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# Gender, Sibship Structure, and Educational Inequality in Taiwan: Son Preference Revisited

*This study examines how sibship characteristics affect educational attainment in Taiwan. Using a multilevel analysis of a sibling sample of 12,715 observations from 3,001 families drawn from a national survey, we investigate the effects of family size, sibship density, birth-order rank, and sibship gender composition. The results support the argument that the effect of son preference on intrafamily educational inequality is conditional on family resources. We also find, however, that male firstborns, who are the ultimate inheritors of paternal authority in Chinese families, have additional leverage in the sibling competition for family resources. The privilege for firstborns does not extend to daughters. Therefore, we argue that culturally defined norms regarding seniority and gender help shape intrafamily resource allocation in Chinese society.*

Research on family structure and educational outcomes has linked children's number of siblings, ordinal position of birth, sibship gender composition, and age differences from siblings with the educational resources that they receive (Blake, 1989; Conley, 2000; Downey, 2001; Steelman & Powell, 1989). One emphasis of this literature has been how sibship configuration is associated with intrafamily gender inequality in educational opportunities (Steelman, Powell, Werum, & Carter, 2002), particularly in developing countries (Buchmann, 2000; Parish & Willis, 1993; Sudha, 1997). Despite this interest in intrafamily educational inequality in less industrialized societies, the effects of a full range of sibship characteristics on sons' and daughters' educational opportunities have not been analyzed systematically. In this study, we used a multilevel analysis of sibling data to investigate how sibship size, the gender composition of siblings, and the age spacing between siblings affect educational inequality between sons and daughters in Taiwanese families. The multilevel modeling approach used in this study extends research in international contexts by taking into account unmeasured family traits that may affect children's schooling (Hauser & Mossel, 1985). This analytical approach allows us to rigorously examine educational inequality within families, which most previous research using nonsibling data has failed to do.

We also examine how the different meanings of birth-order rank between sons and daughters in Chinese culture affect sibling competition for educational resources. Like other Confucianism-influenced societies in East Asia, the cultural context of Taiwanese families highlights not only gender differences but also hierarchical family relations (Stacey, 1983). Prior studies taking cultural norms into account rarely have extended beyond the effect of parental gender preferences on children's schooling (Greenhalgh, 1985; Post,

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2001). We call particular attention to Chinese family norms concerning seniority. As the analysis reveals, the different roles prescribed for children according to both their gender and birth-order rank in Chinese families help explain educational inequality. This research contributes to the sibship literature by providing an understanding of how the cultural context shapes family dynamics.

## BACKGROUND

Arguments for the importance of sibship configuration are based on the notion that a child's structural position in the sibship constrains the parents' ability to invest in the child's schooling. Sibship size is thought to be inversely associated with children's educational achievement because parents become less capable of providing attention and funds for each child as family size increases (Blake, 1985; Downey, 1995, 2001; Steelman & Powell, 1989). Later born children are likely to benefit from their parents' improving financial status over the life course and thus have greater educational attainment (Steelman & Powell, 1991; Van Eijck & De Graaf, 1995). In addition, researchers have argued that how close siblings are in age affects educational outcomes because parents have more time and resources for each child when childbirths are spaced more widely (Powell & Steelman, 1993; Steelman & Powell, 1990).

The association between sibship characteristics and education might differ for sons and daughters because parental investment strategies may depend on the child's gender. Our explanation for intrafamily gender disparity in educational opportunities follows the framework that argues that parents under financial constraints allocate educational funds according to their conscious assessments of sons' and daughters' relative market opportunities (Brinton, 1993; Buchmann, 2000; Parish & Willis, 1993). The lower returns to education for women hence lead to daughters' inferior educational opportunities relative to sons' in East Asia (Brinton). This preferential investment strategy implies that the sibship's gender composition will affect intrafamily gender inequality in schooling. Because families with a higher ratio of daughters to sons can extract more resources from daughters for each son, the gender gap in educational attainment will be larger in such families. Nevertheless, daughters' education will suffer less in richer families because the preferential treatment of sons is conditional on family budgets. Similarly, having fewer children or spacing childbirths more widely may reduce educational inequality by gender within families because it increases the resources available for each child. This framework also predicts that later born children of both genders will achieve higher levels of education because their family budgets are less tight by the time they grow up. In addition, as women's employment opportunities improve in society at large, parental perceptions of the utility of schooling for daughters are expected to change and thus decrease intrafamily gender inequality in educational attainment.

Parents may be willing to provide resources for both sons' and daughters' education when circumstances permit, but their decisions may not be independent of cultural values. Our study expands upon research emphasizing parents' conditional preference for educating sons by bringing in the cultural context (Buchmann, 2000; Parish & Willis, 1993). The norms prescribed for Chinese children are based on both their seniority and gender (Hwang, 1991). Specifically, the ultimate authority within the Chinese family is supposed to be passed from the father to the eldest son (Lin, 1988; Stacey, 1983). The traditional expression "the eldest brother is like a father [to his siblings]" reflects this transfer of authority. The eldest son in a family is expected to lead the patrilineal family, which will later include his married brothers and their families. When the eldest son happens to be the family's firstborn, his seniority in the sibship further legitimizes his entitlement to paternal authority. As the extended family's future leader and decision maker, a firstborn son's educational achievement is assumed to be relevant to the entire family's long-term welfare. For this reason, a firstborn son may be given a greater share of the parental resources and thus may achieve higher educational attainment. Conversely, a firstborn daughter's seniority generally leads to the expectation that she will shoulder family responsibilities like her mother. The eldest sister is especially like a mother to her younger siblings when there are substantial age differences between them. Thus, family burdens can be expected to obstruct firstborn daughters' educational opportunities more when parents space children more widely in age.

