related characteristics (i.e., male/female, younger/older than the focal child, and preschool/school ages) on the focal child's center-based care enrollment, we focused on the sample with no or only one sibling. This was because that only a small number of children (n = 65; about 8% of full sample) had two or more siblings; they might have siblings of both genders and younger/older or preschool-/school-age siblings, which would make it impossible to detect the effects of siblings' gender-/age-related characteristics. In the sample of children with no or one sibling, we conducted separate models with three sets of sibling covariates: male/female, younger/older than the focal child, and preschool/school ages. As regulated in the one-child policy in many localities and discussed above, child's gender and whether or not he/she has one or more siblings, as well as male/female, younger/older, and preschool-/school-age siblings, may be strongly correlated. Therefore, we also included the interactions between child gender and these sibling covariates in their respective models.

To examine the goodness of fit of the models as well as the robustness of results on key variables, we started from an intercept-only model and incrementally added the sets of covariates. Specifically, in the three-level logistic regressions, we first included intercept only in Model 1, which is also known as an unconditional model. In Model 2, we added gender and sibling variables, the key variables of interest, at Level 1. We then further included other child covariates at Level 1 in Model 3. In Model 4, we added household covariates at Level 2. Model 5 additionally included community covariates at Level 3. Finally, in Model 6, we included the interaction terms of child gender and sibling variables.

To assess the goodness of fit of these models, we adopted the Akaike's Information Criterion (AIC). As a tool for model selection and a measure of the goodness of fit of an estimated statistical model, the AIC indicates the tradeoff between bias and variance, or in other words, precision and complexity, in model construction (Akaike 1974; Burnham and Anderson 2002). The formula for the AIC is shown in Eq. 7:

$$\mathrm{AIC} = 2k - 2\ln(L) \quad (7)$$

where k is the number of independently adjusted parameters and L is the maximum likelihood of the estimated model. The model with the lowest AIC value among several competing models best explains the data with a minimum of free parameters and thus is preferred.

In addition, we also examined the intra-class correlation (ICC) of the multilevel logistic models. As a measure of group homogeneity, ICC is the proportion of the between-group variance to the sum of the between- and within-group variance in the outcome variable (Raudenbush and Bryk 2002; Snijders and Bosker 1999). A greater ICC coefficient indicates that the variation in center-based care enrollment is more likely to be due to the features of communities rather than the characteristics of individual children and families. Since the variance at Level 1 in a multilevel logistic model is fixed to $\sigma^2 = \pi^2/3 = 3.29$, the ICC coefficient is calculated using Eq. 8 (Snijders and Bosker 1999):

$$\rho = \tau_{000}/(\tau_{000} + r_{00} + \sigma^2) = \tau_{000}/(\tau_{000} + r_{00} + \pi^2/3) \quad (8)$$

where ρ is the ICC coefficient, τ_{000} is the between-community variance from Level 3, r_{00} represents the between-household variance from Level 2, and σ^2 is the child-level variance. We examined the ICC coefficients in the unconditional models as well as the residual ICC coefficients after controlling for the covariates.

Finally, to further understand the relationships between center-based care enrollment and child gender and sibling variables, we also investigated the moderating roles of household and community resources and the one-child policy by including the interaction terms between child gender and sibling variables and the potential moderators. These potential moderators included whether the household had a low income (defined as below the median category, which was below 1,000 yuan), whether the household was in urban areas, whether child care facilities were available in the community, and whether the local one-child policy allowed only one child.

Results

Descriptive Statistics

Table 1 presents the descriptive statistics of the outcome variable and the covariates in the full sample ($n = 784$), the sample of only children ($n = 467$), and the sample of children with siblings ($n = 317$), respectively. Overall 16% of children in the full sample received center-based care. Only children were more likely to receive center-based care (20%) compared to children with siblings (9%). The findings from multilevel logistic regressions are presented later to show

Table 1 Descriptive statistics of outcome variables and covariates

	All children ($n = 784$)	Only children ($n = 467$)	With siblings ($n = 317$)
Child characteristics			
Receiving center-based care	0.16 (0.37)	0.20 (0.40)	0.09 (0.29)
Boy	0.55 (0.50)	0.52 (0.50)	0.58 (0.49)
Age of child	3.41 (1.79)	3.32 (1.78)	3.56 (1.79)
Ethnic minority	0.20 (0.40)	0.15 (0.36)	0.26 (0.44)
Having one or more siblings	0.40 (0.49)	0.00 (0.00)	1.00 (0.00)
Family characteristics			
Parents were married	0.89 (0.32)	0.86 (0.34)	0.92 (0.27)
Father's age	30.87 (5.12)	29.76 (4.44)	32.52 (5.61)
Mother's age	29.39 (4.67)	28.34 (4.03)	30.94 (5.10)
Father's education			
Primary school or lower	0.18 (0.39)	0.15 (0.36)	0.23 (0.42)
Middle school	0.56 (0.50)	0.54 (0.50)	0.60 (0.49)
High school or higher	0.25 (0.43)	0.31 (0.46)	0.17 (0.38)
Mother's education			
Primary school or lower	0.32 (0.46)	0.25 (0.44)	0.40 (0.49)
Middle school	0.52 (0.50)	0.53 (0.50)	0.50 (0.50)
High school or higher	0.16 (0.37)	0.21 (0.41)	0.09 (0.29)
Father employed	0.84 (0.37)	0.81 (0.39)	0.88 (0.33)
Mother employed	0.72 (0.45)	0.69 (0.46)	0.77 (0.42)
Monthly household income			
Less than 500 yuan	0.23 (0.42)	0.25 (0.44)	0.20 (0.40)
500–1,000 yuan	0.24 (0.43)	0.21 (0.40)	0.29 (0.45)
1,000–2,000 yuan	0.25 (0.43)	0.22 (0.41)	0.30 (0.46)
2,000–3,500 yuan	0.13 (0.34)	0.13 (0.34)	0.14 (0.34)
3,500 yuan and higher	0.15 (0.35)	0.19 (0.39)	0.08 (0.27)
Grandparent in household	0.56 (0.50)	0.58 (0.49)	0.53 (0.50)
Other adult in household	0.37 (0.48)	0.37 (0.48)	0.37 (0.48)
Community characteristics			
Urban residence	0.23 (0.42)	0.30 (0.46)	0.15 (0.35)
Child care facilities available in community	0.78 (0.42)	0.72 (0.45)	0.86 (0.34)
One-child policy in locality			
One child only	0.41 (0.49)	0.47 (0.50)	0.32 (0.47)
Allowing 2nd child if 1st was a girl	0.21 (0.40)	0.21 (0.41)	0.19 (0.39)
Allowing 2nd child in all cases	0.38 (0.49)	0.31 (0.46)	0.48 (0.50)
Local cadre responsibility established	0.72 (0.45)	0.77 (0.42)	0.64 (0.48)
One-child subsidy provided	0.55 (0.50)	0.60 (0.49)	0.48 (0.50)

Note: Means with standard deviations in parentheses; descriptive statistics were provided in the full sample ($n = 784$), the sample of only children ($n = 467$), and the sample of children with siblings ($n = 317$), respectively whether this descriptive result held after controlling for child, household, and community covariates.

