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THE HEALTH EFFECTS OF CESAREAN DELIVERY FOR LOW-RISK FIRST BIRTHS

David Card
Alessandra Fenizia
David Silver

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ABSTRACT

Cesarean delivery for low-risk pregnancies is generally associated with worse health outcomes for infants and mothers. The interpretation of this correlation, however, is confounded by potential selectivity in the choice of birth mode. We use birth records from California, merged with hospital and emergency department (ED) visits for infants and mothers in the year after birth, to study the causal health effects of cesarean delivery for low-risk first births. Building on McClellan, McNeil, and Newhouse (1994), we use the relative distance from a mother's home to hospitals with high and low c-section rates as an instrument for c-section. We show that relative distance is a strong predictor of c-section but is orthogonal to many observed risk factors, including birth weight and indicators of prenatal care. Our IV estimates imply that cesarean delivery causes a relatively large increase in ED visits of the infant, mainly due to acute respiratory conditions. We find no significant effects on mothers' hospitalizations or ED use after birth, or on subsequent fertility, but we find a ripple effect on second birth outcomes arising from the high likelihood of repeat c-section. Offsetting these morbidity effects, we find that delivery at a high c-section hospital leads to a significant reduction in infant mortality, driven by lower death rates for newborns with high rates of pre-determined risk factors.

David Card
Department of Economics
549 Evans Hall, #3880
University of California, Berkeley
Berkeley, CA 94720-3880
and NBER
card@econ.berkeley.edu

Alessandra Fenizia
Department of Economics
549 Evans Hall, #3880
University of California, Berkeley
Berkeley, CA 94720-3880
afenizia@econ.berkeley.edu

David Silver
Department of Economics
Julis Romo Rabinowitz Building
Princeton University
Princeton, NJ 08544
silverdw@gmail.com

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The current rate of cesarean delivery in the U.S. is both high on average and widely variable across hospitals and regions, with little relation to differences in medical need (see e.g., Baicker et al. 2008; Kozhimannil et al., 2014). Though c-sections are clearly beneficial in some situations, a large body of research suggests that cesarean delivery for low-risk pregnancies is associated with *worse* outcomes for infants, mothers and subsequent births.[1] These findings have led many experts – including the American College of Obstetricians and Gynecologists (2014) – to argue for policies to reduce the rate of c-sections for births that could be safely delivered vaginally.

A fundamental problem in assessing the health effects of c-section is the selectivity of observed delivery mode. Even among low-risk first births, where pre-scheduled c-sections are relatively rare (Declercq et al, 2006), a quarter of deliveries end with a cesarean procedure - often when there are indications that labor has stalled or the fetus is under stress (Zhang et al., 2010). Much of the variation in c-section rates appears to be due to differences in how long labor is allowed to progress before patients are recommended for the procedure.[2] What is needed in this setting to evaluate the relative costs and benefits of "marginal" c-sections is exogenous variation in providers' willingness to wait for vaginal deliveries to run their course.[3]

In this paper we exploit the fact that hospitals have different rates of cesarean delivery for low-risk pregnancies, and that many women deliver at the nearest hospital, to derive estimates of the impacts of c-section on health outcomes of infants and mothers whose choice of delivery mode is determined by distance. Building on McClellan, McNeil, and Newhouse (1994), we classify hospitals into two groups based on their average c-section rate for low-risk first births (LRFBs), and use the *relative*

[1] See e.g., Clark and Silver (2011), Gregory et al. (2011), Goer et al. (2012), and Hyde et al. (2012).
[2] Declercq et al. (2006) report that over 90% of mothers who experienced a primary c-section after trial of labor attribute the "idea to have a cesarean" to their care provider.
[3] Two prospective RCTs of "active management of labor" interventions to reduce c-section rates (Lopez-Zeno et al., 1992; Frigoletto et al. 1995) reached different conclusions about whether such programs had an effect on c-section rates. More recently, Gimovsky and Berghella (2016) implemented a small (N=78) RCT to extend labor for women with a prolonged second stage, which substantially reduced c-section rates. These studies were under-powered for studying subsequent health effects on mothers and infants.

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distance from a mother's home zip code to the nearest high c-section (H) hospital versus low c-section (L) hospital as an instrument for the probability of cesarean delivery.

We implement this approach using a California data set that combines hospital discharge records for mothers and newborns, birth certificate information, in-patient and out-patient records for mothers in the year before birth, and similar records for mothers and infants in the year *after* birth. These data provide policy-relevant indicators of postpartum health for a large sample of LRFBs as well as detailed information on demographic characteristics and risk factors. We also link mothers to subsequent births, allowing us to study the effects of primary c-section on fertility and health outcomes at the second birth.

We show that after controlling for local hospital areas and a few maternal characteristics relative distance is uncorrelated with a large set of pre-determined risk factors, but strongly affects the probability of c-section – particularly intrapartum procedures. We then characterize the compliers whose c-sections can be attributed to relative distance. Under "LATE-like" assumptions these are mothers whose hospital choice is affected by distance **and** whose delivery mode is determined by the type of hospital they select. We interpret the latter condition as isolating births with two key features: first, that providers at high c-section hospitals would recommend intervention at an earlier stage in the labor process; and second, that patients would comply.

Empirically we distinguish *hospital compliers* who shift from L to H hospitals when they are relatively closer to an H hospital, and *procedure compliers* who switch from vaginal to cesarean delivery under the same conditions. We find that procedure compliers are more heavily selected: 69% have less than high school education (versus 52% of hospital compliers and 41% of LRFBs); and 83% are covered by Medicaid or other government insurance (versus 61% of hospital compliers and 43% or all deliveries). We conclude that poorer and less-educated women are more likely to select a nearby hospital, and are more likely to undergo a c-section if they present at a high c-section hospital. The flexibility of their delivery outcomes contrasts with the rigidity exhibited by the physician mothers studied by Johnson and Rehavi (2016) and suggests that lower-SES patients may be particularly susceptible to hospital-specific practice differences.[4]

At birth we find that infants delivered by c-section have higher Apgar scores and lower rates of birth-related injuries but are more likely to be placed on mechanical ventilation, consistent with previous studies which point to decreased lung function as a leading unintended effect of c-section.[5] Complying mothers who deliver by c-section have lower rates of trauma to the perineum and vulva, and substantially shorter times between arrival at the hospital and delivery, indicative of the relatively long labors for members of the complier group who deliver vaginally.

In the year after birth we find a large, statistically significant impact of cesarean delivery on the probability of ED admissions, over half of which are attributed to respiratory-related diagnoses. The magnitude of the estimated effect is highly robust to inclusion or exclusion of 17 measured risk factors, including birthweight and prenatal smoking, prenatal care, and mother's health conditions. The effect also cumulates steadily over the follow-up period, suggestive of a persistent health gap between cesarean and vaginal births. For mothers the estimated impacts of cesarean delivery are only one-tenth as big, and far from statistically significant, although we cannot rule out some impact on the probability of ED visits.

An issue for interpreting the effect on ED admissions of infants is that high c-section hospitals may have other practices that affect health independently of the c-section channel.[6] To address this we examine breech presentation births. Hospital

[4] A large literature (see Chandra et al., 2012) studies the relative contribution of supply- and demand-side factors in explaining regional variation in medical treatment. Recent research suggests that supply-side determinants – including physician risk preferences and abilities, supplier-induced demand, organizational factors, and specialization – explain a substantial share of regional variation, while patient preferences are relatively unimportant (e.g., Finkelstein et al. 2016; Cutler, Skinner, Stern, Wennberg 2017). Our results imply that supply-side factors may be even more important for certain patients (e.g., lower SES mothers).

[5] See App. A for an overview of the literature. The causal channel identified in some existing literature (e.g., Hyde et al. 2012) is through the transfer of microbes during labor that lead to differences in immune system development. Jachetta (2014) explores the effect of cesarean delivery on asthma using malpractice insurance premiums as an instrument for c-section.

[6] McClellan et al. (1994) termed this the "correlated treatments" problem. A similar concern arises in other settings where randomization to "gatekeepers" (e.g. physicians, hospitals or judges) with varying treatment styles is used to generate variation in a treatment of interest, e.g., Maestas et al. (2013), Aizer and Doyle (2015).

choices of mothers with breech presentation are affected by relative distance but 98% deliver by c-section, so they are effectively "procedure always-takers". Reassuringly, we find that delivery at an H hospital has no effect on readmission rates of breech deliveries. We also compare the estimated first-stage effects on the probability of c-section and the reduced-form effects on ED admissions across 4 quartiles of the predicted probability of delivering by c-section at H hospitals. We find that the reduced-form effects vary proportionately with the first-stage effects, as would be expected if the causal channel runs through c-section.

We go on to examine two longer-term indicators of maternal health: fertility and second-birth outcomes. Our point estimates show no systematic or significant effect of cesarean delivery on the probability of a second birth up to 4 years after the first birth, though our sample sizes for these impacts are limited (due to our 5-year sample window) and we cannot rule out meaningful impacts of either sign. We find a large positive effect of a first c-section on the risk of c-section at second birth, reflecting the low rate of vaginal birth after c-section in our sample. Consistent with early scheduling of repeat c-sections we find negative effects on gestation and birth weight of the second child after primary c-section. We also find evidence of higher rates of ED visits for the second child, comparable to the effect observed for the first.

Finally, we consider the effects of hospital delivery practices on infant deaths. Delivery at a high c-section hospital is associated with a relatively large reduction in infant mortality -- on the order of 2.5 fewer deaths per 1000 births -- with *p*-values around 0.02. This effect is entirely driven by reductions in deaths for infants with higher predicted death rates (based on pre-determined factors): for the bottom two-thirds of the risk distribution we find no effect of delivery at H hospitals. Given the concentration of the effect among a relatively small subset of births we cannot reliably disentangle how much of the lower death rate is attributable to ***earlier*** c-sections at H hospitals (i.e., an intensive margin effect), versus ***more*** c-sections (an extensive margin effect). In any case, the mortality reductions per c-section at these hospitals are large enough to potentially offset the higher morbidity effects of cesarean delivery, suggesting the need for caution in pursuing policy changes to reduce c-section rates at these hospitals.[7]

II. An Overview of C-Section and Our Modeling Approach

a. Hospital Setting

Figure 1 provides a stylized overview of the pathways leading to c-section. The left side of the figure shows the pathway for mothers with a planned c-section. This group includes women who have had a previous c-section and those with breech pregnancy, multiple fetuses, and risk factors like obesity and eclampsia (Declercq et al, 2006; Zhang et al. 2010). Their c-sections occur with no attempted labor and are commonly classified as "scheduled."

To the right are mothers who reach normal term with no scheduled intervention. Typically, a mother-to-be shows early signs of labor and is admitted to hospital where her progress is monitored and pain relief and labor-augmenting medications are administered.[8] Barring other factors a decision to perform c-section is reached when labor time exceeds the threshold T_H (which depends on maternal characteristics and the specific hospital) resulting in an "unscheduled" or intrapartum c-section. Practices appear to vary widely over how long to allow labor to proceed (Zhang et al., 2010; Kozhimannil et al, 2013, 2014), leading to wide variation across hospitals in the average rate of intrapartum c-section. Similar variation exists in the decision process for mothers whose gestation has exceeded normal bounds, resulting in earlier or later admission to the hospital, more or less aggressive induction, and earlier or later recommendations for c-section.

Given these two very different paths to c-section we focus on low-risk first births, eliminating twins, breech presentations, births to mothers younger than 18 or

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[7] In unreported analysis, we tested for Roy-style selection on survival gains using a generalized control-function approach (following Brinch et al. 2017). We find no evidence of this type of sorting.

[8] Declercq et al. (2006) report that that 76% of all U.S. mothers had epidural anesthesia during labor. Many practitioners believe this slows labor and makes c-section more likely, though the evidence is controversial – see Howell (2000) and Klein (2006).

over 35, and five other risk factors (see below).[9] We identify the causal effect of c-section using a patient's proximity to hospitals with higher and lower average c-section rates for low-risk first births. Our interpretation is that these cross-hospital differences are mainly due to differences in the average time labor is allowed to proceed before performing c-section (i.e., in the mean of T_H in Figure 1).[10] Several factors could play a role in this variation, including financial incentives, malpractice pressures (Baicker et al, 2006), and differences in risk aversion. Rather than try to identify these factors, however, we take a data-driven approach and simply classify hospitals based on their average c-section rates for LRFBs.

b. Econometric Framework

Given the timing of birth events we posit a triangular system for the choice of a high c-section hospital by mother *i* (denoted by H_i), the choice of cesarean delivery (C_i) and a health outcome for the baby or mother (y_i). A linear version of such a system is:

$$H_i = \delta_0 + \delta_1 Z_i + \delta_x X_i + u_i \qquad (1)$$

$$C_i = \lambda_0 + \lambda_1 H_i + \lambda_2 Z_i + \lambda_x X_i + v_i \qquad (2)$$

$$y_i = \beta_0 + \beta_1 C_i + \beta_2 H_i + \beta_3 Z_i + \beta_x X_i + \varepsilon_i \qquad (3)$$

where Z_i is a measure of the relative distance from the mother's home to a low versus high c-section hospital, and X_i is a vector of controls (which we assume for simplicity are indicators for a set of mutually exclusive subgroups). Equation (1) says that the choice of hospital is affected by relative distance. Equation (2) says that the probability of c-section is affected by which type of hospital a patient chooses and possibly by relative distance. Equation (3) says that the health outcome is affected by whether the delivery is performed by c-section or not, and possibly by H_i and Z_i. Our primary interest is in the effect of cesarean delivery, represented by the coefficient β_1.

[9] We do not eliminate "scheduled" c-sections since the classification depends on indicators of labor on the mother's discharge record which are known to be under-reported (Henry et al, 1995).

[10] Consistent with this, we find that relative distance mainly affects the rate of unscheduled c-sections, and that there is a strong effect of relative distance on the elapsed time between the admission of the mother and the birth (see below).

Equations (1) and (2) imply that there is an induced first-stage relationship between relative distance and the probability of c-section:

$$C_i = \pi_0 + \pi_1 Z_i + \pi_x X_i + \eta_i \tag{4}$$

where $\pi_1 = \lambda_1\delta_1 + \lambda_2$. Similarly, equations (1)-(3) imply a reduced-form relationship between relative distance and the health outcome:

$$y_i = \tau_0 + \tau_1 Z_i + \tau_x X_i + \xi_i \tag{5}$$

where $\tau_1 = \beta_1\pi_1 + \beta_2\delta_1 + \beta_3$.

c. Potential Outcomes, Compliers, and Interpretation of IV

Next we use a potential outcomes framework to an interpret IV estimate of the coefficient β_1 in equation (3). For simplicity we focus on a binary version of our relative distance measure, Z_i^B, which indicates whether a mother's home is closer to a high c-section hospital or not.[11] Let H_{0i} and H_{1i} represent indicators for whether mother *i* would choose a high c-section hospital when $Z_i^B = 0$ or $Z_i^B = 1$. Similarly, let C_{0i} and C_{1i} represent indicators for whether she would deliver by c-section when $Z_i^B = 0$ or $Z_i^B = 1$. A given mother's potential responses to changes in Z_i^B are represented by the pairs (H_{0i}, H_{1i}) and (C_{0i}, C_{1i}). In principle, there are 16 different types of mothers, enumerated in App. Table 1. For example, the group with $(H_{0i}, H_{1i}), (C_{0i}, C_{1i})$ = (1,1), (0,0) consists of mothers who are H always-takers and C never-takers.

We impose three assumptions that restrict the possible H and C combinations. First, we assume there are no H defiers. This is a standard LATE assumption and is equivalent to assuming that all patients weakly prefer closer hospitals (similar to Einav et al., 2016). Second, we assume that distance has no direct effect on delivery mode, conditional on the type of hospital (i.e., $\lambda_2 = 0$ in equation 2). This exclusion restriction rules out the possibility, for example, that being closer to the hospital affects the stage

[11] In our analysis below we use both a continuous measure of relative distance and an indicator for being closer to a high c-section hospital, and find they give very similar IV estimates. Characterization of compliers with a continuous instrument is more difficult. We present an extended analysis in App. B.

of labor at arrival, which in turn affects the probability of c-section. To address concerns about the role of travel time we include a measure of the mother's distance to the nearest hospital of any type in the control vector X_i. Empirically, however, we find that distance to nearest hospital does not predict c-section, suggesting that this concern is minimal.[12] A third assumption is that H-compliers never switch from cesarean to vaginal birth. This monotonicity condition rules out rank-reversals in treatment intensity for H compliers. We view this as plausible given the homogeneity of our LRFB sample.[13]

Under these assumptions only seven (H, C) combinations are relevant: four representing groups of mothers whose choice of hospital *and* delivery mode is unaffected by distance, and three subgroups of H-compliers: those who always deliver vaginally (*H complier & C never-taker*); those who always get c-section (*H complier & C always-taker*); and those who switch from vaginal to cesarean delivery when they are induced to choose a high c-section hospital by relative distance (*H&C complier*). Only H&C compliers change delivery mode in response to the value of the instrument, so these are the "procedure compliers".

In addition, we make the standard conditional independence assumption that Z_i^B is as good as randomly assigned conditional on X_i. It is then easy to show that:

$$E[H_i \mid Z_i^B = 1, X_i] - E[H_i \mid Z_i^B = 0, X_i] = \delta_1(X_i) = P(H\ complier \mid X_i) \qquad \text{(6a)}$$

$$E[C_i \mid Z_i^B = 1, X_i] - E[C_i \mid Z_i^B = 0, X_i] = \pi_1(X_i) = P(H\ \&\ C\ complier, X_i) \qquad \text{(6b)}$$

Assuming that X_i is a set of subgroup indicators, the coefficient δ_1 of relative distance in the first-stage model (1) for *H* identifies a weighted average of the fraction of H compliers in each subgroup, where the subgroup's weight equals its relative size times the relative magnitude of its within-group variance in Z_i^B (see App. B). Similarly, the

[12] In our first-stage regressions with a binary indicator for c-section as the dependent variable the estimated coefficient of distance to the nearest hospital is 0.015 per 100 miles (s.e. = 0.039), implying that, if anything, extra travel time to the nearest facility tends to *increase* a mother's likelihood of c-section.

