NBER WORKING PAPER SERIES

PERINATAL HEALTH AMONG 1 MILLION CHINESE-AMERICANS

Douglas Almond
Yi Cheng

Working Paper 27775
http://www.nber.org/papers/w27775

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
September 2020

We thank Janet Currie, Lena Edlund, Suresh Naidu, Cristian Pop-Eleches, Shuang Zhang, Lowell Taylor, and seminar participants at NBER's Education and Children program meeting, Stockholm University's SOFI Institute, Columbia University, UC Santa Barbara, Uppsala University, Pompeu Fabra, Carnegie Mellon University, UVA, American University, NYU-Shanghai, the Federal Reserve Bank of Minneapolis, and Dartmouth College for helpful comments. We thank Jean Roth (NBER) for assistance obtaining the vital statistics and hospital discharge microdata from the Healthcare Cost and Utilization Project (HCUP) and Solveig Asplund for proofreading. We analyze the State Inpatient Databases (SID) provided to HCUP by New York State. The National Association for Public Health Statistics and Information Systems (NAPHSIS) provided mortality micro-data with additional geographic identifiers. The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

NBER working papers are circulated for discussion and comment purposes. They have not been peer-reviewed or been subject to the review by the NBER Board of Directors that accompanies official NBER publications.

© 2020 by Douglas Almond and Yi Cheng. All rights reserved. Short sections of text, not to exceed two paragraphs, may be quoted without explicit permission provided that full credit, including © notice, is given to the source.

Perinatal Health among 1 Million Chinese-Americans
Douglas Almond and Yi Cheng
NBER Working Paper No. 27775
September 2020
JEL No. I1,I12,J13,J16

ABSTRACT

The literature on "missing girls" suggests a net preference for sons both in China and among Chinese immigrants to the West. Perhaps surprisingly, we find that newborn Chinese-American girls are treated more intensively in US hospitals: they are kept longer following delivery, have more medical procedures performed, and have more hospital charges than predicted (by the non-Chinese gender difference). What might explain more aggressive medical treatment? We posit that hospitals are responding to worse health at birth of Chinese-American girls. We document higher rates of low birth weight, congenital anomalies, maternal hypertension, and lower APGAR scores among Chinese Americans girls – outcomes recorded prior to intensive neonatal medical care and relative to the non-Chinese gender gap. To the best of our knowledge, we are the first to find that son preference may also compromise "survivor" health at birth. On net, compromised newborn health seems to outweigh the benefit of more aggressive neonatal hospital care for girls. Relative to non-Chinese gender differences, death on the first day of life and in the post-neonatal period is more common among Chinese-American girls, i.e. later than sex selection is typically believed to occur.

Douglas Almond
Department of Economics
Columbia University
International Affairs Building, MC 3308
420 West 118th Street
New York, NY 10027
and NBER
da2152@columbia.edu

Yi Cheng
Department of Economics
Columbia University
International Affairs Building, MC 3308
420 West 118th Street
New York, NY 10027
yc2969@columbia.edu

1 Introduction

It is well known that sex selection occurs not only in Asia but among Asian immigrants to the West. If parents vary in their degree of son preference, parents having daughters after sex selection becomes routine are likely to be those parents with milder son preference. This compositional effect could lead to an increase in parental investments in girls [Goodkind, 1996, Hu and Schlosser, 2015, Anukriti et al., 2018]. On the other hand, prenatal sex determination allows discrimination against girls to commence before birth, which has been found for parent-reported investments in girls in Asia [Bharadwaj and Lakdawala, 2013]. These competing forces mean that the health of those girls who "survive" prenatal sex selection could either improve or deteriorate, rendering net health impacts an empirical question. To date, the literature has been more focussed on investment measures reported in surveys than on health itself, presumably because of the scarcity of administrative health microdata in Asia.

Immigrants to the West offer the advantage of being observed in universal, administrative data on health [Lhila and Simon, 2008]. Nevertheless, no work to date has shown that Asian girls are systematically in either better or worse early childhood health following the increase in sex selection. Rates of sex selection are lower among Asian immigrants to the West than in Asia, which *per se* would reduce compositional effects that might benefit surviving girls.[1] The seminal study by Lhila and Simon [2008] found no evidence of gender bias in prenatal investments among sex selection "survivors" in the US. Should compromised early childhood health be established for Asian girls in the West, this might be a particular concern given the long-term developmental effects that arise from differences in early childhood health [Barker, 1992]. Furthermore, compromised female health would be a "modern phenomenon" appearing after

[1]We address the question of "survivor bias" in the discussion of Table 3 results. In sum, the relatively modest elevation of sex ratios compared to China circumscribes the compositional effect in the US, especially for first-parity births (sex ratio of 1.07).

the diffusion of prenatal sex determination in the 1980s and thus impacting cohorts reaching adulthood today.

Here we consider the perinatal health and neonatal hospital care of Chinese-Americans, roughly 1.2% of the US population.[2] Son preference has been extensively documented for China, e.g. Sen [1992], Yi et al. [1993], Edlund [1999], DasGupta et al. [2003], Chen et al. [2013]. High sex ratios have also been observed among Asian immigrants to the West [Dubuc and Coleman, 2007, Lhila and Simon, 2008, Almond and Edlund, 2008, Abrevaya, 2009, Almond, Edlund, and Milligan, 2013]. We start by confirming that sex ratios at birth are indeed elevated for Chinese Americans in US natality data, indicating that an average son preference persists. At these magnitudes, roughly 97% of Chinese-American girls were not sex selected prenatally. We turn our attention to female "survivors" and their health.

Although newborn boys offer a ready counterfactual, they are substantially less healthy biologically than girls [Kraemer, 2000]. Rather than expecting equality, the gender benchmark is nuanced: how large a female heath advantage should we see? In considering neonatal care provided by hospitals in the context of better average health of girls, how much less intensive should we expect medical treatment to be?

To account for baseline gender differences in outcomes and medical treatments, we analyze perinatal health and neonatal care in a standard difference-in-differences framework. For binary outcomes, we also show logit specifications which allow for non-additive effects of race and gender. As a placebo check, we also consider whether Chinese girls appear disadvantaged in the US in the years prior to ultrasound diffusion (compared to the pre-ultrasound gender gap among Whites). Were Whites providing a misleading estimate of the gender gap, we should see "impacts" in the the pre-ultrasound period as well. Second, Filipinos in the West may have less son preference and share a

[2]The World Health Organization defines the perinatal period as spanning 22 weeks gestation to 7 days after birth.

"broadly Asian genetic" with Chinese [Almond, Edlund, and Milligan, 2013].[3] While we lose precision (particularly for mortality), we show a similar basic pattern of health results when we compare Chinese Americans to Filipino Americans. We also quantify gender gaps among White and Filipino in our natality sample and overlay these differences in Figure 1. This pattern of similar gender gaps conditional on the specific health outcome considered supports our identifying assumption that biological drivers of gender gaps are similar across race.[4]

Across a wide range of health measures, we find that Chinese American girls are in compromised health. Morbidity outcomes are worse for girls, including lower APGAR score, more congenital anomalies,[5] and a greater likelihood of low birth weight. Stillbirth is significantly more common for Chinese American girls than we should expect. The female health disadvantage tends to be larger in Chinese American families where the mother has less education and where both mother and father are Chinese. "First generation" immigrant parents make up about 90% of our sample and drive the female health disadvantage.

Turning to hospital discharge records for New York, Chinese girls tend to stay longer in the hospital postnatally, have higher charges, and receive more hospital procedures, consistent with medical care attempting to counteract a health disadvantage clinicians observe at birth. Nevertheless, death on the first day of life and childhood deaths from age 1 month to 12 years are more common among girls than expected, particularly for disease-related causes.[6] Following discharge, Chinese American girls are less likely to be brought back

[3]Meyer [1995] also advocates showing results that are benchmarked against multiple control groups.

[4]If the boy-girl differences in perinatal health are race specific, we would expect to see large horizontal distances between gender gaps among White and Filipino. However, Figure 1 shows the opposite: the magnitudes of gender differences are similar for Whites. For example, the gender gap for 5-minute APGAR score is around .025 for both Whites and Filipinos and the gender gap for low birthweight is around .007.

[5]Congenital anomalies result from both the prenatal environment and genetic causes [Weinhold, 2009, Sarmah et al., 2016]. "Congenital" comes from the latin *congenitus*, meaning present from birth.

[6]Deaths not attributed to accidents, suicide, homicide, or other "external" causes.

to the hospital. This re-admission gap grows when we factor in their worse health at birth. To our knowledge, ours is the first evidence that Chinese girls begin life in compromised health.

Our paper is organized as follows. First, we provide a brief literature review (Section 2) and describe our empirical methodology (Section 3) which starts with and builds on a conventional difference-in-differences model. We describe the data (Section 4), present our findings (Section 5), and conclude with a discussion (Section 6).