With regard to child characteristics, as shown in Table 1, 55% of children in the full sample were boys. The average age of all children was 3.4 years old and 20% of them were members of ethnic minority groups. The percentage of boys was slightly higher among children with siblings (58%) than among the only children (52%). Compared to the only children, children with siblings tended to be slightly older (3.6 versus 3.3 years old) and were more likely to be ethnic minorities (26% versus 15%). Overall 40% of all children had one or more siblings. Further descriptive analyses (not reported in Table 1) show that 32% of children in the full sample had one sibling, 6% had two siblings, and less than 2% had three to five siblings. As discussed above, the analyses of the roles of siblings' gender- and age-related characteristics focused on children with no or only one sibling. Among children with only one sibling, 42% had a male sibling and 58% had a female sibling; 21% had a younger sibling and 79% had an older sibling; and 40% had a preschool-age sibling while 60% had a school-age sibling.

In terms of family characteristics, Table 1 shows that, 89% of children in the full sample had parents who were married. On average fathers were about 31 years old and mothers were about 29 years old. More than half (56%) of the parents of all children had middle school education while a larger proportion of fathers (25%) received high school or higher education than mothers (16%). The majority of the parents were employed (84% of fathers and 72% of mothers). Nearly one quarter (23%) of children lived in families with a monthly income less than 500 yuan and 15% with 3,500 yuan or higher. More than half (56%) of children had grandparents living in the household and more than one-third (37%) had other adults in the household.

Compared to the parents of children with siblings, only children's parents were less likely to be married or employed but more likely to be younger and more educated. The family income distribution among only children was more polarized than among children with siblings. Only children were also slightly more likely to live with grandparents than children with siblings, while they were equally likely to have other adults living in the household.

In terms of community characteristics, as presented in Table 1, about 23% of children in the full sample lived in urban areas and 78% of them had child care facilities available in the community. Regarding the local one-child policy, 41% of children lived in communities that allowed having only one child, while 21% lived in communities that allowed for a second child if the first one was a girl and 38% in communities that allowed for a second child in all cases. Almost three quarters (72%) of the children lived in communities that established local cadres' responsibility for family planning and more than half (55%) in communities that provided one-child subsidy to one-child families.

Compared to children with siblings, only children were more likely to live in urban areas and less likely to have child care facilities in their communities. As noted above, the units of communities were different for urban and rural areas (i.e., neighborhoods in urban areas and villages in rural areas). On average, communities in rural areas had a larger population (i.e., 4,433 people per village) than those in urban areas (i.e., 3,037 people per neighborhood), and thus might be more likely to have child care facilities. Rural communities also tended to cover a broader geographic area which required more child care facilities in each community. As a result, children in rural areas were more likely to have child care facilities in their communities than their urban peers. In addition, Table 1 also showed that only children were more likely to live in communities that allowed having only one child, or a second child if the first child was a girl rather than allowing for a second child in all cases, than children with siblings. Also, only children were more likely to live in communities that had local cadres' responsibility of family planning established and provided the one-child subsidy than children with siblings.

Effects of Child Gender and Siblings on Center-based Care

Table 2 shows the results of the effects of child gender and siblings on the odds of receiving center-based care from multilevel logistic regressions. These results were combined from the estimates of log odds in the five datasets generated by multiple imputation. HLM6 provides estimates from both unit-specific models and population-average models in the multilevel logistic regressions, which were almost identical in our analyses. The unit-specific approach is suggested as preferable for investigating the association among clustered observations, especially for constructing AIC or ICC (Agresti 2002; Hardin and Hilbe 2002; Pan 2001). Nevertheless, since our research interest was the effects of child gender and siblings on center-based care in the overall population, based on the suggestions in prior research and by the developers of the HLM6 program (Barr 2008; Raudenbush and Bryk 2002; Raudenbush et al. 2008), we adopted the log odds and their robust standard errors from the population-average models to calculate the final estimates. Compared to the unit-specific model, the population-average model is based on fewer assumptions and thus more robust to misspecification of the random effects in the model (Barr 2008; Heagerty and Zeger 2000; Raudenbush and Bryk 2002). The detailed estimates of odds ratios with z-statistics from these five datasets are presented in Appendix Table 4, which shows that these estimates were very consistent across the imputed datasets in terms of both magnitude and statistical significance. Here we only discuss the combined results presented in Table 2.

Panel A in Table 2 shows the results on the effects of child gender in the sample of only children. Panel B presents results among the full sample that focused on the effects of child gender and whether the child had any siblings. Panel C shows results on the effects of child gender and the number of siblings among children with siblings. Panel D presents results among children with no or one sibling regarding the effects of siblings' gender (Panel D1), younger/older siblings (Panel D2), and preschool-/school-age siblings (Panel D3). Only the findings (odds ratios with z-statistics in parentheses) on key variables of children's gender and siblings are presented in Table 2. The full results of Panels B and D1–D3 from one of the five imputed datasets are shown in Appendix Table 5.

As detailed above, Model 1 in Table 2 was the intercept-only, unconditional model. Child gender and sibling variables were added to Model 2, and other child-level covariates were included in Model 3. Household-level covariates were further

Table 2 Effects of child gender and siblings on center-based care enrollment

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Panel A: Sample of only children (n = 467)						
Intercept	0.25** (−10.58)	0.31** (−7.34)	0.05** (−10.15)	0.03** (−3.15)	0.09+ (−1.91)	—
Boy		0.71 (−1.57)	0.70 (−1.53)	0.72 (−1.33)	0.71 (−1.37)	—
Akaike's information criterion (AIC)	471.03	470.18	425.69	417.01	413.52	—
Intra-class correlation (ICC)	0.18	0.18	0.16	0.13	0.10	—
Panel B: Full sample: having siblings or not (n = 784)						
Intercept	0.21** (−13.27)	0.30** (−7.87)	0.06** (−11.48)	0.04** (−3.94)	0.10* (−2.42)	0.11* (−2.45)
Boy		0.74 (−1.58)	0.73 (−1.56)	0.78 (−1.31)	0.76 (−1.33)	0.77 (−1.10)
Having one or more siblings		0.50** (−3.55)	0.46** (−3.99)	0.50** (−2.79)	0.53** (−2.46)	0.54+ (−1.81)
Boy*siblings						0.96(−0.08)
Akaike's information criterion (AIC)	686.61	669.84	611.64	588.13	584.19	585.17
Intra-class correlation (ICC)	0.23	0.21	0.20	0.17	0.15	0.15
Panel C: Children with siblings: # Siblings (n = 317)						
Intercept	0.12** (−11.06)	0.19** (−4.78)	0.05** (−5.57)	0.05 (−1.41)	0.11 (−0.86)	0.08 (−0.98)
Boy		0.74 (−1.01)	0.68 (−1.14)	0.77 (−0.63)	0.73 (−0.71)	2.51 (0.78)
Number of siblings		0.77 (−1.27)	0.76 (−1.17)	1.06 (0.15)	1.23 (0.54)	1.55 (1.16)
Boy*number of siblings						0.35 (−1.03)
Akaike's information criterion (AIC)	200.53	198.03	186.50	183.02	180.52	181.72
Intra-class correlation (ICC)	0.24	0.23	0.23	0.18	0.13	0.12
Panel D: Children with no or one sibling (n = 719)						
Panel D1: Sibling gender ("no sibling" omitted)						
Intercept	0.21** (−13.16)	0.29** (−7.95)	0.05** (−11.44)	0.05** (−3.44)	0.12* (−2.06)	0.13* (−2.04)
Boy		0.79 (−1.36)	0.76 (−1.49)	0.83 (−0.96)	0.81 (−0.98)	0.77 (−1.09)

