Transmission of Sex Preferences Across Generations: The Allocation of Educational Resources Among Siblings

C. Y. Cyrus Chu
Distinguished Research Fellow
Institute of Economics, Academia Sinica
cyruschu@sinica.edu.tw

Ruey S. Tsay
H. G. B. Alexander Professor of Econometrics and Statistics
Graduate School of Business, University of Chicago
ruey.tsay@gsb.uchicago.edu

Huoying Wu
Associate Research Fellow
Institute of Economics, Academia Sinica
hwu@econ.sinica.edu.tw

November 25, 2002

Correspondence: Ruey S. Tsay.

Abstract

The purpose of this paper is to test whether there is an intergenerational transmission of gender preferences in educational resource allocation among children. The unique data set of Taiwan's Panel Study of Family Dynamics project provides us a rich 3-generation education information and allows us to probe into this question. We performed our analysis along two directions: the first is to see whether the society as whole has any macro change in gender-specific education achievement, and the second is to see whether there is any within-lineage transmission of gender preferences across generations. After carefully reviewing the education system and societal characteristics in Taiwan, we set up an empirical model to estimate and test the hypotheses of intergenerational transmission of gender preferences. As far as the macro pattern is concerned, we found that although there is a clear tendency of differential treatment against females in the old generation, this tendency is significantly weakened and nearly vanishes in the young generation. Furthermore, the supporting effect of senior siblings in the old generation becomes a crowding (resource-dilution) effect in the young generation. However, within each micro lineage, there is a mild "habitus" effect in gender-specific educational resource allocation in the sense that parents who had the experience of gender-specific differential treatment tend to treat their children in a similar fashion. Moreover, this habitus effect is stronger for female respondents (who were the deprived group) than for male respondents (who were the privileged group).

1 Introduction

Ever since the seminal work of Becker (1964), social scientists have long noticed the importance of education in individual earnings and career development. Because the quantity and quality of children's education is very much influenced by the attitude and devotion of their parents, much of the research on this topic has focused on various family-related aspects of children's education. The purpose of this paper is to use a unique data set to probe into, to the best of our knowledge, a new dimension of the research: how the pattern of educational resource allocation among children has changed across generations. Our study contains macro as well as micro aspects of the problem. As to the macro aspect, we are interested in knowing if a particular pattern of within-family discrimination (e.g., against girls) has weakened or disappeared in the young generation. As to the micro aspect, we investigate whether a parent's attitude (e.g., sex preferences) toward his or her children has any lasting impact when the children later on form their own families and have their own children. We begin with a brief review of the related literature.

For the purpose of comparison, we classify, perhaps idiosyncratically, the related literature into three strands. The first strand compares the possibly different achievements, either in attained education or in earnings, among children of different sexes or birth orders. Related literature includes Sewell and Hauser (1977), Greenhalgh (1985), Behrman and Taubman (1986), Kessler (1991), Birdsall (1991), Parish and Willis (1993), Butcher and Case (1994), Hauser and Kuo (1998), and Ermisch and Francesconi (2001). The general finding is that, due to family budget constraints and/or differential preferences, girls or children in the middle birth order tend to receive less educational resources from their parents.

The second strand of literature is concerned with the comparative study

of siblings or twins. The work here seems to put more emphasis on using such sibling or twin data to control the family background and to identify the influence of other variables on individual achievement, rather than to compare siblings' achievement differences *per se*. Important contributions and surveys along this line can be found, just to name a few, in Sewell and Hauser (1977), Griliches (1979), Behrman et al. (1994), and Card (1999).

The third strand of literature concerns the intergenerational mobility of earnings or education. The focus along this line is often on the role of education and family background on a person's mobility parameters. Related studies with respect to different societal scenarios can be found in Bowles (1972), Behrman and Taubman (1985), Lillard and Willis (1994), Dearden et al. (1997), and Solon (1999).

The focus of this paper is related to the intergenerational transmission of sex-based preferential treatment, but it is from an new angle that is different from all of the previous studies. Our main question can be addressed as follows. Suppose we have a set of sibling data of generation t, and are able to identify the effect of birth order, sibling size, and in particular child gender on siblings' education achievement. Suppose further that children of particular characteristics in some families were preferentially or poorly treated in education investment in generation t. When children of the generation-t grew up and had their own children, i.e. generation-$(t+1)$, we would like to know how the pattern of unequal resource allocation has changed in terms of the allocation of educational resources among children of generation-$(t+1)$.

Intuitively, there are two different factors that may change the pattern of resource allocation among children of generation-$(t+1)$. On the one hand, because differential treatment among children are usually due to resource constraints or traditional conceptions of parents, the pattern of differential treatment should be lessened when such constraints or conceptions have relaxed along with economic development. On the other hand, as predicted

by psychological theories, a child being differentially treated in childhood may form a stereotype idea about sex preferences, and continue this attitude toward his or her own children.[1] Our goal then is to study whether the sex-based preferential treatment in education of an earlier generation may transmit, either in macro or micro aspect, to the next generation.

In order to study empirically the problem posed above, a comprehensive data set is necessary. In particular, we need at least *two generations* of *sibling data* in order to identify the possible sex-based differential treatment within each generation; and only with such a good data set are we able to test whether the experience by siblings of generation t may carry over to those of generation $t+1$. Furthermore, if we are to control the parental education background of generation $t-1$ so as to improve the estimation efficiency of the behavioral relationship in generation t, the data requirement is even more restrictive: we need the education data of three generations to accomplish the estimation and test. Lack of comprehensive data sets described above is perhaps a major reason why there has been no study focusing on such an interesting problem in the literature. Fortunately, a recent survey conducted in Taiwan provides us with such a data set and, hence, allows us to perform this study. Details of the data set is given in Section 2.

The remainder of this paper is arranged as follows. In Section 2, we present the panel study of family dynamics (PSFD) data and outline the background characteristics of the education system in Taiwan. The econometric model together with the empirical analysis are presented in Sections 3 and 4, respectively. The final section concludes.

[1]Concerning the correspondence conceptions between parents and children, see for instance the discussion in Goodnow (1992).

