Postpartum Hospital Stay and the Outcomes of Mothers and their Newborns

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Abstract

During the 1980s and 1990s, the lengths of postpartum hospital stays declined for both vaginal and cesarean births. Health professionals and policy makers expressed concern that shorter hospital stays might jeopardize the health of both mothers and infants and the federal government and states responded by passing laws requiring insurance carriers provide coverage for longer postpartum stays. One such law was the California Newborn's and Mother's Health Act of 1997, which went into effect on August 26, 1997, and mandated that insurance carriers provide coverage for at least 48-hour hospital stays for normal deliveries and at least 96-hour hospital stays for cesarean deliveries. If the physician, in consultation with the mother, discharges the patient before these time limits, the law requires that insurers provide coverage for a home or office follow-up visit for these women. A similar federal law called the Newborns' and Mothers' Health Protection Act of 1996 went into effect on January1, 1998. In this paper, we use restricted-use data for all births in California over a six-year period to examine the effect of these early discharge laws. Using an interrupted time series design, we demonstrate these laws reduced considerably the fraction of newborns that were discharged early. In two-stage least square models using the laws as instruments for the postpartum length of stay, we find that an early discharge increases the probability of a readmission within 28 days for newborns from complicated vaginal births, and c-sections. Among newborns from uncomplicated vaginal deliveries, there is no evidence that the early discharges improved health for privately insured newborns but the laws did reduce readmission rates for Medicaid patients. These results suggest that the early discharge laws improved outcomes for those most likely to be readmitted but there is little benefit for the healthiest patients.

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I. Introduction

Between 1970 and 1992, the average postpartum length of stay for mothers who delivered vaginally declined by 46 percent, from 3.9 to 2.1 days. Over the same period, the length of stay for those delivering by cesarean section fell from 7.8 to 4.0 days, a drop of 49 percent (Thilo *et al.* 1998; Hyman 1999). As a result of these trends, health professionals and policy makers expressed concern that shorter hospital stays might jeopardize the health of both mothers and newborns. A number of tragic stories about mothers and newborns discharged early who later developed life-threatening but preventable conditions fueled the desire of legislatures to address this issue. Between 1995 and 1998, 42 states passed laws requiring insurance carriers to provide minimum postpartum length of stays and a similar federal law called the Newborns' and Mothers' Health Protection Act of 1996 went into effect on January 1, 1998.

A number of authors have demonstrated that these laws increased average postpartum hospital length of stay, decreased the fraction of mothers and newborns discharged 'early', and increased hospitalization costs.[1] There is however limited evidence about the impact of these laws on the health of the mothers and their newborns, and estimates from these studies provide conflicting results.[2]

In this paper, we use a restricted-use data set of California births over the 1995-2000 periods to examine the impact of both the California and federal early discharge laws. The California Newborns' and Mothers' Health Act of 1997 (NMHA), which went into effect on August 26, 1997, mandated that insurance carriers provide coverage for at least 48 hour hospital stays for normal vaginal deliveries and at least 96 hour hospital stays for cesarean deliveries. If the mother, in consultation with the physician, agreed to be discharged before the state minimum time limits, the law also required that insurers provide coverage for an early home or office follow-up visit for the new mother and her newborn. The federal law is similar to the California statute but it does not require that insurance carriers provide coverage for follow-up visits. Both the California and Federal laws applied to third-party insurance which implicitly

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[1] For example, see Udom and Betley (1998), Dato *et al.* (1996), and Liu *et al.* (2004).

[2] Madden *et al.* (2004) find little evidence that early discharge places infants at risk for readmission. In contrast, Meara *et al.* (2004) finds that the early discharge law in Ohio reduced readmission rates but the estimates were statistically insignificant. Finally Datar and Sood find that the federal law generated large and statistically significant reductions in readmission in California.

exempted Medicaid births from coverage. However, as we demonstrate below, the unique structure of Medicaid managed care in California provided coverage to some Medicaid patients early on an extension of the law in later year further impacted Medicaid patients.

Using an interrupted time series model, we find that the California and federal law generated substantial and abrupt drops in the early discharge rate among newborns for vaginal and cesarean deliveries for births insured by private carriers. The impact on Medicaid births was not as large but substantial nonetheless. We find that the law had no statistically significant impact on the 28-day readmission rates for privately insured newborns from uncomplicated vaginal deliveries, but there were statistically precise drops in readmission rates for newborns with complicated vaginal deliveries and c-sections. There is scattered evidence that the laws reduced 28-day mortality rates for some newborns but the results are inconsistent and in most cases, not statistically significant.

Because early discharge laws exogenously increased the postpartum length of stay, we use their passage as instrumental variables for length of stay in a two-stage least squares (2SLS) model to obtain consistent estimates of the impact of the length of postpartum stays on medical outcomes. In the samples of vaginal deliveries with complications and c-sections, being discharged early produces large increases in the probability of readmission. Among newborns from uncomplicated vaginal deliveries, there is no evidence that an early discharge negatively impacts health for those with private insurance but there are large and statistically precise benefits for Medicaid patients in this subsample.

These results are qualitatively similar to Datar and Sood (2006) who examined the same questions as here using a public use version of the data we discuss below. Their paper examines the first-stage and reduced form estimates of the Federal law on postpartum length of stay in California. Because the public use data only identifies year of birth, the authors are unable to examine the impact of the State law and they do not consider the impact of expanding the state law to all Medicaid patients. Similarly, they do not combine the data first-stage and reduced-form estimates into a 2SLS estimate as we do below. The results are however similar along many dimensions.

II. Declining Postpartum Length of Stay and Passage of Early Discharge Laws

The trend towards shorter postpartum hospital stays outlined above was brought about by a number of factors including a shortage of hospital beds, cost containment efforts by managed care organizations, and an effort to 'de-medicalize childbirth' (Braverman *et al.* 1995, Eaton, 2000). As an increasing number of new mothers were discharged prior to the medically recommended length of stay, the press took notice and increasingly used terms such as "drive through deliveries" or "drive by deliveries" to describe early discharges.[3]

In the middle of the 1990s, the medical profession and a number of state and federal legislators began recognizing the potential problems of shorter postpartum hospital length of stay. Tragic stories of mothers and newborns discharged early[4] who later developed life-threatening but preventable conditions fueled the desire of legislatures to address this issue. For legislatures, mandating a minimum postpartum hospital length of stay seemed to be a reasonable and direct solution at that time. Declercq (1999) notes "early discharge laws involved incremental changes to an existing policy, a simple solution to a problem whose health consequences are unclear…." The first bill regulating early discharge was passed by the Maryland legislature in 1995. By 1998, early discharge laws had been adopted by 42 states.

The Newborns' and Mothers' Health Protection Act of 1996 (NMHPA) was signed by President Clinton on September 26, 1996 and became effective on January 1, 1998. The federal law mandated insurance carriers provide coverage for at least a 48-hour hospital stays for vaginal deliveries and at least a 96-hour hospital stays for cesarean deliveries. A decision to discharge a patient before these time limits could be made by a physician only after consulting with the mother.

The California Newborns' and Mothers' Health Act of 1997 (NMHA), which was passed and went into effect on the same day, August 26, 1997, adopted similar mandatory minimum stays as the

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[3] According to Declercq (1999), "…a June 1991 article in the Philadelphia Inquirer was the first reference to early postpartum discharge and the phrase "drive through deliveries" first appeared in the headline in a February 14, 1994, editorial in the *New York Times*."

[4] The American College of Obstetrics and Gynecology (ACOG) and the American Academy of Pediatrics (AAP) define "early discharge" as a postpartum stay of less than 48 hours for uncomplicated vaginal births and a stay of less than 96 hours for cesarean deliveries (American College of Obstetrics and Gynecology 1992).

NMHPA. The California law did however require that if the physician, in consultation with the mother, discharged the patient before these time limits, insurers must provide coverage for an early home or office follow-up visit for both the newborn and mother. Both the federal and California statutes specifically exempted Medicaid patients from coverage.[5] However, the unique structure of California's Medicaid managed care plans meant that some Medicaid recipients were in fact covered by these statutes.

According to state policy adopted in 1993, enrollment in Medicaid managed care plans would be done at the county level and each county would be part of one of four types of structures:[6] a County Organized Health System (COHS), a two-plan model, a geographic managed care model (GMC), or no managed care. Under COHS, managed care plans are run by the county. In the Two-Plan model, one of the options would be a county plan and the other would be a private insurance plan that has a contractual obligation to insure county residents in their plan. In the GMC option, there is no county-provided managed care option and residents can only choose a privately contracted Medicaid managed care plan. Based on Duggan (2004), the share of California Medicaid recipients enrolled in a managed care plan increased from 10% in 1991 to 51 percent in 1999.

Given the structure of Medicaid managed care in California, recipients enrolled in a privately-run managed care plan were subject to the Federal and state early discharge laws but enrollees in county-run Medicaid managed care plans and fee-for-service Medicaid were explicitly exempted from the statutes. This distinction based on the source of insurance was confirmed in a memo written by the Department of Health Services on January 16, 1998.[7]

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[5] Declercq (1999) suggests that by exempting Medicaid, the laws required minimal public funds and assured quick passage of the statutes.

[6] Much of this text is a summary of the description of the California Medicaid managed care program found in Aizer, Currie and Moretti (forthcoming) and Duggan (2004).

[7] http://www.dhs.ca.gov/mcs/mcmcd/PDF/PolicyLetters1998_2003/Policy%2098-01.pdf.

To provide equity in coverage for all Medicaid recipients, a bill was introduced in the state legislature in January of 1998 that would extend the California statute to all Medicaid patients in the state.[8] The bill was eventually passed on August 26, 1998, signed into law by the governor on September 20, 1998, and went into effect on January 1, 1999.[9]

III. Literature Review

a. The Health and Medical Consequences of Short Postpartum Hospital Stays

Research results vary widely regarding the consequences of an early postpartum discharge for the mother and newborn. Most of the research on this topic correlates medical outcomes with the length of stay, controlling for observed characteristics of the patient and hospital. Many of these studies have demonstrated that shorter hospital stays for newborns are associated with higher probabilities of hospital re-admissions (Lee *et al*., 1995; Liu *et al.*, 1997; Malkin *et al.*, 2000a), increased non-urgent visits to health centers and primary care providers (Madden *et al.*, 2002; Mandl *et al.*, 2000; Kotagal *et al.*, 1999), increased the risk of jaundice (Liu *et al.*, 1997; Grupp-Phelan *et al.*, 1999) and increased neonatal mortality (Malkin *et al.*, 2000). The magnitudes of these effects are sometimes quite large.[10]

In contrast to these findings, Dalby *et al.* (1996), Kotagul and Tsang (1996), Brumfield *et al.* (1996), Cooper *et al.*, (1996), Gagnon *et al.* (1997), Bragg *et al.* (1997), Mandl *et al.* (1997), Kotagal *et al.* (1999), Danielsen *et al.*, (2000), Johnson *et al.* (2002), and Madden *et al.* (2002), all show little or no relationship between postpartum hospital stays and hospital re-admission rates. Beebe *et al.* (1996) found no relationship between postpartum length of stay and neonatal mortality. Finally, Mandl *et al.* (2000) and Madden *et al.* (2002) found no impact of early discharges on emergency room or urgent care visits

[8] http://info.sen.ca.gov/pub/97-98/bill/asm/ab_1351-1400/ab_1397_bill_19980813_amended_sen.html.
[9] http://www.leginfo.ca.gov/pub/97-98/bill/asm/ab_1351-1400/ab_1397_bill_19980921_history.html.
[10] For example, using Washington State linked birth certificate and hospital discharge abstracts covering 310,578 live births from 1991 to 1994, Liu and Davis (1997) used logistic regressions to assess the impact of an early discharge (a discharge less than 30 hours after birth) on the risk of rehospitalization within one month of birth. They found that newborns discharged early had a 28 percent higher 7-day re-admission rate and a 12 percent higher 28-day rate. Using linked birth and death certificates, plus hospital discharge records for 48,000 births from Washington state in 1989 and 1990, Malkin *et al.*, (2000b) found that neonatal mortality rates (death within 28 days) were 265 percent higher for infants discharged early compared to those with longer hospital stays.

while Bossert *et al.* (2001) and Madden *et al.* (2004) found no link between early discharge and treatment for jaundice. Although many of these studies have small, not all of the results are simply Type II errors. Using Ohio Medicaid Claims data linked to vital statistics files for 102,678 full-term births from July 1, 1991 to June 15, 1995, Kotagal *et al.* (1999) found that the fraction of newborns discharged early increased 185 percent over the period (from 21 to almost 60 percent). However, there was no corresponding increase in the re-hospitalization rates in the same period. In a sample of 1.2 million vaginally-delivered newborns in California over the 1992-1995 period, Danielson *et al.* (2000) found no statistically significant difference in 28-day hospital readmission rates for babies released after a one-night stay compared to those with two or more nights stay.

