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ANALYZING THE HUMAN SEX RATIO AT BIRTH

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To my grandmother, Gladys

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Abstract

The effect of income, education, employment, marital status, age, race, birth order, and national economic conditions on the sex ratio at birth were analyzed for the $N = 21{,}597$ children of the National Longitudinal Survey of Youth 1979 participants. These data were analyzed for individual births using a logistic regression model, treating the sex of each child as the outcome variable, and were analyzed for families using a linear regression model, treating the proportion of male children in each family as the outcome variable. No variable was statistically significantly related to the sex ratio. These findings suggest that the sex ratio at birth may not be affected by the individual- and population-level factors commonly examined in past research.

Analyzing the Human Sex Ratio at Birth

Demographers have considered 105 or 106 to be the standards in the human secondary sex ratio, the ratio of male to female live births (Teitelbaum, 1972). Teitelbaum and many other researchers, however, have argued that the sex ratio at birth varies in response to a list of individual-level factors, such as personal income or level of education, and population-level factors, such as wars or economic downturns. Operating behind these factors are several hypotheses that explain the factors' influences on the live birth sex ratio.

The main focus of the current research is to use a nationally-representative dataset to investigate a number of potential factors' influence on the sex ratio at birth and how those factors are related to theories of the sex ratio.

Introduction

This introduction will review (a) the evidence for and against the effects of both individual- and population-level factors on the sex ratio at birth, (b) the theories explaining variation in human sex ratios at birth, (c) potential avenues for future research as they emerge from past research, and (d) issues and problems arising from the use of multilevel models to analyze human sex ratio data.

Individual-level Factors

Age. The relationship between parents' ages and the sex ratio of their children at birth appears to be (a) generally negative, *i.e.* fewer sons with increasing age, and (b) solely influenced by father's age, with mother's age having no effect (Jacobsen, Møller, & Mouritsen, 1999; Moran, Novitski, & Novitski, 1969; Novitski & Kimball, 1958; Novitski & Sandler, 1956; Russell, 1936). As is common in the sex ratio literature, there are some caveats to be noted. Two studies (Almond & Edlund, 2007; Lowe & McKeown, 1950) showed a negative relationship between mother's age and the sex ratio at birth, though father's age was not measured in either study. Because mother's and father's age at birth are correlated, mother's age may have been a proxy for father's age in these latter two studies. Other research (MacMahon & Pugh, 1953) suggested that mother's age may appear to affect the sex ratio of children through its association with other variables, such as birth order. Novitski and Kimball (1958) found that father's age affects the sex ratio at birth differently for different birth orders. For lower birth orders (roughly three or fewer), there is a negative relationship between father's age and sex ratio; this relationship appears to reverse, though, for higher birth orders. Furthermore, it should also be noted that at least one study (Norberg, 2004) found no effect for either parent's age.

Race. In the United States, the sex ratio of Black births has been shown to be consistently lower than the sex ratio in White births (Ciocco, 1938; Matthews & Hamilton, 2005; Visaria, 1967; Winston, 1931). The sex ratios at birth for both groups, however, have generally favored males. For 1917 births, for instance, Ciocco reported sex ratios of 106.2 and 102.8 for White and Black births, respectively. For 2002 births, Matthews and Hamilton reported similar findings—sex ratios of 105.0 and 103.2. Their data also suggested that the sex ratio at birth of Americans of Asian or Pacific Islander descent are generally the highest. In 2002, the sex ratios were 108.9 for Japanese Americans, 107.5 for Native Hawaiians, and 107.0 for Chinese and Filipino Americans (Matthews & Hamilton, 2002). However, in an analysis that included measures of socioeconomic status, Norberg (2004) found no racial differences in sex ratios.

Income/Social Status. Evidence for the effect of income and social status or socioeconomic status (SES) on the sex ratio at birth is mixed. Researchers (Russell, 1936; Winston, 1931) working with British birth data found a positive relationship between SES and sex ratios, *i.e.* more sons with increasing class, with aristocratic families listed in Burke's *Peerage, Baronetage & Knightage* having the highest sex ratio of all classes. A similarly positive relationship can be found in the income-specific, and less class-based, research of Teitelbaum (1972), although he emphasized a curvilinear relationship between income and sex ratios at birth, with a sharp increase seen between lower and moderate incomes and a more gradual increase between moderate and upper incomes. Cameron and Dalerum (2009), in a study of *Forbes* magazine's billionaires, found that the children of male billionaires had higher sex ratios than the general population. However, this effect did not differ for women who were billionaires in their own right.

Other researchers (Ellis & Bonin, 2002; Norberg, 2004; Rostron & James, 1977; Ruckstuhl, Colijn, Amiot, & Vinish, 2010) found no relationship between income and sex ratios. However, the Ellis and Bonin study may have suffered from low power. In their study, income data were collected by asking college students to report their parents' incomes, potentially causing problems related to reliability and range restriction. Ruckstuhl et al. (2010) argued that, although there may be no direct effect of father's income, it may interact with mother's job stress to affect sex ratios.

Occupation. There does not appear to be a general or strong effect of either occupational type or status on the sex ratio at birth (Chacon-Puignau & Jaffe, 1996; Ellis & Bonin, 2002; Fancher, 1956). Much of the research supporting the effect of occupations comes from the work of Bernstein (1948, 1951, 1954). Bernstein (1948) compared the sex ratios of children born to German men listed in *Who is Who in Commerce and Industry* to the entire German population, and found a higher sex ratio for children born to higher class men than the general population. Bernstein (1951, 1954) expanded on this work by classifying men and women listed in *Who's Who in America* and *American Women*, respectively, by the gender-typicality of their occupation into masculine, feminine, and, for men only, neutral categories. Children born to men or women in more masculine occupations were more likely to be male than to men or women in neutral or feminine occupations. It is important to note that Fancher (1956), using a classification scheme similar to Bernstein's, did not replicate these results. Snyder (1961) did, however, demonstrate an effect of occupation in one specific occupation, fighter pilots, where men have a statistically significant greater number of daughters.

