Does Father Visitation Affect Child Health? The Role of Baby Looks and Maternal Earnings*

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Abstract

Whether a newborn child looks like the putative father is used to construct estimates of the impact of father visitation on the health of one-year old children. These estimates are not biased by the effect of a baby's initial health on father visits, unobserved maternal ability, or measurement errors, thereby eliminating endogeneity biases that plague existing studies. Analysis of data from the first two waves of the Fragile Family and Child Wellbeing study indicates an extra day (per month) of father visitation raises the chance of reported excellent child health by 3 percentage points. Moreover, the impact of father visitation is more important when maternal earnings are low.

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1 Introduction

Income inequality in the US increased substantially since the 1990's. Many dispute the underlying causes, but few can disagree that single-parent households tend to fall at the bottom of the distribution. Further, children in these households are at a disadvantage, which likely affects them throughout their lives.

On December 10, 2014, the White House held a summit on Early Education. The key recommendation emphatically endorsed investing in early childhood development (including health). According to one speaker, "[t]he way parents interact with their children, the amount of time they spend with them and the resources they have . . . greatly affect their children's potential for leading flourishing lives" (Heckman, 2014). Time inputs of both parents appear to be more important than money investments in producing quality children (Del Boca et al., 2014). In this regard, the prevalence of children currently living in low-income single mother homes is of concern.[1] These children are more likely to be food insecure (Nord et al, 2005), and have less access to parental investment especially from absentee fathers (Jones & Mosher, 2013). But, as will be explained, the impact of time, especially father's time on child health and well-being, has not been adequately studied. In this paper, we consider whether promoting non-resident father visitation is effective as an early intervention policy for improving child health. Such a policy is important given that child health is strongly linked to future education, adult health and eventual labor market success (Case et al., 2005; Currie, 2009; Campbell et al., 2014).

A priori, one would suspect the effect of a non-resident father's time is ambiguous. On the one hand, frequent father visits may provide greater parental time for care-giving and supervision, leading to better child health. On the other, a visiting father might stir up parental conflict or put further strain on mother's emotional strength and economic resources. These factors raise the mother's stress and could lead to depression (Slade, 2013), which in turn lowers her parenting quality. Moreover, since household income is important to child health (e.g. Case et al., 2002; Currie, 2009), the magnitude of the visitation-health effect may depend on the earnings of the single mother.

Too little is known about the effect of non-resident father visitation on early childhood health. Correlation-type multivariate analysis yields mixed results. For example, Menning & Stewart (2008) find non-resident father contact is related to a greater risk of obesity. However, Amato & Gilbreth (1999) find no relation between non-resident father contact and a child's general well-being. Similarly, Nepomnyaschy & Donnelly (2015) find no relation of non-resident father contact and a child's risk of injury. Yet, somewhat contradictory, other studies find that frequent contact is related to fewer food acquisition problems (Garasky & Stewart, 2007) and healthier eating habits (Stewart & Menning, 2009). Simply controlling for demographic and socioeconomic factors limits these studies' facility to control for endogeneity. As such, they do not establish a causal effect of father visitation on child health. Further, no attention is given to maternal income in moderating the causal effect of father visitation.

[1] In 2014, 17.4 million (25%) children under the age of 18 are living with mother only and 45% live below the poverty line (US Census Bureau, Tables C2 and C8).

Typically, there are three sources of endogeneity: reverse causality, omitted variables and measurement error. First, a newborn's health may deter a non-resident father's involvement rather than the reverse.[2] Ignoring such reverse causation creates an upward estimation bias. Second, a protective self-reliant and competent mother may discourage ('gatekeep') father engagement in child-rearing activities (Allen & Hawkins, 1999; Gaertner et al., 2007). Such omitted mother variables cause a downward bias. Third, a mother who views herself as the sole contributor to her child's health may be more likely to under-report a father's participation while a mother who merely wishes the father participated more might overstate father's involvement. So, a mother-reported father visitation measure may be measured with error, which can create a downward bias. The overall bias is ambiguous, making it impossible to advise on public policy.

Our study focuses on eliminating these endogeneity biases when estimating the effect of non-resident father visitation on child health. Our innovation is to incorporate a novel instrument for a father's time involvement to identify the causal effect. Specifically, we use reported father-child resemblance (whether the baby looks like the father), determined by both parents in separate and private interviews. To the best of our knowledge, this study is the first to use 'baby looks' as an instrumental variable (IV). As justification, several studies from evolutionary psychology, biology, and human and animal behavior (e.g. Platek et al., 2002, 2003, 2004; Volk & Quinsey, 2002, 2007; Alvergne et al., 2009, 2010) find experimental and empirical support that father-child resemblance predicts a (putative) father's time-investment. Furthermore, Alvergne et al. (2010) show that actual resemblance to the father is consistent with both parents reporting father-child facial similarity. Finally, we note that this IV should not violate the exclusion restriction that father-child resemblance not affect maternal time-investment, which is the case because by giving birth the mother already knows the child is hers.

We utilize the Fragile Families and Child Wellbeing data based on US families with one year old infants born out of wedlock who have a non-resident father. As in other studies, we use mother-reported child health as our primary outcome variable (e.g. Case et al., 2002; Currie & Stabile, 2003). Our study is the first to confirm on a national scale, in accord with evolutionary theory, that non-resident fathers react positively to baby looks. Typically, such fathers spend over 2 days (per month) longer when the child looks like him. Based on our IV estimate, an added visitation day raises the probability by 3 percentage points that a child is in excellent health. This result is robust when specific health outcomes (e.g. asthma attack episodes and emergency room visits) are used. Further, when employing a flexible varying-coefficient method, we find father visitation results in greater child health benefits as maternal earnings fall. As such, this paper presents a case for greater father time-investment in early childhood health capital, especially for children of low-earning single mothers. Thus, promoting father visitation is effective as an early intervention policy to reduce disparities in child health and thereby secure future educational and career success.

The remainder of the paper is structured as follows. The next section discusses baby looks in relation to paternal investment. In section 3, we describe in detail the data and define key variables. We then conduct estimation using the method outlined in section 4, report the

[2] Hawkins et al (2007) find this to be the case for adolescents with behavioral or academic problems.

results in section 5 and, in two subsequent sections, provide some checks on the robustness of our results and the validity of our IV. The final section summarizes and makes some concluding remarks.

2 Baby looks and paternal investment

Evolutionary theory predicts parents will provide preferential care to genetically related children to advance their genetic success (Hamilton, 1963; Trivers, 1972; Alexander, 1974; Rushton, 1989). Therefore, investments in unrelated children are deemed wasteful and costly since they reduce investment in other genetically related offspring. Indeed, stepchildren and adoptees tend to receive lower parental investment and greater mistreatment (e.g. Bertram (1975) for evidence on lions; and Daly & Wilson (1996); Gibson (2009) for evidence on humans). Similar discriminatory practices also prevail when males suspect infidelity. The evidence based on both animals and humans indicate that males invest in parental care in response to paternity certainty (birds: Moller & Birkhead (1993); fish: Neff & Gross (2001); nonhuman primates: Buchan et al (2003); Langos et al (2013); humans: Fox & Bruce (2001); Anderson et al (2007)).

