# After Midnight: A Regression Discontinuity Design in Length of Postpartum Hospital Stays*

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*Estimates of moral hazard in health insurance markets can be confounded by adverse selection. This paper considers a plausibly exogenous source of variation in insurance coverage for childbirth in California. We find that additional health insurance coverage induces substantial extensions in length of hospital stay for mother and newborn. However, remaining in the hospital longer has no effect on readmissions or mortality, and the estimates are precise. Our results suggest that for uncomplicated births, minimum insurance mandates incur substantial costs without detectable health benefits.*

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**Abstract**

*Estimates of moral hazard in health insurance markets can be confounded by adverse selection. This paper considers a plausibly exogenous source of variation in insurance coverage for childbirth in California. We find that additional health insurance coverage induces substantial extensions in length of hospital stay for mother and newborn. However, remaining in the hospital longer has no effect on readmissions or mortality, and the estimates are precise. Our results suggest that for uncomplicated births, minimum insurance mandates incur substantial costs without detectable health benefits.*

## I. Introduction

The US spends substantially more on healthcare than any other country: 15.3% of GDP compared to less than 9% for the median OECD country; Switzerland, the next highest-spending country, spends only 11.3% of GDP [OECD, 2008]. High US health expenditures are not accompanied by superior rankings of aggregate health outcomes. The health of US infants compares especially poorly: infant mortality is more than one-third higher than in Portugal, Greece, Ireland, and Britain, and double the rate in Japan and the Nordic countries. Meanwhile, life expectancy is roughly similar across these countries.

This combination of much higher spending and roughly similar health outcomes has been found across markets within the US as well, which has led to the conclusion that healthcare may have reached the "flat of the curve": diminishing returns may have set in such that additional spending yields little benefit (see, e.g., Peter Zweifel, Friedrich Breyer and Mathias Kifmann (2009) and references therein). In particular, there are concerns that moral hazard problems have led to costly overconsumption in the US (Katherine Baicker and Amitabh Chandra 2008, Alan M. Garber and Jonathan Skinner 2008). Such overuse may include hospital care for childbirth: Carol Sakala and Maureen P. Corry (2008) describe the "casual application" of costly medical procedures to healthy labor and delivery patients.

This paper tests for moral hazard in the care of newborns by comparing treatment and health outcomes across essentially identical newborns who happen to differ in their insurance coverage. This focus on US newborns offers three advantages. First, childbirth is the most common reason for hospitalization: mothers and newborns account for roughly one quarter of all patients discharged from US hospitals (Sakala and Corry 2008). While the mortality rate for newborns is generally much lower than previous study populations (most commonly, patients suffering heart attacks), the gains from survival in terms of additional years are an order of magnitude greater. Second, nearly all deliveries are performed in hospitals, and, as such, there is little potential for selection bias. This strength is analogous to that of studies of heart attack patients (M. McClellan, B.J. McNeil and J.P Newhouse 1994, David Cutler, Mark McClellan, Joseph Neewhouse and Dahlia Remler 1998, Amitabh Chandra and Douglas Staiger 2007). Finally, US postpartum hospital stays are significantly shorter than in other countries.[1] Given diminishing marginal returns, US postpartum hospitalizations constitute a strong test of moral hazard.

A main limitation when testing for moral hazard – and the effects of healthcare treatment on health outcomes more generally – is that insurance coverage and treatment levels are chosen and are likely endogenous to the underlying health of the patients (i.e., adverse selection). We address this endogeneity by exploiting the coarseness of hospital reimbursement schedules: hospitals are reimbursed based on the number of days a patient is in the hospital. These days are counted as the number of midnights in care. A newborn delivered at 12:05 a.m. will have nearly a full day in care before being "logged" as present in the hospital, whereas a newborn delivered at 11:55 p.m. will be counted as present only 5 minutes after delivery. Coupled with insurance coverage of at least 1 or 2 days, those born just after midnight are covered for an additional night of care compared to those born just prior to midnight. This type of variation lies at the heart of recent policy debates over the appropriate minimum level of coverage, which led 42 US states and ultimately the federal government to mandate minimum stays of 2 days (William N. Evans, Craig Garthwaite and Heng Wei 2008).

Our analysis uses hospital discharge data linked to birth and death certificates for all births in California from 1991 to 2002, including nearly 100,000 births within 20 minutes of midnight. These data report the hour and minute of birth. An additional dimension of our empirical analysis is made possible by the introduction of California's minimum-insurance mandate on August 26, 1997. The Newborns' and Mothers' Health Act, which required insurance coverage for at least two days of hospitalization following childbirth, allows us to trace out the effect of an increase in stay length using the midnight discontinuity from two different baselines. Prior to the law, the midnight threshold primarily induced variation between 0 and 1 additional midnights (i.e. one versus two total nights in the hospital). Following the law, the midnight threshold primarily induced patients to switch between 1 and 2 additional midnights (or two versus three total nights in hospital). We can therefore directly test for diminishing returns to stay length and assess both the current two-day minimum mandated by law (and recommended by the American Academy of Pediatrics) as well as a further expansions in minimum stay length: stays that are closer to averages in Europe and typical of US stays during the early 1980s.

We find that the discontinuity in insurance coverage associated with the minute of birth generates a substantial difference in average stay length, despite nearly identical observable characteristics. Infants born shortly after midnight spend an additional 0.25 nights in the hospital, on average (i.e. 1 in 4 newborns born after midnight spent an extra night in the hospital). This difference is similar to the change in length of stay when the California mandate increased the minimum stay by one day.

[1]US postpartum hospital stays following a vaginal birth average close to 2 days, "distinctly shorter than that of Australia (3.0), Great Britain (3.0), Sweden (4.0), Norway (4.5), or Japan (6.5)" (Eudene Declercq and Diane Simmes 1997). Similarly, Graham W. S. Scott (2006) found average postpartum stay lengths were more than twice as long in Germany and France as in the US in 2002.

If health benefits are absent, longer stay lengths induced by the availability of a "free night" constitute moral hazard. And indeed, for conditions that usually dominate welfare calculations – serious health problems and mortality – we find no effect of a post-midnight birth. Both visual inspection of the raw data and regression models that control for patient characteristics reveal estimates close to zero for hospital readmissions and infant mortality. The absence of major health effects from exogenous variation in hospital stays – even when the baseline stay length is fairly short – calls into question the welfare benefit of insurance mandates governing stay length. While the expanded choice set of reimbursable stays *per se* benefitted patients, the cost to insurers of providing this choice was apparently not offset by an appreciable reduction in insurer outlays for follow-up care (and may thereby have increased premiums). Conversely, we infer the push toward shorter lengths of stay during the early 1990s before mandates were enacted presumably lowered insurer costs (who would have been responsible for follow-up care necessitated by shorter initial stays).

Our results apply to a population that is induced to have a longer hospital stay as a result of a post-midnight birth, a group that we show is broadly similar to the overall population in terms of background characteristics, but faces low health risks. Likewise, low-risk newborns are presumably those most likely affected by minimum stay-length mandates. In addition to evaluating exogenous variaton in stay length, we also consider the demographic and health characteristics of those who extended their hospital stays in response to being offered an additional "free night." To our knowledge, ours is the first study to estimate the average characteristics of those exhibiting moral hazard.

The rest of the paper is organized as follows. Section II considers empirical challenges to assessing moral hazard, the background that led to early discharge laws in the US, and the role that minute of birth plays in determining the length of stay. Section III describes the data. Section IV describes how we estimate complier characteristics and *LATE*. Section V presents the results and Section VI interprets the results in relation to the costs of an additional night in care. Section VII concludes.

## II. Background

### *A. Assessing Moral Hazard*

In observational studies, adverse selection tends to confound estimates of moral hazard (Pierre-André Chiappori, Franck Durand and Pierre-Yves Geoffard 1998, Amy Finkelstein and James Poterba 2006). David M. Cutler and Richard J. Zeckhauser (2000) concluded that: "the data nearly uniformly suggest that adverse selection is quantitatively large." To the extent this selection process is overlooked, demand for healthcare will appear higher among those with more insurance. Differentiating moral hazard from adverse selection is of "considerable interest" to the design of efficient public policies because "the government can potentially ameliorate inefficiencies produced by adverse selection, but it is unlikely to have any comparative advantage relative to the private sector in redressing inefficiencies caused by moral hazard" (Finkelstein and Poterba 2006).

Ideally, moral hazard would be assessed when "a given set of agents experience a sudden and exogenous change in the incentive structure they are facing" (Pierre-André Chiappori and Bernard Salanié 2000). The highly influential Rand Health Insurance Experiment (HIE) short-circuited the self-selection process by randomly assigning different insurance policies to households (Joseph P. Newhouse 1993). The generosity of the assigned policies had a significant impact on the consumption of healthcare, although the price elasticity of -0.2 was on the low end of previous estimates, consistent with elimination of upward bias from adverse selection. Nevertheless, health *outcomes* did not vary with these insurance and induced healthcare differences. The lack of a response in health outcomes has been interpreted as a finding of significant moral hazard from health insurance (Austan Goolsbee 2005, Baicker and Chandra 2008).[2]

Despite the fact that childbirth is the most common reason for hospitalization in the US, the extent of moral hazard in postnatal/postpartum hospitalizations is unknown. In the Rand HIE, demand for healthcare by children appeared less sensitive to the insurance plan assigned: pediatric hospitalizations were "insensitive to insurance plan" (Peter Zweifel and Willard G. Manning 2000). Taken at face value, moral hazard may be absent where exigent care of children is concerned. Similarly, Robert J. Haggerty (1985) concluded from the HIE evidence: "Clearly there is little cost-saving to be found by imposing copayment for hospital care for children."