Similar expectations for firstborn sons and daughters may exist in other cultures. In the United States, however, research has found almost no difference between boys and girls in the effects of birth order on educational outcomes (Blake, 1989; Zajonc, 2001). Although Van Eijck and De Graaf (1995) found some benefit of being firstborn on schooling among Hungarian families, they attributed this effect to firstborns having fewer competing siblings during the early years and suggested that the results are the same for both genders. The prominence of seniority-based family norms in Confucianism-influenced societies could lead to different results from those of previous studies conducted in other cultural contexts. If traditional Chinese family norms play a part in parental resource distribution, being firstborn will benefit the son's education but hinder the daughter's in Taiwan. The firstborn son's privilege will be more pronounced in larger families, where the eldest son's future role as a family leader bears more importance. Moreover, if the culture legitimizes the extra leverage for firstborn sons in competing for family resources, higher family economic status will not reduce their advantage. For the same reason, firstborn sons' relative advantage should change rather slowly over time.

Here we conceptualize the firstborn son's privilege as a gain *after* taking into consideration a general disadvantage for earlier born children. On the whole, earlier born children may have lower educational levels because their parents face greater financial constraints when such children are school aged. Parish and Willis (1993) even showed this association for both genders in Taiwan. We acknowledge that a firstborn son may not fare better than his siblings by an absolute measure. If firstborn sons are privileged, however, then we should find that the firstborn status helps boys' education after controlling for the family's increasing budgets (proxied by birth order in a linear measure). Parish and Willis's study did not explore this possibility.

To briefly introduce the context for the present study, Taiwan began industrializing just before World War II and experienced rapid economic development in the postwar decades. In 1955, 61% of Taiwan's labor force was in agricultural employment. This declined to 20% by 1980 and to near 10% in the 1990s. An increase in industrial employment not only raised the returns to education for both men and women but also affected the government's educational policies. Education through primary school became mandatory for children born after 1945. The children born in 1956 and later further benefited from the extension of mandatory education to 9 years.

Prior studies on educational attainment in Taiwan suggest the importance of family background characteristics, including socioeconomic status and the father's and mother's levels of schooling (Parish & Willis, 1993; Tsai, Gates, & Chiu, 1994). Children from single-parent families may have lower educational levels because their families tend to have fewer resources (McLanahan & Sandefur, 1994). Whether the mother is a homemaker also may be a critical factor because an unemployed mother can spend more time on her children's education. Two other potential predictors for Taiwanese children's schooling are ethnic background and rural residence. Research has documented that Mainlanders, immigrants to Taiwan after the Kuomintang regime lost China's civil war in 1949, had a significant advantage in educational attainment over other ethnic groups, such as Fukien and Hakka (Tsai et al., 1994). Such ethnic differences may have persisted to shape educational outcomes. Educational opportunities also may have been worse in rural areas, especially during the early period of Taiwan's development, and its lingering effects may still be experienced today. In addition, because of Taiwan's rapid development and social change, family norms and values may have varied over time. For example, parental perceptions of their sons' and daughters' proper educational levels should have risen with the extensions of mandatory education. The government's highly effective family planning programs that began in the late 1960s may also have encouraged more parents to trade off the quantity for the "quality" of children (Chang, Freedman, & Sun, 1987), which implies different educational investment strategies. Controlling for cohort differences among families is therefore essential because families formed in different time periods may have developed different expectations and values that affect intrafamily resource allocation.

## METHOD

Data for the present study came from the first wave of the Panel Study of Family Dynamics in Taiwan conducted by the Institute of Economics at Academia Sinica. The Panel Study of Family Dynamics collected longitudinal data from a nationally representative sample of 4,105 men and women born 1935 – 1976. The first wave was carried out in three time periods during 1999 – 2003 because of budgetary constraints.

The survey asked respondents to provide detailed information on their family of origin, including the age, gender, ordinal position, education, and occupation of up to five live siblings. These reports on respondents' siblings provided us with quasi-sibling data for examining educational attainment. Sibling samples have been argued to be particularly useful for studying educational mobility because they allow researchers to control for unobserved interfamily differences that may affect schooling (Hauser & Mossel, 1985). Although our data were not collected from multiple children of one family, we have information regarding parental characteristics, age, gender, education, sibship size, birth order, and sibship gender composition for up to six children in each informant's family of origin. We took advantage of these quasi-sibling data by treating all the respondents and their reported siblings as individual observations in the statistical analysis.