It may be no surprise that the results of the single-equation models vary considerably from study to study. Infants assigned longer postpartum hospital stays are not a random selection of newborns but rather, tend to be children with more complications. Therefore, if we expect longer stays to reduce readmissions and there is positive selection on unobserved variables as well, then the expected negative relationship between length of stay and hospital readmission rates should be biased upwards (closer to zero). Positive selection bias on observed characteristics is easy to establish in our data set. As we outline below, the data for this project includes all hospital discharges for childbirth in California over the 1995-2000 period. Using data from the pre-California law period (1995 and 1996), in a simple OLS model, we regress postpartum length of stay (LOS) on observed characteristics. Likewise, we estimate a logistic model with the same covariates but use as the dependent variable a dummy indicating whether the newborn was readmitted within 28 days. We estimate these models for all vaginal and c-section births covered by private insurance and Medicaid and results for these models are reported in Table 1. We report coefficients from the length of stay regression and marginal effects from the 28-day readmission logit.

The results from these models indicate that for both vaginal and c-section deliveries, there is positive selection, that is, children we expect to have longer hospital stays tend to have observed characteristics that predict higher readmission rates. Focusing on the results for vaginal deliveries, those with private insurance, those admitted to non-profit and for profit hospitals (compared to government-owned hospitals), children whose mothers had fewer complications during pregnancy and delivery, those with one or two previous births and girls, all have lower length of stays and lower odds of being readmitted to the hospital. There is less of a consistent story about the selection bias for some of the demographic variables such as the marriage, race/ethnicity, age, and education. Children with married mothers have shorter stays but the coefficient on marriage in the readmission logit is statistically insignificant. Likewise, the coefficients on white and black children are of opposite signs in the two models.

b. Analyses of Early Discharge Laws

Udom and Betley (1998), Dato *et al.* (1996), and Liu *et al.* (2004) examined the impact of early discharge laws on the postpartum length of stay and costs. All three studies show the laws increased the postpartum length of stay and increased costs, while the final two studies demonstrate the laws increased length of stays for those not impacted by the law such as Medicaid recipients and the self-insured.

Madden *et al.* (2004) examined the effects of two policies affecting length of stay of mothers and newborns in Massachusetts: An HMO protocol adopted in 1994 requiring a one-night hospital stay plus a nurse home visit after a vaginal delivery, and a 1996 Massachusetts early discharge law that was similar in scope to the 1998 federal statute. The authors used data on 20,366 mother-newborn pairs with normal vaginal deliveries between October 1990 and March 1998. They found that the reduced length of stay in this HMO and the increase in stay generated by the state law had little impact on subsequent medical encounters for jaundice or newborn feeding problems.[11]

Meara *et al.* (2004) examined the impact of the Ohio early discharge law on the health of newborns in Medicaid. Unlike the California and federal statutes, the Ohio law covered Medicaid

[11] Madden *et al.* (2002) examined the impacts of the same two interventions as in the previous paragraph, but this article considered the impacts for more vulnerable subgroups such as mothers enrolled in Medicaid and mothers from low-income or low-education census tracts. This work generated results similar to their earlier work.

patients. Using Medicaid claims over the 1991 through 1998 period, the authors establish that the law decreased considerably the fraction of short postpartum stays but they showed a noticeable but not statistically significant drop in hospital readmission rates. The most innovative portion of this study was an examination of the efficacy of early post-discharge office visits. Using the fact that there is variation in the delay of an office follow-up visit generated by the day of the week the birth occurred, the authors find that a follow-up visit that occurred within three days of discharge generated statistically significant reductions in hospital re-admission probabilities.

Despite the large volume of research, there is still no consensus regarding the impact of short hospital stays on mothers and their newborns. A review of the literature by Britton *et al.* (1994) concludes that "heterogeneity and limitations of methodology and study design substantially limit conclusions that may be drawn from published studies (p. 291)." Braverman *et al.* (1995) concludes that "there is no clear evidence for the safety, efficacy, and effectiveness of the hospital and post-hospital practices that were previously standard. The current available literature provides little scientific evidence to guide discharge planning for most apparently well newborns and their mothers (p. 724)". A third review by Grullon and Grimes (1997) concludes that "The safety of early discharge is unclear (p. 860)" and "the current data do not support or condemn widespread use of early postpartum discharge in the general population (p. 860)."

There are three persistent problems noted by the authors of the literature reviews discussed above. First, samples sizes are often of an insufficient size to detect effects on outcomes that are rare in the population. Second, many of these studies lacked detailed control variables, especially measures of pregnancy complications. And third, few studies had experimental variation in the covariate of interest (postpartum length of stay). Our work addresses all three of these shortcomings.

Although there have been numerous studies on the impacts of short postpartum hospital stays, much of the literature has one or more of the shortcomings listed above. Studies such as Marbella *et al.* (1998) have large samples but limited controls and no quasi-experimental variation. Studies such as Malkin *et al.* (2000), Danielson *et al.* (2000), Kotagal *et al.* (1999), and Liu and Davis (1997), have large samples, excellent control variables, but no quasi experimental variation postpartum length of stay. Studies such as Meara *at al.* (2004), Madden *et al.* (2002) or Madden *et al.* (2004) have quasi-experimental variation and excellent control variables but samples that may be too small to make meaningful predictions about rare events like hospital re-admissions and neonatal mortality.

Our study is conceptionally similar to Datar and Sood (2006) who use public-use versions of the data we use in this project to examine the impact of the Federal early discharge law. Using data on all births in California over the 1995-2000 period, the authors find that the Federal law reduced the odds of a 28-day readmission rates among newborns by a 9.3 percent in the first year and by nearly 20 percent in the third year. Both of these results are statistically significant at conventional levels. However, as we demonstrate below, these results dramatically overstate the effectiveness of these laws. Because Datar and Sood only have public use data, they are unable to effectively control for month to month variation in readmission rates. As we show below, moving to a restricted-use sample generates substantially smaller and statistically insignificant results.

Our study deals with the three major concerns listed above. First, the law change occurred in California, a state with a large population, and we utilize data for approximately 3 million births in total, with more than half of these births occurring in the treatment period. Our study and that of Datar and Sood are therefore be the largest sample ever used to analyze the impact of postpartum length of stay on health outcomes. Second, although random assignment clinical trials are the gold-standard for inferring causal relationships, the number of observations necessary to eliminate Type II error concerns rate for many low incidence outcomes makes clinical trails impractical for some questions. The best that one can hope to obtain is quasi-experimental variation in field data that mimics a clinical trial. As we demonstrate below, the California law change generated a large and immediate decline in the fraction of newborns and mothers discharged early provides just such variance. Third, in contrast to most previous studies, we examine the impact of early discharge for infants whose mothers experienced complications either during pregnancy or labor. These results are instructive in that early discharge appears to have little clinical risks for newborns that *a priori* have a low risk of readmission.

IV. Constructing the Analytic File

a. Data

The data set for this analysis is a specially linked administrative record data sets of all mothers and newborns discharged from non-federal hospitals in California from January 1, 1995 to December 31, 2000. The data set is generated and maintained by the State of California Office of Statewide Health Planning and Development (OSHPD) and created by linking patient discharge data sets with birth, death and fetal death certificate information.

Public-use versions of the patient discharge dataset contain demographic information such as the age, race, and sex of the mothers and newborns, information about the admission such as the length of stay, procedures used, diagnoses codes, the hospital charges, the type of insurance, and whether the patient died in the hospital. These discharge data sets also contain a code that identifies the hospital.

The linked patient discharge dataset with vital statistics birth file is a restricted-use version of the discharge data that contains all the information in the public use discharge record, plus the exact date and time of birth (and therefore the newborn's admission to the hospital), the zip code of residence, a scrambled Social Security number, information from the birth file that identifies when and where a baby was born and information from the death file that identifies when and where a newborn died for up to one year after discharged. The scrambled Social Security number can be used to link the discharge record over time so as to construct re-admission rates for both mothers and newborns. We have the ability to measure re-admission rates for up to one year after discharge. The information from the birth and death files will allow us to identify whether a newborn died within a fixed time period after admission and discharge, not just whether they died in the hospital. Also, the developers of the data file have matched mothers to newborns allowing us to use characteristics of both mother and the newborn as covariates in multivariate regressions. During the six years in our data set, there are approximately 3 million births in total, with almost 1.7 million births occurring after the passage of the California law.

Although the early discharge laws impact the length of stay for both mothers and their newborns, in this analysis, we focus primarily on the outcomes of infants. We do this because adverse outcomes like readmission and mortality rates are higher for infants than mothers and as we demonstrate below, our samples have just enough power to produce statistical significance in two-stage models for these adverse events.

b. Outcome Variables

There are several outcome variables that we can utilize in the linked Hospital Discharge Data/Vital Statistics birth files that directly or indirectly measure the health of newborns. The most obvious outcomes is whether the newborn was discharged early from the hospital. The federal and state laws provide coverage for a 48-hour hospital stay following a vaginal birth and a 96-hour stay following a c-section. The hospital discharge data set however measures length of hospital stay in days and as a result, we cannot calculate the length of stay in hours. The intent of the law was to provide mothers with the ability to stay an extra night in the hospital if they desired. Subsequently, the key outcome in our analysis is whether the infant was *Discharged early* which is a dummy variable that equals 1 if the newborn spent less than two days in the hospital if the mother delivered vaginally. For newborns whose mother delivered by c-section, this outcome equals 1 if the newborn spent less than four nights in the hospital after a c-sections.

As we noted in the literature review, one concern with early discharges is that health care providers may not have had sufficient time to detect certain conditions. We will exploit the linked nature of our data and construct a measure of whether the newborns were re-admitted to the hospital within a specified time period. In the Hospital Discharge Data/Vital Statistics birth files, the scrambled Social Security number can be used to link the discharge records of newborns over time. Researchers typically measure re-admissions within 7, 14 and 28-days of birth, and we will follow this convention. Initially however, we will focus on the 28-day readmission rates. We will measure the re-admission rates with a dummy variable that equals 1 if a person is re-admitted within a particular number of days.

We will also use neonatal mortality rates as an outcome. In the linked data set, death records for the newborn have been linked to the discharge and birth record. For each newborn, we know whether they died within one year of birth, plus the cause and place of death. Following previous literature, we will use 28-day mortality rates for newborns.