However, the HIE considered pediatric health care for ages 0 to 18; it did not evaluate postnatal and postpartum outcomes following childbirth (Newhouse 1993). Thus, the HIE offers no direct evidence on a substantial portion of the health insurance market: postpartum and postnatal hospitalizations account for 15 percent of total private insurance charges and 27 percent of Medicaid charges (Sakala and Corry 2008). Because the HIE was conducted thirty-five years ago, it predates sweeping changes in the healthcare and health insurance industry that attempted to reduce costs (and presumably mitigate moral hazard). While recent empirical work has analyzed postpartum stay laws enacted during the 1990s, it has not directly addressed the issue of moral hazard. An identifying assumption of the interrupted time series designs utilized by recent studies is that patient characteristics did not happen to depart from trend at the same time of the law, which we discuss below (Sections II II.C and II II.D).[3] In any event, the findings of health benefits from the laws (see discussion of William N. Evans and Craig L. Garthwaite (2009) below) would imply that moral hazard alone cannot explain longer stay lengths.

[2]That said, the Rand Health Insurance Experiment did not randomly assign anyone to receive *no* health insurance (Helen Levy and David Meltzer 2008), nor does our identification strategy.

[3]Although *a priori* we might not expect significant adverse selection in hospital births (on the extensive margin), estimates of moral hazard are as credible as the identifying assumptions and "are more convincing when there is a reference sample for which [insurance] incentives did not change" (Chiappori and Salanié 2000), generally absent in the analysis of law changes.

### B. Postpartum Hospital Care in the US

According to a 2001 report to Congress, "Current scientific knowledge does not provide conclusive evidence about ideal delivery length of stay...."[4] Nevertheless, the average stay length for childbirth has changed dramatically over time, and prompted a flurry of legislative mandates entitling newborns to a minimum stay length.

Between 1970 and 1995, the average length of stay for a vaginal delivery fell from 3.9 to 1.7 days. For cesarean births, the corresponding numbers are 7.8 to 3.6 days (Sally C. Curtin and Lola Jean Kozak 1998). In the mid-1990s, this decrease was halted, and a slight increase in average stays followed.[5] For short-stays, the pattern is much more stark. Figure 1A plots the share of vaginal births with stays under 2 days from 1970-2004. There is a doubling of these "early discharges" from 1990 to 1995, followed by a sharp and sustained reduction.

The practice of "drive-through delivery" formed a rallying point against cost-saving measures imposed by third-party payers (David A. Hyman 2001). In 1995, the official journal of the American Academy of Pediatrics ran a commentary entitled: "Early discharge, in the end: maternal abuse, child neglect, and physician harassment." It warned that inadequate screening of newborns was the "most dangerous and potentially long-term effect of early discharge," especially the "re-emergence" of jaundice as a cause of hospital re-admission (Seymour Charles and Barry Prystowsky 1995). Postpartum stays afford the opportunity to monitor the newborn, preform screening tests for problems, and instruct parents on infant care. In 2005, the American Academy of Pediatrics published criteria for the discharge of newborns, noting it is"unlikely that fulfillment of these criteria and conditions can be accomplished in 48 hours" even for healthy newborns (A.A.P. 2004).

The decrease in the frequency of early discharges is largely the result of legislative responses to this outcry. The legislation invoked longer average stay lengths observed in Europe and Japan as motivation (Nancy Kassebaum 1996). Between 1995 and 1998, 42 states passed laws requiring insurers to cover a minimum postpartum stay (Evans, Garthwaite and Wei 2008). In January 1998, the federal government followed suit, mandating a minimum stay of two days.

On August 26, 1997 the California Newborns' and Mothers' Health Act came into effect in California that entitled newborns to 48 hours of inpatient care, as well as coverage for early follow-up care if the newborn is discharged early. Figure1B shows that the fraction of vaginal births in California that had an early discharge increased to 75% prior to the law change, and decreased from October 1997 to February 1998 to 50%.[6,7]

[4]As summarized by Antoinette Parisi Eaton (2001).

[5]NCHS (2000) and various additional years.

[6]The spikes seen occur on December 23 and 24 each year when short stays are particularly common. Such short stays (and potential under-staffing) do not appear to affect health outcomes in our data.

[7]Using the California data described below, but for children born at all times of the day and discharged from the hospital, we find that relatively short stays of zero or one nights are not associated with an increase in readmissions (we find a 2% reduction in readmissions within 28 days of discharge), although we do find a 2% increase in one-year mortality. We hesitate to focus on these results because the mortality

### *C. Analysis of Minimum Stay Laws*

Estimates of the return to stay length may be biased if patients in worse underlying health spend more time in the hospital. Previous work has attempted to estimate the causal effect of stay length by using passage of state laws as natural experiments, which can eliminate this bias (JM Madden, SB Soumerai, TA Lieu, KD Mandl, F Zhang and D Ross-Degnan 2002, Ellen Meara, Uma R. Kotagal, Harry D. Atherton and Tracy A. Lieu 2004, Ashlesha Datar and Neeraj Sood 2006, Evans, Garthwaite and Wei 2008, Evans and Garthwaite 2009). While these studies have consistently found that those born shortly after the laws stayed longer in the hospital, evidence on health impacts has been mixed. According to Evans and Garthwaite (2009), the "effect of these laws on health outcomes, however, has been less clear." Some evidence suggests improved outcomes for those covered by Medicaid and for more complicated births, e.g., Evans, Garthwaite and Wei (2008). To assess whether the laws encouraged moral hazard requires the assumption that the pre-law trend provides the correct counterfactual, as discussed below (Section II.D).

Evans and Garthwaite (2009) focus on the potential for heterogeneity in the health impacts of stay length. In the spirit of Marianne P. Bitler, Jonah B. Gelbach and Hilary W. Hoynes (2006), a small average effect from a broad-based treatment may hide large treatment effects for important sub-populations. Evans and Garthwaite (2009) explore this issue systematically in the context of postpartum stays, finding that those with a high *a priori* propensity for longer stay lengths received substantial health benefits from the laws. If "smart laws" that targeted these sub-populations could be adopted, then greater health benefits might be realized. As such, our identification strategy speaks more to the denser portion of the propensity distribution where health risks are relatively low. For this sub-population, our results conform with Evans and Garthwaite (2009)'s finding that the "vast majority of individuals...received no benefit from the increased length of stay generated by the law."

Several recent papers have emphasized the persistence of health complications related to early discharge despite implementation of laws mandating coverage of minimum hospital stays. This was found especially for deliveries paid for by Medicaid or where the mother was Hispanic. In an attempt to buffer the presumed effects of early discharge that occurred despite the legislation, coverage for early follow-up visits were mandated in these cases. Previous evidence suggests that such a mandate is unlikely to affect the take up of such services, however. Meara et al. (2004) found no effect of Ohio minimum stay-early follow-up visit legislation on the take up of early follow-up care among the state's Medicaid population, although the timing of the follow-up care after birth was slightly delayed due to the longer stays in the hospital. Meanwhile, Alison A. Galbraith, Susan A. Egerter, Kristen S. Marchi, Gilberto Chavez and Paula A. Braveman (2003) found no difference in early follow-up care for newborns who were discharged early versus those who were not. These results suggest that our comparisons of effects before and

result is not measured precisely, the inconsistency in the sign, and the problem that short stays could be an effect of the mortality outcome rather than a cause.

after the California law are unlikely to be affected by differences in early follow-up care.

### D. The "After Midnight" Approach

Our Innovation. — There are four main innovations in this paper compared to the earlier law-change analyses.

First, we demonstrate that simple billing rules can have large effects on the treatment of patients, in keeping with previous work that found large effects of reimbursement rules (Leemore Dafny 2005). To our knowledge, we are the first to exploit administrative rules governing stay length to identify utilization and health effects.

Second, because timing vaginal births to occur just after midnight is difficult, our identification strategy plausibly eliminates adverse selection.[8] We present empirical evidence below suggesting that such strategic behavior is indeed absent. In this respect, we have replicated a key feature of the Rand HIE that enabled isolation of moral hazard, while yielding contemporary evidence for a large patient population the Rand HIE overlooked.

Third, we utilize the introduction of insurance mandates in a new way. Interacting the midnight discontinuity in length of stay with the law change allows us to consider variation in stay length from two baselines: increasing length of stay principally from zero to one additional nights in hospital versus increasing from one to two additional nights. This allows us to assess whether moral hazard is less evident when hospital stays become shorter and to investigate diminishing returns to postpartum care.

Last, the identifying assumption underlying the interrupted time-series approach is that the trend in length of stay and outcomes prior to the law change describes the counterfactual length of stay and outcomes after the law change. If other interventions happened at the same time as the law change, then the before-after effects may reflect both the law change and the complementary interventions. For example, the policy changes were accompanied by warnings of the increase in jaundice in the early 1990s (Charles and Prystowsky 1995), and policies that aimed to reduce jaundice infections may have happened at the same time discharge policies were amended to comply with the laws. This is consistent with an increase in jaundice readmissions found in California after the law extended stays (Datar and Sood 2006). In addition, our analysis of California data finds that the daily number of newborns in the hospital increased by approximately 10% following the law change. The effect on newborns will depend on the way hospitals responded to this anticipated increase with additional staff either at the time of the reform or over time.[9] In contrast, our analysis considers increases in length of stay when these short-run costs of additional volume are not present and thus arguably provides a better measure of the long-run or "steady-state" effects of stay-length after the initial adjustment period. Lastly, if hospitals began responding to the popular outcry prior to

[8] Furthermore, knowledge of this feature of hospital billing rule is required.