To construct a sibling sample for the analysis, we excluded the respondents who supplied insufficient information on their siblings. Among the informants for the Panel Study of Family Dynamics, 515 (12.6%) did not answer the year of birth for any of their siblings and 279 of them (6.8%) did not provide the age of the family's firstborn. We omitted these respondents and their siblings because such missing data made it impossible for us to estimate the period during which the family was formed and to examine cohort differences. We also omitted 116 respondents (2.8%) with inconsistent reports on siblings' birth orders and ages. Furthermore, 105 respondents (2.6%) were eliminated from the sample because they missed educational levels or other key information for all siblings mentioned. The respondents excluded for missing critical data on siblings tended to be older, less educated, and from larger families. Because of this selection pattern, we expect the effects of sibship size and belonging to earlier family cohorts on educational attainment to be attenuated.

In addition, our focus on Chinese family norms led to the exclusion of observations from 89 aboriginal families (2.2%) that might have experienced somewhat different cultural influences. The reports from the remaining respondents were used to construct the sibling sample for the analysis. We then omitted 196 individuals who were 25 years or younger from the sibling sample to ensure that all observations had completed schooling at the time of the survey. Our final sample consisted of 12,715 observations from 3,001 families. Because the Panel Study of Family Dynamics asked respondents to report only up to five siblings, we do not have the entire set of siblings for 951 (31.7%) families. We should be able to analyze intrafamily educational inequality with the sample of siblings from these families as long as the selection of siblings reported by respondents was not biased with respect to schooling. Our exploratory analysis also indicated that the results were similar after controlling for whether the full set of siblings is included (not shown).

The dependent variable for the statistical analysis is educational attainment, measured by the years of formal schooling reported for each observation in the survey. The independent variables contain both individual- and family-level indicators. Measures of individual characteristics are a person's gender (1 = *female*), ordinal position in the sibship (in a continuous measure), a dummy variable indicating whether one is a firstborn son, and another dummy for a firstborn daughter. We treated the size, gender composition, and density of the sibship as family-level variables (i.e., variables universal for all children in a family). Sibship size was measured as the total number of children in a family. Sibship gender composition was measured by a dummy variable indicating whether the family had more sons than daughters. Because the informants for the Panel Study of Family Dynamics were asked to report the numbers of their brothers and sisters, we were able to correctly measure the gender composition of siblings for all observations. We coded 0 for both families with equal numbers of sons and daughters and families with more daughters than sons because our exploratory analysis suggested similar effects for these two types of sibship gender composition. Sibship density, which is how closely children are spaced in age, was measured by the age difference between the oldest and the youngest children available from the family divided by the difference in their ordinal positions. Although measuring sibship density as one's age differences relative to the neighboring siblings may be more precise, our data did not always include the ages of an observation's adjacent siblings. We argue that the mean age spacing estimated from the children available is nevertheless a sufficient proxy for parents' fertility behavior and will show whether child-spacing strategies affected the children's average schooling. We took the natural logarithm of the average age spacing (+1) to adjust for skewness because the distribution of average age spacing in the sample tilted toward smaller values.

Other family-level variables consist of measures for one's background, including the father's and mother's years of education, family socioeconomic status, and a dummy indicating whether the mother was mostly homemaking when the child was growing up. Family socioeconomic status was measured using the longest lasting occupation the father had during the survey informant's childhood. We used the mother's longest occupation as a replacement when the father was reported to be absent. We converted the three-digit occupational codes into the International Socioeconomic Index of Occupational Status proposed by Ganzeboom and Treiman (1996). We also constructed a proxy for children from single-parent or no-parent families according to survey informants' reports regarding absent fathers or mothers when they were school aged. This dummy variable is referred to as single-parent family structure hereafter because there were few no-parent families in the sample (<1%). To reduce statistical bias, the multiple imputation technique was employed to handle the missing values on the family-level control variables (Acock, 2005). The variable requiring the most imputed values was father's education, which 4.4% of survey informants failed to report.

Other family traits introduced as control variables include ethnic background (Fukien, Hakka, or Mainlander), rural residence, and family cohort. Measuring each observation's childhood residence and its level of urbanization was impossible because the Panel Study of Family Dynamics did not ask respondents to report such information for their siblings. Hence, we used an indicator that the father was a farmer as a proxy for rural residence because children from farming families are likely to have resided in rural areas while they were school aged. Family cohort was measured using the year when the oldest child was born into the family, which should be close to the time when the family was formed. We divided the sample into five family cohorts corresponding to changes in educational and population policies in Taiwan: 1935 and earlier, 1936 – 1945, 1946 – 1955, 1956 – 1965, and 1966 – 1978. Ideally, we also should have included individual birth cohorts in the models because family cohorts may not completely capture changes in individuals' educational opportunities across different periods. An earlier diagnostic analysis, however, indicated that introducing both sets of cohort variables caused a problem of excessive multicollinearity. We chose to include only the family cohorts because we are primarily interested in how changes in parental expectations and family practices over time affect intrafamily educational inequality. The results were very similar regardless of whether we used family cohorts, individual cohorts, or both in the models (analyses not shown).