Ruckstuhl et al. (2010) examined the effect of mother's occupational stress on the sex ratio at birth, finding that increased occupational stress was related to lower sex ratios at birth. They noted, however, an interaction between mother's occupational stress and father's income. With lower partner incomes, mother's occupational stress had a large, negative effect on sex ratios. With higher incomes, however, the relationship begins to reverse, *i.e.* more sons with increasing job stress.

Education. In two studies (Chacon-Puignau & Jaffe, 1996; Almond & Edlund, 2007), researchers suggested that there is a positive relationship between education and the sex ratio at birth, *i.e.* more sons with increasing levels of education. Almond and Edlund only measured mother's level of education. In the Venezuelan sample from

Chacon-Puignau and Jaffe, this effect is seen in both mother's and father's levels of education. Norberg (2004), however, reported no relationship between levels of education for either parent and sex ratios.

Marital/Cohabitation Status. In three studies (Almond & Edlund, 2007; Chacon-Puignau & Jaffe, 1996; Norberg, 2004), married mothers or mothers cohabiting with a partner had a statistically significantly greater proportion of sons. Chacon-Puignau and Jaffe also found a positive relation between the length of the cohabitation period and the sex ratio at birth, *i.e.* more sons with increasingly longer cohabitation periods.

Birth Order. The effect of birth order on the sex ratio at birth, when it does appear, is universally negative, *i.e.* fewer sons with increasing birth order (Ciocco, 1938; McMahon & Pugh, 1953; Novitski & Kimball, 1958; Novitski & Sandler, 1956; Renkonen, 1964; Russell, 1936; Winston, 1931). Furthermore, this negative relationship appears when mother's (McMahon & Pugh, 1953) and father's age (Novitski & Kimball, 1958; Novitski & Sandler, 1956) are controlled. Other studies (Cann & Cavalli-Sforza, 1968; Jacobsen et al., 1999; Norberg, 2004), however, found no effect for birth order.

Population-level Factors

Although most studies in the literature have either focused solely on individual-level factors or solely on population-level factors, this rigid segregation might not be sound. First, there is evidence that both individual- and population-level factors can potentially affect the sex ratio at birth. Secondly, as parents and the reproductive success of their offspring are affected by both the parents' own conditions and the conditions of the surrounding environment, it is unreasonable to assume that sex ratios are not simultaneously affected by individual- and population-level factors.

Environment. Several environmental events, including air pollution (Lyster, 1974), flooding (Lyster, 1974; Lyster & Bishop, 1965), and earthquakes in Japan (Fukuda, Fukuda, Shimizu, & Møller, 1998) and Iran (Saadat, 2008), have been shown to decrease sex ratios in affected populations. Other researchers (Helle, Helama, & Lertola, 2009) have also found a positive relationship between sex ratios and temperatures in Finland, *i.e.* more sons with warmer temperatures. However, Ciocco (1938), in his state-by-state analysis of sex ratios, argued that the effect of climate or geography on sex ratios is weak, as several similar states—Illinois, Indiana, and Ohio or Massachusetts and Connecticut, for example—have very different sex ratios. However, it should be noted that Ciocco's argument is not exhaustive and that unexamined, state-specific environments could have been driving those differences.

Economics. Evidence appears mixed for the effect of economic conditions on sex ratios. Some research studies in the United States (Ciocco, 1938) and Finland (Helle et al., 2009) found no evidence for an effect of economic conditions on sex ratios at birth. Other studies suggested a relationship between the two in East Germany following unification (Catalano, 2003) and in Sweden (Catalano & Brickner, 2005).

War. The evidence for the effect of war on sex ratios at birth is mixed. Sex ratios increased, *i.e.* more male births, following World War I in most belligerent countries (Bernstein, 1948; James, 2009; Russell, 1936), but not in the United States (Ciocco, 1938). Sex ratios did increase in the United States following World War II and the Korean War (James, 2009). They decreased, however, in Iran following the Iran-

Iraq War and in Slovenia following the Yugoslav civil war (James, 2009). Other wars, including the Russo-Swedish War, the Napoleonic Wars, and the Franco-Prussian War (James, 2009) and the Finnish civil war (Helle et al., 2009), appeared to have no effect whatsoever.

Manmade Disasters. Two studies examining the effect of terrorism on sex ratios found that sex ratios declined following the September 11, 2001 terrorist attacks in California (Catalano, Bruckner, Gould, Eskenazi, & Anderson, 2005) and in New York City (Catalano, Bruckner, Marks, & Eskenazi, 2006). Scherb, Kusmierz, and Voigt (2013) found that sex ratios actually increased in Russia and Cuba, which received the majority of its food imports from Russia in the late 1980s, following radiation exposure from the Chernobyl nuclear plant disaster.

Sex Ratio Theories

There are several theories to explain why sex ratios might be skewed higher (male bias) or lower (female bias) due to individual-level factors. This study will examine two theories—the Trivers-Willard hypothesis and the fixed phenotypes hypothesis.

Trivers-Willard hypothesis. According to the Trivers-Willard hypothesis (Trivers & Willard, 1973) mothers will skew the sex ratio of their offspring in such as a way as to maximize their probability of having grand-offspring. Their argument hinges upon differences in the variability of reproductive success among males and females. If there is greater variability among one sex in terms of reproductive success than there is among the other, mothers in good condition should have more grand-offspring by having more offspring of whichever sex has more reproductive variability. Conversely, mothers in poor condition should have more grand-offspring by having more offspring of whichever sex has less reproductive variability.

In the case of humans, a species in which males have greater variability in reproductive success, mothers in good condition who have more sons than daughters should have more grand-offspring, as the daughters of mothers in good condition should benefit less from inheriting their mothers' traits than their brothers will. The sons of mothers in good condition should be able to increase their reproductive success relative to other males and produce more offspring than their sisters, whose reproductive success will be relatively more similar to other females. The daughters of mothers in poor condition, however, should produce more offspring than their brothers, as females' reproductive success is less variable than their brothers, who would be out-reproduced by other males in better condition.