[9] Treatment choices do not appear to change at the time of the law (August 26, 1997), although there is suggestive evidence that cesarean section rates increased beginning in January 1998; they were close to 21% for 3 years prior to the federal mandate, increased to 22% in March 1998, and continued to increase to 27% by the end of the sample period (2002).

the law's enactment, the change in outcomes before and after the law may include the effect of the change in length of stay as well as the response to the law. Indeed, Figure 1B indicates that the upward trend in early discharges was halted in 1995 and that early discharge became less common in 1996, approximately one year *before* California's law was enacted.

Time of Birth and Stay Length. — Essentially, all births in the US are insured, either by private coverage or Medicaid (Rebecca Russell, Nancy Green, Claudia Steiner, Susan Meikle, Jennifer Howse, Karalee Poschman, Todd Dias, Lisa Potetz, Michael Davidoff, Karla Damus and Joann Petrini 2007), and postpartum care is generally reimbursed for a predetermined number of days in care, with longer stays requiring physician approval. These days are counted as the number of midnights. For example, the Medicaid program in California, known as "Medi-Cal", issues the following guidelines regarding prior authorization for obstetric admissions:

> Welfare and Institutions Code, Section 14132.42, mandates that a minimum of 48 hours of inpatient hospital care following a normal vaginal delivery and 96 hours following a delivery by cesarean section are reimbursable without prior authorization. For [Treatment Authorization Requests (TARs)] and claims processing purposes, it is necessary to use calendar days instead of hours to implement these requirements. Therefore, a maximum of two consecutive days following a vaginal delivery or four consecutive days following a delivery by cesarean section is reimbursable, without a Treatment Authorization Request. The post-delivery TAR-free period begins at midnight after the mother delivers. (Medi-Cal 2007)

For a birth occurring at 11:59 p.m., the number of reimbursable days in care begins one minute later, whereas births just after midnight are afforded nearly a full 24-hour period before the number of reimbursable days in care begin to be counted.

The minimum coverage is effectively extended by one night for those born after midnight. In a setting of costless bargaining, this increase in the property right to the hospital room might not be expected to have any impact on stay length (Ronald H. Coase 1960). If the insurer wanted to bargain for shorter stay lengths, however, they would face political costs. Before the minimum stay law, policy and practice entitled newborns to only 1 day in the hospital. After the law change, insurers are limited to the use of early follow-up care as an incentive for an early discharge, although previous evidence suggests that early follow-up visits are not determined by length of stay.

An underlying assumption of our approach is that for uncomplicated deliveries, the minute of birth around midnight is effectively random. There are ways of increasing or decreasing the speed of the delivery, however, and physicians may have some discretion when recording the exact time of birth. In terms of incentives, the patient and the insurer would prefer a post-midnight birth, as billing for time in the hospital would not begin until one day later. The hospital would likely prefer a pre-midnight birth to begin billing for the time in care sooner. The cost to the hospital of supervising the child would also decline if the birth occurred before midnight and the discharge time was sooner. To the extent that the physician's interest is aligned with the patient (or the insurer), there may be a tendency to record births after midnight, whereas if the incentives are aligned with the hospital, the tendency would be to record births just before midnight. We will consider the frequency of births around midnight for all births, observable characteristics of the newborns around the midnight threshold, as well as births that occur in Kaiser Hospitals–hospitals where the insurer owns the hospital.

Previous work most closely related to ours is Jesse D. Malkin, Michael S. Broder and Emmett Keeler (2000), who used 4-hour categories in the time of birth (as well as cesarean section) as instruments for length of stay. However, births scheduled for "business hours" may have different baseline characteristics compared to births later in the day (reflecting, e.g., the scheduling of high-risk deliveries). Indeed, we find that baseline health and demographic characteristics are substantially different during "business hours" compared to the overnight hours in California.[10] For this reason, we will restrict comparisons to births just before and after midnight when observable characteristics are similar across the newborns.

## III. Data

### A. Description

Our data include the universe of live births in California from 1991-2002, some 6.6 million records. We focus on the 270,000 births occurring between 11 p.m. and 1 a.m. The California Office of Statewide Health Planning and Development created a research database that includes hospital discharge records linked to birth and death certificate records. For a given birth, discharge data are available nine months prior to delivery so as to capture the course of antepartum, inpatient care. In addition, hospital admissions up to one year after delivery are matched to the birth record for both mothers and infants. Death certificate data provide a measure of mortality, while the birth certificate includes a wealth of information about the parents and the circumstances of the birth itself.

The hospital discharge data include the patient's age, procedure and diagnosis codes, primary payer, day of the week, hospital ownership information, and admission and discharge date. Beginning in 1995, whether the birth was scheduled or unscheduled is reported as well.

[10]We replicated the results in Malkin, Broder and Keeler (2000) using our California data, finding that infants born between 8 a.m. and 12 noon have a length of stay that is 7 hours longer than those born between 8 p.m. and 12 midnight (discharge time was assumed to be 5 p.m. for those with same day discharges and 1 p.m. for those who stay at least one night in the hospital as in Malkin, Broder and Keeler (2000), who cite a survey conducted by the Rand Institute (D.S. Gifford, S.C. Morton, M. Fiske, J. Keesey, E.B. Keeler and K.L. Kahn 2000)). Like Malkin, Broder and Keeler (2000), we also find statistically significant decreases in readmissions with an increase in length of stay using these 4-hour blocks as instruments. The groups differ substantially, however, with the largest differences found for mother's first birth (34% for morning births vs. 44% for 8-midnight births), induced labor (only 6% for 8 a.m. to noon vs. 12% for those occurring later in the evening), and cesarean section (30% vs. 14% for those born between 8 a.m. and noon vs. midnight to 4 a.m.).

Over 98% of the births in the hospital discharge data, and approximately 96% of all births in the vital statistics records, were successfully linked together. The birth-certificate data report pregnancy and birth characteristics that are not available in hospital discharge data, including pregnancy and birth complications, birth weight and gestational age, as well as parents' age, education, and place of birth.[11] While Race/Ethnicity of the newborn will be considered, Hispanic births were no longer separately identified in our data after 1995. The mother's day of admission and the precise time of birth, required to implement our design, are recorded. A list of the variables used can be found in Appendix Table A1.

Length of stay is reported in the data as the discharge date minus the admission date: the total number of midnights in care. The admission date is the date of birth for the newborn.[12] The main measure of resource usage in the analysis will be the number of *additional* midnights in care: the number of midnights in the hospital not counting the initial one that defines the threshold. That is, our treatment measure for those born after midnight is the usual one; for those born before midnight, we subtract one from the usual length of stay so as to remove the mechanical midnight that is not related to the true length of stay in the hospital.[13]

Discharge time is not recorded, so our length of stay measure is in days. While we would prefer a more detailed measurement of stay length, in the spirit of "not biting the hand that feeds," we note that its absence is presumably what makes our discontinuity design possible. It is possible that infants born prior to midnight may stay later on the day of discharge compared to newborns who stay an additional night in the hospital due to the post-midnight birth. For example, Malkin, Broder and Keeler (2000) report that in a survey of Los Angeles and Iowa hospitals, the median time of discharge for infants with a same-day discharge is 5 p.m., whereas it is 1 p.m. for those with at least one night in the hospital. In total, this difference is bounded by the time a physician can see the patients in the morning and the end of business hours, or approximately 8 hours. As such, the results below consider the effect of having an extra night in the hospital, representing an increase in time for supervision, screening, and education of between 16 and 24 hours.

### B. *Analysis Sample*

The main analysis will consider vaginal births within 20 minutes of midnight; births by cesarean section will be examined separately in part because the time of birth is more

[11]When gestational age is used as a control, it is measured as the number of days not including the midnight that defines the threshold.

[12]Despite the potential incentive on behalf of the patient or the insurer to record the admission date of a pre-midnight birth as after midnight, this does not seem to occur. Indeed, there are 160 twin pairs where birth times straddle the midnight cutoff: the accounting length of stay is greater for the infant born before midnight (by 1 day), and charges associated with that infant are $1024 higher, on average.

[13]Total hospital charges are available in the data, but those born just before midnight are billed for a night in care almost immediately and have slightly higher accounting charges despite spending less time in the hospital.

likely to be a choice variable and the insurance coverage differs for these births. When constructing the sample, only births in hospitals that are matched to the vital statistics data will be used, (a loss of approximately 5% of births in California). Second, to focus on patients where the births are most likely to have random variation with regard to the timing of the birth, unscheduled births are considered after 1995, which excludes 16% of the remaining births. Scheduled births will be considered separately as well. The 1% of remaining newborns with lengths of stay of more than 28 days are excluded. These outliers are less likely to be affected by the accounting rules and may skew the mean differences before and after midnight. Another 4% of the remaining data have missing covariate information; results will be shown with and without these births.[14]

Figures of raw means by minute of birth and local linear regressions will be shown using data for every minute of the day. Models are also estimated with two samples: a 40-minute sample and a 2-hour sample, which include data within those windows around midnight. A spike in births on the hour typically reflects births that occurred at any time during that hour. Our analysis aims to compare newborns born just before or after midnight, however, so the main analysis excludes the births from 11:56 p.m. to 12:04 a.m. This leaves a "donut hole" in the OLS and probit estimation, but we take these points as the first reliable points before and after midnight to accurately represent the boundary. Results are similar, and the interpretation is identical, when these minutes were included. The 40-minute sample, then, includes births fewer than 20 minutes from 11:55 p.m. or 12:05 a.m. Figures (including additional figures in the online appendix) will show data for every minute, however.

