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Parental Condition and Infant Sex at Birth in the Japan Environment and Children's Study: A Test of the Trivers–Willard Hypothesis

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The Trivers–Willard hypothesis predicts that females in good condition should bear more sons rather than daughters in certain mammals, including humans. This study tests the hypothesis by using 66,638 childbirth records from a national birth cohort survey in current Japan. Our analyses showed that, contrary to the hypothesis, indicators of parental condition, such as mother's age, body mass index, job status, education level, medical history, or household income, had few statistically significant effects on infant sex at birth. In previous studies investigating the Trivers–Willard hypothesis, the results have been quite mixed and inconclusive. We discuss some theoretical and methodological challenges towards a precise understanding of the hypothesis in human populations.

Keywords
birth sex ratio, birth cohort study, reproductive strategy, evolutionary demography, secondary data analysis, Japan

Introduction

According to evolutionary ecological theories, parental (especially maternal) condition will affect child sex ratio at birth. Trivers and Willard (1973) hypothesized that females in good condition should bear more sons, rather than daughters, in polygamous mammals in certain ecological situations. This hypothesis is widely known as the Trivers–Willard hypothesis (hereinafter called T–WH). Here, social rank, nutritional status, and accessibility to resources are often used as measures of the parental condition. The T–WH could be supported under the following three assumptions: (1) good maternal condition leads to good condition of her offspring; (2) one's good condition at child stage results in his/her good condition at adult stage; and (3) males compete for their mating and reproductive opportunities with each other, where those in better condition have higher reproductive success than individuals in worse condition. There exist numerous studies on the T–WH, but the results have been quite mixed (e.g., Douhard, 2017; West, 2009, see also Veller, Haig, & Nowak, 2016 for the T–WH on sex-biased parental investment).

The T–WH has been also applied to humans (e.g., Kolk & Schnettler, 2016; Luo, Ding, Gao, Sun, & Zhao, 2017). It is well recognized that humans satisfy the aforementioned three ecological assumptions for the T–WH, even for populations in modern developed and socially-monogamous settings (see Hopcroft, 2005; Kolk & Schnettler, 2016 for details). As is the case of nonhuman mammals, inconclusive results have been reported in humans too (e.g., James, 2012). This ambiguity needs and prompts further investigation. Most of the previous studies were based on data in the United States, Europe, and some traditional African societies. On the other hand, this study analyzes modern Japanese data from a national birth cohort survey. We believe that cross-cultural and descriptive studies will take a substantial role in evaluating the T–WH in human populations.

In our analyses, candidate measures of the parental condition are: mother's age; body mass index (BMI); medical history; and socioeconomic status. We take account of not only demographic and anthropometric measurements, but also some medical information of each subject, which has been studied rarely. It can be assumed that being the appropriate age for reproduction, a healthy physical and medical condition, and a high socioeconomic status are indicative of good parental condition.

Subjects and Methods

Dataset

We analyzed the source data from the Japan Environment and Children's Study (JECS), a national birth cohort survey involving about 100,000 parent-child pairs (specifically, 104,102 pregnancy and childbirth records of 97,460 recruited mothers, as of May 2017, see also Michikawa et al., 2015 and the supplementary material). Participants' information was mainly collected by questionnaires given to the parents (mostly mothers), and medical examinations by obstetricians and nurses. This study used the dataset released in June 2016 and revised in October 2016 (Dataset Name: *jecs-ag-20160424*).

Primary subjects of the analysis were 66,638 childbirth records from 2011 to 2014 that included complete information for all variables described below and the child's ID number (i.e., complete case analyses were

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performed). We restricted the sample for data analysis to spontaneous/natural pregnancies (clinical record) and singleton live childbirths, and then focused on each mother's first childbirth during the survey period only, due to a smaller number of cases outside of the criteria and for straightforward statistical analyses. To accumulate empirical evidence, we supplementarily examined the T–WH for subjects who had no experience of pregnancy before participating in the JECS only (n = 20,973). These cases meant the mother's first pregnancy and childbirth in her life course.