To detect genetic relatedness, doubtful males can use a child's facial likeness (Daly & Wilson, 1998), in lieu of paternity testing.[3] Experiment-based evidence for humans corroborates this. For example, based on video images, Volk & Quinsey (2002, 2007) find men are more prone than women to adopt an infant looking like them. Similarly, based on pictures morphed with at least 25% of respondent's facial traits, Platek et al. (2002, 2003, 2004) find men (but not women) hypothetically spend more time, adopt, or provide financial support to toddlers similar to themselves. Moreover, according to Platek et al. (2004, 2005) neurocognitive processes drive these gender differences. Studies on actual biological parents also support that men use resemblance as a paternity cue (e.g. Apicella & Marlowe, 2004; Alvergne et al., 2009; Heijkoop et al, 2009). Even fathers convicted of family violence treat children better based on similar facial resemblance (Burch & Gallup, 2000). Finally, with respect to the reliability of reported resemblance, Alvergne et al. (2010), present three key findings. First, resemblance is consistent when reported in private by each parent. Second, both parents more likely to report father-child similarity the more the child actually resembles the father (as determined by external judges). Thirdly, actual facial resemblance is predictive of paternal investment.

To conclude, many studies support using baby looks to explain a putative father's time investment. Because women do not use facial resemblance cues to decide whether to provide care, we expect baby looks affect child developmental outcomes only through a father's investment. When data is collected in private for each parent, reported resemblance is reliable and matches actual resemblance.

[3] Even recent studies on nonhuman primates (e.g. Parr et al (2010); Kazem & Widdig (2013); Pfefferle et al (2014)) find evidence that facial similarities is used to detect kinship.

3 Data and variable descriptions

We base our empirical analysis on data from the Fragile Families and Child Wellbeing (FFCW) study. The study follows nearly 5000 children born (between 1998 and 2000) to married or unmarried parents in large U.S. cities. The data oversample births to unmarried parents to assess living conditions and wellbeing of such children.[4] Mothers and putative fathers were interviewed separately at birth (mainly within the first 3 days in the hospital) and then at four follow-ups. [5] This study uses data collected from both parents' interviews at birth (baseline) and from mother's interview at 1-year follow-up.[6] We focus on 830 unmarried families in which at 1-year follow-up the newborn resides with the mother, and the father lives elsewhere, is not in jail, and not mandated (through a child support agreement) to visit the child. Therefore, the father's visitation is not legally restricted. In such fragile families, infidelity and sexual distrust abound (Hill, 2007); and so, paternal confidence is likely to be low.

3.1 Baby looks measure

The FFCW collects baby looks data at birth from both parents' responses to the question: "Who does the baby looks like?" Omitting non-responses, the final study sample is 713 unmarried and non-cohabiting parents.[7] We classify a baby as resembling the father if both parents reported in private that the baby looks either like the father only or like both parents.[8] Otherwise, a baby is classified as having little or no resemblance to the father. Nearly two-thirds of parents (n=456) agreed on baby looks. Of these, 56% (n=255) agreed to a father-child resemblance. The remaining parents (n=257) could not agree on whether or not the baby looks like the father. To alleviate 'subjectivity' bias, these latter observations are not included in our estimation sample (n=456); but instead, they are used in section 4

[4] See Reichman et al. (2001) for more detail information about the design of the FFCW sample.

[5] Mothers were asked to identify and locate the father by providing his visiting schedules or to pass on interviewer's business card to the father (FFCW User Guide). Hereafter, these putative fathers are simply referred to as fathers.

[6] These nearly 15% of fathers who responded at baseline were lost to 1-year follow-up. We focus on the first year of birth because that time period is critical for a non-resident father to create and maintain a bond with his child; otherwise, the father tends to have a consistently low level of involvement over time (Cheadle et al., 2010). Furthermore, since paternal care is costly, the decision about whether or not to invest is better made from infancy.

[7] In two cities, this question was not asked because it was added to a later version of the survey during fielding. Missing values still exist for some variables. However, these are found to be missing completely at random (MCAR) based on Little (1988) MCAR test (χ^2=423.71, df =427, p.=0.536). As such, analyses are done without concern of bias due to listwise deletion of missing values.

[8] Men are no better than women in detecting resemblance Platek et al. (2003). So, rather than rely on one parent's perception of resemblance, we use the consistency in both parents' perception, each of which were reported in private. Moreover, both parents are likely to see father-child resemblance, if indeed the child looks like the father (Alvergne et al., 2010). The grouping of the categories 'father-only' and 'both parents' as father-child resemblance is in line with Platek et al. (2003) finding that men are more likely to invest in children with at least 25% of their facial characteristics.

as an evaluation sample to validate the out-of-sample performance of our model. Table 1 provides summary statistics on the rest of the variables.

3.2 Father visitation index

Father visitation is defined as the time a non-resident father spends participating in activities with the child. We rely on two measures of a father's time-investment: (1) the number of days in the last month the father has seen the child (father-child contact measure); and (2) the number of days in a typical week the father engages in parenting activities (father engagement measure). Both are based on the mother's response at a 1-year follow-up.[9] The latter measure of fathers' parenting activities include changing diapers, preparing/feeding a bottle, reading/telling stories, singing songs/nursery rhymes, and playing (games and toys).[10] Combining these two sets of contact and engagement measures yields a single time-investment visitation index of days per month the father engages in parenting activities.[11] The index ranges from 0 to 30 and has a mean and standard deviation of 5.8 days and 8 days, respectively (Table 1). Father visitation is on average 2.5 days longer in a given month when the child reportedly looked like him.

3.3 Child health indicators

The FFCW denotes child health subjectively with a 5-point scale ranging from 1=excellent to 5=poor, based on mothers' responses. Due to the low proportion of children in the poor, fair and good categories, we reclassify subjective child health status as 0=poor/fair/good, 1=very good and 2=excellent, with a frequency of 14.0%, 22.5% and 63.5%, respectively. Table 1 indicates an average subjective child health of 1.5, which is significantly higher for children that resemble their father.[12] Consistent with other studies, we use subjective child health as our primary outcome variable.

We also use data on four specific health outcomes. These include: whether the child experienced an asthma episode/attack since birth; number of visits to a health care professional for illness since birth; number of emergency room visits since birth; and the longest overnight stay in hospital. Asthma, in particular, is a common and growing childhood chronic health illness that is associated with high health care utilization such as emergency room visits

[9] We rely on mothers' reports because many fathers were lost at the 1-year follow-up (22% of study sample, n=713). In any case, Hernandez & Coley (2007) find that either mother or father report is a reliable measure of father involvement within low-income families.

[10] The responses are consistent based on a 0.943 Cronbach's alpha reliability coefficient. Jones & Mosher (2013) use similar activities for children under 5 years to define father engagement over the last month. They are also used as parental engagement indicators in other national surveys such as the Panel Study of Income Dynamics.

[11] We obtain this index by dividing the father's contact days/month (measure 1) by 7 days/week yielding contact weeks/month, which we then multiply by the average number of days/week engaged in each activity (measure 2) to generate the index denominated in activity days per month.

[12] As a validity check, we find that height-for-age and BMI objective measures from a 3-year FFCW follow-up agree significantly with subjective child health.

and hospitalization (Akinbami, 2006).[13] We also compute a composite score, through principal component analysis, which effectively summarizes the underlying commonalities among these four child health indicators as well as the subjective overall child health measure used above. While this score links the health measures, its unknown scale makes the economic significance of the father visitation effect hard to assess. Later in section 6, we show our results are robust with respect to all these other child health measures. Table 1 shows the incidence of asthma is 10.4%. Also, since birth, there were on average four health care visits for illness, less than two visits to emergency rooms and no more than a day stay in hospital. In addition, children that resemble their father have more favorable health conditions, based on a composite child health score.