The final cut of the data considers infants born before and after the law change in California. The law came into effect on August 26, 1997. Births from January 1, 1991 to July 31, 1997 will be used to estimate the models before the law change, and births from September 1, 1997 to December 31, 2002 will be considered for post-law-change births. The resulting 40-minute samples include over 60,000 observations prior to the law change and over 35,000 observations after the law change.

## IV. Compliers and Estimation

Our main analysis first compares length of stay across the midnight threshold to establish whether the change in minimum insurance coverage affected consumption. We then compare health outcome measures to test whether any change in consumption led to health improvements. Specifically, we are estimating effects for those who stay an extra night at the hospital because the infant is entitled to an extra day without charge due to the time of birth around midnight. Presumably, those induced to stay longer are not a random draw from the population. In particular, infants born around midnight likely

[14]In the 40-minute sample pooled over all of the years, post-midnight births are less likely to result from a c-section (17.7% vs. 19%) and less likely to have missing hospital data (3.7% vs. 5.2%). If missing data are more likely in the case of death, then the results would be biased toward finding reductions in mortality with longer stays associated with post-midnight births. Differences in home births, scheduled births, and missing covariates are small and not statistically significant across the groups.

differ from those planned to be born earlier in the day. For this reason, *a priori* our estimates may be expected to differ from those previously estimated using the minimum stay-length mandates. Further, infants induced to stay an extra night as a result of the time of birth around midnight exclude families who wish to leave the hospital soon after birth or newborns with serious health complications who will stay in the hospital much longer than two nights regardless (these groups are also excluded from the time-series estimates that consider the law changes).

### A. *Description of Compliers*

The local average treatment effect is the average treatment effect for "compliers": those who are induced to have a longer stay as a result of the post-midnight birth.[15] In contrast, "always takers" or "never takers" have stay lengths that are unaffected by the minute of birth. It is not possible to identify the compliers, but it is possible to describe their observable characteristics (Alberto Abadie 2003).[16]

First, we define the binary variable $Z$:

$$Z = \begin{cases} 0 & \text{born just before midnight} \\ 1 & \text{born just after midnight} \end{cases}$$

Next, define the binary variable, $D$, to be an indicator for a long stay (e.g. more than 1 night prior to the law change and more than 2 nights after the law change):

$$D = \begin{cases} 0 & \text{short stay} \\ 1 & \text{long stay} \end{cases}$$

Also define $D_Z$ as the value $D$ would take if $Z$ were either 0 or 1. $Z$ is again assumed to be independent of $D$. Compliers in this context are such that $D_1 = 1$ and $D_0 = 0$.

Consider $E(X|D_1 = 1)$, which represents the characteristics of those with long stays who are born after midnight. This group is comprised of always takers and compliers as shown in the following two terms:

$$E(X|D_1 = 1)$$
$$= E(X|D_1 = 1, D_0 = 1)P(D_0 = 1|D_1 = 1)$$

[15] We highlight the binary nature of the compliers and always takers to more easily demonstrate the difference in the average observable characteristics below. This could be relaxed, however. John DiNardo and David S. Lee (2010) describe a more general interpretation, where the instrumental-variable estimate "can be viewed as the *weighted* average treatment effect for the entire population where the weights are proportional to the increase in the probability of treatment caused by the instrument."

[16] Abadie (2003) showed that characteristics of compliers can be described using his kappa weighting scheme. It is also known that for binary characteristics and a binary instrument, the relative likelihood that an individual in a particular group is a complier is the ratio of the first-stage coefficient on the instrument estimated on that group's subsample to the first-stage coefficient for the full sample.

$$+ E(X|D_1 = 1, D_0 = 0)P(D_0 = 0|D_1 = 1). \tag{1}$$

Always takers can be described by the characteristics of individuals who were born before midnight ($Z = 0$) yet have longer stays ($D = 1$). That is, $E(X|D_1 = 1, D_0 = 1) = E(X|D_0 = 1)$, by the monotonicity condition ($D_1 - D_0 \geq 0$).[17]

By the independence of $Z$, the proportion of the population that represents always takers, $\pi_A$, is $P(D_0 = 1)$, and never takers, $\pi_N$, is $P(D_1 = 0)$. The proportion of the population who are compliers is then $\pi_C = 1 - \pi_A - \pi_N$, as defiers are assumed away by the monotonicity condition. Among the group born after midnight, the fraction with longer stays, $P(D_1 = 1)$, is $\pi_C + \pi_A$. Meanwhile, $P(D_0 = 1|D_1 = 1) = \pi_A/(\pi_A + \pi_C)$, and $P(D_0 = 0|D_1 = 1) = \pi_C/(\pi_A + \pi_C)$.

(1) can then be re-arranged, and the expected characteristics of the compliers can be written as:

$$E(X|D_1 = 1, D_0 = 0)$$

$$= \frac{\pi_C + \pi_A}{\pi_C}\left[E(X|D_1 = 1) - \frac{\pi_A}{\pi_C + \pi_A}E(X|D_1 = 1, D_0 = 1)\right]$$

$$= \frac{\pi_C + \pi_A}{\pi_C}\left[E(X|D = 1, Z = 1) - \frac{\pi_A}{\pi_C + \pi_A}E(X|D = 1, Z = 0)\right]. \tag{2}$$

Each of the terms in (2) can be estimated using sample means.

## *B. Estimation*

We begin by examining the first stage – the relationship between the time of birth and the length of stay – and then proceed to the reduced form – the relationship between the time of birth and the health outcomes. Local linear regressions before and after midnight are estimated using a triangle kernel (Ming-Yen Cheng, Jianqing Fan and JS Marron 1997). Asymptotic standard errors are also reported (Jack Porter 2003).[18]

In addition, we estimate OLS models that include covariates and linear trends (in the minutes from midnight) that are allowed to vary before and after midnight. This is a simple local linear estimator with a rectangle kernel where the weights do not decay as the distance from midnight increases (Guido Imbens and Thomas Lemieux 2007). For outcomes $Y$ (including length of stay, readmissions, and mortality), the models for infant

[17] For an expanded discussion of our analysis in terms of a potential outcomes framework, the conditions for *LATE* estimation, and the multi-valued treatment, please see the working paper version of our paper (Douglas Almond and Joseph J. Doyle, Jr. 2008). That discussion follows Joshua D. Angrist and Guido W. Imbens (1995) closely.

[18] Thanks go to Doug Miller for providing code from Jens Ludwig and Douglas Miller (2007)

$i$ born at minute $t$ from a midnight ($t$=0) cutoff are as follows:

$$Y_{it} = \beta_0 + \beta_1 \mathtt{1}(t \geq 0) + \beta_2 \mathtt{1}(t \geq 0) * t + \beta_3 \mathtt{1}(t < 0) * t + \beta_4 X_i + \varepsilon_{it}, \tag{3}$$

where X is a vector of observable birth characteristics.[19]

This basic regression discontinuity model is estimated using Ordinary Least Squares for length of stay. Additional tests of robustness with regard to the estimation are also reported. For the binary outcomes of readmissions and mortality, probit models are estimated and marginal effects at the mean of the control variables are reported. Heteroskedasticity-robust standard errors are reported for these models.

Choice of Bandwidth. — Across the outcomes, bandwidths of close to 10 minutes were found to minimize the sum of squared errors between the local linear estimator and a fourth-degree polynomial model estimated within two hours of midnight. Local estimation at a boundary is generally thought to require a larger bandwidth compared to interior points, and we applied a rule of thumb used in density estimation at a boundary of two times the cross-validation bandwidth (S. Zhang and R.J. Karunamuni 1998): a 20-minute pilot bandwidth in this case. Shorter and longer bandwidths were used as well with similar results. Relatively wider bandwidths appear appropriate given that we find a sustained increase in length of stay following midnight, and if this length of stay affects outcomes, a shift in the readmission or mortality rate should be sustained throughout the first hour as well. The tradeoff is that infants born far from midnight are more likely to differ from one another. We do not expect infants born at 11:00 p.m. to be so very different from those born at 1:00 a.m.

Average Causal Response. — The length of stay treatment takes on multiple values, and the average causal response of the health outcome measures to changes in the length of stay is the weighted average of effects from increases in the number of nights from one to two; from two to three; and so on. Passage of California Newborns' and Mothers' Act in 1997 changed these weights dramatically. Prior to the law change, the midnight threshold increased the number of additional midnights primarily from zero to one (or one to two total "nights" in the hospital). After the law change, stay length increased from one to two additional midnights due to the midnight rule (see Section V.C for details). By estimating results separately for births before and after the law change, the potential for diminishing returns to length of stay can be examined.

[19]In practice, the analysis samples exclude births within 10 minutes of midnight. $t = -1$ just prior to the cutoff (11:55 p.m.) and $t = 0$ at the cutoff (12:05 a.m.).

## V. Results

### A. Frequency of Births Around Midnight

Do physicians systematically misreport time of birth around the midnight threshold? As noted above, physicians may have an incentive to record births as occurring earlier or later than midnight. Figure 2 provides a visual check on this behavior (Justin McCrary 2008), plotting the number of births by minute of the day. Births are more frequent during "business hours" of 7:00 a.m. to 5:00 p.m. The frequency declines until midnight and remains fairly stable until around 7:00 a.m. The time of birth is more likely to be reported on the even hour and additionally at times ending in 0 or 5, due to rounding.