The final outcome measure we will examine is total charges for both the mother and infant's hospitalization.[12] This variable is closely connected with postpartum hospital length of stay of mothers and newborns and the results will allow us to conduct a cost-benefit analysis of the early discharge law. In regression models, given the skewness of hospital charges, we will examine models where the dependent variable is measured in natural logs.[13]

c. Analysis Samples

Many previous analyses of the impact of early discharge on the health of newborns have restricted their attention to uncomplicated vaginal and c-section deliveries. The author surmise that few complicated deliveries will be discharged early so they focus on the deliveries most likely impacted by the early discharge laws. However, as we demonstrate below, we found that these laws significantly impacted the length of stay for more complicated deliveries as well. Therefore, we work with three distinct subsamples: complicated and uncomplicated vaginal deliveries and c-section deliveries. There are a variety of ways to define complicated deliveries. One popular way is to use a specific DRG code for uncomplicated deliveries and there are codes for both mothers and newborns.[14] In our data set, we also have data from the birth record that can be used to define a complicated pregnancy/delivery as any one where the mother presented any one of 24 complications during pregnancy[15] or 23 complications during

[12] Total charges were converted into real 2000 dollars using the Consumer Price Index.
[13] Because some hospital stays are incredibly long and expensive, we will also experiment with using median regressions to eliminate the influence of outliers in the data set.
[14] DRG 370 represents cesarean deliveries with complications, and DRG 372 represents vaginal deliveries with complications.
[15] These include such factors as pre-eclampsia, chronic hypertension, renal disease, Rh sensitivity, premature labor, sexually transmitted disease, Hepatitis B, low or high birth weight, less than 37 weeks gestation, plus others.

labor.[16] Although many patients whose DRG codes indicating a complicated delivery also identified complications on the birth record, the overlap was not perfect. Subsequently, we define a birth as uncomplicated if neither the DRG code of the mother nor the birth record identifies a complication.

We use ICD-9 procedure codes and data on the birth record to identify whether the mother delivered vaginally or by c-section. We restrict our attention to births covered by Medicaid or private insurance carriers, which captures about 95 percent of all births in the state over the period of analysis. Table 2 reports basic demographic information of mothers in our three samples, vaginal without and with complication and and c-section deliveries. Within these subsamples, we report sample characteristics for privately insured and Medicaid deliveries and numbers in parentheses are standard deviations. The inpatient costs (newborn plus mother) measured in real 2000 dollars of complicated vaginal deliveries generate about 60 percent more costs than an uncomplicated birth and c-sections are 70 percent more expensive than complicated vaginal deliveries. There is little difference within each subsample in the costs of Medicaid and non-Medicaid births. On average, based on our samples, women who had a cesarean delivery were older, more educated, and more likely to be black than women who had a vaginal delivery. Across all three samples, Medicaid mothers are more likely to be younger, less educated, racial and ethnic minorities, and more likely to have a previous birth.

V. Graphical Analysis of California Law and Federal Law

a. The Change in Postpartum Length of Stay

In Figure1, we plot the monthly fraction of newborns delivered vaginally without complications that were released early in each month. For reasons that become apparent later, we provide data from July 1, 1995 through December 31, 2000. The vertical line in September of 1997 indicates the first full month the California law was in effect, the line at January 1998 indicates when the federal law became effective, and the third line represents when the state law was expanded to include all Medicaid recipients

[16] These include such factors as seizure during labor, premature rupture of membrane, breech presentation, excessive bleeding, sepsis, cord prolapse, fetal distress, anesthetic complications, maternal blood transfusion, plus others.

Note that in Figure 1, there was an abrupt change in the private insurance time series during a short period between September 1997 and January of 1998. Before September of 1997, the fraction of newborns with private insurance that had a length of stay less than 2 days was relatively stable with a small drift downward. In August of 1997, 82 percent of newborns whose deliveries were paid for by private insurance had a postpartum length of stay less than 2 days. By February of the next year, just six months later, this number had fallen to 50 percent, a 32 percentage point decline and a 39 percent reduction. Early on in our sample, most insurance carriers knew the federal law would take effect in a 1998, but from Figures 1, it appears that few were adjusting 6 months prior to the law change. Therefore, the state law change caught insurance carriers by surprise and as a result, it took some time to adjust. Most carriers had a long lead time to prepare for the federal law and it appears from Figure 1 (and subsequent figures), that the adjustments to a longer length of stay occurred by the 1st quarter of 1998.

Notice also in Figure 1 that although the California law only applied to a Medicaid recipients in particular managed care plans, there is also a noticeable but less dramatic drop in the fraction of newborns from uncomplicated vaginal births discharged early during the 6 months after passage of the California early discharge law. There is however another sharp decline in early discharges for this sample starting in January 1999 when the California law was extended to all Medicaid births. In this simple time series, the drop in early discharges from August of 1997 to February 1998 was 16 percentage points and the early discharge rate fell another 6 percentage points over the next 12 months.

Figure 2 plots the early discharge rate for newborns after complicated vaginal deliveries. In this figure, we see that before the laws, there is a more pronounced downward trend in early discharges among those with private insurance than for uncomplicated vaginal deliveries, with rates falling from 71 to 58 percent over the July 1995 through August 1997 period. Between August 1997 and February 1998 however, early discharge rates among the privately insured fell by an additional 27 percentage points. Over the same period, the early discharge rate among Medicaid recipients in this subsample fell by almost 16 percentage points. Although the rates of early discharge among uncomplicated and complicated vaginal deliveries were very different before the California law, the absolute change in rates was similar for both groups.

In Figure 3 we graph the early discharge rate for newborns who were delivered via cesarean section. In this figure, an early discharge is defined as a postpartum length of stay that was less than four days. Notice that starting in September of 1997 for those covered by private insurance, there was a noticeable drop in the fraction of early discharges. In August of 1997, 90 percent of newborns delivered by cesarean section were discharged in less than 4 days. By the middle of 1999, this number had fallen anywhere from 14 to 16 percentage points. Notice again the sharp decline in early discharge rates in the Medicaid subsample starting in January of 1999 when the state law was expanded to all Medicaid recipients. In contrast to the results for the two vaginal delivery samples, the numbers in this graph show continued declines in the early discharge rates after the effective date for the federal law.

These figures highlight a number of important facts. First, if a short postpartum stay does affect some outcomes such as hospital re-admissions, then the sharp and dramatic drop in short stays generated by the California law should provide an excellent opportunity to precisely estimate the magnitude of this effect. Second, the extension of the state law to all Medicaid births appears to been effective. Third, the timing of the change in time trends corresponded exactly with the effective dates of the state and federal law. Fourth, among vaginal deliveries, the federal law reduced early discharge rates for complicated deliveries by the same rate as for uncomplicated deliveries, even though the former group has substantially lower early discharge rates prior to passage of the federal law.

b. Changes in Health Outcomes

In the Figures 4 through 6, we plot the 28-day readmission rate of newborns covered by private insurance for uncomplicated vaginal deliveries, complicated vaginal deliveries, and c-section deliveries, respectively. In each of these figures, the month-to-month variation in 28-day readmission rate dwarfs any systematic change in readmission rates produced by the law, so it is difficult in these graphs to tell whether the law improved birth outcomes. Therefore, also plot as dotted line the pre-state law and post-federal law means of the outcomes.

Notice that in Figure 4 for uncomplicated vaginal deliveries, there is virtually no time series trend in the readmission rate and there is a small drop in mean rates after January, 1998. The difference in means between the two periods is only 0.07 percentage points, indicating the federal early discharge law had little if any impact on the health of these infants. In contrast, we find more noticeable drops in newborn readmission after the enactment of the federal law in the complicated vaginal and c-section subsamples. In Figure 5, notice a slight downward trend in readmissions throughout the period and in Figure 6, there is a slight upward trend. The vertical distance between the pre/post mean lines in Figure 5 is roughly 4 tenths of a percentage point (almost six times the impact as in the uncomplicated vaginal birth samples) and the same value for Figure 6 is 8 tenths of a percentage point, a number more than 10 times as large as the same value in Figure 4.

In Figures 7-9, we report the times series for the 28-day readmission rates among Medicaid newborns for vaginal deliveries without complications, vaginal with complications and c-section births, respectively. In contrast to Figure 4, we see a more pronounced drop in newborn readmission rates in Figure 7 for uncomplicated vaginal deliveries paid for by Medicaid. Notice that in all three of these graphs, there is a slight drop in the readmission rate after the law was extended to all Medicaid patients in January 1999.

Figures 4-9 point out that it will be difficult to determine whether the state law improved birth outcomes during the four months it was in effect before the federal law became effective. Notice in Figures 4 and 7 there is a noticeable increase in readmission rates in December of 1997, the fourth full month that the California law was in effect and the last month before the federal law took effect. One may be tempted to attribute this sharp increase in readmissions to the California law. But a closer inspection of the data suggests that something else was occurring. Notice that readmission rates always spike in the late autumn and early winter months during the flu season. The winter of 1997/98 was a particularly heavy flu season in California as was pointed out in a report released by OSHPD[17] and flu complications are a common reason for readmission among infants. This cyclic nature of newborn readmissions suggests that we must control for day-to-day environmental conditions that may alter readmission rates. For infants born on a particular day, we calculate the 28-day admission rate (defined as admissions per number of children) for those born 90 to 180 days ago. Our use of children 90-180 days old does however mean that we lose the first six months of data in our analysis,[18] which is why our analysis samples begin on July 1, 1995.

In Figure 10, we average the 28-day readmission rate of newborns and older infants up to the month level and plot the series over time for vaginal deliveries under private insurance. These series demonstrate that the admission rates of slightly older infants is highly correlated with the 28 day readmission rate of newborns and both indexes demonstrates a sharp rise in admissions in December of 1997. We will use the admission index for older infants as a control variable in our model to capture the day to day variation in unobserved conditions that lead to changes in readmission rates.

In Table 3, we provide numeric estimates that coincide with the graphical presentation in Figures 1-9. These estimates are simple difference estimates that compare the post-federal law period (January 1, 1998 and after) with the pre-state law period (July 1, 1995-August 31, 1997). The numbers in parentheses are standard errors.

VI. Econometric Model

a. A Reduced-Form Model

To estimate the impact of the law on birth outcomes, we would ideally use a difference-in-difference model, where we could compare the average length of stay of mothers and newborns before and after the California early discharge law went into effect, using a comparison group to identify what

[17] The title of this report is *CALIFORNIA HEALTH CARE SYSTEM: OVERVIEW OF THE HOSPITAL/EMS CRISIS WINTER OF 1997-98*

[18] Use of the index as a covariate assumes that changes in early discharge rates produced by the state and federal law will have no impact on admission rates after 90 days of age. Results in Table 5 provide some evidence in support of this hypothesis.

the time path of outcomes would have been in the absence of the intervention. Unfortunately for our research purposes, all states were treated at the same time by the Federal law. Within California, the largest potential control group, Medicaid patients, were indirectly impacted by the law and eventually included in the coverage. Datar and Sood (2006) also found that the federal law increased the length of postpartum stays for the self-insured as well. Given the lack of any control group, we therefore use an interrupted time series model instead. Interrupted time series models are sometimes difficult to implement for a variety of reasons. It is often not clear when an intervention became effective. Second, other events may contaminate the treatment effect. For this particular study, we are fortunate that the timing of the law changes are exact, and the immediate and large impact of the law makes it difficult to argue that some other event explains the sudden and precipitous change in hospital length of stays for newborns.

The unit of analysis for this study will be a mother/newborn pair, and the key patient of interest is the newborn because prior studies have suggested their health is more likely to be affected by an early discharge. In the newborn models, we label the outcome variable of interest Y. Outcomes vary across patients, hospitals and time, which are indexed by i, k and t, respectively. The basic interrupted time series model is of the form

$$
\begin{aligned}
(1)\quad Y_{ikt} = {} & X_{ikt}\beta + PAYER(j)_{ikt}{*}MTREND_t{*}\delta(j) \\
& + STATELAW_t{*}PAYER(j)_{ijt}\,\alpha(j)_1 + FEDLAW_t{*}PAYER(j)_{ikt}\alpha(j)_2 + \\
& + PAYER(2)_{ikt}{*}EXPANDED_t\,\alpha_3 + \varepsilon_{ikt}
\end{aligned}
$$

Where X is a vector of characteristics of the newborn (insurance PAYER, sex, race, ethnicity, hour, day and month of birth), the mother (such as age, education, and the number of previous births) and the hospital (hospital size,[19] ownership status[20], and hospital service area[21]) and the admission index for

[19] We break hospitals up into 6 groups based on average monthly number of deliveries. The groups are <20, >20 and ≤ 50, >50 and 100, >100 and ≤ 150, >150 and ≤300, and >300.

children 90 to 180 days old, which controls for the day-to-day conditions that may generate readmissions.[22] The variable ε_{ikt} is an additive error.