[13] Rank invariance is routinely assumed in the analysis of quantile treatment effects (e.g., Chernozhukov and Hansen, 2005). Currie and MacLeod (2017) consider a setting where rank-reversal could be important. They document substantial heterogeneity in physician diagnostic ability, which could lead to rank reversals if better diagnosticians are concentrated at certain hospitals. Our low-risk first birth sample excludes most of the higher-risk births considered by Currie and MacLeod (2017).

coefficient π_1 in the first stage model (4) for c-section identifies a weighted average of the fraction of H&C compliers in the various covariate subgroups.

Although conditional independence is not directly testable, we evaluate its plausibility by "holding back" a set of observable risk factors from our vector of basic controls X_i. We then check whether these risk factors are orthogonal to relative distance after taking account of the basic controls (which consist of maternal demographics and neighborhood identifiers).

Finally we consider the potential health outcomes for a given birth. Let $Y_i(c,h,z)$ represent the outcome that would be observed for birth *i* conditional on delivery mode, hospital type and relative distance. As a baseline we assume that

$$Y_i(c,h,z) = Y_i(c) \in \{Y_i(0), Y_i(1)\} \tag{7}$$

i.e., that the health outcome depends only on delivery mode, with no dependence on H or Z_i^B. This assumption, plus our assumptions on the possible (H, C) combinations and on the conditional independence of Z_i^B, imply that:[14]

$$E[y_i \mid Z_i^B = 1, X_i] - E[y_i \mid Z_i^B = 0, X_i] = \pi_1(X_i) E[Y_i(1) - Y_i(0) \mid H \,\&\, C\, complier, X_i]. \tag{8}$$

In other words, the reduced-form difference in average outcomes, conditional on X_i, is proportional to the fraction of H&C compliers times the average treatment effect on H&C compliers. Together equations (6b) and (8) imply that an IV estimate of the coefficient β_1 in equation (3) using Z_i^B as an instrument for C_i yields an estimate of a *weighted average* of the treatment effects of c-section on H&C complying mothers in each X_i −subgroup, where the weights combine the relative size of the group, the relative within-group variance in Z_i^B, and the relative size of the first stage effect on the group (see App. B).

A more general assumption is that the potential outcomes depend on H and C:

$$Y_i(c,h,z) = Y_i(c,h) \in \{Y_i(0,0), Y_i(0,1), Y_i(1,0), Y_i(1,1)\}.$$

[14]The only mothers in a given covariate subgroup who switch potential outcomes when the instrument changes from 0 to 1 are the H&C compliers. For these mothers we observe $Y_i(0)$ when $Z_i^B = 0$ and $Y_i(1)$ when $Z_i^B = 1$. Since the share H&C complier is $\pi_1(X_i)$ equation (8) follows immediately.

In this case proximity to an H hospital can affect the health outcomes of all three subgroups of H-compliers: H&C compliers (who switch hospital type and delivery mode when Z_i^B changes from 0 to 1); H-complier/C-always takers (who switch hospitals but always deliver by cesarean); and the H-complier/C-never takers (who switch hospitals but always deliver vaginally). Let $\rho_1(X_i)$, $\rho_2(X_i)$ and $\rho_3(X_i)$ represent the shares of these 3 groups among the H-compliers (conditional on X_i), and let

$$\mu_1(X_i) = E[Y_i(1,1) - Y_i(0,0) \mid H \,\&\, C\,complier, X_i)$$

$$\mu_2(X_i) = E[Y_i(1,1) - Y_i(1,0) \mid H\,complier \,\&\, CAT, X_i)$$

$$\mu_3(X_i) = E[Y_i(0,1) - Y_i(0,0) \mid H\,complier \,\&\, CNT, X_i)$$

represent the corresponding treatment effects of delivering at an H hospital.[15] For a given X_i – group an IV procedure using Z_i^B as an instrument for H_i yields an estimate of:

$$\rho_1(X_i)\mu_1(X_i) + \rho_2(X_i)\mu_2(X_i) + \rho_3(X_i)\mu_3(X_i). \tag{9a}$$

Using Z_i^B as an instrument for C_i, on the other hand, yields an estimate of:

$$\frac{1}{\rho_1(X_i)}\big(\rho_1(X_i)\mu_1(X_i) + \rho_2(X_i)\mu_2(X_i) + \rho_3(X_i)\mu_3(X_i)\big). \tag{9b}$$

In the baseline case where $\mu_2(X_i) = \mu_3(X_i) = 0$, treating C_i as the endogenous variable gives rise to a consistent estimator of $\mu_1(X_i)$. More generally, if delivery at an H hospital also affects C-always takers or C-never takers, then treating C_i as the endogenous variable could over- or under-state the effect on H&C compliers.

We evaluate the dependence of health outcomes on hospital type by studying outcomes for breech presentations.[16] These infants are nearly always delivered by c-section, owing to the perceived risks of and lack of current expertise in delivering breech

[15] μ_1 is a combined treatment effect of c-section and delivery at an H hospital for H&C compliers, while μ_2 and μ_3 represent treatment effects of delivery at an H hospital for mothers whose delivery mode is independent of hospital type.

[16] To ensure comparability to our primary sample, our sample of breech babies consists only of births meeting all criteria for being low-risk first births *except* for their presentation.

babies vaginally (Hannah et al, 2000; ACOG, 2006).[17] Thus, an IV procedure using Z_i^B as an instrument for H_i yields an estimate of

$$E[Y_i(1,1)-Y_i(1,0)\,|\,breech, H\ complier, X_i)$$

A finding that this effect is close to zero suggests that any independent effect of delivery at a high c-section hospital is also likely to be small for LRFBs.

We also perform a second check by stratifying our sample into groups with different first-stage effects on the probability of delivering by c-section, and on the probability of delivering at an H hospital. We then check whether the reduced-form health effects for different subgroups vary proportionately with their first-stage effects on the probability of c-section, or with their first-stage effects on the probability of delivery at an H hospital.

III. Data Sources, Sample Overview, Relative Distance Instrument

a. Data Sources

We use a linked data set created by the California Office of Statewide Health Planning and Development (OSHPD) that combines information from patient discharge (PD) records, emergency department (ED) records, ambulatory surgery (AS) records, and vital statistics (VS) records for all in-hospital births in the state between 2007 and 2011.[18] Specifically, PD records for the birth stay of the mother and the infant are linked with birth certificate data and PD/ED/AS and VS records over the following year for mothers and infants and with PD/ED/AS records for mothers in the year prior to the birth. The resulting data set includes birth certificate information on the mother (e.g., demographics, weight gain and smoking during pregnancy) and the infant (gestation, birthweight, Apgar score), as well as PD-derived information on diagnoses at the delivery. The pre-birth PD/ED/AS records provide additional measures of maternal health (such as the number of ED visits in the year prior to birth). The post-partum

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[17] Thus breech deliveries provide "identification at infinity" – see Chamberlain (1986), Heckman (1990).
[18] This is known as PDD/ED/AS/Linked Birth Cohort data, and is available to researchers through OSHPD. See App. C for more information on the characteristics of the data and the derivation of our samples.

PD/ED/AS and VS records provide our main health outcomes (hospital visits and infant death). Importantly, we can link later births for the same mother, enabling us to study effects on the probability of additional births and on the health outcomes associated with these births.

This data set has two key limitations. First, we do not observe physician office visits. This means that we miss some fraction of less urgent health problems for newborns and mothers than are treated in an office setting rather than at a hospital or ED/AS center. As discussed below, we suspect that the compliers for our distance-based instrument tend to use the ED more for routine care than other mothers, so we arguably capture a higher fraction of such problems than would be detected in a representative sample of all births.

A second limitation is that we have no direct information on several important pieces of clinical information, including whether a c-section occurred before or after a trial of labor.[19] In addition the reporting of certain secondary diagnoses (such as uterine inertia) appears to be endogenously related to the decision to perform c-section (e.g., for billing purposes). The offsetting benefit is that we have large sample sizes, allowing us to detect plausibly sized effects with an IV research design.

b. Sample Overview

Table 1 provides an overview of the characteristics of all 2.7 million births in California during our 5-year sample window (column 1) and of all low-risk first births (column 2). We define LRFBs as singleton non-breech first births delivered at 37+ weeks of gestation. These restrictions correspond to the two lowest risk groups in Robson's (2001) widely used classification. In addition we eliminate births from mothers under 18 or over 35, and with 5 other risk factors: eclampsia, pre-eclampsia, growth restrictions, mother's BMI >90th percentile, and >20 pre-natal visits. We do not condition on other

[19] We follow the existing literature (e.g., Gregory et al., 2002; Johnson and Rehavi, 2017) and classify labor as having occurred prior to c-section based on the presence of at least one of a set of ICD-9-CM diagnosis codes devised by Henry et al. (1995) that indicate fetal distress during labor or dystocia. Henry et al.'s analysis showed that these indicators are measured with some error relative to clinical indications.

observable risk factors (such as birthweight), allowing us to ***test*** for orthogonality of our distance-based instrument with factors that are correlated with the health of mothers or infants.

Column 1 shows that about one-half of all California mothers are Hispanic, one-half have no more than a high school education, and one-half have their delivery stay paid by Medi-Cal. All three rates fall to around 40% among LRFB mothers. LRFB mothers are also younger and weigh less. LRFB mothers are similar to all mothers in their average number of prenatal visits (about 12) and their probability of an ED visit in the year before delivery (20%), but have slightly longer mean gestation (40 vs. 39 weeks). Overall about one third of all California births were delivered by c-section during our sample period, compared to 25% for LRFBs. These fractions are very similar to national averages reported by Osterman and Martin (2014).

c. Construction of Relative Distance Instrument

Our distance-based IV strategy relies on a prior classification of hospitals. Since c-section rates vary systematically across regions of California, we elected to define high and low c-section hospitals *within* Health Referral Regions (HRRs).[20] As detailed in App. C, we fit a logit model for c-section on our LRFB sample, including hospital dummies and a set of risk factors. We classify a hospital as "high c-section" (H) if its risk-adjusted c-section rate (i.e., the hospital effect in the logit) is above the patient-weighted mean rate for all hospitals in its HRR. Otherwise it's classified as low c-section (L).

App. Table 2 shows that 29% of LRFBs were delivered by cesarean at H hospitals, compared to 22% at L hospitals. About two-thirds (0.048/0.075) of the difference is attributable to a higher rate of unscheduled (intrapartum) c-sections. H hospitals also differ in other ways: they are more likely to be for-profit (18% vs. 9%) and less likely to have a NICU unit (74% vs. 86%). Nevertheless, H and L hospitals have very similar

[20] Hospital referral regions (HRRs) represent regional health care markets for tertiary medical care, defined by the Dartmouth Atlas (see Dartmouth Atlas, undated). There are 25 HRRs in our sample of LRFBs. If we classified hospitals on a statewide basis, we would have many more high c-section hospitals in Southern and Central California and many more low c-section hospitals in Northern California.

average numbers of deliveries per year (3,695 versus 3,635), suggesting that the classification is not driven by volume-related differences.

Using this classification, we then calculate the distance from the centroid of a patient's home zip code to the centroid of the zip code of the nearest H hospital (d_{Hi}) and the nearest L hospital (d_{Li}). We define the relative distance $Z_i \equiv d_{Li} - d_{Hi}$ and an indicator for being closer to an H hospital $Z_i^B = 1[Z_i \geq 0]$. We also define the distance to nearest hospital $d_i^{MIN} = \min[d_{Li}, d_{Hi}]$.

The third column of Table 1 presents the characteristics of the subsample of LRFBs that have patient home zip code information, non-missing values for all the control variables we use in our main specifications, and have $d_{Hi} \leq 20$ miles, $d_{Li} \leq 20$ miles, and ≤ 20 miles between the mother's home zip code and the actual hospital she delivered in. These restrictions eliminate just over 20% of LRFBs, leaving us with a final analysis sample of 491,604 births. A comparison of columns 2 and 3 suggests that this sample is quite similar to the overall LRFB sample.

Figure 2 illustrates the strong relationship between relative distance and hospital choice for LRFB mothers. Here we plot the fraction of mothers in each zip code who deliver at an H hospital against the value of Z_i. The data suggest a nearly symmetric S-curve relationship, tending toward a minimum of about 10% when $Z_i < -15$ and a maximum of about 90% when $Z_i > 15$. We interpret this symmetry as evidence that most mothers treat H and L hospitals as exchangeable, though some are H always-takers and others are H never-takers.

d. Instrument Validity

A concern with any IV strategy is that the instrumental variable is correlated with unobserved determinants of the outcome of interest. To assess this concern we estimated a series of OLS models for each of a set of held-back risk factors that are predetermined at the delivery date, looking for evidence of a correlation with relative distance. Specifically for the jth risk factor we fit a model of the form:

$$R_i = \psi_0 + \psi_1 Z_i + \psi_X X_i + \zeta_i$$

where X_i is the set of basic control variables we include in all our outcome models. This includes year effects, controls for mother's demographics[21], an indicator for whether the mother had any visits to the ED in the year before the birth, mean income in the mother's home ZIP code, distance to the nearest hospital (d_i^{MIN}), and a set of Health Service Area (HSA) dummies.[22] The latter ensure that all our models compare mothers and infants from the same narrow geographic area.

We considered a set of 16 potential risk factors: length of gestation, birthweight, an indicator for low birth weight, 3 measures of prenatal care, indicators for maternal diabetes, herpes, and asthma, five measures of maternal smoking, and counts of the number of ED visits and total medical facility visits (hospital+ED+AS) by the mother in the year prior to the birth. We also used these 16 variables and X_i to estimate logit models for the probability of c-section and the probability of an ED visit by the infant in the year after birth, then formed predicted probabilities of these two outcomes that combine the risk factors in a potentially more powerful way.

The resulting estimates of the ψ_1 coefficients are reported in App. Table 3. Only two risk factors – the number of ED visits by the mother in the year prior to the birth and the indicator for maternal herpes – have significant partial correlations with Z_i . Interestingly, the former is positively related to relative distance (with t=2.0) while the latter is negatively related (with t=-2.2).

To summarize our findings we multiplied the estimated coefficient for each risk factor by 10 and divided by its standard deviation (i.e., $10\hat{\psi}_1 / std(R_i)$), providing an estimate of the "effect size" of a 10-mile change in relative distance. The resulting effect sizes are plotted in Figure 3 along with their confidence intervals. We order the coefficients by their precision, resulting in a "forest plot"-style graph (Hedges and Olkin,

[21]These are dummies for mother's age, education, race, insurance type, and presence of father, and continuous controls for height (cubic), weight (cubic), and BMI.
[22] Hospital service areas (HSAs) are designed to be relatively self-contained with respect to hospital care, and are defined by the National Center for Health Statistics: there are 176 in our sample.

1985). For reference, we also highlight the ±0.01 effect size range. The pattern of estimated effect sizes suggests that relative distance has a small and unsystematic effect on the risk factors, controlling for the X_i variables. In particular we find a precisely estimated zero effect for birthweight (a standard marker of infant health), and relatively small and opposite-signed effect sizes for the probability of a c-section delivery or an ED visit in the year after birth.

IV. First Stage Relationships and Complier Characteristics

The first two rows of Table 2 present the estimated first-stage relationships between relative distance and the probabilities of giving birth in an H hospital (H_i) or delivering by cesarean (C_i). Both the continuous measure of relative distance (in the first column) and the binary version (second column) have strong effects on H_i, with partial F-statistics of 104 and 69, respectively. The estimate the first row of the first column means that a mother living 10 miles closer to an H-hospital is 15.9 percentage points (ppt's) more likely to deliver at an H hospital (controlling for X_i), while the estimate in the second column 2 means that a mother who is closer to an H hospital has a 10.1 ppt higher probability of delivering at such a hospital. The relative distance variables also have a strong effect on C_i, with F statistics of 37 and 26, respectively. A mother living 10 miles closer to an H hospital has a 1.8 ppt higher probability of c-section; while a mother who is simply closer to an H-hospital than an L hospital has a 1.1 ppt higher probability of c-section.

The third and fourth rows of Table 2 show the separate effects of relative distance on scheduled and unscheduled c-sections. These estimates imply that 73% (using Z_i) or 80% (using Z_i^B) of the extra c-sections attributable to proximity to H hospitals are intrapartum procedures. Since indications of labor appear to be significantly under-reported, however, we believe these fractions should be interpreted as ***lower bounds*** on the unscheduled shares. Indeed, the 16% under-reporting rate found by Henry et al. (1995) suggests that 60-80% of the scheduled c-sections attributed to distance are actually intrapartum procedures with no reported indications of labor.[23]

The next rows of Table 2 show the estimated effects of relative distance on the probability of cesarean delivery at H and L hospitals, respectively. The estimates imply that the net increase in overall c-sections attributable to proximity to H hospitals is a result of a relatively large rise in the rate of c-sections at H hospitals, offset by a reduction in c-sections at L hospitals. The reduction reflects the behavior of H-complier/C-AT's, who switch hospital types in response to relative distance but have a cesarean delivery regardless of where they present.

Finally, in the bottom panel of the table we show the implied breakdown of the overall H-complier population into its three constituent subgroups (H&C compliers, H compliers/C-AT's, and H compliers/C-NT's). To calculate the fractions using the continuous distance instrument we use the changes in probabilities associated with a 10-mile reduction in relative distance to an H hospital. Regardless of the choice of instrument we find that H&C compliers represent about 11% of the overall H complier population, while C-AT's represent about 20%. This means that a shift in relative proximity to an H hospital induces an 11-ppt increase in the overall c-section rate of the H-compliers, from 20% to 31%.[24]

To further explore the characteristics of the extra c-sections attributable to relative distance, we used the method suggested by Abadie (2003) to estimate the mean characteristics of the H (or distance) compliers and the H&C (or procedure) compliers. The results are summarized in Table 3. We show the distributions of the two complier groups across race/ethnicity groups, mother's education categories and insurance types in Panels A and B. Panel C shows a number of maternal and infant characteristics, including mother's height, maternal use of the ED prior to the birth, and

[23] Henry et al. (1995)'s data show that of 831 primary c-sections for non-breech births that were clinically coded as having trial of labor, 701 had indications of labor on the discharge record, implying a 15.6% under-reporting rate. If we assume that distance only affects unscheduled c-section rates we would expect to find that *scheduled* c-sections account for 15.6% of the overall first stage effect.