Statistical Analyses

To study the effects of parental condition on infant sex at birth, the generalized linear mixed models (GLMMs) with the binominal error distribution were adopted. The dependent variable was infant sex (0: girl, 1: boy). The independent variables were the following items: mother's age at the early pregnancy period (unit = years; continuous); the age squared (age^2; to examine the intermediate optimum); mother's BMI before the pregnancy (kg/m^2, clinical record; continuous); the BMI squared (BMI^2); mother's highest education level (continuous; five levels from low to high, see Table 1); mother's job status in the early pregnancy period (continuous; three levels, 1: permanent or self-employed including the assistant, 2: temporary/contractual, 3: housewife or suspension/unemployed); household annual income in the mid pregnancy period (continuous; nine levels from low to high, see Table 1); and mother's medical history during the past one year from when the subject answered in the early pregnancy period (binary;

Table 1. Descriptive statistics of the sample used in the main analysis.

Variable	Range	Mean	S.D.	*n*
Infant sex				
Girl	-	-	-	32,589
Boy	-	-	-	34,049
Mother's age (years old)	15–48	30.73	4.89	-
Mother's BMI	13.34–52.77	21.22	3.26	-
Mother's highest education level				
Junior high school	-	-	-	2,936
High school	-	-	-	20,766
Technical college / vocational school / junior college	-	-	-	28,056
University	-	-	-	13,920
Graduate university	-	-	-	960
Mother's job status (type of employment)				
Permanent or self-employed including the assistant	-	-	-	25,849
Temporary/contractual	-	-	-	15,756
Housewife or suspension/unemployed	-	-	-	25,033
Household annual income (JPY)				
< 2 M	-	-	-	3,727
≥ 2 to < 4 M	-	-	-	23,161
≥ 4 to < 6 M	-	-	-	22,123
≥ 6 to < 8 M	-	-	-	10,547
≥ 8 to < 10 M	-	-	-	4,376
≥ 10 to < 12 M	-	-	-	1,556
≥ 12 to < 15 M	-	-	-	616
≥ 15 to < 20 M	-	-	-	340
≥ 20 M	-	-	-	192
Mother's medical history (experience of illnesses)				
Absence	-	-	-	58,811
Presence	-	-	-	7,827
Past childbirth experience (times)	0–11	0.86	0.90	-
Year of childbirth				
2011	-	-	-	6,953
2012	-	-	-	19,809
2013	-	-	-	23,689
2014	-	-	-	16,187
Number of people who earn a salary in the family	0–9	1.72	0.77	-
Father's highest education level				
Junior high school	-	-	-	4,780
High school	-	-	-	24,557
Technical college / vocational school / junior college	-	-	-	15,104
University	-	-	-	19,267
Graduate university	-	-	-	2,930

0: none of the following illnesses, 1: had measles, rubella, tuberculosis, influenza, and/or mumps). With regards to the medical history, we focused on the five illnesses because these reflected recent health condition that should have strong influences on pregnancy and childbirth.

As control variables, the following items were also incorporated in the statistical model: past childbirth experience (clinical record; unit = times); year of the childbirth (categorical; from 2011 to 2014); the number of people who earned a salary in the family (unit = person); and father's highest education level. The regional code (20 categories) was incorporated as a random effect in the GLMMs. We did not examine interactions between the independent variables. Table 1 is the descriptive statistics of the sample used in the main analysis. In the supplementary analysis for subjects who had no prior experience of pregnancy, the GLMM included information of past pregnancy experience instead of past childbirth experience (see Table S1 for the descriptive statistics).

We assumed that parental condition throughout one's pregnancy period should reflect, to some extent appropriately, both the mother and her family's condition at the timing of conception that would be a more accurate indicator. The GLMMs were performed by using R software version 3.2.4 for Mac OS with the *glmmML* function (family = binomial). This study has been approved by the Kyoto University Graduate School and Faculty of Medicine, Ethics Committee (#R0628).