Because we do not have a natural control group that was unaffected by the law changes, we must capture the time series in postpartum length that would have occurred in the absence of the California with time trends. We are aided by the fact that the pre-law and post-law trends are very similar as Figures 1-9 demonstrate. [23] To eliminate any secular trend in the data, we include a monthly trend $MTREND_t$ that equals 1 if the birthday of the newborn is in July of 1995, 2 in August, etc. We allow the monthly trend to vary by insurance PAYER status (where j=1 for Private and j=2 for Medicaid). Later in this paper, we allow the trends to vary based on more observed characteristics and we allow for higher order terms.

The key variables in the model are the vectors α_1 and α_2 and the parameter α_3 that measure the impact of the state and federal respectively. The variable STATELAW equals 1 from August 27, 1997 through December 31, 1997, while FEDLAW equals 1 from January 1, 1998 and on. These variables vary by payer status (Private or Medicaid). The variable EXPANDED equals 1 in January 1, 1999 and on and it captures the expansion of the California law to all Medicaid patients and given the nature of the expansion in 1999, this variable is only interacted with Medicaid.

There are two key outcomes: whether the newborn was discharged early (< 2 days for vaginal births and < 4 days for c-sections). Although these outcomes are discrete, we estimate linear probability models. We relax this assumption in later sections and demonstrate that even for low incidence events, linear probability estimates are very similar to estimates from probit models. In all models, we control for possible autocorrelation in errors by allowing for arbitrary correlation in errors within a hospital over

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[20] For hospital ownership variable, it is classified into 11 categories: 1=church, 2=non profit corporation, 3=no profit other, 4=individual investor, 5=partnership investor, 6=corporation investor, 7=state, 8=county, 9=city/county, 10=city, 11=district.

[21] We use 14 health service areas defined by the U.S. Department of Health and Human Services for the state of California. Health service areas are sometimes single counties (e.g., Los Angeles or Orange County) but in many cases, areas include multiple counties.

[22] For an infant born on day t, we use as an index the number of admissions per child over the next 28 days for all children aged 90-180 days who were alive on that birth date.

[23] Taking the aggregate data for Figure 1 and regressing the fraction of early discharges in private insurance on a time trend, dummies for the state and federal law, plus a time trend for the four months of the state law, we obtain an R^2 of 0.98.

time. This procedure also allows for arbitrary forms of heteroskedasticity which is present, by construction, in our linear probability models.

We titled this section 'a reduced-form model' for a particular reason. The California law changed two things at once. First, it required insurance carriers to provide coverage for longer postpartum hospital stays. Second, it required insurance carriers to provide coverage for a follow-up visit for mothers who, after consulting with their physician, were discharged from the hospitals early. Therefore, the estimated impact of the law contained in coefficients (the α's) captures both of these changes. This distinction is potentially important. Suppose that a) early discharge increases the chance of a hospital readmission, b) early follow-ups of patients released early eliminate this risk, and c) everyone released early has a follow-up visit. In this case, the coefficients on the α's will both be zero since the harm from an early discharge was compensated for by the office follow-up visit. Previous research from California has demonstrated, however, that this is probably not a concern.[24] Although in principal the α's capture both effects, any impact we estimate is likely to be driven primarily by the change in length of stay and not an increase in early follow-up visits.

b. Two-Stage Least-Squares Estimate

We argued in the previous section that the health benefits from the law are likely to be driven by longer postpartum length of stays and not the mandated coverage for early follow-up visits after an early discharge. We quantify the size of the relationship between postpartum length of stay and adverse outcomes by using the information from the reduced-form models in a more structured way. Specifically, we use the adoptions of the federal and state laws as instrumental variables for length of stay in a two-stage least squares (2SLS) model to obtain consistent estimates of the impact of length of stay on medical outcomes.

[24] Galbraith *et al*. (2003) surveyed 2,828 mothers in California in 1999 and found there was no difference in the percentage of newborns with an early follow-up (within 2 days of discharge) between those discharge early and those discharged later.

To more formally outline this model, note that one question of interest is the impact of length of stay on 28-day readmission rates. We can model this statistically with the following 'structural' equation of interest:

$$(2)\ 28DAY_{ikt} = X_{1ikt}\beta_1 + \text{Discharged Early}_{ikt}\ \delta_1 + \varepsilon_{1ikt}$$

where for simplicity, we include all the covariates from equation (1) into one vector X_1. The key covariate in this regression is *Discharged early*. As we noted above, equations such as (2) have been estimated by a number of authors but we suspect that cov($Early_{ikt}$, ε_{1ikt}) >0 so OLS estimates of δ_1 will be biased down.

Two-stage least squares estimation requires that a researcher identify a variable that exogenously changes the endogenous covariate of interest (LOS) but has no direct impact on health. In this case, the instruments are the enforcement dates for the state and federal law. As Figures 1 – 9 indicate, the federal law clearly changed hospital length of stay and for the reasons mentioned above it is plausible that this change was exogenous. For this particular study, we are fortunate that the timing of the law change is exact, and the immediate and large impact of the law makes it difficult to argue that some other event was explaining the sudden and precipitous change in hospital length of stays for newborns. Likewise, so long as we properly control for the secular trends in the outcome, the instruments should not generate any omitted variables bias in the outcome equation of interest.[25]

[25] Our key outcomes (re-admissions and neonatal mortality) are both discrete and the key covariate of interest (length of stay in days) is continuous. A 2SLS model where we instrument for LOS will therefore not mimic the data generating process well. This model could be estimated by maximum likelihood models where the outcome is discrete and the endogenous variable of interest is continuous (Evans, Oates and Schwab, 2002; Evans, Farrelly and Montgomery, 1998). However, given the size of the data set and the number of covariates in the model, this model will be difficult to estimate. Angrist (2001) has however demonstrated that two-stage least squares models applied to limited and discrete dependent variables replicate treatment effect parameters from more complicated maximum likelihood models.

VII. Results

a. Reduced Form Model Regression Results

Table 4 presents the OLS estimates results of equation (1) – the reduced-form model for newborns using three subsamples: vaginal without and with complications and cesarean deliveries. We initially report linear probability results for two dichotomous outcomes: whether the newborn was discharged early and whether they were readmitted within 28 days. We only report the coefficients on the legal variables and for space considerations, suppress the coefficients on the other key variables.

For privately insured vaginal deliveries without complications, we find that the California and federal law reduced early discharge rates of newborns by 16 and 31 percentage points, respectively, with the later result being 38 percent of the pre-law rates. The standard errors on these estimates are incredibly small and the results are statistically significant at conventional levels.[26] Interestingly, the federal law had only a slightly smaller impact on vaginal deliveries with complications although the mean discharged early is 14 percentage points lower in this later group. For c-section deliveries, the federal law is estimated to reduce early discharge rates by 13.5 percentage. In general, the effect of the state law is about half the impact of the federal law for those with private insurance. The Federal law is estimated to have reduced early discharge rates for Medicaid births by 12.3 and 14.5 percentage points for vaginal deliveries without and with complications, respectively, and both estimates are precisely estimated. Expansion of the state law to cover all Medicaid births increased this effect by nearly 18 percentage points in the uncomplicated vaginal sample but only by an addition 6 percentage points in the complicated vaginal group. By the time the state law was expanded, the early discharge statutes are estimated to have reduced early discharge rates for Medicaid-covered c-section newborns by about 7 percentage points.

In the next column of results for each subsample, we report the impact of the laws on 28-day readmission rates. If longer stays reduce readmission rates, then we should see reductions in rates in these rates after passage of the state and federal laws. For vaginal deliveries without complications, we observe a drop in readmissions rates of 0.07 percentage point after the passage of the federal law for

[26] References to statistically significant estimates assume a p-value of 0.05 or below.

privately insured, a number identical to the difference generated graphically in Figure 4, and the simple difference estimate in Table 2. The similarity of these three estimates is no surprise given he lack of a trend in graph in Figire 4. Among uncomplicated vaginal deliveries paid for by Medicaid, the federal law is estimated to reduce readmissions by 0.12 percentage points. Both results are statistically insignificant at conventional levels. We do however see a statistically significant reduction in readmission rates for Medicaid patients after the state law was expanded to include all Medicaid patients. In contrast, for the same subsample, we observe a statistically significant *increase* in readmissions after passage of the state law for both insurance types. This could be attributed to the law but more likely, this is due to the fact we cannot control perfectly for the large spike in admissions that is observed in December of 1997 during the severe flu season in California.[27]

For vaginal deliveries with complications, we observe reductions in readmission rates after the state and federal law and for both insurance types, but the results are only statistically significant for Medicaid patients after the federal law and the total effect on readmissions of the expansion of the state law of -0.01 percentage point is also statistically significant.. Among those with private insurance delivered by c-section, there was a drop in readmission rates of a little more than a half of a percentage point (p-value<0.08) after the passage of the federal law and the estimate for the Medicaid subsample is -0.0033 with a t-statistic of roughly -1.

Given the concerns that the spike in readmissions during the December 1997 period is due to some other cyclic shock not related to the passage of the state statute, we also report in Table 3 estimates where we delete data from September 1, 1997 through December 1, 1997. Estimates from these models are reported in the lower half of the table. This model specification only allows us to identify the estimate on the Federal law coefficients and the impact of the expansion of the state law to all Medicaid births. These estimates are nearly identical to the results found in the top half of the table.

[27] The coefficient (standard error) on the hospital admission index for 90 to 180 day old infants in this sample is 0.65 (0.059) but the string correlation between this and 28-day readmission rates in this sample cannot capture the huge spike in newborn readmissions during December of 1997.

Given differences in model specification, these results are not directly comparable to those in Datar and Sood (2006) who use public use versions of the data here to examine the same question. In their data set, the authors pool all births in one model and estimate the reduced-form relationship between readmission rates and the passage of the federal law. The public use version of the data set does not identify month of birth and the authors do not examine either the impact of the passage of the state law or the extension of the state law to all Medicaid births. The only trend included in the model is a linear trend in years and the trend is common to all groups. In logit models, the authors estimate that the federal law generated statistically significant reduction in the odds of readmission of 9.3, 11.8 and 19.7 percent during the first three years.

Using our data set, not controlling for the month of admission, adding a time trend that is linear in years and controlling for many of the same variables in Datar and Sood, in a logit model, we estimate that federal law changes the log odds of 28-day readmission by 8.7 percent (standard error of 2.6 percent). Exploiting the month of birth by adding a time trend that is linear in months and controlling for the month of birth reduces slightly the estimated change in the odds of readmission to -7.73 percent (standard error of 2.7 percent). However, once we control for the passage of the state law, the estimated impact of the federal law falls in magnitude to a -5.55 percent change in the odds with a standard error of 3.5 percentage points. By not controlling for the state law, it appears that pre-federal law readmission rates were substantially higher than average, increasing by a large amount the estimated treatment effect of the federal law. If for example, we delete the data during the final four months of 1997 (the period when only the state law was in effect) and re-estimate the model just outlined, we estimate that the federal law reduces the odds of a readmission by 5.4 percent (standard error of 3.6 percent). In this case, lack of restricted-use data greatly overstates the impact of the laws.[28]

[28] We were unable to replicate one set of results in Datar and Sood (2006). The authors estimate that the impact of the federal law on readmission rates doubles between the first and third years of implementation. This is surprising since the authors find only a 40% change in the impact of the law on length of stay from the first to the third year. When we estimate models similar to those in Datar and Sood, we find a uniform impact across years. Specifically, the percentage change in the odds ratio of a 28-day readmission (standard error) in the first three years of the law are estimated to be -8.3 (2.9), -10.4, (3.3), and -8.9 (4.0) respectively and the p-value on the null hypothesis that all three

Much of the literature in this area examines 28-day readmission rates as a key outcome but we can calculate 7 and 14 day readmission rates as well. As we shorten the follow up after birth, the incidence rate falls considerably. Linear probability models tend to generate marginal effects similar to estimates from probit or logit models but only when the mean outcome is some distance from zero or one. Therefore, one must be concerned with whether a linear probability model is appropriate for these low-incidence outcomes. In Table 5, we report the reduced-form probit regression models where the outcomes of interest are initially the 7, 14, and 28 day readmission rate for infants and in the table, we report the 'average treatment effect' which is the estimated change in the logistic CDF when the law dummies are turned on and off. In the fourth column of results, we also report the estimates from the linear probability model for the 28-day readmission rate for comparison. In this table, we use the model that excludes the last four months of 1997.