Much of the analysis will focus on births between 11 p.m. and 1 a.m. and Figure 2B shows roughly 2500 births are recorded each minute, while 3000 births are found at the 5 minute marks. The largest spikes occur at 11:00 p.m. and 12:01 a.m. 12:00 midnight – the only time that uses the number 24 as the hour – has fewer observations (N=734), possibly due to physicians making clear that the birth occurred the following day. The spike at 12:01 is similar to the spike at 11:00 p.m., though it is slightly larger than the spike at 1:00 a.m. These spikes likely reflect births that occurred at any time during that hour–one reason to exclude births at 12:01 a.m., as most of these births occurred later in the hour. The number of observations is similar in the hour before and after midnight. 153,180 births are recorded from 11:00 p.m. to 11:59 p.m. and 147,113 are reported from 12:00-12:59, or 4% more prior to midnight. This is similar to the 3% difference between the 10:00 p.m. hour and the 11:00 p.m. hour, though smaller than the midnight versus 1:00 a.m. comparison of less than 0.5 percent. In the 40-minutes around midnight from 11:36 to 12:24 (excluding the 10 minutes around midnight), the fraction born after midnight is 0.493 before the law change, and 0.497 after the law change, with the second proportion not statistically significantly different from 0.5 (p-value = 0.31).[20]

### B. Comparison of Covariates

Another test to consider whether there is selection into the post-midnight sample compares observable characteristics across the threshold. Table 1 reports means of selected covariates for the 40 minute sample. This is for the pooled sample of births from 1991-2002. Means are reported for each sub-period in the appendix.

Mean differences are small across the covariates. For example, the average maternal age of 26.8 is identical across the two groups.[21] Some small differences are found. 20.4%

[20] The fraction born after midnight is close to 0.5 across hospitals and dates. One date of interest is December 31, when a pre-midnight birth is subject to a tax deduction (Stacy Dickert-Conlin and Amitabh Chandra 1999). We find that births tend to be pushed toward January 1, reflecting possible benefits of being a "baby new year." In the two-hours around midnight, the highest proportion found to be born "after midnight" was on January 1, 2000 (72%), followed by January 1, 2001 (67%). January 1 1995, 1997, and 1998 are also in the top 100 dates in terms of the highest fraction of post-midnight births.

[21] No difference was found for particular age groups as well.

of women with births after midnight had fewer than 9 prenatal visits compared to 19.9% for births before midnight, whereas the means for 9-15 prenatal visits are 68.9% and 69.5%, respectively. Educational characteristics, are nearly identical, including missing data for fathers which can be seen as an indicator of single-parent births. An indicator for the mother's first birth is slightly smaller for births after midnight (39.4% vs. 40%), largely due to a difference in the post-1997 period. Those born after midnight are slightly less likely to be white (53.7% vs. 54.4%), although the differences are not statistically significant in the two sub-periods. The use of forceps or vacuum to speed the delivery is slightly less likely after midnight (9.5% vs. 9.9%), although other measures of labor being stimulated or induced are not different, especially in the pre-1997 time period. Births prior to midnight are slightly higher in for-profit hospitals (14.2% vs. 13.6%). Out of the 56 characteristics listed in Appendix Table A1, 6 have statistically significant differences (at the 5% level) in the pre-law period, and 5 have statistically significant differences in the post-law period. Most of these differences do not appear economically significant (often indistinguishable out to 2 significant digits), despite the statistical significance due to the large sample size. When the post-midnight indicator is regressed on the observable characteristics, the F-test fails to reject that all of the coefficients are zero (F-statistics of 1.13 and 1.05 for the two time periods ; p-values of 0.17 and 0.35).[22]

In addition to the mean differences, it is also possible to consider the characteristics for each minute close to midnight. Appendix Figure A1 reports local linear regression estimates, as well as means for each minute, for an indicator that the newborn is classified as low birthweight ($\leq$ 2500 grams), as well as a predicted 28-day readmission rate based on a probit model with full controls. The Figures show some variation over the day, but little change is seen at midnight and no change is consistently seen in the two time periods. In sum, it appears that births just after midnight are similar to those just before midnight.

### *C. Length of Stay*

Length of stay is measured as the number of midnights in care. If the minute of birth were unrelated to the timing of discharge, newborns with a time of birth just prior to midnight should have a length of stay recorded as one midnight longer than newborns born just after midnight, by definition.[23]

To consider the number of nights in the hospital rather than the accounting length of stay, Figure 3 shows the number of additional midnights in care by minute of the day. Additional midnights are simply the raw length of stay minus 1 for births prior to midnight (and after noon). Panel A describes the relationship prior to the 1997 law

[22]These tests exclude 1% of the observations with "missing admission day of the week for the mother": a variable that is associated with post-midnight births. Results are identical when these cases were excluded from the main analysis.

[23]Appendix Figure A2 reports the the raw accounting data, and a magnified view with a smaller bandwidth is whon in Figure A3. As expected, the accounting length of stay–the number of midnights in care–is higher for those born just before midnight, although the difference is significantly less than one. (Were just the mechanical effect at play, this difference should be close to unity.)

change, and Panel B considers births after the law came into effect. Prior to the law change, the average length of stay is close to 1.5 days for births at noon, increasing to 2 days by midnight followed by an increase of 0.28 additional nights in the hospital for those born just after midnight. The upward slope in this and later length-of-stay Figures is due to the mechanical relationship that as a birth approaches midnight, they are more likely to achieve a greater number of midnights in the hospital in the future. As long as this relationship is smooth, however, we can control for this trend in the main analysis, and these trends are much less of an issue in the outcomes of interest. After the law change, Panel B shows a similar picture, although the average length of stay is shifted upwards by roughly 0.35 days in care. The increase in stay length after-midnight remains approximately the same, albeit from a higher baseline. That is, approximately one in four births after midnight experience an additional night in the hospital compared to what they would have spent in the hospital had they been born minutes earlier.

Appendix Figure A4 shows the proportion of births with greater than 1, or greater than 2, additional midnights. Before the law change, the local-linear regressions show that 57% of those born just before midnight stay at least one more night in the hospital compared to 72% of those born after midnight. After the law change, 83% of those born just before midnight stay at least one extra night, with a smaller increase to 90% for those born after midnight. By comparison, the proportion of newborns staying at least two more nights increases from 11% to 17% before the law change and doubles from 16% to 32% after the law change. Once the newborn has stayed two nights, the post-midnight birth has a smaller effect.

To summarize the data shown in Figure 3 (and Appendix Figures A2-A4) and consider one of the estimation samples, Table 2 reports means for the 40-minute sample used in the estimation below.[24] The increase in length of stay after the law change is evident: the average length of stay increases from 1.99 to 2.29 for those born before midnight, and from 1.23 to 1.58 for those born after midnight.

When births before and after midnight are compared, the average number of additional midnights increases from 0.99 to 1.23 before the law change and from 1.29 to 1.58 after the law change.[25] During the pre-law period, then, roughly 1 in 4 infants born after midnight are afforded an extra night in the hospital due to the minute of birth; after the law change the additional night is found for slightly more than 1 in 4 infants.[26] This change is remarkably similar to the change in average length of stay following the law change, i.e. leaving aside the midnight discontinuity. This suggests that our use of the

[24]The sample mimics the nonparametric estimates that include births less than 20 minutes: births from -19 $\leq t \leq$ 19. With the threshold at minute $t = 0$, this leads to 1 more minute post midnight than before midnight: 11:37 p.m. to 11:55 p.m. and 12:05 a.m. to 12:24 a.m. This results in slightly more observations in the post-midnight columns.

[25]When the length of stay for those born in the 11:00 hour was recorded as zero (0.6%), this likely reflects measurement error and the number of additional midnights was set to zero.

[26]In the pre-law period, 1 in 4 newborns "taking up" the additional reimbursable night made possible by post-midnight birth (and thereby spending .99+1 = 1.99 additional nights in the hospital) would increase the mean stay length from .99 to 1.23. Similarly for the post-law period, a take-up rate of .29 on 1.29+1=2.29 additional nights would yield the observed 1.58 additional nights for the post-midnight births.

midnight accounting rule mimics the law's mandate that entitled newborns to 48-hour minimum stays when insurance providers routinely reimbursed only 24-hours in care.

The next three rows of Table 2 report the proportion of newborns who stay at least one additional midnight, at least two, and at least three additional midnights. For each category, these measures are larger for those born after midnight, which is consistent with the monotonicity condition. As reflected in Appendix Figures A3A and A3B, before the law change, the increase in the number of additional midnights is most pronounced between 0 and 1, whereas the jump after the law change is seen primarily for newborns staying 2 additional midnights as opposed to 1. In terms of the local average treatment effect weights described in Section IV, prior to the law change, the weight on treatment increases from zero to one additional night is 73%, while after the law change the weight on treatment increases from one to two additional nights is 65%.[27] This provides further evidence that it is the discontinuity in insurance coverage that is leading to the variation in hospital care, as opposed to other rules of thumb used by healthcare providers.

Table 1 suggests that controlling for observable characteristics should have little effect on the results, and this is confirmed in Table 3. Table 3 shows the results for the first-stage relationship between additional midnights in care and an indicator that the birth occurred after midnight. Column 1 reports the difference in local linear regression estimates just before and after the midnight boundary. A bandwidth of 20 minutes was used and the models were separately estimated above and below the threshold. The estimates are similar before and after the law change: 0.27 vs. 0.26, although they represent increases from different baselines. The average number of additional midnights for births prior to midnight is 1.00 before the law and 1.30 afterwards. The estimates are highly significant, with standard errors of close to 0.04.