2 Literature Review

Four literatures are most relevant to our study.

First, that girls have better health than boys has been extensively documented in public health. Women are "hardier than men and, given similar care, survive better at all ages – including in utero" [Sen, 1992]. For example, boys in the US are 21% more likely to die in the first year of life than girls [Xu et al., 2016]. Kraemer [2000] argues that the male is more "fragile" during prenatal and neonatal periods. The sex ratio decreases from the initial excess of around 120 male conceptions per female to 105 at birth. Perinatal brain damage, congenital deformities, and stillbirth are also more common in boys.

Second, son preference exists in China and persists among some Chinese immigrants to the West. (These studies focus on the "extensive" margin and do not consider the health of sex selection "survivors".[7]) Empirical evidence for this preference comes chiefly from "missing girls" [Sen, 1992], which typically follow an older sister(s) [Yi et al., 1993, Duflo, 2012]. Indeed, a demographic regularity is that sex selection is muted for the first child, but becomes pronounced at higher birth order children. Almond and Edlund [2008] document

[7]We are aware of two exceptions. Gonzlez [2018] considered both sex selection and early childhood health among Indian immigrants to Spain, finding no evidence of a excess gap in infant health. A working paper by Muchomba and Chatterji [2020] considers gender gaps in disability among Asian immigrants to the US.

male-biased sex ratios among US-born children of Chinese, Korean, and Asian Indian parents in a 5% sample of the 2000 US Census. These heightened sex ratios appear at higher birth orders if there was no previous (elder) son. Almond, Edlund, and Milligan [2013] document a similar pattern for Asians in Canadian Census data, and argue that Filipinos (along with Whites) provide a suitable control group because of similar biology and the relative absence of son preference.[8,9]

Third, a smaller literature has analyzed prenatal *investment* measures (*cf.* administrative health outcomes data) in the context of son preference. Bharadwaj and Lakdawala [2013] document preferential treatment of boys *in utero* in India following ultrasound, including more frequent prenatal visits and vaccinations. Their context is one of substantially stronger son preference as reflected by a sex ratio of 1.22 versus 1.08 among Chinese-Americans (and a biological benchmark of 1.05-1.06). Our analysis can be viewed as assessing whether female "survivors" exhibit worse *health* – as opposed to investments – in a less extreme context. We do not think Bharadwaj and Lakdawala [2013]'s finding tells us what we would find for Chinese American health, particularly as Hu and Schlosser [2015], Anukriti et al. [2018] reach the opposite conclusion for India: surviving girls do better thanks to ultrasound (see below).

A seminal study by Lhila and Simon [2008] considers Asian immigrants to the US. First, they document that sex ratios increase at higher parities for Chinese and Asian Indian births in the US. Leveraging information from US birth certificates on reported ultrasound usage, they find no gender bias in prenatal investments. As Lhila and Simon [2008] acknowledge, reporting an ultrasound procedure is an imperfect indicator for knowing fetal gender, in part because reported ultrasound use may include obstetrical ultrasounds that

[8]Among Chinese in Canada, the sex ratio of the third child was 1.38 following 2 elder daughters [Almond, Edlund, and Milligan, 2013, Endnote #13].

[9]Economists have modeled son preference in Asia as child gender providing utility directly to parents [Edlund, 1999]. If sex-selective abortion is costly (including its psychological costs), some parents who prefer sons will continue to have daughters even if prenatal sex determination is relatively inexpensive [Almond, Li, and Zhang, 2019].

occur at delivery (too late for prenatal investments to respond). Additionally, parents may be less likely to report truthfully having received an ultrasound if it was motivated by son preference (a second source of measurement error, and one highlighted by Bharadwaj and Lakdawala [2013] for India where prenatal sex determination is illegal). Similar concerns may exist for other self-reported investment measures, such as the number of parental care visits. Third, requesting an ultrasound may be endogenous to son preference, and even absent measurement error may introduce bias to regression parameters from endogenous control [Angrist and Pischke, 2009].[10]

As we do not have an instrumental variable which might address these econometric issues,[11] we ignore individual-level information on reported ultrasound use and focus instead on administrative and clinical measures of health that are not reported by parents.[12] In the absence of a natural experiment, our results may be viewed as descriptive. That said, we will take a stronger stand insofar as the analysis of firstborn children are concerned, as their sex ratio is relatively normal.

Finally, we note that in the handful of studies when health outcomes instead of investments have been considered, ultrasound has been found to improve girls' outcomes. The increase in ultrasound availability in India led to substantial improvements in the relative survival, i.e. under-5 mortality, of girls after birth [Anukriti et al., 2018]. Hu and Schlosser [2015] show that an increase in the practice of prenatal sex selection, approximated by sex ratios at birth, leads to a reduction in the prevalence of malnutrition among surviving girls in India. Consistent with the compositional effect posited by Goodkind

[10] Lhila and Simon [2008] acknowledge: "The identification of the effect of knowing fetal gender rests on the assumption that ultrasound receipt is not correlated with unobserved factors, specifically gender preference" and "Admittedly, measurement error may have biased our results downward to the point that we fail to capture the true effect of knowing fetal gender on prenatal investments."

[11] In annual American Hospital Association data on facilities, ultrasound machines that would be used for prenatal exams are not distinguished from other ultrasound equipment, *cf.* Chen et al. [2013].

[12] By ignoring the parent-reported health investment measures, we depart from both Bharadwaj and Lakdawala [2013] and Lhila and Simon [2008].

[1996], these findings highlight the importance of considering the magnitude of sex selection when analyzing gender gaps in health.

3 Methodology

We begin with a standard difference-in-difference (DD) comparison:

$$y_i = \gamma + \gamma_c \cdot Chinese_i + \theta \cdot Female_i + \theta_c \cdot Female_i \cdot Chinese_i + \varepsilon_i. \quad (1)$$

Estimates of γ, γ_c, and θ are reported in Tables 2 and 5. The coefficient of interest, θ_c, captures the additional gender difference among Chinese-Americans, and is reported tables 2, 4, and 5

Assuming the gender gap among Whites (θ) is biological, we might worry that differences in average birth outcomes by race complicate the (additive) extrapolation of θ to Chinese. For example, because congenital anomalies are about half as common among Chinese as among Whites, the White female advantage θ might overstate the expected biological gender difference among Chinese, making θ_c an overestimate of gender discrimination among Chinese. The magnitude (and even sign) of this DD "mistake" will differ depending on the particular health outcome considered (we consider sixteen different outcomes y_i).

The simplest way to address this potential issue is to estimate logit models for binary outcome variables. For non-binary outcomes, we could log the dependent variable. As shown in Appendices A and B, logit and logging avoid this potential DD "mistake" by adopting a multiplicative relationship. We note the issue being addressed here is distinct from the conventional motivation for a logit model over a linear probability model. In Tables 2, 4, and 5 we report both with conventional DD (linear probability model) results alongside logit results.

That said, it is not necessary to adopt a logit model (or log the dependent variable) to address this issue. In Table 6, we present "calculated effects" which

build more directly on equation (1). For these, we assume θ is a scalar of the White male average: $\theta = \alpha \cdot \gamma$ (with different α's for different health outcomes). Carrying this assumption to Chinese implies θ_c can be decomposed into two parts: a biological part that depends on race ($\alpha \cdot \gamma_c$) and an unexplained residual, which we denote Θ[13] and refer to it as the "calculated effect". That is:

$$\theta_C = \alpha \cdot \gamma_C + \Theta. \tag{2}$$

If this simple multiplicative model is correct, $\alpha \cdot \gamma_c$ is the implied adjustment to the ordinary DD estimate θ_c. This adjustment is larger for outcomes with: i) larger biological gender gaps α; ii) larger racial differences. Under the above multiplicative assumption, the conventional DD estimate θ_c remains "correct" (i.e. interpretable as discrimination) if biological gender differences *or* racial differences are 0.[14] Estimates of Θ using the above procedure are equivalent to those obtained if one normalized White outcomes by the White male mean and Chinese outcomes by the Chinese male mean, and then regressed these on the Chinese and gender dummy and their interaction.[15,16]

This adjusted DD model is appropriate for both binary and continuous outcome measures and returns estimates in absolute terms (units of the untransformed dependent variable) and can viewed as a decomposition of *any* DD model estimated in levels. Nevertheless, we focus on the unadjusted DD model and logit estimates due to their familiarity. Furthermore, our results

[13]Standard errors are estimated using the δ-method.

[14]We construct a matched control group in an attempt to reduce racial differences in male average outcomes using a subsample of White babies based on geographic information (mother's state and county of residence) and maternal characteristics (marital status, education, and age). We obtain similar effect estimates without applying our (gratuitous) adjustment to DD estimates. That said, for most outcomes it is difficult to find White subgroups with comparably good outcomes as Chinese using the available control variables. Furthermore, we only conduct this exercise for 1989-2002, when the more precise geocodes are available in the natality data.

