NBER WORKING PAPER SERIES

# CHILDREN ARE BRIDGES TO HEAVEN: THE EFFECTS OF FERTILITY ON LATER-LIFE MORTALITY

Jason Fletcher
Hamid Noghanibehambari

Working Paper 34378
http://www.nber.org/papers/w34378

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
October 2025

The authors would like to acknowledge financial support from NIA grants (R01AG060109, R01AG076830). The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

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NBER Working Paper No. 34378
October 2025
JEL No. I1, J13, N0

**ABSTRACT**

Several competing theories in a number of disciplines point to a possible influence of fertility history on health and mortality. However, the direction of effects is theoretically a-priori unknown and the empirical evidence is also inconclusive. This paper examines the effects of fertility during midlife on later-life longevity using Social Security Administration death records linked with the full-count 1940 census. We tackle endogeneity and selection concerns in the longevity-fertility association by implementing an instrumental variable based on the sex composition of first two children. Our findings indicate that having an extra child is linked to a decrease in women's longevity by approximately 5 months. Men, on the other hand, experience smaller and insignificant reductions in longevity of about 3.3 months. This divergence in effects suggests that biological factors may play a small role in the relationship between fertility and later-life longevity.

Jason Fletcher
University of Wisconsin - Madison
and NBER
jfletcher@lafollette.wisc.edu

Hamid Noghanibehambari
Austin Peay State University
noghanih@apsu.edu

## 1. Introduction

There has been a growing body of empirical [a]evidence to support life-course theories that highlight the relevance of lifecycle exposures on later-life and old-age health and mortality (Almond et al., 2018; Fletcher, 2012; Kim et al., 2017; Lupien et al., 2005; Montez & Hayward, 2014; Puterman et al., 2016; Yang et al., 2016). One important aspect of lifecycle experiences is childbearing. Although several mechanisms are proposed to explain the biological and social pathways between fertility and health, and specifically longevity, the direction of effects are a-priori unknown and remain an empirical question. Evolutionary theories of aging posit a trade-off between reproduction and longevity. Therefore, reproduction competes for resources that could otherwise be utilized in body maintenance and repair. Based on the so-called "*disposable soma theory*", the body investment in somatic maintenance can be hindered by childbearing, leading to faster aging (Hawkes, 2003; Kaptijn et al., 2015; Thomas et al., 2000; Westendorp & Kirkwood, 1998). Childbearing and childrearing can also be a stressful process with economic burden and financial distress. Based on the so-called "*maternal depletion*" theory, the cumulative demands and rigors of fertility and childrearing over the years has delayed impacts on health and raises risks of mortality. All these factors contribute to adverse health later in life, resulting in lower longevity (Currie, 2009; Elo et al., 2014; Fagundes et al., 2013; Graham et al., 2006). On the other hand, parenthood could result in changes in health behavior such as smoking cessation to adopt with social norms. Further, childbearing may also provide old age assistance and mental support for the elderly population. Social programs to support poor families with children such as Earned Income Tax Credit may also benefit parents by providing fertility-tied transfers. Thus, these pathways suggest a positive fertility-longevity relationship (Fenelon & Preston, 2012; Holt-Lunstad et al., 2010; Salm, 2011).

Studies that empirically explore the later-life health impact of fertility find nonlinear effects that usually reveal a U-shape pattern with positive effects for lower fertility families and negative impacts at the higher levels of fertility. Grundy & Tomassini (2005) employ data from England and Wales and find that women with no child and with 5-and-more children have elevated mortality risks than other women. Spence (2008) fails to find an effect of child-parity on later-life physical and psychological health. Yang et al. (2022) reports a U-shape relationship between fertility history and cognitive functioning among older adults in China. While the association is positive for individuals with 0-3 children, cognitive functioning declines for those with 5-and-more children. Hank (2010) employ data from East and West Germany and finds mixed evidence of fertility history for health and longevity at older ages. He shows positive associations for parents in West Germany and negative association in East Germany, suggesting that social and institutional pathways may be more important determinants than biological mechanisms. Kaptijn et al. (2015) use historical data from Netherlands and fail to find an association between fertility and longevity among women. Hsu et al. (2021) employ genealogical data across several European countries spanning 16$^{th}$ -20$^{th}$ centuries and find a negative fertility-mortality association that weakens across consecutive cohorts. Dribe (2006) employs data from Sweden and shows that the negative fertility-longevity relationship holds among women but not men, suggesting the relevance of biological pathways. Barclay et al. (2016) use data from Sweden for cohorts born between 1932 – 1960 and implement sibling fixed effects to minimize confounding factors and examining the effects of parity on mortality. They find that childless parents and parents with more than 4 children reveal elevated mortality. Einiö et al. (2016) find that all-cause mortality is higher among Finnish men and women with 0-1 children compared with 2-child parents. Several other papers using European data also provide surprisingly mixed results with no clear conclusions (Alter et al., 2007;

Doblhammer, 2010; Kravdal et al., 2012; Poulain et al., 2016). Two studies examine later-life health effects of family structure. Henretta (2007) uses Health and Retirement Study (HRS) with a focus on men and women born between 1931-1941 and observed at 1994. While he finds negative impacts of teenage pregnancy and nonmarital childbearing for health in old age, his results do not provide any evidence of the association between children ever born and later-life health and mortality. Barclay et al. (2020) examine the impacts of parental and child's death on mortality of women in the Utah Population Database. They find negative effects of bereavement on longevity. Although family size and fertility history are not their main variable of interest, they find suggestive gains of higher fertility for mortality outcomes. The findings of the latter studies are difficult to interpret due to the confounding influence of other unaccounted and unobserved factors. Moreover, the direct impacts of fertility-longevity association using US data remains limited to studies with small sample sizes and cross-sectional analyses. Our paper aims to extend this literature and fills these gaps.