Columns (2)-(5) are estimated by OLS with controls for linear trends in minutes from the midnight cutoff, trends that are allowed to vary before and after midnight, as described above. Note that minutes from the cutoff are positive after midnight and negative before midnight. Using the same 40-minute window as the local linear results, but with a sample that includes nonmissing covariates, the results are similar: 0.29 before and 0.24 after the law change. A full set of birth characteristics listed in Appendix Table A1, as well as individual indicators for mother's age, father's age, year of birth, and month of birth are included in models reported in Column (3).[28] The main coefficients are similar, however (0.27 and 0.23). The robust standard errors are similar to the asymptotic standard errors calculated in Column (1), although they are slightly smaller. The estimates based on the two-hour sample are 0.22 and 0.25 before and after the law change, respectively, regardless of the use of controls, or close to 20% of the pre-midnight means.

[27] Table 1 shows that the change in the proportion of infants born close to midnight who stay at least three additional midnights is smaller (on the order of 1-2 percentage points). Excluding these differences in stays of greater than 2 nights, the weights are proportional to the differences in proportions of children staying at least 1 vs. at least 2 additional midnights. Prior to the law change, the weight on stays increasing from zero to one additional night is (73-57)/((73-57)+(17-11))=73%. After the law change, the weight on stays increasing from one to two nights is (32-15)/((32-15)+(91-82))=65%.

[28] An indicator for mother's age being less than 16, each age, and then greater than 40, as well as a missing age indicator is included. Similar indicators for father's age are included as well.

The results suggest that approximately one quarter of births after midnight experience an extra night in the hospital. To place the estimate of 0.25 in context, Appendix Table A2 includes the covariates for a pooled sample from 1991-2002. Similar differences in length of stay are found for 1st births (0.23), 30-year old mothers compared to 20-year olds (0.28), missing father's information (0.20), and labors that were over 20 hours (0.26). Infants in the lowest birthweight quintile ($\leq$ approximately 3000 grams) had longer stays (coefficient = 0.7), and government hospitals had longer stays than for-profit hospitals (coefficient = 0.44).[29]

### D. *Newborn Outcomes*

The health outcomes we can measure are readmissions to a California hospital and mortality. In particular, 7-day readmissions, 28-day readmissions, as well as 28-day and 1-year mortality are considered. The 7- and 28-day measures are calculated from the midnight that defines the threshold.[30]

Table 2 shows that health outcomes are similar for those born before or after midnight, with statistically and economically insignificant differences. The readmission rates and associated hospital charges are slightly larger for those born after midnight (the group with longer spells in the hospital), although the result is statistically significant only for the 28-day readmission rate in the time period before the law change.

Mortality is less frequently observed, with 28-day mortality rates for this analysis sample of 3 per 1000 and 1-year mortality rates of 4-5 per 1000. Lower mortality rates after the law change are largely due to the exclusion of scheduled, and potentially riskier, births (these births are excluded beginning in 1995 due to data availability). The mortality rates are similar before and after midnight, with differences that are not statistically significant. These results are perhaps better interpreted in the context of variation in mortality rates across the day.

Figure 4 (along with Appendix Figure A5) report means of the health outcome measures for each minute of the day, as well as local-linear regression results using a pilot bandwidth of 20 minutes. The left two panels of Figure 4 consider births before the law change, whereas the right panels report the results after the law change. Little change is found before and after midnight for outcomes. In particular, the readmissions are flat across the minutes of birth (Figure 4A). If anything, there appears to be an increase in

[29] Similar estimates are found when length of stay is treated as a count variable and a negative binomial model is estimated with full controls, with the marginal effect of a post-midnight birth estimated to be 0.203 (s.e.= 0.015) before the law change and 0.244 (s.e.= 0.020) after the law change. Further, the length of stay may be considered censored when a newborn is discharged to another facility or when a newborn dies in the hospital (1.5% of the sample). When a Cox proportional hazard model of the additional midnights in care + 0.5 was estimated with full controls and taking into account this possible censoring, the estimated change is slightly smaller with the hazard ratio estimated to be 0.858 (s.e. = 0.005) before the law change and 0.831 (s.e. = 0.007) afterwards. When the censored observations were excluded from the analysis, the estimates were similar to the main results as well.

[30] For example, 28-day readmission is coded to 1 if the difference between the readmission date and the date of birth were less than or equal to 28 for those born after midnight, and less than or equal to 29 for those born just prior to midnight.

readmissions for births that occur after midnight–infants with a longer length of stay–prior to the law change. This increase is not sustained, however, and is not seen in the post-law-change period.

Considering mortality in the bottom two Panels of Figure 4, the data exhibit more noise as these outcomes are less frequently observed. This is especially the case at the boundaries when data from both sides of the point of interest cannot be used in the estimation, as well as the lack of reporting of births at exactly midnight. More generally, an increase in the mortality rate after business hours is observed, as in previous research that questions whether such increases are due to changes in staffing (Z.C. Luo and J. Karlberg 2001).[31] No sustained change in mortality is seen just before and after midnight, however.

For a magnified view, Appendix Figures A6 and A7 report the results from 8 p.m. to 4 a.m., and for further comparison a bandwidth of 10 minutes was used. Figure A6C shows a slight increase in 28-day readmissions following the midnight birth, and Figure A6B shows that the 1-year mortality rate is close to 6 per 1000 in the minutes from 11 p.m. to 1 a.m., with noisier measurements at the boundary. After the law change, any differences in mortality rates shown in Table 2 for the 40-minute sample are not found to be sustained in the minutes after 12:25 a.m.

The lack of an effect shown in the figures is evident in Tables 4 and 5, which report results for readmissions and mortality with and without controls for newborn characteristics. Columns 1 to 4 of Table 4 consider 28-day readmission rates by time of birth. Readmission is modestly more common for post-midnight births prior to the law, and slightly less common post-midnight following the law change, although the differences are not statistically significant. The small magnitudes and the instability of the signs, which are usually contrary to a diminishing returns possibility given the positive point estimates in the pre-law period and negative coefficients in the post-law period, are consistent with Figure 4 that outcomes look similar before and after midnight.[32] Columns 5 and 6 of Table 4 incorporate the substantial differences in stay length on either side of the midnight threshold into an IV estimate of stay length on readmissions. Unsurprisingly, the effect of stay length on 28-day readmission cannot be distinguished from zero, but these IV estimates are less precise than the corresponding reduced form estimates.

Similar results are obtained for neonatal mortality in Table 5. The coefficients for birth after midnight are close to zero in both time periods and are of unstable sign. In both the 40-minute and 2-hour samples, the lower limit on the 95-percent confidence interval is -0.0002, or 5% of the pre-midnight mean. After the law change, the lower limits are -0.00007 and -0.00016 (or 2-5% of the mean). For 1-year mortality, some of the coefficients found are fairly large, but they are not robust (Appendix Table A4). For

[31]When comparing births just before and after midnight, it is possible that the staff changes shifts at the same time. California hospital advertisements for nursing services suggest that 12-hour shifts that end at 7 p.m. are common. Midnight shift changes are possible in the case of 8-hour shifts, although the increase in the mortality rate in our data appears to occur between 7 and 9 p.m.

[32]Similarly, virtually no difference is found in 7-day readmissions in both time periods for those born before or after midnight, with an estimated increase in readmissions of 0.04%, or 1.4% of the pre-midnight mean, despite longer stays in care (shown in Appendix Table A3).

example, the local-linear estimation prior to the law change, and the probit models using the 2-hour samples before and after the law change, yield coefficients that are essentially zero. The results are less precisely estimated, however. Using the 2-hour samples, the lower limit on the 95% confidence intervals is -0.0007 and -0.0006 for the two time periods (or 13% of the pre-midnight means). While fairly large effects are within the confidence interval for 1-year mortality, the lack of robustness of any beneficial effect of longer stays associated with an after-midnight birth again confirms the visual evidence from Figure 4 that outcomes appear remarkably similar despite the difference in length of stay for the two groups. Likewise, the IV estimates reported in the last two columns of Table 5 again find no significant effect of stay length on mortality, though again precision is compromised.

To place these results in context, Table A2 includes the estimated marginal effects of the covariates evaluated at the sample mean. In terms of statistical significance, patients with few prenatal visits, boys, newborns with a birthweight in the lowest quintile, and Medicaid patients tend to have worse outcomes. Newborns to new mothers were more likely to have a readmission (14% higher than the mean), but little difference is found in terms of mortality. Other covariates, such as maternal education, are found to have little relation to infant mortality (controlling for the other covariates).

### *E. Complier Characteristics*

In a local average treatment effect setting such as this, the estimated effects apply to a population of compliers: those who are induced to have a longer stay as a result of the post-midnight birth. Compliers are likely to differ from a random draw from the population. In particular, the results are most likely to apply to uncomplicated births where the minute of birth is plausibly exogenous and the stay length is not expected to be especially long so that coverage for one or two days is more likely to bind.

While it is not possible to identify individual compliers in the data, it is possible to estimate their mean observable characteristics, as described in Section IV. Given the different effects on length of stay before and after the law change, births of at least one additional midnight before the law change and at least 2 additional midnights after the law change were coded as receiving the longer-stay "treatment" ($D = 1$). The estimated fraction of compliers is similar in the two time periods (16% prior to the law change and 17% after the law change). Always takers are more common prior to the law change, when the threshold for a longer stay is lower (57% vs. 15%). Given these proportions and the average characteristics of always takers, $E(X|D = 1, Z = 0)$, along with the average characteristics of patients who are either always takers or compliers, $E(X|D = 1, Z = 1)$, we calculated the implied means of the complier characteristics (means reported in Appendix Table A5).