[15]With the caveat that normalization returns coefficients interpreted as percents. See Appendix Section C for details.

[16]Both approaches implicitly allow for Chinese girls to be discriminated against. If one thought instead Chinese boys are discriminated for, then one would "correct" male outcomes and normalize by female means. This alternative normalization yields similar estimates, see Appendix Table S6.

are similar across the two sets of approaches.

4 Data

We analyze births occurring in the US using US Vital Statistics microdata from the National Center for Health Statistics (NCHS) and New York (NY) state hospital discharge data (1993-2013) from the Healthcare Cost and Utilization Project (HCUP) State Inpatient Databases (SID). We combine fetal death files (1989-2005)[17] and mortality files (1989-2015) to assess the gender composition of mortality. We use natality files (1989-2013) for morbidity outcomes and linked birth/infant death files (1989-1991, 1995-2010)[18] for mortality. One limitation of our data is that we cannot link natality data to hospital discharge records. Nor can we match siblings in any dataset.

Maternal race is reported in the natality, linked birth/infant death, and fetal death data. Mortality data and hospital discharge records only report race of the decedent (patient). We use race of mother where available, and the child's race where it is unavailable.[19]

We include Chinese, Whites,[20] and Filipinos in our analysis. We exclude: *i)* multiple births in natality, linked birth/infant death, and NY hospital discharge data, for whom birth order has a distinct interpretation; *ii)* foreign resident mothers in US natality and linked birth/infant death data, so as to remove "birth tourists" who might base their travel decision in part on knowl-

[17] Chinese is not separately identified from other Asian races in fetal death data after 2005.

[18] NCHS did not produce Linked Birth/Infant Death Files for 1992-1994. No decedent age is reported after 2010.

[19] NCHS likewise adopted this race assignment rule in summary statistics beginning in 1989.

[20] We take a 10% random sample of White births to make the sample size more manageable for analysis.

edge of fetal gender;[21] and iii) mortality records for foreign-born decedents.[22] Because no subcategories of "Asian" race are reported in the hospital discharge data, we omit the Filipino group and "isolate" Chinese by including only Asian residents in zip codes with a Chinese to Asian ratio above 0.5.[23] Our final analysis sample consists of: a) 7,649,124 White, 822,825 Chinese, and 785,189 Filipino births in US Vital Statistics data, and b) 2,415,040 White and 136,828 "Chinese" (out of 282,147 Asian) births in NY hospital discharge data.

5 Results

We organize results by health outcome: death from mortality microdata (Table 1), perinatal outcomes from birth certificates (Table 2, 3, and 4), and hospital discharge outcomes (Table 5). We take logit estimates as our preferred measure as they address the issue that arises from differences in average outcomes by race. To facilitate the interpretation of effect magnitudes, we provide estimates in absolute terms (units of the un-transformed dependent variable), i.e. our "calculated effects" introduced in Section 3, in Table 6. Across the three datasets, we see Chinese females appear to be at a disadvantage. First, we note that the sex ratio at birth among Chinese in the US is 1.08, versus 1.05 for Whites (Appendix Table S1).

[21] We are unaware of any previous analysis of this question. Foreign resident mothers who give birth in the US are a small fraction of total US births: .17% for Whites, .09% for Chinese, and .05% for Filipino mothers. White non-resident births are 0.4 percentage point more likely to be male (not statistically significant), Chinese non-resident births are 2.8 percentage points more likely to be male (not statistically significant), and Filipino non-residents 3.1 percentage points *less* likely to be male (not statistically significant) when compared with the fraction of male births among US-resident mothers (of the same race). Future research might consider the sex ratio of additional non-resident subpopulations giving birth in the US.

[22] Because gender-biased immigration of children could also confound our analysis.

[23] California also has a large population of Chinese immigrants. However, California's hospital discharge data do not report patients' zip code, so we are unable to narrow California's "Asian" patients toward Chinese.

5.1 Death certificate microdata

Table 1 reports the fraction of male deaths during the late fetal (i.e. stillbirth), neonatal, and post-neonatal periods. Compared to US Whites, the fraction male among Chinese American decedents is 2-4 percentage points lower.[24] Among 1,603 fetal deaths, the fetus is 3.8 percentage points less likely to be male among Chinese (significant at the 1% level) than the corresponding White gender difference. The gap in overall neonatal death (-.021, not reported) is driven by death on the first day of life: the decedent is 4.1 percentage points less likely to be male in their first 24 hours of life among 838 Chinese deaths (significant at the 5% level) compared to the White gender gap. We define post-neonatal childhood deaths as running from 1 month to 12 years of age (1,587 Chinese deaths). The gender gap in post-neonatal death is driven by 1,289 disease-related deaths.[25]

Our finding that excess female mortality occurs after birth is broadly consistent with Anderson and Ray [2010], who argue that in developing countries, "missing women" manifest later in life than commonly believed. However, for China this gender gap emerges early in life and reappears at substantially older ages, driven by cardiovascular and respiratory diseases. Thus, *postnatal* childhood deaths in China play a smaller role than in India [Anderson and Ray, 2010]. Moreover, to the extent that diffusion of diagnostic ultrasound in China enabled prenatal discrimination and thereby worse health later in life, affected cohorts were too young to drive the increased cardiovascular and respiratory diseases observed by Anderson and Ray [2010]. Finally, we depart from Anderson and Ray [2010] by analyzing excess female morality in a developed country, along with morbidity outcomes.

When comparing Chinese Americans to Filipino Americans instead of Whites,

[24] Among Asian racial groups in the US, Chinese shows the lowest fraction male for perinatal mortality [Almond and Cheng, 2016, Table 1]. To the extent this reflects average son preference, we argue Chinese-Americans are of particular interest when analyzing mortality and perinatal health outcomes (in the subsequent morbidity analysis).

[25] See footnote 6.

we observe similar patterns but less precise point estimates due to smaller sample sizes in the control group. Nevertheless, the gender gap in fetal deaths and death on day 1 of life differs statistically for Chinese and Filipinos. These excess female fetal deaths among Chinese Americans contribute to the high sex ratio at birth above and beyond the effect of sex-selective abortion. Live-born Chinese American girls also die too often, which is a new finding suggesting morbidity at birth.

5.2 Birth certificate microdata

We examine perinatal outcomes during the post-ultrasound period (1989-2013) using natality microdata. The DD coefficients are presented in Table 2's top panel and logit estimates[26] in the corresponding bottom panel. Linked birth/infant death data allow us to revisit neonatal deaths with birth as the unit of analysis.[27] Consistent with the patterns in Table 1, our logit estimates in Table 2 show that the probability of death within 1 day increases by 33.2% among Chinese-American girls.[28] The female advantage in neonatal deaths among US Whites is eliminated and indeed reversed among Chinese. However, we find no effect on death days 1 to 28 of life.[29] Beyond mortality, female Chinese American neonates have a 3.0% increase in probability of low birth weight than projected by the US White gender gap. Low birth weight may be a better measure than average birthweight because of the non-linear relationship between birthweight and health [Almond, Chay, and Lee, 2005] and because the differential adoption of c-sections will confound our estimates

[26]Logit estimates are to be interpreted as percent change by its multiplicative feature. See Appendix Section B for details.

[27]We cannot perform the same analysis on fetal deaths and post-neonatal deaths (up to age 12) due to lack of linked live-birth data.

[28]We also explored following Clarke, Oreffice, and Quintana-Domeque [2019] and adopting adjusted critical values for t statistics to counter over-rejection of null hypotheses with large sample sizes [Leamer, 1978]. Our effects on perinatal outcomes remain significant except for low birth weight.

[29]Results are presented in Appendix Table S3

for relatively normal birthweights.[30] Five-minute APGAR scores and congenital anomalies show similar disadvantages among female Chinese Americans. The logit estimates in Table 2 show that Chinese American girls are 4.05% more likely to have APGAR scores below 8 and their probability of having congenital anomalies is 16.1% higher. The last two columns show effects on pregnancy-related and chronic (pre-pregnancy) maternal hypertension. We find that Chinese mothers in the US who are pregnant with a baby girl are 5.75% more likely to have pregnancy-related hypertension. We do not observe a corresponding difference in chronic (pre-pregnancy) hypertension, a "placebo" outcome. While a cortisol-based measure of stress would be preferred, we hypothesize that Chinese mothers with a baby girl might experience more stress or less support during pregnancy from their families.