2 The Data Set and the Social Background

The data set we use is from the PSFD survey recently conducted in Taiwan.[2] Since Taiwan is a well-known area undergoing rapid economic and demographic transitions,[3] parents' attitude toward gender differences is expected to change significantly in the past four decades. Against this changing background, it is more likely to observe coexisting samples with contrasting socioeconomic characteristics, which are particularly appropriate for the study of intergenerational changes in family behavior.

The PSFD data set starts with roughly 3,000 respondents of a random survey from cohorts born between 1934 – 1964. The interviewed questions cover detailed socio-economic information about family members of the sampled individual as well as their relations with each other. In particular, for each randomly sampled respondent, information concerning the educational background of almost all of his or her siblings was asked.[4] Moreover, related information of the main respondent's parents, such as their ethnicity and school years, was also included in the interview. As such, we have enough information to estimate the differential education achievement of siblings for the interviewee's generation, call them generation o (old).

[2]The data used is the third-year output of an attempt to develop a unique panel data set in a Chinese society. The project, entitled PSFD, was conducted with the support of the Chinh Ching-Kou Foundation and the National Science Council of Taiwan, and under the auspices of Gary Becker, Angus Deaton, Robert Hauser, James Heckman, Cheng Hsiao, Ronald Lee, William Parish, George Tiao, Jim Vaupel, Arthur Wolf, Cyrus Chu, and other local collaborators. The data set is free of charge for all academic uses. For details, see http://psfd.sinica.edu.tw.

[3]See Chu and Lee (2000) for more details.

[4]Up to 6 siblings' education information was asked. This is about to exhaust all possibilities; in our two-generation sample points, 15.7% of them have sibling sizes larger than or equal to 7, in which most (15.5%) belong to the old generation. When the sibling size is larger than 7, the education information is restricted to the eldest 6.

To enlarge the information of the family, PSFD also interviewed one (randomly chosen) sibling of the main respondent in year 2000, using the same questionnaire as the main respondent's. For these two interviewed siblings, the education information of their children is also asked. This allows us to study the differential education achievement of the generation of interviewees' children, referred to as generation y (young). The sampling structure is drawn in Figure 1.

Since part of our purposes is to analyze the causes of differential education achievement among siblings, we feel obliged to introduce the education system in Taiwan. There are five main tiers of regular schools in Taiwan, namely elementary (6 years), junior high (3 years), high school (3 years), college (4 years) and graduate schools, together with some supplementary vocational schools. Although various schools used to screen their own students, starting from 1950 most schools in Taiwan have participated in the *joint entrance examinations* (JEE) to admit students. Before 1968, for the entrance from elementary to junior high, from junior high to high school, or from high school to college, a student needs to go through a respective JEE. The high school to college JEE is nationwide, whereas the others are held in separate districts, within which there are thousands of students joining the competition. After 1968, the mandatory education extends from six to nine years, and hence the JEE from elementary to junior high was abolished. As one can see, nearly all the sampled respondents are subject to the JEE system.

In Taiwan, because (i) the training of teachers of all tiers of schools except the colleges were monopolized by national Normal colleges, (ii) the salary scales of teachers and professors are seniority-based, and (iii) the University professor licensure is uniformly regulated by the Ministry of Education in most relevant periods of our study,[5] there are no *a priori* reasons to expect

[5]The monopoly of training teachers was finally changed in 1997, and the uniform pro-

quality differences among school teachers. Moreover, the tuition upper bound of private schools regulated by the government also renders the quality improvement of private schools impossible. Thus, most parents and students prefer to go to the less-expensive public schools and universities rather than the private ones. A JEE ranks all participating students according to their test scores, and higher-score students are allowed to choose schools to enter before lower-score students do. Eventually, there are always some disappointed students who do not have any desirable match.[6]

The JEE in Taiwan is basically a written exam, and therefore the criterion of screening students is very uniform. Given the above-mentioned rigid JEE system, whether a student can enter a higher tier school or college depends on his or her ability as well as the resources devoted by his or her parents (e.g., to after school tutoring). The resource devotion from parents to their children of course depends on the parents' education background, ethnicity, budget constraint, and in particular their sex preferences. For instance, if the parents have finance constraints and are only able to afford one child to go to college, then their preferences with respect to child gender or birth order may be important. In sum, the uniform JEE system in Taiwan makes a student's upward moving ladders relatively standard, and hence is convenient for our econometric analysis.

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fessor licensure system was decentralized in 1991; but these recent changes could not have affected the previous decisions of the respondents. For related discussion of controlling school quality, see Behrman and Birdsall (1983).

[6]For instance, in year 2000, 125,498 students registered the JEE of college entrance. The overall entrance rate from high school to college was 59.98%. The most-preferred college in general is the Natinal Taiwan University, which only admitted 3,244 students in year 2000. Students whose scores lower than the rank criterion of various departments of National Taiwan University would have to choose other universities to study. In the same year, there were 22,115 students participating the JEE from junior high to high schools in the Taipei area; corresponding figures in other areas are omitted.

Some descriptive statistics of the sample are presented in Table 1. As one can see from the table, although most statistics are roughly of the same size for males and females, there are several interesting exceptions. First, the average year of schooling for men is longer than that of women, revealing a possible pattern of sex discrimination. This is particularly so for the old generation where the difference is about 2 years. The difference shrinks markedly to 0.3 year in the young generation. The same phenomenon is also shown in the difference of education years between the father and mother of the respondent. Second, the average of schooling years increases substantially across generations for both men and women. It is then not clear whether the reduction in difference between the gender-specific schooling years is due to the improved economic resources of the parents or a change in their attitude.

We also notice from Table 1 that the sibling size reduces significantly across generations, revealing the pattern of Taiwan's demographic transition. Finally, it is very interesting to observe that, although the probabilities of being the first-born are roughly the same for male and female children, the probabilities of being the last-born are substantially larger for males. Indeed, if there is a general preference for sons that induces many parents to have "at least one son," their optimal stopping rule of fertility would indeed wind up with a large macro proportion of boys being the last-born.

[Insert Table 1 about here.]