A hypothetical example that might explain how the Trivers-Willard hypothesis could work in humans—although this example is not always supported by data—would be to consider the mother's social class as defining her condition. In some monogamous societies, wealthy sons have a much larger potential mating pool than their sisters, as sons, unlike daughters, can often marry within their own class or in classes below them. Similarly, in polygynous societies, wealthier men can afford to take on multiple wives, simultaneously increasing their own chances of fathering children and decreasing other men's chances of doing the same.

The Trivers-Willard hypothesis is deceptively simple to understand; correctly testing it is much harder. As Cronk (2007) notes, many studies either create insufficient tests of the Trivers-Willard hypothesis or actually test entirely different sex ratio hypotheses. In animals, a Trivers-Willard effect can be experimentally induced, as researchers have done by clipping female tree swallows' wings (Whittingham, Dunn, & Nooker, 2005) and by feeding seabird couples (Merkling et al., 2012). Cameron, Lemons, Bateman, and Bennett (2008), building upon Cameron's (2004) earlier review of the literature, went further, experimentally inducing a Trivers-Willard effect in mice by changing dietary glucose levels and also providing a potential mechanism in the process.

In humans, testing the Trivers-Willard hypothesis is much more difficult for several reasons, including operationalizing maternal condition. In a study examining differences in how the Roma (Gypsies) and Hungarians treat their children, Bereczkei and Dunbar (1997) noted that the Roma, considered an underclass in Hungary, have a lower sex ratio and spend more time nursing and more money educating their daughters than the Hungarians do. Cagnacci, Renzi, Arangino, Alessandrini, and Volpe (2004), studying mother's weight and weight gain around the time of conception found that low-weight mothers had lower sex ratios than higher-weight mothers and that change in weight was an even better predictor than weight itself. Other studies have found an inverse relationship between the odds of male birth and psychological stress (Obel, Henriksen, Secher, Eskenazi, & Hedegaard, 2007) and sex ratios and women's antidepressant and anxiolytic use (Catalano, Bruckner, Hartig, & Ong, 2005). In a previously mentioned study (Cameron & Dalerum, 2009), it was noted that male billionaires had offspring with higher sex ratios than the general population (60 percent versus 51 percent, respectively). In further analyses to better test for a Trivers-Willard effect, Cameron and Dalerum also found that women married to billionaires had, on average, 1.33 more grandchildren from their sons than from their daughters and that the sons of billionaires were wealthier, on average, than their sisters.

However, it should be noted that not all researchers believe that the Trivers-Willard effect is operating in the United States. Keller, Nesse, and Hofferth (2001), for example, found that parents of lower status do not invest resources, such as time, more in daughters, nor do parents of higher status invest more resources in sons, as Trivers and Willard (1973) predict.

Fixed phenotypes hypothesis. This hypothesis (proposed here, though not represented in the literature) suggests that sex ratios run in families, due either to heritable or environmental causes. The word "fixed" may be somewhat misleading, however, as it may suggest that families, by some means of their biology, are forced to have children of only one sex. Instead, the hypothesis should be concerned with only a differential propensity, inclination, or probability of a woman (or a couple) having children of one sex relative to the other. Put differently, the fixed phenotypes hypothesis would predict some sort of bias within families to have children of one sex or another above and beyond all other factors affecting their children's sex ratios.

There is mixed evidence to suggest that sex ratios do run in families. Edwards (1958) argued that families have different expectations of male births. In a later paper, Edwards (1961) also found a positive correlation (r = .0261) between successive births. However, this correlation only appears in some countries, but not all (Edwards, 1962). Trichopoulos (1967) attempted to determine if the relative sex ratio in the father's and mother's families had any relationship to the sex ratio of their children. A relationship held with the father's family, *i.e.* when a man's family had relatively more males than females, his children were also predominately male. No relationship held with the mother's family. However, other researchers (Jacobsen et al., 1999; Rodgers & Doughty, 2001) have found no evidence for a correlation between successive births, nor any heritable or environmental components to the sex ratio at all. In their study, Rodgers and Doughty (2001) used the NLSY, which this study also will use. However, it should be noted that their data came from an earlier and less complete version of the dataset.

Past Research and this Dissertation

Most of the past research on the factors affecting the sex ratio have focused on single-time, aggregate measures. For example, a typical study might simultaneously measure a family's income at a single time point and the proportion of male children in that family. This type of research, though, is limited in that it can only examine between-family differences. With the exception of Norberg (2004), no studies have included a longitudinal component that would allow a researcher to examine potential within-family differences.

However, Norberg's study suffers from some limitations, likely owing to its timing. As a part of her study, Norberg used participants from the National Longitudinal Survey of Youth's 1979 cohort (NLSY 1979). Her results, though, are conditioned upon the most recent data available to her at the time, which came from 1998, before many of the participants exited their child-bearing years. Thus, birth order (and other time-related) main effects and interactions are not fully realized and represented within her data. The last round of data from the NLSY to be released come from the 2010 interviews, adding 12 more years of fertility data and allowing most participants to complete their fertility histories.

A second limitation can be found in Norberg's analyses. Norberg used a conditional logistic regression approach. In her analyses, data were stratified by birth order, so that all first births are assumed to share their own unique probability of male birth, all second births are assumed to share their own unique probability of male birth, and so forth. These probabilities are not directly estimated, being conditioned out of the model. The effect of the factors thought to affect sex ratios, such as income, could then be estimated. However, a hierarchical linear modeling approach can provide a conceptually cleaner way of considering within-family changes regarding the sex ratios of their children, as individual births can be nested within families.

The Problem of Multilevel Modeling

It was the original goal of this study to analyze the sex of the NLSY respondents' children using a multilevel model. This approach was, at least conceptually, attractive because the nature of families matched the analytic framework. Children are inherently nested within parents in a multilevel hierarchy, i.e. generations. Using a multilevel model would permit the inclusion of within-family correlations concerning the sex of siblings at birth and would help avoid making assumptions about those correlations' existence or, probably more important, their non-existence.