[24] Among all LRFBs, those delivering at L hospitals have a 22% c-section rate, while those delivering at H hospitals have a 28.9% c-section rate (see App. Table 2). The somewhat larger gap among H-compliers likely reflects H-compliers being more subject to hospital policies than the broader LRFB group.

the baby's gender and birthweight. We also show the fraction of infants at high risk to use the ED, based on a simple logit model that includes maternal characteristics and other predetermined factors.[25] Finally, Panel D shows the shares of births from low air quality areas, as measured by ozone and particulate (PM2.5) pollution. To the extent that cesarean delivery leads to reduced lung function, the net effect on infant health is likely to be magnified for families that live in areas with worse air pollution.

A comparison of the demographic characteristics of all LRFB mothers (column 1) and the H-compliers (column 2) shows that the latter group are more likely to be Hispanic and to have at most a high school education. They are also more likely to have government insurance (mainly Medi-Cal).[26] They are not differentially selected from low-income ZIP codes but they do tend to reside in areas with worse air quality, and their babies have a higher predicted risk of visiting the ED. Consistent with evidence from other settings – e.g., Beckert et al.'s (2012) study of hip replacement patients – these comparisons suggest that less-advantaged families put more weight on distance in deciding which hospital to use.

The H&C compliers (column 3) are even more highly selected in all these dimensions: for example, only 14% have a college degree. H&C-complying mothers are also much more likely to live in lower income areas, and to have been users of the ED prior to the birth. Nearly all (91%) of these mothers' infants have above-median risk of a postpartum ED visit. H&C-complying mothers are also relatively likely to have short stature and to deliver a male baby.

An implication of the results in Table 3 is that differences in hospital practices have a larger impact on the probability of c-section for lower-SES mothers. This is the "flip side" of the findings reported by Johnson and Rehavi (2017), who conclude that the delivery modes of physician mothers are *less* responsive to the financial incentives faced

[25] The index combines risk factors for lower health and predictors of the use of the ED for care.

[26] The very low fraction of distance compliers with Kaiser insurance is consistent with the fact that Kaiser insured mothers would be expected to deliver at a Kaiser hospital in all but emergency situations. We have estimated our main models excluding Kaiser insurees and find that the first stage and reduced form effects are typically a little larger in magnitude, while the associated IV estimates are of very similar size. This is as expected given there are so few Kaiser insurees among the compliers.

by their doctors.[27] A second implication is that since the procedure compliers are relatively heavy predicted ED-users, our IV estimates of the effect of cesarean delivery on ED use may overstate the impacts for other groups who rely more on physician office visits for routine and non-emergent care.

V. Impacts of C-Section Delivery on Infant and Maternal Health

a. Outcomes at the Delivery

With this background, we turn to an analysis of the health effects of cesarean delivery. We begin in Table 4 by focusing on outcomes associated with delivery itself, including Apgar scores, incidence of a birth injury, admission to the NICU, and use of ventilation for the infant. For mothers, we focus on perineal laceration and other injuries incurred during labor, as well as the length of the hospital stay at delivery. For each outcome, we show the mean value among all LRFBs (in the first column), the OLS coefficient from a regression of the outcome on C_i and our basic controls (in the second), then the estimated reduced-form effects and IV estimates based on the continuous instrument Z_i and the discrete instrument Z_i^B.

Apgar scores are widely used as indicators of newborn health (see Casey et al., 2001).[28] We focus on the 5-minute test, which in an OLS regression has a small but significantly negative correlation with cesarean delivery. In IV models the correlation becomes relatively large and positive (0.5 to 0.65 points on a 1-10 scale), suggesting that there are unobserved determinants of Apgar scores that are negatively correlated with c-section, but that on average, cesarean delivery has a positive impact on newborns.[29] For birth injuries the OLS and IV estimates are both negative (i.e., cesarean delivery lowers the risk of injury) but the IV estimates are substantially larger in magnitude. The

[27] An earlier study Grytten et al. (2011) finds that in Norway, where c-section rates are among the lowest in the OECD, physician mothers are ***more*** likely to have c-section. They attribute this to enhanced agency of these mothers in the hospital setting.
[28] The Apgar is based on 5 components (breathing, heart rate, muscle tone, reflexes, and skin color) each of which is scored 0 1 or 2. See Finster and Wood (2005) for a brief history and discussion of the test.
[29] Altman et al. (2015) show that prolonged second stage of labor is associated with low Apgar scores. The positive IV effect of c-section we measure could be attributable to a shortening of labor.

IV estimates suggest that c-section reduces injuries associated with difficult labor (e.g., branchial plexus injury) at a rate of 1-2 per 100 c-sections, without offsetting rises in other types of injuries (e.g., lacerations attributable to the procedure itself).[30]

Comparing OLS and IV estimates of the effect of c-section on NICU admission we also find a switch from a positive (OLS) to a negative (IV) effect. In this case, however, the IV estimates are potentially confounded by the lower fraction of H hospitals with a NICU unit (74% vs. 86% at L hospitals). To the extent that NICU admission decisions are influenced by the availability of a unit in the same hospital, we might expect to see lower NICU admissions for deliveries at H hospitals. To assess this channel, we examined NICU admissions for breech births. Since nearly all breech births are delivered by c-section, with no differential between H and L hospitals, any reduced form effect of being closer to an H hospital for these babies is arguably attributable to supply-side considerations. In fact, we find that the reduced form effects for breech deliveries are comparable to the effects for LRFBs, though imprecise. Thus, the reduction in NICU admissions suggested by the IV models in Table 4 could be explained by the lower availability of NICU's at H hospitals, rather than a causal effect of c-section.

Last, we look at the incidence of ventilation (use of mechanical devices to aid the newborn in breathing). OLS models show that ventilation is more likely for newborns delivered by c-section, with an effect that is large (+1 percentage point) relative to the baseline rate (1.5% across all births).[31] This pattern is consistent with a large literature suggesting that breathing problems are more likely for babies delivered by c-section (See App. A). In contrast to the patterns we see for the 5-minute Apgar and the risk of birth injury, the IV models suggest even larger (though somewhat imprecisely estimated) effects. Even at the lower bound of its 95% confidence interval, the IV estimate based on the continuous distance instrument implies that about 3 per 100

[30] Alexander et al. (2006) study fetal injuries after c-section and note that the highest rates of injuries are for fetuses born after an unsuccessful trial of forceps or vacuum delivery. To the extent that doctors at H hospitals use these procedures less, and opt for c-section earlier, these injuries will be avoided.

[31] The 1.5% incidence rate we measure in our sample is comparable to the rate of 1.8% measured by Angus et al. (2001) using 1994 discharge data for California and New York.

babies delivered by c-section are placed on a ventilator. The point estimate implies that one-quarter of all cesarean deliveries for complying mothers end up on ventilation.

The next two rows of Table 4 present OLS and IV effects for two (related) measures of maternal injuries during labor: trauma to the perineum and vulva, and more serious (2nd degree or higher) perineal laceration (PL).[32] The former category includes the latter, as well as less serious (1st degree) PL's, vulvar and perineal hematomas, and anal sphincter tears. The rate of trauma injuries is relatively high for first-birth mothers (46% on average); about two-thirds of these are 2nd degree or higher PL's. As shown by the OLS estimates, trauma and PL are substantially lower for cesarean deliveries: indeed, the rates of both injuries are under 1% among c-section births. Interestingly, the IV estimates are even larger in magnitude.

To interpret the IV estimate for the rate of more serious PLs based on the binary instrument we used Abadie's (2003) method to estimate the "potential outcomes" for PL among the complier population. This approach suggests that for compliers who deliver vaginally at L hospitals the rate of PL is essentially 100% (the numerical estimate is 1.22, with a standard error of 0.27), while for those who deliver by c-section at H hospitals the rate is 0 (the numerical estimate is 0.01, with a standard error of 0.01). Our interpretation is that, for compliers who deliver vaginally because they are closer to L hospitals, the rate of more serious delivery-related injuries is high. This elevated rate is potentially indicative of a long and difficult labor, which also has been linked by other researchers to bad infant outcomes such as asphyxia (Chandra et al. 1997; Maghoma and Buchmann 2002; Aslam et al. 2014), and highlights a tradeoff between performing intrapartum c-sections and waiting for labor to progress.

In the final three rows of Table 4, we investigate the effects of c-section on the mother's length of stay (LOS) during the delivery episode. On average mothers undergoing c-section have longer recovery periods and thus longer LOS, a fact that is

[32] We have investigated a few other outcomes at birth, namely unplanned hysterectomy and asphyxia of the neonate. In our sample, the mean rate of unplanned hysterectomy is only 1 in 10,000; we find no evidence of an effect of c-section but the precision of our estimates is low. Asphyxia is also rare (3.2 per 1000), and we find very weak evidence (t=0.5) that c-sections reduce asphyxia. There is somewhat stronger evidence that c-section reduces asphyxia-related infant deaths.

reflected by our OLS estimate of an additional 1.4 days in the hospital for mothers undergoing c-section relative to vaginal delivery. We then split a mother's LOS into two components: (1) the number of days from a mother's admission to the birth, i.e. the length of labor,[33] and (2) the number of days from birth until the mother is discharged home, i.e. the post-birth LOS. Most of the gap in total length of stay for mothers undergoing c-section occurs ***after*** birth.

Turning to the IV estimates, we find that marginal c-sections among the procedure compliers have a somewhat shorter LOS, though the estimates are somewhat imprecise. Consistent with our discussion in Section IIa we find that the time from mother's admission to birth is significantly reduced (by around 0.6 days) when the delivery is by c-section. The duration of the post-birth stay is longer for compliers who deliver by c-section rather than vaginally (by about 0.4 days) but the difference is substantially smaller than the 1.3 days implied by the OLS estimate.[34] We conclude from these estimates that the complier group have relatively long labors when they deliver vaginally and require a long post-birth recovery – nearly as long as a post-cesarean recovery.[35]

b. Post Deliver Admission Outcomes

Table 5 summarizes the estimated impacts of cesarean delivery on admissions to hospitals, ambulatory surgical centers (ASCs) and EDs in the year after birth by infants and mothers. Focusing first on infants, the estimated OLS coefficients imply that cesarean delivery is associated with no change in the combined risk of any type of visit (in-patient, ASC or ED), but with an elevated risk of ED visits (with an effect of about 6

[33] We use this terminology somewhat loosely, as we do not observe exactly when the mother went into different stages of labor. We only know the days elapsed between her admission to the hospital and the birth date of the baby. According to Declercq et al. (2006), the mean time in labor for first-time mothers is around 11 hours. Consistent with this, the delay from admission to birth is 0 or 1 day in 95% of cases.
[34] The mean potential outcome of post-birth stays for the compliers is 2.7 days for a cesarean delivery and 2.4 days for a vaginal delivery. By comparison the average post birth stay for *all* vaginal births is 1.8 days.
[35] We have also investigated other indicators of prolonged labor, including a code on the birth certificate, which yields qualitatively similar results. However, average reported rates for prolonged labor on the birth certificate vary widely across hospitals (from 0 to 16%) so we are reluctant to attach much weight to this variable and do not use it elsewhere in the paper.

per 1000 c-sections), mainly attributable to visits for respiratory conditions. These patterns mean there is actually a **negative** OLS effect of c-section on in-patient and ASC visits, suggesting that babies delivered by c-section have less need for procedures that would be performed in these settings.

The reduced-form estimates for both the continuous and binary measures of relative distance show significant positive effects on ED visits in the year after birth. Scaling by the first stage, the IV estimates imply that *complier-driven* c-section deliveries have a substantially elevated risk of ED use in the year after birth, 60-70% of which is attributable to visits for respiratory-related conditions.

The OLS and IV results for mothers' readmissions present an interesting contrast to these estimates. OLS models show that LRFB mothers who deliver by c-section have a roughly 3 ppt higher probability of visiting a hospital, ED, or ASC in the following year. Given the average rate of about 15%, this is a reasonably large effect. In contrast to the case for infants, however, the reduced-form estimates show negligible effects of relative distance on combined in-patient and out-patient visits or ED visits, so the IV point estimates are 5-20 times smaller than the corresponding estimates for infants. Nevertheless, the imprecision in these estimates means we cannot reject the OLS results, or even somewhat larger effects.

c. Evaluating the Robustness of the Estimated Impacts on Infant ED Admissions

The estimated impacts of cesarean delivery on ED use by infants in Table 5 are quite large, but also somewhat imprecise. For example, using the continuous measure of distance a 95% confidence interval for the IV estimate extends from +0.2 to 1. To probe the robustness of these estimates we performed a number of checks. First, we estimated the reduced-form impacts on the probability of at least one ED visit using a series of time windows from 1 month to 12 months after the birth. The reduced-form estimates and associated confidence intervals based on the continuous distance measure are reported in App. Table 4 and plotted in Figure 4. For both versions of the instrument we find that the cumulative effect rises smoothly, as would be expected if there is a systematic health gap between cesarean and vaginal deliveries.

As a second check we developed a simple procedure to evaluate the sensitivity of the estimated reduced-form effect of relative distance to the inclusion or exclusion of other controls. As noted above we have 16 risk factors that are excluded from our basic control set X_i. In addition we included an additional measure of pre-birth medical use by the mother (a dummy for any in-patient visit in the year before birth), yielding a total of 17 extra risk factors. For each integer *J=0...*17 we performed a simple Monte Carlo exercise, randomly selecting subsets of *J* risk factors to be added to our baseline reduced form model.[36] We then calculated the minimum, maximum, mean and median value of the estimated reduced form effect for each *J*. These statistics are plotted in Figure 5.

The figure illustrates two important points. First, adding any subset of the extra controls has at most a small effect on the magnitude of the estimated reduced-form effect. Second, although extra controls can lead to slightly smaller or slightly larger reduced-form estimates, the range is symmetric. Following the logic of Altonji, Elder and Taber (2005) this suggests that the addition of other (unmeasured) risk factors would not be expected to lead to a systematic change in the estimated reduced form estimate.[37]

d. Testing Multiple Channels Using Breech Deliveries

As discussed in Section II, a concern with our distance-based IV strategy is that high c-section hospitals may have other treatment practices that contribute to the

[36] Note that *J=0* corresponds to our baseline model, while *J=17* corresponds to a model with all 17 factors.
[37] We also used Oster's (2018) approach to estimate a lower bound on the reduced form effect. This starts with an assumption on the maximal R-squared that could be achieved with all possible controls, then extrapolates from the change in the estimated coefficient from no controls to the available controls to form an extreme bound. In our case, if we start from the reduced form estimate with our baseline covariates and consider the effect of adding the 17 extra risk factors, and follow her suggestion of a 30% maximal increase in R-squared, her approach implies a bound of 0.13 on the reduced form effect of relative distance. If instead we start from the reduced form effect with only HSA fixed effects, her approach implies a bound of 0.10. The main factor in our baseline covariate set that has an impact on the reduced form coefficient estimates is maternal education.

reduced-form effect of relative distance on infant health. To evaluate this concern we examine breech presentation pregnancies. App. C describes the derivation of the sample and the mean characteristics of the 12,744 low-risk breech presentation first births (BPFBs) in our data set. Relative to LRFBs, mothers with breech presentation are less likely to be Hispanic, have higher education and are less likely to be covered by Medi-Cal. Most importantly, 98% of BPFBs are delivered by c-section. BPFBs are also less likely to visit EDs or hospitals in the year after birth: their mean probability of any inpatient or outpatient visit is 0.355, while their mean probability of an ED visit is 0.308 – both rates are about 3 ppt's ***below*** the corresponding rates for LRFBs.

The first 6 columns in Panel A of Table 6 compare the estimated first-stage and reduced-form models for LRFBs and BPFBs, utilizing our continuous measure of relative distance. Relative distance has a slightly ***larger*** effect on the choice of an H hospital by BPFBs than LRFBs, though we cannot reject similar effects (t=1.35). As expected, however, the first-stage effect on the probability of c-section for BPFBs is small in magnitude, suggesting that H hospitals are no more likely to perform c-section for BPFBs. Importantly, the reduced-form effect of relative distance on the probability of an ED visit in the year after birth is also small in size (-0.030), though somewhat imprecise. We infer that delivery at an H hospital has no large effect on the health of distance-complying BPFBs, nearly all of whom are c-section always-takers.

If one assumes that the health impacts of cesarean delivery and delivery at an H hospital are ***additive*** and the same size for LRFBs and BPFBs then it is possible to fit a pooled model on the two groups that includes two endogenous variables: H_i (delivery at an H hospital) and C_i (cesarean delivery) and has two instrumental variables: Z_i (relative distance) and the interaction of Z_i with an indicator for breech presentation. The estimated first-stage and reduced-form coefficients for this pooled model are shown in the 3 right-most columns of Panel A, while estimates from three alternative second-stage models are shown in the 3 right-most columns of Panel B: one which assumes that only H_i matters, one that assumes that only C_i matters, and one that allows both effects.[38] For reference we also show the estimated IV coefficients from the two single-channel models fit separately to LRFBs and BPFBs.

The pooled first-stage models essentially reproduce the first stages for the subgroups, while the pooled reduced-form model shows a positive effect of relative distance and a negative effect of the interaction between relative distance and breech presentation. The IV estimates from the specification that includes both channels show a large positive effect of cesarean delivery (0.730) – very similar in magnitude to the estimated effect in our baseline model for LRFBs – coupled with a very small and statistically insignificant effect of delivery at an H hospital (0.019). These results provide some assurance that the effects of proximity to an H hospital work through c-section, rather than through unobserved differences in treatment practices at these hospitals.

e. Additional Checks

We conducted two additional checks to evaluate the possibility that delivery at H hospitals has an independent effect on ED admissions in the year after birth. First, we compared the first-stage and reduced-form effects of relative proximity to an H hospital across various subgroups of infants. Equation (8) implies that the reduced-form difference in average health outcomes for a subgroup defined by a specific set of $X_i's$, say $X_i = X_g$, is proportional to the fraction of H&C compliers in group *g* times the average treatment effect on H&C compliers in the group. If the average treatment effect is ***constant*** across subgroups then we should observe a relationship like:

$$RF(g) = \pi_1(g)\beta_1 \tag{13a}$$

where $RF(g)$ is the estimated reduced-form effect of Z_i^B on the health outcome for subgroup *g,* $\pi_1(g)$ is the estimated first-stage effect of Z_i^B on the probability of c-section for subgroup *g,* and β_1 is the treatment effect of c-section. On the other hand,

[38] In the pooled models, we include a breech dummy and interactions of the HSA dummies with the breech dummy, which capture any unobserved differences in the latent health of BPFB's versus LRFB's across HSA's. Estimated models without these interactions are quite similar, and not much more precise.

if delivery at an H hospital is the mediating channel then we would expect to observe a relationship like:

$$RF(g) = \delta_1(g)\beta_2 \tag{13b}$$

where $\delta_1(g)$ is the estimated first-stage effect of Z_i^B on the probability of delivery at an H hospital for subgroup *g,* and β_2 is the treatment effect of H delivery on the outcome.