The current study re-evaluates the fertility-longevity relationship using newly released Social Security Administration death records linked to the full-count 1940 census. The linked data allows us to observe family structure in mid-life and longevity in later-life for hundreds of thousands of individuals.[4] We tackle the endogeneity and selection issues by exploiting an instrumental variable (IV) strategy based on the sex composition of the first two children in the household. Our choice of instrument relies on parental preference towards having a sex-mix portfolio of children rather than a same-sex composition, hence they are more likely to have a third child if the gender of the first two is the same (Angrist & Evans, 1996; Bonsang & Skirbekk, 2022;

[4] This is in sharp contrast with the sample size of previous works, e.g., Henretta (2007) employs a sample of about 4,000 observations.

Cools & Kaldager Hart, 2017; Nguyen, 2019; Westoff & Potter, 2015). We show that the gender composition of children is a strong predictor of future fertility. The ordinary least square (OLS) and two-stage least-square instrumental-variable (2SLS-IV) strategies both provide a negative fertility effect on later-life longevity. We find IV estimates that are substantially larger than OLS estimates, suggesting that other confounding factors, such as selection into parenthood, underestimate the cross-sectional OLS results. Further, the coefficients for men are slightly smaller than women and statistically insignificant, suggesting that both socio-biological pathways are relevant. For women, an additional child is associated with a significant reduction of about 5 months in longevity.

We should note that although the 2SLS-IV estimates plausibly account for selections and confounders, it produces Local Average Treatment Effect (LATE) estimates and limits the external validity of the findings across the entire distribution of fertility. More specifically, the LATE estimates provide support of the impacts of having more than two children versus having only two children since the instrument is based on sex composition of the first two children. Therefore, we cannot infer similar interpretations for similar changes for parents with lower number of children.

Our paper contributes to two strands of research. First, to our knowledge, this is the first study to establish a causal link between fertility and later-life longevity using US data. Prior work employed small sample sizes and attempted to account for confounders by including other observable measures (Barclay et al., 2020; Henretta, 2007). In this work, we implement an IV approach and find negative fertility-mortality associations. Second, our work also contributes to the ongoing and growing literature on the influence of lifecycle exposures and later-life health outcomes.

The rest of the paper is organized as follows. Section 2 discusses the econometric method. Section 3 introduces data sources. Section 4 reviews the results. Finally, section 5 concludes the paper.

## 2. Econometric Method

Estimates of fertility-mortality associations could be confounded for various reasons. For instance, individuals may expect future rises in income and wealth and respond accordingly by changing their childbearing behavior as studies point to income-fertility relationship (Becker, 1960; Becker & Lewis, 1973; Black et al., 2013; Lindo, 2010; Willis, 1973). On the other end, income and wealth are important contributors to longevity (Chetty et al., 2016; B. Kim & Ruhm, 2012). Another potential source of endogeneity is regarding innate determinants of education and skill formation which are correlated with their desired family size (Brand & Davis, 2011; Cygan-Rehm & Maeder, 2013; McCrary & Royer, 2011). Moreover, there is evidence of positive impacts of education on longevity (Buckles et al., 2016; Lleras-Muney, 2005; Meghir et al., 2018). Another problem is the reverse causality concern, where changes in expected longevity affect parents' human capital, marriage, and fertility decisions (Cervellati & Sunde, 2005, 2013; Chen, 2009; Soares, 2005). To tackle these issues, we rely on the fact that the distribution of children's gender is close to random. Moreover, it is relatively well documented that families tend to prefer a mixed-gender profile of children. They are more likely to consider having additional children if their first two children are of the same sex (Angrist & Evans, 1996; Bonsang & Skirbekk, 2022; Cools & Kaldager Hart, 2017; Nguyen, 2019; Westoff & Potter, 2015). We exploit these two facts and posit that the variations in fertility caused by sex composition of the first two children are arguably exogenous to other determinants of health and longevity. Therefore, we employ a same-sex instrumental variable in the following forms:

$$Fertility_{ib} = \alpha_0 + \alpha_{11}BB_i + \alpha_{12}GG_i + \alpha_2 X_i + \gamma_b + \epsilon_{ib} \quad (1)$$

$$DA_{ib} = \beta_0 + \beta_1 \widehat{Fertility_{ib} + \beta_2} X_i + \phi_b + \varepsilon_{ib} \quad (2)$$

Where $DA$ and $Fertility$ are death-at-age and fertility of parent $i$ who belongs to birth cohort $b$. Fertility in these equations is a continuous variable measuring the number of children of an individual. The first stage instruments are $BB$ which is a dummy indicating the first two child are boys and $GG$ which equals one if the first two are girls and zero otherwise.[5] In $X$, we include race and ethnicity dummies. We specifically avoid including education and other socioeconomic measures as they are likely endogenously determined by fertility decisions.[6] However, later in robustness analyses, we show that the results are surprisingly insensitive to including these measures. Birth cohort fixed effects control for cross-cohort secular changes in longevity. They are represented by $\gamma$ and $\phi$ in the first and second stage, respectively. Finally, $\epsilon$ and $\varepsilon$ are error terms of the first and second stage, respectively. We cluster standard errors at birth-year level to account for serial correlations in longevity across cohorts.

## 3. Data and Sample Selection

The primary source of data is death records of the Social Security Administration that are linked to the full-count 1940 census. Specifically, we employ Numerical Identification (Numident) data extracted from the CenSoc Project (Goldstein et al., 2021). The Numident data covers records

[5] In Appendix D, we show the results of regressions that employ these two instruments separately. We find heterogenous results based on the sex composition of children. The negative impacts of fertility on maternal longevity are primarily concentrated in estimations that use a boy-boy instrument and the negative effects of fertility on paternal longevity comes from estimations that use a girl-girl instrument.

[6] In Appendix E, we explore the heterogeneity by educational attainments and find that the negative impacts of education on longevity are primarily concentrated among low educated mothers and high educated fathers.

of males and females who died between 1988-2005.[7] Although the data sources are limited in death window coverage, it has two primary advantages that allow for a unique and unprecedented setting to explore the fertility-mortality relationship. First, the data is linked at the individual-level to the full-count 1940 census, which allows access to a wide array of family characteristics and family structure information. Second, the initial sample covers millions of observations and allows for enough variation and number of observations even after restrictive sample selections.