Table 3 divides our analysis sample into subgroups. Among families where both parents are Chinese, effects on low birth weight, APGAR score, congenital anomalies, and pregnancy-related hypertension increase in magnitude. The effect on death within 1 day have a slightly smaller magnitude but remains significant. The elevated sex ratio among Chinese Americans generates concern of compositional effects due to sex selection. But the sex ratio among Chinese first parity births is 1.07, which is close to the biological normal ratio of 1.05 or 1.06. We view the first child's sex as relatively exogenous and see how parents respond to it [Ben-Porath and Welch, 1976].[31] Most of the effects among female Chinese Americans remain statistically significant among first parity births. The effect on low birth weight doubles in magnitude compared to the full sample. One exception is that the increase in the probability of

[30] We find Chinese American boys have a 1.95% higher probability of c-section, see Almond, Chee, Sviatschi, and Zhong [2015]. This leads to a mechanical reduction in gestation length (estimated as 4.5 hours) and average birth weight. See Appendix Table S3. If we ignore the issue of endogenous sample selection and restrict the sample to vaginal births, we find similar results.

[31] We know of no evidence that *boys* are selectively aborted among Chinese, which could offset selective abortion of females to yield normal sex ratios. Therefore, we interpret relatively normal sex ratios as reflecting low or no sex selection, and thereby a natural experiment in child sex at first parity.

death within 1 day is entirely driven by higher parity births. We also investigate whether our findings differ by maternal education levels and find larger effects among less educated Chinese mothers across all outcomes except for APGAR score below 8. Finally, we compare effects among first generation (foreign born) and second generation (US born) Chinese mothers (Appendix Table S4). Not surprisingly, negative effects among Chinese American girls are mostly concentrated among first generation families. That said, US-born Chinese mothers are only 10% of Chinese mothers in our sample. Separately, we lose significance on low birth wright in all groups for this shorter sample period.[32]

We repeat our analysis using natality data prior to ultrasound diffusion (1968-1980) as a falsification test. Although coefficients in Table 4 column 1 and 2 are less precisely estimated due to the smaller sample size (the number of Chinese births grows over time), all coefficients are substantially smaller and indeed take the opposite sign from what "son-preference" would suggest. This pattern is consistent with our hypothesis that female Chinese American babies were protected in the womb due to the possibility they might be male. Moreover, this raises confidence that our effects are not from biological differences between White and Chinese. Another approach to the same issue is to use Filipino Americans as an alternative control group, following Almond, Edlund, and Milligan [2013]. The coefficients show a very similar pattern in Table 4 column 5 and 6 except that effects on congenital anomalies and pregnancy-related hypertension lose their significance. Our estimates are also robust to including maternal covariates (maternal age, education, marital status) and birth order, as shown in Appendix Table S5.[33]

[32]Since mother's birth country is reported in birth data up to 2004, we adopt a shorter sample period in Appendix Table S4.

[33]For example, Ahsan and Maharaj (2018) found that higher maternal education (which differs by race) has a positive effect on neonatal health.

5.3 Hospital discharge microdata

If health among female Chinese Americans is worse at birth than expected, how do hospitals react? Using discharge data from New York, we see that hospitals appear to try and counteract such adverse health. Table 5 shows that "Chinese"[34] female babies in New York stay 0.59% longer in hospital, incur 5.08% higher total charges, and receive 25.2% more procedures when compared to Whites.[35,36] However, after discharge from hospital, they are less likely to be re-admitted despite higher than expected mortality.[37] The re-admission probability during the neonatal period (28-day) is 11.0% lower for Chinese American girls. The magnitude of this gap increases to 18.2% when we take into account their observed health disadvantage at birth. Turning to re-admission probability during the birth year, Chinese girls are 8.01% less likely to be readmitted, and adding at-birth controls increases this negative effect to 14.9%. In terms of the intensive margin, Chinese American girls have fewer readmissions during their birth year than projected by US Whites, reported in Appendix Table S3. This frequency gap is statistically significant and larger after controlling for health at birth.

5.4 "Calculated effects"

Section 3 highlights that the unadjusted DD estimates may be compromised by differential average birth outcomes by race. Logit model or log dependent variable could address this issue and provide effect estimates in terms of percentage changes. To provide effect estimates in absolute terms (units of the un-transformed dependent variable), we summarize our "calculated effects" on

[34] See Section 4 for race designations and assignment in New York's hospital discharge data.

[35] Estimates are robust to including hospital fixed effects. This helps address the possibility that our effects are driven by parents' gender-biased choice of the birth hospital. It also helps alleviate concern about compositional effects from "birth tourists" to the extent that mothers arriving from abroad tend to give birth at certain hospitals. See also footnote 21.

[36] We log these outcomes because they are right-skewed [Manning and Mullahy, 2001].

[37] Analyses of readmission are performed on data with readmission linkage, i.e. 2003-2013.

birth outcomes in Table 6.

Column 1 summarizes adjusted effects from the DD estimates in Table 2. Chinese American girls have a 0.030 percentage points increase in probability of death within 1 day, and the probability of low birth weight increases by 0.135 percentage points. APGAR score (non-binary measure) is 0.0128 lower on average, and the probability of congenital anomalies is 0.065 percentage points higher among Chinese American girls than projected by the US White gender gap. Similar to the patterns in Table 3, effect magnitudes are larger among families where both parents are Chinese except for death within 1 day and pregnancy-related hypertension. Comparing first parity births, where we believe the child's sex as relatively exogenous, to higher parity births, the effect on low birth weight is completely driven by the first parity. It is worth noting that estimated effects on APGAR score are significant and similar in magnitude among the two birth groups despite the small point estimates on the binary APGAR score measure. Analyses by maternal education levels indicate larger effects among less educated Chinese mothers across most of birth outcomes, same as shown in Table 3.

6 Discussion

Son preference in the US is not confined to sex-selective abortion. Among Chinese Americans, female death is more common both in the perinatal period and in childhood than we should expect. Moreover, perinatal health outcomes among Chinese American girls are worse. Hospitals seem to respond to worse neonatal health with longer stays, higher charges, and more medical procedures. But these countermeasures do not eliminate the unexpected gender gap in health (as captured by subsequent mortality). After delivery, girls are less likely to be re-admitted to the hospital, a decision about which parents have some discretion. This re-admission deficit grows when we factor in their worse health at birth.

The strong biological advantage enjoyed by females in early childhood health helps mask our finding.[38] Gender equality is simply the wrong counterfactual. Compromised perinatal health has persistent effects later in life [Barker, 1992] even for small differences in health around birth [Almond, Currie, and Duque, 2018] and for outcomes other than adult health, e.g. wages. The pre-ultrasound veil of ignorance that once protected girls from son preference has been eliminated, and future gender inequality may be "programmed" as a result. Now that prenatal sex determination is routine, sex selection does allow those with particularly strong son preference to abort girls, which would tend to improve outcomes for surviving girls [Goodkind, 1996, Hu and Schlosser, 2015, Anukriti et al., 2018]. Our effects on perinatal health overwhelm any selection effect. Rates of sex selection among Chinese Americans (as inferred from sex ratios) are lower than in China or India, which circumscribes the compositional effect of sex selection.[39] At first parity, the sex ratio of 1.07 is within 2% of the biological norm, further reducing the scope for compositional effects.

A limitation of our study is that we do not identify the mechanism(s) by which this detriment to female health emerges in the US. Our data do not include administrative measures of health investments, only self-reported ones. Moreover, some of the gender differential investment may be unconscious, including greater feelings of stress if a girl is expected.[40] Nor do we know the share of Chinese parents that respond to fetal sex in a way that harms girls. It would be consistent with our results if a majority of Chinese-American parents did *not* react in this way.

We hypothesize son preference extends beyond sex-selective abortion (and

[38]For most outcomes, Chinese girls still do better than Chinese males and White females.

[39]Nor do we leverage variation in sex selection for identification, *cf.* Hu and Schlosser [2015], which would tend to accentuate the compositional effect [Bharadwaj and Lakdawala, 2013].

[40]Bereavement stress during pregnancy, which Persson and Rossin-Slater [2018] find compromises the later-life mental health of the fetus, has a similar impact on low birth weight as our preferred estimate in Table 3: .00293 versus .00358 for the death of a close relative in [Persson and Rossin-Slater, 2018, Table 2].

postnatal discrimination) to perinatal health in China. Absent corresponding administrative data for China, we are unable to evaluate this hypothesis directly. Our prior is that son preference would not be strongest among Chinese who emigrate to the US. That sex ratios at birth are higher in China than among Chinese Americans is consistent with attenuated son preference in the US. Compositional effects from sex selection aside, the average effects we estimate may then be a lower bound for the 8 million girls born each year in China.

References

Jason Abrevaya. Are there missing girls in the united states? evidence from birth data. *American Economic Journal: Applied Economics*, 1(2):1–34, April 2009.

Douglas Almond and Yi Cheng. Perinatal health among 1 million American-born Chinese. December 2016. Columbia University, Department of Economics Discussion Paper #1617-04.

Douglas Almond and Lena Edlund. Son biased sex ratios in the 2000 United States census. *Proceedings of the National Academy of Sciences (PNAS)*, 105(15):5681–5682, April 2008.