3 Changing Pattern of Sex Preferences: A Macro Analysis

There are several approaches to estimating and testing the intergenerational transmission of sex preferences. The first approach is described as the following. Let the subscript tij refer to the jth child of family (actually lineage) i

in generation t, and let Y be the schooling years of the child, α the family fixed or random effect parameter, S the sex indicator of the child in question, Z_{ti} a vector of lineage-specific variables, X a vector of other explanatory variables, and ϵ the error term satisfying all regular assumptions, especially being independent of S. The first econometric model of the paper is

$$Y_{tij} = \alpha_{ti} + \gamma_0 S_{tij} + \gamma_1 (I_t \times S_{tij}) + \beta_t X_{tij} + \eta Z_{ti} + \epsilon_{tij}, \quad t = y, o \qquad (1)$$

where I_t is a generational dummy variable with $I_t = 1$ if $t = y$ and $I_t = 0$ if $t = o$. In Equation (1), we use various interaction terms to allow possibly different influences of variables across generations. In particular, γ_0 captures the possible existence of parents' gender preferences, and γ_1 characterizes the weakening or strengthening of this gender effect for the young generation relative to the old generation.

In Equation (1), because of the existence of a common family effect, the errors ϵ_{tij} are not independent for data from the same lineage. Any least squares estimation failing to take into account this dependence will result in inefficient estimation. Furthermore, as pointed out by Griliches (1979), applying the fixed effect model to the family context may exacerbate other econometric problems such as measurement errors and variable endogeneity and may interfere with the estimation of common-to-all-sibling variables. Care must be exercised. To overcome these potential problems in estimation, we follow Parish and Willis (1993) and adopt the approach of least square estimation with Huber (1967) adjustment.[7]

To test whether there is any across-generation change in discrimination against a female child, we test the hypothesis $\gamma_1 = 0$. Following the common practice in the literature, the explanatory variable X should include the (sex- and seniority-specific) sibling size, the ethnicity background, the education

[7]The random effect model requires the assumption that family specific effects be uncorrelated with other explanatory variables, which is not satisfied in our context either.

levels of the parents, and other relevant variables. The exogenous variables adopted in our regression are by and large compatible with those in Parish and Willis (1993), Lillard and Willis (1994), and Ermisch and Francesconi (2001). Table 2 gives the regression results when the data from the two generations are pooled together.[8]

[Insert Table 2 about here.]

Consider the pooled estimates with Huber's adjustment in Table 2. The numbers of younger or older siblings refer to the child in question, but the birth cohort dummies refer to that of the child's father. The reference group for the father's birth cohort is "father born before 1920."[9] The regression result in Column 3, marked by model 2, includes the same variables as those of model 1 (Column 2) and the dummy variable "born after 1956", which signifies the structural change of mandatory education from 6 to 9 years as mentioned in Section 2. Comparing models 1 and 2, we see that allowing for the structural change improves the fit, but does not alter significantly the impacts of other variables on schooling years of a child. The fourth column (model 3) of Table 2 adds the interaction terms between certain explanatory variables and the generation dummy in order to capture the coefficient change

[8]There are nearly 3,000 interviewees born between 1934 – 1964. For these interviewees, most of them and their siblings have finished their education by the time of the interview. But because our goal is to estimate the family resource allocation across generations, we can employ only interviewees who are old enough so that some of their children have completed the education. Thus, interviewees who do not have children older than 22 years old are deleted. This reduces the effective sample families to roughly 1,500. In addition, several observations contain missing information such as father's birth year. Deleting these points, we have 1,364 families of data used in the estimation of Table 2.

[9]For the old generation, there are 2,496 observations with their fathers born before 1920, 1,088 observations with their fathers born in 1920-1929, and 178 observations with their fathers born after 1929. For the young generation, there are 57 observations with their fathers born before 1929, 732 observations with their fathers born between 1930-1939, and 1,870 observations with their fathers bron after 1940.

associated with the young generation.[10]

From Table 2, we observe that the gender coefficient is always negatively significant, indicating a clear pattern of unfavorable education achievement against the female. As expected, this gender difference is most pronounced in the old generation and it is significantly lessened in the young generation. Indeed, as one can see from the coefficient of the product term of generation and gender; the net effect is even slightly positive (1.99-1.97 = 0.02), albeit statistically insignificant, for the young-generation females. The reason may be either the relaxed budget constraints, or parents' equalizing preferences, toward the young generation.

As to the effect of sibling sizes of different sexes and orders, we find that *older* brothers or sisters always have a positive effect on the child's education. This positive effect is particularly significant for older sisters, a result consistent with the finding in Greenhalgh (1985). However, such a positive effect decreases substantially in the young generation, as one can see from the negatively significant coefficient of the product terms of generation dummy with older siblings. This phenomenon is consistent with the general pattern of economic development: In Taiwan's early development period with general pro-boy perceptions, parents tend to make female older children to join the labor market early so that their incomes can support the education of younger children, especially younger boys. As the family income increases along with economic development, parental budget constraints relax, and hence the original dependence of younger children's education on older siblings' incomes is reduced. Of course, parents' attitude toward the gender of

[10]The father's cohort captures the social as well economic background of an observation. We have tried to replace this cohort dummy by the per capita GDP by the time when the child in question finishes his or her elementary school. But this is not possible because 1) Taiwan's GDP accounting starts only since 1952 and 2) some old-generation children finished their elementary school in Mainland China, and hence their corresponding per capita GDPs cannot be compared with that of Taiwan.

their children may be more equal for the young generation, which may also help explain the phenomenon. In summary, the macro pattern shows that there is little disadvantage for female children of the young generation, but we shall investigate in more details the *micro* changes behind such a macro pattern.

On the other hand, the number of *younger* sisters or brothers always have a negative effect on the education of the child, revealing the crowding (i.e., resource dilution) effect of younger siblings. This is consistent with the evidence found in most previous literature, e.g. Parish and Willis (1993). Although the coefficients of the product term of younger sibling size and the generation dummy also have reverse signs (indicating a weakening of the crowding effect), they are not statistically significant.