The 1940 census data is extracted from Ruggles et al. (2020). We merge this with the DMF data, which restricts the sample to males only with the advantage of having more death years coverage. We build another subsample for females only and merge it with the Numident data. We restrict both samples to those with at least two children in the household since our instrument is based on sex composition of the first two offspring. Further, we limit the sample to those aged 25-45 to allow enough years to have two children and to provide a better proxy for completed fertility.[8] We further limit the sample to parents whose eldest child is at most 18 years old to avoid cross-household migration of children. Moreover, the sample is restricted to individuals with at least two children.

Table 1 reports summary statistics of the final samples. The average age-at-death is 1044.7 (1054.7) months for fathers (mothers) sample. For both samples, we focus on birth cohorts of 1895-1915 who died between 1988-2005. Compared with mothers' sample, the fathers' sample has relatively higher share of whites (0.95 versus 0.93) and lower share of blacks (0.045 versus 0.067). Compared with fathers' sample, mothers' sample contain individuals with higher number

[7] One concern regarding the data is truncation issues. In Appendix B, we employ an alternative data source and provide empirical evidence that truncation may induce a slightly upward bias in our results and the larger effects are observed at older ages.
[8] This age restriction limits birth cohorts to those of 1895-1915. In Appendix A, we show the robustness of the results to narrower birth cohorts, e.g., 1900-1915 and 1900-1910.

of children (3.21 versus 2.76). However, both samples reveal almost identical values for boy-boy and girl-girl composition of first two children (0.27 and 0.24, respectively).

## 4. Results

### 4.1. Endogeneity Concerns

One criterion to evaluate the validity of our instruments is to explore the exclusion restriction assumption. This assumption posits that sex composition of children is orthogonal to other observable and unobservable determinants of parents' longevity. In this section, we explore the associations between instruments and a series of observable characteristics of parents as an indirect test of exclusion restriction. Specifically, we use regressions of equation 1 and replace the outcome with observable parental characteristics. We report the results in Table 2 and Table 3 for mothers and fathers samples, respectively. There are small differences by race: white mothers are more (less) likely than other racial groups to have a boy-boy (girl-girl) composition of children. However, the associations are very small in magnitude. Compared with the mean of the outcome, the coefficients represent a 0.2-0.3 percent change. Moreover, these significant associations are not consistent with other outcomes. When we look at maternal education dummies, the results do not provide significant associations, which further lends to the validity of selected instruments.

In Table 3, we look at associations with fathers' characteristics. We do not find consistent, robust, and significant correlations with race, ethnicity, education, and socioeconomic status dummies. Although some coefficients are statistically significant, they remain economically small and inconsistent with the overall picture of the table. Since we fail to find associations with observable features, we argue that it is unlikely that there are discernible correlations between instruments and unobservables (Altonji et al., 2005; Fletcher et al., 2021).

Another concern is the selection of individuals into the final sample. If there are certain characteristics among parents in the final sample that are correlated with their selection from the original population into the final sample as well as with their childbearing decisions, then the effects could be capturing those features and mislead the interpretations. We empirically test whether successful merging between the original population and the final sample is correlated with the number of children. In so doing, we implement sample selections of section 3 for the full population of 1940 census for males and females separately. We then merge these two samples with mothers and fathers samples. We generate a successful merging dummy that equals one if two records are merged and zero otherwise. We then implement regressions using the 2SLS-IV estimation strategy derived from equations 1 and 2. We report the results in panels A and B of Table 4. The first column reports simple OLS correlations, and the second column reports the IV results. While the OLS estimates suggest significant associations between probability of merging and number of children, this issue is not present in the IV estimates. Therefore, the model specification of IV accounts for endogenous factors that results in selective merging into the final sample.[9]

### 4.2. Main Results

We report the first stage results in Table 5. In the mothers sample, we find that having a boy-boy composition for the first two children raises the total number of children by 0.06. The coefficient for girl-girl reveals an insignificant and small decrease of 0.002 (column 1). In column 2, we show the results for a dummy variable indicating the number of children is 3 or greater. The

[9] In Appendix C, we use all observations from the original cohorts in combination with observations in the final sample of the paper to implement two-sample two-stage least-square estimations. We observe comparable estimates to the main results of the paper.

coefficients of boy-boy and girl-girl suggests 2.2 and 1.2 percentage-points rise in the probability of having more than 3 children, off a baseline of 0.56.

In columns 3-4, we replicate the results for the fathers sample. The coefficients of boy-boy and girl-girl sex composition suggest significant increases in fertility by about 0.03 and 0.04 children, respectively. We also observe an increase in the likelihood of having 3 or more children by about 1.9 and 2.8 percentage-points, off a mean of 0.42.

The main results of the paper are reported in Table 6. Based on OLS estimates, an additional child is associated with a reduction of about 1.3 months in longevity of women (column 1). However, the IV estimate suggests an effect of 4.9 months (column 2). This sharp discrepancy in marginal effects implies that the endogeneity biases the OLS coefficients towards zero. Although this number appears small with respect to the outcome's mean (a 0.5 percent reduction), we argue that it is a relatively large effect as we compare it with other determinants of longevity. For instance, Fletcher & Noghanibehambari (2021) explore the effects of college openings on college education and longevity. They estimate a treatment-on-treated effect of college education on longevity of about 1 additional year of life. Halpern-Manners et al. (2020) implement twin fixed effect strategy to examine the education-longevity nexus. They find that an additional year of education leads to about 3.4 months additional lifespan. Therefore, the negative effects of an additional child is equivalent to 41 percent of positive impacts of college education and might be offset by 1.4 more years of schooling.

Chetty et al. (2016) examine the income-longevity relationship and document that each additional income percentile (an increase of $8,000 from the sample mean, in 2020 dollars) is correlated with about 1.6 months higher longevity. Therefore, the impact of an additional child is equivalent to roughly a drop of $24.5K in household income. Dribe (2006) examines the fertility-longevity relationship in 19th century Sweden and documents that women with 4-5 children compared with those with 0-1 children live about 3.5 years shorter lives. Our findings imply that a 4-child difference is associated with about 1.6 years shorter lives. The discrepancy in these two estimates can be explained by weakening fertility-longevity relationships over cohorts and centuries that are suggested by several studies (Hsu et al., 2021; Kaptijn et al., 2015).