Results

The main GLMM analysis did not find any factors that affected infant sex at birth significantly (here $p < .05$, Table 2). There existed no strong correlation (Pearson's correlation coefficient $|r| > .5$) between the independent variables that could produce statistical multicollinearity (note that the year of childbirth, a categorical factor, was not included in the calculation). Predictions from the T–WH were not supported. Candidate measures of good parental condition that we assumed (i.e., appropriate age for reproduction, healthy physical and medical condition, and high socioeconomic status) did not lead to a higher probability of bearing sons in the Japanese sample. With regards to the mother's age and BMI, these quadratic effects were not significant. We also analyzed a model without the age squared term, and confirmed that the linear effect of age was not significant ($p = .165$, details not shown). Our analysis treated household income, education level, and job status as indicators of socioeconomic status, and the experience of certain illnesses as an indicator of medical/health condition. None of these had significant effects on infant sex at birth.

As for the supplementary analysis for subjects who had no prior experience of pregnancy, the GLMM analysis found few effects that affected infant sex at birth significantly (Table 3). Here, the mother's age and the age squared had marginal effects ($p = .072$ and $p = .065$, respectively). According to the regression coefficients of these, the probability of bearing sons was estimated to be the lowest at the age of 29.07. This result would be in contrast with the T–WH predicting that mothers of the appropriate age for reproduction should bear more sons. Additionally, the effect of the number of people who earned a salary in the family had a significant effect on infant sex. However, the effect of household income itself was not significant.

Table 2. Factors affecting infant sex at birth in the main GLMM.

Dependent variable: infant sex (0: girl, 1: boy)				
Independent variable	Coefficient	S.E.	*z*	*p*
Mother's age (continuous)	–0.0158460	0.0163178	–0.97109	.332
Age^2	0.0002961	0.0002624	1.12836	.259
Mother's BMI (continuous)	–0.0072245	0.0174003	–0.41519	.678
BMI^2	0.0002084	0.0003570	0.58357	.560
Mother's education level (continuous)	0.0073916	0.0107910	0.68498	.493
Mother's job status (continuous)	–0.0125929	0.0104830	–1.20128	.230
Household income (continuous)	–0.0039486	0.0070532	–0.55984	.576
Mother's medical history (0: absence, 1: presence)	–0.0271148	0.0241480	–1.12286	.261
Past childbirth experience (continuous)	–0.0101651	0.0094683	–1.07360	.283
Year of childbirth (categorical, reference: 2013)				
2011	–0.0012233	0.0273121	–0.04479	.964
2012	–0.0021125	0.0192804	–0.10957	.913
2014	0.0013403	0.0204230	0.06563	.948
Number of people who earn a salary (continuous)	–0.0105771	0.0117265	–0.90198	.367
Father's education level (continuous)	–0.0015409	0.0085459	–0.18030	.857

Two large values of *r* were −.4675 between mother's job status and the number of people who earn a salary family, and .4464 between mother's education level and father's education level.
The residual deviance was 92,340 on 66,622 degrees of freedom.

Table 3. Factors affecting infant sex in the supplementary GLMM for subjects who had no prior experience of pregnancy.

Dependent variable: infant sex (0: girl, 1: boy)				
Independent variable	Coefficient	S.E.	*z*	*p*
Mother's age (continuous)	–0.0506793	0.0281710	–1.79899	.072
Age^2	0.0008717	0.0004720	1.84665	.065
Mother's BMI (continuous)	0.0313796	0.0337334	0.93022	.352
BMI^2	–0.0004386	0.0007055	–0.62165	.534
Mother's education level (continuous)	0.0197826	0.0193687	1.02137	.307
Mother's job status (continuous)	–0.0053435	0.0189736	–0.28163	.778
Household income (continuous)	0.0040869	0.0127308	0.32103	.748
Mother's medical history (0: absence, 1: presence)	–0.0094144	0.0440060	–0.21393	.831
Year of childbirth (categorical, reference: 2013)				
2011	–0.0240796	0.0502088	–0.47959	.632
2012	0.0047050	0.0351498	0.13386	.894
2014	–0.0187870	0.0353693	–0.53117	.595
Number of people who earn a salary (continuous)	–0.0450058	0.0216472	–2.07906	.038
Father's education level (continuous)	–0.0001918	0.0151570	–0.01265	.990

Two large values of r were −.4247 between mother's job status and the number of people who earn a salary in the family, and .4159 between mother's education level and father's education level.
The residual deviance was 29,050 on 20,958 degrees of freedom.