Table 1 provides descriptive statistics of the individual- and family-level variables for the sample. On average, those included in the analysis had nearly 10 years of schooling, which greatly surpassed the means of their parents' education. This intergenerational difference corroborates that Taiwan's educational opportunities have improved considerably over time. The proportion of firstborn sons in the sample is slightly over one tenth, similar to that of firstborn daughters. The mean value of sibship size is 5.50, with a standard deviation of 1.99, indicating a substantial variation in size among Taiwanese families. This variation reflects Taiwan's rapid fertility declines since the 1960s. Around 40% of the families in the sample had more boys than girls. The average age difference between adjacent siblings among the families sampled was 3.60 years.

The analysis of the effects of sibship characteristics used hierarchical linear models, which controlled for the dependence among the educational outcomes of siblings resulting from their shared family experiences (Raudenbush & Bryk, 2002). Conceptually, what the hierarchical modeling does is similar to estimating the individual-level equation of educational attainment separately for each of the families included in the sample and then treating the coefficients from each family as the dependent variables in the family-level equations. The individual-level model is specified as:

$$Y_{ij} = \beta_{0j} + \beta_{1j} O_{ij} + \beta_{2j} G_{ij} + r_{ij} \tag{1}$$

where $Y_{ij}$ is the educational attainment of individual $i$ in family $j$, $\beta_{0j}$ is the individual-level intercept, $O_{ij}$ indicates an ordinal position in the sibship, and $G_{ij}$ indicates an individual's gender. $r_{ij}$ is the error term, which is assumed to be normally distributed, with a mean of 0 and a variance of $\sigma^2$. The family-level model was used to determine whether overall sibship traits and other family characteristics explain the model

Table 1. *Descriptive Statistics for Individual-Level and Family-Level Variables*

| | *M* | *SD* | Range |
|---|---|---|---|
| Individual level | | | |
| Years of education | 9.95 | 4.55 | 0 – 22 |
| Ordinal position of birth | 2.96 | 1.64 | 1 – 10 |
| Gender[a] | .51 | .50 | 0 – 1 |
| Firstborn son | .11 | .32 | 0 – 1 |
| Firstborn daughter | .12 | .32 | 0 – 1 |
| Family level | | | |
| Father's education | 5.23 | 4.41 | 0 – 22 |
| Mother's education | 3.15 | 3.73 | 0 – 16 |
| Family socioeconomic status | 33.65 | 14.76 | 16 – 88 |
| Mother homemaker | .52 | .50 | 0 – 1 |
| Single-parent family | .08 | .27 | 0 – 1 |
| Farm origin | .37 | .48 | 0 – 1 |
| Ethnicity | | | |
| Fukien | .79 | .41 | 0 – 1 |
| Hakka | .12 | .33 | 0 – 1 |
| Mainlander | .09 | .29 | 0 – 1 |
| Family cohort | | | |
| ≤1935 | .12 | .32 | 0 – 1 |
| 1936 – 1945 | .23 | .42 | 0 – 1 |
| 1946 – 1955 | .29 | .45 | 0 – 1 |
| 1956 – 1965 | .20 | .40 | 0 – 1 |
| 1966 – 1978 | .16 | .36 | 0 – 1 |
| Sibship size | 5.50 | 1.99 | 1 – 14 |
| Male-dominant composition[b] | .39 | .49 | 0 – 1 |
| Mean age spacing logged | 1.28 | .28 | 0 – 2.94 |

*Note:* $N = 12{,}715$ individuals from 3,001 families. Imputed values are excluded from the descriptive statistics reported. Mother homemaker and farm origin coded 1 = *yes*, 0 = *no*.

[a]1 = *female*, 0 = *male*. [b]1 = *family has more sons than daughters*, 0 = *family has no more sons than daughters*.

intercept ($\beta_{0j}$) and the gender gap in schooling ($\beta_{2j}$) among individuals. The family-level model is:

$$\beta_{0j} = \gamma_{00} + \gamma_{01}(\mathrm{SS}_j) + \gamma_{02}(\mathrm{SC}_j) + \gamma_{03}(\mathrm{SD}_j) + \gamma_{04}(\mathrm{FC}_j) + \gamma_{05}(\mathrm{SES}_j) + U_{0j} \quad (2)$$

$$\beta_{1j} = \gamma_{10} + U_{1j} \quad (3)$$

$$\beta_{2j} = \gamma_{20} + \gamma_{21}(\mathrm{SS}_j) + \gamma_{22}(\mathrm{SC}_j) + \gamma_{23}(\mathrm{SD}_j) + \gamma_{24}(\mathrm{FC}_j) + \gamma_{25}(\mathrm{SES}_j) + U_{2j} \quad (4)$$