Columns 3-4 of Table 6 report the results for fathers sample. While the IV effect is still larger than OLS estimate, it is statistically insignificant. The IV estimate of fathers' longevity is also smaller than that of mothers', suggesting that biological mechanisms are likely an important driver in fertility-longevity equation of mothers sample.

### 4.3. Robustness Checks

We explore the robustness of the results to alternative specifications in Table 7. Panels A and B report the results for mothers and fathers samples, respectively. Within each panel, the first sub-panel reports the OLS estimates and the second sub-panel shows the IV results. For comparison purposes, we start by reporting the baseline results of Table 6 in the first column of Table 7.

In the second column, we add county fixed effects to account for unobserved time-invariant features of the location of residence. The IV effects of mothers remain robust in magnitude although statistically becomes insignificant which limits further interpretations. The OLS effects of both samples, however, reveal a very similar effect compared with those of column 1. In column 3, we add birth-state by birth-year fixed effects to account for all time-variant characteristics of birth-states that might affect individuals' longevity. For instance, state-level economic conditions as well as many policies such as compulsory schooling, child labor laws, suffrage law movements, and prohibition movements are documented to influence longevity outcomes (Fletcher, 2015;

Lleras-Muney, 2005; Noghanibehambari & Fletcher, 2023, 2021; Schmitz & Duque, 2022). We observe marginal effects that are quite comparable to those of column 1.

Another source of concern in the results is the potential influence of seasonality. Studies document the relevance of season of birth in predicting later-life mortality outcomes (Vaiserman, 2021; Vaiserman et al., 2002). Further, there are seasonality in mortality that may confound longevity equations (Marti-Soler et al., 2014; Seretakis et al., 1997; Simmerman et al., 2009). In column 4, we add birth-month and death-month fixed effects to fully absorb the unobserved factors related to seasonality. The estimated effects are virtually identical to those of column 1.

In columns 5-6, we explore the sensitivity of the results to the functional form. In column 5, we replace the outcome with a dummy variable indicating age-at-death being greater than 88 years, the median of age-at-death in the final sample. We observe a reduction of 4.2 (mean=0.45) and 7.2 (mean=0.38) percentage-points for mothers and fathers samples, respectively. In column 5, we replace the outcome with the log of age-at-death. The IV estimate of mothers suggests a significant reduction of about 0.44 percent. This change is very similar to the implied percentage change with respect to the mean of the outcome (0.46 percent) as in column 1. The IV estimate of fathers sample suggests an insignificant reduction of 0.3 percent, which is almost identical to 0.3 percent extracted from column 1.

Finally, we explore the statistical significance of the results to alternative standard errors correction methods. Specifically, we use Huber-White robust standard errors to account for heteroscedasticity in error terms. . The results, reported in column 7, reveal a similar pattern as those reported in column 1.

## 5. Conclusion

Understanding the fertility-longevity relationship is important for policies that contain components of family planning and family structure. For instance, policies such as Earned Income Tax Credit may include sub-optimal levels of payments in the presence of negative fertility-longevity relationships. Moreover, if there are later-life negative health effects of fertility, we need to understand the mechanisms to mitigate these impacts with policy instruments. These policies might not be impactful if the results imply primarily biological pathways. Therefore, it is important for both family planning purposes and for policy designs to understand the associations between fertility history and later-life health and longevity. This paper aimed to do so.

We employed Social Security Administration death records linked with the full-count 1940 census and implemented an IV estimate based on sex composition of children to rely on the plausibly exogenous changes resulted from sex-mix-children preference of parents. We find that an additional child is associated with about 5 months reduction in longevity of women. Our results suggested smaller and statistically insignificant reductions of about 3.3 months for men. This difference in effects suggests that part of the fertility-longevity relationship could operate through biological mechanisms.

To put these effects into perspective, we compare the influence of fertility change over the 20$^{th}$ century on observed life expectancy rise. Between 1900-2000, total fertility rate decreased from 4.2 to 2.0 children per women (Boldrin et al., 2015). During the same period, life expectancy experienced a change of 29.6 years, from 47 to 76.6 years (O'Neill, 2021). Among women, the longevity rises induced by reductions in fertility is roughly 11 months. This increase in longevity is about 3.1 percent of overall rise in life expectancy over the 20$^{th}$ century.

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## Tables

**Table 1 - Summary Statistics**

| | Mean | SD | Min | Max |
|---|---|---|---|---|
| ***Fathers Sample:*** | | | | |
| Age at Death (Months) | 1044.7258 | 67.3059 | 865 | 1313 |
| Birth Year | 1907.0011 | 4.8687 | 1895 | 1915 |
| Death Year | 1994.0811 | 4.5459 | 1988 | 2005 |
| White | .9513 | .2153 | 0 | 1 |
| Black | .0455 | .2085 | 0 | 1 |
| Hispanic | .0173 | .1304 | 0 | 1 |
| Education: None | .0093 | .0961 | 0 | 1 |
| Education: Less than High School | .5116 | .4999 | 0 | 1 |
| Education: High School | .3632 | .4809 | 0 | 1 |
| Education: College | .1088 | .3114 | 0 | 1 |
| Education: Missing | .0164 | .1271 | 0 | 1 |
| Socioeconomic Score Missing | .0259 | .1589 | 0 | 1 |
| Socioeconomic Score Below Median | .5467 | .4978 | 0 | 1 |
| Socioeconomic Score Above Median | .4274 | .4947 | 0 | 1 |
| Number of Children | 2.7649 | 1.1782 | 2 | 9 |
| First Two Children Male | .2649 | .4413 | 0 | 1 |
| First Two Children Female | .2406 | .4274 | 0 | 1 |
| Observations | | 333,690 | | |
| ***Mothers Sample*** | | | | |
| Age at Death (Months) | 1054.6675 | 70.59 | 865 | 1314 |
| Birth Year | 1907.8728 | 5.227 | 1895 | 1915 |
| Death Year | 1995.7796 | 4.8807 | 1988 | 2005 |
| White | .9302 | .2548 | 0 | 1 |
| Black | .0672 | .2504 | 0 | 1 |
| Hispanic | .0115 | .1068 | 0 | 1 |
| Education: None | .0062 | .0788 | 0 | 1 |
| Education: Less than High School | .4905 | .4999 | 0 | 1 |
| Education: High School | .412 | .4922 | 0 | 1 |
| Education: College | .0793 | .2702 | 0 | 1 |
| Education: Missing | .0182 | .1337 | 0 | 1 |
| Number of Children | 3.2134 | 1.5235 | 2 | 9 |
| First Two Children Male | .2728 | .4454 | 0 | 1 |
| First Two Children Female | .2363 | .4248 | 0 | 1 |
| Observations | | 418,883 | | |