Table 2 continued

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Male sibling		0.26** (−4.18)	0.22** (−4.15)	0.24** (−3.73)	0.25** (−3.49)	0.16** (−2.69)
Female sibling		0.65 (−1.55)	0.57 (−1.58)	0.65 (−1.43)	0.66 (−1.37)	0.64 (−1.19)
Boy*male sibling						1.99 (0.82)
Boy*female sibling						1.08 (0.15)
Akaike's information criterion (AIC)	647.22	633.47	578.15	557.42	548.49	550.15
Intra-class correlation (ICC)	0.21	0.20	0.18	0.16	0.15	0.15
Panel D2: Younger/older sibling ("no sibling" omitted)						
Intercept	0.21** (−13.16)	0.29** (−7.91)	0.06** (−11.27)	0.05** (−3.54)	0.11* (−2.15)	0.11* (−2.17)
Boy		0.78 (−1.43)	0.74 (−1.49)	0.81 (−1.03)	0.80 (−1.04)	0.77 (−1.09)
Younger sibling		0.82 (−0.55)	0.46 (−1.55)	0.52 (−1.56)	0.57 (−1.31)	0.60 (−1.17)
Older sibling		0.41** (−3.95)	0.40** (−3.72)	0.44** (−2.82)	0.44** (−2.81)	0.38* (−2.13)
Boy*younger sibling						0.92 (−0.13)
Boy*older sibling						1.30 (0.46)
Akaike's information criterion (AIC)	647.22	634.99	581.16	561.23	551.35	553.11
Intra-class correlation (ICC)	0.21	0.20	0.18	0.16	0.15	0.15
Panel D3: Preschool-/school-age sibling ("no sibling" omitted)						
Intercept	0.21** (−13.16)	0.29** (−7.91)	0.05** (−11.32)	0.04** (−3.61)	0.10* (−2.23)	0.10* (−2.28)
Boy		0.77 (−1.47)	0.75 (−1.53)	0.81 (−1.03)	0.80 (−1.05)	0.77 (−1.06)
Preschool-age sibling (ages 0–6)		0.55 (−1.56)	0.48 (−1.59)	0.53 (−1.61)	0.59 (−1.49)	0.74 (−0.70)
School-age sibling (ages 7–18)		0.44** (−3.44)	0.38** (−3.72)	0.42** (−2.85)	0.41** (−2.87)	0.29** (−2.70)

Table 2 continued

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Boy*preschool-age sibling						0.66 (−0.71)
Boy*school-age sibling						1.80 (0.99)
Akaike's information criterion (AIC)	647.22	636.71	581.03	561.09	550.99	552.75
Intra-class correlation (ICC)	0.21	0.19	0.18	0.16	0.15	0.15

Notes: Model 1 included intercept only, Model 2 added gender and sibling variables, Model 3 included additional other child-level covariates, Model 4 added household-level covariates, Model 5 included additional community-level covariates, and Model 6 added the interactions between gender and sibling variables; Panel A shows the effects of child gender in the sample of only children, Panel B presents the effects of child gender and whether the child had any siblings in the full sample, Panel C shows the effects of child gender and the number of siblings among children with siblings, and Panel D presents results among children with no or one sibling regarding the effects of siblings' gender (Panel D1), younger/older siblings (Panel D2), and preschool-/school-age siblings (Panel D3); only the findings on key variables of children's gender and siblings are presented in the above table, while the full results of Panels B and D1–D3 from one of the five imputed datasets are shown in Appendix Table 5; odds ratios with *z*-statistics in parentheses combining results from five datasets generated by multiple imputation detailed in Appendix Table 4; ** $p < 0.01$, * $p < 0.05$, + $p < 0.10$

added to Model 4. Model 5 additionally included community-level covariates. In Model 6, the interaction terms of gender and sibling variables were added. As shown in Table 2, overall the values of AIC in all the panels gradually declined from Model 1 to Model 5, indicating that the goodness of fit of models was improving when the additional sets of covariates were added in the models. Model 5 had the lowest AIC value compared to those of Models 1–4 in the same panels and thus would be preferred. The AIC values in Model 6 in Panels B, C, and D were slightly larger than those of Model 5 in the same panels, which implied that the inclusion of interaction terms tended to make the models overly complex. Since the differences in AIC values between Models 5 and 6 were trivial and it was necessary to check the robustness of findings on gender and sibling variables, we kept the findings from both Models 5 and 6 and discussed them below.

Table 2 also shows the ICC coefficients across models. The ICC coefficients in the intercept-only, unconditional models (i.e., Model 1) were 0.18 in Panel A, 0.23 in Panel B, 0.24 in Panel C, and 0.21 in Panel D, which indicated that approximately 18% to 24% of the variance in center-based care enrollment was attributable to the different characteristics of communities. This also provided evidence for conducting multilevel analyses (Raudenbush and Bryk 2002; Snijders and Bosker 1999). Similar to the trend in AIC values, the residual ICC coefficients also gradually declined from Model 2 to Models 5 and 6, suggesting that the variance in children's center-based care enrollment due to the different features of communities declined as the covariates at the levels of children, household, and communities were included in the models.

As shown in Table 2, overall children's gender did not have significant effects on their odds of receiving center-based care across all four panels (i.e., Panels A–D), while having one or more siblings (in Panel B), especially male siblings, siblings older than the focal children, or school-age siblings (in Panel D), tended to reduce the focal children's odds of enrolling in center-based care. The coefficients and their statistical significance of child gender and sibling variables were quite consistent across models, indicating that although the sample size was relatively small in our study, the findings on the effects of child gender and siblings on center-based care enrollment were robust to the choice of covariates and interactions. Since Models 5 and 6 had the best goodness of fit, as discussed above, we only focus on the results from these two models for discussion below.

Results in Panel B in Table 2 suggest that having one or more siblings had significant negative effects on children's center-based care enrollment. Specifically, compared to only children, having one or more siblings reduced children's odds of receiving center-based care by about 50% (significant at $p < 0.01$ in Model 5 and marginally significant at $p < 0.10$ in Model 6). The interactions between child gender and whether the child had any siblings were not statistically significant, which implied that the effects of whether having any siblings on center-based care enrollment were not statistically different for boys or girls.

Further analyses among children with siblings (presented in Panel C, Table 2) did not find significant effects of the number of siblings on the odds of focal children's receiving center-based care. This finding suggested that whether or not children had siblings mattered (as shown in Panel B) but the number of siblings—as long as they had any—did not matter (as shown in Panel C).

Given that whether or not children had siblings mattered, Panel D in Table 2 further shows the effects of sibling characteristics on focal children's center-based care enrollment. Focusing on siblings' gender, Panel D1 shows that compared to only children, having a male sibling significantly lowered focal children's odds of receiving center-based care by 75% (in Model 5) to 84% (in Model 6). In contrast, having a female sibling did not show significant effects on children's enrollment in child care centers. Panel D2 shows that having an older sibling significantly decreased focal children's odds of receiving center-based care by 56% (in Model 5) to 62% (in Model 6) than only children, while having a younger sibling did not matter. Panel D3 shows that compared to only children, having a school-age sibling significantly reduced focal children's odds of enrolling in child care centers by 59% (in Model 5) to 71% (in Model 6), while having a preschool-age sibling did not matter. None of the interaction terms between child gender and sibling characteristics were statistically significant in any models in Panel D, suggesting that while these sibling characteristics mattered, their effects on the enrollment of center-based care were not statistically different for boys or girls.

As shown in Appendix Table 5, some other covariates also showed consistently significant effects on focal children's odds of receiving center-based care. In particular, older children had higher odds of being enrolled in child care centers, after accounting for all other covariates. Compared to their less advantaged peers, children whose mothers had middle school or more education and those in families with a monthly income of 500 yuan or more had higher odds of receiving center-based care. As expected, having a grandparent in the household lowered the odds of children's center-based care enrollment. Compared to children in communities that allowed having only one child, children living in communities that allowed for more than one child had lower odds of being enrolled in child care centers.