3.4 Control variables

Table 1 also presents summary statistics for family demographic and socioeconomic control variables. Unless otherwise specified, the controls are taken from FFCW baseline survey and measured in percentages. The variables unavailable at baseline, and therefore measured at 1-year follow-up, are appended with 'year 1'.

At the first follow-up, the children are approximately one year old, of which about 55% are boys. Some children were disabled (3.8%) or of low birth weight (12.8%); and about 20% of births were covered by private health insurance, which indicates that some parents had better access to physicians and health care. The unmarried parents in this sample knew each other for typically 4.1 years before pregnancy, 58.6% of which had a visiting or friendly relationship, and majority (72.5%) had their first child together.[14] Parents were mainly Black (nearly 70%); and about 25% of mothers gave birth as a teenager, whereas only 13.5% of fathers were teenagers. Over a third of parents had kids in a previous relationship at baseline and some mothers (11%) lived with a new partner when asked at 1-year follow-up. Fathers were more likely to be in very good/excellent health (70.8%) and to have at least high school education (64%), relative to mothers (64.7% and 57.5%, respectively). Just over 75% of fathers had a job at child's birth. Mothers earned on average $242 weekly, when they last worked prior to birth; and almost one-third of them were homeowners. Nearly half of mothers had received welfare prior to birth. When measured at 1-year follow-up, mothers tend to have strong social support; and 62.4% percent of children were being cared for by a non-parent on a regular basis (for typically 40 hours per week). [15]

Importantly, aside from child gender and whether father has other children, the sample characteristics generally do not differ significantly by baby looks. Thus, these descriptive

[13] The three specific child health measures are not perfect as they may reflect both access to health care and true health quality. In our study sample though, all child health measures have statistically significant (p. < 0.001) correlations in the expected direction.

[14] Married parents knew each other for about twice as long before pregnancy.

[15] Social support is measured as the average of the mother's responses (yes = 1, no = 0) to whether the mother could count on someone to (1) loan her $200, (2) provide her with a place to live, and (3) help her with emergency child care, in the next year. The score range from 0 to 1 with higher scores indicating greater access to social support.

results give credence to using father-child resemblance as an exogenous indicator. Exogeneity is even more likely if all these characteristics are partialled out.

4 Estimation method

Our objective is to estimate a visiting father's impact on the underlying health quality of his child. Child health quality in family i (H_i^*) is reported by the mother as a categorical variable H_i, where $H_i = h$, if $\tau_h < H_i^* \leq \tau_{h+1}$, for $h = 0, 1, \ldots, J-1$. The unknown ordered cutoff values, $\tau_0 = -\infty$, $\tau_1 = 0$, $\boldsymbol{\tau} = (\tau_2, \tau_3, \ldots, \tau_{J-1})$ and $\tau_J = \infty$, define J categories of child health quality. In our application, $J = 3$ and the categories are ordered as good/fair/poor, very good, and excellent. Child health quality is modeled as

$$H_i^* = \alpha_0 + \alpha_1 f_i + \boldsymbol{X}_i'\boldsymbol{\beta} + \epsilon_i \tag{1}$$

The variable f_i is our measure of father visitation; $\boldsymbol{X}_i$ is a vector of demographic and socioeconomic controls chosen to be uncorrelated to the standard normal error term ϵ_i. The coefficient α_1 captures the impact of father visitation on latent child health. Given the concerns raised in section 1 about reported father visitation being endogenous (that is, $\mathrm{E}[\epsilon_i|f_i, \boldsymbol{X}_i] \neq 0$), estimation of (1) will result in biased and inconsistent estimates. We therefore use baby looks bl_i as an IV. This variable should be exogenously related to father visitation (and not directly to child health outcome). The relationship is given by

$$f_i = \eta_0 + \eta_1 bl_i + \boldsymbol{X}_i'\boldsymbol{\eta_2} + \upsilon_i \tag{2}$$

where we assume the error term $\upsilon_i \sim N(0, \sigma^2)$. In equation (2), η_1 is the coefficient of interest. It should be both economically and statistically significant for baby looks to be an IV. First, we examine the reduced form model defining the impact of baby looks on child health. We substitute (2) in (1) to obtain:

$$H_i^* = \mathrm{a}_0 + \mathrm{a}_1 bl_i + \boldsymbol{X}_i'\mathrm{b} + \mathrm{e}_i \tag{3}$$

where the coefficients and error terms in (3) are combinations of those in (1) and (2). As long as bl_i and $\boldsymbol{X}_i$ are uncorrelated with ϵ_i, a statistically significant estimate of a_1 yields the causal effect of baby looks. Second, we estimate the causal effect of fathers visitation. If f_i is endogenous, it must be driven by a correlation of unobservables ϵ_i and υ_i, which we assume to be bivariate normally distributed with correlation coefficient ρ. As such, we partial out the effect of υ_i on ϵ_i, by using υ_i as a control variable in equation (1). This yields

$$H_i^* = \alpha_0 + \alpha_1 f_i + \boldsymbol{X}_i'\boldsymbol{\beta} + (\rho \backslash \sigma)\upsilon_i + \kappa_i \tag{4}$$

Here $\kappa_i \sim N(0, 1-\rho^2)$ and $\rho\backslash\sigma$ is the coefficient on υ_i. A significant ρ implies endogeneity; and its sign indicates the omitted variables at play.[16] If f_i is uncorrelated with the error term κ_i, estimation of (1) produces unbiased and consistent estimates. See Technical appendix A.1 for the likelihood function that is maximized to produce these estimates.

A large literature documents how family income improves child health (see Currie (2009) for a review). For a single mother especially, greater earnings enables the purchase of better quality child health inputs such as high quality food, better medical care, child care, safer toys and quality housing. Maternal income also reduces stress and depression and increases general psychological well-being, all of which affect the quality of parental care (Gregg et al., 2007). Some studies (e.g. Alaimo et al., 2001; Hofferth & Curtin, 2005; Doyle et al., 2007; Gregg et al., 2007) link income and child health non-linearly, so that the effect of income on child health gets smaller as income rises. In this vane, a visiting father's role in a child's health may be less important for relatively high-income mothers. Ignoring such nonlinearity, based on a mother's ability to independently provide for her child, could then overestimate the visiting father's contribution. To capture this non-linearity, we generalize (4) to the following semivarying-coefficient specification:

$$H_i^* = \alpha_0(U_i) + \alpha_1(U_i) f_i + \tilde{\boldsymbol{X}}_i' \boldsymbol{\delta} + \kappa_i \tag{5}$$

where $\tilde{\boldsymbol{X}}_i = (\boldsymbol{X}_i, \upsilon_i)$, $\boldsymbol{\delta} = (\boldsymbol{\beta}', \rho\backslash\sigma)$, and the coefficients $\alpha_l(U_i)$, for $l = 0, 1$, are unknown functions of maternal weekly earnings (U_i).[17] The coefficient function $\alpha_1(U_i)$ is of direct interest because it measures how father's visitation effects relate to maternal earnings. Since poor child health may lower maternal earnings, we use mother's weekly earnings prior to birth as a proxy for her weekly earnings at 1-year follow-up.[18] Lin et al (2015) outline an approach for estimating $\alpha_1(U_i)$ in equation (5), without committing to a specific functional form. The general idea is: (i) find consistent estimates for all fixed parameters, and (ii) given these estimates, for a specific value of U_i (say, u), the coefficient $\alpha_1(u)$ is estimated by observations of U_i that are close to u, such that $|U_i - u| \leq \lambda$, where λ is a bandwidth parameter. We show the steps of this estimation approach (including the method for estimating λ) in Technical appendix A.2.