Overall, it appears that the compliers are quite similar to the population of births. The main differences are that the compliers are less likely to be low birthweight and more likely to be full term, as expected. Across the two time periods, we also find that the complier group is slightly less likely to be the result of a stimulated labor, and the mother is more likely to have been admitted on a weekend. Before the law change, we generally find that those who are more likely to be disadvantaged are also more likely to be compliers (mothers who are high school drop outs, missing father's education, and Medicaid recipients). After the law change, the reverse tends to be found, with compliers more likely to be privately insured. There are exceptions in both time periods, however, and the differences tend to be small and are rarely statistically significantly different.[33] As a summary, the predicted 28-day readmission rate from a probit model estimated with the full set of control variables was estimated, and the measure was evenly distributed for compliers.[34]

The compliers from a before-after estimator using the law change as the source of variation in length of stay can also be described using all data from January-August in 1997, 1998, and 1999 (excluding stays of more than 28 days). Evans, Garthwaite and Wei (2008) noted that some Medicaid recipients were excluded from the law for the middle time period while all were covered from January 1999 onwards. To define compliers, the "treatment" is a stay of 2 or more days in the hospital or 4 days for c-section births (D=1), and the estimated proportion of compliers is 0.21 and 0.25 for the 1998 and 1999 time periods. Similar to our characteristics, this group is also less likely to be low birthweight (constituting approximately 2.5% of compliers vs. 6% overall for this time period). Appendix Table A6 reveals that other characteristics show much larger differences: compliers are much less likely to be births to mothers with less than a high school education (approximately 18% vs. 31% overall) and more likely to be college graduates (32% vs. 19%). As expected, the compliers are less likely to receive Medicaid in the middle time period (19% vs. 42%), but also in the period when all births are covered (27% vs. 42%).

### *F. Maternal Outcomes*

When we consider instead maternal outcomes, findings are similar to those for newborns. Appendix Table A7 reports the results for maternal length of stay and readmissions, although adverse outcomes are less common among mothers.[35] The mother's length of stay was calculated as the number of additional midnights after the birth of the child. The post-midnight increase is similar to that for newborns (0.30 and 0.23),

[33] 95% confidence intervals were constructed using a bootstrap procedure, where the sample was redrawn 300 times and the weights for compliers and always takers were re-estimated each time to reflect variation in these estimates. Prior to the law change, compliers are (statistically significantly) more likely to be Hispanic and have a mother admitted on a weekend. Meanwhile, parents' ages are younger, and the newborns are less likely to be low birthweight. After the law change, there are differences in prenatal visits (compliers are more likely to have fewer than 9), compliers are also less likely to be a first birth or an induced birth.

[34] The measure was broken into quartiles for the full sample, and in the pre-law period, the complier means of these quartiles are 0.257, 0.236, 0.274, and 0.233; after the law change they are 0.214, 0.245, 0.278, 0.263.

[35] The death certificate data were linked only for newborns. When in-hospital mortality was considered within 1 year for mothers, that mortality rate was 8 per 100,000 and the estimates before and after midnight were much noisier.

although it is larger relative to the (smaller) mean length of stay for mothers in both time periods. Despite the longer length of stay for mothers who give birth after midnight, little relationship is found for readmissions. 28-day readmissions are rare (8 per thousand), and a birth after midnight is associated with a small decline in readmissions prior to the law change, and a small increase after the law change, although neither difference is statistically insignificant.

### G. Robustness

The data were explored to test the robustness of the main results and to consider subgroups that have been identified in previous research to benefit from longer stays. These results are listed in Appendix Table A8 and the main results are remarkably stable.[36] Similar results were found for cesarean section births, as well as scheduled births–two of the sample restrictions in the main analysis. As a summary, similar results were found for births with high or low predicted mortality based on the observable characteristics of the newborns. Of note, Medicaid patients have a larger increase in length of stay at midnight, whereas mothers with low education levels had smaller increases. Results were similar for Friday/Saturday births, when postpartum care may be compromised for early discharges during the weekend.

Results were also similar when alternative measures of readmission timeperiods were used (Figure A8), and when readmissions were weighted by hospital charges accrued during the readmission. Other models with similar results included mother fixed effects (over time), hospital fixed effects, hospital-by-date fixed effects, a triple difference strategy considering midnight jumps across hospitals before and after the law change, and models that incorporated the 10 minutes immediately around midnight. In terms of the standard errors, the results are qualitatively similar when we cluster by minute-of-birth as suggested by David Card and David S. Lee (2006).[37] Last, the results are robust to bandwidth selection. This has been described above, with the Figures shown with bandwidths of 20 minutes and 10 minutes (in the Appendix), and the tables that essentially use bandwidths of 20 minutes for the 40-minute sample and 55 minutes for the "2-hour" sample.

Hospital characteristics allow an examination of effects at hospitals where incentives to limit length of stay might differ. For-profit hospitals are found to have larger increases in length of stay for those born after midnight. It appears that the 48-hour rule may be

[36] These results are discussed in more detail in the working paper version of this paper. For 7-day and 28-day readmissions, and 28-day and 1-year mortality, across 15 subgroups of interest (60 probit models), 3 were found to be statistically significant at an 5% level (uncorrected for the number of tests conducted), and all were in the pre-1997 time period. Post-midnight births in teaching hospitals were associated with a lower 28-day readmission rate; post-midnight c-sections were associated with a lower 1-year mortality rate; and post-midnight births when there was a labor complication were associated with higher 28-day readmissions.

[37] Some estimates are slightly larger and others slightly smaller. For the two-hours sample prior to the law change, for example, the clustered standard error of interest for 28-day mortality is 0.0001268 compared to 0.0001222 when no clustering is used; for 1-year mortality the clustered standard error is 0.0002946, smaller than 0.000349 calculated without clusters.

more likely to bind in these hospitals for those born just before the cutoff. Meanwhile, Kaiser hospitals–hospitals where the insurer owns the hospital and the billing rules may be expected to be less salient in terms of hospital incentives to extend the length of stay–tend to have shorter lengths of stay for everyone, by approximately 0.1 nights on average. They also have smaller jumps at midnight (0.19 before the law change and 0.09 after the law change).

For-profit hospitals have slightly higher rates of births prior to midnight in the 40-minute sample. Do differing incentives to deliver prior to midnight lead to different obstetric procedures by hospital type? One way that the time of birth can be manipulated is by stimulating labor and by the use of forceps or a vacuum. We find suggestive evidence that for-profit hospitals are modestly more likely to use forceps or vacuum procedures at times when the reimbursement is tied to the minute of birth. In particular, after the law change the forceps/vacuum rate is on the order of 20% higher at for-profit hospitals from 9 p.m. to 1 a.m. (results available from authors).[38] Before the law change the difference is smaller, however. Due to the low rates of forceps/vacuum use, the absolute magnitude of the difference by hospital type is modest: for profit hospitals are less than 1.5 percentage points more likely to accelerate delivery near midnight.

## VI. Interpretation

Our results suggest that extending the length of stay by an additional night provides little health benefit for uncomplicated births in terms of rehospitalizations and mortality.[39] The main welfare benefit would come from reductions in mortality attributable to longer hospital stays.[40] The point estimates are generally small, however, and even at the lower limits of the 95% confidence interval for 28-day mortality (mortality potentially most likely to respond to an extra night in the hospital), the implied cost of saving a statistical life ranges from $2 to $6 million, which are approaching typical value-of-statistical-life estimates (such as $1.9 million by Orley Ashenfelter and Michael Greenstone (2004)).[41] Overall, it appears that longer lengths of stay associated with minimum-stay mandates are not worth the extra expense for uncomplicated births, at least as reflected by readmission and mortality outcomes.

[38] Unsuccessful accelerations (insofar as the midnight reimbursement rule is concerned) may be reflected in deliveries shortly after midnight.

[39] Of course, whether a birth is "uncomplicated" is known prior to when length of stay is generally determined.

[40] There may be other benefits to stay length, although we argue these do not appear to be particularly valued as described in the Conclusions. Costs of readmission could also be considered. Using our design, facility charges associated with readmissions are not found to be related to the time of birth. Even at the lower end of the 95% confidence interval, the readmission charges are $300 lower for post-midnight births prior to the law change and only $40 lower after the law change.

[41] The lower limits range from -0.0002 to -0.0007. $400/0.00002 = $2m. When 1-year mortality is considered, the point estimates are again zero, although the confidence intervals widen to include cost of saving a statistical life on the order of $500,000. Similarly, when an IV model was estimated, the point estimates are zero, although confidence intervals are wider and a $250,000 cost of saving a statistical life cannot be rejected.

By applying to all deliveries – including relatively routine deliveries most responsive to our instrument – it appears minimum stay-length legislation incurred substantial unnecessary cost over the last decade. Healthcare expenditures associated with an extra night in the hospital are generally in the range of $1000 in the mid-1990s. With 4.6 million births per year, an increase of 0.25 days would be on the order of $1.1 billion per year (or $11 billion since 1997).[42] As argued above, our identification strategy credibly precludes adverse selection and reveals no health benefits. Therefore, if the expenditures reflect the opportunity cost of the resources used, we interpret $11 billion to the the legislation's cost from moral hazard. In California, our data suggest that the budgetary cost of an extra night in the hospital is roughly $1500. An increased length of stay of 0.25 days on average would cost roughly $400 per birth or $200 million per year in California.