Potential Moderation of Household and Community Resources and One-child Policy

Table 3 presents the results regarding the moderating roles of household and community resources and the one-child policy in the relationships between center-based care enrollment and child gender and sibling variables. The analyses were based on Model 5 of Panel B in Table 2 from the full sample ($n = 784$), using the variable of whether the focal child had one or more siblings. We also conducted supplemental analyses by including the interactions between the potential moderators and specific sibling variables (i.e., male/female sibling, younger/older sibling, and preschool-/school-age sibling in Panels D1–D3) and found similar results. The models in Table 3 included the interactions of child gender and sibling variables with household low income (in Model 1), urban residence (in Model 2), child care facility availability (in Model 3), and one-child-only policy (in Model 4). Model 5 included all the interaction terms in Models 1–4.

As shown in Table 3, overall we did not find evidence that the effects of child gender and siblings on center-based care enrollment were moderated by household

Table 3 Potential moderation of household and community resources and one-child policy in the relationships between center-based care enrollment and child gender and sibling variables

	Model 1	Model 2	Model 3	Model 4	Model 5
Boy	0.77 (−0.95)	0.56 (−1.55)	1.07 (0.14)	0.74 (−1.02)	0.81 (−0.36)
Having one or more siblings	0.51* (−2.04)	0.52* (−2.45)	0.46+ (−1.72)	0.39* (−2.52)	0.36+ (−1.75)
Household low income	0.58+ (−1.73)	0.52* (−2.46)	0.52* (−2.51)	0.53* (−2.47)	0.51+(−1.82)
Household low income*boy	0.84 (−0.40)				1.09 (0.18)
Household low income*sibling	0.99 (−0.02)				0.90 (−0.22)
Urban residence	1.12 (0.39)	0.77 (−0.77)	1.15 (0.49)	1.13 (0.44)	0.80 (−0.62)
Urban residence*boy		1.14 (1.56)			1.26 (1.57)
Urban residence*sibling		0.91 (−0.15)			0.76 (−0.43)
Child care facilities available in community	1.16 (0.55)	1.14 (0.47)	1.27 (1.00)	1.16 (0.54)	1.24 (0.75)
Child care facility*boy			0.60 (−1.00)		0.68 (−0.71)
Child care facility*sibling			1.13 (0.21)		1.14 (0.21)
One-child-only policy	2.68** (3.76)	2.72** (3.78)	2.66** (3.73)	2.32* (2.33)	2.56** (2.58)
One-child-only policy*boy				0.97 (−0.09)	0.77 (−0.62)
One-child-only policy*sibling				1.69 (1.04)	1.85 (1.25)

Notes: Results were based on the full model in the full sample (i.e., Model 5 of Panel B in Table 2; $n = 784$) from one of the five datasets generated by multiple imputation; models included the interactions of child gender and sibling variables with household low income (in Model 1), urban residence (in Model 2), child care facility availability (in Model 3), and one-child-only policy (in Model 4), while Model 5 included all the interactions terms in Models 1–4; odds ratios with z-statistics in parentheses; ** $p < 0.01$, * $p < 0.05$, $^{+}$ $p < 0.10$

and community resources or the local one-child policy. Consistent with the findings above and across models, boys did not show advantages in center-based care enrollment, whereas having one or more siblings significantly reduced children's odds of receiving center-based care regardless of their gender. Children in low-income households were less likely to attend child care centers, but the effects of household low income did not differ for boys and girls or whether they had one or more siblings. Similarly, neither urban residence nor the availability of child care facilities in the communities showed significant effects on center-based care enrollment, and their effects did not differ for boys and girls or only children and children with siblings. Lastly, children in communities that allowed only one child tended to have significantly higher odds of attending in center-based care compared to their peers in communities that allowed two or more children. Nevertheless, the effects of one-child policy on center-based care enrollment did not show significant differences regarding child gender or the presence of siblings.

Discussion and Conclusion

This study aimed to examine the effects of child gender and siblings on center-based care enrollment, a topic that has been of great concern but largely understudied in the literature. Using data from the China Health and Nutrition Survey (CHNS) 2000 wave and multilevel logistic regression models, we found that overall children's gender did not have significant effects on their odds of receiving center-based care. This finding was robust across analyses in the full sample and the samples of only children and children with siblings. Meanwhile, having one or more siblings, especially male siblings, siblings older than the focal children, or school-age siblings, tended to reduce focal children's odds of receiving center-based care. Additional analyses did not find evidence that the effects of child gender and siblings were moderated by household and community resources or the local one-child policy.

The findings clearly showed that children without siblings (i.e., only children) were more likely to receive center-based care than their peers who had siblings. The findings on child gender and siblings' gender- and age-related characteristics were also interesting. On the one hand, there was no evidence that child gender played a determining role in the center-based care enrollment of preschoolers (0–6 years old). On the other hand, the presence of siblings and siblings' gender mattered. Having one or more siblings, especially male, older, or school-age siblings, significantly reduced the focal children's chance of receiving center-based care. Furthermore, the effects of siblings on center-based care enrollment were not moderated by the local one-child policy. In fact, due to intra-household resource constraints, sibling rivalry is neither because of the one-child policy, nor a unique phenomenon in China, but exists across nations and cultures. For example, among the limited number of studies on the effects of siblings on child care arrangements in the U.S., Lehrer (1989) found that preschoolers' probabilities of enrolling in center-based care were significantly reduced when they had siblings, but the ages of siblings (i.e., preschool or school ages) did not matter. Similarly, Liang et al. (2000) showed that children who were first-born in the family and children with fewer siblings were more likely to be enrolled in center-based care. In contrast, Joesch et al. (2006) found that, compared to preschoolers who had younger siblings, those who had older siblings were more likely to receive non-parental care (i.e., relative care, non-relative care, and center-based care) but had no significant differences in the probabilities of attending center-based care in particular.

In terms of the effects of household and community resources, we found that children from low-income households had significantly lower odds of receiving center-based care, whereas urban residence and the availability of child care facilities in the community did not show significant effects on focal children's center-based care enrollment. These findings were robust across models in different samples and with different specifications, including those with interaction terms between these variables and child gender and sibling variables. This suggests that, after controlling for child and household characteristics as well as other community-level variables, including the local one-child policy indicators, the urban or rural residence of focal children and the availability of child care facilities in their communities did not significantly affect their enrollment in child care centers. In addition, these variables of household and community resources, including household income, urban residence, and the availability of child facilities in the community, did not show significant moderating roles in the relationship between center-based care enrollment and child gender and sibling variables. This finding suggests that the effects of child gender and siblings on center-based care enrollment were robust even with variations in the level of household resource, urban or rural residence, or the availability of child care facilities after controlling for child gender, sibling, and other characteristics as well as other household- and community-level covariates.

Furthermore, as discussed above, influenced by Confucianism, education traditionally has been highly valued in the Chinese culture. Nevertheless, in this study overall only 16% of preschool children were enrolled in child care centers. In contrast, more than 40% of 3-year-olds and nearly 70% of 4-year-olds in the U.S. are in some form of center-based care on a regular basis (U.S. Census Bureau 2008; Waldfogel 2006). The low preschool enrollment rates in China suggest that center-based care is probably still considered less important than elementary and higher schooling (the so-called "formal education") by many Chinese parents. However, this indicates a laggard condition in China as compared to many other countries and needs to be addressed. As discussed earlier, cross-national studies have consistently reported long-term positive effects of high-quality center-based care programs. Enrolling in center-based care, especially in high quality child care centers, may indeed lay a solid foundation for these children and give them a head start in multiple dimensions of their development.