To distinguish whether (13a) or (13b) provides a better description of the pattern of reduced-form effects we need a stratification such that $\delta_1(g)/\pi_1(g)$ varies across groups. Since this ratio is just the relative fraction of H&C compliers among all H-compliers in a subgroup, we elected to form groups based on the predicted probability of delivering by c-section at an H hospital (i.e., H_iC_i). Specifically, we used a combination of X_i plus the 16 risk factors we used to test the validity of relative distance as an instrument. As shown in App. Table 5, classifying our sample into 4 quartiles based on predicted probabilities from this model leads to groups with $\pi_1(g)$ ranging from 0.007 to 0.016, $\delta_1(g)$ ranging from 0.075 to 0.118, and the ratio $\delta_1(g)/\pi_1(g)$ ranging from 0.085 to 0.137.

Figure 6 plots the estimates of $RF(g)$ against $\pi_1(g)$ for the four groups. We also show the fit from a simple regression (with an unrestricted constant) of $RF(g)$ on $\pi_1(g)$. The figure reveals three important facts. First, the reduced form impacts on ED admission are highly correlated with the first stage effect on c-section, with a squared correlation of 0.71. Second, the two quantities move proportionally – the estimated constant is essentially 0. Third, the slope of the between-group regression is 0.67, which is very close to our direct IV estimate of 0.697 (see Table 5). In contrast, as shown in App. Figure 1 the relationship between $RF(g)$ and the estimated first-stage effect on delivery at an H hospital is much weaker (R-squared = 0.12) and has a relatively large intercept. We conclude that the variation in the reduced form effect of Z_i^B on ED readmission rates is better explained by differences in first-stage effects on c-section rates than by differences in first stage effects on H-delivery.

As a second check, we developed bounds on the effects of H delivery for infants whose mothers were H-compliers but either c-section always takers (H-compliers/C-AT)'s or c-section never takers (H-compliers/C-NT)'s. The bounds are derived by examining the reduced-form effects of relative distance on ED admission rates for infants who are delivered by c-section or delivered vaginally, regardless of hospital type. Such reduced-form comparisons are confounded by selection bias since the H&C complier group switches from vaginal to cesarean delivery when an H hospital is closer. Since the probability of ED admissions has to range from 0 to 1, however, it is possible to bound the size of the selection-bias component. The derivation of the bounds is presented in App. B; the results are reported in App. Table 6. Unfortunately, the bounds are relatively wide so we are not able to rule out potentially important effects of delivery at an H hospital on the postpartum ED admission rates of the two groups of H-compliers that have the same mode of delivery regardless of where they are born.

f. Effects on Subsequent Fertility and Second-Birth Outcomes

Next, we turn to the impacts of cesarean delivery on subsequent fertility and the health outcomes of later births. For this analysis we link first-time mothers with any second delivery we observe during our sample period. We study the effects on fertility using subsamples of mothers we observe for 2, 3 or 4 years after their first birth. Panel A of Table 7 shows the number of observations in each subsample, the mean c-section rate, and the first-stage coefficients for c-section using the continuous and discrete instruments (both measured using the mother's address ***at the first birth***). Panel B presents the estimated reduced-form effects of the instruments on the probability of a second birth in each of the 3 follow-up windows, as well as OLS and IV estimates of the effects of cesarean delivery at the first birth on this probability.

Looking across the rows of Panel A we see that the mean rate of cesarean delivery at first birth is very stable across our follow-up windows. The estimated first-stage coefficients are also relatively stable, and similar to the corresponding estimates based on the entire sample of births. The entries in the first column of Panel B show that about 13% of LRFBs have a second child within 2 years after their first birth; 27% have a second birth within 3 years and 36% have a second birth within 4 years.

The OLS coefficients in the second column of Panel B suggest that cesarean delivery is associated with a roughly constant 2 percentage point reduction in the probability of a second birth. This represents a 13% effect on fertility within 2 years, a 9% effect within 3 years, and a 5% effect within 4-years. These effects are at the lower end of the range in the existing literature summarized in the meta analyses by O'Neill et al. (2013) and Gurol-Urganci et al. (2013), though they are comparable to OLS estimates obtained by Norberg and Pantano (2016) using survey-based US data.[39]

Observational comparisons of fertility differences may yield biased estimates of the causal effect of c-section if there are unobserved factors that affect both primary c-section and future fertility. Our IV estimates are purged of these factors and give a more mixed picture, with the signs of the effects over a 2-year or 3-year period varying with the choice of instrument. Over the longest possible 4-year window both IV estimates are positive but imprecise. Given the limited sample sizes for the various follow-up windows, our design is underpowered to detect modest sized fertility effects.

An alternative design for studying the causal effect of c-section on fertility is proposed by Halla et al. (2016), who focus on differences in c-section rates for mothers who deliver on weekends and Fridays. We tried their approach in our data and found that the probability of c-section is significantly lower on weekends (by 3.5-3.7 ppt's, mostly attributable to a reduction in scheduled c-sections), and that weekend deliveries of the first birth are associated with slightly higher subsequent fertility rates. (The Friday effect in our data is very small). The implied IV estimates (and standard errors) on the probability of birth after 2, 3, and 4 years are -0.026 (0.044), -0.078 (0.068) and -0.105 (0.105), respectively. These are not significantly different from our distance based estimates but they are systematically negative (and also more precise).

[39] As emphasized by Battacharya et al. (2005) and Norberg and Pantano (2016) any effect of primary c-section on later fertility can represent a combination of physiological and behavioral responses. Both these studies find some evidence that women who deliver a birth by c-section have a higher rate of subsequent contraception use.

Nevertheless, we are reluctant to place much weight on these estimates since the weekend effect is largely driven by scheduled c-sections, and we suspect that these mothers may have other health issues that are correlated with future fertility.

Next, we turn to the health outcomes of the second child. For this analysis, we focus on all ***second*** births to mothers in our LRFB sample – a total of around 93,500 births. Descriptive statistics for this sample are reported in App. C. Unlike our main sample we include breech presentations, preterm births (i.e., <37 weeks gestation), and pregnancies with other risk factors in our second-birth sample.

The first row of Table 8 presents the mean rate of primary c-section and the estimated first-stage coefficients for the probability of c-section at ***first birth*** in our second-birth sample. A slightly lower fraction of mothers in the second birth sample had a c-section at the first birth than in the overall LRFB sample (24.5% vs. 25.6%). The first-stage coefficients of our two distance instruments are also a little different in the 2-birth sample. In particular, the effect of the binary version of the instrument is substantially attenuated (estimated effect =0.006 versus 0.011 in the LRFB sample) and is not statistically different from 0 at conventional levels (t=1.53), indicating that the resulting IV estimates may be unreliable.

The next row presents OLS and IV estimates of the effect of a cesarean delivery at the first birth on the probability of a cesarean delivery at the second birth. The OLS estimate shows that a primary c-section increases the probability of a subsequent c-section by 81 ppt. The IV estimates are even larger (and in fact bigger than 1, though not significantly so) suggesting that proximity to an H hospital at first birth creates a permanent split in delivery mode within the complier population.[40]

[40] We used Abadie's (2003) method with our binary distance instrument to estimate the mean probabilities of c-section at second birth for members of the complier group who delivered by c-section or vaginally at first birth. The estimated means are 1.08 and -0.08, respectively. By construction the difference in these two estimates is the IV estimate of the increase in probability of c-section at second birth for the compliers who delivered by c-section at first birth. We also looked at models for having two cesarean deliveries: the effect of relative distance at the first birth on this outcome is essentially the same as the effect on cesarean delivery at first birth, confirming that essentially all mothers in the complier group who are induced to have a primary c-section have a second c-section.

A potential concern with using relative distance at first birth as an instrument for primary c-section is that this variable is highly correlated with relative distance at the second birth, since many mothers remain in the same zip code. To the extent that delivery mode at the *second* birth is affected by relative distance at that time, there may be an upward bias in the IV estimates. To examine this issue, in the next row we show a specification that includes relative distance measured at *second* birth as an extra control variable, so that identification relies on mothers who move between births.[41] Adding this extra control reduces our relative-distance IV estimate for the effect of c-section at first birth on c-section delivery at second birth from 1.160 (std. error = 0.189) to 0.864 (std. error = 0.302), though the estimated coefficient for the control itself is not significant. For the binary version of the instrument, our standard errors unfortunately make this test uninformative (point estimate for effect of c-section at first birth on c-section at second birth = -0.187, std. error = 2.743).

Another way to evaluate concerns over the effect of distance at the second birth is examine the impact of c-section at first birth on *unscheduled* c-sections. Arguably, the "state dependence" effect of a primary c-section should affect *scheduled* c-sections, whereas proximity to an H hospital at the second birth should mainly affect the probability of an unscheduled c-section. The fourth row of Table 8 presents estimates of the effect of c-section at first birth on the probability of an *unscheduled* c-section at second birth. The OLS effect is small and positive (but highly significant) whereas the IV estimates are small and negative (but insignificantly different from 0), implying that the causal effects we estimate for c-section of the second birth are all driven by effects on scheduled c-section rates.

The next three rows of Table 8 present OLS and IV estimates of the effect of a cesarean delivery at the first birth on the probability of an inpatient or outpatient visit by the mother in the year before the second birth, and on the gestation length and birthweight of the second child. The OLS estimates suggest that mothers with a primary

[41] The between-birth mobility rate (defined as changing ZIP codes between births) in our LRFB sample is 38.8%. This rate is comparable in annualized terms to the 26.4% 1-year mobility rate in the 2007-2011 American Community Survey for women aged 18-35 who report having a child in the past year.

c-section are slightly more likely to visit hospital facilities in the period before their second birth, have a slightly shorter gestation, and have slightly heavier second babies. The IV estimates show a larger negative effect on gestation – equivalent to about a 2-week earlier deliver – and a larger negative effect on birthweight, equal to about 600 grams. These impacts are in accord with the nearly 100% rate of scheduled c-section following a primary c-section, and the fact that pre-scheduled c-sections are typically performed a couple of weeks ahead of the expected delivery date, leading to a reduction in birthweight of 220-240 grams per week of pre-term delivery (the slope of a standard fetal growth curve at 35 weeks gestation).

Next, we examine the incidence of two important risk factors at the second birth: placenta previa (placenta partially covering the cervix) and other placenta-related issues; and hypertension (including pre-eclampsia and eclampsia).[42] As discussed in App. A, a sizeable literature has documented that previous c-section is associated with an elevated risk of placental previa and related problems. There are also a few studies linking previous c-section to eclampsia/pre-eclampsia (e.g., Cho et al, 2015; Mbah et al, 2012). Both groups of risk factors have incidence rates around 2.5% in our second birth sample. An OLS model shows a small but highly significant positive effect of prior c-section on the risk of placenta-related problems. The IV estimate using our continuous distance measure is about 2 times larger, but too imprecise to rule out much larger effects, or a zero effect. The OLS estimate of the effect of primary c-section on risk of hypertension is essentially 0.[43] The corresponding IV estimates are positive, but again are too imprecise to provide much information.

The last three rows of Table 8 examine post-partum hospital and ED visits by the mother and second child. OLS models show an elevated rate of hospital visits by

[42]We measure incidence of placenta previa and related diagnoses using by the presence of at least one ICD9 code of 641 among the primary and secondary diagnoses on the mother's discharge record for the second birth. This includes placenta previa, premature separation of the placenta (placental abruption) and placenta-related hemorrhage. We measure hypertension by the presence of at least one ICD9 code of 642. This includes pre-existing hypertension conditions as well as transient hypertension of pregnancy (eclampsia and pre-eclampsia).

[43] We note that the mothers in our second birth sample did not have eclampsia or pre-eclampsia at their first birth, and had BMI less than the 90th percentile among first birth mothers, so two important risk factors for subsequent hypertension-related complications are removed.

mothers who had a previous c-section, similar in size to the effect in the year after the first birth (0.026 versus 0.031 in Table 5). The reduced-form and IV estimates based on the continuous distance instrument are small in magnitude, but the standard errors are large so we cannot rule out a range of causal effects.

The OLS, reduced form, and IV results for hospital and ED visits of the second child in the year after birth are also similar in magnitude to the corresponding effects on the first child. For example, the IV-based point estimates of the effect of cesarean delivery of the first birth are in the range of 0.7-0.9 extra ED visits for the first birth and 0.4-0.8 extra ED visits for the second. Given that the primary c-section leads to a nearly 100% rate of c-section at the second birth, these findings support the hypothesis that cesarean delivery causes health problems for the infant that result in increased ED admissions in the year after birth.

g. Impacts on Infant Death

In the final step of our analysis we turn to the effects of hospital practices and cesarean delivery on infant death in the year after birth. A concern here is that the risk of death may be related to the ***timing*** of c-section for births in which the fetus was under severe stress.[44] If so, then the tendency of practitioners at H hospitals to reach a decision to perform c-section earlier in the labor process could have an "intensive margin" effect on infants who would ultimately be delivered by c-section regardless of hospital type (H-complier/C-always takers) in addition to any impact on H&C compliers who are only delivered by c-section at an H hospital (the "extensive margin" effect we believe is most relevant for other outcomes). To deal with this possibility we mainly focus on measuring the causal effect of delivery at an H hospital, interpreting the impact in the framework of equation (9a).

[44] Tolcher et al. (2014) present a meta-analysis of the literature on the health impacts of the delay between reaching a decision to perform c-section and delivery. They find an inconclusive link. Nevertheless, standard practice guidelines aim to keep this delay time to under 30 minutes -- see e.g., Dunphy et al. (1991); National Institute for Health and Care Excellence (2011) -- and some studies find that an extended delay is associated with worse outcomes -- e.g., Thomas et al. (2004).

A second issue is that for most LRFBs the risk of death is extremely low. Arguably, any true effect of hospital practices or cesarean delivery should be concentrated on the subset of births with an elevated risk of death. To isolate high-risk deliveries, we developed an index for the risk of death using the full set of control variables X_i and the 16 additional risk factors discussed in Section IIId. We also included one additional risk factor based on the presence of a secondary diagnosis code indicating an irregular heart rate or rhythm during labor.[45] This indicator of fetal stress is uncorrelated with relative distance to an H hospital, but is highly correlated with both the probability of c-section and the risk of death.[46] We used a logistic regression to model the risk of death, then stratified deliveries into predicted risk groups.[47] Ultimately, we settled on a simple two-group classification: a low-risk group with predicted risk scores in the bottom two-thirds of the overall distribution, and a high-risk group with predicted risk scores in the top third of the distribution.

To set the stage for our analysis, Appendix Table 7 shows the characteristics of all LRFBs and those that lead to infant death within a year of birth, and a parallel comparison between all high-risk deliveries and those that lead to infant death. We note that there are only 596 deaths in our entire sample, representing a combined neonatal and post-neonatal death rate of 0.121% or 1.2 per 1,000 births. (By comparison the overall infant mortality rate in California during our sample period was about 5.5 per 1,000 births.) 358 of these deaths occur in the high-risk sample, representing a death rate of 2.2 per 1,000.

A few important characteristics stand out as highly correlated with the risk of death. Infants that die are more likely to be low birth weight: 14.4% of all deaths and 23.2% of the deaths among the high-risk group were <2,500 grams at birth versus only

[45] Specifically, we use any report of ICD9 65971 (abnormality in fetal heart rate or rhythm).

[46] Nelson, Sartwelle and Rouse (2016) argue that the use of electronic fetal heart monitors during labor leads to a high rate of false positive detection of fetal stress, which in turn contributes to the high rate of c-section delivery in the U.S., in part because of fear of litigation in the event of a problem such as cerebral palsy attributed to birth asphyxia.

[47] We form risk groups using a 10-fold sample-splitting technique to preclude over-fitting or "endogenous stratification" (Harvill et al. 2013; Abadie et al. 2017). In particular, we randomly select 10 equal-sized folds; then estimate the model 10 times, leaving out one of the folds; and finally, we predict risk of death in the left-out fold.

2.3% of our overall LRFB sample. They are also more likely to experience abnormal heart rate/rhythm during delivery; more likely to a 5-minute Apgar score below 7; more likely to be transferred to a NICU unit; more likely to have an inpatient hospital stay after birth; and more likely to have been delivered by cesarean. Interestingly, the higher rate of c-section is due to *scheduled* c-sections: infants that die are notably ***less likely*** to have been delivered by an unscheduled c-section.

Panel A of Table 9 shows the OLS, reduced form, and IV results for models of the risk of death for all LRFBs, while panels B and C present parallel analyses for the low- and high-risk subgroups, respectively. For each group we show estimated 1st stage models assuming that the endogenous variable of interest is delivery at an H hospital, along with the corresponding reduced form and IV estimates. We then show a parallel set of models assuming that the endogenous variable is c-section delivery.

Beginning with panel A, a simple OLS model with basic controls shows that delivery at an H hospital has a small, insignificant negative effect on the death rate. The reduced-form models, however, show a relatively large negative effect of proximity to an H hospital on death rates – with a magnitude of -0.38 deaths per 1,000 births for a move 10 miles closer to an H hospital, or -0.27 deaths per 1,000 births for being closer to an H hospital. Both estimates are significant at conventional levels, with a t-statistic of 2.35 for the continuous measure of relative distance (p=0.019), and a t-statistic of 2.14 (p=0.033) for the discrete distance measure. Scaling by the first-stage effect for delivery at an H hospital, the estimated IV coefficients imply a life-saving effect of 2.4-2.6 infant deaths per 1,000 births to H-complying mothers. Scaling by the first stage for c-section – which is ***only*** appropriate if there is no effect of H delivery on H-compliers who have the same delivery mode at H or L hospitals – the effect is 21 to 23 infant deaths prevented per 1,000 c-sections performed on H&C compliers.