where $\gamma_{00}$ is the family-level intercept, $\gamma_{01}$ the effect of sibship size on the model intercept ($\beta_{0j}$), $\gamma_{02}$ the effect of sibship gender composition on $\beta_{0j}$, $\gamma_{03}$ the effect of sibship density (i.e., the average age spacing between siblings) on $\beta_{0j}$, $\gamma_{04}$ the effect of family cohort on $\beta_{0j}$, and $\gamma_{05}$ the effect of family socioeconomic status on $\beta_{0j}$. Similarly, $\gamma_{20}$ is the intercept for the predicted gender effect, $\gamma_{21}$ the effect of sibship size on $\beta_{2j}$, $\gamma_{22}$ the effect of sibship gender composition on $\beta_{2j}$, $\gamma_{23}$ the effect of sibship density on $\beta_{2j}$, $\gamma_{24}$ the effect of family cohort on $\beta_{2j}$, and $\gamma_{25}$ the effect of family socioeconomic status on $\beta_{2j}$. Alternatively, $\gamma_{20}$ can be understood as the main effect of gender, and $\gamma_{21}$, $\gamma_{22}$, and $\gamma_{23}$ as the cross-level interaction effects between gender and sibship characteristics. $U_{0j}$, $U_{1j}$, and $U_{2j}$ are error terms, which are assumed to be normally distributed, with a mean of 0 and a variance of $\sigma^2$. Next, we introduce the dummies for being a firstborn son and being a firstborn daughter to the individual-level model to examine whether the firstborn status leads to additional gain or loss. This model is specified as:

$$Y_{ij} = \beta_{0j} + \beta_{1j}O_{ij} + \beta_{2j}G_{ij} + \beta_{3j}M_{ij} + \beta_{4j}F_{ij} + r_{ij} \quad (5)$$

where $M_{ij}$ is a dummy variable for being a firstborn son, $F_{ij}$ is a dummy variable for being a firstborn daughter, and the rest of the parameters are the same as in Equation 1. The family-level model is specified to examine whether contextual factors, such as family socioeconomic status, family cohort, and other sibship characteristics, explain the effects of being a firstborn son ($\beta_{3j}$) and being a firstborn daughter ($\beta_{4j}$) on schooling. The equations predicting $\beta_{3j}$ and $\beta_{4j}$ are similar to Equations 2 and 4.

## RESULTS

Table 2 presents a series of hierarchical linear models with increasing complexity. Model 1 is the baseline model that uses family-level variables to predict the model intercept and includes only birth order at the individual level. This model demonstrates how differences in family traits affect children's average education, after controlling for individuals' birth-order rank. The results for the family-level control variables are consistent with those found in the general literature on educational attainment. Children of those with higher education and better socioeconomic status obtained more years of schooling. Children with a homemaking mother also had higher educational levels. Conversely, children from single-parent families achieved lower educational levels. In addition, the average years of education was greater among children from Mainlander or Hakka families, nonfarm families, and families formed more recently.

The coefficient of sibship size indicates that adding a child to a family reduced everyone's schooling by 0.21 year. This finding is consistent with the argument that the family resources available for each child's education decrease as the sibship size rises. Holding sibship size and density constant, children from families with more sons than daughters had 0.34 year more schooling on average than those from other families. The fact that having more sons raises the family mean is not surprising if women generally have lower education than men. To the extent that a child's schooling is directly related to the parents' educational investment, however, this finding implies that parents are capable of providing more resources when there are more sons in a family. Thus, family resources for children's education are somewhat flexible at the parents' discretion. Model 1 also shows that spacing childbirths more widely did not raise a family's average educational attainment. The positive effect of ordinal position in the sibship reveals that later born children had higher educational attainment, suggesting that a family's improving finances over the life cycle helps later born children's educational opportunities. Because the model did not control for individuals' birth cohort, one may suspect that some of this birth-order effect had to do with cohort differences between earlier and later born children in the same family. In other words, later born children may fare better because the educational opportunities for their birth cohorts have increased from those of their elder siblings' cohorts. A separate analysis nevertheless showed that the effects of birth order remained positive and significant when individuals' birth cohorts were included (not shown).

Model 2 adds gender as an individual-level predictor. The result shows that women had 1.13 years less schooling, holding family traits constant. The effects of family-level variables in Model 2 are similar to those in Model 1, except that sibship gender composition no longer affects the family's average schooling once gender is controlled. Models 3 and 4 examine how family characteristics explain the gender gap in schooling. As noted previously, the effects of family-level variables on the gender coefficient estimated in the individual-level model can be more intuitively understood as the interaction effects between the child's gender and family traits on educational attainment. We therefore label them as interaction terms in Table 2. When examining interactions with gender that affect education in the models, we used only sibship size, gender composition, sibship density, family socioeconomic status, and family cohort because other interactions do not pertain to our research interest. Besides, a separate analysis found that the interactions between gender and other family traits had only nonsignificant effects (not shown).

Although Model 3 shows that the gender gap in educational attainment was greater in larger families and in families spacing children more widely, these interaction effects became nonsignificant after taking into account family budgets and period differences in Model 4. A further exploration indicated that the addition of family cohort is related to the vanishing effects of sibship size and density on intrafamily gender inequality in education. The fact that fertility behavior has varied considerably among family cohorts in Taiwan explains this relationship. More recent families tend to be smaller and to have shorter average age spacing between children. Within each family cohort, however, the relatively small variations in sibship size and density appear not to affect the gender gap in schooling. According to the coefficients for family cohort, not only did the discrepancy between boys' and girls' schooling virtually disappear but girls also had higher educational levels than boys within families established from 1966 onward ($-1.03 + 2.40 > 0$; $p < .05$). In addition, daughters' disadvantage in education became smaller as a family's socioeconomic status increased. Thus, the effect of son preference on gender inequality in education is mediated by family income.