However, and as was feared, the small number of births per family resulted in cluster sizes that were too small to support sound analysis. The average number of children per NLSY respondent was 2.32 children, creating problems in three areas. In the first, the average family size was far below the 50 observations per cluster recommended in Moineddin, Matheson, and Glazier (2007), which could lead to potentially biased random and fixed effects estimates. In the second, the small cluster sizes could also lead to estimation problems for some of the random effects, with SAS being unable to estimate some variance parameters, automatically setting them to zero. In the third, some of the statistical tests could only be completed using data from the larger families, as smaller families lack the degrees of freedom necessary to contribute to those tests. And it would be unsound to assume no difference between small and large families.

In light of these problems, the multilevel approach was dropped. This change, from a multilevel analysis to a classical, non-nested analysis, might create problems if individual's scores within a cluster are not independent, leading to inappropriately estimated standard errors and grossly liberal statistical tests. However, there is evidence to suggest that this might not be an issue with these data. Using NLSY data, Rodgers and Doughty (2001) found no relationship between the sexes of siblings, nor did they find any heritable or shared environmental effects in a behavior genetic analysis of individuals' proportion of children who were male. These findings provide some evidence that births are independent within parents and even within the larger family structures that are included within the NLSY. This suggests that a non-nested approach could be used to analyze this study's data without any risks to its statistical conclusion validity. It may be important to note that assuming uncorrelated sexes among the children within the same family should not, by necessity, lead to null findings between any given variable and the sex ratio, as the Trivers-Willard hypothesis is not really driven by relationships between the sexes of children born within the same family. In fact, it is probably driven by non-random, differences between mothers concerning their probabilities of having male offspring over the scale of a given variable. Or, putting it differently, a Trivers-Willard effect could still occur, even if one child's sex is independent from his or her siblings', as maternal condition only varies between mothers, and it is maternal condition that drives Trivers-Willard effects.

Methods

Participants

Participants for this study were taken from the NLSY79, a national probability sample originally consisting of 12,686 adolescents who were between the ages of 14 and 21 on December 31, 1978, when the household sampling occurred. A few respondents turned 22 before interviews were conducted in spring, 1979, and so the first-round age range is 14-22. To note, the overall sample was originally comprised of three subsamples. The first subsample (N=6,111) was a cross-sectional household probability sample intended to represent children and young adults between the ages of 14 and 21 on December 31, 1978 who were not institutionalized. The second subsample (N=5,295) was an oversample of Hispanic, Black, and economically disadvantaged non-Hispanic-non-Black children and young adults. Follow-up interviews for economically disadvantaged non-Hispanic-non-Black youth were dropped after 1990, however, due to funding limitations. The third subsample (N=1,280) was a sample of young men and women between the ages of 17 and 21 enlisted in active service in one of the nation's four military branches . The number of continuing interviews in this subsample was limited to a random sample of 201 after 1984, due to funding limitations.

The interviews for the NLSY79 were completed every year from 1979 until 1994, with interviews occurring every two years thereafter. The latest available data come from the 2010 round of interviews. The NLSY79, which is funded through the Bureau of Labor Statistics (BLS), collects a large amount of data on work-related variables, including employment, education, and income. The study also collects extensive information on fertility-related variables, including data on each participant's children.

Data

The sex of each biological child born to each NLSY respondent is recorded in the NLSY data files as a dichotomous variable. For this study, the data were recoded (Female = 0, Males = 1), so that all parameters in the logistic regression equation indicate the increased or decreased likelihood of having a male child. The parent's ages at the birth of each child were calculated by subtracting each child's birth date from his or her mother's or father's birth dates. For most children (N=21,554), parent's age could be calculated in years and months. However, the birth months for some children (N=43) were missing, as some parents did not provide the information. In those cases, parent's age could only be calculated in years. Only data from the parent responding to the NLSY interview were included in the analyses.

The race variable that was used is trichotomously defined, with groups representing Black, Hispanic, or Non-Black-Non-Hispanic respondents. From these, two dichotomous variables were created, comparing Black births to the Non-Black-Non-Hispanic births and comparing Hispanic births to the Non-Black-Non-Hispanic births, with the last group serving as a reference group (Non-Black-Non-Hispanic = 0).

Income was measured as annual family net income in U.S. dollars in the survey year closest to the child's birth. The dollar amounts, in all interview rounds, were right truncated. Four different truncation methods have been used since the interviews began. Between 1979 and 1984, any earners with incomes above $75,000 were assigned incomes of $75,001. Between 1985 and 1988, the truncation point was moved to $100,000, with families earning above that amount being assigned incomes of $100,001. Between 1989 and 1996, researchers averaged the net family income of any family earning more than $100,000 and assigned that average to any family falling above the truncation point. Beginning in 1996, a similar averaging procedure was used, with the truncation point being changed to that year's 98th percentile of income earners, instead of the fixed $100,000. To account for changes in the buying power of the U.S. dollar, all dollar amounts were converted to 2012 U.S. dollars using data from the BLS's Consumer Price Index (CPI).

Education was measured as the highest reported grade level completed at the time of the child's birth, with values ranging from 1st grade to 8th year of college or more. Employment was measured as the number of weeks (out of the 40 weeks preceding the birth of the child) in which the parent was employed. Marital status was coded as a dichotomous variable (Unmarried = 0, Married = 1) to ascertain the increased, or decreased, likelihood of having a male child conferred by being married. Birth order was determined by the children's birth dates within each family. Only the biological children of each NLSY respondent was used in birth order determinations.

Population-wide economic data included annual real gross domestic product (GDP) per capita for the entire United States and monthly adjusted unemployment rates for the entire United States. All economic data were obtained from the Federal Reserve Bank of St. Louis Federal Economic Research database.