As a goodness-of-fit check, we examine the out-of-sample performance of equation (5) relative to equation (4). We use an evaluation sample, comprised of those parents who could not agree on whether or not the child looks like the father (n=257), to compute the mean absolute deviation (MAD) of H_i from its predicted frequency distribution. Importantly, the evaluation sample is completely independent of our estimation sample.

[16] For instance, $\rho < 0$ may imply an unobservable in ϵ_i, such as the degree of mother's protective gatekeeping behavior, is negatively related to the unexplained amount of father visitation in υ_i.

[17] Like Fan & Huang (2005), we take the square-root of U_i so that it has a near symmetric distribution.

[18] There is a highly significant positive association (partial correlation of 0.21) between pre-birth weekly earnings and weekly earnings at 1-year follow-up, controlling for several maternal characteristics including age and education.

5 Estimation results

Table 2 reports the estimation results corresponding to equations (1) to (4). Column (1) shows that when we estimate equation (1), father visitation has a positive but insignificant effect on latent child health.[19] This insignificance indicates a visiting father provides little or no child health benefits, and is consistent with the general findings of Amato & Gilbreth (1999). However, as discussed above, this estimate is likely to be biased and inconsistent, most likely due to the influence of unobservables.

5.1 Do baby looks matter?

To derive causal effects, we rely on baby looks as an IV. Our discussion in section 2 provides theoretical arguments and empirical support from other disciplines for the use of baby looks as an IV. The descriptive statistics presented in section 3.4 provide encouraging results regarding both the relevance and exogeneity of the baby looks measure. Column (2) of table 2 shows OLS estimation of equation (2). We control for a number of parental and child characteristics, of which child gender, access to private insurance, relationship/cohabitation status, father employment and additional children with another mother are statistically important determinants of father visitation. Consistent with the literature, we still find compelling evidence that visiting fathers react to father-child resemblance. In particular, the 2.3 difference in days/month of father visitation between those children resembling their father and those that do not is statistically significant. Therefore, the non-resident fathers in our sample use resemblance to assess paternity, and hence to decide whether to devote investment resources, particularly time, to their reputed child.[20]

Before presenting the IV estimate, we provide the effect of father-child resemblance on child health, from which we may infer whether there is a causal impact of a visiting father. There is no reason why child resemblance should matter for child health, other than encouraging paternal time-investment. We expect that looking like the father raises child fitness because a visiting father's time-investment is higher for a child that resembles him. Our 0.28 statistically significant estimate (column (3) of table 2) corroborates this conjecture. In terms of the marginal probability effect (not shown in the table), we find that on average a child that resembles the father is 5.3 percentage points (pp) less likely to be reported to have low health; 4.1 pp less likely to be in reportedly very good health; and 9.3 pp more likely to be in reportedly excellent health. These results implicitly support a causal effect of father visitation in response to father-child resemblance.

[19] Alternatively, the marginal effect (not shown in the table) on the probability of a child reportedly having low (good/fair/poor) health is -0.17 percentage points (pp) and 0.3 pp for reportedly excellent health.

[20] As a placebo test, we also regress mother's engagement in parental activities on father-child resemblance, using all the controls in Table 1. If such resemblance is used as a paternal cue, it should have no effect on mother's involvement. We find that a child's resemblance to the father has a very insignificant effect on maternal care (0.302, s.e. = 0.585).

5.2 Are there health benefits from father visitation?

We now present the estimated effect of father visitation on child health quality in column (4) of Table 2 and the corresponding marginal probability effects in Table 3. The estimate is based on equation (4). We find that an additional day (per month) of father visitation (in response to the child's resemblance to him) increases latent health by 0.093 units. Based on Table 3, this one extra day decreases the probability of low and very good child health by 2.3 pp and 0.6 pp (columns (1) and (2)) respectively, and increases the probability of excellent health by 2.9 pp (column (3)). Recall that the incidence of children in the low health category of our sample is 14%. For such children, there appears to be substantial benefits from a typical visiting father (one who engages with child for about 6 days per month) providing an additional day of involvement (per month).[21] Thus, a typical nonresident father can be encouraged to visit about twice weekly to reduce the risk of low health by 2.3 pp.

5.3 What is the role of maternal earnings?

To predict the effect of encouraging a day of involvement by visiting fathers, it is important for policy purposes to consider the mother's own ability to provide for the child and compensate for any loss in paternal time-investment. In doing so, we estimate how father's visitation effects relate to maternal earnings by estimating equation (5) without presuming a particular functional form. The coefficient function $\alpha_1(U_i)$ measures the interaction between father's visitation and mother's pre-birth weekly earnings using independent data on parents who differ in their assessment of their baby's resemblance (evaluation data).[22] Equation (5) has better predictive performance than (4), which assumes constant effects, since the former's MAD statistic is lower (12.42 versus 19.10).

Figure 1 presents the estimated interaction effect (solid black line) and its 95% confidence interval (dotted grey lines).[23] It shows that, while the marginal health effect of father visitation is always non-negative, the effect is a nonlinear function of maternal earnings. As the level of earnings increase, the marginal health benefits decline steadily from 0.10, levels off in the neighborhood of $U_i = 2$ (about $400 weekly earnings), and decline sharply to 0.04 (insignificant) thereafter. Therefore, the contribution of a visiting father to child health is more important for single mothers with relatively low earnings.

[21] It should be noted that, due to the nonlinear nature of our ordered probit framework, the marginal probability effect gets lower (in absolute value) for fathers with more-than-typical visitation days.

[22] The coefficient function $\alpha_0(U_i)$, which measures the relationship between latent child health and maternal pre-birth weekly earnings, is not of direct interest. We do find a non-linear effect of maternal earnings on child health quality. The income-health gradient is generally positive in the area of statistical significance and largest at lower levels of earnings. The findings support the positive income-health gradient reported by other studies (e.g. Case et al., 2002; Currie & Stabile, 2003); and also support the nonlinearity thereof reported by other studies (e.g. Doyle et al., 2007; Gregg et al., 2007).

[23] Since semiparametric estimates at the tails of the distribution tend to be the least accurate, we exclude from the discussion and Figures 1 and 2 mothers with no pre-birth earnings and also those with the highest pre-birth earnings.

6 Robustness checks

To test the robustness of these estimates, we repeat the analysis (Table 4) using the specific and the composite health measures discussed in Section 3.3. As before, we find no significant effect of father's visitation when using equation (1) which does not account for endogeneity (Panel 1). Column (1) contains the previous insignificant 0.009 coefficient. Columns (2) through (5) contain the impact on specific health measures, again indicating no statistical significance. Column (6) uses the composite health index (where higher values indicate better health) derived by extracting the principal component from all five previously used health indicators of child health. Here too father's visitation has no statistically significant impact.

On the other hand, accounting for endogeneity using baby looks as an instrument (Panel 2) yields significant father effects. One extra day of father's time greatly reduces the incidence of asthma (from -.023 to -.115 or about 2.7 percentage points), decreases visits to health practitioners for illness (from -.006 to -.057), cuts visits to the emergency room (from -.003 to -.093), and decreases hospital stays (from 0.008 to -.169). The composite child health score, created from all five child health indicators, confirms an overall positive and statistically significant effect of father visitation (column (6)). Thus, we find consistent effects regardless of how child health is measured.

In section 5.3, we found children benefit less (when using the subjective health measure) from father visitation as maternal income rises. We check this result by re-estimating equation (5) using the composite child health score.[24] Figure 2 plots the father's effect as a function of mother's pre-birth weekly earnings. Our previous findings are upheld: the visiting father's contribution is higher for relatively low-income mothers.