The results in Panel B show that there is essentially no effect of proximity to high c-section hospitals on the death rate of lower-risk infants. Thus, the overall effects in Panel A are driven by responses for the higher-risk group. Focusing on this group

(Panel C) we note four salient facts. First, the average c-section rate for these infants is actually slightly ***below*** the rate for infants with lower risk of death (24.9% versus 27.5%). Second, within the high-risk sample there is a significant positive relationship between c-section and death – the OLS coefficient implies that infants delivered by c-section have a 1.2/1,000 higher death rate. Third, mothers of high-risk infants are somewhat more responsive to relative distance than other mothers: the first-stage effects on the probability of H-delivery are about 20% bigger in magnitude than the corresponding effects for mothers of low-risk infants. [48] Fourth, the probability of c-section is also more responsive to relative distance, implying that hospital practice patterns have a larger effect on delivery mode of these mothers.

The estimated reduced form coefficients in Panel C are large in magnitude: a 10 mile drop in the relative distance to an H-hospital leads to a decline in infant mortality of about 1 per 1,000 births (t=2.76), while simply being closer to an H hospital leads to a reduction of 0.8 per 1,000 births (t=2.73). We investigated the robustness of these estimates to controlling for the 17 extra risk factors included in our risk of death model. We find that the reduced-form effects are essentially the same when we include all these factors, or any random subset of extra risk factors.[49] We also investigated the reduced form effects using logit models, which assume a proportional effect of proximity to H hospitals rather than an absolute effect. We find that the *average* reduced-form effects are very similar to the estimated effects reported in Table 9.

A key question is whether these reduced-form effects should be scaled by the first-stage coefficients for delivery at an H hospital or by the first-stage coefficients for cesarean delivery. To explore this we tried the approach in Figure 6 of dividing high-risk deliveries into subgroups with different first stage effects for H-delivery and c-section, but were unable to find any stratification that led to significant non-proportionality in

[48] The bigger responses to distance may be due in part to the characteristics of mothers of high-risk infants, which are similar to the characteristics of the distance and procedure compliers (i.e., lower-education, and insured by Medi-Cal).

[49] Adding all 18 risk factors the reduced-form effect of relative distance on death for the overall sample becomes -0.384 (std. error=0.160) while the reduced-form effect for the high-risk sample becomes -1.113 (std. error=0.390). The IV estimates are similarly highly robust.

the two first stages. Given this uncertainty we focus on the implied IV coefficients that treat H-delivery as the causal channel. Among hospital-complying mothers in the high-risk sample, 13-15% are H&C compliers, 18-19% are H-complier/C always takers, and 65-69% are H-complier/C never takers. The treatment effect per delivery at an H hospital can be interpreted as a weighted average of the treatment effects on these 3 groups, using their relative shares as weights.

The IV estimates in panel C imply that delivery at a high c-section hospital prevents about 6 to 7 deaths per 1,000 births by distance-complying mothers, though the 95% confidence interval for even the more precise estimate based on the continuous measure of relative distance is relatively wide, extending from 1.6 to 10.5 deaths per 1,000. Assuming that the delivery practices at H hospitals have no lifesaving effect on H-complier/C never-takers, the point estimate of the causal effect *per combined marginal and infra-marginal c-section* is around 3 times larger than the effect per delivery at an H hospital. Under this assumption, the effect is 18 prevented deaths per 1,000 marginal and inframarginal c-sections, with a 95% confidence interval from 4 to 32 per 1,000.

VI. Discussion

To summarize our main results, we find:

1. **at birth**: cesarean delivery is associated with shorter labor times, lower rates of birth-related injuries for mothers, higher APGAR scores, lower rates of birth injuries for infants, but a substantial increase in use of mechanical ventilation.
2. **in the year after birth**: each c-section delivery is associated with 70-80 ppt. increase in the probability of an ED visits by the infant, mostly for respiratory ailments, but no higher rate of in-patient hospital visits or ASC use.
3. **longer-term**: cesarean delivery has no clear effect on fertility. Subsequent births after a primary c-section are virtually all delivered by scheduled c-section, with a 2-week shorter gestation and 500-600 gram lower birth weight. These second infants appear to have a higher risk of ED visits in the year after birth.

4. **infant death**: delivery at a hospital with an above-average rate of cesarean deliveries is associated with a significant reduction in neonatal and post-neonatal death. The point estimate of the effect amounts to a saving of roughly 18 deaths per 1,000 marginal and infra-marginal c-sections at high c-section hospitals.

Together these findings paint a nuanced picture of the costs and benefits of "marginal" c-sections that are attributable to distance-based hospital choices by patients and systematic variation in hospital delivery practices. We confirm the widely held presumption in the existing literature that cesarean deliveries have reduced lung function that results in the need for extra care after birth. Moreover, these health problems appear to spill over to later births due to the very high rate of repeat c-section, creating a multiplier effect. On the other hand, we find evidence that delivery practices at high c-section hospitals have a relatively large lifesaving effect. Whether the positive mortality effects are large enough to offset the negative morbidity effects is a question we leave for further work.

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Appendix A: An Overview of the Literature on the Health Effects of Cesarean Delivery

(i) Infant Outcomes

Table A1 summarizes a selection of recent studies on the short and medium-run health effects of cesarean delivery for infants. We review studies on injury or death of the baby; lung function and respiratory problems; asthma; immune system; and breastfeeding. Not included in the table are several other active areas of research that study impacts of cesarean delivery on longer-term outcomes such as the probability of adult obesity (see the recent review by Darmasseelane et al., 2014).

Across the board a general finding is that babies delivered by c-section fare worse: higher neonatal and post-neonatal death; elevated risks of respiratory system problems including asthma; evidence of digestive system disorders, and lower rates of breastfeeding. An unusually detailed prospective study by Villar et al. (2007) of births in eight Latin American countries illustrates the general nature of these findings and the difficulty in interpreting the results as causal. [1] The authors show that neonatal death rates for cephalic fetuses delivered by c-section after trial of labor are substantially higher than rates for those delivered vaginally (0.65% versus 0.38%). Eliminating the roughly 30% of intrapartum c-sections performed after indications of fetal distress, the neonatal death rate of the remaining c-section group falls to 0.51% -- not statistically different from the rate for the vaginal births (but still higher), and indicative of a potentially large endogeneity bias in the overall comparison.

Our reading of the literature is that the most widely documented correlation is between c-section delivery and respiratory problems. Such a pattern has been documented in large-scale

[1] The study is unusual in collecting detailed data on reasons for c-section, gathered immediately after the birth by trained survey staff.

cohort studies in several Nordic countries (e.g., Hansen et al., 2008; Tollanes et al., 2008) and in meta analyses of the literature (e.g., Thavagnanam et al., 2008). As discussed in a recent review by Hyde et al. (2012), there is clinical evidence that babies born by c-section have worse lung function immediately after birth -- possibly attributable to a therapeutic effect of the labor process (including release of hormones and clearance of lung liquid). A number of researchers also hypothesize that there is a transfer of microbes from mother to infant during labor that aid in the development of the immune and digestive systems (e.g., Neu and Rushing, 2012).

(ii) Maternal Outcomes

Table A2 presents a parallel summary of the literature on the health effects of cesarean delivery on mothers. Here the literature is less numerous: our reading is that the major health risks include complications at birth and maternal death; reduction in future fertility; abnormal placentation in subsequent pregnancies; and risk of future stillbirths. Most studies find that mothers who deliver by c-section have higher risk of birth-related complications (such as need of a blood transfusion), higher risk of severe morbidity and mortality in the period after the birth, reduced future fertility, higher risk for placenta previa (placenta near or covering the cervix) and placenta accreta/increta/percreta (abnormal placental attachment). Evidence on future stillbirths is less clear.

As with the literature on infant health effects, most of these studies are based on observational designs, making it difficult or impossible to assert causality, though some of the potential effects are grounded in clinic evidence (see for example the review of studies on abnormal placentation by Clark and Silver, 2011). An interesting exception is the study by Halla et al. (2016) on future fertility, which uses day of the week of the birth as an instrument for c-section. We discuss this design in Section V.f, where we report the same basic pattern as Halla et al. (2016) in our data, but argue that day of the week may not be a valid instrument in our setting because of the greater presence of pre-scheduled c-sections on weekdays.

Table A1: Summary of Literature on Infant Health Effects of C-Section

Health Issue	Study authors; design; main findings
1. Delivery injuries and death	a. Rouse and Owen (1999) prophylactic CS's for large fetuses (>4000g) have small impact on permanent brachial plexus injury
	b. Alexander et al (2006): 1.1% of CS babies have some birth injury - mostly cuts from the incision.
	c. Villar et al (2007): CS might decrease death for cephalic pregnancies, definitely for breech; increased NICU, but rupturing of membranes may be protective
	d. MacDorman et al (2008): CS has 1.7-2.4 higher risk of infant neonatal mortality for primary, low-risk births. Intention to treat analysis combines CS after TOL with vaginal births as intended vaginal.
	e. Molina et al. (2015); cross-national analysis of C-section and infant morality; neonatal mortality rates decline until C-Section rate of 20%, then stable across countries
2. Lung Function and Repiratory Problems	a. Hansen et al. (2008); Danish cohort study (cov. adj.); scheduled C-Section increases risk of respiratory illness 200-400%
	b. Moore et al. (2011); Australian register study (cov. adj.); elective CS increases risk of hospitalization for bronchiolitis by 10% in first year of life
	c. Hyde et al. (2012); review of clinical literature; CS without TOL associated with reduced lung function after birth
	d. Kristensen et al. (2015); Danish register study (cov. adj.); elective CS associated with 20% higher risk of pneumonia and other mucosal system disorders
3. Asthma	e. Salam et al (2006): retrospective study of California youth; CS raises incidence of allergy by 26% (cov. adj.)
	b. Roduit et al. (2008); Dutch cohort study (cov. adj.). CS associated with 20% increase in risk of childhood asthma, higher effect for allergic parents
	c. Thavagnanam et al. (2008); meta analysis of 23 studies of CS and asthma; CS associated with 45% increase in risk at age 8
	d. Tollanes et al. (2008); Norwegian register study (cov. adj.); CS raises risk of asthma by age 18 by 50%
	e. Jachetta (2014); IV study using MSA-level malpractice premiums instrument; CS associated with higher rate of hospitalization for asthma and lung disease
4. Immune System	a. Neu and Rushing (2012); review of clinical literature; CS without TOL affects micobial colonization/immune response
	b. Sevelsted et al. (2015); Danish register study (cov. adj.); CS associated with higher risk of immune deficiency, inflammatory bowel disorders
	c. Stokholm et al. (2016); prospective study of Copenhagen births; CS associated with different gut microbes in first year
5. Breastfeeding	Prior et al (2012); meta-analysis of 48 studies; CS without TOL associated with lower rate of early initiation of breastfeeding; CS after TOL same as vaginal births

Notes: CS = c-section delivery; OR = odds ratio; TOL=trial of labor; cov-adj = covariate adjustment

Table A2: Summary of Literature on Maternal Health Effects of C-Section

Health Outcome	Study authors; design; main findings
1. Complications at birth; mortality	a. Lydon-Rochell et al. (2000); cohort of primiparous women in Washington State; mean effect = 80% higher rate of rehospitalization in 60 days following CS
	b. Deneux-Tharaux et al. (2006); 3.5 times more likely for mom to die in CS
	c. Villar et al (2007); WHO-supported study of Latin American births; incidence of mother injury/death increases in CS
	d. Kuklina et al (2009) - rise in CS explains rise in maternal morbidity at birth
	e. Curtain et al. (2015); US births in 2013; (no cov. adj.); higher rates of tranfusion, ICU admission
	f. Molina et al. (2015); cross-national analysis of CS and maternal morality; mortality rates decline until CS rate of 20%, then stable across countries
2. Fertility	a. Hall et al. (1989); U.K. cohort study (cov. adj); 23% lower fertility
	b. Kjerulff et al. (2013); U.S. cohort study (covariate adustment); 16% lower fertility
	c. Gurol-Urganci et al. (2013); meta analysis of 18 cohort studies; mean effect = 9% reduction in fertility following CS
	d. Halla et al. (2016); IV based on day of delivery; 17% lower fertility
3. Abnormal Placentation (previa, accreta, etc.)	a. Hemminki et al. (2005); Finish register (cov. adj.); 90% higher risk
	b. Getahun et al. (2006); U.S. linked cohorts (cov. adj.); 30-100% higher risks
	c. Gurol-Urganci et al. (2011); U.K. cohort study and meta analysis of 37 studies CS at first birth raises risk of placenta previa in second by 50-60%
	d. Clark and Silver (2011); review of previous studies; increased risks
4. Future Stillbirth	Bahtiyar et al. (2006); large U.S. cross-section study (cov. adjustment); no effect

Note: CS = cesarean delivery; cov. adj = covariate adjustment.

Appendix B: Methods

1. Interpretation of First Stage, Reduced Form and IV Estimates

Consider the case where individuals (indexed by i) belong to mututally exclusive subgroups. Let X_i represent a vector of indicators for membership in each of J subgroups, let y_i represent an outcome of interest, let D_i represent an endogenous treatment indicator, and let Z_i represent an instrumental variable.

Suppose we estimate a pooled first stage model for D_i that includes Z_i and the vector X_i:

$$D_i = \pi_0 + \pi_1 Z_i + \pi_X X_i + v_i.$$

By standard Frisch-Waugh arguments the OLS estimate of π_1 is:

$$\widehat{\pi}_1 = \frac{\sum_i (D_i - \overline{D}_{j(i)})(Z_i - \overline{Z}_{j(i)})}{\sum_i (Z_i - \overline{Z}_{j(i)})^2}$$

where $j(i)$ is $i's$ subgroup, and $\overline{D}_j$ and $\overline{Z}_j$ represent the means of D and Z within subgroup j. Let N represent the combined sample size and N_j the sample size for group j. Then

$$\begin{aligned}
\widehat{\pi}_1 &= \frac{\sum_j \sum_{i\in j} (D_i - \overline{D}_{j(i)})(Z_i - \overline{Z}_{j(i)})}{\sum_j \sum_{i\in j} (Z_i - \overline{Z}_{j(i)})^2} \\
&= \sum_j \left(\frac{N_j}{N}\right) \left(\frac{\frac{1}{N_j}\sum_{i\in j}(Z_i - \overline{Z}_j)^2}{\frac{1}{N}\sum_j \sum_{i\in j}(Z_i - \overline{Z}_{j(i)})^2}\right) \frac{\sum_{i\in j}(D_i - \overline{D}_{j(i)})(Z_i - \overline{Z}_{j(i)})}{\sum_{i\in j}(Z_i - \overline{Z}_{j(i)})^2} \\
&= \sum_j \left(\frac{N_j}{N}\right) \frac{V_{Zj}}{V_Z} \widehat{\pi}_{1j}
\end{aligned}$$

where V_{Zj} is the variance of Z within group j, V_Z is the overall variance of Z and $\widehat{\pi}_{1j}$ is the first stage regression coefficient for group j.

By the same argument if we estimate a pooled reduced form model for y_i that includes Z_i and the vector X_i:

$$y_i = \delta_0 + \delta_1 Z_i + \delta_X X_i + u_i.$$

the OLS estimate of δ_1 is

$$\widehat{\delta}_1 = \sum_j \left(\frac{N_j}{N}\right) \frac{V_{Zj}}{V_Z} \widehat{\delta}_{1j}$$

where $\widehat{\delta}_{1j}$ is the reduced form coefficient for group j. Finally, the pooled IV estimate of the effect of D on y using Z as an instrument and controlling for X is:

$$\widehat{\beta}_1 = \frac{\widehat{\delta}_1}{\widehat{\pi}_1}$$

$$
\begin{aligned}
&= \sum_j \left(\frac{N_j}{N}\right)\left(\frac{V_{Zj}}{V_Z}\right)\left(\frac{\widehat{\pi}_{1j}}{\widehat{\pi}_1}\right)\frac{\widehat{\delta}_{1j}}{\widehat{\pi}_{1j}} \\
&= \sum_j \left(\frac{N_j}{N}\right)\left(\frac{V_{Zj}}{V_Z}\right)\left(\frac{\widehat{\pi}_{1j}}{\widehat{\pi}_1}\right)\widehat{\beta}_{1j}
\end{aligned}
$$

where $\widehat{\beta}_{1j} = \widehat{\delta}_{1j}/\widehat{\pi}_{1j}$ is the IV estimate within subgroup j.

2. Bounds on the Treatment Effects of H-delivery on H-complying C-section Always-takers and Never-takers

For purposes of this section, assume that the instrument Z_i is dichotomous.

a) Effect on H-compliers/C-always takers. Consider a comparison of

$$E[y_i|Z_i = 1, C_i = 1] - E[y_i|Z_i = 0, C_i = 1]$$

From Appendix Table 1, the group of mothers with $Z_i = 0, C_i = 1$ is: $G_1 = \{(H-NT, C-AT), (H-AT, C-AT), (H-comp, C-AT)\}$. The group with $Z_i = 1, C_i = 1$ consists of G_1 **plus** $(H\&C\ comp)$. Now:

$$
\begin{aligned}
E[y_i|Z_i = 1, C_i = 1] &= (1-\theta_1)E[y_i|Z_i = 1, i \in G_1] \\
&\quad +\theta_1 E[y_i|Z_i = 1, i \in (H\&C\ comp)\}]
\end{aligned}
$$

where

$$\theta_1 \equiv \frac{P(H\&C\ comp)}{P(G_1) + P(H\&C\ comp)}$$

is the relative fraction of H&C compliers in the group with $Z_i = 1, C_i = 1$. Therefore

$$
\begin{aligned}
&E[y_i|Z_i = 1, i \in G_1] \\
&= \frac{E[y_i|Z_i = 1, C_i = 1] - \theta_1 E[y_i|Z_i = 1, i \in (H\&C\ comp)]}{1-\theta_1}
\end{aligned}
$$

Moreover, since

$$E[y_i|Z_i = 0, i \in G_1] = E[y_i|Z_i = 0, C_i = 1]$$

if we knew $E[y_i|Z_i = 1, i \in (H\&C\ comp)]$ we could construct an estimator of the difference:

$$
\begin{aligned}
D_1 &\equiv E[y_i|Z_i = 1, i \in G_1] - E[y_i|Z_i = 0, i \in G_1] \\
&= \frac{E[y_i|Z_i = 1, C_i = 1] - \theta_1 E[y_i|Z_i = 1, i \in (H\&C\ comp)]}{1-\theta_1} \\
&\quad -E[y_i|Z_i = 0, C_i = 1]
\end{aligned}
$$

Since only the $(H-comp, C-CAT)$ group switches hospital types when Z_i switches, the ratio D_1/S_1 provides an estimator of the causal effect of switching from an $H=0$ to an $H=1$ hospital on this group, where

$$S_1 = \frac{P(H-comp, C-AT)}{P(G_1)}$$

is the share of the group that switches hospital.