Douglas Almond, Kenneth Y. Chay, and David S. Lee. The costs of low birth weight. *The Quarterly Journal of Economics*, 120(3):1031–1084, August 2005.

Douglas Almond, Lena Edlund, and Kevin Milligan. Son preference and the persistence of culture: Evidence from south and east asian immigrants to canada. *Population and Development Review*, 39(1):75–95, 2013. ISSN 1728-4457. doi: 10.1111/j.1728-4457.2013.00574.x. URL `http://dx.doi.org/10.1111/j.1728-4457.2013.00574.x`.

Douglas Almond, Christine Pal Chee, Maria Micaela Sviatschi, and Nan Zhong. Auspicious birth dates among chinese in california. *Economics & Human Biology*, 18:153 – 159, 2015. ISSN 1570-677X. doi: https://doi.org/10.1016/j.ehb.2015.05.005. URL http://www.sciencedirect.com/science/article/pii/S1570677X15000404.

Douglas Almond, Janet Currie, and Valentina Duque. Childhood circumstances and adult outcomes: Act ii. *Journal of Economic Literature*, 56 (4):1360–1446, December 2018. doi: 10.1257/jel.20171164. URL http://www.aeaweb.org/articles?id=10.1257/jel.20171164.

Douglas Almond, Hongbin Li, and Shuang Zhang. Land reform and sex selection in china. *Journal of Political Economy*, 127(2):560–585, 2019. doi: 10.1086/701030. URL https://doi.org/10.1086/701030.

Siwan Anderson and Debraj Ray. Missing women: Age and disease. *The Review of Economic Studies*, 77(4):1262–1300, 2010. doi: 10.1111/j.1467-937X.2010.00609.x. URL http://restud.oxfordjournals.org/content/77/4/1262.abstract.

Joshua D. Angrist and Jörn-Steffen Pischke. *Mostly Harmless Econometrics: An Empiricist's Companion*. Princeton University Press, Princeton, New Jersey, 2009.

S Anukriti, Sonia Bhalotra, and Hiu Tam. On the Quantity and Quality of Girls: Fertility, Parental Investments, and Mortality. Boston College Working Papers in Economics 950, Boston College Department of Economics, January 2018. URL https://ideas.repec.org/p/boc/bocoec/950.html.

D.J.P. Barker, editor. *Fetal and Infant Origins of Adult Disease*. British Medical Journal, London, 1992.

Yoram Ben-Porath and Finis Welch. Do sex preferences really matter?*.

The Quarterly Journal of Economics, 90(2):285–307, 1976. doi: 10.2307/1884631. URL +http://dx.doi.org/10.2307/1884631.

Prashant Bharadwaj and Leah K. Lakdawala. Discrimination begins in the womb: Evidence of sex-selective prenatal investments. *Journal of Human Resources*, 48(1):71–113, 2013. URL http://ideas.repec.org/a/uwp/jhriss/v48y2013i1p71-113.html.

Yuyu Chen, Hongbin Li, and Lingsheng Meng. Prenatal sex selection and missing girls in china: Evidence from the diffusion of diagnostic ultrasound. *Journal of Human Resources*, 48(1):36–70, 2013. doi: 10.3368/jhr.48.1.36. URL http://jhr.uwpress.org/content/48/1/36.abstract.

Damian Clarke, Sonia Oreffice, and Climent Quintana-Domeque. The demand for season of birth. *Journal of Applied Econometrics*, 34(5):707–723, 2019.

Monica DasGupta, Zhenghua Jiang, Bohua Li, Zhenming Xie, Woojin Chung, and Bae Hwa-OK. Why is son preference so persistent in east and south asia? a cross-country study of china, india and the republic of korea. *Journal of Development Studies*, 40(2):153–187, 2003.

Sylvie Dubuc and David Coleman. An increase in the sex ratio of births to india-born mothers in england and wales: Evidence for sex-selective abortion. *Population and Development Review*, 33(2):383–4000, June 2007.

Esther Duflo. Women empowerment and economic development. *Journal of Economic Literature*, 50(4):1051–79, December 2012. doi: 10.1257/jel.50.4.1051. URL http://www.aeaweb.org/articles?id=10.1257/jel.50.4.1051.

Lena Edlund. Son preference, sex rations, and marriage patterns. *Journal of Political Economy*, 107(6):1275–1304, December 1999. URL http://ideas.repec.org/a/ucp/jpolec/v107y1999i6p1275-1304.html.

Libertad Gonzlez. Sex selection and health at birth among indian immigrants. *Economics & Human Biology*, 29:64 – 75, 2018. ISSN 1570-677X. doi: https://doi.org/10.1016/j.ehb.2018.02.003. URL http://www.sciencedirect.com/science/article/pii/S1570677X17300527.

Daniel Goodkind. On substituting sex preference strategies in east asia: Does prenatal sex selection reduce postnatal discrimination? *Population and Development Review*, 22(1):111–125, 1996. ISSN 00987921, 17284457. URL http://www.jstor.org/stable/2137689.

Luojia Hu and Anala Schlosser. Prenatal sex selection and girls' well-being: Evidence from india. *The Economic Journal*, 125(587):1227–1261, 2015. ISSN 1468-0297. doi: 10.1111/ecoj.12259. URL http://dx.doi.org/10.1111/ecoj.12259.

Sebastian Kraemer. The fragile male. *BMJ*, 321(7276):1609–1612, 2000. doi: 10.1136/bmj.321.7276.1609. URL http://www.bmj.com.

Edward E Leamer. *Specification searches: Ad hoc inference with nonexperimental data*, volume 53. Wiley New York, 1978.

Aparna Lhila and Kosali I. Simon. Prenatal health investment decisions: Does the child's sex matter? *Demography*, 45(4):885–905, November 2008.

Willard G Manning and John Mullahy. Estimating log models: to transform or not to transform? *Journal of Health Economics*, 20(4):461 – 494, 2001. ISSN 0167-6296. doi: https://doi.org/10.1016/S0167-6296(01)00086-8. URL http://www.sciencedirect.com/science/article/pii/S0167629601000868.

Bruce Meyer. Natural and quasi-experiments in economics. *Journal of Business and Economics Statistics*, 13(2):151–161, April 1995.

Felix M. Muchomba and Sangeeta Chatterji. Child gender and disability among children of immigrants from India and China: Is there excess fe-

male disability? 2020. Rutgers University School of Social Work, working paper.

Petra Persson and Maya Rossin-Slater. Family ruptures, stress, and the mental health of the next generation. *American Economic Review*, 108(4-5):1214–52, April 2018. doi: 10.1257/aer.20141406. URL `http://www.aeaweb.org/articles?id=10.1257/aer.20141406`.

Swapnalee Sarmah, Pooja Muralidharan, and James A Marrs. Common congenital anomalies: Environmental causes and prevention with folic acid containing multivitamins. *Birth Defects Research Part C: Embryo Today: Reviews*, 108(3):274–286, 2016.

A. Sen. Missing women. *BMJ*, 304(6827):587–588, 1992. ISSN 0959-8138. doi: 10.1136/bmj.304.6827.587. URL `http://www.bmj.com/content/304/6827/587`.

Bob Weinhold. Environmental factors in birth defects: what we need to know. *Environmental Health Perspectives*, 117(10):A440, 2009.

Jiaquan Xu, Sherry L. Murphy, Kenneth D. Kochanek, and Brigham A. Bastian. Deaths: final data for 2013. *National vital statistics reports*, 64(2), 2016.

Zeng Yi, Tu Ping, Gu Baochang, Xu Yi, Li Bohua, and Li Yongping. Causes and implications of the recent increase in the reported sex ratio at birth in china. *Population and Development Review*, 19(2):283–302, June 1993.

Table 1: Fraction Male among Deaths

Sample	(1) Chinese	(2) White	(3) Filipino	(4) (1)-(2)	(5) (1)-(3)	N (Chinese)	N(Filipino)
Panel A: Fetal Deaths (1989-2005)							
Gestation ≥20 Weeks	0.490	0.528	0.526	-0.0384***	-0.0362**	1603	2763
				(0.0125)	(0.0157)		
Panel B: Neonatal (within 28 days) Deaths (1989-2015)							
Within 1 Day	0.512	0.553	0.553	-0.0410**	-0.0414*	838	1153
				(0.0172)	(0.0226)		
In 1-28 Days	0.568	0.566	0.570	0.00225	-0.00134	695	739
				(0.0188)	(0.0262)		
Panel C: Post-neonatal (beyond 1 month) Deaths (1989-2015)							
Up to 12 Years	0.555	0.576	0.565	-0.0214*	-0.0105	1587	1747
(All Causes)				(0.0124)	(0.0172)		
Up to 12 Years	0.535	0.563	0.555	-0.0276**	-0.0200	1289	1401
(Disease-related Causes)				(0.0139)	(0.0192)		

$***\ p<0.01$, $**\ p<0.05$, $*\ p<0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data.