7 Validity checks

To be a valid IV, our reported baby looks must be exogenous and must affect health only through father visitation, and not via any other path (the exclusion restriction). We now provide evidence that our reported baby looks meets these two criteria.[25]

First, we investigate whether there is a violation of the IV's exclusion restriction. Suppose a father provides more financial support (in addition to his time) when a baby looks like him. This is done by establishing paternity, which increases child support payments (Mincy et al.,

[24] We use the estimation procedure outlined in Technical Appendix A.2, with two changes: (1) $m(h)$ becomes a normal density function, and (2) the cross-validated bandwidth is chosen by minimizing the out-of-sample mean squared error. The bandwidth is 0.41. Equation (5) still has better predictive performance than equation (4); the formers root mean square error computed using the evaluation data is lower (1.997 versus 2.057).

[25] Another possible validity check is to rerun the analysis using the two components of baby looks (whether mother believes the father looks like the child and whether the father thinks he looks like the child) as separate IVs. Doing so yields comparable results (available upon request). Further, we obtain a statistically insignificant Hansen J statistic (0.383) indicating valid instruments.

2005).[26] We find that a father is 8.1 pp more likely to establish paternity, although only at the 10% level, when the baby looks like him. Because child support increases child health (Amato & Gilbreth, 1999; Garasky & Stewart, 2007) in addition to father visitation, the exclusion restriction would be violated. The problem is alleviated if we can show father-child resemblance increases father visitation independent of establishing paternity. Figure 3 establishes this. It shows that looking like the baby increases father's time with the baby whether or not paternity is established

Second, we check whether the reported baby looks IV is exogenous. To be exogenous, reported baby looks must not be determined by other variables. In other words, there should be no ulterior motive for either the mother or putative father to indicate the baby looks like the father, when it fact it does not. Though lying is unlikely because responses are for an academic survey not viewed by either parent, we still can test for this possibility, especially a subconscious lie. Suppose the mother perceives the purported father to look like the baby because she instinctively seeks to secure the maximum amount of paternal care for her child, independent of the purported father's biological status. Similarly, suppose the purported father's perception is influenced by his desire to spend time with the mother and her child. Such false perceptions, if they exist in our original sample, should lead to an overestimate of the proportion of children that look like the father. But suppose we find another group of parents whose true perceptions are not similarly inhibited. Respondents with the purported father in prison would be an example. Incarcerated fathers cannot spend time with the mother or child and the mothers know this. In this case, parents' responses should be accurate. Thus a simple test to determine exogeneity of baby looks is to test whether the responses from the two groups, namely our original sample and the sample of prisoners, are equal. That is, do both groups have the same proportion of parents reporting father-child resemblance?

The FFCW study contains 97 unmarried purported fathers in prison, who were asked about the baby's looks. Of these, 62 sets of purported parents agree on baby looks; and of these, 35 (56.5%) parents agree that the baby looks like the father. Our original sample contains 456 mothers and purported fathers that agree on baby looks. Of these, 255 (55.9%) indicate the baby looks like the father. Based on a 0.937 p-value, the odds of reporting father-child resemblance is not significantly different between the two groups. This finding supports the exogeneity of our baby looks measure.

8 Conclusion

Early childhood development is an important precursor for future success. Parental inputs are potentially crucial. In this regard, children in single parent households are most vulnerable. This study estimates the causal effect of nonresident father visitation on child health

[26]For unmarried families, establishing paternity means voluntarily signing a paper legally acknowledging the father for the child. In most cases, this is done in-hospital. This then guarantees the father's parental rights and allows his name to be on the birth certificate. The child in return is eligible for child support and any other derived benefits (e.g. veteran survivor benefits).

using baby looks as an IV. The idea is that, due to paternal uncertainty, men assess genetic relatedness based on whether the child resembles him. This prediction is supported in experimental studies based on animal/human families in evolution-related disciplines, but to date has not been used in economics research. In our study, we employ rich US national data on 713 unmarried parents with nonresident fathers taken from the first two waves of the Fragile Family and Child Wellbeing study. This data is appropriate since paternal uncertainty is more likely to prevail among fragile families.

Our findings support father-child resemblance as a paternity cue used by men to determine whether to make time-investments. Moreover, we find, even after several robustness and validity checks, that father visitation, in response to baby looks, causes an increase in child health. Specifically, our basic result is that an extra visitation day reduces the incidence of reportedly low health by about 2 percentage points and increase the chance of reportedly excellent health by about 3 percentage points. Furthermore, the health contribution of the visiting father increases with lower maternal earnings. This study therefore supports father visitation policies for delivering improvements in early childhood health, especially for single working mothers with relatively low earnings. Indeed, the National Quality Improvement Center for Non-Resident Fathers was created in 2006 by the US child welfare system to promote non-resident father engagement with children in the welfare system.

A Technical Appendix

A.1 IV-ordered probit (IV-OP) estimation

To perform maximum likelihood estimation of (4), we require the conditional joint density $g(\cdot)$ for H_i and f_i (since they are jointly related). For $h = 0, 1, \ldots, J-1$, $g(\cdot)$ is expressed as

$$g(f_i, H_i = h|\ \boldsymbol{X}_i, \boldsymbol{Z}_i;\ \boldsymbol{\theta}, \boldsymbol{\alpha}) = \frac{1}{\sigma}\phi\left(\frac{f_i - \boldsymbol{Z}_i'\boldsymbol{\eta}}{\sigma}\right) \cdot m(h)$$

where $\boldsymbol{Z}_i = (1, bl_i, \boldsymbol{X}_i)$; $m(h) = (\Phi_{i,h+1} - \Phi_{i,h})$; $\Phi_{i,h} = \left(\frac{\tau_h - (\alpha_0 + \alpha_1 f_i + \boldsymbol{X}_i'\boldsymbol{\beta} + (\rho\backslash\sigma)(f_i - \boldsymbol{Z}_i'\boldsymbol{\eta}))}{\sqrt{1-p^2}}\right)$ is the standard normal cumulative distribution function for $H_i = h$; $\Phi_{i,h+1} - \Phi_{i,h}$ is the probability of child i being in health category h; and $\phi(\cdot)$ is the standard normal density for f_i; $\boldsymbol{\theta} = (\boldsymbol{\eta}', \sigma, \tau, \beta')$; $\boldsymbol{\eta}' = (\eta_0, \eta_1, \eta_2')$; and $\boldsymbol{\alpha} = (\alpha_0, \alpha_1)$. The estimates of $(\boldsymbol{\theta}, \boldsymbol{\alpha})$ are maximizers of the log-likelihood function $\ell(\boldsymbol{\theta}, \boldsymbol{\alpha}) = \sum_{i=1}^{n}\sum_{h=0}^{J-1} 1(H_i = h) \cdot \ln\ g_i(\cdot)$, where $1(\cdot)$ is an indicator function.