Assuming the outcome y_i is dichotomous, $0 \leq E[y_i|Z_i=1, i \in (H\&Ccomp)] \leq 1$. So an upper bound on the causal effect is

$$U_1 = \frac{1}{S_1}\left(\frac{E[y_i|Z_i=1, C_i=1]}{(1-\theta_1)} - E[y_i|Z_i=0, C_i=1]\right)$$

and a lower bound is

$$L_1 = \frac{1}{S_1}\left(\frac{E[y_i|Z_i=1, C_i=1]-\theta_1}{(1-\theta_1)} - E[y_i|Z_i=0, C_i=1]\right)$$

b) Effect on H-compliers/C-never takers. Consider a comparison of

$$E[y_i|Z_i=1, C_i=0] - E[y_i|Z_i=0, C_i=0]$$

From Appendix Table 1, the group of mothers with $Z_i=1, C_i=0$ is: $G_0 = \{(H-NT, C-NT), (H-AT, C-NT), (H-comp, C-NT)\}$. The group with $Z_i=0, C_i=0$ is G_0 plus the $(H\&C\ comp)$ group. Using the same argument as above, we can decompose

$$\begin{aligned} E[y_i|Z_i=0, C_i=0] &= (1-\theta_0)E[y_i|Z_i=0, i \in G_0] \\ &\quad +\theta_0 E[y_i|Z_i=0, i \in (H\&C\ comp)\}] \end{aligned}$$

where

$$\pi_0 = \frac{P(H\&C\ comp)}{P(G_0)+P(H\&C\ comp)}$$

is the relative fraction of H&C compliers in the group with $Z_i=0, C_i=0$. Thus

$$\begin{aligned} &E[y_i|Z_i=0, i \in G_0] \\ &= \frac{E[y_i|Z_i=0, C_i=0] - \theta_0 E[y_i|Z_i=0, i \in (H\&C\ comp)]}{1-\theta_0} \end{aligned}$$

Moreover, since

$$E[y_i|Z_i=1, i \in G_0 = E[y_i|Z_i=1, C_i=0]$$

if we knew $E[y_i|Z_i=0, i \in (H\&C\ comp)]$ we could construct the difference:

$$\begin{aligned} D_0 &= E[y_i|Z_i=1, i \in G_0] - E[y_i|Z_i=0, i \in G_0] \\ &= E[y_i|Z_i=1, C_i=0] \\ &\quad - \frac{E[y_i|Z_i=0, C_i=0] - \theta_0 E[y_i|Z_i=0, i \in (H\&C\ comp)]}{(1-\theta_0)} \end{aligned}$$

Moreover, since only the $(H-comp, C-NT)$ group swithes hospital types when Z_i switches, the ratio D_0/S_0 provides an estimator of the causal effect of switching from an $H=0$ to an $H=1$ hospital for this group, where

$$S_0 = \frac{P(H-comp, C-NT)}{P(G_0)}$$

As before, we know $0 \leq E[y_i|Z_i=0, i \in (H\&C\,comp)] \leq 1$. So an upper bound on the causal effect is

$$U_0 = \frac{1}{S_0}\left(E[y_i|Z_i=1, C_i=0] - \frac{E[y_i|Z_i=0, C_i=0]-\theta_0}{(1-\theta_0)}\right)$$

and a lower bound is

$$L_0 = \frac{1}{S_0}\left(E[y_i|Z_i=1, C_i=0] - \frac{E[y_i|Z_i=0, C_i=1]}{(1-\theta_0)}\right)$$

Appendix Table 6 presents point estimates of the lower and upper bounds for the two groups of H-complier/C-AT and H-complier/C-NT. We form confidence intervals for these point estimates by a Monte Carlo approach in which we randomly draw new samples, with replacement. To retain the clustered structure of the data we draw samples based on maternal home zip codes.

3. Characterizing Compliers with a Continuous Instrument

Assume the causal model of interest is

$$y_i = \beta_0 + \beta_1 D_i + \beta_2 X_i + \beta_3 W_i + \varepsilon_i$$

where X_i is a particular covariate of interest, W_i is a set of other controls, and D_i is an endogenous treatment dummy. Denote the first stage model by:

$$D_i = \pi_0 + \pi_1 Z_i + \pi_2 X_i + \pi_3 W_i + \upsilon_i$$

Then the coefficient π_1 from the population regression model is:

$$\pi_1 = \frac{Cov[D_i, \widetilde{Z_i}]}{var[\widetilde{Z_i}]}$$

where $\widetilde{Z_i}$ refers to the part of Z_i that remains after projecting on X_i, W_i, and a constant.

Consider the "generalized complier" IV:

$$D_i X_i = \lambda_0 + \lambda_1 D_i + \lambda_2 X_i + \lambda_3 W_i + \upsilon_i$$

which is estimated by IV using the first stage model above for the endogenous variable D_i. The population IV estimator is

$$\lambda_1^{IV} = \frac{cov[D_i X_i, \widetilde{Z_i}]}{cov[D_i, \widetilde{Z_i}]}.$$

We will show that under a monotonicity condition this provides an estimate of an interpretable weighted average of X_i.

The numerator of the expression for λ_1^{IV} is:

$$\begin{aligned} cov[D_i X_i, \widetilde{Z_i}] &= E[D_i X_i \widetilde{Z_i}] \\ &= E[X_i E[D_i \widetilde{Z_i} | X_i]] \end{aligned}$$

Using this expression, the IV estimator can be written as:

$$\begin{aligned} \lambda_1^{IV} &= \frac{E[X_i E[D_i \widetilde{Z_i} | X_i]]}{E[D_i \widetilde{Z_i}]} \\ &= \int_x X \, w(X) f(X) dX \end{aligned}$$

where

$$w(X_i) = \frac{E[D_i \widetilde{Z_i} | X_i]}{E[D_i \widetilde{Z_i}]}$$

and $f(X)$ is the density of X. Notice that the weights $w(X_i)$ sum to 1, since $E[E[D_i \widetilde{Z_i} | X_i]] = E[D_i \widetilde{Z_i}]$. Moreover, provided that $E[D_i \widetilde{Z_i} | X_i] = cov[D_i, \widetilde{Z_i} | X_i]$ has the same sign for all X_i (i.e., that the instrument either raises or lowers the probability of $D_i = 1$ for each value of X) the weights are all positive. Finally, notice that

$$\begin{aligned} w(X_i) &= \frac{cov[D_i, \widetilde{Z_i} | X_i]}{cov[D_i, \widetilde{Z_i}]} \\ &= \frac{cov[D_i, \widetilde{Z}_i | X_i] / var[\widetilde{Z_i} | X_i]}{cov[D_i, \widetilde{Z_i}] / var[\widetilde{Z_i}]} \times \frac{var[\widetilde{Z}_i | X_i]}{var[\widetilde{Z}_i]} \\ &= \frac{\pi_1(X_i)}{\pi_1} \times \frac{var[\widetilde{Z}_i | X_i]}{var[\widetilde{Z}_i]} \end{aligned}$$

where $\pi_1(X_i) \equiv cov[D_i, \widetilde{Z}_i | X_i] / var[\widetilde{Z_i} | X_i]$ is the first stage coefficient of the instrument conditional on X_i. So the weight can be interpreted as a product of the relative first stage effect for observations for a given $x - group$ and the relative variance of the instrument for the $x - group$.

To summarize: Under a monotonicity assumption the "complier IV" λ_1^{IV} estimand is a weighed average of X_i, where the weights reflect the relative size of the first-stage effect of the instrument on the particular $x - group$, multiplied by the relative variance of the instrument within the group.

Appendix C: Data

a. Overview of PDD/ED/AS/Linked Birth Cohort Data

California OSHPD has created a linked file that combines in-patient discharge records for delivering mothers and newborns with Vital Statistics (VS) data (i.e., information collected from birth certificates and death records) and information on in-patient, Emergency Department (ED), and Ambulatory Surgery Center (ASC) records for each mother in the period from one year before to one year after the birth, and for each infant in the period up to one year after the birth. We use a version of this file that has information on live hospital delivered births for the period from 2007 to 2011.

Appendix D of the data base gives the name, address, zip code, and Hospital Service Areas (HSA) for each hospital, ED, and ASC in the state. We also use external information from the Dartmouth Atlas website to assign HSA's and Health Referral Regions (HRR's). We add data from the US Census Bureau on average income in each zip code.

b. Construction of relative distance instruments

The procedure for constructing a mother's relative distance to high and low c-section hospitalsconsists of 3 steps:

1. We estimate each hospital's risk-adjusted c-section rate among low-risk first births;
2. We classify hospitals as low (L) or high (H) c-section hospitals based on their risk-adjusted c-section rates from (1);
3. We calculate each mother's distances to the nearest L and H hospitals, from which we calculate our main relative distance measure.

In step 1 we fit a logistic regression model to our sample of low-risk first births that includes a baseline set of case risk factors $X_i$and indicators for the hospital $h(i)$ at which mother i delivered. Specifically, using our LRFB sample, we estimate the model:

$$P(C_i = 1|X_i) = \Lambda(\alpha + \mathbf{X}_i'\beta + \gamma_{h(i)})$$

where Λ is the logistic CDF.

In step 2 we compare hospital h's estimated logit coefficient $\hat{\gamma}_h$ to the birth-weighted average hospital coefficient in each Hospital Referral Region (HRR) $\bar{\gamma}_{HRR} = \left[\sum_{j\in HRR} N_j\right]^{-1} \sum_{j\in HRR} N_j\hat{\gamma}_j$ (where N_h is the number of low risk first births delivered at hospital h in our analysis sample). We define a hospital to be a "high c-section hospital" (or H hospital) if $\hat{\gamma}_h \geq \bar{\gamma}_{HRR}$ and otherwise a "low c-section hospital."

In step 3 we use information on the centroid of each mother's home zip code and on the centroids of the zip codes for each hospital to define the distance from each mother to each hospital. We then define the distance to the nearest H hospital and the nearest L hospital.

c. Breech Birth Sample

We use the variable "fetpres" from VS records as our indicator of fetal presentation. We define our breech first birth (BPFB) sample using the same restrictions as we impose on our main low risk first birth analysis sample, with the exception that we focus on breech presentation fetuses. Column 2 of the Data Appendix Table shows the characteristics of the resulting sample. For reference column 1 reproduces the third column of Table 1 and shows the characteristics of our main LRFB analysis sample.

d. Second Birth Sample

We select our second birth as follows: using the mother id variable we find any later birth to a mother who is included in our LRFB sample for which the recorded parity is 2 and for which information on delivery mode is available. Column 3 of the Data Appendix Table shows the characteristics of the first births that can be linked to a second birth. Among the second births, 1.7% are breech presentation, 5.9% are delivered at less than 37 weeks gestation, and mean birth weight is 3392 grams.

Data Appendix Table: Characteristics of Main Analysis Sample, Breech Sample, and Second Birth Sample

	LRFB Analysis Sample	Low Risk Breech Births	LRFB with 2nd Birth
Mother's characteristics			
Mean age	25.6	27.4	26.4
At most high school education (%)	41.2	31.7	31.4
Mean weight (pounds)	137	137	137
Race/eth: Hispanic (%)	44.2	35.1	33.5
Asian (%)	17.6	19.6	19.4
Nonhispanic white (%)	31.7	41.6	41.5
Nonhispanic black (%)	5.6	2.9	4.7
Insurance: Medi-Cal (%)	39.8	30.8	26.6
private non-Kaiser (%)	39.9	50.6	53.8
private Kaiser (%)	14.9	14.3	17.3
Birth risk factors and characteristics			
Breech presentation (%)	0.0	100.0	0.0
Mean number prenatal care visits	12.2	12.2	12.5
Mother had ED visit year prior to birth (%)	19.5	17.9	20.7
Mean gestation (weeks)	39.9	39.4	39.9
Mean birthweight (grams)	3348	3278	3364
Delivery outcomes			
C-section delivery (%)	25.6	97.9	24.5
Scheduled c-section (%)	9.2	79.2	8.6
Delivered at H hospital (%)	51.5	51.9	52.8
Postpartum outcomes			
Infant re-admitted to ED (%)	33.8	30.9	31.0
Infant re-admitted as in-patient (%)	8.2	7.3	7.9
Mother readmitted (any type) (%)	14.9	13.8	17.1
Another birth within 4 years (%)	36.4	39.5	100.0
Sample size	491,604	12,749	93,575

Notes: See notes to Table 1. Low risk breech births are births with same low risk criteria as LRFB sample, but with breech presentation. Second birth sample describes charactersistics of first birth for those mothers that are in the LRFB sample and are observed having a second birth.

Figure 1: Pathways to C-Section Delivery

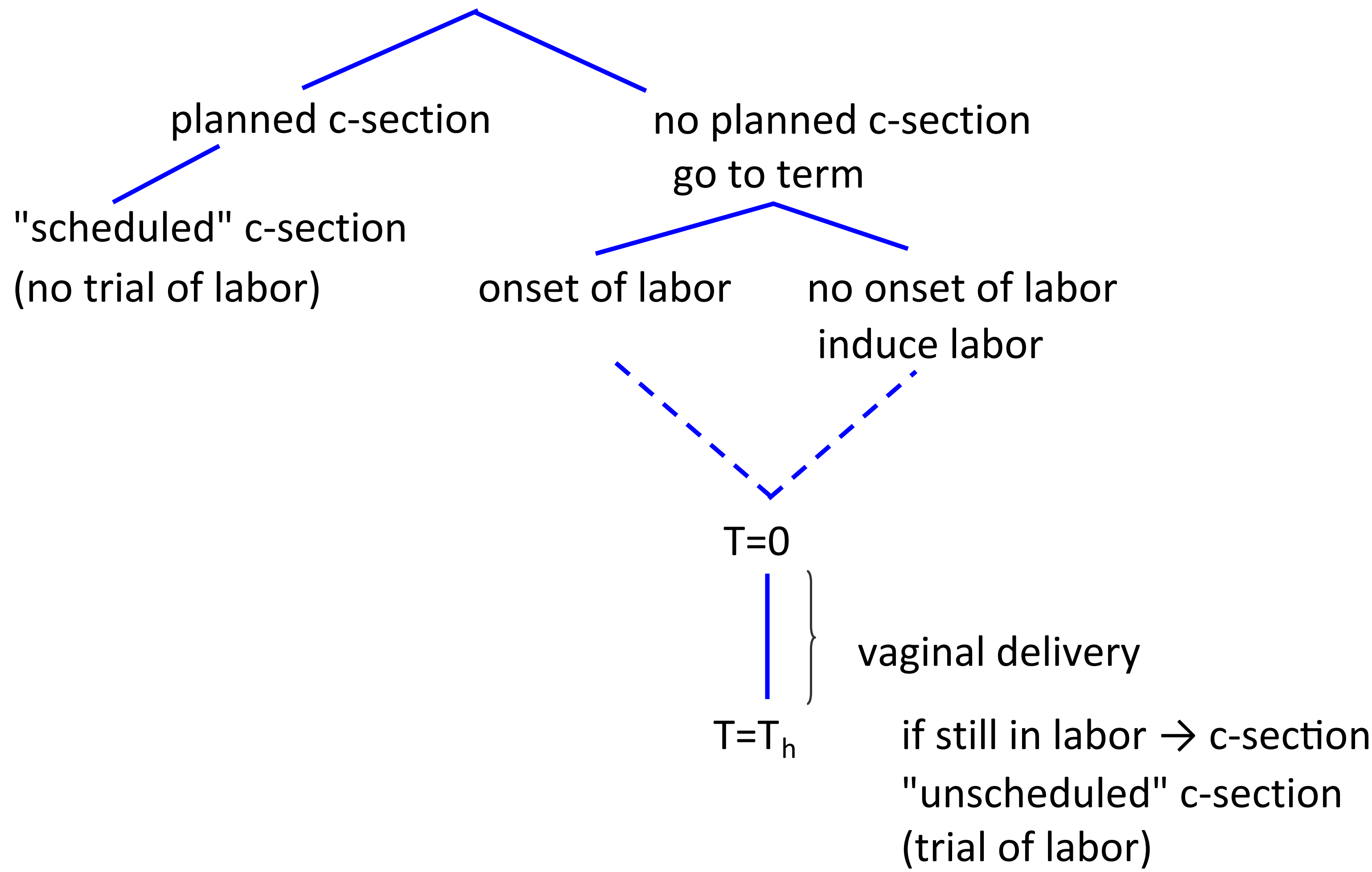
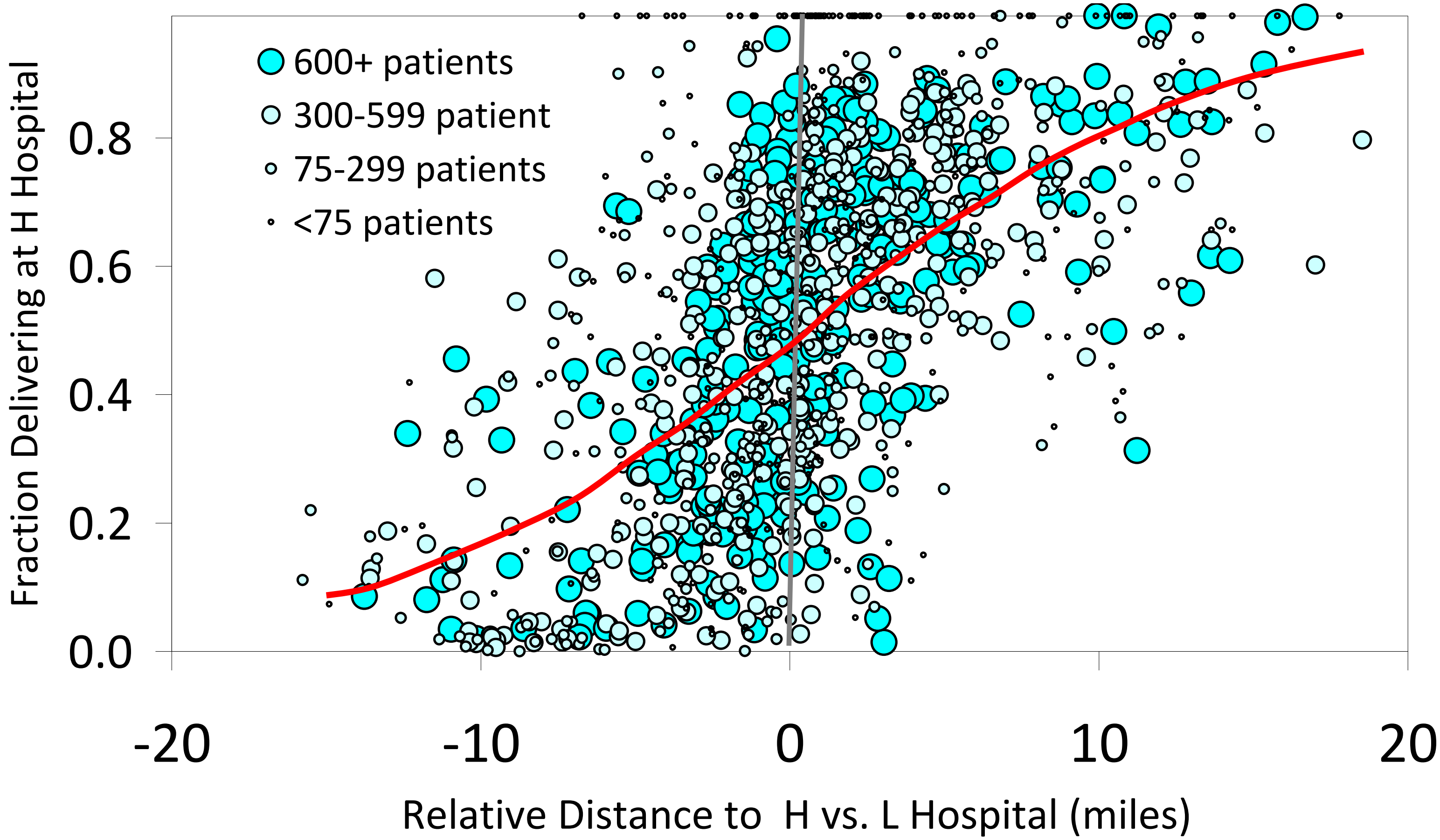


Figure 2: Relative Distance and Probability of Delivery at High C-section (H) Hospital



Note: each point represents a home zip code. Fitted logistic shown in red.