There are a number of key results in Table 5. First, in general, the marginal effects from the 28-day probit and the liner probability estimates of the same equation produce strikingly similar results in all subsamples. The linear probability model appears to do an adequate job in this case. Second, marginal effects in the 7-day readmission equation are, in general, a high fraction of the value of the same coefficient in the 28-day equation, indicating that the bulk of the readmissions prevented by the law would have happened within the first week.

In the final column of each section of results, we also report probit estimates where the outcome of interest is the 28-day mortality rate for newborns. Across the three samples, neonatal mortality is anywhere from 6 per 10,000 to 32 per 10,000 so this adverse outcome is very rare. Among privately insured newborns from uncomplicated pregnancies, the marginal effect for the federal law is a positive 0.00005 which is one-tenth the sample mean with a standard error twice the parameter size. Across all three samples, we estimate statistically significant drops in mortality for uncomplicated vaginal births and c-sections impacted by the expansions of the law to all Medicaid patients, and reduction in mortality of

coefficients are the same is 0.26. When we delete the data from the last four months of 1997 when only the state law was in effect, the coefficients are -3.5 (4.0), -3.7 (4.8), and -0.4 (5.7).

5.6 deaths per 10,000 among privately insured complicated vaginal births (p-value < 0.08). The results for this outcome are not very clear. Although there is some evidence that the drop in early discharges reduces mortality, the statistically significant results do not line up with the statistically significant results from the 28-day readmission equations, nor do the results line up with the large declines in early discharge from the first stages.

b. 2SLS Regression Results

The reduced-form estimates presented in the previous three tables represent the 'intention to treat' impacts of the state and federal law. In this section, we generate estimates of the impact of the law on those treated by calculating 2SLS estimates of equation (2). As we noted above in Table 1, those most likely to have longer hospital stays are also those most likely to experience a 28-day readmission. If the unobserved characteristics of mothers and their infants that predict 28-day readmission rates are correlated with length of stay in the same direction as these observed characteristics, then OLS estimates of equation (2) are biased towards zero.

In the first third of Table 6, we report OLS estimates of equation (2) to form a baseline to which we can compare 2SLS results. In this model, the 28-day readmission dummy variable is the outcome of interest and key covariate is whether the newborn was discharged early. The results across the three samples tell a fascinating story. The estimated impact of being discharged early among uncomplicated vaginal deliveries is incredibly small at a statistically significant three tenths of a percentage point, which is less than tenth of the pre-treatment sample mean. In contrast, an early discharge increases the chance of a 28-day readmission by 1.8 percentage points in both the complicated vaginal delivery and c-section samples, a number six times as large as the estimate for uncomplicated vaginal deliveries. All of these estimates are statistically significant at conventional levels.

If those with the longest hospital stays are those most likely to be readmitted, as the results in Table 1 suggest, then the OLS presented above should be biased downward. In the next third of the table, we report 2SLS estimates that use the five instruments for postpartum length of stay: the federal and state laws interacted with Medicaid and private insurance, thus the extension of the state law to Medicaid patients in 1999. In all three samples, 2SLS estimates of the discharged early variable are positive and larger than the OLS estimate. The estimate in the uncomplicated vaginal delivery sample is still small and a statistically insignificant .4 percentage points. In the complicated vaginal delivery sample, the 2SLS coefficient on the discharged early variable is 1.9 percentage points with a p-value of about 0.15, and estimate slightly larger than the OLS value. Finally, the same estimate in the c-section sample is 4.5 percentage points which slightly smaller than the pre-treatment sample mean and statistically significant at conventional levels. The F-tests that the instruments can be excluded from the first-stage equation of interest are all large indicating that finite sample bias is not a concern here. Finally, we do reject the test of over-identifying restrictions at the 0.05 level in the uncomplicated vaginal delivery sample. The test can be thought of as a test of the null that the 2SLS estimates are the same regardless of the instruments we use and since the reduced-form estimates are all over the map, is therefore no surprise we can reject the null in this case.

Given the imperfect controls for cyclic variation in the readmission rates, we find in reduced-form models that the state law actually increased readmission rates. Since the test of over-identifying restrictions can also be though of as a test of the null hypothesis that the 2SLS estimates are identical regardless of the instrument set used, we would expect to reject the null with test since the federal instruments is predicting a positive benefit of an additional day of stay whereas the other instruments predict a negative effect.

The reduced form estimates in Table 4 suggest that the early discharge laws improved outcomes for newborns insured by Medicaid and delivered by uncomplicated vaginal deliveries. There is however no evidence that these laws improved outcomes for privately insured newborns in the same subsample. In Table 7, we estimate separate 2SLS estimates for each subsample and by payer method. As with Table 6, we report both the OLS and 2SLS estimates but we only consider the models when the period the state law was in effect is deleted from the sample. Therefore the private insurance models are exactly identified and the Medicaid models have two instruments (Federal law x Medicaid and Expanded state law x Medicaid).

In these results we see that for complicated vaginal deliveries and c-section samples, there are large estimated positive impacts of early discharges on readmission rates, yet few of these results are statistically significant at conventional levels. Among uncomplicated vaginal deliveries, there is limited evidence that early discharges negatively impact health for privately insured newborns. However, the 2SLS coefficient on the early discharge variable in the Medicaid group for this subsample generates a large impact of early discharge on readmission and the coefficient has a p-value of roughly 0.08. Note that from Table 3, privately insured newborns from uncomplicated deliveries have the lowest readmission rate among all the subsamples we consider. These results suggest that early discharge have little negative health consequences (as measured by readmission rates) for those who we expect a priori to have the lower chance of a readmission anyway.

The results are robust to a number of alterations to the model. These results are presented in Table 8. In Figure 11, we plot the time series of the fraction of births in our sample that are to Hispanic mothers. Note the persistent nature of the cycle throughout the months. On the same graph, we report estimates of farm workers per month and the cycle in percent births to Hispanics is about two months out of phase of the farm worker cycle. For this reason, we allow the time trends to vary by insurance status and Hispanic origin (row 2), allow the month effects to vary by insurance status and ethnic origin (row 3), and allow the time trends to vary by insurance status, ethnicity and health service region in California (row 4). None of these models alter the qualitative results.

In rows 6 and 7 we alter the basic estimates in row 1 by allowing for quadratic and cubic time trends. There is mixed evidence of the robustness of the model. Adding quadratic does not change the qualitative nature of the results. There is concern however that adding these additional terms is 'over-fitting' the model because a p-value of the null hypothesis that the additional terms are all zero is in excess of 0.05 in all three models. However, when we add a cubic term, the p-values fall between 0.01 and 0.05 in two samples and both the qualitative and statistical significance of the results is eliminated.

VIII. Conclusion

In this paper, we use large and sudden changes in postpartum length of stay generated by the passage of a state and federal law to examine the effect of longer stays on the health of newborns. State and federal laws worked as intended in that the fraction of newborns discharged early fell dramatically for both privately insured and Medicaid newborns. However, changes in readmission rates generated by the passage of the laws were not nearly as uniform. We estimate the law had small if any benefit for the group with the lowest 28-day readmission rate, privately insured newborns from uncomplicated vaginal deliveries, a group that represents 30 percent of births in our time period of analysis. However, for all other subsamples, there is some evidence that early discharges increase readmission rates. Among c-sectional deliveries, complicated vaginal births and Medicaid patients with complicated vaginal deliveries, we estimate early discharges increase readmission rates by a substantial amount, with many of these results being statistically significant at conventional levels. These results suggest that for routine pregnancies, early discharge of newborns pose little health concern, yet those with the highest risk of readmission benefited enormously from passage of the early discharge laws.

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Figure 1: % Newborn Discharged Early, Vaginal Deliveries without Complications