**Table 2 - Balancing Tests of Mothers Sample**

| | *Outcomes:* | | | | | | |
|---|---|---|---|---|---|---|---|
| | White | Hispanic | Education: None | Education: Less than High School | Education: High School | Education: College | Education: Missing |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| Boy-Boy | .00197** | .00077* | .00031 | .00264 | -.00201 | -.00148 | .00085** |
| | (.00091) | (.00042) | (.0003) | (.00214) | (.00199) | (.00116) | (.00033) |
| Girl-Girl | -.00363*** | .00111*** | -.00012 | -.00233 | .00126 | .00114 | -.00007 |
| | (.00098) | (.00027) | (.00035) | (.00195) | (.00162) | (.0013) | (.00039) |
| Observations | 408446 | 408446 | 408446 | 408446 | 408446 | 408446 | 408446 |
| R-Squared | .28325 | .3205 | .18001 | .20477 | .18061 | .13836 | .39799 |
| Mean DV | 0.930 | 0.011 | 0.006 | 0.489 | 0.414 | 0.079 | 0.018 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively.
*** p<0.01, ** p<0.05, * p<0.1

**Table 3 - Balancing Test of Fathers Sample**

| | *Outcomes:* | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | White | Hispanic | Education: None | Education: Less than High School | Education: High School | Education: College | Education: Missing | Socioeconomic Score Below Median | Socioeconomic Score Above Median | Socioeconomic Score Missing |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
| Boy-Boy | -.00062 | .00021 | -.00091* | -.00211 | .00343 | -.0011 | -.00021 | -.00412** | .00387* | .00025 |
| | (.00084) | (.00056) | (.00045) | (.00167) | (.00215) | (.00182) | (.00042) | (.00194) | (.00204) | (.0008) |
| Girl-Girl | -.0024** | .00044 | -.00024 | .00536** | -.0029 | -.00207 | -.00039 | -.00024 | -.00083 | .00107 |
| | (.00088) | (.00066) | (.00042) | (.00217) | (.00197) | (.00127) | (.00066) | (.00169) | (.00175) | (.00103) |
| Observations | 323415 | 323415 | 323415 | 323415 | 323415 | 323415 | 323415 | 323415 | 323415 | 323415 |
| R-Squared | .32852 | .18503 | .1601 | .18155 | .16594 | .11471 | .32767 | .21636 | .21189 | .11084 |
| Mean DV | 0.954 | 0.017 | 0.009 | 0.506 | 0.367 | 0.110 | 0.016 | 0.540 | 0.434 | 0.026 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively.
*** p<0.01, ** p<0.05, * p<0.1

**Table 4 - Exploring Endogenous Numident-Census Merging**

| | *Outcome: Successful Merging between Numident and Original Cohorts of 1940 Census* | |
|---|---|---|
| | OLS | 2SLS-IV |
| | (1) | (2) |
| ***Mothers Sample:*** | | |
| Number of Children | .00498*** | .00101 |
| | (.00005) | (.00116) |
| Observations | 4570588 | 4570588 |
| R-Squared | .01141 | .00155 |
| Mean DV | 0.031 | 0.031 |
| | | |
| ***Fathers Sample:*** | | |
| Number of Children | -.00338*** | -.00238 |
| | (.00011) | (.00638) |
| Observations | 4356142 | 4356142 |
| R-Squared | .01961 | .00234 |
| Mean DV | 0.068 | 0.068 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

**Table 5 - First Stage Effects**

| | ***Samples and Outcomes:*** | | | |
|---|---|---|---|---|
| | Mothers Sample DV: Number of Children | Mothers Sample DV: Number of Children≥3 | Fathers Sample DV: Number of Children | Fathers Sample DV: Number of Children≥3 |
| | (1) | (2) | (3) | (4) |
| Boy-Boy | .06429*** | .0229*** | .02893*** | .01945*** |
| | (.00541) | (.00181) | (.00468) | (.00204) |
| Girl-Girl | -.00228 | .01221*** | .03713*** | .0284*** |
| | (.00547) | (.00197) | (.00484) | (.00241) |
| Observations | 418883 | 418883 | 333690 | 333690 |
| R-Squared | .14623 | .08032 | .11214 | .06885 |
| Mean DV | 3.213 | 0.560 | 2.765 | 0.425 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

**Table 6 - Main Results**

| | *Outcome: Age at Death (Months), Samples:* | | | |
|---|---|---|---|---|
| | Mothers Sample | | Fathers Sample | |
| | OLS | 2SLS-IV | OLS | 2SLS-IV |
| | (1) | (2) | (3) | (4) |
| Number of Children | -1.34093*** | -4.95652** | -.71444*** | -3.31507 |
| | (.10301) | (2.51389) | (.09125) | (4.21749) |
| Observations | 418883 | 418883 | 333690 | 333690 |
| R-Squared | .38502 | .37982 | .39796 | .39612 |
| Mean DV | 1054.668 | 1054.668 | 1044.726 | 1044.726 |
| %Change | -0.127 | -0.470 | -0.068 | -0.317 |
| First-Stage F-Stat | | 28.235 | | 24.940 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