The presence of male, older, or school-aged siblings (rather than female, younger, or preschool-aged siblings) significantly reduced children's odds of receiving center-based care. These findings together suggest that gender may still play an important role in families' investment and a school-age male sibling may spend a proportion of family resources that is large enough to significantly reduce his preschool-age sibling's chance to attend child care centers. Therefore, son preference does not prevail in the care arrangements of preschoolers, but probably still exists among school-age children. As reviewed above, derived from cultural customs and social institutions, son preference has been popular in many regions of the world, especially those of Asian and African societies. Recent cross-national studies have shown mixed evidence regarding whether son preference, or parental gender preference in general, decreases along with modernization (e.g., as measured by urbanization, industrialization, female education and employment, and household wealth). Some studies found that modernization reduced son preference (Burgess and Zhuang 2002; Chung and Das Gupta 2007; Lin 2009), while others showed that this either only happened slowly or did not happen at all (Andersson et al. 2006; Das Gupta et al. 2003; Filmer et al. 2008). As long as the cultural and social roots of son preference (e.g., relying on males to perpetuate family surname and lineage, perform rituals, and provide elderly care) exist, it may not disappear automatically over time. As researchers (Chung and Das Gupta 2007; Das Gupta et al. 2003) indicate, in addition to the provision of equal opportunities for females in education and employment, policymakers should also focus on the change of traditional beliefs and social norms on gender roles and equity through interventions such as media campaigns, legislation, women's organizations, and financial incentives. The Chinese government has been making efforts to adopt these measures, yet it will be a long way to change the belief of son preference that has been prevalent for thousands of years in this most populous country in the world.

There are several limitations to the present study. First, although this study made efforts to control for covariates related to the characteristics of children, households, and communities, some omitted factors might still play important roles in moderating the relationships between center-based care enrollment and child gender and sibling characteristics. For example, parents' preference for the gender and the number of children they have may have important effects on child care. Due to the mandatory one-child policy and the traditional preference to have many children, particularly sons, couples may not have had the specific gender or number of children they desired to, which could have played a role in their child care choices. Unfortunately, the CHNS data do not include information related to parents' preferred number and gender of children.

Second, caution should be taken in generalizing the findings of this study because of data and sampling limitations. In the CHNS survey, the nine provinces included are all in coastal and near-inland areas, which are far more developed than those in the western areas. Moreover, although the survey generally used a multistage, random cluster process to draw the sample, both the provincial capital city and the lower income city in each province were intentionally selected. Since we only focused on households with one or more preschool children (i.e., age of six and younger), our analysis sample tended to have higher proportions of households in rural areas (i.e., 77%) and ethnic minority groups (i.e., 20%) compared to the national statistics in the same time period (i.e., 64% and 8%, respectively) (National Bureau of Statistics 2001). Thus with a sample of 784 preschool children narrowed down from 4,064 households in the CHNS survey, the findings of this study may not be generated to children in other areas or the national population. The small sample also made it impossible to conduct further analyses to investigate the effects of the more specified combinations of siblings' gender- and age-related characteristics, including male/female older/younger siblings or male/female preschool-/school-aged siblings. In addition, this study was based on a cross-sectional survey, which could not address the issue of potential endogeneity or reverse causality. Recently the CHNS study released the new data of 2004 and 2006 waves to the public and corrected the problems in the identification numbers across survey years. This makes it possible for researchers to conduct longitudinal studies, which could not be done previously, to examine the effects of child gender and siblings on child care arrangements. Thus our future research will test whether the findings in this study hold using the CHNS longitudinal data and newer waves.

Despite the limitations, the findings of this study provide important implications for future research and policymaking on child care and education in the context of the one-child policy and the traditional preference to have many children, especially sons, in China. This is one of the first empirical studies to directly focus on the effects of gender and siblings on center-based care enrollment in China. Son preference was not found in the center-based care enrollment of preschoolers, but probably still existed among school-age children. This could be due to parents' lack of recognition of the importance of early childhood education and care, rather than the actual non-existence of a gender bias during the preschool years. Policymakers, as they have already been doing, should continue to make efforts to improve the enrollment and quality of center-based care and to increase the equity of education for both boys and girls as well as for only children and children with siblings.

Acknowledgement We are grateful to the editor and anonymous reviewers for very helpful comments and suggestions. We thank the funding support for this study from the Columbia University Weatherhead East Asian Institute and the Chiang Ching-Kuo Foundation for International Scholarly Exchange. We also thank the China Health and Nutrition Survey, funded by NIH (R01-HD30880, DK056350, and R01-HD38700), and the Carolina Population Center and the Chinese CDC for providing data used in this study.

Appendix

See Tables 4 and 5.

Table 4 Effects of child gender and siblings on center-based care enrollment

	No interactions					Interactions with child gender				
	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
Panel A: Sample of only children (n = 467)										
Boy	0.70 (−1.41)	0.71 (−1.33)	0.71 (−1.38)	0.70 (−1.42)	0.72 (−1.30)	—	—	—	—	—
Panel B: Full sample: having siblings or not (n = 784)										
Boy	0.76 (−1.33)	0.74 (−1.45)	0.78 (−1.21)	0.75 (−1.41)	0.76 (−1.31)	0.76 (−1.12)	0.76 (−1.18)	0.78 (−1.02)	0.76 (−1.17)	0.77 (−1.05)
Having one/more siblings	0.53** (−2.51)	0.54* (−2.38)	0.53** (−2.50)	0.54** (−2.54)	0.53* (−2.43)	0.53+ (−1.88)	0.56+ (−1.75)	0.54+ (−1.82)	0.55+ (−1.85)	0.54+ (−1.77)
Boy*siblings						0.98 (−0.04)	0.94 (−0.14)	0.99 (−0.03)	0.97 (−0.07)	0.94 (−0.13)
Panel C: Children with siblings: # Siblings (n = 317)										
Boy	0.71 (−0.79)	0.66 (−0.94)	0.78 (−0.55)	0.72 (−0.78)	0.79 (−0.54)	2.22 (0.73)	2.75 (0.84)	3.25 (0.88)	2.10 (0.70)	2.37 (0.80)
Number of siblings	1.21 (0.51)	1.20 (0.45)	1.19 (0.45)	1.18 (0.49)	1.40 (0.82)	1.49 (1.04)	1.58 (1.24)	1.50 (1.07)	1.44 (1.12)	1.76 (1.49)
Boy*number of siblings						0.37 (−1.04)	0.31 (−1.15)	0.29 (−1.06)	0.42 (−0.96)	0.40 (−1.03)
Panel D: Children with no or one sibling (n = 719)										
Panel D1: Sibling gender (“no sibling” omitted)										
Boy	0.81 (−0.97)	0.80 (−1.05)	0.82 (−0.91)	0.80 (−1.04)	0.82 (−0.94)	0.76 (−1.12)	0.76 (−1.15)	0.78 (−1.04)	0.77 (−1.11)	0.78 (−1.03)
Male sibling	0.26** (−3.43)	0.25** (−3.44)	0.25** (−3.47)	0.25** (−3.55)	0.25** (−3.59)	0.18** (−2.56)	0.17** (−2.66)	0.16** (−2.64)	0.16** (−2.82)	0.16** (−2.84)
Female sibling	0.66 (−1.38)	0.68 (−1.30)	0.67 (−1.32)	0.67 (−1.35)	0.63 (−1.52)	0.61 (−1.30)	0.65 (−1.14)	0.65 (−1.14)	0.66 (−1.10)	0.62 (−1.30)
Boy*male sibling						1.82 (0.71)	1.96 (0.80)	2.02 (0.83)	2.08 (0.87)	2.09 (0.89)
Boy*female sibling						1.17 (0.31)	1.07 (0.14)	1.08 (0.15)	1.03 (0.07)	1.05 (0.09)
Panel D2: Younger/older sibling (“no sibling” omitted)										
Boy	0.80 (−1.04)	0.79 (−1.13)	0.81 (−0.98)	0.79 (−1.09)	0.81 (−0.99)	0.76 (−1.12)	0.76 (−1.16)	0.78 (−1.04)	0.76 (−1.12)	0.78 (−1.03)
Younger sibling	0.56 (−1.36)	0.55 (−1.33)	0.58 (−1.33)	0.59 (−1.23)	0.58 (−1.31)	0.60 (−1.20)	0.58 (−1.20)	0.59 (−1.24)	0.61 (−1.19)	0.63 (−1.04)