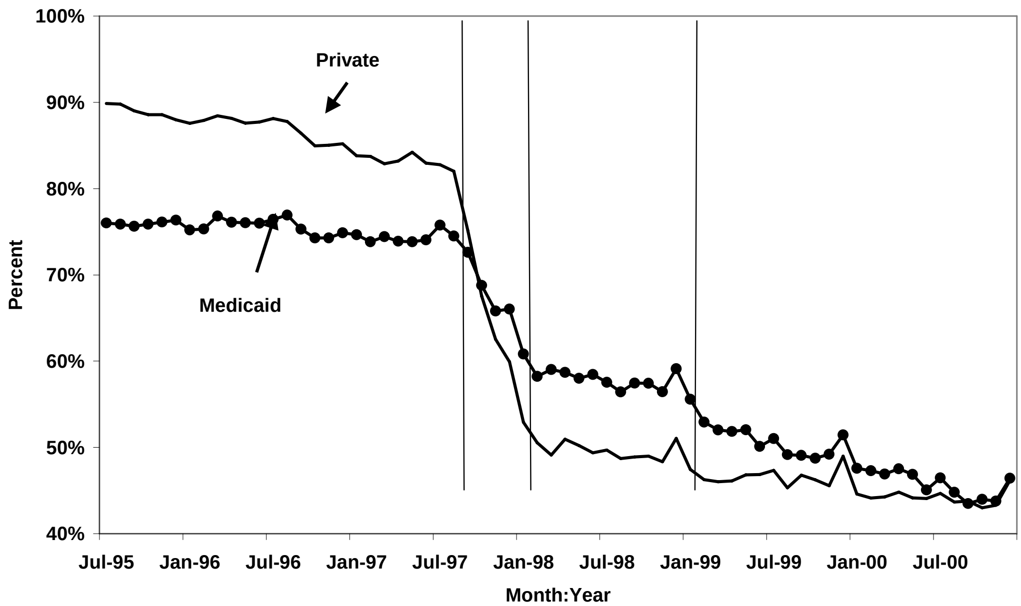

Figure 2: % Newborn Discharged Early, Vaginal Deliveries with Complications

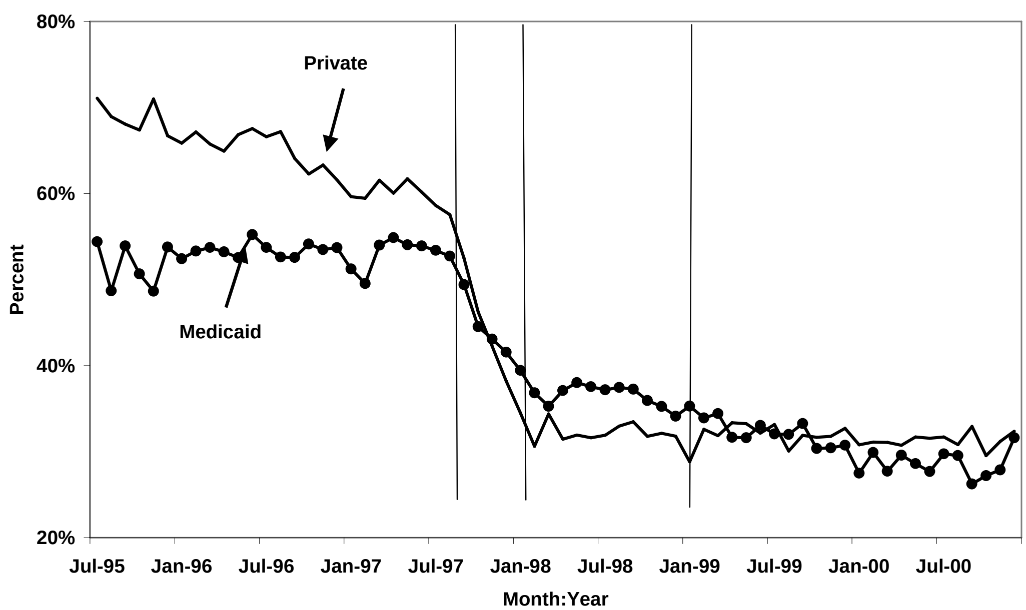

Figure 3: % Newborn Discharged Early,
C-Section Deliveries

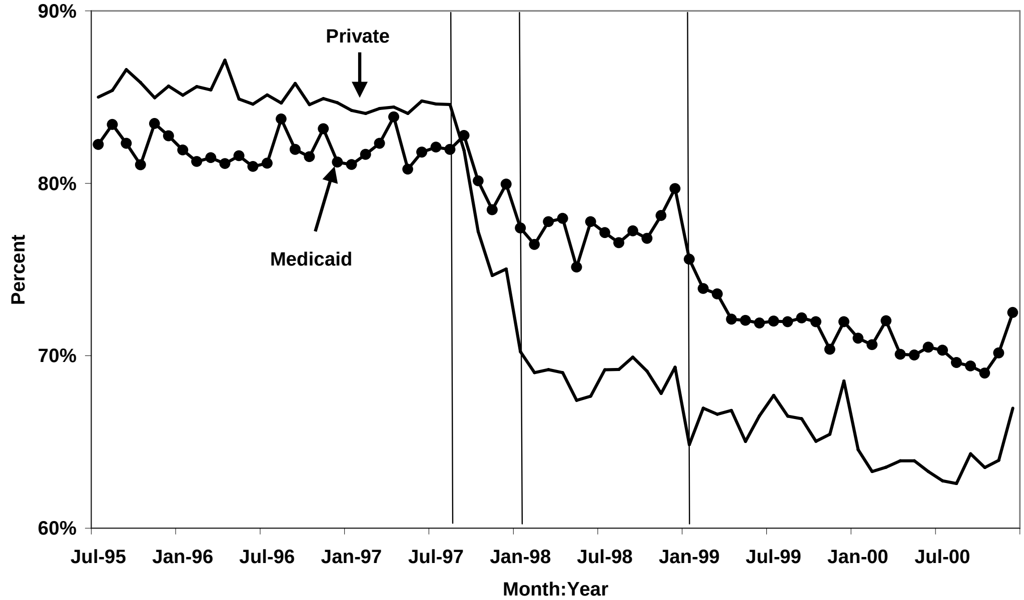

Figure 4: % Newborns Readmitted within 28-Days,
Vaginal Deliveries without Complications, Private Insurance

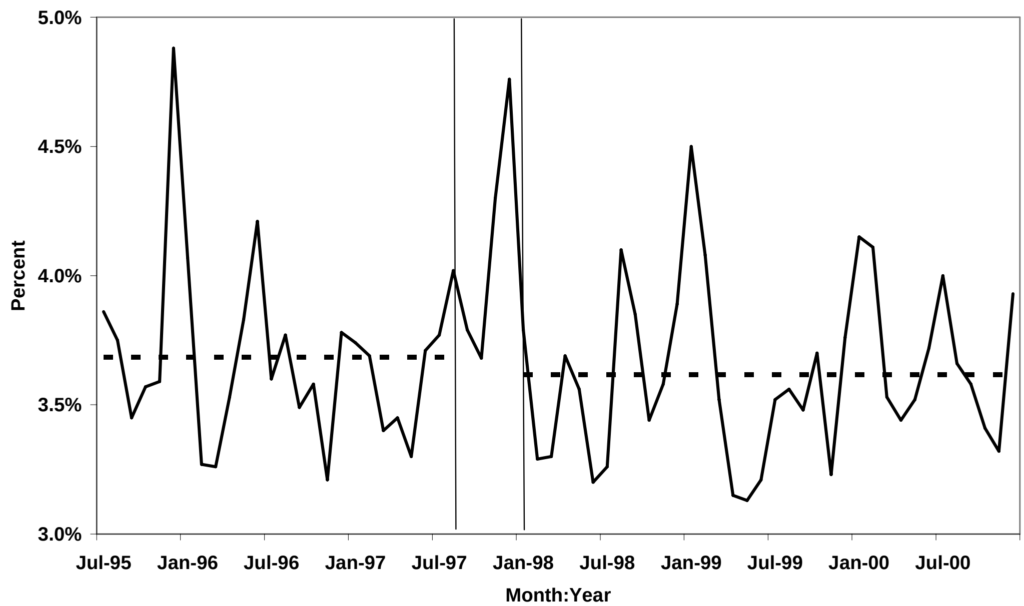

Figure 5: % Newborns Readmitted within 28-Days,
Vaginal Deliveries with Complications, Private Insurance

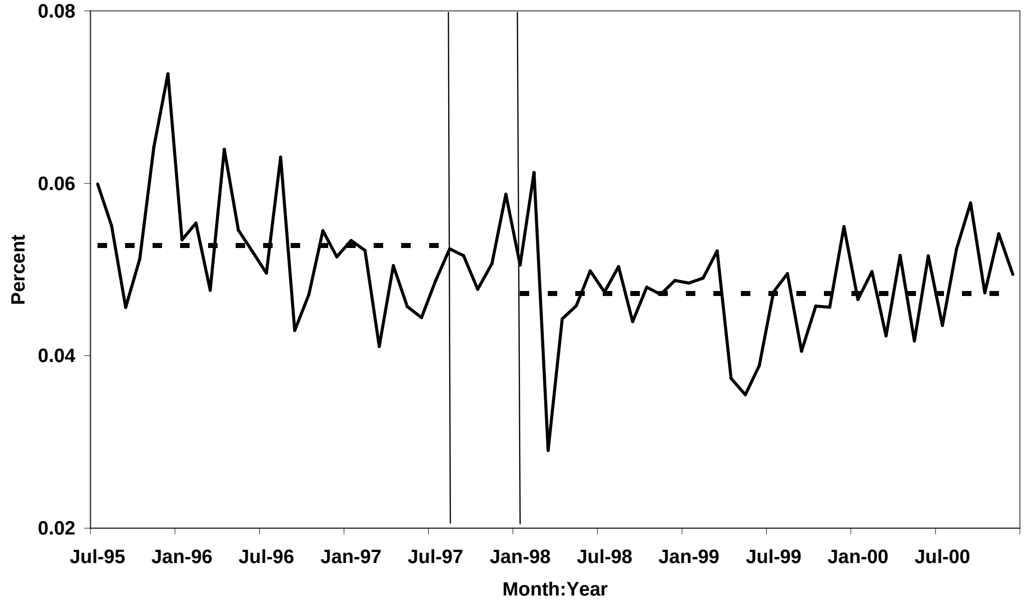

Figure 6: % Newborns Readmitted within 28-Days,
C-Section Deliveries, Private Insurance

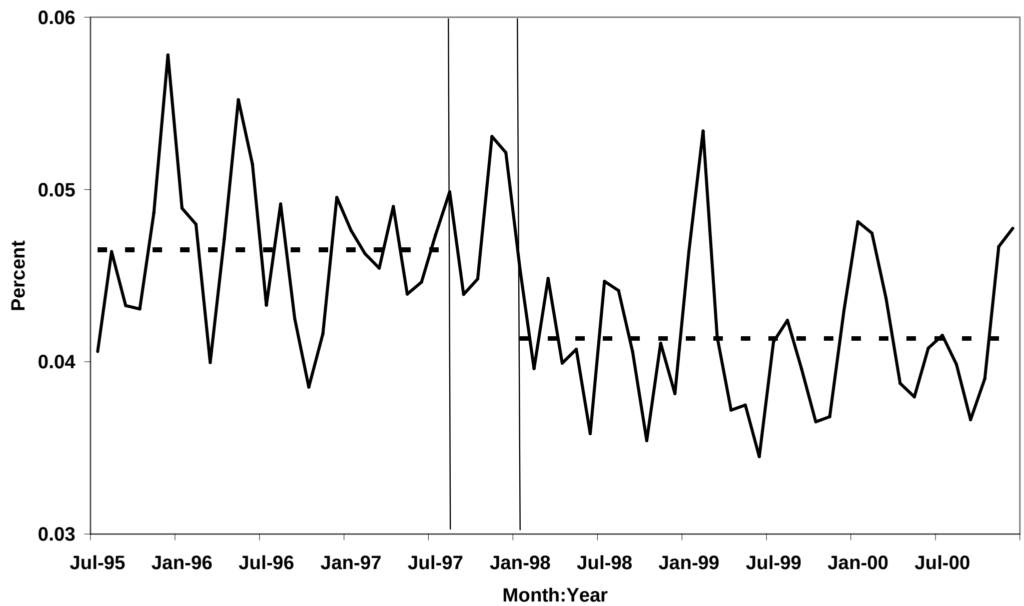

Figure 8: % Newborns Readmitted within 28-Days, Complicated Vaginal Deliveries, Medicaid Insurance

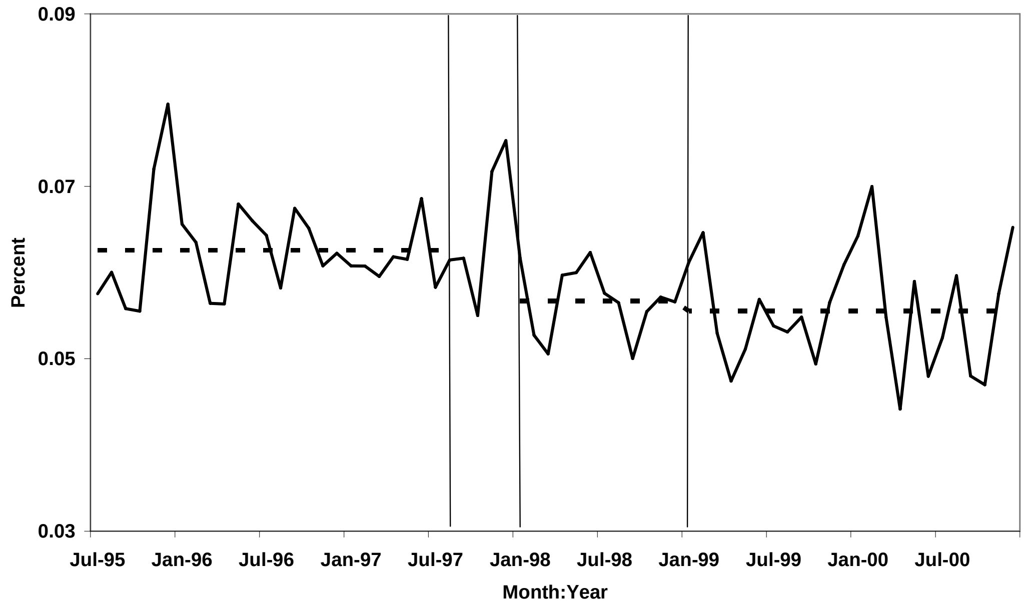

Figure 9: % Newborns Readmitted within 28-Days, C-Section Deliveries, Medicaid Insurance

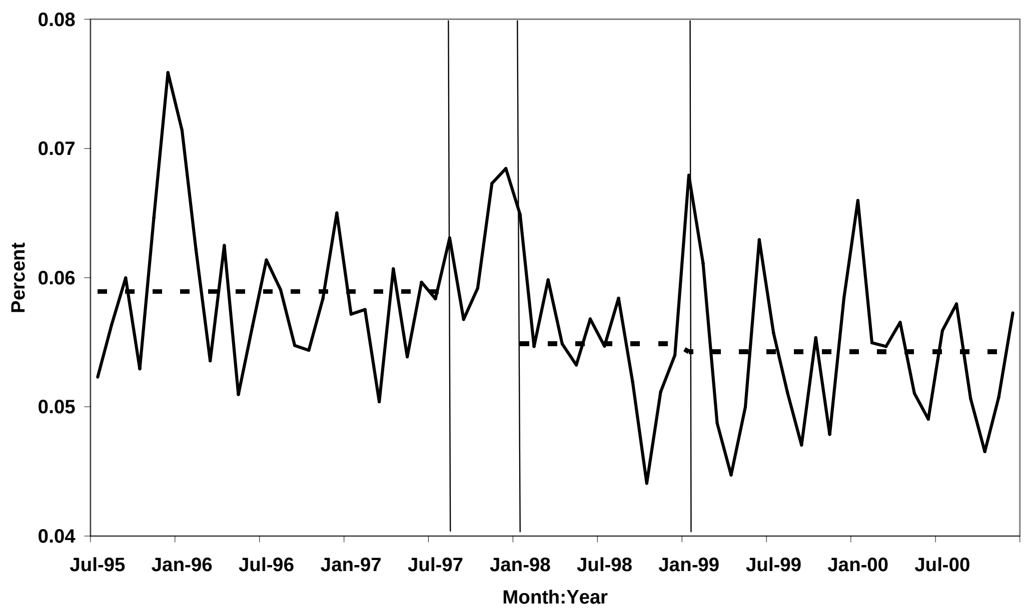

Figure 10: % Newborns Readmitted in 28-Days, Private Vaginal Uncomplicates, and 28-Day Hospital Admission Rate for 90-180 Day Old Infants