**Table 7 - Robustness Checks**

| | *Outcome: Age at Death (Months)* | | | | | | |
|---|---|---|---|---|---|---|---|
| | From Table 6 | Adding County FE | Adding Birth-State by Birth-Year FE | Adding Birth-Month and Death-Month FE | Outcome: Age at Death > 88 (Median) | Outcome: Log Age at Death | Clustering SE at Birth-Year Level |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| ***Panel A. Mothers Sample*** | | | | | | | |
| ***Panel A1. OLS*** | | | | | | | |
| Number of Children | -1.34093*** | -1.34513*** | -1.29546*** | -1.3834*** | -.00847*** | -.00125*** | -1.34093*** |
| | (.10301) | (.11102) | (.10385) | (.10546) | (.00109) | (.0001) | (.05805) |
| Observations | 418883 | 418879 | 418867 | 418883 | 418883 | 418883 | 418883 |
| R-Squared | .38502 | .39042 | .38722 | .38703 | .25616 | .38129 | .38502 |
| Mean DV | 1054.668 | 1054.668 | 1054.665 | 1054.668 | 0.445 | 6.959 | 1054.668 |
| ***Panel A2. 2SLS-IV*** | | | | | | | |
| Number of Children | -4.95652** | -4.2353 | -4.52299 | -4.97904** | -.04167* | -.00438* | -4.95652* |
| | (2.51389) | (2.97085) | (2.71905) | (2.53322) | (.0232) | (.00233) | (2.95598) |
| Observations | 418883 | 418879 | 418882 | 418883 | 418883 | 418883 | 418883 |
| R-Squared | .37982 | -.00395 | -.0054 | .38189 | .24731 | .37703 | .37982 |
| Mean DV | 1054.668 | 1054.668 | 1054.667 | 1054.668 | 0.445 | 6.959 | 1054.668 |
| ***Panel B. Fathers Sample*** | | | | | | | |
| ***Panel B1. OLS*** | | | | | | | |
| Number of Children | -.71444*** | -.73455*** | -.75074*** | -.76472*** | -.00067*** | -.71444*** | -.80242*** |
| | (.07554) | (.07821) | (.07731) | (.07541) | (.00007) | (.09124) | (.08235) |
| Observations | 333690 | 333673 | 333526 | 333690 | 333690 | 333690 | 443790 |
| R-Squared | .39796 | .40465 | .40256 | .40014 | .39881 | .39796 | .43091 |
| Mean DV | 1044.726 | 1044.727 | 1044.714 | 1044.726 | 6.949 | 1044.726 | 1022.805 |
| ***Panel B2. 2SLS-IV*** | | | | | | | |
| Number of Children | -3.31507 | -3.31448 | -3.64755 | -3.42388 | -.0723 | -.00308 | -3.31507 |
| | (4.21749) | (4.92191) | (4.30877) | (4.19242) | (.04396) | (.004) | (5.42865) |
| Observations | 333690 | 323415 | 333526 | 333690 | 333690 | 333690 | 333690 |
| R-Squared | .39612 | -.00187 | -.00334 | .39821 | .23318 | .3971 | .39612 |
| Mean DV | 1044.726 | 1043.932 | 1044.714 | 1044.726 | 0.378 | 6.949 | 1044.726 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