Table 4 continued

	No interactions					Interactions with child gender				
	(1)	(2)	(3)	(4)	(5)	(1)	(2)	(3)	(4)	(5)
Older sibling	0.45** (−2.73)	0.46 ** (−2.72)	0.44** (−2.79)	0.44** (−2.93)	0.42** (−2.99)	0.38* (−2.16)	0.40* (−2.05)	0.39* (−2.03)	0.39* (−2.13)	0.36* (−2.35)
Boy*younger sibling						0.90 (−0.17)	0.93 (−0.11)	0.96 (−0.06)	0.96 (−0.07)	0.86 (−0.24)
Boy*older sibling						1.36 (0.54)	1.27 (0.41)	1.29 (0.44)	1.27 (0.41)	1.33 (0.50)
Panel D3: Preschool-/school-age sibling ("no sibling" omitted)										
Boy	0.80 (−1.05)	0.78 (−1.13)	0.81 (−0.98)	0.79 (−1.10)	0.81 (−0.99)	0.77 (−1.09)	0.76 (−1.13)	0.78 (−1.01)	0.77 (−1.09)	0.78 (−1.00)
Preschool-age sibling	0.58 (−1.57)	0.58 (−1.48)	0.59 (−1.51)	0.61 (−1.38)	0.59 (−1.54)	0.72 (−0.78)	0.74 (−0.69)	0.73 (−0.75)	0.78 (−0.57)	0.73 (−0.75)
School-age sibling	0.42 ** (−2.78)	0.43 ** (−2.78)	0.41** (−2.87)	0.41** (−3.02)	0.39** (−3.01)	0.30** (−2.65)	0.31** (−2.61)	0.30** (−2.57)	0.29** (−2.89)	0.28** (−2.84)
Boy*preschool sibling						0.67 (−0.71)	0.65 (−0.73)	0.69 (−0.67)	0.65 (−0.72)	0.67 (−0.70)
Boy*school sibling						1.87 (1.05)	1.77 (0.94)	1.77 (0.95)	1.81 (1.02)	1.79 (0.98)

Notes: Results were based on Models 5 (without interactions) and 6 (with interactions) in Table 2; odds ratios with z-statistics in parentheses from five datasets (1)–(5) generated by multiple imputation; ** $p < 0.01$, * $p < 0.05$, + $p < 0.10$

Table 5 Full results from multilevel logistical regressions

	Having sibling or not		Sibling gender		Younger/older sibling		Preschool-/school-age sibling	
	No Int.	W/Int.	No Int.	W/Int.	No Int.	W/Int.	No Int.	W/Int.
Child gender and siblings								
Boy	0.76 (−1.33)	0.76 (−1.12)	0.81 (−0.97)	0.76 (−1.12)	0.80 (−1.04)	0.76 (−1.12)	0.80 (−1.05)	0.77 (−1.09)
Having siblings or not								
Yes	0.53** (−2.51)	0.53+ (−1.88)						
Boy*sibling		0.98 (−0.04)						
Sibling gender (no sibling omitted)								
Male sibling			0.26** (−3.43)	0.18** (−2.56)				
Female sibling			0.66 (−1.38)	0.61 (−1.30)				
Boy*male sibling				1.82 (0.71)				
Boy*female sibling				1.17 (0.31)				
Younger/older sibling (no sibling omitted)								
Younger sibling					0.56 (−1.36)	0.60 (−1.20)		
Older sibling					0.45** (−2.73)	0.38* (−2.16)		
Boy*younger sibling						0.90 (−0.17)		
Boy*older sibling						1.36 (0.54)		
Preschool-/school-age sibling (no sibling omitted)								
Preschool age sibling (ages 0–6)							0.58 (−1.57)	0.72 (−0.78)
School-age sibling (ages 7–18)							0.42** (−2.78)	0.30** (−2.65)
Boy*preschool-age sibling								0.67 (−0.71)
Boy*school-age sibling								1.87 (1.05)
Other child and family characteristics								
Child age	1.63** (6.87)	1.63** (6.86)	1.63** (6.25)	1.63** (6.23)	1.61** (6.04)	1.61** (6.03)	1.62** (6.20)	1.62** (6.21)
Ethnic minority	1.06 (0.18)	1.06 (0.18)	1.19 (0.49)	1.20 (0.52)	1.15 (0.39)	1.15 (0.40)	1.15 (0.39)	1.15 (0.39)
Parents were married	0.59 (−1.30)	0.60 (−1.30)	0.63 (−1.20)	0.63 (−1.21)	0.66 (−1.08)	0.66 (−1.09)	0.67 (−1.05)	0.66 (−1.08)