A.2 Semivarying-coefficient IV-OP estimation

We derive estimators for the functional coefficients $\boldsymbol{\alpha}(U_i)$ based on the local constant approximation method. Similar to Lin et al (2015), we observe the following profile maximum likelihood procedure. First, define:

$$\ell_i(\boldsymbol{\theta}, \boldsymbol{\alpha}) = \sum_{h=0}^{J-1} 1(H_i = h) \cdot \ln\ g\Big(f_i, H_i = h|\ \boldsymbol{X}_i, \boldsymbol{Z}_i;\ \boldsymbol{\theta}, \boldsymbol{\alpha}\Big)$$

Step 1: Maximize a local (weighted) log-likelihood function, with respect to $(\boldsymbol{\theta}, \boldsymbol{\alpha})$, to derive initial estimates for any given u:

$$\Big(\bar{\boldsymbol{\theta}}(u), \bar{\boldsymbol{\alpha}}(u)\Big) = \arg\max_{(\boldsymbol{\theta}, \boldsymbol{\alpha})} \sum_{i=1}^{n} \ell_i(\boldsymbol{\theta}, \boldsymbol{\alpha}) K_\lambda(U_i - u)$$

where K is the univariate gaussian kernel, which depends on the bandwidth parameter λ and assigns greater weights to observations of U_i in the neighborhood of u.

Step 2: Derive a pseudo log-likelihood function for observation i by replacing $\boldsymbol{\alpha}$ in $\ell_i(\boldsymbol{\theta}, \boldsymbol{\alpha})$ with $\bar{\boldsymbol{\alpha}}(U_i)$ from step 1. Maximizing the sum of these pseudo log-likelihood contributions, with respect to $\boldsymbol{\theta}$, provides the following consistent estimates:

$$\hat{\boldsymbol{\theta}} = \arg\max_{\boldsymbol{\theta}} \sum_{i=1}^{n} \ell_i(\boldsymbol{\theta}, \bar{\boldsymbol{\alpha}}(U_i))$$

Step 3: Derive a pseudo log-likelihood function for observation i by replacing $\boldsymbol{\theta}$ in $\ell_i(\boldsymbol{\theta}, \boldsymbol{\alpha})$ with $\hat{\boldsymbol{\theta}}$ from step 2. We maximize a weighted sum of these pseudo log-likelihood contributions, with respect to α, to derive a final estimate of $\boldsymbol{\alpha}(u)$, for any given u:

$$\hat{\boldsymbol{\alpha}}(u) = \arg\max_{\boldsymbol{\alpha}} \sum_{i=1}^{n} \ell_i(\hat{\boldsymbol{\theta}}, \boldsymbol{\alpha}) K_\lambda(U_i - u)$$

We take $(\hat{\boldsymbol{\theta}}, \hat{\boldsymbol{\alpha}}(u))$ as the estimates for $(\boldsymbol{\theta}, \boldsymbol{\alpha}(u))$. To make inferences about the functional coefficient of interest $\boldsymbol{\alpha}(u)$, like Hwang (2013) and Lin et al (2015), we compute robust sandwich-type standard errors, which are then used to construct point-wise 95% confidence intervals.

Bandwidth Selection. Note that we constrain the bandwidths in steps 1 and 3 to be equal to λ, so as to reduce the high computational costs associated with selecting possibly two different optimal bandwidths. Furthermore, we use a low-intensive, stratified 10-fold cross-validation (CV) procedure for estimating λ, as recommended by Kohavi (1995). This procedure involves: (i) taking each child health status H_i as a stratum of the estimation sample and (ii) within each stratum, randomly assigning observations to 10 groups (or folds) of data, labeling them as d_k, for $k = 1, \ldots, 10$.[27] In each of the 10 data groups, there is about the same number of observations and the categories of H_i are in roughly the same proportion as in the full estimation sample. For each k, the data group d_k is excluded from the estimation sample and steps 1-3 are executed to get estimates $(\hat{\boldsymbol{\theta}}_{-k}, \hat{\boldsymbol{\alpha}}_{-k}(U_i))$ of $(\boldsymbol{\theta}, \boldsymbol{\alpha}(U_i))$. This procedure allows us to compute the following statistic:

$$CV(\lambda) = 10^{-1} \sum_{k=1}^{10} |d_k|^{-1} \sum_{i \in d_k} 1\left(\arg\max_{0 \le h \le J-1} p_{ih}(\hat{\boldsymbol{\theta}}_{-k}, \hat{\boldsymbol{\alpha}}_{-k}(U_i)) \neq H_i \right)$$

where $|d_k|$ is the number of observations in the k^{th} data group; $p_{ih}(\hat{\boldsymbol{\theta}}_{-k}, \hat{\boldsymbol{\alpha}}_{-k}(U_i))$ is the probability that child i in d_k is in health class h, based on estimates computed without using data d_k. The function $CV(\lambda)$ is a measure of the average out-of-sample forecast classification error. As in Hwang (2013), the value of λ in the interval $[0.5\hat{\sigma}_{U_i}, 3\hat{\sigma}_{U_i}]$ that minimizes $CV(\lambda)$ is taken as its estimate, where $\hat{\sigma}_{U_i}$ is the standard deviation of U_i. Our estimated value of λ is approximately 0.30.

[27] The random seed for reproducing our random assignments is set at 123456789 in R.

B Appendix: Figures and Tables

Figure 1: Plot of the estimated functional coefficient (solid lines) and its 95% confidence intervals (dotted grey lines). Outcome variable is the subjective child health measure.

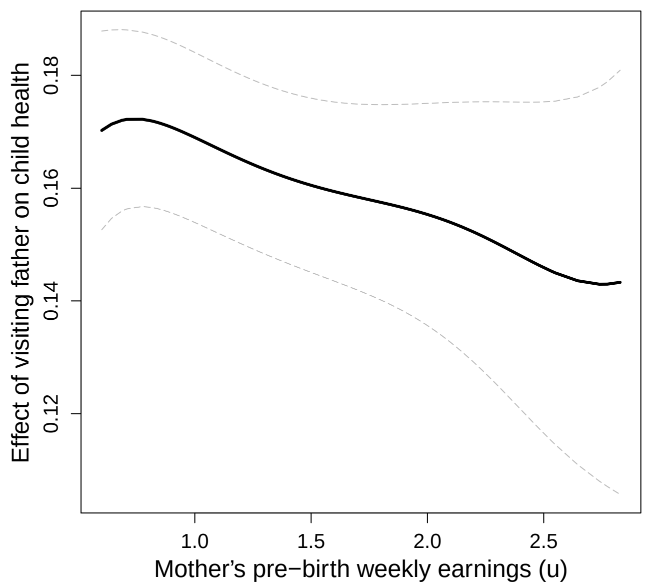

Figure 2: Plot of the estimated functional coefficient $\hat{\alpha}_1(u)$ (solid lines) and its 95% confidence intervals (dotted grey lines). Outcome variable is the composite child health score.