## Figures

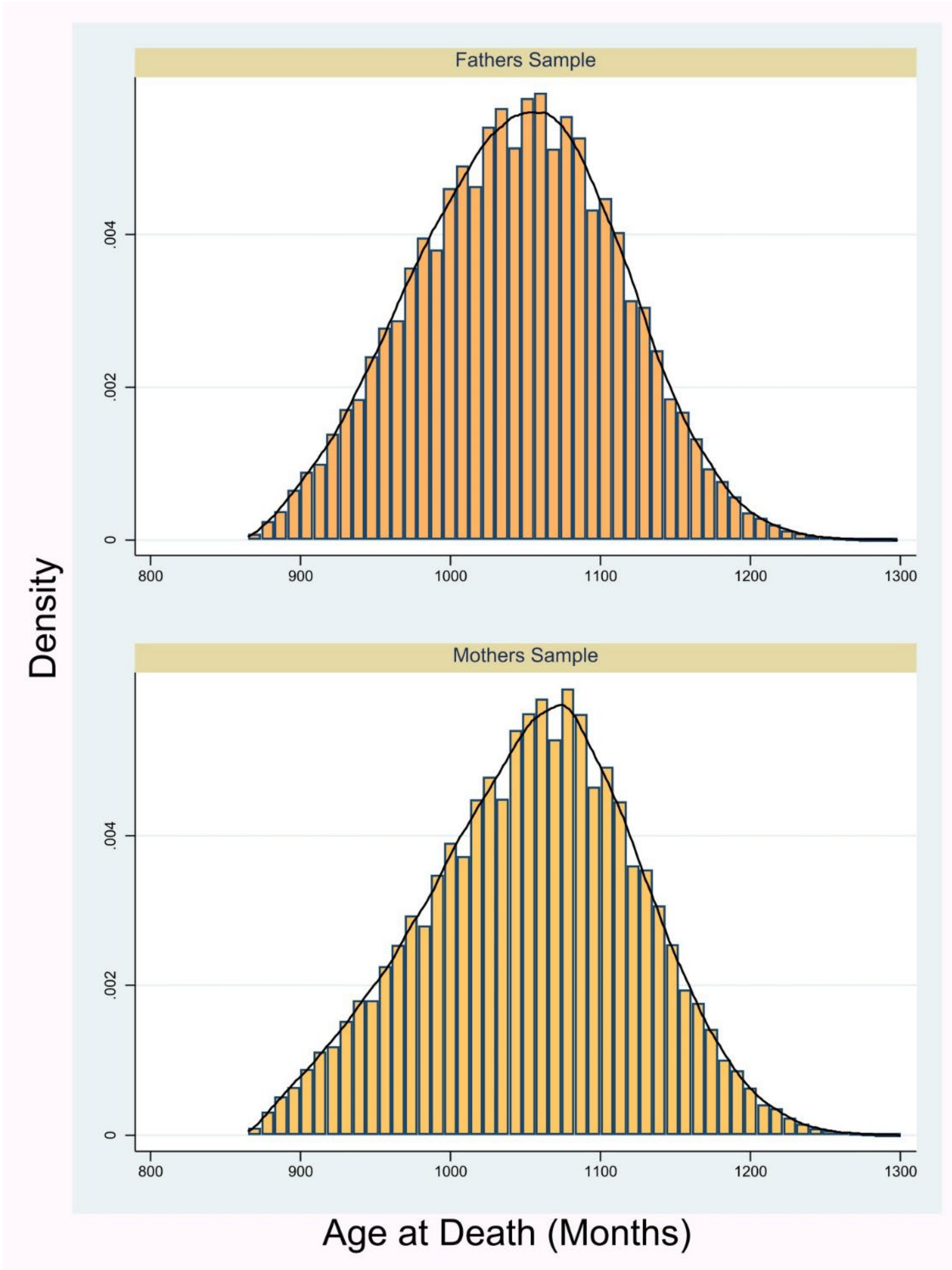


**Figure 1 - Density Distribution of Age at Death**

## Appendix A

The analysis in the main results contain birth cohorts of 1895-1915. This selection is due to age selection of 1940 census to avoid very old people whose children may have left the household in 1940 and very young people with uncompleted fertility. However, one concern is the presence of earlier birth cohorts who, between 1988-2005, are much older than earlier birth cohorts. In this appendix, we restrict the sample to cohorts of 1900-1915 and 1900-1910 and replicate the main results. These estimates are reported in Appendix Table A-1 and Appendix Table A-2 for mothers and fathers samples, respectively. For mothers sample, we observe slightly smaller coefficients. The 2SLS estimate for 1900-1915 cohorts become statistically insignificant but still comparable to the main results in magnitude. The results of 1900-1910 cohorts reveal larger effects but statistically insignificant. For fathers sample, we observe similar pattern but both coefficients are statistically insignificant.

**Appendix Table A-1 - Replicating the Main Results of Mothers Sample for Narrower Birth Cohorts**

| | ***Outcome: Age at Death (Months), Samples:*** | | | |
|---|---|---|---|---|
| | Mothers Sample 1900-1915 Cohorts | | Mothers Sample 1900-1910 Cohorts | |
| | OLS | 2SLS-IV | OLS | 2SLS-IV |
| | (1) | (2) | (3) | (4) |
| Number of Children | -1.47455*** | -3.54446 | -1.47945*** | -5.48441 |
| | (.09875) | (3.60607) | (.11175) | (4.00226) |
| Observations | 386003 | 386003 | 227128 | 227128 |
| R-Squared | .31194 | .31023 | .15672 | .14557 |
| Mean DV | 1047.621 | 1047.621 | 1074.445 | 1074.445 |
| %Change | -0.141 | -0.338 | -0.138 | -0.510 |
| First-Stage F-Stat | | 23.582 | | 24.433 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.

*** p<0.01, ** p<0.05, * p<0.1

**Appendix Table A-2 - Replicating the Main Results of Fathers Sample for Narrower Birth Cohorts**

| | ***Outcome: Age at Death (Months), Samples:*** | | | |
|---|---|---|---|---|
| | Fathers Sample 1900-1915 Cohorts | | Fathers Sample 1900-1910 Cohorts | |
| | OLS | 2SLS-IV | OLS | 2SLS-IV |
| | (1) | (2) | (3) | (4) |
| Number of Children | -.79928*** | -2.86053 | -.79778*** | -3.92117 |
| | (.10057) | (4.29249) | (.09899) | (4.97654) |
| Observations | 307468 | 307468 | 215597 | 215597 |
| R-Squared | .31777 | .31664 | .19249 | .18837 |
| Mean DV | 1037.741 | 1037.741 | 1056.587 | 1056.587 |
| %Change | -0.077 | -0.276 | -0.076 | -0.371 |
| First-Stage F-Stat | | 38.349 | | 26.828 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

## Appendix B

One concern in the main results is truncation of the final sample. Death window is limited to 1988-2005 and is truncated from both right and left. Life expectancy in US increased from 41.2 to 54.1 years from 1895 to 1915 (birth cohorts of the final sample). For an average person in the sample to reach post-2005 years, they should live at least 105 years, almost twice their average life expectancy. Therefore, the right truncation is less of a concern in this setting. However, left truncation could bias the results in two ways. It could be the case that the mortality effects appear in younger ages and our results underestimate the true overall impacts. It is also possible that there are mortality benefits for early old age years (prior to 1988) and inclusion of earlier deaths drive the coefficients toward zero. We address this concern using an alternative data source that covers several earlier death years. Specifically, we employ Social Security Death Master Files (DMF) extracted from Goldstein et al. (2021). The DMF data reports deaths to male individuals only but covers death years of 1975-2005. We implement the same sample restrictions as in the paper and replicate the main results in Appendix Table B-1. The OLS estimate of column 1 suggests a significant drop in longevity of men by about 2 months. This is considerably larger than the OLS effect of Numident (-0.7). However, we observe both an IV estimate that is smaller in magnitude relative to the Numident estimate, by about 20 percent. This fact suggests that truncation may overestimate our main results and inclusion of earlier death will reduce the magnitude of coefficients.

**Appendix Table B-1 - Replicating the Main Results of Fathers Sample Using DMF Data for Death Years 1975-2005**

| | ***Outcome: Age at Death (Months)*** | |
|---|---|---|
| | OLS | 2SLS-IV |
| | (1) | (2) |
| Number of Children | -1.93648*** | -2.65028 |
| | (.20839) | (4.06529) |
| Observations | 856970 | 856970 |
| R-Squared | .13147 | .13137 |
| Mean DV | 974.903 | 974.903 |
| %Change | -0.199 | -0.272 |
| First-Stage F-Stat | | 79.671 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

## Appendix C

In this appendix, we use the original cohorts of the 1940 census (after sample restrictions and before merging with Numident data) and implement two-sample two-stage least square (TS2SLS) estimation strategy. The outcome of the first stage comes from the calculated fertility of the 1940 census and the outcome of the second stage (longevity) is extracted from the sample of merged individuals with Numident data. The first stage effects are reported in Appendix Table C-1. We observe larger effects of boy-boy and a negative association between girl-girl and fertility for both samples.