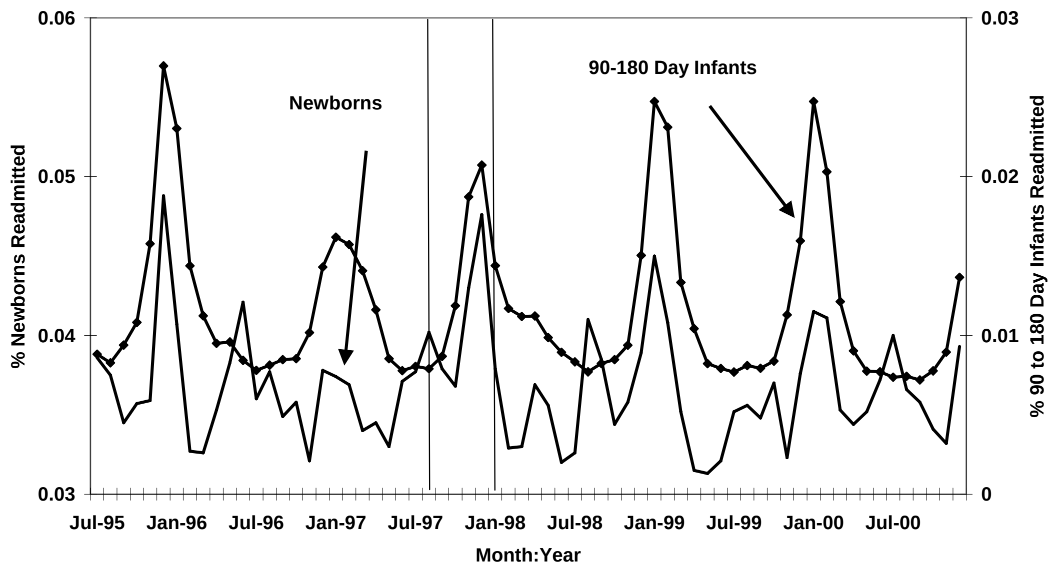

Figure 11: % of Births to Hispanic Mothers and Farm Employment in California

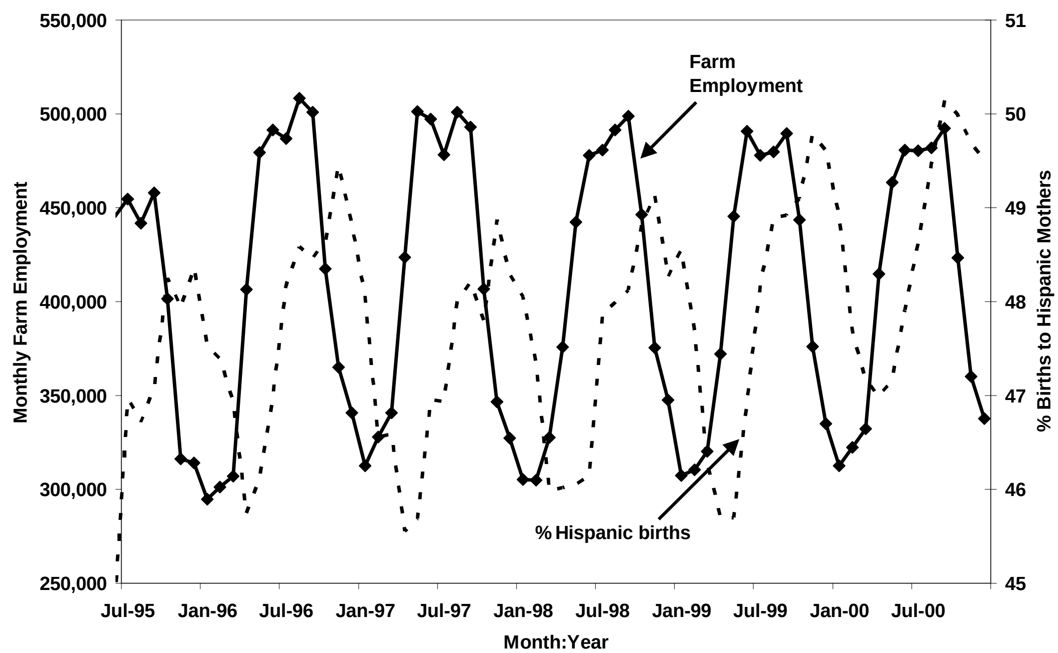

Table 1
OLS Estimates of Length of Stay and Logit Estimates of 28-Day Readmission Equations
Newborns in California Private/Medicaid Sample, 1995-1996

OLS Parameter Estimate or Marginal Effect

	Vaginal Deliveries		C-sections			Vaginal Deliveries		C-Sections	
		28-day		28-day			28-day		28-day
	LOS	Readmit.	LOS	Readmit.		LOS	Readmit.	LOS	Readmit.
Covariates	(1)	(2)	(1)	(2)		(1)	(2)	(1)	(2)
Private	-0.3079	-0.0055	-0.3353	-0.00641	Married	-0.1525	0.0002	-0.2247	0.00003
Insurance	(0.0129)	(0.0003)	(0.0514)	(0.0007)		(0.0119)	(0.0003)	(0.0472)	(0.0006)
Non-profit hospital	-0.3609	-0.0002	-0.1389	0.0006	Mother ≤ 20 years	-0.4392	-0.00005	0.2129	0.0103
	(0.0157)	(0.0004)	(0.0640)	(0.0007)	of age	(0.0242)	(0.0006)	(0.0898)	(0.0014)
For-profit hospital	-0.4290	-0.0019	-0.6868	-0.0016	Mother 21-25 years	-0.3836	-0.0006	-0.2005	0.00003
	(0.0184)	(0.0005)	(0.0723)	(0.0009)	of age	(0.0211)	(0.0005)	(0.0730)	(0.0010)
Boy	0.1191	0.0056	0.0562	0.0045	Mother 26-30 years	-0.2630	-0.0007	-0.2979	-0.0023
	(0.0100)	(0.0002)	(0.0395)	(0.0005)	of age	(0.0197)	(0.0005)	(0.0651)	(0.0009)
0 complications	-3.5362	-0.0151	-7.7511	-0.0255	Mother 31-35 years	-0.1497	-0.00006	-0.1765	-0.0045
during pregnancy	(0.0814)	(0.0009)	(0.2162)	(0.0013)	of age	(0.0201)	(0.0005)	(0.0642)	(0.0010)
1 complication	-2.2407	0.0008	-4.7291	-0.0004	White, non-Hispanic	-0.1172	0.0010	-0.2165	0.0026
during pregnancy	(0.0829)	(0.0010)	(0.2213)	(0.0014)		(0.0130)	(0.0004)	(0.0513)	(0.0009)
2 complications	-1.9378	0.0026	-3.0720	0.0057	Black, non-Hispanic	0.4360	-0.0025	1.1868	-0.0026
during pregnancy	(0.0901)	(0.0014)	(0.2438)	(0.0020)		(0.0222)	(0.0008)	(0.0819)	(0.0015)
0 complications	-1.7962	-0.0087	-1.4601	-0.0089	Other race, non-	0.0068	0.0024	0.2498	-0.0003
during delivery	(0.0399)	(0.0006)	(0.0704)	(0.0009)	Hispanic	(0.0169)	(0.0006)	(0.0689)	(0.0013)
1 complication	-1.7728	-0.0037	-1.8587	-0.0053	Mother < HS	0.1261	-0.0017	-0.0899	0.0002
during delivery	(0.0496)	(0.0011)	(0.1065)	(0.0017)	education	(0.0198)	(0.0005)	(0.0759)	(0.0010)
2 complications	-1.2634	0.0009	-1.2082	0.0006	Mother HS education	0.0660	-0.0010	-0.2032	-0.0006
during delivery	(0.0428)	(0.0007)	(0.0731)	(0.0009)		(0.0175)	(0.0004)	(0.0651)	(0.0009)
No previous births	0.2152	0.0053	-0.9158	-0.0070	Mother some college	0.0448	0.0001	-0.1615	-0.0024
	(0.0185)	(0.0004)	(0.0720)	(0.0009)	education	(0.0175)	(0.0005)	(0.0643)	(0.0010)
1 previous birth	-0.0222	-0.0017	-0.8449	-0.0027	Observations	754,107	754,107	205,625	205,625
	(0.0176)	(0.0004)	(0.0700)	(0.0009)	Mean of outcome	1.6950	0.0450	4.4162	0.0535
2 previous births	-0.0757	-0.0027	0.6587	0.0007	R^2/-2 Log like.	0.0222	0.0086	0.0344	0.0169
	(0.0187)	(0.0005)	(0.0746)	(0.0010)					

Note: Standard errors in parenthesis. Other covariates include month and year dummy variables. The reference categories are Medicaid insurance, government hospital, girls, 3 or more complications during pregnancy or delivery, unmarried mothers, and mothers aged 36 or more who are Hispanic and with a college education.

Table 2
Descriptive Statistics, Privately and Medicaid Insured Births in California,
July 1 1995 – August 31, 2005

Means and (Standard Deviations)

	Uncomplicated Vaginal Deliveries		Complicated Vaginal Deliveries		C-Section Deliveries	
Variable	Private	Medicaid	Private	Medicaid	Private	Medicaid
Total charges (mother + infant)	$8,885 ($36,391)	$8,487 ($35,741)	$13,950 ($39,957)	$14,407 ($45,089)	$24,213 ($57,954)	$24,557 ($58,581)
% Mother's age<20	0.087 (0.282)	0.294 (0.456)	0.082 (0.274)	0.275 (0.447)	0.042 (0.202)	0.187 (0.390)
% Mothers' age 20-24	0.200 (0.400)	0.339 (0.473)	0.177 (0.382)	0.309 (0.462)	0.129 (0.335)	0.286 (0.452)
% Mother's age 24-30	0.318 (0.466)	0.216 (0.411)	0.300 (0.458)	0.218 (0.413)	0.283 (0.450)	0.252 (0.435)
% Black	0.051 (0.219)	0.073 (0.260)	0.061 (0.240)	0.084 (0.284)	0.066 (0.248)	0.093 (0.251)
% Other race	0.169 (0.375)	0.079 (0.270)	0.162 (0.357)	0.094 (0.292)	0.160 (0.367)	0.065 (0.247)
% Hispanic	0.279 (0.448)	0.662 (0.473)	0.256 (0.437)	0.594 (0.491)	0.256 (0.436)	0.642 (0.479)
% < high school education	0.137 (0.344)	0.571 (0.495)	0.128 (0.334)	0.554 (0.497)	0.113 (0.317)	0.531 (0.499)
% with a high school degree	0.532 (0.499)	0.404 (0.491)	0.535 (0.499)	0.417 (0.493)	0.533 (0.499)	0.435 (0.496)
% w/ previous birth	0.603 (0.489)	0.624 (0.484)	0.562 (0.496)	0.603 (0.489)	0.571 (0.495)	0.647 (0.478)
Observations	734,257	625,054	297,446	206,583	324,296	240,433

Numbers in parentheses are standard deviations. Dollar costs are measured in real 2000 dollars using the Consumer Price Index-all products index as a deflator.