Parents' education shows a positive effect on those of their children, a result consistent with the general perception. This positive correlation is weakened in the young generation, perhaps due to the trend of public and mandatory education, which weakens the importance of parental background. Fathers having more professional occupations are generally richer and tend to provide better educational opportunity and support for their children; this is revealed in the positive coefficient of the father's occupation variable.[11] Again, the impact of father's occupation on children's education is significantly reduced in the young generation. Finally, the reference group of ethnicity is Taiwan's aborigines; the positively significant coefficients for all three new migrant groups listed show that the aborigines indeed have inferior educational achievement. The relative difference in educational achievement among new migrants will be further discussed later.

In Table 3, we rerun the regressions separately for the old and young generations, and for children of different sexes. The reference group of father's

[11]The occupation variable used here is similar to that in Erikson and Goldthorp (1992), where the reader can find more detailed explanation.

birth cohort is "father born before 1920" for the old generation, and "father born between 1920-1929" for the young generation. Since there are only 57 observations with their fathers born in the 1920-1929 cohort, it is not surprising that the corresponding coefficient is insignificant. As we can see from Table 3, for the young generation, the negative (crowding) effect of younger siblings remains the same, whereas the originally positive effect of older siblings (on younger ones' education) disappears or reverses, a result consistent with the finding in Table 2. For instance, the impact of older brothers and sisters on a (younger) child's education changes from positive to negative or insignificantly different from zero. This suggests that the original *supporting* effect of older sibling, thanks to the improved economic environment, has become a *crowding* effect to junior children in the young generation. Notice that whenever we find significant crowding sibling effects in Table 3, a junior female child always receives a larger impact (in absolute value) than a junior male child. This remains true for the young generation, indicating that gender-specific preferential treatment may still exist in the young generation. We shall return to this point in Section 4. Table 3 also confirms that the importance of parental education and occupation also reduces in the young generation, again indicating the increasing importance of public education and the declining role of family background.

[Insert Table 3 about here.]

Among the three groups of new migrants, it is observed from Table 3 that, for the old generation, educational achievement for mainlanders (the most recent migrants) are significantly larger than that of Fukien and Hakka (the earlier migrants). But for the young generation, the difference is either blurred or reversed, indicating that the superiority of mainlanders in educational achievement disappears in the young generation. This is indeed intuitively appealing. For recent migrants who mostly fled the chaotic environment of Mainland China during the Chinese civil wars in the 1940s, they

certainly realized that "carryable capital goods" are not physical assets, but human resources. Thus, they tended to invest more on their children's education. As time passes and as the young generation gradually loses the memory of the chaotic past, they tend to behave like old migrants and decrease their educational investment on children.

4 The Changing Pattern of Sex Preferences: A Micro Analysis

In Table 2, the coefficient of $generation \times gender$ being positively significant only says that parents' gender discriminations against girls are weaker for the young generation, perhaps due to the slackness of budget constraints in modern economic environment, or the more mature conception of gender equality in general. However, further analysis from Table 3 tells us that, as far as the crowding effect is concerned, female children of both generations seem to be affected more acutely. In order to have a better understanding about the scenario, we need a micro study to investigate whether there is any change of attitude at the family level.

Suppose a member in generation o has experienced unfair treatment against girls in educational opportunity. We would like to know how would this experience affect the educational resource allocation toward his or her own children in generation y. In this section, we propose two ways to characterize and test the existence and direction of such a micro intergenerational carry-over effect.

Consider a modified version of Equation (1) as follows:

$$Y_{sij} = \alpha_{si} + \gamma_{si} S_{sij} + \beta_s X_{sij} + \eta Z_{si} + \epsilon_{sij}, \quad s = y, o, \tag{2}$$

where γ_{si} is the sex-bias parameter of family i in generation s. The major difference between Equations (1) and (2) is that the gender effect is allowed

to be distinct across families (actually lineages) indexed by various subscript i's in (2). We can then test several hypotheses of interest.

1. **Compensation** hypothesis. It suggests that parents who have experienced unfavorable treatment to girls when they were young tend to treat their own daughters better.

2. **Habitus** hypothesis. It implies that parents who have experienced unfavorable treatment to girls when they were young tend to treat their own children in a similar fashion.

3. **Reenforcing** hypothesis. It suggests that parents' habitus preferences against females are strengthened in the young generation.

Although Equation (2) for the old generation is easy to understand conceptually, it involves estimation of 971 sex-bias parameters $\hat{\gamma_{oi}}$ (the lineage size 971 is explaned in footnote 12 below). This requires significant computer work and is formidable until recently. The newest version of STATA published in the first quarter of 2002 allows us to estimate regression with up to 11,000 parameters. A simplified but less efficient approach for those who do not have the sophisticated software is given in the Appendix.

One way to test the possible existence of a lagged effect of sex-based unfair resource allocation is the following. We run Equation (2) separately for generations y and o respectively, and obtain a set of paired gender effect parameters $\{(\hat{\gamma}_{oi}, \hat{\gamma}_{yi})\}$ for all the families with sufficient data points. If the sex-based unfair allocation of resources has a habitus (compensation) effect from generation o to generation y, then we should observe a positively (negatively) significant correlation between these two estimates. A simple test using correlation coefficients can then be exercised to see if the lagged effect exists. While this approach is intuitively appealing, it treats the estimates $\{(\hat{\gamma}_{oi}, \hat{\gamma}_{yi})\}$ as data in testing the correlation coefficient. In other words, this

approach uses a two-stage procedure to make inference and appears to be indirect. We shall use the following alternatives.

The second way to test the possible existence of a carry-over effect of sex-based unfair allocation is to run Equation (2) first for the old generation to obtain an estimate $\hat{\gamma}_{oi}$, and then run the following equation for generation y:

$$Y_{yij} = \alpha_{yi} + \delta(\hat{\gamma}_{oi} \times S_{yij}) + \beta_y X_{yij} + \eta Z_{yi} + \epsilon_{yij}. \tag{3}$$

We then test the significance of δ according with the various hypotheses listed above. Specifically, when $\delta < 0$ ($0 < \delta \leq 1$, $\delta > 1$), it suggests that the compensation (habitus, reenforcing) hypothesis applies. Evidently, this second approach uses the lineage-specific information of sex preferences in an earlier generation to infer the possible influence on individuals of the same lineage.