Figure 3: Effect Sizes of Being 10 Miles Closer to a High C-Section Hospital on Predetermined Risk Factors

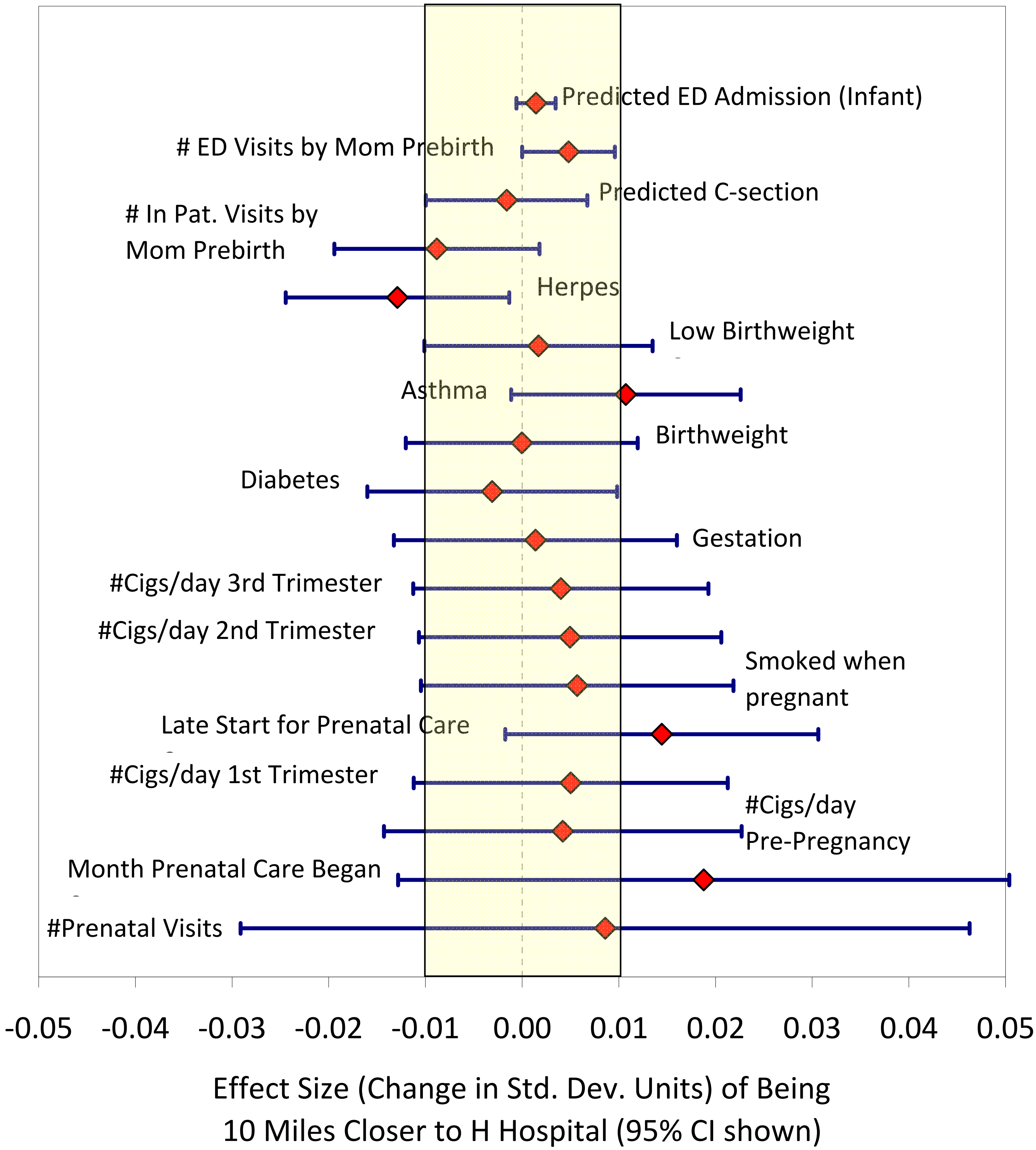


Figure 4: Estimated Reduced Form Effects of Relative Distance on Probability of ED Visit for Infant at Different Followup Horizons

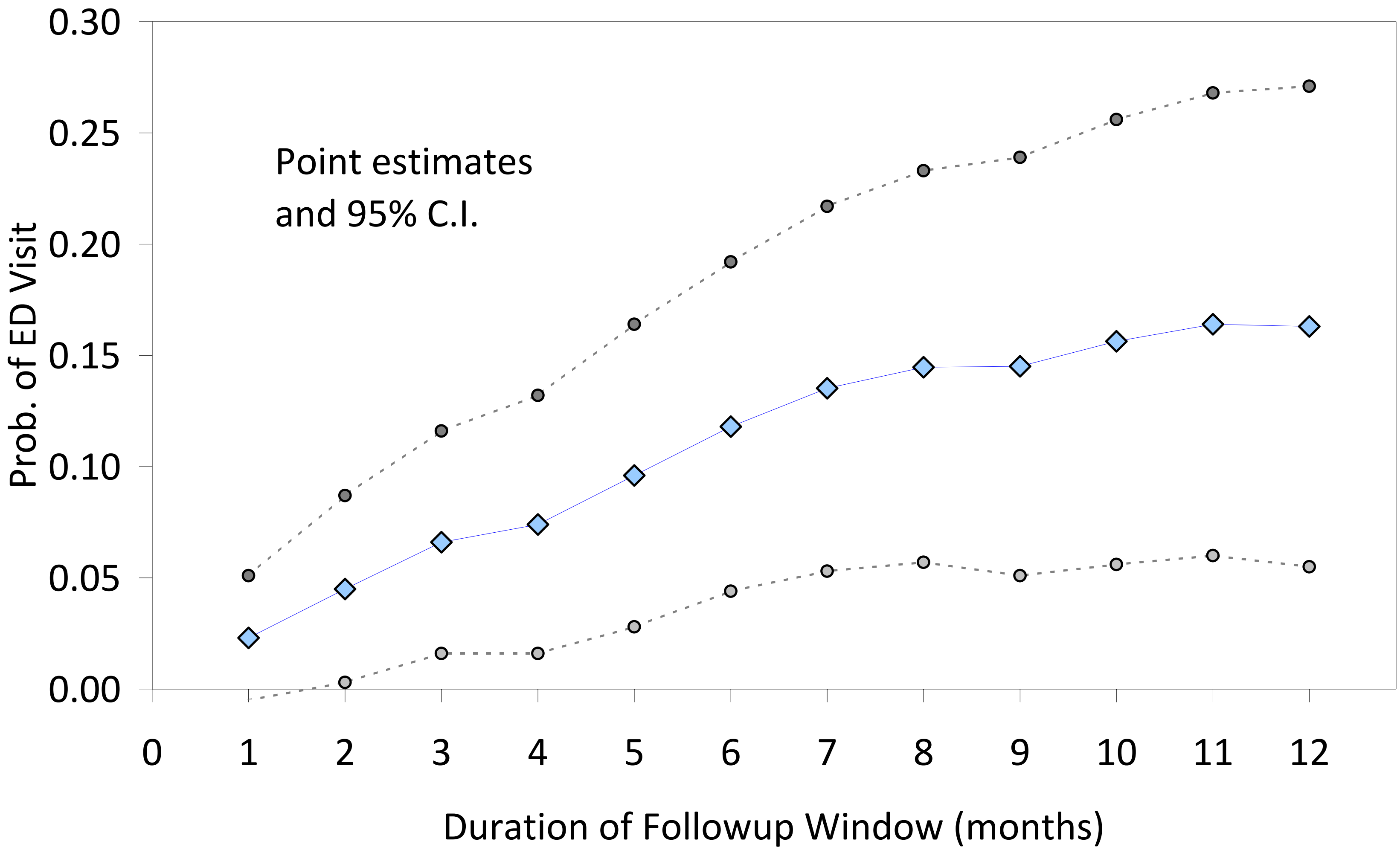


Figure 5: Sensitivity of Estimated Reduced Form Effect of Relative Distance on Probability of ED Visit

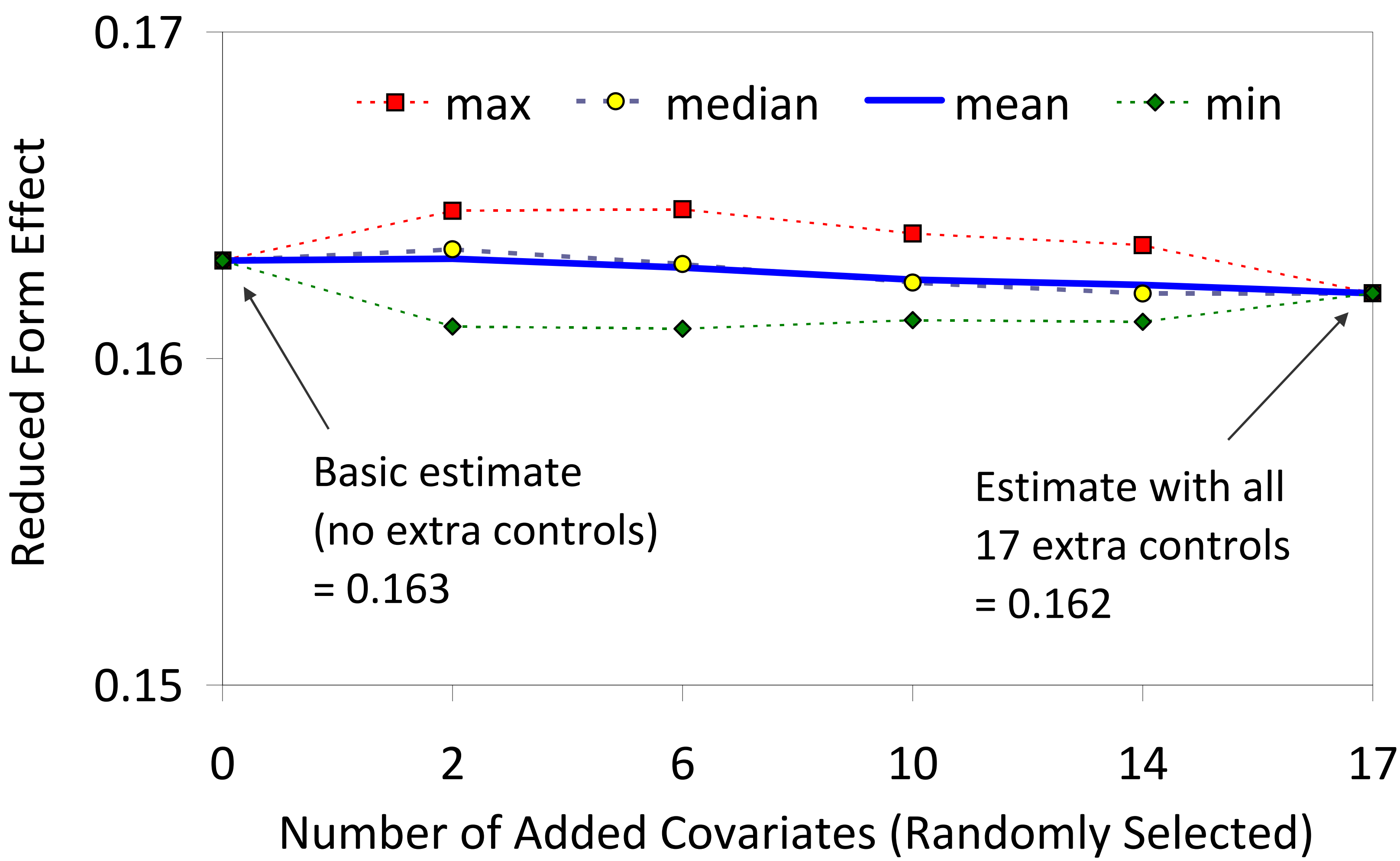
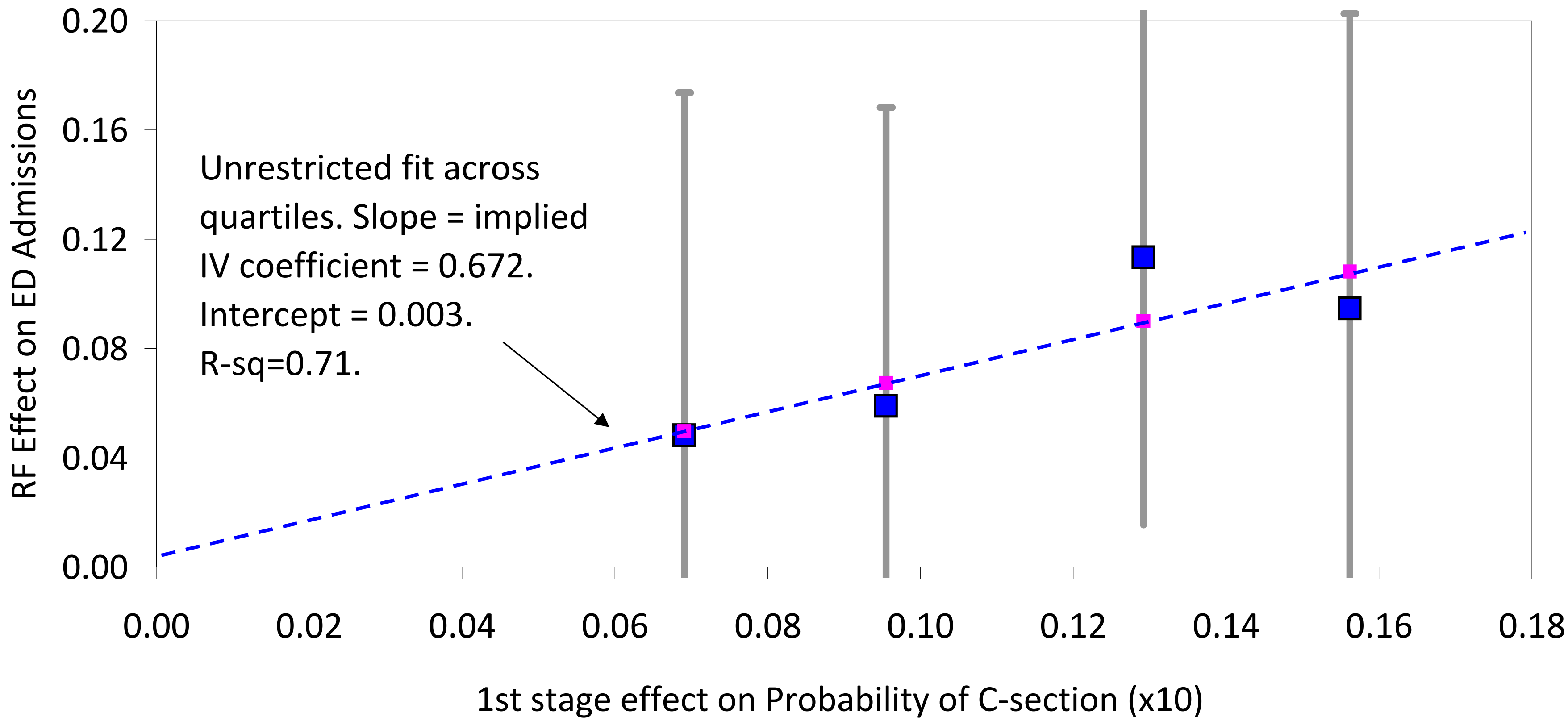


Figure 6: First Stage and Reduced Form Effects of Proximity to H-Hospital on Probability of C-Section and Infant ED Admission



Note: Points = fitted pairs of 1st stage and reduced form coefficients (with 95% CI for RF coefficients) by quartile of predicted probability of c-section at H-hospital.