Note: We exclude accidents, suicide, homicide, and all other external death causes in defining disease-related causes according to ICD codes.

Table 2: Gender Gaps in Perinatal Outcomes

	Death within 1 Day	Low Birth Weight	APGAR Score<8	APGAR Score	Congenital Anomalies	Pregnancy Hypertension	Chronic Hypertension
Chinese	-0.00105***	-0.00894***	-0.0136***	0.0482***	-0.00395***	-0.0251***	-0.00445***
	(0.0000781)	(0.000352)	(0.000301)	(0.00135)	(0.000144)	(0.000290)	(0.000138)
Female	-0.000343***	0.00692***	-0.00477***	0.0253***	-0.00274***	-0.00169***	-0.000140**
	(0.0000344)	(0.000158)	(0.000126)	(0.000566)	(0.0000642)	(0.000131)	(0.0000623)
Chinese×Female	**0.000484*** **	**0.0000589**	**0.00271*** **	**-0.0126*** **	**0.00182*** **	**0.00178*** **	**0.0000611**
	(0.000113)	**(0.000508)**	**(0.000434)**	**(0.00194)**	**(0.000207)**	**(0.000418)**	**(0.000199)**
Constant	0.00198***	0.0478***	0.0295***	8.902***	0.00927***	0.0365***	0.00787***
	(0.0000240)	(0.000110)	(0.0000882)	(0.000395)	(0.0000448)	(0.0000912)	(0.0000435)
Logit	**0.332*** **	**0.0299*** **	**0.0405***	**-**	**0.161*** **	**0.0575*** **	**-0.00537**
Chinese×Female	**(0.0840)**	**(0.0115)**	**(0.0222)**	**-**	**(0.0338)**	**(0.0212)**	**(0.0391)**
N	6470115	8464639	6957570	6957570	8149587	8351362	8351362

$***\ p<0.01$, $**\ p<0.05$, $*\ p<0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data (1989-2013).

Table 3: Logit Estimates by Subgroups

	Mother Chinese	Both Parents Chinese	First Parity	Higher Parity	High Maternal Education	Low Maternal Education
Death within 1 Day	0.332***	0.269**	0.148	0.507***	0.232**	0.485***
	(0.0840)	(0.104)	(0.123)	(0.117)	(0.118)	(0.131)
Low Birth Weight	0.0299***	0.0461***	0.0594***	-0.00263	0.0238	0.0454**
	(0.0115)	(0.0133)	(0.0152)	(0.0176)	(0.0149)	(0.0183)
APGAR Score<8	0.0405*	0.0567**	0.0192	0.0761**	0.0657**	0.00437
	(0.0222)	(0.0270)	(0.0283)	(0.0362)	(0.0289)	(0.0368)
Congenital Anomalies	0.161***	0.203***	0.147***	0.170***	0.136***	0.190***
	(0.0338)	(0.0394)	(0.0469)	(0.0491)	(0.0422)	(0.0562)
Pregnancy Hypertension	0.0575***	0.0672**	0.0839***	0.00504	0.0652**	0.0665*
	(0.0212)	(0.0267)	(0.0263)	(0.0362)	(0.0266)	(0.0370)
N Chinese	822825	620246	418635	401085	486851	295101
% of Total Chinese	100%	75%	51%	49%	59%	36%

$***\ p < 0.01$, $**\ p < 0.05$, $*\ p < 0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data (1989-2013).

Note: The subgroups by maternal education do not add up to one due to missing values in maternal education.

Table 4: Effect Estimates by Sample Periods and Control Groups

	D-in-D	Logit	D-in-D	Logit	D-in-D	Logit
Death within 1 Day	-	-	0.000484***	0.332***	0.000479***	0.313***
	-	-	(0.000113)	(0.0840)	(0.000141)	(0.100)
Low Birth Weight	-0.00247	-0.0360	0.0000589	0.0299***	0.000295	0.0746***
(Starting from 1968)	(0.00188)	(0.0393)	(0.000508)	(0.0115)	(0.000734)	(0.0140)
APGAR Score<8	0.00246	0.0203	0.00271***	0.0405*	0.00268***	0.0622**
(Starting from 1978)	(0.00282)	(0.108)	(0.000434)	(0.0222)	(0.000527)	(0.0286)
APGAR Score	0.00228	-	-0.0126***	-	-0.00957***	-
(Starting from 1978)	(0.0149)	-	(0.00194)	-	(0.00240)	-
Congenital Anonalies	0.00172	0.0468	0.00182***	0.161***	0.000941***	0.0557
(Starting from 1980)	(0.00210)	(0.292)	(0.000207)	(0.0338)	(0.000253)	(0.0422)
Pregnancy Hypertension	-	-	0.00178***	0.0575***	0.000564	0.0240
	-	-	(0.000418)	(0.0212)	(0.000454)	(0.0246)
Control Group	White		White		Filipino	
Sample Period	Pre-Ultrasound		Post-Ultrasound		Post-Ultrasound	

$***\ p < 0.01$, $**\ p < 0.05$, $*\ p < 0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data (1968-1980, 1989-2013).

Note: The pre-ultrasound period includes births 1968-1980, post-ultrasound 1989-2013.

Table 5: Gender Gaps in Hospital Outcomes in New York

	Length of Stay (log)	Total Hospital Charges (log)	Number of Procedures (log)	Readmission within 28 Days		Readmission within Birth Year	
Chinese	0.000245	0.217***	0.0647***	0.00280***	-0.000281	0.00606***	-0.00150
	(0.00161)	(0.00329)	(0.00166)	(0.000673)	(0.000682)	(0.000928)	(0.000940)
Female	-0.0253***	-0.107***	-0.418***	-0.00389***	-0.00191***	-0.00712***	-0.00303***
	(0.000538)	(0.00110)	(0.000553)	(0.000240)	(0.000258)	(0.000331)	(0.000356)
Chinese×Female	**0.00594****	**0.0508*****	**0.252*****	**-0.00227****	**-0.00349*****	**-0.00361*****	**-0.00585*****
	(0.00232)	**(0.00475)**	**(0.00239)**	**(0.000970)**	**(0.000971)**	**(0.00134)**	**(0.00134)**
Constant	1.841***	7.627***	0.747***	0.0169***	0.0112**	0.0833***	0.0310***
	(0.0110)	(0.0232)	(0.0112)	(0.00433)	(0.00505)	(0.00597)	(0.00695)
Logit	**-**	**-**	**-**	**-0.110***	**-0.182*****	**-0.0801***	**-0.149*****
Chinese×Female	**-**	**-**	**-**	**(0.0587)**	**(0.0588)**	**(0.0419)**	**(0.0421)**
N	2542492	2536746	2550520	1193211	1193022	1193211	1193022
At-birth Covariates	NA	NA	NA	No	Yes	No	Yes

$***\ p<0.01$, $**\ p<0.05$, $*\ p<0.1$ Standard errors in parentheses

Data: HCUP New York State Inpatient Database (1993-2013).

Note: Analyses of readmission are performed on data with readmission linkage, i.e. 2003-2013. We include birth year fixed effects in all regressions. At-birth covariates include insurance types, total hospital charges, length of stay, and the number of procedures. We use log(y+1) in all log outcomes to include zero values. This will attenuate slightly estimated magnitudes, more so for outcomes with a low mean (number of procedures). We "isolate" Chinese from other Asian races by including only Asian residents in zip code areas with dominantly Chinese.