Model 4 also shows that the gender gap in education was narrower among families with male-dominant sibship compositions. This is to say, boys have a smaller advantage relative to girls when they represent the majority gender. To put it differently, girls' educational opportunities are closer to their brothers' in families with more sons than daughters. This finding is consistent with the hypothesis that parents with more daughters than sons can squeeze more resources out of

Table 2. *Multilevel Model Results Predicting Years of Education*

| Predictors | Model 1 *B* | Model 1 *SE* | Model 2 *B* | Model 2 *SE* | Model 3 *B* | Model 3 *SE* | Model 4 *B* | Model 4 *SE* | Model 5 *B* | Model 5 *SE* | Model 6 *B* | Model 6 *SE* | Model 7 *B* | Model 7 *SE* | Model 8 *B* | Model 8 *SE* |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Family level | | | | | | | | | | | | | | | | |
| Intercept | 10.41** | .04 | 10.39** | .04 | 10.43** | .04 | 10.43** | .04 | 10.44** | .04 | 10.44** | .04 | 10.44** | .04 | 10.43** | .04 |
| Father's EDU | .24** | .02 | .24** | .02 | .24** | .02 | .24** | .02 | .24** | .02 | .24** | .02 | .24** | .02 | .24** | .02 |
| Mother's EDU | .14** | .02 | .14** | .02 | .14** | .02 | .14** | .02 | .14** | .02 | .14** | .02 | .14** | .02 | .14** | .02 |
| Family SES | .03** | .00 | .03** | .00 | .03** | .00 | .03** | .00 | .03** | .00 | .03** | .00 | .03** | .00 | .03** | .00 |
| Mother homemaker | .34** | .09 | .34** | .09 | .34** | .09 | .34** | .09 | .34** | .09 | .34** | .09 | .34** | .09 | .34** | .09 |
| Farm origin | −.76** | .13 | −.73** | .13 | −.74** | .13 | −.72** | .13 | −.73** | .13 | −.72** | .13 | −.73** | .13 | −.73** | .12 |
| Ethnicity (reference Fukien) | | | | | | | | | | | | | | | | |
| Hakka | .50** | .13 | .49** | .13 | .50** | .13 | .49** | .13 | .49** | .13 | .50** | .13 | .49** | .13 | .49** | .13 |
| Mainlander | .50** | .15 | .49** | .14 | .49** | .14 | .50** | .14 | .50** | .14 | .51** | .14 | .50** | .14 | .50** | .14 |
| Family cohort (reference ≤1935) | | | | | | | | | | | | | | | | |
| 1936 – 1945 | 1.19** | .20 | 1.19** | .19 | 1.21** | .19 | 1.18** | .19 | 1.19** | .19 | 1.18** | .19 | 1.19** | .19 | 1.19** | .19 |
| 1946 – 1955 | 2.97** | .20 | 3.03** | .19 | 3.03** | .19 | 2.98** | .19 | 3.01** | .19 | 3.01** | .19 | 3.02** | .19 | 3.05** | .19 |
| 1956 – 1965 | 3.89** | .21 | 3.97** | .21 | 3.98** | .21 | 3.92** | .21 | 3.96** | .21 | 3.96** | .21 | 3.96** | .21 | 4.00** | .21 |
| 1966 – 1978 | 4.52** | .23 | 4.63** | .23 | 4.64** | .23 | 4.60** | .23 | 4.64** | .23 | 4.64** | .23 | 4.63** | .23 | 4.68** | .23 |
| Sibship size | −.21** | .03 | −.18 | .03 | −.18** | .03 | −.19** | .03 | −.20** | .03 | −.19** | .03 | −.20** | .03 | −.19** | .03 |
| Male-dominant composition | .34** | .09 | −.06 | .09 | .01 | .09 | .00 | .09 | .00 | .09 | .00 | .09 | .00 | .09 | .00 | .09 |
| Mean age spacing logged | .10 | .16 | .10 | .16 | .07 | .16 | .12 | .16 | .15 | .16 | .14 | .16 | .16 | .17 | .28 | .17 |
| Single-parent family | −.71** | .20 | −.68** | .20 | −.67** | .20 | −.67** | .20 | −.69** | .20 | −.66** | .20 | −.68** | .20 | −.68** | .20 |
| Individual level | | | | | | | | | | | | | | | | |
| Ordinal position of birth | .48** | .02 | .46** | .02 | .46** | .02 | .47** | .02 | .53** | .02 | .54** | .02 | .54** | .02 | .53** | .02 |
| Gender, female | | | −1.13** | .06 | −1.02** | .06 | −1.03** | .05 | −.87** | .06 | −.87** | .06 | −.89** | .06 | −.91** | .06 |
| Female × Sibship Size | | | | | −.22** | .03 | .03 | .04 | .01 | .04 | .01 | .04 | .05 | .04 | .04 | .04 |
| Female × Male-Dominant Composition | | | | | .19 | .11 | .29** | .11 | .30** | .11 | .30** | .11 | .37** | .12 | .29* | .13 |
| Female × Mean Age Spacing Logged | | | | | −1.04** | .20 | .07 | .21 | .04 | .21 | .03 | .21 | −.09 | .24 | .42 | .26 |
| Female × Family SES | | | | | | | .01** | .00 | .01** | .00 | .01** | .00 | .02** | .00 | .01** | .00 |
| Female × Family Cohort (reference ≤1935) | | | | | | | | | | | | | | | | |
| Female × 1936 – 1945 | | | | | | | .21 | .23 | .20 | .23 | .20 | .23 | .13 | .25 | .31 | .25 |
| Female × 1946 – 1955 | | | | | | | .99** | .22 | .97** | .22 | .96** | .22 | 1.00** | .24 | 1.05** | .25 |