Analyses

Two sets of analyses were run on the NLSY sex ratio data. In the first, each child's birth was treated as the unit of analysis, with no nested structure imposed on the data. In this analysis, the outcome variable of interest was the sex of the child (Female = 0, Male = 1) and each predictor variable represents the parent's score on the variable at the time of the child's birth. As such, this analysis is essentially the same as the abandoned multilevel logistic model, without the multilevel structure. This analysis was to proceed in a hierarchical regression fashion. First, only an intercept term was included. In the next step, the block of all variables that could vary for children of the same family were entered. This block included all variables except for parent’s sex and race. Finally, a second block of variables that could not vary for children of the same family—parent’s sex and race—were entered. Follow-up analyses were also included to better understand the data structure. This type of analysis does run the risk of violating the independent observations assumption, which could lead to potentially liberally statistical tests. As such, any significant findings should be treated cautiously.

In a second analysis, data were aggregated within families, making families the unit of analysis. The outcome variable of interest here would be the proportion of children in a family who are male, ranging from 0 to 1, and each predictor variable represents the average of the parent's scores across all of their children's birth years on each variable. It should be noted that aggregating across birth order within a family changes that variable’s interpretation. As an average birth order only discriminates among families of different sizes, I opted to include number of children directly to make interpretation simpler. All variables, save for parent's sex and race, were mean centered to make the intercept interpretation more appropriate. This analysis was completed using PROC LOGISTIC in SAS, implementing the procedure's EVENTS/TRIALS coding option for the outcome variable. This coding option reproduces the proportion of male births in a given family by treating a male birth as the event of interest and any birth as a trial in a binomial probability distribution.

Results

Descriptive statistics for the outcome variable, children's sex, can be found by NLSY responding parent's sex (Table 1) and by parent's education (Table 2). Descriptive statistics for the independent variables can be found in Table 3.

The results for the logistic regression analyses, where the log-odds of a male birth were regressed onto the predictor variables with each child as a replication, can be found in Table 1. Eight different models were analyzed. For the first model, only the intercept term was included, which was statistically significant, $b_0 = .0492$, Wald $\chi^2(1) = 13.07$, $p = .0003$. The intercept estimate suggests that male births are more likely than female births, with an odds ratio of male-to-female births of $e^{(.0492)} = 1.05$. This odds ratio, when rescaled by multiplying by 100, also provides an estimate for the sex ratio at birth for the entire NLSY sample, and because it was a probability sample, also an estimate of the population (as the intercept in an intercept-only model is not conditional upon any other predictors). A sex ratio of 105 is considered to be a baseline value by Teitelbaum's (1972) suggestion.

In the second model, the block of predictors that could vary between children from the same parent was added to the model. The intercept and all predictor variables failed to reach statistical significance, except for education, $b_{\text{Education}} = -.0155$, Wald $\chi^2(1) = 4.25$, $p = .0392$. This finding suggests that the odds of a male birth increase by .985 with each year of additional education. This is the opposite effect predicted from most of the past research (Almond & Edlund, 2007; Chacon-Puignau & Jaffe, 1996). However, there are some problems with interpreting the education parameter's statistical significance directly. As stated earlier, the statistical tests may be liberal due to a potential violation of the independent observations assumption. Furthermore, as multiple tests were run, Type I error rates might not be adequately controlled. Implementing a conservative Bonferroni correction would suggest that the education effect is not statistically significant.

In the third model, all statistically non-significant predictors were removed from the model, leaving only education. A second block of predictors, those that could not vary between children from the same parent, including parent sex and race, were also added to the model. No predictors reached statistical significance.

In the fourth model, parent's sex and race were added to the entire first block of predictor variables. This was the equivalent of adding the level-2 predictors from the multilevel model to the level-1 predictors. This was the fullest model analyzed. In this analysis, no parameter estimates reached statistical significance.

In the fifth model, all predictors except for education were dropped by the full model. Only the intercept was statistically significant, $b_0 = .0498$, Wald $\chi^2(1) = 12.12$, $p = .0005$. To examine why education was no longer a statistically significant predictor, the other predictors were each dropped from the model one at a time. When two variables, parent's age and parent's income, were taken out of the model, parent's education remained statistically non-significant.

For the sixth model, parent's education, age, and income were included as predictor variables. On the predictor side, only parent's age, $b_{\text{Age}} = .0070$, Wald $\chi^2(1) = 5.44$, $p = .0197$, and parent's education, $b_{\text{Education}} = -.0185$, Wald $\chi^2(1) = 6.80$, $p = .0091$, were statistically significant. As before, a Bonferroni correction dividing α by the number of tests would result in both of these tests being non-significant.

In the seventh model, parent's income was taken out of the model, with parent's age and education remaining as the only predictors. In this instance, only parent's age was statistically significant, $b_{\text{Age}} = .0059$, Wald $\chi^2(1) = 5.19$, $p = .0227$. In the eighth model, only age was entered as a predictor, but did not quite reach statistical significance, $b_{\text{Age}} = .0040$, Wald $\chi^2(1) = 3.42$, $p = .0645$.

The results for the linear regression analysis, where the proportion of male children was regressed onto the predictor variables, can be found in Table 2. None of the predictor variables were statistically significant. This analysis does not suffer from liberal bias as the previous one did. However, given the absence of significant results, there is no concern over using liberal tests, as a corrected test would result in each test remaining statistically nonsignificant.

Discussion

The results from this study suggest there is no strong evidence for any of the variables' relationships with the sex ratio at birth. Parent's level of education was weakly related to the odds of a male birth when individual births were the unit of analysis, but this relationship disappeared when parent's sex and race were accounted for. There is no evidence for the effect of any variable on the sex ratio when data were aggregated across children within the same family.