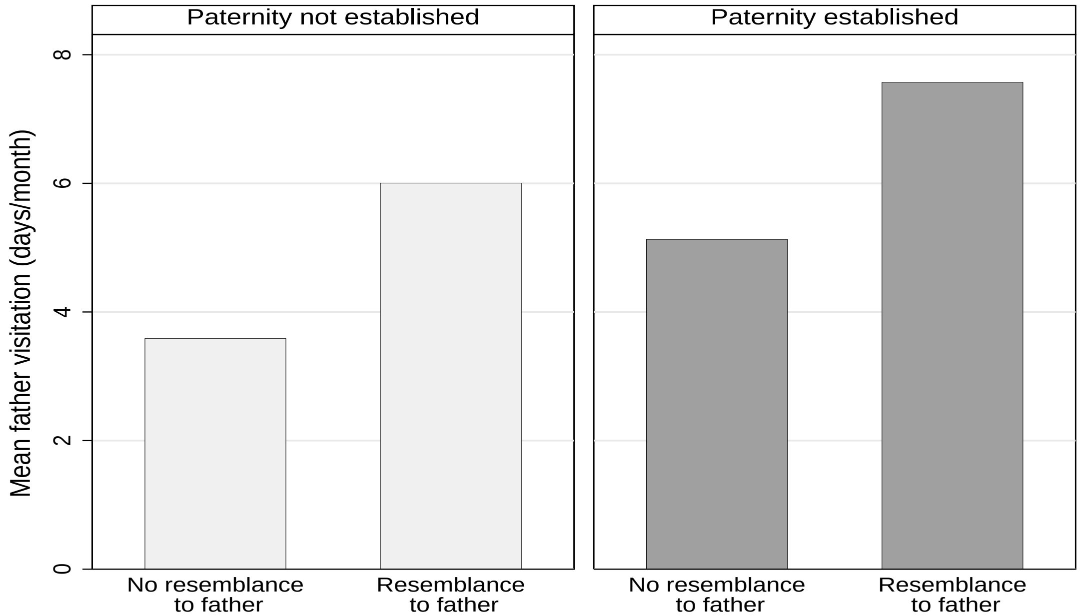

Figure 3: Father time-investment by baby looks and legal paternity status

Table 1: Descriptive statistics by reported father-child resemblance (FCR)†

		Estimation sample		
	Overall (n=713)	**Agreed FCR (n=255)**	**Agreed No FCR (n=201)**	**Disagreed on FCR (n=257)**
Father visitation index (days/month)	5.8 (8.0)	7.1 (8.7)	4.6 (6.7) *	4.7 (7.4)
Subjective child health (0-2)	1.50 (0.73)	1.55 (0.68)	1.40 (0.79)*	1.51 (0.72)
Asthma episodes/attacks	10.4	8.3	12.4	10.9
No. of health care visits for illness	3.8 (5.1)	3.5 (3.9)	4.3 (6.2)	3.6 (5.1)
No. of emergency room visits	1.6 (2.2)	1.6 (2.4)	1.8 (2.4)	1.4 (2.0)
Longest stay in hospital (days)	0.9 (4.6)	0.7 (1.8)	1.4 (7.5)	0.7 (3.4)
Composite child health score	0.00 (1.40)	0.10 (1.32)	-0.21 (1.53)*	0.06 (1.36)
Control variables				
Child is boy	54.6	60.0	49.3 *	53.3
Child age in months (year 1)	15.3 (3.5)	15.7 (3.5)	15.2 (3.4)	15.0 (3.5)
Child disabled (year 1)	3.8	4.3	4.0	3.1
Child Low birth weight	12.8	12.9	9.5	15.2
Child birth covered by private insurance	20.2	20.0	22.9	18.3
Years mother knew father pre-pregnancy	4.1 (4.4)	3.7 (4.0)	4.1 (4.7)	4.4 (4.6)
Parent relations: cohabited [1]	34.2	38.4	31.3	32.3
Parent relations: visiting/friends [1]	58.6	55.7	61.2	59.5
Parents first kid together	72.5	77.3	73.1	70.0
Father race: Black [2]	68.2	67.1	72.6	65.8
Father race: Hispanic [2]	19.2	18.4	17.9	21.0
Father teenager	13.5	14.5	15.4	10.9
Father has kids with other mother	37.9	33.1	42.8 *	38.9
Father's subj. health: excellent/very good	70.8	73.2	65.7	72.4
Father's education: high school [3]	42.9	44.3	39.3	44.4
Father's education: some college or more [3]	21.2	21.2	19.9	22.2
Father employed	76.1	78.7	74.1	75.1
Mother's race: Black [2]	67.0	65.5	69.7	66.5
Mother's race: Hispanic [2]	18.5	19.6	17.9	17.9
Mother teenager	25.7	25.1	27.4	24.9
Mother has kids with other father	39.7	42.7	37.3	38.5
Mother lives with new partner (year 1)	11.0	12.2	11.5	9.4
Mother's subj. health: excellent/very good	64.7	65.9	60.7	66.5
Mother's education: high school [3]	31.8	32.9	30.8	31.5
Mother's education: some college or more [3]	25.7	26.3	23.4	26.8
Mother's weekly earnings ('00)	2.42 (1.42)	2.46 (1.43)	2.37 (1.40)	2.40 (1.44)
Mother owns home	32.2	30.7	30.5	34.9
Mother on welfare/TANF	47.4	46.7	44.8	50.2
Mother has social support access (year 1)	0.84 (0.30)	0.83 (0.31)	0.84 (0.28)	0.85 (0.30)
Non-parental caregiver (year 1)	62.4	62.0	62.2	63.0

Source: Fragile Families and Child Wellbeing Study
† Variables measured at baseline unless otherwise indicated. Figures are sample means. Standard deviation is reported in parenthesis. Unit is percentage unless otherwise indicated.
* denotes a significant difference (p.<0.05) in the sample means between those kids agreed to look like father and those kids agreed to not look like father.
Reference groups: [1] Little or no relations (hardly or never talk). [2] White. [3] Some high school or less.

Table 2: Subjective child health (H) and visiting father time-investment (f)

Outcome variable	**(1)** **H**	**(2)** $\boldsymbol{f}$	**(3)** **H**	**(4)** **H**
Father visitation	0.009 *(0.008)*			0.093 *(0.029)*
Father-child resemblance		2.327 *(0.727)*	0.280 *(0.130)*	
Boy	-0.196 *(0.129)*	1.546 *(0.783)*	-0.204 *(0.131)*	-0.301 *(0.120)*
Child age	-0.008 *(0.020)*	-0.021 *(0.109)*	-0.011 *(0.020)*	-0.006 *(0.017)*
Child disabled	-0.867 *(0.282)*	-0.116 *(2.332)*	-0.878 *(0.280)*	-0.664 *(0.357)*
Low birth weight	-0.351 *(0.207)*	-0.777 *(1.116)*	-0.381 *(0.206)*	-0.221 *(0.197)*
Private insurance	0.423 *(0.179)*	-2.214 *(0.897)*	0.438 *(0.177)*	0.542 *(0.159)*
Years mother knew father	0.040 *(0.016)*	0.011 *(0.080)*	0.042 *(0.016)*	0.031 *(0.017)*
Cohabited [1]	0.297 *(0.259)*	4.315 *(1.391)*	0.316 *(0.254)*	-0.157 *(0.291)*
Visiting/friends [1]	0.060 *(0.245)*	3.025 *(1.227)*	0.078 *(0.243)*	-0.221 *(0.237)*
First kid together	0.211 *(0.172)*	0.340 *(1.017)*	0.206 *(0.172)*	0.123 *(0.164)*
Father Black [2]	0.281 *(0.258)*	-1.231 *(1.608)*	0.297 *(0.262)*	0.342 *(0.240)*
Father Hispanic [2]	0.115 *(0.268)*	-1.556 *(1.796)*	0.134 *(0.269)*	0.246 *(0.260)*
Father teenager	0.236 *(0.206)*	0.465 *(1.370)*	0.254 *(0.210)*	0.149 *(0.201)*
Father: kids by other mom	0.064 *(0.137)*	-2.219 *(0.809)*	0.087 *(0.136)*	0.272 *(0.140)*
Father's health	0.078 *(0.141)*	0.026 *(0.836)*	0.060 *(0.144)*	0.045 *(0.133)*
Father: high school [3]	-0.208 *(0.164)*	0.397 *(1.026)*	-0.230 *(0.165)*	-0.213 *(0.161)*
Father: >high school[3]	-0.374 *(0.211)*	1.161 *(1.205)*	-0.391 *(0.210)*	-0.408 *(0.205)*
Father employed	0.275 *(0.158)*	-2.163 *(0.979)*	0.247 *(0.156)*	0.390 *(0.152)*
Mother Black [2]	-0.037 *(0.232)*	2.337 *(1.491)*	-0.019 *(0.237)*	-0.232 *(0.226)*
Mother Hispanic [2]	-0.288 *(0.253)*	1.929 *(1.598)*	-0.290 *(0.253)*	-0.402 *(0.235)*
Mother teenager	-0.145 *(0.175)*	-0.338 *(1.060)*	-0.160 *(0.176)*	-0.091 *(0.160)*
Mother: kids by other dad	0.070 *(0.149)*	1.647 *(0.908)*	0.070 *(0.147)*	-0.098 *(0.156)*
Mother with new partner	0.233 *(0.206)*	-3.469 *(0.974)*	0.203 *(0.204)*	0.479 *(0.182)*
Mother health	0.529 *(0.133)*	0.481 *(0.805)*	0.524 *(0.133)*	0.358 *(0.168)*
Mother: high school [3]	-0.184 *(0.158)*	-1.684 *(0.973)*	-0.200 *(0.158)*	0.005 *(0.166)*
Mother: >high school [3]	-0.192 *(0.195)*	-0.753 *(1.194)*	-0.215 *(0.193)*	-0.093 *(0.191)*
Mother's weekly earnings	-0.026 *(0.053)*	0.332 *(0.313)*	-0.028 *(0.052)*	-0.053 *(0.047)*
Mother owns home	0.003 *(0.138)*	1.185 *(0.929)*	0.023 *(0.138)*	-0.091 *(0.134)*
Mother on welfare	-0.079 *(0.142)*	1.016 *(0.859)*	-0.084 *(0.143)*	-0.160 *(0.130)*
Social support access	0.154 *(0.234)*	1.648 *(1.628)*	0.193 *(0.233)*	-0.005 *(0.238)*
Non-parental caregiver	-0.188 *(0.139)*	1.137 *(0.860)*	-0.180 *(0.138)*	-0.244 *(0.129)*
Estimation Method	OP	OLS	OP	IV-OP