The estimation results of Equations (2) and (3) are presented in Table 4. As one can see from the Table, coefficients of most variables are the same as those in Table 3: i) the size of senior siblings has a supporting effect for the juniors' education in generation o, but has a crowding effect (i.e. resource dilution) in generation y; ii) the size of junior siblings has a crowding effect in both generation o and generation y; iii) the influence of parents' education and occupation on child education reduces, either in scale or in significance, in the y generation; iv) the influence of ethnicity background (among new migrants) on child education decreases in the y generation, but the difference between new migrants and aborigines still exists.[12]

[12]As one can see, there are 1,364 families in the estimation of Equation (2). But to estimate the regressions in Table 4, we have to delete lineages which have a missing value in *any* generation. This additional restriction limits the number of sampled families to 971. To obtain columns 3 and 4 of Table 4, we only keep lineages which have children of *both* sexes in order to do the gender discrimination analysis. The lineage size then becomes 750. Notice that equations (2) and (3) involve panel data, hence the actual individuals *observations* used in the regression are much larger than the size of families.

[Insert Table 4 about here.]

The main focus of Table 4 is the coefficient of [previous generation's family specific gender effects * gender] (i.e. δ) in Equation (3). The result shows that the habitus hypothesis cannot be rejected at the (one-tail) 5% level. Specifically, the estimate 0.0357 is significantly different from zero at the one-side 5% level and it is obvious less than unity. This result can be compared with what is found in Table 2 to gain insight into the intergeneration effect. Combining the coefficients of [gender] and [generation * gender] in Table 2, we see that there is essentially no macro gender effect for the young generation. Table 4, on the other hand, shows that the sex-based discrimination persists within many lineages into the young generation, even though the average habitus coefficient is fairly small in magnitude with only .0357 year in the educational achievement.

Some further improvement can be made for the estimation in Table 4. The first improvement is to fit Equation (3) only using data from families in which there were significant gender discrimination in the old generation; i.e., when their corresponding $\hat{\gamma}_{oi}$'s are significant. To do this, we modify Equation (3) as

$$Y_{yij} = \alpha_{yi} + \delta(\hat{\gamma}^*_{oi} \times S_{yij}) + \beta_y X_{yij} + \eta Z_{yi} + \epsilon_{yij}, \tag{4}$$

where

$$\hat{\gamma}^*_{oi} = \begin{cases} \hat{\gamma}_{oi}; & \text{if } \hat{\gamma}_{oi} \text{ is significant} \\ 0; & \text{otherwise.} \end{cases}$$

Equation (4) assumes that the intergenerational carry-over effect exists only in families where the old generation *did* have a gender discrimination. Results (listed in the second column of Table 5) show that this improvement only has some minor impacts.

The second improvement is to run regression (3) for male and female old-generation children separately. The reason is as follows. If there is a gender discrimination in an old-generation family, the psychological imprints on an

old-generation boy (the privileged) and an old-generation girl (the deprived) should be different. Intuitively, for male respondents (the privileged), since they did not experience the pain of "losing" support, their psychological imprint may not be very strong. For the deprived female respondents, since they had less education compared with their male siblings, they may be more likely to preserve the habitus imprint of their families and treat their children in a similar pattern.

The third improvement is to consider subsamples of the young generation with different cutoff ages. Since there is a 2-year mandatory military service for young men in Taiwan, many males finish their college study at the age of 24. In Table 4, we only include young generation respondents aged older than 24. To test the sensitivity of this cutoff age, we also tried other cutoff thresholds, such as 26 and 28.

We take into account these improvements and combine the results in Table 5. In the table, $\hat{\gamma}^*_{oi}$ instead of $\hat{\gamma}_{oi}$ is used so that across-generation influence for families with insignificant old-generation gender discrimination is assumed to be zero. From the Table, we see that female children from a family with gender discrimination do have significant habitus effect as they treat their own children, whereas male children in such families do not show such a tendency. This suggests that the psychological imprints are indeed more likely to be on the "deprived" rather than on the "privileged" children. This is a very interesting result which may deserve more attention and discussion from psychologists.

As far as the intergenerational carry-over effect of gender discrimination is concerned, the change of age-group thresholds does not have much impact on the coefficients and their significance levels. However, as the age threshold becomes older, we are moving toward older cohorts, which also makes the coefficients moving toward the "old generation". For instance, the coefficient of fathers' ethnicity moves to the benefit of mainlanders as the age of children

in question becomes older. This is of course consistent with our previous discussion.

[Insert Table 5 about here.]

5 Conclusions and Remarks

The purpose of this paper is to test whether there is an intergenerational transmission of gender preferences in educational resource allocation among children. The unique data set of Taiwan's Panel Study of Family Dynamics project provides us a rich 3-generation education information and allows us to probe into this question. We performed our analysis along two directions: the first is to see whether the society as whole has any macro change in gender-specific education achievement, and the second is to see whether there is any within-lineage transmission of gender preferences across generations.

After carefully reviewing the education system and societal characteristics in Taiwan, we set up an empirical model to estimate and test the hypotheses of intergenerational transmission of gender preferences. As far as the macro pattern is concerned, we found that although there is a clear tendency of differential treatment against females in the old generation, this tendency is significantly weakened and nearly vanishes in the young generation. Moreover, the supporting effect of senior siblings in the old generation becomes a crowding (resource-dilution) effect in the young generation. However, within each lineage, there is a mild habitus effect in gender-specific educational resource allocation in the sense that parents who had the experience of gender-specific differential treatment tend to treat their children in a similar fashion. Moreover, this habitus effect is stronger for female respondents (who were the deprived group) than for male respondents (who were the privileged group).

As we know, the preferential treatment of child education is more related

to parental attitude, and is also more under the control of parents. One type of gender-specific differential treatment can be observed in the job market, as the literature has extensively studied. Another line of research worthy of exploring is to study the education-earnings relationship, and to investigate whether we can find any gender-specific preferential treatment in the old and young generations and whether such a preferential treatment has any pattern of intergenerational transmission. The PSFD data set does not have sufficient information at the current stage; we hope that more earnings data will be available in the future so that research projects along this line can be investigated.