This lack of significant findings requires some reconciliation with the past research, and can be somewhat difficult to interpret when considering the past literature. Few studies, if any, report so many negative findings. One possible explanation may be a bias against publishing articles with null results. However, most of the variables examined in this study have been shown to not have a statistically significant relationship with the sex ratio in one or more studies from the literature. For instance, Jacobsen et al. (1999) found a negative relationship between paternal age and odds of a male birth, but they found no relationship with maternal age or birth order. In another study, Norberg (2004) found that parents cohabiting were more likely to have a son than parents who were not, but she found no relationship with parent's age, income, or education. There is evidence for and against the existence of a relationship between the sex ratio and most of the variables examined in this study. As such, this study's complete lack of findings might not be considered anomalous, as the relationships examined appear only inconsistently in the past literature.

It might be helpful, however, to review some of the differences between this study and past ones to determine if any artifacts inherent to either might have lead to the difference in results. A good deal of the earlier sex ratio research (some examples include Novitski & Sandler, 1956; Russell, 1936; Winston, 1931) relied upon national vital statistics for their data. In these studies, data were aggregated across all individuals within the same age range, the same state, the same marital status, or by some other means. For example, Novitski and Sandler examined the proportion of male children born to men in five-year categories: 15 to 19, 20 to 24, and so forth. Ignoring a large number of other variables might make patterns in the sex ratio more apparent, but could also lead to potentially misleading results as many of the variables studied, as well as important variables omitted, are correlated.

Other studies have focused on different subpopulations or used smaller, non-probability samples to collect more selected individual-level data, such as American and Canadian college students (Ellis & Bonin, 2002); births at a hospital in Cambridge, United Kingdom (Ruckstuhl et al., 2010), University of Glasgow medical students (Thomas, 1951), or the Collaborative Study on Cerebral Palsy, Mental Retardation, and Other Neurological and Sensory Disorders of Childhood (Teitelbaum, 1972). It is possible, and should at least be considered, that the findings from these studies hold in the specific subpopulations they examined but are absent in the general population.

The studies that have used large probability samples or censuses also reveal some findings, albeit generally fewer than seen in the other studies. For example, Almond and Edlund (2007) used the United States' National Center for Health Statistics' Vital Statistics Birth Cohort Linked Birth/Infant Death Data set, with roughly 48,000,000 births and found that the sex ratio was positively related to marital status and education and negatively related to age. Chacon-Puignau and Jaffe (1995) used Venezuelan birth registries (N=577,976) from a period spanning several years and found that parents who lived together were more likely to have sons than parents who did not and that education was positively related to the sex ratio. Jacobsen et al. (1999) used the 815,891 birth records from the Danish Fertility Database and found a negative relationship between paternal age and the sex ratio. And Norberg (2004), used five national databases: the National Collaborative Perinatal Project, the Panel Study of Income Dynamics, the National Longitudinal Survey of Young Women, the National Longitudinal Survey of Youth (also used in this study), and the National Survey of Family Growth to obtain records for 86,436 births and found that parents who live together are more likely to have sons than those who do not.

Of the four studies that used large probability samples, three (Almond & Edlund, 2007; Jacobsen et al., 1999; Norberg, 2004) used logistic regression. Two of the studies (Almond & Edlund, 2007; Norberg, 2004) also used datasets from the United States. This suggests that neither the type of analysis used nor the population from which the data came are likely to have played a role in this study's lack of findings. This study, however, had a smaller sample size (N = 21,557), which may help explain my relative lack of findings (though the probability sample status of this sample is important). In samples with N's in the hundreds of thousands, even very small effect sizes will be statistically significant.

As this study was very similar to Norberg's (2004), it may prove useful to discuss differences between the two studies' results. Norberg found that parents who were living together were more likely to have sons than parents who lived apart. This study did not examine if parents lived together or not, focusing instead on marital status. This was partially the result of the NLSY data. Originally, parents were to be separated into three groups: those married at the time of their child's birth, those unmarried but cohabiting at the time of their child's birth, and those neither married nor cohabiting at the time of their child's birth. Determining who was married and who was not was relatively easy using NLSY data. However, determining who was cohabiting and who was not cohabiting among the unmarried respondents was much more difficult and would have led to a large amount of missing data. It was feared that classifying married couples relatively well and classifying unmarried couples very poorly might bias results. As such, respondents were only classified on marital status and not on their living arrangements, and this may have contributed to the difference in findings.

This study benefited from using the NLSY79 as it is a national probability sample. Another important benefit of using the dataset at this time is that the NLSY79 respondents have largely exited their childbearing years, providing a complete record of the respondents' childbirth history. These results are not conditional upon people having children earlier in life, which is an important selection bias in most previous research using the NLSY. However, several limitations should be discussed. As with any longitudinal study, attrition can be a problem. As of 2010, 75.9 percent of the original respondents were still completing interviews (BLS, 2010). It should be noted that retention rates do not vary widely for different races or sexes. Hispanic men had the lowest retention rate (71 percent); black women had the highest (82.1 percent).

A second problem involved a large amount of missing data for the income variable. Approximately 30 percent of the children in the NLSY79 were missing their parent's income data at the time of their birth. These data are likely not missing completely at random, and so some caution should be urged in interpreting this study's income results. However, income has been shown to not be related to the sex ratio in other studies (Rostron & James, 1977; Ruckstuhl et al., 2010), so these results are likely valid.

It should also be noted that abandoning the multilevel model and ignoring potential within-family correlations might lead to inappropriately estimated standard errors for the model's parameters. There was evidence (Rodgers & Doughty, 2001) to suggest that the sex of siblings were independent of one another. However, even if this were not the case, ignoring those within-family correlations should have resulted in liberal statistical tests, which would have made this study's results even more statistically nonsignificant.