Table 1: Characteristics of All Births, Low Risk First Births and Analysis Sample

	All Births	Low Risk First Births	Analysis Sample
Mother's characteristics			
Mean age	28.3	25.4	25.6
At most high school education (%)	49.6	41.5	41.2
Mean weight (pounds)	149	137	137
Race/eth: Hispanic (%)	50.9	44.2	44.2
Asian (%)	12.2	15.5	17.6
Nonhispanic white (%)	27.7	32.2	31.7
Nonhispanic black (%)	5.8	5.5	5.6
Insurance: Medi-Cal (%)	46.0	41.1	39.8
private non-Kaiser (%)	34.3	38.0	39.9
private Kaiser (%)	13.1	14.2	14.9
Birth risk factors and characteristics			
Mean Parity	2.1	--	--
Previous c-section (%)	20.7	--	--
Breech presentation (%)	3.4	--	--
Mean number prenatal care visits	12.0	12.1	12.2
Mother had ED visit year prior to birth (%)	20.6	19.9	19.5
Mean gestation (weeks)	39.2	39.9	39.9
Mean birthweight (grams)	3309	3347	3348
Delivery outcomes			
C-section delivery (%)	32.7	25.5	25.6
Scheduled c-section (%)	23.7	9.4	9.2
Delivered at H hospital (%)	51.8	51.6	51.5
Postpartum outcomes			
Infant re-admitted to ED (%)	33.9	34.0	33.8
Infant re-admitted as in-patient (%)	9.3	8.2	8.2
Mother readmitted (any type) (%)	16.8	15.3	14.9
Another birth within 4 years (%)	24.3	35.8	36.4
Sample size	2,699,302	631,506	491,604

Notes: All births include all live in-hospital births in California, 2007-2011. Low risk first births include singleton nonbreech full term (37+ weeks) first births with no indications of eclampsia or pre-eclampsia, mother's BMI < 33.83 (90th percentile), and ≤ 20 prenatal visits. Analysis sample includes mothers with valid home zip code, distance to nearest high or low c-section hospital ≤ 20 miles, and distance to actual hospital of delivery ≤ 20 miles.

Table 2: Estimated First Stage Effects of Relative Distance on Place of Delivery and C-Section

Outcome Variable	Mean	Instrument=Relative Distance 1st Stage Coefficient (x 100)	Instrument=Closer to H Hospital 1st Stage Coefficient (x 10)
Deliver at High C-Section (H) Hospital	0.515	1.586	1.013
		(0.155)	(0.122)
C-section Delivery	0.256	0.182	0.114
		(0.030)	(0.022)
Scheduled C-section (no trial of labor)	0.092	0.051	0.023
		(0.020)	(0.013)
Unscheduled C-section (trial of Labor)	0.163	0.132	0.091
(trial of labor)		(0.026)	(0.018)
Breakdown of C-Section Deliveries:			
C-Section at at High C-Section Hospital	0.149	0.489	0.317
		(0.050)	(0.040)
C-Section at at Low C-Section Hospital	0.106	-0.307	-0.203
		(0.034)	(0.122)
Fractions of Complier Subgroups --Moving Family 10 miles closer to H hospital (col. 2) or closer to H hospital (col. 3)			
P(H Complier)		0.159	0.101
P(C&H Complier)		0.018	0.011
P(H Complier & C Always Taker)		0.031	0.020
P(H Complier & C Never Taker)		0.110	0.070
P(C Complier|H Complier)		0.115	0.113
P(C-section| H Complier, further from H hosp.)		0.193	0.200
P(C-section| H Complier, closer to H hosp.)		0.308	0.313

NOTE: Sample=491,604 low-risk first births. Standard errors in parentheses clustered at 5-digit ZIP code level. All models include controls for Hospital Service Area, year of birth, distance to closest hospital, mother's age (17 dummies), mother's education (8 dummies), race, father present, insurance type, cubic in mother's height, cubic in mother's weight, pre-pregnancy BMI, and mean income in ZIP code. Estimates of fractions of complier subgroups in bottom panel are only interpretable under specific assumptions described in text.

Table 3: Characteristics of Compliers

	Means for Low Risk First Births (1)	Means for Hospital Compliers (2)	Means for Procedure Compliers (3)
A. Socio Economic Characteristics of Mother			
Race/Ethnicity			
White	0.32	0.27	0.19
Black	0.06	0.03	0.07
Asian	0.18	0.11	0.12
Hispanic	0.44	0.59	0.62
Education			
High School or Less	0.41	0.52	0.69
Some College	0.20	0.15	0.17
BA or Higher	0.39	0.33	0.14
Zip code avg. inc. < median	0.50	0.49	0.92
B. Mother's Insurance Coverage at Delivery			
Medi-Cal or other Gov.	0.43	0.61	0.83
Private Kaiser	0.15	0.02	-0.05
Private non-Kaiser	0.38	0.33	0.22
Other	0.04	0.04	0.00
C. Other Maternal/Infant Characteristics			
Mother height < 5 ft.	0.04	0.05	0.09
Mother visit ED prepartum	0.19	0.21	0.32
Number prepartum ED visits	0.26	0.29	0.46
Male baby	0.51	0.53	0.59
Birth weight < median	0.50	0.51	0.49
Low birth weight (<2500 g)	0.02	0.03	0.03
Pred. infant readm. > median	0.50	0.63	0.83
D. Local Air Quality (based on Home Zip Code)			
Ozone > median	0.50	0.64	0.77
PM-25 > median	0.50	0.85	0.84

Notes: column 1 shows estimated means for overall analysis sample of LRFB's. Column 2 shows means for births that are delivered at H hospitals as a result of being relatively closer to such hospitals; column 3 shows means for births that are delivered by c-section as a result of being closer to an H-hospital. Models used to estimate complier characteristics include basic controls plus characteristic itself.

Table 4: Estimated Effects of C-Section Delivery on Outcomes of Infant and Mother at Birth

Outcome Variable	Mean	OLS Coeff. C-Section	Instrument=Relative Distance: RF Coefficient (x 100)	Instrument=Relative Distance: IV Estimate	Instrument=Closer to H Hopsital: RF Coefficient (x 10)	Instrument=Closer to H Hopsital: IV Estimate
C-section (1st stage)	0.256	--	0.182	--	0.114	--
			(0.030)		(0.022)	
Infant Outcomes:						
Apgar (5 minute)	8.915	-0.022	0.088	0.482	0.075	0.654
		(0.002)	(0.030)	(0.183)	(0.020)	(0.210)
Birth Injury (×100)	0.094	-0.019	-0.142	-0.779	-0.293	-2.570
		(0.009)	(0.129)	(0.713)	(0.114)	(1.125)
NICU	0.034	0.021	-0.026	-0.140	-0.032	-0.282
		(0.001)	(0.014)	(0.074)	(0.010)	(0.095)
Ventilation	0.015	0.010	0.045	0.248	0.006	0.051
		(0.001)	(0.021)	(0.111)	(0.010)	(0.088)
Maternal Outcomes:						
Trauma to Perineum and	0.461	-0.608	-0.153	-0.837	-0.161	-1.491
Vulva During Labor		(0.005)	(0.057)	(0.271)	(0.035)	(0.288)
Perineal Laceration	0.290	-0.391	-0.147	-0.809	-0.139	-1.215
(2nd degree or higher)		(0.005)	(0.039)	(0.199)	(0.028)	(0.268)
Length of Stay (days)	2.637	1.384	-0.033	-0.177	-0.046	-0.404
		(0.007)	(0.091)	(0.501)	(0.067)	(0.630)
Length of Labor (days)	0.530	0.084	-0.112	-0.610	-0.087	-0.770
(birth - admission)		(0.003)	(0.044)	(0.252)	(0.030)	(0.314)
Post-birth Stay (days)	2.105	1.294	0.084	0.454	0.041	0.361
(discharge-birth)		(0.006)	(0.073)	(0.384)	(0.058)	(0.484)

NOTE: Sample=491,604 first births, except models for 5 minute Apgar, which includes 487,643 observations, and models for length of stay, length of labor and length of post-birth stay, which have 482,187 observations. Length of labor is measured by number of days from mother's admission to birth, censored at maximum of 3 days. Length of stay is censored at maximum of 5 days. Post birth stay is length of stay minus length of labor. Standard errors in parentheses clustered at 5-digit ZIP code level. All models (OLS and IV) include the set of controls described in note to Table 2.

Table 5: Estimated Effects of C-Section Delivery on Subsequent In-Patient and Out-Patient Visits

			Instrument=Relative Distance		Instrument=Closer to H Hopsital	
Outcome Variable	Mean	OLS Coeff. C-Section	RF Coefficient (x 100)	IV Estimate	RF Coefficient (x 10)	IV Estimate
C-section (1st stage)	0.256	--	0.182 (0.030)	--	0.114 (0.022)	--
Infant Outcomes:						
Any in-patient or out--patient visit	0.385	0.000 (0.002)	0.123 (0.057)	0.673 (0.332)	0.056 (0.042)	0.492 (0.392)
Any ED visit	0.338	0.006 (0.002)	0.163 (0.054)	0.892 (0.335)	0.080 (0.040)	0.698 (0.390)
ED visit for respiratory related conditions	0.126	0.005 (0.001)	0.090 (0.030)	0.496 (0.190)	0.051 (0.023)	0.448 (0.232)
Maternal Outcomes:						
Any in-patient or out--patient visit	0.149	0.031 (0.001)	0.019 (0.026)	0.103 (0.142)	0.003 (0.019)	0.025 (0.166)
Any ED visit	0.129	0.027 (0.001)	0.020 (0.026)	0.108 (0.139)	0.007 (0.018)	0.062 (0.158)

NOTE: Sample=491,604 first births. Standard errors in parentheses clustered at 5-digit ZIP code level. All models (OLS and IV) include the set of controls described in note to Table 2. In-patient and out-patient visits are measured over year following birth.

Table 6: Estimated Models for Breech Births and Pooled Models for Low Risk and Breech Births

A. First Stage and Reduced Form Models

	Low-risk First Births			Breech First Births			Pooled		
	1st stages:		Red. Form:	1st stages:		Red. Form:	1st stages:		Red. Form:
	Deliver at H Hospital	*C-section*	*Visit ED (12 mo.)*	*Deliver at H Hospital*	*C-section*	*Visit ED (12 mo.)*	*Deliver at H Hospital*	*C-section*	*Visit ED (12 mo.)*
Rel. Distance to H Hosp.	1.586 (0.155)	0.182 (0.030)	0.163 (0.054)	1.965 (0.228)	-0.057 (0.039)	-0.030 (0.133)	1.586 (0.155)	0.182 (0.030)	0.163 (0.054)
Rel. Distance × Breech	--	--	--	--	--	--	0.386 (0.170)	-0.259 (0.055)	-0.182 (0.129)

B. Second Stage Models for ED Visit in 12 Months After Birth

	Low-risk First Births		Breech First Births		Pooled		
	Channel = *Deliver at H Hospital*	Channel = *C-Section*	Channel = *Deliver at H Hospital*	Channel = *C-Section*	Channel = *Deliver at H Hospital*	Channel = *C-Section*	Both Channels
Deliver at H Hospital	0.103 (0.035)	--	-0.015 (0.067)	--	0.098 (0.035)	--	0.019 (0.058)
C-Section	--	0.892 (0.335)	--	0.520 (2.316)	--	0.888 (0.329)	0.730 (0.431)

Note: Standard errors clustered by ZIP code in parentheses. Low risk first birth (LRFB) sample has 491,604 observations with mean c-section rate=0.256 and mean rate of ED visit=0.338. Breech first birth sample has 12,749 observations with mean c-section rate=0.979 and mean rate of ED visit=0.309 (see App. C). Pooled models are estimated on combined sample. Models for LRFB and breech first births include same control variables listed in note to Table 2. Pooled models include these controls plus a dummy for breech births and interactions of the breech dummy with Health Services Area dummies.

Table 7: Estimated Effects on C-section Delivery on Probability of Second Birth

	Mean	OLS Coeff. On C-section at 1st birth	Instrument=Relative Distance: RF Coeff. on Rel. Distance (x 100)	Instrument=Relative Distance: IV Estimate	Instrument=Closer to H Hospital: RF Coeff. on Closer to H (x 10)	Instrument=Closer to H Hospital: IV Estimate
A. First Stage Coefficients for C-section at First Birth:						
2 year followup sample	0.257	--	0.183	--	0.106	--
(n=299,203)			(0.036)		(0.026)	
3 year followup sample	0.255	--	0.155	--	0.090	--
(n=200,742)			(0.040)		(0.030)	
4 year followup sample	0.252	--	0.187	--	0.110	--
(n=100,570)			(0.052)		(0.041)	
B. Reduced Form and IV Effects:						
Second Birth within	0.127	-0.017	0.017	0.095	-0.007	-0.069
2 Years		(0.001)	(0.029)	(0.158)	(0.023)	(0.214)
Second Birth within	0.272	-0.025	-0.029	-0.183	0.003	0.037
3 Years		(0.002)	(0.050)	(0.322)	(0.040)	(0.436)
Second Birth within	0.364	-0.020	0.033	0.174	0.036	0.325
4 Years		(0.003)	(0.068)	(0.360)	(0.054)	(0.501)

NOTE: Standard errors in parentheses clustered at 5-digit ZIP code level. All models include controls described in Table 2.

Table 8: Estimated Effects of C-Section Delivery at First Birth on Outcomes at Second Birth

			Instrument=Relative Distance		Instrument=Closer to H Hospital	
Outcome Variable	Mean	OLS Coefficient	RF Coefficient (x 100)	IV Estimate	RF Coefficient (x 10)	IV Estimate
C-section at 1st birth (first stage)	0.245	--	0.188 (0.080)	--	0.060 (0.038)	--
Outcomes for Second Birth:						
Delivered by c-section	0.275	0.811 (0.003)	0.217 (0.052)	1.160 (0.189)	0.062 (0.041)	1.035 (0.420)
Delivered by CS, control for 2nd-birth rel. distance	0.275	0.800 (0.003)	0.110 (0.059)	0.864 (0.302)	-0.003 (0.042)	-0.187 (2.743)
Del. by unscheduled c-section	0.027	0.032 (0.002)	-0.004 (0.017)	-0.023 (0.093)	-0.005 (0.015)	-0.078 (0.256)
In-patient/ASC/ED visits by mother, year **pre-birth**	0.222	0.015 (0.003)	0.045 (0.048)	0.241 (0.255)	0.023 (0.036)	0.388 (0.642)
Length of gestation (days) related conditions	275.6	-2.40 (0.10)	-2.94 (1.44)	-15.68 (7.78)	-0.85 (1.17)	-14.21 (19.34)
Birthweight (grams)	3390	16.9 (3.7)	-111.6 (58.6)	-595.0 (345.8)	-27.6 (42.3)	-460.6 (750.4)
Placenta previa and related problems	0.027	0.004 (0.001)	0.001 (0.014)	0.007 (0.076)	-0.008 (0.011)	-0.142 (0.216)
Hypertension (including eclampsia/pre-eclampsia)	0.024	0.000 (0.001)	0.007 (0.016)	0.035 (0.085)	0.008 (0.013)	0.140 (0.246)
In-patient/ASC/ED visits by **mother**, year after birth	0.167	0.026 (0.003)	0.009 (0.040)	0.045 (0.213)	0.034 (0.033)	0.559 (0.634)
ED visits by **infant**, year after birth	0.285	0.002 (0.003)	0.082 (0.057)	0.438 (0.330)	0.050 (0.045)	0.835 (0.937)
In-patient/ASC/ED visits by **infant**, year after birth	0.335	0.000 (0.004)	0.081 (0.059)	0.432 (0.339)	0.047 (0.047)	0.785 (0.959)

NOTE: Sample=93,575 second births to mothers observed having low-risk first birth (see App. C). Standard errors in parentheses clustered at 5-digit ZIP code level. All models include the controls described in Table 2. Placenta previa and related problems includes ICD9 641.xx. Hypertension includes ICD9 642.xx.

Table 9: Estimated Effects of Hospital Policies on Neonatal and Postneonatal Mortality

			Instrument=Relative Distance		Instrument=Closer to H Hospital	
Outcome Variable	Mean	OLS coeff. of Mediator	RF Coeff. on Rel. Distance (x 100)	IV Estimate	RF Coeff. on Closer to H (x 10)	IV Estimate
A. All Births						
Deliver at H Hospital	0.515	--	1.586	--	1.013	--
(Potential First Stage)			(0.155)		(0.122)	
C-section delivery	0.256	--	0.182	--	0.114	--
(Potential First Stage)			(0.030)		(0.022)	
Death (x100) by age 1	0.121	-0.005	-0.377	-0.238	-0.267	-0.264
(Mediated by H delivery)		(0.013)	(0.161)	(0.100)	(0.125)	(0.130)
Death (x100) by age 1	0.121	0.064	-0.377	-2.067	-0.267	-2.339
(Mediated by c-section)		(0.013)	(0.161)	(0.878)	(0.125)	(1.163)
B. Births with Low Risk of Death (Lowest 67% of Predicted Risk of Death)						
Deliver at H Hospital	0.525	--	1.432	--	0.931	--
(Potential First Stage)			(0.156)		(0.124)	
C-section delivery	0.275	--	0.137	--	0.080	--
(Potential First Stage)			(0.035)		(0.025)	
Death (x100) by age 1	0.068	-0.001	0.014	0.010	0.021	0.022
(Mediated by H delivery)		(0.011)	(0.147)	(0.103)	(0.115)	(0.124)
Death (x100) by age 1	0.068	0.038	0.014	0.102	0.021	0.262
(Mediated by c-section)		(0.012)	(0.147)	(1.079)	(0.115)	(1.455)
C. Births with High Risk of Death (Top 33% of Predicted Risk of Death)						
Deliver at H Hospital	0.506	--	1.801	--	1.130	--
(Potential First Stage)			(0.175)		(0.141)	
C-section delivery	0.249	--	0.243	--	0.169	--
(Potential First Stage)			(0.044)		(0.033)	
Death (x100) by age 1	0.218	-0.012	-1.091	-0.606	-0.843	-0.745
(Mediated by H delivery)		(0.030)	(0.395)	(0.221)	(0.309)	(0.289)
Death (x100) by age 1	0.218	0.119	-1.091	-4.492	-0.843	-4.986
(Mediated by c-section)		(0.031)	(0.395)	(1.645)	(0.309)	(1.974)

NOTE: Sample of all births has 491,604 observations for low risk first births. Sample of births with low risk of death has 327,736 observations. Sample with high risk of death has 163,868 observations. Risk of death is predicted by a logit model using basic controls plus additional risk factors (see text). Predictions made using 10-fold cross-validation of logit model to avoid overfitting. Standard errors in parentheses clustered at 5-digit ZIP code level. All models include controls described in Table [illegible]