Appendix: A Simplified Approach of Estimating (2) and (3)

For boys and girls of generation o, equation (2) can be written respectively as

$$Y_{oij} = \alpha_{oi} + \gamma_{oi} + \beta_o X_{oij} + \eta Z_{oi} + \epsilon_{oij}, \quad j \in boys \tag{A1}$$

$$Y_{oij} = \alpha_{oi} + \beta_o X_{oij} + \eta Z_{oi} + \epsilon_{oij}, \quad j \in girls \tag{A2}$$

Take marginal means of Y's, X's and ϵ's with respect to j over different sexes, and denote them respectively by $(\bar{Y}_{oib}, \bar{Y}_{oig})$, $(\bar{X}_{oib}, \bar{X}_{oig})$, and $(\bar{\epsilon}_{oib}, \bar{\epsilon}_{oig})$, where b and g refer to boys and girls respectively. Subtract such means from both sides of $(A1)$ and $(A2)$, we have the following deviation forms:

$$(Y_{oij} - \bar{Y}_{oib}) = \beta_o(X_{oij} - \bar{X}_{oib}) + (\epsilon_{oij} - \bar{\epsilon}_{oib}), \quad j \in boys,$$

$$(Y_{oij} - \bar{Y}_{oig}) = \beta_o(X_{oij} - \bar{X}_{oig}) + (\epsilon_{oij} - \bar{\epsilon}_{oig}), \quad j \in girls.$$

Since ϵ is independent of sexes by assumption, the above two equations can be pooled together to obtain a consistent estimator for $\hat{\beta}_o$. Given $\hat{\beta}_o$, the gender effect of family i in generation o can be obtained algebraically as follows:

$$\hat{\gamma}_{oi} \equiv [(\bar{Y}_{oib} - \bar{X}_{oib}\hat{\beta}_o)] - (\bar{Y}_{oig} - \bar{X}_{oig}\hat{\beta}_o)].$$

The same task can be done for generation y, and therefore we can generate a series $\hat{\gamma}_{yi}$.

The above approach can obtain consistent estimators for γ_{oi} and γ_{yi}, but some efficiency is lost. In particular, in the process of differencing equations $(A1)$ and $(A2)$, the family-specific variables $(Z_{si},\ s = o, y)$ are sacrificed. Readers are suggested to apply the more efficient approach in the text if they can obtain the 2002 STATA/SE software.

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Table 1. Means and Standard Deviations for Selected Variables

	Both Generations		Old Generation		Young Generation	
	Male	Female	Male	Female	Male	Female
	1	2	3	4	5	6
age	40.7	41.5	52.3	52.2	30.2	30.5
	(13.0)	(12.9)	(8.09)	(7.77)	(5.82)	(5.91)
Sibling size + 1	4.91	5.37	6.39	6.68	3.56	4.01
	(2.17)	(2.17)	(2.06)	(2.00)	(1.16)	(1.34)
Birth order	2.68	2.72	3.20	3.21	2.20	2.22
	(1.59)	(1.56)	(1.79)	(1.72)	(1.20)	(1.19)
Number of elder brothers	0.806	0.818	1.09	1.06	0.543	0.567
	(0.997)	(0.968)	(1.16)	(1.09)	(0.725)	(0.746)
Number of elder sisters	0.870	0.903	1.11	1.15	0.654	0.648
	(1.09)	(1.10)	(1.20)	(1.19)	(0.923)	(0.922)
Number of younger brothers	1.12	1.35	1.60	1.75	0.691	0.935
	(1.23)	(1.23)	(1.42)	(1.39)	(0.796)	(0.859)
Number of younger sisters	1.10	1.29	1.58	1.70	0.670	0.866
	(1.29)	(1.40)	(1.51)	(1.54)	(0.828)	(1.09)
Born after 1956 (%)	59.5	56.9	16.5	16.2	98.6	98.9
First-born (%)	29.1	26.7	21.9	19.1	35.6	34.6
Last-born (%)	20.0	14.0	13.2	8.90	26.0	19.2
Years of schooling	10.7	9.39	8.81	6.83	12.3	12.0
	(4.16)	(4.69)	(4.39)	(4.52)	(3.10)	(3.14)
Father's education	5.68	5.51	3.76	3.72	7.43	7.35
	(4.56)	(4.43)	(4.25)	(4.12)	(4.11)	(3.95)
Mother's education	3.71	3.49	1.83	1.78	5.42	5.25
	(4.09)	(4.04)	(3.32)	(3.26)	(3.96)	(4.00)
Father's occupation	5.53	5.50	4.28	4.34	6.67	6.69
	(3.01)	(2.93)	(1.83)	(1.79)	(3.40)	(3.37)
Father's ethnicity (%)						
Aborigines	1.97	2.59	1.39	2.48	2.48	2.72
Fukien	80.4	78.3	81.8	80.1	79.2	76.4
Hakka	11.7	12.7	12.2	13.1	11.3	12.2
Mainlander	5.92	6.45	4.67	4.32	7.06	8.64
Number of observations	3917	3738	1864	1898	2053	1840
Number of families	1304	1250	831	823	1168	1054

Notes：

1. The second generation's children are confined to those who are aged 22 or older in year 2000.
2. Father's occupation is measured on a scale of 0-10, where 10 is professional and technical occupation while 0 is out of labor force. For the old generation, this is their father's longest occupation. For the young generation, this is their father's current occupation during the survey year.