This study, in light of Cronk (2007), does not provide the strongest test of the Trivers-Willard hypothesis, by using poor predictors of maternal condition and by not making use of any post-natal measures of parental investment. However, it should be noted that this study's scope was concerned with finding Trivers-Willard effects, *i.e.* patterns consistent with the Trivers-Willard hypothesis but without any concern for how those effects lead to different probabilities for sons and daughters for different mothers. This was done as an initial step in better understanding human sex ratios at birth. As previously stated, though, no variables were statistically significant predictors of the sex ratio in this study and, therefore, showed no Trivers-Willard effect. This is likely due to three reasons. Data were collected around the time of each child's birth, and Cameron's (2004) review suggests that it is better to measure each variable as close to conception as possible. Furthermore, many of the variables measured here are relatively new to the human experience, *e.g.* education or national unemployment rate, and not likely to have the evolutionary history or importance of other variables, *e.g.* maternal diet or weight. And finally, many of the variables used, such as national unemployment rate, likely have little direct relationship to the reproductive success of one's children (Cronk,

2007) and so have little bearing on mother's ability to increase their probabilities of having grandchildren.

Establishing a line of future research concerning the Trivers-Willard hypothesis and the human sex ratio at birth is simultaneously somewhat easy and somewhat difficult. Newer research, completed within the last two decades, is increasingly suggesting that many of the variables once thought to be related to the sex ratio, in fact, are likely not. Over time, fewer and fewer variables remain to be discussed (which could help researchers focus in on a few key factors, if such factors do indeed exist). The most promising candidates from the past research is probably the mother's cohabitation status, stress levels, and diet.

However, even the mother's cohabitation status may just be a proxy variable. And this is where future research will become more difficult. Related to mother's cohabitation status and probably most other factors that have been examined in the past, is maternal stress, which is likely to have a much more direct effect on the sex ratio and is consistent with the Trivers-Willard hypothesis. Maternal stress has rarely been directly studied. Hansen, Møller, and Olsen (1999) found that women who went through severe life events, such as the death of a partner or a partner's cancer diagnosis, before their second trimester were much less likely to have a son than women who did not. And in their study concerning women's occupational stress, Ruckstuhl et al. (2010) found that women who worked in occupations considered more stressful were less likely to have sons than women working in less stressful occupations. Obel and colleague's (2007) study comes much closer in terms of directly studying stress levels, finding that having more psychological stress decreases women's probability of having a son.

The best future research studies will begin to incorporate measures of maternal stress and nutrition during pregnancy into their models. Stress measures could be physiological in nature, but they do not necessarily need to be. Proxy variables focused on maternal stress, such as losing or not having a partner, job loss, losing a home or residence, or a clinical diagnosis of malnutrition are all possible, if potentially weak, avenues to follow. Another potential line of research would involve using subjective measures of maternal stress, including asking mothers about their level of general stress, their level of job insecurity, their level of relationship insecurity, and their level of difficulty obtaining basic resources, including food and shelter. It is possible that whatever evolutionary mechanism is behind the Trivers-Willard hypothesis would be sensitive to women picking up on stress-related cues in their environment and that stressed women would increase their number of grandchildren by having daughters, as opposed to sons. At any rate, it should be strongly noted that the additional emphasis on finding and differentiating between mothers in good and bad situations is no accident, as that is the comparison from which the Trivers-Willard effect is most likely to appear and the comparison that most past research, this study included, has likely not made.

Changes in maternal diet or weight also provide excellent avenues for research as both have been shown to be related to the probability of male births, as Cameron et al. (2008) experimentally induced by changing glucose levels in mice and as Cagnacci et al. (2004) found by examining women's changing weights during pregnancy. Studies using dietary or weight changes also have the added benefit of, in the presence of statistically significant hypothesis tests, providing us with more than just a Trivers-Willard effect, which already abound in the literature. These studies can also provide us with specific mechanisms that help explain how the Trivers-Willard effect occurs and potentially how it developed.

In conclusion, the current study used a national probability sample and relevant statistical methods to test for a number of links identified in previous literature as relating to human sex ratio. None were statistically stable, a finding that raises doubts about certain previous studies in the sex ratio literature, and that reinforces the idiosyncratic and inconsistent nature of much of that literature.

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Table 1

Children's Sex, Overall and by Parent's Sex

	Overall		NLSY79 Mothers		NLSY79 Fathers	
Sex	*N*	Percent	*N*	Percent	*N*	Percent
Females	10,533	48.77	5,628	48.93	4,905	48.58
Males	11,064	51.23	5,873	51.07	5,191	51.42
Total	21,597	100.00	11,501	100.00	10,096	100.00

Table 2

Children's Sex by Parent's Last Year of Education

Year	Males	Females	Proportion Male
0	4	2	0.67
1	4	3	0.57
2	.	.	.
3	17	18	0.49
4	26	12	0.68
5	29	21	0.58
6	66	62	0.52
7	131	106	0.55
8	345	340	0.5
9	622	521	0.54
10	816	815	0.5
11	1,006	969	0.51
12	3,744	3,638	0.51
13	770	724	0.52
14	714	655	0.52
15	347	319	0.52
16	975	897	0.52
17	182	203	0.47
18	183	189	0.49
19	86	73	0.54
20	99	92	0.52
Missing	898	874	0.51

Table 3

Independent Variable Descriptive Statistics

Variable	N	Mean	Standard Deviation	Skew
Birth order	21,597	0.98	1.17	1.70
Parent's age	21,597	26.80	6.25	0.56
GDP Growth	21,597	2.01	2.14	-0.60
US unemployment rate	21,554	6.73	1.47	0.69
Income Log_{10}	16,204	4.56	0.65	-4.21
Education	19,825	12.30	2.60	0.23
Personal Employment	20,001	26.31	16.65	-0.69

Table 4

Logistic Regression Analyses of Sex for Individual Child Data (Male = 1)

	Model 1		
Parameter	*B*	Odds Ratio	Percent Change in Probability of Male Birth
Intercept	0.0492*	1.05	
Birth order			
Age			
GDP growth			
Unemployment rate			
Education			
Employment			
Marital status			
Log_{10} Income			
Parent's Sex			
Race (B v. NHB)[a]			
Race (H v. NHB)[a]			
-2 Log Likelihood	29,926.742		

* *p*< .05

a B = Black, H = Hispanic, NHB = Non-Black, Non-Hispanic

Table 4

Logistic Regression Analyses of Sex for Individual Child Data (Male = 1)