Table 6: Summary on Calculated Effects

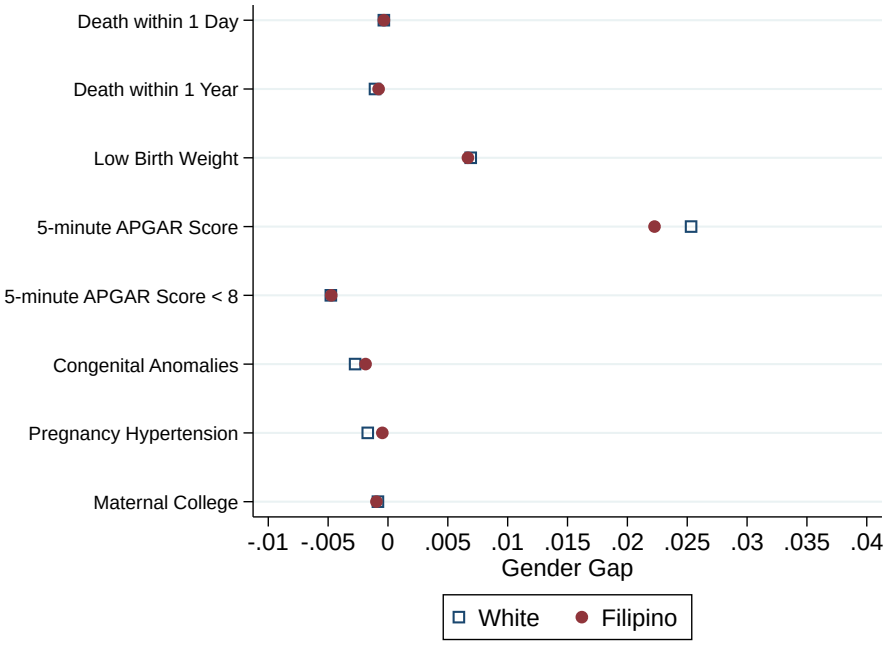
	Mother Chinese	Both Parents Chinese	First Parity	Higher Parity	High Maternal Education	Low Maternal Education
Death within 1 Day	0.000302***	0.000205*	0.000120	0.000503***	0.000169	0.000479***
	(0.0000998)	(0.000115)	(0.000143)	(0.000139)	(0.000128)	(0.000166)
Low Birth Weight	0.00135**	0.00198***	0.00293***	-0.0000485	0.00106	0.00208**
	(0.000535)	(0.000610)	(0.000804)	(0.000722)	(0.000687)	(0.000880)
APGAR Score<8	0.000510	0.000644	0.000256	0.000765	0.000880*	0.0000234
	(0.000389)	(0.000450)	(0.000608)	(0.000504)	(0.000524)	(0.000632)
APGAR Score	-0.0128***	-0.0155***	-0.0136***	-0.0136***	-0.0119***	-0.0139***
	(0.00195)	(0.00223)	(0.00294)	(0.00261)	(0.00258)	(0.00309)
Congenital Anomalies	0.000646***	0.000789***	0.000615**	0.000653***	0.000567**	0.000765***
	(0.000174)	(0.000201)	(0.000250)	(0.000246)	(0.000225)	(0.000296)
Pregnancy Hypertension	0.000619	0.000599	0.00115*	0.0000347	0.000747	0.000644
	(0.000390)	(0.000451)	(0.000646)	(0.000470)	(0.000514)	(0.000663)
N Chinese	822825	620246	418635	401085	486851	295101
% of Total Chinese	100%	75%	51%	49%	59%	36%

$***\ p<0.01$, $**\ p<0.05$, $*\ p<0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data (1989-2013).

Note: The subgroups by maternal education do not add up to one due to missing values in maternal education.

Figure 1: Gender Gaps by Maternal Race



Appendices

A Log DD Model

We derive the correspondence between estimates of Θ introduced in Section 3 and coefficients from a log DD model.

Consider a DD model of log outcomes:

$$log(y_i) = \beta + \beta_c \cdot Chinese_i + \phi \cdot Female_i + \phi_c \cdot Female_i \cdot Chinese_i + \varepsilon_i. \quad (3)$$

Applying the approximation $e^x \approx 1 + x$:

$$\begin{aligned}
y_i &= e^{\beta+\beta_c \cdot Chinese_i+\phi \cdot Female_i+\phi_c \cdot Female_i \cdot Chinese_i+\varepsilon_i}, \\
y_i &\approx e^{\beta}(1+\beta_c \cdot Chinese_i)(1+\phi \cdot Female_i)(1+\phi_c \cdot Female_i \cdot Chinese_i) + e_i, \\
y_i &= (e^{\beta}+e^{\beta}\beta_c \cdot Chinese_i)(1+\phi \cdot Female_i)(1+\phi_c \cdot Female_i \cdot Chinese_i) + e_i, \\
y_i &= (e^{\beta}+e^{\beta}\beta_c \cdot Chinese_i)(1+\phi \cdot Female_i), \\
&\quad +\phi_c(e^{\beta}+e^{\beta}\beta_c \cdot Chinese_i)(1+\phi \cdot Female_i) \cdot Female_i \cdot Chinese_i + e_i, \\
y_i &= e^{\beta}+e^{\beta}\beta_c \cdot Chinese_i + e^{\beta}\phi \cdot Female_i + e^{\beta}\beta_c\phi \cdot Female_i \cdot Chinese_i, \\
&\quad +\phi_c((e^{\beta}+e^{\beta}\beta_c)(1+\phi)) \cdot Female_i \cdot Chinese_i + e_i.
\end{aligned}$$

We have the correspondence between equations (1) and (3): $\beta = log(\gamma)$, $\beta_c = \frac{\gamma_c}{\gamma}$, $\phi = \alpha = \frac{\theta}{\gamma}$, and $\Theta = \phi_c(\gamma + \gamma_c)(1 + \alpha)$. That is, ϕ_c identifies the effect as a percentage of the Chinese female average.

B Logit DD Model

Consider the standard logistic DD regression model on a binary outcome Y:

$$P(Y_i = 1) = \frac{1}{1 + e^{-(\beta + \beta_c \cdot Chinese_i + \phi \cdot Female_i + \phi_c \cdot Female_i \cdot Chinese_i)}}. \tag{4}$$

We derive the race- and gender-specific predicted odds,

$$odds_{wm} = \frac{P(Y_{wm} = 1)}{P(Y_{wm} = 0)} = e^{\beta}, \qquad odds_{wf} = \frac{P(Y_{wf} = 1)}{P(Y_{wf} = 0)} = e^{\beta + \phi},$$
$$odds_{cm} = \frac{P(Y_{cm} = 1)}{P(Y_{cm} = 0)} = e^{\beta + \beta_c}, \qquad odds_{cf} = \frac{P(Y_{cf} = 1)}{P(Y_{cf} = 0)} = e^{\beta + \beta_c + \phi + \phi_c}.$$

That is, we derive a log DD model 3 on odds:

$$log(odds_i) = \beta + \beta_c \cdot Chinese_i + \phi \cdot Female_i + \phi_c \cdot Female_i \cdot Chinese_i.$$

Any interpretation of coefficients derived in Appendix A applies to logit DD model in terms of odds.[41]

[41] If the probability of occurrence is small, $\text{odds}_i \approx P(Y_i = 1)$.

C Normalized DD Model

We develop an equivalent DD specification by normalizing outcomes based on the race-specific male average, i.e. a normalized DD model.

Consider DD specification (1) in Section 3:

$$y_i = \gamma + \gamma_c \cdot Chinese_i + \theta \cdot Female_i + \theta_c \cdot Female_i \cdot Chinese_i + \varepsilon_i.$$

Assuming $\theta = \alpha \cdot \gamma$ and $\theta_c = \alpha \cdot \gamma_c + \Theta$, we derive race-gender specific averages

$$\begin{aligned} \bar{y}_{wm} &= \gamma, & \bar{y}_{wf} &= \gamma + \theta = \gamma(1+\alpha), \\ \bar{y}_{cm} &= \gamma + \gamma_c, & \bar{y}_{cf} &= \gamma + \gamma_c + \theta + \theta_c = (\gamma + \gamma_c)(1+\alpha) + \Theta. \end{aligned}$$

Denote race-specific normalized outcomes

$$\begin{aligned} \tilde{y}_{iw} &\equiv \frac{y_{iw}}{\bar{y}_{wm}} = 1 + \alpha \cdot Female_{iw} + \tilde{\varepsilon}_{iw}, \\ \tilde{y}_{ic} &\equiv \frac{y_{ic}}{\bar{y}_{cm}} = 1 + \alpha \cdot Female_{ic} + \frac{\Theta}{\gamma + \gamma_c} \cdot Female_{ic} + \tilde{\varepsilon}_{ic}. \end{aligned}$$

That is

$$\tilde{y}_i = \beta + \beta_c \cdot Chinese_i + \phi \cdot Female_i + \phi_c \cdot Female_i \cdot Chinese_i + \tilde{\varepsilon}_i, \qquad (5)$$

where $\beta = 1$, $\phi = \alpha$, $\beta_c = 0$, and $\Theta = \phi_c \cdot (\gamma + \gamma_c)$. That is, ϕ_c identifies the effect as a percentage of the Chinese male average.