Table 2. Effects of Sibling and Parental Education on Schooling (Pooling Two Generations Together): LS estimation with Huber's adjustment

Model	1	2	3
Father's birth cohort			
1920-1929	1.06	0.753	0.632
	(4.97)	(3.35)	(2.79)
1930-1939	2.21	1.32	1.78
	(11.1)	(5.39)	(5.74)
After 1940	2.57	1.60	2.30
	(13.1)	(6.36)	(7.22)
Born after 1956 (Yes=1)		1.07	1.09
		(5.49)	(5.15)
Gender (Female=1)	-0.948	-0.977	-1.97
	(-11.4)	(-11.8)	(-14.9)
Generation*Gender			1.99
(young generation=1)			(12.1)
Number of elder brothers	0.082	0.058	0.076
	(1.32)	(0.93)	(1.00)
Generation* # of elder brothers			-0.153
			(-1.41)
Number of elder sisters	0.160	0.120	0.177
	(3.19)	(2.39)	(2.53)
Generation* # of elder sisters			-0.202
			(-2.19)
Number of younger brothers	-0.294	-0.259	-0.287
	(-5.08)	(-4.42)	(-4.26)
Generation* # of younger brothers			0.081
			(0.71)
Number of younger sisters	-0.262	-0.236	-0.259
	(-5.02)	(-4.53)	(-3.91)
Generation* # of younger sisters			0.058
			(0.68)
Father's education	0.264	0.264	0.303
	(14.0)	(14.1)	(9.29)
Generation* father's education			-0.116
			(-3.10)
Mother's education	0.228	0.227	0.281
	(11.2)	(11.2)	(7.57)
Generation* mother's education			-0.079
			(-1.85)
Father's occupation	0.057	0.053	0.170
	(2.97)	(2.75)	(2.94)
Generation* father's occupation			-0.134
			(-2.29)
Father's ethnicity			
Fukien	1.70	1.74	1.86
	(5.20)	(5.32)	(5.69)
Hakka	2.14	2.18	2.33
	(6.21)	(6.30)	(6.69)
Mainlander	2.63	2.55	2.73
	(6.48)	(6.19)	(6.66)
R^2	0.4732	0.4766	0.4990
Number of observations	7655	7655	7655
Number of families	1364	1364	1364

Note: Asymptotic t-values are in parentheses.

Table 3. Effects of Sibling and Parental Education on Schooling: By Generation and Sex

	1	2	3	4
	Old Generation		Young Generation	
	Male	Female	Male	Female
Father's birth cohort				
1920-1929	0.389	1.15		
	(1.37)	(4.15)		
1930-1939	0.670	1.46		0.676
	(1.52)	(3.08)		(1.39)
After 1940			0.084	1.23
			(0.48)	(2.45)
Born after 1956 (Yes=1)	0.974	0.751	0.133	3.14
	(3.33)	(2.58)	(0.14)	(3.87)
Number of elder brothers	0.154	0.216	-0.107	-0.504
	(1.68)	(1.89)	(-1.02)	(-4.15)
Number of elder sisters	0.250	0.304	-0.170	-0.091
	(2.56)	(3.39)	(-2.11)	(-1.11)
Number of younger brothers	-0.114	-0.334	-0.279	-0.507
	(-1.33)	(-3.76)	(-2.06)	(-4.47)
Number of younger sisters	-0.151	-0.238	-0.233	-0.385
	(-1.66)	(-3.01)	(-2.62)	(-5.25)
Father's education	0.296	0.284	0.206	0.142
	(7.80)	(6.70)	(8.84)	(5.54)
Mother's education	0.233	0.319	0.193	0.193
	(5.42)	(6.78)	(7.87)	(6.93)
Father's occupation	0.228	0.220	-0.018	0.052
	(3.15)	(2.91)	(-0.83)	(2.04)
Father's ethnicity				
Fukien	1.65	2.41	1.18	1.95
	(2.38)	(3.40)	(2.29)	(5.22)
Hakka	2.46	3.04	1.07	2.46
	(3.30)	(4.09)	(1.86)	(5.93)
Mainlander	4.27	3.99	0.916	2.30
	(4.82)	(4.84)	(1.54)	(4.36)
R^2	0.3373	0.3860	0.2560	0.2899
Number of observations	1864	1898	2053	1840

Note: Asymptotic t-values are in parentheses.

Table 4. Estimates of Intergenerational Transmission Effects of Sex Preference on Schooling

	Old Generation Equation (2)	Young Generation Equation (3)
Father's birth cohort		
1920-1929	1.18	
	(3.81)	
1930-1939	1.57	0.722
	(3.04)	(1.11)
After 1940		0.644
		(0.99)
Born after 1956 (Yes=1)	0.545	1.82
	(2.14)	(1.68)
Previous generation's family specific gender effects * gender		0.0357
		(1.75)
Number of elder brothers	0.131	-0.320
	(1.17)	(-2.60)
Number of elder sisters	0.289	-0.156
	(2.96)	(-1.88)
Number of younger brothers	-0.299	-0.307
	(-3.07)	(-2.30)
Number of younger sisters	-0.247	-0.280
	(-2.84)	(-3.35)
Father's education	0.286	0.175
	(5.99)	(6.44)
Mother's education	0.320	0.233
	(6.02)	(7.83)
Father's occupation	0.216	0.0288
	(2.53)	(1.03)
Father's ethnicity		
Fukien	2.41	2.04
	(3.05)	(5.13)
Hakka	3.04	2.18
	(3.66)	(4.70)
Mainlander	3.97	2.53
	(4.29)	(4.57)
F-value of family specific gender Dummies	2.38 d.f. (831, 2917)	
R^2	0.6095	0.2848
Number of observations	3762	2049
Number of families	971	750

Note: Asymptotic t-values are in parentheses.

Table 5. Estimates of Sex-Preferences Transmission Effects (using γ^* which is significant at 1% level)