Table 5 continued

	Having sibling or not		Sibling gender		Younger/older sibling		Preschool-/school-age sibling	
	No Int.	W/Int.	No Int.	W/Int.	No Int.	W/Int.	No Int.	W/Int.
Father's age	0.98 (−0.50)	0.98 (−0.50)	1.00 (−0.02)	1.00 (−0.05)	1.00 (0.03)	1.00 (0.04)	1.00 (0.01)	1.00 (0.08)
Mother's age	1.00 (−0.04)	1.00 (−0.04)	0.97 (−0.57)	0.97 (−0.53)	0.98 (−0.47)	0.98 (−0.48)	0.98 (−0.42)	0.98 (−0.49)
Father's education (primary or lower omitted)								
Middle school	0.81 (−0.59)	0.81 (−0.60)	0.83 (−0.49)	0.82 (−0.55)	0.82 (−0.52)	0.81 (−0.57)	0.82 (−0.53)	0.82 (−0.54)
High school or higher	0.78 (−0.58)	0.78 (−0.58)	0.74 (−0.68)	0.72 (−0.73)	0.74 (−0.65)	0.73 (−0.68)	0.74 (−0.66)	0.75 (−0.65)
Mother's education (primary or lower omitted)								
Middle school	2.42** (2.66)	2.41** (2.67)	2.07* (2.07)	2.08* (2.09)	2.06* (2.06)	2.07* (2.10)	2.06* (2.08)	2.06* (2.09)
High school or higher	2.34+(1.91)	2.34+ (1.91)	2.17+ (1.66)	2.20+ (1.70)	2.11+ (1.63)	2.13+ (1.66)	2.12+ (1.64)	2.10+ (1.64)
Father employed	0.72 (−0.99)	0.72 (−0.99)	0.63 (−1.50)	0.62 (−1.49)	0.61 (−1.56)	0.61 (−1.50)	0.60 (−1.57)	0.62 (−1.46)
Mother employed	1.12 (0.39)	1.12 (0.39)	1.29 (0.89)	1.29 (0.89)	1.27 (0.82)	1.26 (0.82)	1.26 (0.80)	1.26 (0.80)
Monthly household income (<500 yuan omitted)								
500–1,000 yuan	1.78+ (1.69)	1.78+ (1.68)	2.01* (2.10)	2.01* (2.12)	1.97* (2.03)	1.97* (2.05)	1.95* (2.01)	1.95* (2.01)
1,000–2,000 yuan	1.85 (1.61)	1.85 (1.61)	1.97+ (1.83)	1.96+ (1.83)	2.01+ (1.97)	2.01+ (1.92)	2.01+ (1.91)	1.99+ (1.89)
2,000–3,500 yuan	3.24** (2.92)	3.24** (2.91)	3.87** (3.56)	3.87** (3.55)	3.79** (3.51)	3.79** (3.52)	3.77** (3.51)	3.77** (3.49)
3,500 yuan and higher	5.32** (3.84)	5.32** (3.84)	5.90** (4.22)	5.99** (4.30)	6.04** (4.27)	6.10** (4.33)	5.99** (4.28)	6.05** (4.34)
Grandparent in household	0.61* (−2.08)	0.61* (−2.08)	0.58* (−2.17)	0.59* (−2.14)	0.57* (−2.21)	0.58* (−2.18)	0.58* (−2.21)	0.59* (−2.17)
Other adult in household	0.88 (−0.44)	0.88 (−0.44)	0.85 (−0.57)	0.83 (−0.62)	0.84 (−0.59)	0.83 (−0.61)	0.83 (−0.63)	0.82 (−0.67)
Community characteristics								
Urban residence	0.92 (−0.27)	0.92 (−0.27)	0.96 (−0.15)	0.96 (−0.16)	0.94 (−0.21)	0.93 (−0.23)	0.94 (−0.20)	0.94 (−0.21)
Child care facilities in community	1.15 (0.49)	1.15 (0.49)	1.12 (0.41)	1.13 (0.44)	1.14 (0.47)	1.14 (0.47)	1.14 (0.46)	1.14 (0.46)
One-child policy in locality (one child only omitted)								
Allowing 2nd child if 1st was a girl	0.48** (−2.57)	0.48** (−2.57)	0.50* (−2.33)	0.50* (−2.33)	0.51* (−2.27)	0.51* (−2.28)	0.50* (−2.30)	0.50* (−2.38)
Allowing 2nd child in all cases	0.27** (−3.84)	0.27** (−3.85)	0.32** (−3.21)	0.33** (−3.18)	0.31** (−3.34)	0.31** (−3.32)	0.31** (−3.36)	0.31** (−3.33)
Local responsibility established	0.74 (−1.05)	0.74 (−1.05)	0.73 (−1.10)	0.72 (−1.16)	0.74 (−1.06)	0.74 (−1.06)	0.74 (−1.04)	0.74 (−1.05)
One-child subsidy provided	1.13 (0.45)	1.13 (0.45)	1.02 (0.08)	1.02 (0.09)	1.03 (0.11)	1.04 (0.13)	1.04 (0.16)	1.06 (0.21)

Table 5 continued

	Having sibling or not		Sibling gender		Younger/older sibling		Preschool-/school-age sibling	
	No Int.	W/Int.	No Int.	W/Int.	No Int.	W/Int.	No Int.	W/Int.
Intercept	0.10* (−2.42)	0.10* (−2.45)	0.12* (−2.06)	0.13* (−2.04)	0.11* (−2.15)	0.11* (−2.17)	0.10* (−2.23)	0.10* (−2.28)
Observations	784	784	719	719	719	719	719	719
Variance Components at Level 3	0.59	0.59	0.57	0.58	0.56	0.56	0.56	0.57
Intra-class correlation (ICC)	0.15	0.15	0.15	0.15	0.15	0.15	0.15	0.15
Akaike's information criterion (AIC)	584.19	585.17	548.49	550.15	551.35	553.11	550.99	552.75

Notes: Odds ratios with z-statistics in parentheses based on Models 5 (without interactions) and 6 (with interactions) in Table 2 from one of the five datasets generated by multiple imputation; ** $p < 0.01$, * $p < 0.05$, + $p < 0.10$

References

Agresti, A. (2002). *Categorical data analysis* (2nd ed.). New York: Wiley.

Akaike, H. (1974). A new look at the statistical model identification. *IEEE Transactions on Automatic Control, 19*, 716–723.

Allison, P. D. (2001). *Missing data*. Thousand Oaks, CA: Sage Publications, Inc.

Andersson, G., Hank, K., Rønsen, M., & Vikat, A. (2006). Gendering family composition: Sex preferences for children and childbearing behavior in the Nordic countries. *Demography, 43*, 255–267.

Attane, I. (2002). China's family planning policy: An overview of its past and future. *Studies in Family Planning, 33*, 103–133.

Barr, D. J. (2008). Analyzing "visual world" eye-tracking data using multilevel logistic regression. *Journal of Memory and Language, 59*, 457–474.

Bernman, J. (1999). China attempts to soften its one-child policy. *The Lancet, 353*, 567.

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

Burgess, R., & Zhuang, J. (2002). Modernization and son preference in People's Republic of China. *Economics and Research Department Working Paper Series No. 20*. Manila, Philippines: Asian Development Bank.

Burnham, K. P., & Anderson, D. R. (2002). *Model selection and multimodel inference: A practical information-theoretical approach* (2nd ed.). New York: Springer Science+Business Media.

Chow, E. N., & Zhao, S. M. (1996). The one-child policy and parent-child relationships: A comparison of one-child with multiple-child families in China. *International Journal of Sociology and Social Policy, 16*, 35–62.

Chung, W., & Das Gupta, M. (2007). Why is son preference declining in South Korea? The role of development and public policy, and the implications for China and India. *Policy Research Working Paper 4373*. Washington D.C.: The World Bank.

Clarke-Stewart, A., & Allhusen, V. D. (2005). *What we know about childcare*. Cambridge, MA: Harvard University Press.

Cohen, J., & Cohen, P. (1983). *Applied multiple regression/correlation analysis for the behavioral sciences*. Hillsdale, NJ: Lawrence Erlbaum Associates, Inc.

Currie, J. (2001). Early childhood intervention programs: What do we know? *Journal of Economic Perspectives, 15*, 213–238.

Das Gupta, M., Jiang, Z., Li, B., Xie, Z., Chung, W., & Bae, H. (2003). Why is son preference so persistent in East and South Asia? A cross-country study of China, India and the Republic of Korea. *The Journal of Development Studies, 40*, 153–187.

Ejrnaes, M., & Portner, C. C. (2004). Birth order and the intra-household allocation of time and education. *The Review of Economics and Statistics, 86*, 1008–1019.

Filmer, D., Friedman, J., & Schady, N. (2008). Development, modernization, and son preference in fertility decisions. *Policy Research Working Paper 4716*. Washington D.C.: The World Bank.

Fong, V. L. (2002). China's one-child policy and the empowerment of urban daughters. *American Anthropologist, 104*, 1098–1109.

Garg, A., & Morduch, J. (1998). Sibling rivalry and the gender gap: Evidence from child health outcomes in Ghana. *Journal of Population Economics, 11*, 471–493.

Hardin, J. W., & Hilbe, J. M. (2002). *Generalized estimating equations*. Boca Raton, Florida: Chapman and Hall/CRC Press.