Table 3
Early Discharge and Readmission Rates for Newborns,
Before and After Federal Law

		Uncomplicated Vaginal Deliveries		Complicated Vaginal Deliveries		C-Section Deliveries	
		Private	Medicaid	Private	Medicaid	Private	Medicaid
Discharged early							
(1)	7/1/95 – 8/31/97	0.863	0.754	0.719	0.602	0.850	0.821
(2)	1/1/98-12/31/00	0.469	0.516	0.385	0.390	0.663	0.733
	Difference (2)-(1)	-0.394	-0.238	-0.334	-0.213	-0.187	-0.087
		(0.0011)	(0.0012)	(0.0018)	(0.0024)	(0.0016)	(0.0019)
Readmitted with 28 days							
(1)	7/1/95 – 8/31/97	0.0368	0.0445	0.0524	0.0625	0.0465	0.0589
(2)	1/1/98-12/31/00	0.0361	0.0421	0.0448	0.0559	0.0413	0.0544
	Difference (2)-(1)	-0.0007	-0.0023	-0.0076	-0.0066	-0.0052	-0.0045
		(0.0004)	(0.0005)	(0.0008)	(0.0011)	(0.0008)	(0.0010)
Observations							
(1)	7/1/95 – 8/31/97	278,096	268,247	107,841	90,683	114,674	97,619
(2)	1/1/98-12/31/00	411,032	318,619	171,597	103,084	190,599	128,802

Standard errors are reported in parentheses.

Table 4
OLS Estimates, Impact of Early Discharge Laws on Early Discharge and 28-Day Readmission Rates for Newborns, California Births, July 1995-December 2006

	Uncomplicated Vaginal Deliveries (1,359,150 obs.)		Complicated Vaginal Deliveries (503,795 observations)		C-Section Deliveries (564,248 observations)	
	Discharged early	Readmitted w/in 28 days	Discharged Early	Readmitted w/in 28 days	Discharged early	Readmitted w/in 28 days
			All data, July 1, 1995 – December 31, 2000			
Federal law x	-0.3091	-0.0007	-0.2696	-0.0040	-0.1335	-0.0058
private insurance	(0.0177)	(0.0012)	(0.0197)	(0.0036)	(0.0128)	(0.0032)
Federal law x	-0.1228	-0.0012	-0.1449	-0.0063	-0.0279	-0.0032
Medicaid	(0.0119)	(0.0014)	(0.152)	(0.0031)	(0.0077)	(0.0031)
Expanded state law x	-0.1794	-0.0043	-0.0585	-0.0037	-0.0431	-0.0009
Medicaid	(0.0144)	(0.0021)	(0.0075)	(0.0022)	(0.0097)	(0.0020)
State law x	-0.1624	0.0035	-0.1591	-0.0027	-0.0576	0.0009
private insurance	(0.0134)	(0.0012)	(0.0118)	(0.0031)	(0.0090)	(0.0028)
Statel law x	-0.0368	0.0029	-0.0613	-0.0010	-0.0069	0.0022
Medicaid	(0.0079)	(0.0016)	(0.0109)	(0.0036)	(0.0062)	(0.0030)
R^2	0.2362	0.0027	0.1958	0.0053	0.0859	0.0061
			Delete data from September – December 1997			
			(472,971 observations)		(531,213 Observations)	
Federal law x	-0.3095	-0.0007	-0.2700	-0.0040	-0.1337	-0.0058
private insurance	(0.0178)	(0.0012)	(0.0197)	(0.0035)	(0.0128)	(0.0033)
Federal law x	-0.1232	-0.0012	-0.1455	-0.0062	-0.0282	-0.0031
Medicaid	(0.0120)	(0.0014)	(0.0152)	(0.0031)	(0.0077)	(0.0031)
Expanded state law x	-0.1809	-0.0043	-0.0602	-0.0035	-0.0440	-0.0007
Medicaid	(0.0147)	(0.0022)	(0.0075)	(0.0022)	(0.0097)	(0.0020)
R^2	0.2416	0.0026	0.2002	0.0054	0.0880	0.0052

Standard errors are reported in parentheses. Other covariates include the hospital admission index for 90 to 180 day infants, controls for payer status, mother's education, race, ethnicity, age, and previous births, the size, location and type of hospital, the hour, day and month of birth, plus time trends that vary by insurance status. Standard errors are calculated allowing for arbitrary correlation in errors within a hospital.

Table 5
Reduced-Form Regressions Logistic Regressions of Infant Readmission and Mortality Models, California Privately Insured and Medicaid Patients, July 1995 through December 2000

	Probit 7-day readmission	Probit 14-day Readmission	Probit 28-day readmission	OLS 28-day readmission	Probit 28-day mortality
	Vaginal deliveries without complications (1,359,308 Observations)				
Federal law x private insurance	-0.00066 (0.00100)	-0.00042 (0.00111)	-0.00073 (0.00128)	-0.0007 (0.0012)	0.00005 (0.00011)
Federal law x Medicaid	0.00034 (0.00118)	-0.00041 (0.00117)	-0.00118 (0.00131)	-0.0012 (0.0014)	-0.00011 (0.00009)
Expanded state law x Medicaid	-0.00128 (0.00091)	-0.00205 (0.00092)	-0.00292 (0.00109)	-0.0043 (0.0022)	-0.00023 (0.00008)
Mean of dependent var.	0.0234	0.0293	0.0395	0.0395	0.00055
	Complicated Vaginal Deliveries (472,971 observations)				
Federal law x private insurance	-0.0024 (0.0027)	-0.0028 (0.0029)	-0.0036 (0.0033)	-0.0040 (0.0035)	-0.00056 (0.00031)
Federal law x Medicaid	-0.0033 (0.0024)	-0.0039 (0.0022)	-0.0053 (0.0024)	-0.0062 (0.0031)	0.00054 (0.00044)
Expanded state law x Medicaid	-0.0022 (0.0017)	-0.0034 (0.0017)	-0.0035 (0.0019)	-0.0035 (0.0022)	0.00014 (0.00036)
Mean of dependent var.	0.0324	0.0416	0.0523	0.0523	0.00186
	C-section deliveries without complications (531,213 Observations)				
Federal law x private insurance	-0.0043 (0.0026)	-0.0049 (0.0028)	-0.0059 (0.0030)	-0.0058 (0.0033)	0.00003 (0.00040)
Federal law x Medicaid	-00018 (0.0021)	-0.0017 (0.0022)	-0.0025 (0.0025)	-0.0031 (0.0031)	-0.00028 (0.00047)
Expanded state law x Medicaid	-0.0021 (0.0015)	-0.0028 (0.0017)	-0.0008 (0.0018)	-0.0007 (0.0020)	-0.00075 (0.00037)
Mean of dependant var.	0.0318	0.0378	0.0488	0.0488	0.00317

Standard errors are reported in parentheses. Other covariates include the hospital admission index for 90 to 180 day infants, controls for payer status, mother's education, race, ethnicity, age, and previous births, the size, location and type of hospital, the hour, day and month of birth, plus time trends that vary by insurance status. Standard errors are calculated allowing for arbitrary correlation in errors within a hospital.

Table 6
2SLS Estimates of Newborn 28-Day Readmission Equation,
Privately Insured and Medicaid Deliveries in California,
July 1995 through December 2000

	Uncomplicated vaginal Deliveries	Complicated vaginal deliveries	C-section Deliveries
	OLS Estimates, Newborn 28-day readmission equation		
Discharged early	0.0030	0.0179	0.0180
	(0.0011)	(0.0036)	(0.0029)
Mean of dep. variable, 7/1/95-8/31/97	0.040	0.057	0.052
Observations	1,359,150	507,795	564,248
	2SLS Estimates, Newborn 28-day readmission equation (Using Federal law*private, Federal law*Medicaid, Expanded state law x Medicaid, State law x Private, State law x Medicaid as instruments)		
Discharged early	0.0040	0.0190	0.0456
	(0.0037)	(0.0130)	(0.0227)
F-test (p-value) for 1st stage	78.4	57.8	28.3
	(<0.0001)	(<0.0001)	(<0.0001)
P-value, Test of over-identifying restrictions	<0.0001	0.094	0.1020
Observations	1,359,150	507,795	564,248
	2SLS Estimates, Newborn 28-day readmission equation (Using Federal law*private, Federal law*Medicaid, and Expanded state law x Medicaid as instruments)		
Discharged early	0.0036	0.0193	0.0477
	(0.0031)	(0.0131)	(0.0249)
F-test (p-value) for 1st stage	52.3	86.2	45.5
	(<0.0001)	(<0.0001)	(<0.0001)
P-value, Test of over-identifying restrictions	0.0187	0.199	0.4135
Observations	1,275,833	472,971	531,213

Standard errors are reported in parentheses. Other covariates include the hospital admission index for 90 to 180 day infants, controls for payer status, mother's education, race, ethnicity, age, and previous births, the size, location and type of hospital, the hour, day and month of birth, plus time trends that vary by insurance status. Standard errors are calculated allowing for arbitrary correlation in errors within a hospital.

Table 7
2SLS Estimates of Newborn 28-Day Readmission Equation,
Privately Insured and Medicaid Deliveries in California,
July 1995 through December 2000

Covariate	Uncomplicated vaginal Deliveries		Complicated vaginal deliveries		C-section Deliveries	
	Private	Medicaid	Private	Medicaid	Private	Medicaid
	OLS Estimates, Newborn 28-day readmission equation					
Discharged early	0.0026	0.0037	0.0184	0.0179	0.0184	0.0190
	(0.0016)	(0.0014)	(0.0049)	(0.0036)	(0.0037)	(0.0027)
Observations	689,047	586,786	279,318	193,653	304,998	226,215
	2SLS Estimates, Newborn 28-day readmission equation (Using Federal law*private, Federal law*Medicaid as instruments)					
Discharged early	0.0017	0.0187	0.0142	0.0436	0.0442	0.0316
	(0.0040)	(0.0110)	(0.0137)	(0.0131)	(0.0239)	(0.0469)
Observations	689,047	586,786	279,318	193,653	304,998	226,215

Standard errors are reported in parentheses. Other covariates include the hospital admission index for 90 to 180 day infants, controls for payer status, mother's education, race, ethnicity, age, and previous births, the size, location and type of hospital, the hour, day and month of birth, plus time trends that vary by insurance status. Standard errors are calculated allowing for arbitrary correlation in errors within a hospital. The F-test for the 1st stage is the test of the null hypothesis that all the instruments are zero.

Table 8
2SLS Estimates of Newborn 28-Day Readmission Equation,
Privately Insured and Medicaid Deliveries in California,
July 1995 through December 2000

	Uncomplicated vaginal Deliveries	Complicated vaginal deliveries	C-section Deliveries
(1) Baseline model, Table 4	0.0036	0.0193	0.0477
	(0.0031)	(0.0131)	(0.0249)
(2) Allow trends to vary by insurance status and	0.0021	0.0194	0.0425
Hispanic origin	(0.0039)	(0.0131)	(0.0231)
(3) Allow trends and month effects to vary by	0.0022	0.0194	0.0435
insurance status and Hispanic origin	(0.0040)	(0.0132)	(0.0231)
(4) Allow trends to vary by insurance status,	0.0022	0.0187	0.0436
Hispanic origin and health service area	(0.0031)	(0.0126)	(0.0231)
(5) add to model (1) quadratic time trends that	0.0036	0.0156	0.0392
vary with insurance status and Hispanic origin	(0.0039)	(0.0148)	(0.0240)
[p-value on F-test that the additional variables are zero]	[0.180]	[0.449]	[0.282]
(6) add to model (5) cubic time trends that vary	0.0025	-0.0090	0.0028
with insurance status and Hispanic origin	(0.0042)	(0.0194)	(0.0298)
[p-value on F-test that the additional variables are zero]	[0.298]	[0.023]	[0.0171]

Standard errors are reported in parentheses. Other covariates include the hospital admission index for 90 to 180 day infants, controls for payer status, mother's education, race, ethnicity, age, and previous births, the size, location and type of hospital, the hour, day and month of birth, plus time trends that vary by insurance status. Standard errors are calculated allowing for arbitrary correlation in errors within a hospital.