The results of TS2SLS are reported in Appendix Table C-2. Among mothers, an additional child is associated with 4.7 months lower longevity, almost identical to the estimates of Table 6. However, the effect is statistically insignificant. Among fathers, we observe a slight increase in magnitude of about 15 percent relative to those of Table 6, though the coefficient remains statistically insignificant.

**Appendix Table C-1 - First Stage Effects of Two-Sample Two-Stage Least-Square Estimation**

| | ***Outcome: Number of Children*** | |
|---|---|---|
| | Mothers Sample | Fathers Sample |
| | (1) | (2) |
| Boy-Boy | .0861*** | .0322*** |
| | (.0016) | (.0013) |
| Girl-Girl | .0080*** | .0417*** |
| | (.0017) | (.0014) |
| Observations | 5,117,853 | 4,058,126 |
| R-Squared | 0.0330 | 0.0465 |
| Mean DV | 3.2819 | 2.8098 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

**Appendix Table C-2 – The Results of Two-Sample Two-Stage Least-Square Estimation**

| | *Outcome: Age at Death (Months)* | |
|---|---|---|
| | Mothers Sample | Fathers Sample |
| | (1) | (2) |
| Number of Children | -4.7404 | -3.8904 |
| | (2.9262) | (5.1927) |
| Observations | 5,117,853 | 4,058,126 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

## Appendix D

In the main analysis of the paper, we used an instrumental variable based the first two children having the same sex. In this appendix, we replicate the results using two alternative instruments for the first two children being boy (boy-boy instrument) and the first two children being girl (girl-girl instrument). We report the results in Appendix Table D-1. The first two columns show the results for the mothers sample and the second to columns report the results for the father sample. Panel A reports the results for the boy-boy instrument and panel B for the girl-girl instrument. For mothers, the boy-boy instrument provides a strong first stage and suggests a reduction in longevity of about 4.7 months for an additional child. This fact is consistent with studies that examine the effect of sex composition of children on maternal longevity (Harrell et al., 2008). Although the girl-girl instrument provides substantially larger coefficient, the F-statistics of the first stage is so small that makes the resulting coefficient unreliable.

Amount fathers, this pattern reverses. The boy-boy instrument has a much smaller F-statistics from the first stage than the girl-girl instrument (10 versus 29). However, the boy-boy instrument points to an insignificant increase in longevity for fathers. On the other hand, the estimation based on girl-girl instrument suggests 6.3 months reduction in longevity for an additional child, the coefficient that is slightly larger than the main results of the paper.

**Appendix Table D-1 - Replicating the Main Results Using Boy-Boy and Girl-Girl Instruments**

| | ***Outcome: Age at Death (Months), Samples:*** | | | |
|---|---|---|---|---|
| | Mothers Sample | | Fathers Sample | |
| | OLS | 2SLS-IV | OLS | 2SLS-IV |
| | (1) | (2) | (3) | (4) |
| ***Panel A. Boy-Boy IV*** | | | | |
| Number of Children | -1.34269*** | -4.72945* | -.71444*** | 2.38218 |
| | (.10325) | (2.54188) | (.09125) | (12.39909) |
| Observations | 418576 | 418576 | 333690 | 333690 |
| R-Squared | .38503 | .38046 | .39796 | .39535 |
| Mean DV | 1054.697 | 1054.697 | 1044.726 | 1044.726 |
| %Change | -0.127 | -0.448 | -0.068 | 0.228 |
| First-Stage F-Stat | | 31.217 | | 9.995 |
| ***Panel B. Girl-Girl IV*** | | | | |
| Number of Children | -1.34269*** | -32.28065** | -.71444*** | -6.25111 |
| | (.10325) | (16.20879) | (.09125) | (8.04738) |
| Observations | 418576 | 418576 | 333690 | 333690 |
| R-Squared | .38503 | .00399 | .39796 | .38962 |
| Mean DV | 1054.697 | 1054.697 | 1044.726 | 1044.726 |
| %Change | -0.127 | -3.061 | -0.068 | -0.598 |
| First-Stage F-Stat | | 4.546 | | 28.758 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.
*** p<0.01, ** p<0.05, * p<0.1

## Appendix E

In this appendix, we examine the heterogeneity in the results based on education. We replicate the two-stage least square estimations using the subsample of low educated (education less than high school) and high educated (education of at least high school) individuals. We report these results in Appendix Table E-1. The evidence is mixed among mothers and fathers. In the mothers sample, the negative effects are primarily concentrated among low educated mothers. However, in the fathers sample, the effects are confined to high educated fathers.

**Appendix Table E-1 - The Heterogeneity in the Results Based on Education**

| | ***Outcome: Age at Death (Months), Samples:*** | | | |
|---|---|---|---|---|
| | Mothers Sample | | Fathers Sample | |
| | Education < High School | Education ≥ High School | Education < High School | Education ≥ High School |
| | (1) | (2) | (3) | (4) |
| Number of Children | -5.97326 | -1.35794 | .44119 | -3.57353 |
| | (5.07586) | (3.56625) | (8.41988) | (7.10214) |
| Observations | 205404 | 213172 | 170713 | 162977 |
| R-Squared | .38468 | .37478 | .41858 | .37481 |
| Mean DV | 1055.696 | 1053.733 | 1047.076 | 1042.264 |
| %Change | -0.566 | -0.129 | 0.042 | -0.343 |
| First-Stage F-Stat | 23.291 | 20.269 | 18.550 | 16.988 |

Notes. Standard errors, clustered on birth year, are in parentheses. Regressions include birth year fixed effects. The variables boy-boy and girl-girl represents parents whose first two child are all boys and all girls, respectively. Regressions also include race and ethnicity dummies.

*** p<0.01, ** p<0.05, * p<0.1