Table 2. *Continued*

| Predictors | Model 1 B | Model 1 SE | Model 2 B | Model 2 SE | Model 3 B | Model 3 SE | Model 4 B | Model 4 SE | Model 5 B | Model 5 SE | Model 6 B | Model 6 SE | Model 7 B | Model 7 SE | Model 8 B | Model 8 SE |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Female × 1956 – 1965 | | | | | | | 1.66** | .24 | 1.64** | .24 | 1.63** | .24 | 1.69** | .26 | 1.73** | .27 |
| Female × 1966 – 1978 | | | | | | | 2.40** | .26 | 2.41** | .26 | 2.41** | .26 | 2.58** | .29 | 2.59** | .30 |
| FS[a] | | | | | | | | | .65** | .09 | .66** | .09 | .60** | .10 | .60** | .10 |
| FS × Sibship Size | | | | | | | | | | | | | .13* | .05 | .12* | .05 |
| FS × Male-Dominant Composition | | | | | | | | | | | | | .31 | .17 | .29 | .17 |
| FS × Mean Age Spacing Logged | | | | | | | | | | | | | −.51 | .29 | −.43 | .29 |
| FS × Family SES | | | | | | | | | | | | | .00 | .01 | .00 | .01 |
| FS × Family Cohort (reference ≤1935) | | | | | | | | | | | | | | | | |
| FS × 1936 – 1945 | | | | | | | | | | | | | −.34 | .33 | −.31 | .33 |
| FS × 1946 – 1955 | | | | | | | | | | | | | .16 | .33 | .15 | .33 |
| FS × 1956 – 1965 | | | | | | | | | | | | | .23 | .35 | .21 | .35 |
| FS × 1966 – 1978 | | | | | | | | | | | | | .60 | .38 | .56 | .38 |
| FD[b] | | | | | | | | | | | .02 | .09 | .03 | .09 | .10 | .10 |
| FD × Sibship Size | | | | | | | | | | | | | | | .06 | .05 |
| FD × Male-Dominant Composition | | | | | | | | | | | | | | | .37* | .18 |
| FD × Mean Age Spacing Logged | | | | | | | | | | | | | | | −1.42** | .32 |
| FD × Family SES | | | | | | | | | | | | | | | .01 | .01 |
| FD × Family Cohort (reference ≤1935) | | | | | | | | | | | | | | | | |
| FD × 1936 – 1945 | | | | | | | | | | | | | | | −.84** | .29 |
| FD × 1946 – 1955 | | | | | | | | | | | | | | | −.06 | .27 |
| FD × 1956 – 1965 | | | | | | | | | | | | | | | −.02 | .31 |
| FD × 1966 – 1978 | | | | | | | | | | | | | | | .04 | .32 |
| Estimated parameters | 21 | | 25 | | 28 | | 33 | | 38 | | 39 | | 43 | | 51 | |
| Deviance | 6,3559.0 | | 6,2822.9 | | 6,2729.9 | | 6,2575.1 | | 6,2510.7 | | 6,2514.4 | | 6,2492.4 | | 6,2426.1 | |

*Note:* $N = 12{,}715$ individuals from 3,001 families. FS = firstborn son; FD = firstborn daughters; EDU = years of education; SES = socioeconomic status. All variables are centered at their grand means; models use the full maximum likelihood estimate.

[a]Specified as fixed in Models 6 – 8. [b]Specified as fixed in Models 7 and 8.

$^{*}p < .05$. $^{**}p < .01$.

the daughters and contribute more funds to their sons' education.

Model 5 introduces the dummy for a firstborn son to test the hypotheses concerning the influence of cultural norms on firstborn children. A firstborn son had a 0.65-year net gain of education, after controlling for gender, a linear term of ordinal position, and family traits. In other words, although being born earlier typically leads to worse educational prospects among Taiwanese children, the firstborn status offsets some of the negative effect of growing up at an earlier stage of the family life cycle for sons. This finding is consistent with our expectation that the firstborn son's privilege prescribed by Chinese culture protects his share of educational resources. Model 5 includes a random error term in the family-level equation predicting the effect of being a firstborn son, but diagnostic statistics indicate that this error term is not significantly different from 0 ($p < .21$). Therefore, in the subsequent models, we specify the firstborn-son coefficient as fixed and remove the random error term from the family-level equation predicting the firstborn-son effect (Raudenbush & Bryk, 2002).