	Age>24			Age>26			Age>28		
	Children of all respondents 1	Children of male respondents 2	Children of female respondents 3	Children of all respondents 4	Children of male respondents 5	Children of female respondents 6	Children of all respondents 7	Children of male respondents 8	Children of female respondents 9
Born after 1956 (Yes=1)	1.83	2.06	1.73	1.77	2.00	1.63	1.69	2.00	1.56
	(1.68)	(2.93)	(1.02)	(1.69)	(2.63)	(1.01)	(1.69)	(2.79)	(1.01)
Previous generation's family specific gender effects * gender	0.0218	-.0114	0.0441	0.0386	0.00923	0.0564	0.0423	-.00298	0.0709
	(1.49)	(-.52)	(2.15)	(2.44)	(0.38)	(2.58)	(2.38)	(-0.11)	(2.88)
Number of older brothers	-0.305	-0.416	-0.186	-0.328	-0.401	-0.259	-0.350	-0.406	-0.293
	(-2.41)	(-2.60)	(-1.02)	(-2.40)	(-2.29)	(-1.32)	(-2.25)	(-1.97)	(-1.33)
Number of older sisters	-0.154	-0.257	-0.0628	-0.163	-0.240	-0.0893	-0.172	-0.259	-0.0899
	(-1.86)	(-2.19)	(-0.54)	(-1.86)	(-1.86)	(-0.75)	(-1.79)	(-1.86)	(-0.67)
Number of younger brothers	-0.295	-0.273	-0.300	-0.305	-0.241	-0.351	-0.321	-0.280	-0.352
	(-2.16)	(-1.73)	(-1.50)	(-2.05)	(-1.41)	(-1.60)	(-1.93)	(-1.46)	(-1.40)
Number of younger sisters	-0.276	-0.375	-0.159	-0.326	-0.381	-0.253	-0.354	-0.422	-0.261
	(-3.31)	(-3.50)	(-1.21)	(-3.74)	(-3.32)	(-1.88)	(-3.89)	(-3.54)	(-1.80)
Father's education	0.175	0.206	0.157	0.181	0.212	0.163	0.175	0.203	0.161
	(6.47)	(5.08)	(4.28)	(6.28)	(5.00)	(4.15)	(5.30)	(4.16)	(3.64)
Mother's education	0.236	0.236	0.234	0.240	0.257	0.227	0.253	0.260	0.245
	(7.93)	(5.45)	(5.69)	(7.28)	(5.40)	(4.98)	(6.64)	(4.87)	(4.60)
Father's occupation	0.0299	0.0440	0.00850	0.0440	0.0491	0.0343	0.0572	0.0586	0.0536
	(1.07)	(0.99)	(0.23)	(1.41)	(1.00)	(0.81)	(1.61)	(1.06)	(1.11)
Father's ethnicity									
Fukien	2.06	2.09	2.02	1.95	1.99	1.86	1.60	1.70	1.47
	(5.01)	(4.68)	(3.27)	(4.36)	(5.75)	(2.58)	(3.81)	(5.98)	(1.89)
Hakka	2.21	2.27	2.08	2.13	2.04	2.10	1.61	1.49	1.60
	(4.63)	(4.18)	(2.95)	(4.02)	(3.95)	(2.54)	(2.92)	(2.74)	(1.72)
Mainlander	2.56	1.55	2.83	2.55	1.29	2.81	2.36	1.03	2.54
	(4.55)	(2.02)	(3.73)	(4.16)	(2.10)	(3.25)	(3.69)	(1.52)	(2.61)
R^2	0.2843	0.3191	0.2666	0.2969	0.3233	0.2862	0.2926	0.3021	0.2954
Number of observations	2049	903	1146	1727	755	972	1406	605	801
Number of families	750	328	424	638	277	362	536	237	299

Note: Asymptotic t-values are in parentheses. Other explanatory variable includes father's birth cohort.

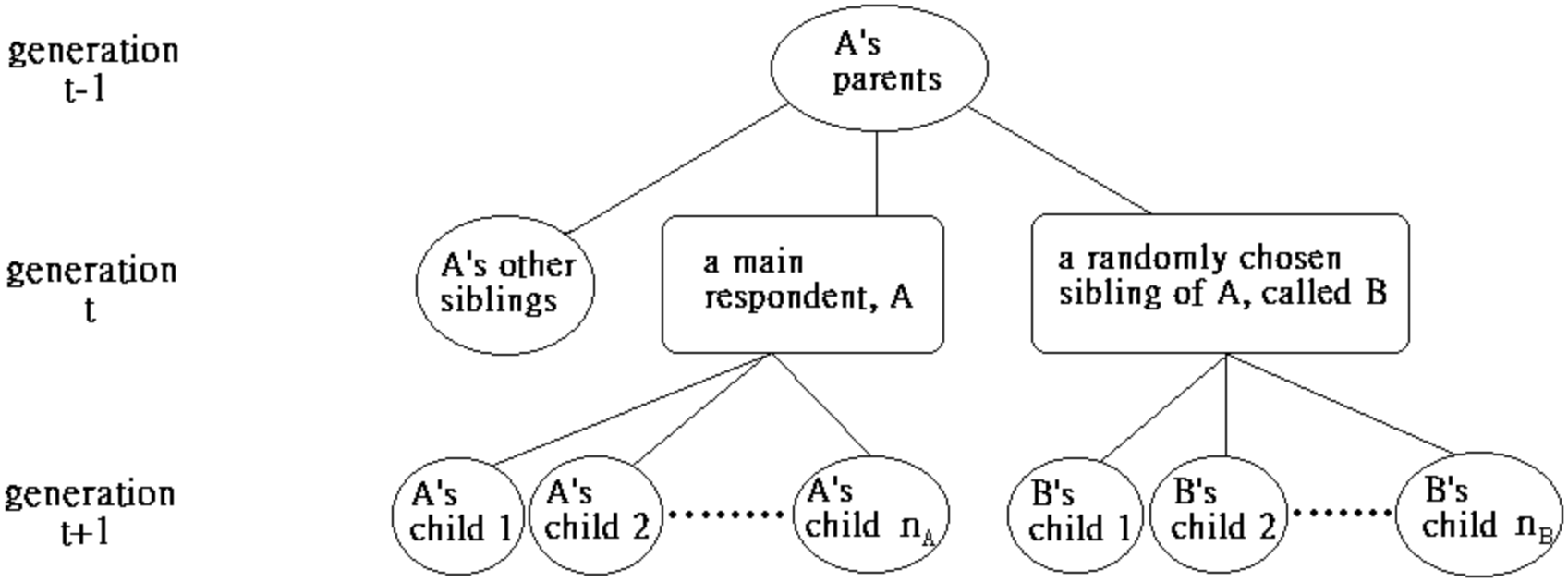

Figure1 : The Sample and Information Structure.
Squares indicate face-to-face interviews, whereas circles indicate information covered by interview questionaires.