Notes: Sample size n= 456. Specifications (1)–(4) correspond to those described in section 4: estimation method. Robust standard errors are in parentheses. Correlation coefficient ρ for equation (4) is -0.641 (s.e.= 0.216). IV-OP is the instrumental variable ordered probit model.
Reference groups: [1] Little or no relations (hardly or never talk). [2] White. [3] Some high school or less.

Table 3: Marginal probability effects on subjective child health (H)

	(1) $Pr(H_i = 0)$	**(2)** $Pr(H_i = 1)$	**(3)** $Pr(H_i = 2)$
Father visitation	-0.023 *(0.011)*	-0.006 *(0.003)*	0.029 *(0.008)*
Boy	0.109 *(0.034)*	0.007 *(0.015)*	-0.116 *(0.037)*
Child age	0.002 *(0.004)*	0.000 *(0.001)*	-0.002 *(0.006)*
Child disabled	0.260 *(0.106)*	-0.060 *(0.026)*	-0.200 *(0.114)*
Low birth weight	0.087 *(0.052)*	-0.010 *(0.016)*	-0.077 *(0.064)*
Private insurance	-0.214 *(0.041)*	0.043 *(0.029)*	0.171 *(0.046)*
Years mother knew father	-0.008 *(0.004)*	-0.002 *(0.002)*	0.010 *(0.005)*
Cohabited [1]	0.062 *(0.077)*	0.006 *(0.014)*	-0.056 *(0.090)*
Visiting/friends [1]	0.081 *(0.061)*	0.003 *(0.015)*	-0.084 *(0.072)*
First kid together	-0.048 *(0.041)*	0.004 *(0.013)*	0.044 *(0.052)*
Father Black [2]	-0.135 *(0.065)*	0.020 *(0.020)*	0.115 *(0.075)*
Father Hispanic [2]	-0.090 *(0.060)*	-0.004 *(0.021)*	0.094 *(0.078)*
Father teenager	-0.056 *(0.046)*	-0.001 *(0.019)*	0.056 *(0.063)*
Father: kids by other mom	-0.099 *(0.039)*	-0.005 *(0.012)*	0.105 *(0.042)*
Father's health	-0.017 *(0.033)*	-0.003 *(0.009)*	0.016 *(0.042)*
Father: high school [3]	0.079 *(0.040)*	0.003 *(0.015)*	-0.081 *(0.050)*
Father: >high school[3]	0.161 *(0.056)*	-0.027 *(0.021)*	-0.135 *(0.063)*
Father employed	-0.138 *(0.044)*	-0.013 *(0.019)*	0.152 *(0.048)*
Mother Black [2]	0.085 *(0.056)*	0.003 *(0.016)*	-0.089 *(0.068)*
Mother Hispanic [2]	0.159 *(0.068)*	-0.026 *(0.020)*	-0.133 *(0.073)*
Mother teenager	0.034 *(0.040)*	-0.000 *(0.012)*	-0.034 *(0.050)*
Mother: kids by other dad	0.038 *(0.041)*	-0.003 *(0.009)*	-0.036 *(0.048)*
Mother with new partner	-0.166 *(0.046)*	-0.021 *(0.024)*	0.187 *(0.052)*
Mother health	-0.128 *(0.036)*	-0.011 *(0.026)*	0.138 *(0.057)*
Mother: high school [3]	-0.002 *(0.041)*	0.000 *(0.011)*	0.002 *(0.052)*
Mother: >high school [3]	0.036 *(0.048)*	-0.003 *(0.014)*	-0.033 *(0.060)*
Mother's weekly earnings	0.013 *(0.012)*	0.004 *(0.004)*	-0.017 *(0.015)*
Mother owns home	0.035 *(0.035)*	-0.000 *(0.009)*	-0.034 *(0.042)*
Mother on welfare	0.062 *(0.034)*	-0.006 *(0.011)*	-0.057 *(0.041)*
Social support access	0.001 *(0.059)*	0.000 *(0.016)*	-0.002 *(0.074)*
Non-parental caregiver	0.096 *(0.031)*	-0.011 *(0.015)*	-0.085 *(0.040)*

Notes: Sample size n= 456. Marginal probability effects are based on column (4) of Table 2. Robust standard errors are in parentheses.
Reference groups: [1] Little or no relations (hardly or never talk). [2] White. [3] Some high school or less.

Table 4: Robustness checks

	(1) Subjective health	(2) Has asthma	(3) Visits for Illness	(4) Emergency visits	(5) Hospital stay	(6) Composite health score
	Panel 1: Using no IV					
Father visitation index	0.009	-0.023	-0.006	-0.003	0.008	0.013
	(0.008)	(0.013)	(0.008)	(0.008)	(0.016)	(0.008)
Estimation method	OP	Probit	Poisson	Poisson	OLS	OLS
	Panel 2: Using father-child resemblance as IV					
Father visitation index	0.093	-0.115	-0.057	-0.093	-0.169	0.172
	(0.029)	(0.024)	(0.029)	(0.044)	(0.114)	(0.080)
Estimation method	IV-OP	IV-Probit	IV-Poisson	IV-Poisson	2SLS	2SLS

Notes: Sample size n=456. Robust standard errors are in parentheses below coefficients. The coefficients indicate the effect of one-extra day of father visitation on the indicated child health outcome. Controls are the same as those used in Tables 2 and 3. IV-OP is the instrumental variable ordered probit model.

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