Appendix Tables

Table S1: Summary Statistics: Birth Certificate Micro-data

	Mean	Stdev	Min	Median	Max	N
Panel A: White Sample						
Fraction Male	0.513	0.500	0	1	1	7649124
Maternal Age	27.36	5.987	14	27	50	7649124
Marital Status	0.710	0.454	0	1	1	7649124
High School Graduate	0.786	0.410	0	1	1	7250569
College Graguate	0.250	0.433	0	0	1	7250569
Death within 28 Days	0.00324	0.0568	0	0	1	5865911
APGAR Score	8.914	0.725	0	9	10	6366740
Birth Weight	3377.2	555.4	227	3402	8165	7642464
Low Birth Weight	0.0512	0.220	0	0	1	7642464
Congenital Anomalies	0.00794	0.0887	0	0	1	7365996
Pregnancy Hypertension	0.0357	0.185	0	0	1	7535059
Panel B: Chinese Sample						
Fraction Male	0.520	0.500	0	1	1	822825
Maternal Age	31.53	4.872	14	32	50	822825
Marital Status	0.904	0.294	0	1	1	822825
High School Graduate	0.875	0.331	0	1	1	781952
College Graguate	0.553	0.497	0	1	1	781952
Death within 28 Days	0.00187	0.0432	0	0	1	604204
APGAR Score	8.956	0.569	0	9	10	590830
Birth Weight	3292.2	481.4	227	3290	7977	822175
Low Birth Weight	0.0422	0.201	0	0	1	822175
Congenital Anomalies	0.00487	0.0696	0	0	1	783591
Pregnancy Hypertension	0.0114	0.106	0	0	1	816303
Panel C: Filipino Sample						
Fraction Male	0.517	0.500	0	1	1	785189
Maternal Age	29.93	5.916	14	30	50	785189
Marital Status	0.802	0.399	0	1	1	785189
High School Graduate	0.937	0.243	0	1	1	737816
College Graguate	0.430	0.495	0	0	1	737816
Death within 28 Days	0.00334	0.0577	0	0	1	596638
APGAR Score	8.898	0.680	0	9	10	494726
Birth Weight	3215.4	543.7	227	3232	7777	784258
Low Birth Weight	0.0737	0.261	0	0	1	784258
Congenital Anomalies	0.00765	0.0871	0	0	1	765365
Pregnancy Hypertension	0.0311	0.174	0	0	1	776528

Data: U.S. Vital Statistics micro data (1989-2013).

Table S2: Summary Statistics: Hospital Discharge Micro-data

	Mean	Stdev	Min	Median	Max	N
Panel A: White Sample						
Length of Stay	3.012	5.708	0	2	462	2407257
Total Charges	5080.1	27775.4	1	1760	4695603	2401653
Number of Procedures	0.924	1.259	0	1	15	2415040
Readmission within 28 Days	0.0163	0.127	0	0	1	1119889
Readmission within Birth Year	0.0315	0.175	0	0	1	1119889
Number of Readmissions within Birth Year	0.0357	0.217	0	0	12	1119889
Panel B: Chinese Sample						
Length of Stay	2.871	4.633	0	2	234	136583
Total Charges	5921.6	27341.0	10	2585	2236011	136441
Number of Procedures	1.417	1.288	0	1	15	136828
Readmission within 28 Days	0.0180	0.133	0	0	1	73322
Readmission within Birth Year	0.0358	0.186	0	0	1	73322
Number of Readmissions within Birth Year	0.0399	0.228	0	0	9	73322

Data: HCUP New York State Inpatient Database (1993-2013). Readmission link is only available after 2003. We "isolate" Chinese from other Asian races by including only Asian residents in zip code areas with a Chinese to Asian ratio above 0.5.

Table S3: Gender Gaps in Additional Outcomes

	Death in 1-28 Days	Ventilator Use	C-Section	Number of Readmissions within Birth Year	
Chinese	-0.000623***	-0.0131***	-0.00570***	0.00540***	-0.00390***
	(0.0000695)	(0.000263)	(0.000698)	(0.00107)	(0.00109)
Female	-0.000300***	-0.00370***	-0.0184***	-0.00812***	-0.00295***
	(0.0000306)	(0.000115)	(0.000315)	(0.000383)	(0.000411)
Chinese×Female	**0.000150**	**0.00206*** **	**-0.00322*** **	**-0.00299***	**-0.00578*** **
	(0.000100)	**(0.000379)**	**(0.00101)**	**(0.00155)**	**(0.00155)**
Constant	0.00157***	0.0279***	0.259***	0.103***	0.0297***
	(0.0000214)	(0.0000802)	(0.000220)	(0.00692)	(0.00804)
Logit	**0.0395**	**0.0270**	**-0.0195*** **	**-**	**-**
Chinese×Female	**(0.0903)**	**(0.0203)**	**(0.00544)**	**-**	**-**
N	6470115	8121554	8378796	1192929	1192741
At-birth Covariates	NA	NA	NA	No	Yes

$***\ p < 0.01$, $**\ p < 0.05$, $*\ p < 0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data (1989-2013), HCUP New York State Inpatient Database (1993-2013). In HCUP New York Inpatient Database, we "isolate" Chinese from other Asian races by including only Asian residents in zip code areas with Chinese to Asian ratio above 0.5.

Table S4: Effects by Mother's Birth Country

	Calculated Effect	Logit Effect	Calculated Effect	Logit Effect	Calculated Effect	Logit Effect
Death within 1 Day	0.000478***	0.489***	0.000563	0.536*	0.000468***	0.483***
	(0.000129)	(0.104)	(0.000415)	(0.313)	(0.000136)	(0.109)
Low Birth Weight	0.000724	0.0159	-0.00245	-0.0442	0.00102	0.0236
	(0.000715)	(0.0157)	(0.00227)	(0.0429)	(0.000749)	(0.0167)
APGAR Score	-0.0102***	-	0.00701	-	-0.0118***	-
	(0.00299)	-	(0.00974)	-	(0.00313)	-
Congenital Anonalies	0.00105***	0.177***	0.000916	0.108	0.00106***	0.187***
	(0.000275)	(0.0384)	(0.000865)	(0.0995)	(0.000289)	(0.0413)
Pregnancy Hypertension	0.000840	0.0819***	0.000622	0.0330	0.000838	0.0899***
	(0.000515)	(0.0296)	(0.00170)	(0.0695)	(0.000542)	(0.0326)
Chinese Sample	All Chinese Mothers		US Born Chinese Mothers		Foreign Born Chinese Mothers	
N Chinese	445934		42973		402961	

$***\ p < 0.01$, $**\ p < 0.05$, $*\ p < 0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data (1989-2004). Mother's birth country is not reported after 2004.

Reported Chinese sample size is not adjusted for missing values in outcomes.

Table S5: Estimates with and without Covariates

	Death within 1 Day		Congenital Anonalies		APGAR Score	
Chinese	-0.00105***	-0.000826***	-0.00395***	-0.00311***	0.0482***	0.0771***
	(0.0000781)	(0.0000799)	(0.000144)	(0.000147)	(0.00135)	(0.00137)
Female	-0.000343***	-0.000342***	-0.00274***	-0.00275***	0.0253***	0.0253***
	(0.0000344)	(0.0000344)	(0.0000642)	(0.0000641)	(0.000566)	(0.000563)
Chinese×Female	**0.000484*** **	**0.000479*** **	**0.00182*** **	**0.00182*** **	**-0.0126*** **	**-0.0123*** **
	(0.000113)	**(0.000113)**	**(0.000207)**	**(0.000207)**	**(0.00194)**	**(0.00193)**
Constant	0.00198***	0.00296***	0.00927***	0.0138***	8.902***	8.874***
	(0.0000240)	(0.000473)	(0.0000448)	(0.000898)	(0.000395)	(0.00848)
N	6470115	6470115	8149587	8149587	6957570	6957570
Covariates	No	Yes	No	Yes	No	Yes

$***\ p < 0.01$, $**\ p < 0.05$, $*\ p < 0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data (1989-2013).

Covariates include maternal age, education, marital status, infant's birth order, and birth year FE.

Table S6: Normalized DD Estimates: alternative gender benchmark

	Death within 1 Day		Congenital Anonalies		APGAR Score	
	Male Base	Female Base	Male Base	Female Base	Male Base	Female Base
Chinese	5.32e-08	9.43e-08	3.86e-08	-2.44e-08	1.25e-07	1.02e-07
	(0.0430)	(0.0513)	(0.0164)	(0.0237)	(0.000151)	(0.000157)
Female	-0.173***		-0.296***		0.00284***	
	(0.0190)		(0.00734)		(0.0000635)	
Chinese×Female	**0.325*** **		**0.121*** **		**-0.00143*** **	
	(0.0621)		**(0.0237)**		**(0.000218)**	
Male		0.210***		0.420***		-0.00284***
		(0.0218)		(0.0102)		(0.0000633)
Chinese×Male		**-0.342*** **		**-0.209*** **		**0.00142*** **
		(0.0712)		**(0.0329)**		**(0.000217)**
Constant	1.000***	1.000***	1.000***	1.000***	1.000***	1.000***
	(0.0132)	(0.0156)	(0.00513)	(0.00730)	(0.0000443)	(0.0000454)
Calculated Effect	**0.000301**	**-0.000366**	**0.000644**	**-0.000918**	**-0.0128**	**0.0127**
N	6470115	6470115	8149587	8149587	6957570	6957570

$***\ p < 0.01$, $**\ p < 0.05$, $*\ p < 0.1$ Standard errors in parentheses

Data: U.S. Vital Statistics micro data (1989-2013).

Only point estimates without significance levels are reported for calculated effect.