Discussion

We found no clear evidence that supported the T–WH (Tables 2 and 3) as with a recent study by Kolk and Schnettler (2016). Although the predictions were not supported, this study analyzing a Japanese birth cohort dataset has provided a new piece of fundamental information and descriptive results. Until now, the T–WH and its application to humans have been tested in numerous studies. However, arguments as to whether the hypothesis can be supported or not are quite inconclusive. Below, we briefly present two possible explanations why the hypothesis was (often) unsupported with some future directions.

First, it may need to reconsider the central assumption that humans satisfy the essential ecological situations for the T–WH. In particular, there exist several findings that did not match with the third assumption in terms of the relationship between socioeconomic status and reproductive success for men, especially for populations in modern developed, low-fertility, and socially-monogamous settings (e.g., Morita, in press). On the other hand, as Hopcroft (2005), and Kolk and Schnettler (2016) carefully discussed, psychological mechanisms acquired in ancestral environments through human evolutionary history can be more significant than behavioural patterns and phenotypic outcomes observed in the current environment. Additional and further synthesized discussions on the prerequisites of the T–WH is required now.

Second, there is also a possibility that our selection and treatment of each independent variable could not detect parental condition appropriately. For example, this study analyzed the linear effects of socioeconomic status, because we believe that exploring how higher socioeconomic status can affect reproductive outcomes simply is the primary step to better understand human reproductive strategies (see also Morita, Ohtsuki, & Hiraiwa-Hasegawa, 2016). In the next step, it will be also necessary to examine the intermediate optima or hump-up effects by treating these variables as not *continuous* but *categorical* factors with specific and reasonable biological meanings. Moreover, this study did not take account of any physiological or genetic mechanisms including detailed information on medical history, psychological stress, and exposure to chemical substances. Extensive analyses, such as factors affecting stillbirth and sex ratio at conception, should be desirable to study infant sex fully. We leave these theoretical/methodological limitations and potential research topics for future work.[1]

Note that the T–WH predicts not only *infant sex ratio at birth* but also *sex-biased postnatal parental investment*, for example in educational investment and breastfeeding (e.g., Du & Mace, in press; Fujita et al. 2012; Hopcroft, 2005; Hopcroft & Martin, 2016). Parental investment after birth has a large impact on child survival and quality. We have suggested that Japanese parents did not manipulate infant birth sex, but their postnatal investment may differ between sons and daughters. It is important to examine the hypothesis from different angles in the broader context of sex allocation. The combination of two aspects (i.e., sex ratio and sex-biased investment) can lead to a comprehensive understanding. We hope that longitudinal birth cohort surveys, like the JECS, have the potential to provide rich empirical material for that. Future studies will be able to investigate the T–WH with integrated approaches based on a sample from the same dataset.

Lastly, another consideration to the statistical procedure is given. This study adopted complete case analyses strictly and therefore excluded many childbirth records from our analyses. The number of subjects who had at least one not available (NA: missing; indistinguishable; or excluded from the research design explained above) item for the main analysis was 37,464 (approximately 36% of the total). We understand that the complete case analysis is a proper method for this study based on a large dataset, but some limitations on such a method has been suggested (e.g., whether missing cases occurred at random completely or not) particularly in epidemiology and population health sciences. We could elaborate the statistical method to show more robust results and conclusions.

[1]It has been also suggested that the Trivers–Willard effect will be weak in modern homogenous societies, where the variation in individual's condition is relatively small like current Japan (cf. Luo et al., 2017). Comparative analyses based on populations that can be divided into several sub-populations may find clear effects (see also Mace, 2008).

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Data Accessibility

See the supplementary material.

Supplementary Material

Electronic supplementary material is available online.

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