Parameter	Model 2		
	B	Odds Ratio	Percent Change in Probability of Male Birth
Intercept	0.0211		
Birth order	0.0154	1.016	0.79%
Age	0.0046	1.005	0.25%
GDP growth	-0.0058	0.994	-0.30%
Unemployment rate	-0.0109	0.989	-0.56%
Education	-0.0155*	0.985	-0.76%
Employment	-0.0001	1.000	0.00%
Marital status	-0.0217	0.958	-2.19%
Log_{10} income	0.0220	1.022	1.08%
Parent sex (M v. F)			
Race (B v. NHB)			
Race (H v. NHB)			
-2 Log Likelihood	21,267.147		
Hosmer-Lemeshow $\chi^2(8) = 9.0625$, p = .3371			

*p<.05

Table 4

Logistic Regression Analyses of Sex for Individual Child Data (Male = 1)

	Model 3		
Parameter	*B*	Odds Ratio	Percent Change in Probability of Male Birth
Intercept	0.0487*		
Birth order			
Age			
GDP growth			
Unemployment rate			
Education	-0.0049	0.995	-0.25%
Employment			
Marital status			
Log_{10} Income			
Parent sex (M v. F)	-0.0111	0.978	-1.12%
Race (B v. NHB)	0.0139	1.006	0.30%
Race (H v. NHB)	-0.0216	0.971	-1.49%
-2 Log Likelihood	27,467.973		
Hosmer-Lemeshow $\chi 2(8) = 3.9819$, p = .8588			

* $p < .05$

Table 4

Logistic Regression Analyses of Sex for Individual Child Data (Male = 1)

	Model 4		
Parameter	*B*	Odds Ratio	Percent Change in Probability of Male Birth
Intercept	0.0211		
Birth order	0.0178	1.018	0.88%
Age	0.0041	1.004	0.20%
GDP growth	-0.0058	0.994	-0.30%
Unemployment	-0.0112	0.989	-0.56%
Education	-0.0142	0.986	-0.71%
Employment	-0.0004	1.000	0.00%
Marital status	-0.0137	0.973	-1.39%
Log_{10} Income	0.0213	1.022	1.08%
Parent sex (M v. F)	-0.0108	0.979	-1.07%
Race (B v. NHB)	0.0128	0.996	-0.20%
Race (H v. NHB)	-0.0299	0.954	-2.41%
-2 Log Likelihood	21,265.681		

Hosmer-Lemeshow $\chi^2(8) = 14.1073$, $p = .0790$

Table 4

Logistic Regression Analyses of Sex for Individual Child Data (Male = 1)

	Model 5		
Parameter	*B*	Odds Ratio	Percent Change in Probability of Male Birth
Intercept	0.0498*		
Birth order			
Age			
GDP growth			
Unemployment rate			
Education	-0.0027	0.997	-0.15%
Employment			
Marital status			
Log_{10} Income			
Parent sex (M v. F)			
Race (B v. NHB)			
Race (H v. NHB)			
-2 Log Likelihood	26,777.921		
Hosmer-Lemeshow $\chi^2(6) = 6.2207$, $p = .3989$			

* $p < .05$

Table 4

Logistic Regression Analyses of Sex for Individual Child Data (Male = 1)

Parameter	**Model 6** *B*	Odds Ratio	Percent Change in Probability of Male
Intercept	0.0478*		
Birth order			
Age	0.0070*	1.007	0.35%
GDP growth			
Unemployment rate			
Education	-0.0185*	0.982	-0.92%
Employment			
Marital status			
Log_{10} Income	0.0315	1.032	1.55%
Parent sex (M v. F)			
Race (B v. NHB)			
Race (H v. NHB)			
-2 Log Likelihood	21,761.448		
Hosmer-Lemeshow $\chi^2(8) = 11.0466$, $p = .1991$			

* $p < .05$

Table 4

Logistic Regression Analyses of Sex for Individual Child Data (Male = 1)

	Model 7		
Parameter	*B*	Odds Ratio	Percent Change in Probability of Male Birth
Intercept	0.0468*		
Birth order			
Age	0.0059*	1.006	0.30%
GDP growth			
Unemployment rate			
Education	-0.0098	0.990	-0.51%
Employment			
Marital status			
Log_{10} Income			
Parent sex (M v. F)			
Race (B v. NHB)			
Race (H v. NHB)			
-2 Log Likelihood	27,464.470		
Hosmer-Lemeshow $\chi^2(8) = 6.9444$, $p = .5426$			

* $p < .05$

Table 4

Logistic Regression Analyses of Sex for Individual Child Data (Male = 1)

	Model 8		
Parameter	*B*	Odds Ratio	Percent Change in Probability of Male Birth
Intercept	-0.0587		
Birth order			
Age	0.0040	1.004	0.20%
GDP growth			
Unemployment rate			
Education			
Employment			
Marital status			
Log_{10} Income			
Parent sex (M v. F)			
Race (B v. NHB)			
Race (H v. NHB)			
-2 Log Likelihood	29,923.321		
Hosmer-Lemeshow $\chi^2(8) = 9.6041$, $p = .2939$			

Table 5

Logistic Regression Analyses of Proportion of Male Births within Families

Parameter	*B*	Odds Ratio	Percent Change in Probability of Male Birth
Intercept	0.0715		
Family size	0.0040	1.004	0.20%
Age	0.0020	1.002	0.10%
GDP growth	-0.0146	0.985	-0.76%
Unemployment rate	-0.0107	0.989	-0.56%
Education	-0.0051	0.995	-0.25%
Employment	-0.0015	0.999	-0.05%
Marital status	0.0055	1.006	0.30%
Log_{10} Income	0.0224	1.023	1.12%
Parent's Sex	-0.0167	0.967	-1.71%
Race (B v. NHB)	0.0051	0.993	-0.35%
Race (H v. NHB)	-0.0171	0.971	-1.49%
-2 Log Likelihood	26,965.884		

Hosmer-Lemeshow $\chi^2(8) = 1.4167$, $p = .9940$