Model 6 shows a nonsignificant effect of being a firstborn daughter. This finding suggests that other than the general disadvantage for earlier born children, there is no additional gain or loss attached to the firstborn status for daughters, unlike for sons. Despite this result, the positive effect of birth-order rank and the negative effect of being a girl, taken together, do predict that female firstborns will be the least educated of all children. The results from Model 6 also suggest that we should specify the firstborn-daughter coefficient as fixed rather than as random ($p < .34$; $\chi^2$ test for the significance of the variance component). Thus, we alter the specification for the next two models.

Models 7 and 8 examine how the effects of being firstborn for sons and daughters vary with family-level factors, including sibship traits, family socioeconomic status, and family cohort. Other cross-level interactions with family attributes are not included because they are neither theoretically relevant nor statistically significant in shaping educational attainment. Interestingly, socioeconomic status fails to explain the positive effect of being a firstborn son. The firstborn son's privilege also did not decrease significantly within more recent families. Thus, a firstborn son can secure a relatively large share of educational resources, regardless of family wealth. Furthermore, the firstborn-son status led to more gain when the sibship size was larger, supporting our argument that the eldest son's leadership means more in larger families.

Regarding the interactions between family characteristics and being a firstborn daughter on education, the results for sibship gender composition and density are notable. A firstborn daughter's relative educational level was higher when she was from a family with more sons, as compared with one with a greater or equal number of daughters. This finding suggests that being firstborn may help girls when there is no "surplus" of daughters in a family. Moreover, firstborn daughters from families where children were spaced more widely had a greater relative disadvantage in educational attainment, holding sibship size constant. This finding is consistent with the argument that larger age distances from younger siblings strengthen firstborn daughters' mother-like role and thus hamper their own achievement.

## CONCLUSIONS

Results from the analysis are largely consistent with the framework that attributes intrafamily gender inequality in education to both family resource constraints and parental perceptions of the utilities of sons' and daughters' schooling. In addition, we find that the particularly privileged status of a firstborn son offsets some of the general disadvantage associated with having an earlier birth-order rank in educational attainment. Conversely, a firstborn daughter does not have any additional leverage. Rather, her schooling is likely to suffer more when her family context strengthens the expectation that she will care for the younger ones like the mother. Thus, this study contributes to the family literature by showing that parental investment strategies are conditioned not only by family budgets but also by culturally defined family norms. Another contribution of this study is to show that the ordinal birth position's implications for educational opportunities differ for male and female children. The extant literature on sibship configuration has tended to consider the effects of the child's gender and ordinal position on educational outcomes independently (Steelman et al., 2002). This research calls attention to how gender mediates the effects of sibship characteristics on educational success.

The findings regarding intrafamily gender inequality in educational opportunities from Taiwan also support Knodel and Jones's (1996) claim that policies aimed at reducing the gender gap in schooling are not always the most effective way to improve girls' education in developing countries. In Taiwan, intrafamily gender inequality in schooling disappeared with time as families changed their fertility behavior and their perceptions of the utility of women's education with economic development. This trend has also been observed in other developing countries (e.g., Knodel, 1997). Moreover, because the intrafamily gender gap in educational attainment is mediated by family socioeconomic status, improving poor children's access to education naturally would narrow the gender gap.

Also worth highlighting are two findings that present an interesting contrast with those from U.S.-based research. First, unlike in the United States (Steelman & Powell, 1989, 1990), the hypothesis regarding the harm of sibship density receives little support in the case of Taiwan. In particular, we find that a wider age spacing among siblings tends to hinder firstborns' educational achievement, whereas some U.S. research has suggested that firstborns benefit when their competing siblings arrive more slowly (Zajonc, 2001). A possible explanation for this difference is that, in developing countries, firstborns are more likely to share the parents' responsibilities for caring for or supporting the younger children. Second, our analysis indicates that girls benefit from being the minority gender in the sibship, whereas Conley (2000) found the gender majority to gain in the United States. Perhaps when parents have restricted resources and son preferences, as in many less industrialized societies, a girl's best hope for fairer treatment is to belong to the family's rarer and thus more precious gender group. This study's different findings underscore the importance of taking the social context into account.

One limitation of this study is that we cannot completely rule out that parental predisposition may shape the sibship structure and children's educational outcomes simultaneously. For instance, parents who value children's education highly may decide to have fewer children in the first place, which may explain the inverse association between family size and schooling. For most of the period examined in this study, however, Taiwanese parents were not equipped to select their first child's gender. Thus, it is difficult to explain the effects we found for being firstborn sons and daughters with an argument of parental predisposition. Further, even though parents with a male firstborn may have fewer subsequent children than those with a female firstborn as a result of son preferences, the smaller family size would only weaken the comparative advantage of male firstborns because they benefit more in larger families.

A second limitation is the use of sibling data collected from only one informant per family, which necessitated our multilevel analysis. A potential problem of reliability naturally arises from this method of data collection. In addition, we faced a disproportionately large number of missing observations from earlier and larger families, whose informants were older and often less able to recall detailed sibling information. Our main finding regarding the firstborn son's privilege, however, is more pronounced in larger families. Thus, our case is likely to be even stronger without such sample bias. Nevertheless, future research on family contexts and individuals' life chances would be enhanced by data collected from matched sibling samples.

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