Hatzitheologou, E. (1997). Reading achievement, birth order and family size. *International Journal of Early Childhood, 29*, 14–21.

Haughton, J., & Haughton, D. (1995). Son preference in Vietnam. *Studies in Family Planning, 26*, 325–337.

Heagerty, P. J., & Zeger, S. L. (2000). Marginalized multilevel models and likelihood inference. *Statistical Science, 15*, 1–26.

Hesketh, T., & Zhu, W. (1997). The one child family policy: The good, the bad, and the ugly. *British Medical Journal, 314*, 1685–1687.

Hill, L. J., Reiter, J. P., & Zanutto, E. L. (2004). A comparison of experimental and observational data analyses. In A. Gelman & X. Meng (Eds.), *Applied Bayesian modeling and causal inference from incomplete-data perspectives: An essential journey with Donald Rubin's statistical family* (pp. 49–60). West Sussex, UK: Wiley.

Hossain, M. M., & Glass, R. I. (1988). Parental son preference in seeking medical care for children less than five years of age in a rural community in Bangladesh. *American Journal of Public Health, 78*, 1349–1350.

Hutzler, C., & Chang, L. T. (2004). China weighs easing its harsh 'one child' rule: family-planning policies have long drawn flak. *Wall Street Journal, Oct 4, 2004:A.1.*

Joesch, J. M., Maher, E. J., & Durfee, A. (2006). Child care arrangements for toddlers and preschoolers: Are they different for youngest children? *Early Childhood Research Quarterly, 21*, 253–266.

Jones, M. P. (1996). Indicator and stratification methods for missing explanatory variables in multiple linear regression. *Journal of the American Statistical Association, 91*, 222–230.

Jowett, A. J. (1991). China: The one, two, three, four and more child policy. *Focus, 41*, 32–36.

Kilburn, M. R., & Datar, A. (2002). The availability of child care centers in China and its impact on child care and maternal work decisions. *Labor and Population Program*, *Working Paper Series 01-12, DRU-2824-NIH*. Santa Monica, CA: RAND.

Lee, L. E. (1992). Day care in the People's Republic of China. In M. E. Lamb, K. J. Sternberg, C. P. Hwang, & A. G. Broberg (Eds.), *Child care in context: Cross-cultural perspectives* (pp. 355–392). Hillsdale, New Jersey: Lawrence Erlbaum Associates.

Lehrer, E. L. (1989). Preschoolers with working mothers: An analysis of the determinants of child care arrangements. *Journal of Population Economics, 4*, 251–268.

Leung, J. C. (2003). Social security reforms in China: Issues and prospects. *International Journal of Social Welfare, 12*, 73–85.

Li, J. (2004). Gender inequality, family planning, and maternal and child care in a rural Chinese county. *Social Science and Medicine, 59*, 695–708.

Li, Y., Shi, A., Wan, Y., Hotta, M., & Ushijima, H. (2001). Child behavior problems: Prevalence and correlates in rural minority areas of China. *Pediatrics International, 43*, 651–661.

Liang, X., Fuller, B., & Singer, J. D. (2000). Ethnic differences in child care selection: The influence of family structure, parental practices, and home language. *Early Childhood Research and Quarterly, 15*, 357–384.

Lin, T. (2009). The decline of son preference and rise of gender indifference in Taiwan since 1990. *Demographic Research, 20*, 377–402.

Little, R. J. A., & Rubin, D. B. (1987). *Statistical analysis with missing data*. New York: Wiley.

Magnuson, K. A., & Waldfogel, J. (2005). Early childhood care and education: Effects on ethnic and racial gaps in school readiness. *The Future of Children, 15*, 169–196.

Meulenberg, C. (2004). Definitely probably one: A generation comes of age under China's one-child policy. *World Watch, 17*, 31–33.

Mishra, V., Roy, T. K., & Retherford, R. D. (2004). Sex differentials in childhood feeding, health care, and nutritional status in India. *Population and Development Review, 30*, 269–295.

National Bureau of Statistics. (2001). *China statistical yearbook 2001*. Beijing: China Statistics Press.

Obermeyer, C. M., & Cardenas, R. (1997). Son preference and differential treatment in Morocco and Tunisia. *Studies in Family Planning, 28*, 235–244.

Organization for Economic Co-Operation and Development (OECD) (2001). *Starting strong: Early childhood education and care*. Paris, France.

Organization for Economic Co-Operation and Development (OECD) (2006). *Starting strong II: Early childhood education and care*. Paris, France.

Pan, W. (2001). Akaike's information criterion in generalized estimating equations. *Biometrics, 57*, 120–125.

Park, C. B., & Cho, N. H. (1995). Consequences of son preference in a low-fertility society: Imbalance of the sex ratio at birth in Korea. *Population and Development Review, 21*, 59–84.

Raudenbush, S. W., & Bryk, A. S. (2002). *Hierarchical linear models: Applications and data analysis methods*. Thousand Oaks, CA: Sage Publications Inc.

Raudenbush, S. W., Bryk, A. S., & Congdon, R. (2008). *HLM6 for windows*. Lincolnwood, IL: Scientific Software International, Inc.

Riley, N. E. (2004). China's population: New trends and challenges. *Population Bulletin, 59*, 1–36.

Royston, P. (2005). Multiple imputation of missing values: Update of ICE. *Stata Journal, 5*, 527–536.

Rubin, D. B. (1987). *Multiple imputation for nonresponse in surveys*. New York: Wiley.

Rubin, D. B. (1996). Multiple imputation after 18+ years. *Journal of the American Statistical Association, 91*, 473–489.

Schafer, J. L. (1997). *Analysis of incomplete multivariate data*. London: Chapman & Hall.

Short, S. E., Zhai, F., Xu, S., & Yang, M. (2001). China's one-child policy and the care of children: An analysis of qualitative and quantitative data. *Social Forces, 79*, 913–943.

Smolensky, E., & Gootman, J. A. (Eds.). (2003). *Working families and growing kids: Caring for children and adolescents*. Washington, D.C.: The National Academies Press.

Snijders, T. A. B., & Bosker, R. J. (1999). *Multilevel analysis: An introduction to basic and advanced multilevel modeling*. London: Sage Publications.

United Nations Educational, Scientific and Cultural Organization (UNESCO) and Organization for Economic Co-operation and Development (2002). *Financing education—investments and returns: Analysis of the world education indicators*. Paris, France.

U.S. Census Bureau. (2008). *Statistical abstract of the United States: 2009* (128th ed.). Washington, DC: U.S. Census Bureau.

van Buuren, S., Boshuizen, H. C., & Knook, D. L. (1999). Multiple imputation of missing blood pressure covariates in survival analysis. *Statistics in Medicine, 18*, 681–694.

Waldfogel, J. (2006). *What children need*. Cambridge, MA: Harvard University Press.

Wang, D., Kato, N., Inaba, Y., Tango, T., Yoshida, Y., Kusaka, Y., et al. (2000). Physical and personality traits of preschool children in Fuzhou, China: Only child vs. sibling. *Child: Care, Health and Development, 26*, 49–60.

Winckler, E. A. (2002). People's Republic of China law on population and birth planning. *Population and Development Review, 28*, 579–585.

Wong, M. N. C., & Pang, L. J. (2002). Early childhood education in China: Issues and development. In L. K. S. Chan & E. J. Mellor (Eds.), *International developments in early childhood services* (pp. 53–70). New York: Peter Lang Publishing.

Zhai, F., & Gao, Q. (2008). Center-based early childhood education and care in China: Policies, trends, and implications. *Asian Social Work and Policy Review, 2*, 127–148.