Estimating the implied costs permits a comparison with other health initiatives. Nevertheless, marginal *social* costs of an additional night in the hospital may be low given the availability of hospital staff regardless of the number of births on a particular day. We considered times when the hospital had an unusually large number of births around the midnight in question to test the effects of a post-midnight birth when the marginal cost of a bed is likely higher. We found small decreases in length of stay for all of the newborns associated with these busier times, but the estimated discontinuity at midnight was not affected.[43]

## VII. Conclusions

This paper makes use of a rule of thumb in patient billing which approximates the length of California hospital stays with the number of midnights in hospital following delivery. In apparent response to the discontinuous financial incentives, infants born just after midnight remain in the hospital about 0.25 additional nights, on average, compared to infants born just prior to midnight. This implies that 1 in 4 infants born after midnight stayed an additional night in the hospital compared to similar infants born minutes earlier. In addition, the 1997 early discharge law in California allows us to consider estimates that are drawn mainly from newborns induced to stay one additional midnight prior to the law change and a second midnight in the post-law period. In the

[42] Madden et al. (2002) found an HMO's expenditure related to an extra night to be roughly $1000. Similarly, Kristiana Raube and Katie Merrell (1999) found extra charges on the order of $1000 in the mid 1990s. A lower estimate would come from Russell et al. (2007), who used the 2001 Nationwide Inpatient Survey and found average (facility) costs of $600 per delivery (for births > 2500g). Susan K. Schmitt, LaShika Sneed and Ciaran S. Phibbs (2006) used 2000 California data and found the average cost for mothers and newborns of $4750 ($3100 for mothers and $1650 for infants), again for newborns > 2500g. As noted above, when twins whose times of birth straddled the midnight threshold are considered with our data, the infant born prior to midnight has $1000 higher charges. Last, we conducted an analysis of hospital costs associated with length of stay for uncomplicated births and arrived at an estimate of $1500. This is described in greater detail in a working paper version of the paper.

[43] The main results were unaffected when we controlled for the number of births in the infant's hospital 5 days before and after the midnight used to define the threshold. We also calculated the the maximum number of 3-day birth counts per hospital and controlled for the fraction of this capacity used in the 2 days before and 1 day after the midnight in question. Again, the results were unchanged.

presence of diminishing returns to stay length, we would expect to observe a larger health benefit of being born shortly after midnight prior to the law.

We find no outcome differences associated with post-midnight birth, even among births prior to the minimum-stay law. Our impact estimates are fairly precise: at the endpoints of our confidence intervals, the most "optimistic" impact estimates would not justify the greater expenditures associated with longer stays. This finding suggests that physicians can identify newborns who require additional time in the hospital and that the technology of postpartum care is such that it can be administered effectively on the first day of life. This finding is consistent with profit-maximizing HMOs in the early 1990s driving down stay lengths, despite having to reimburse hospitals for (relatively costly) readmissions.These results apply to uncomplicated births where the exact minute of birth is plausibly exogenous. When the complier characteristics were compared, they appear similar to the universe of births, unlike previous estimates that used the timing of state law changes.

The outcomes considered here, while they are particularly costly and tend to dominate welfare analyses, do not capture other potential benefits to parents and infants. For example, breastfeeding initiation may respond to postpartum stay length, but is not recorded in the California natality data until 2006 (after our analysis period). Additionally, longer stays may provide benefits in terms of additional rest and the comfort of professional supervision. Future research might seek to capture benefits we have not been able to measure by considering the longer-term effects of postnatal stay length (e.g. on adolescent health or test scores).

The market for longer stays is somewhat unusual in that it requires permission from the physician. Nevertheless, the brevity of stays for uncomplicated births (in the cross section) suggests that an additional day is not particularly valued on average. Finally, to the extent that early discharge is particularly unpleasant or prevents detection of chronic problems such as postpartum depression, we might expect the next birth to be postponed. Our sibling-matched California data allow us to examine this question directly, and we find that a post-midnight birth is not associated with a difference in the time interval to the next birth.[44] For uncomplicated births – births where the early-discharge laws were most likely to bind – it appears that longer hospital stays yield little health gains in terms of hospital readmissions and mortality rates, suggesting substantial moral hazard.

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[44]In the pooled 2-hour sample, the fraction of Mother's with another infant within 3 years is 0.171 for mothers with a pre-midnight birth and 0.167 for mothers with a post-midnight birth. Even this difference vanishes when the full set of controls are used. When a hazard model of the time to the next birth is used, a post-midnight birth is associated with a hazard ratio of 1.0008 before the law change and 1.007 after the law change with neither estimate statistically significantly different from 1.

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**Table 1: Selected Characteristics: 40 Minutes Around Midnight Sample**

| | | All Years | | |
|---|---|---|---|---|
| | | Before Midnight | After Midnight | p-value |
| Pregnancy Characteristics | At least one pregnancy complication | 0.585 | 0.589 | (0.264) |
| | <9 prenatal visits | 0.199 | 0.204 | (0.052) |
| | 9-15 prenatal visits | 0.695 | 0.689 | (0.043)* |
| | >15 prenatal visits | 0.088 | 0.089 | (0.655) |
| | Prenatal visits missing | 0.019 | 0.019 | (0.850) |
| Mother's Characteristics | Born in California | 0.390 | 0.391 | (0.753) |
| | Born outside U.S. | 0.472 | 0.475 | (0.411) |
| | 1st Birth | 0.400 | 0.394 | (0.047)* |
| | Age | 26.82 | 26.79 | (0.489) |
| | High school drop out | 0.355 | 0.356 | (0.652) |
| | High school | 0.287 | 0.288 | (0.684) |
| | Some College | 0.184 | 0.181 | (0.250) |
| | College+ | 0.164 | 0.164 | (0.907) |
| Father's Characteristics | Age | 29.755 | 29.725 | (0.529) |
| | High school drop out | 0.302 | 0.303 | (0.848) |
| | High school | 0.287 | 0.285 | (0.465) |
| | Some College | 0.156 | 0.155 | (0.650) |
| | College+ | 0.181 | 0.181 | (0.898) |
| | Missing education data | 0.075 | 0.078 | (0.086) |
| Newborn Characteristics | Boy | 0.509 | 0.507 | (0.530) |
| | White | 0.544 | 0.537 | (0.029)* |
| | African American | 0.065 | 0.066 | (0.570) |
| | Hispanic | 0.186 | 0.188 | (0.443) |
| | Asian | 0.092 | 0.094 | (0.198) |
| Birth Characteristics | Birthweight < 2500 grams | 0.048 | 0.047 | (0.506) |
| | Gestational age >= 37 weeks | 0.909 | 0.906 | (0.170) |
| | Vaginal birth after C-section | 0.025 | 0.026 | (0.353) |
| | Forceps or vacuum | 0.099 | 0.095 | (0.022)* |
| | Less than 3 hours | 0.018 | 0.019 | (0.228) |
| | More than 20 hours | 0.007 | 0.006 | (0.310) |
| | Labor stimulated | 0.119 | 0.115 | (0.053) |
| | Labor induced | 0.099 | 0.095 | (0.065) |
| | Admitted on a Weekend | 0.237 | 0.241 | (0.141) |
| Primary Payer | Medicaid | 0.463 | 0.463 | (0.898) |
| | Self pay/unknown | 0.041 | 0.039 | (0.182) |
| | Private | 0.480 | 0.482 | (0.458) |
| | Government | 0.223 | 0.227 | (0.147) |
| Hospital Characteristics | Private nonprofit | 0.634 | 0.636 | (0.554) |
| | Private for-profit | 0.142 | 0.136 | (0.012)* |
| | Observations | 45807 | 47046 | |

Data are pooled 1991-2002. Calculations from the "40 minute" sample includes births from 11:37pm-11:55pm & 12:05am-12:24am. Race and Ethnicity is not broken out separately after 1994. Prenatal visits is missing in approx. 2% of the cases. Mother's education is missing in approximately 1% of the cases. Full set of variables and separate time periods are listed in the appendix. The p-value is calculated from a test of the difference in means before and after midnight. ** = significant at 1%; * = significant at 5%.

## Table 2: Time of Birth, Length of Stay, and Infant Outcomes: 40 Minutes Around Midnight Sample

| | | Before 1997 Law Change | | | After 1997 Law Change | | |
|---|---|---|---|---|---|---|---|
| | | Born Before Midnight | Born After Midnight | | Born Before Midnight | Born After Midnight | |
| **Variable** | | **Mean** | **Mean** | **p-value** | **Mean** | **Mean** | **p-value** |
| Length of Stay | Raw Length of Stay | 1.99 | 1.23 | (0.000)** | 2.29 | 1.58 | (0.000)** |
| | Additional Midnights | 0.99 | 1.23 | (0.000)** | 1.29 | 1.58 | (0.000)** |
| | >=1 Additional Midnights | 0.57 | 0.73 | (0.000)** | 0.82 | 0.91 | (0.000)** |
| | >=2 Additional Midnights | 0.11 | 0.17 | (0.000)** | 0.15 | 0.32 | (0.000)** |
| | >=3 Additional Midnights | 0.06 | 0.08 | (0.000)** | 0.07 | 0.09 | (0.000)** |
| Infant Outcomes | 7-Day Readmission | 0.027 | 0.028 | (0.428) | 0.028 | 0.027 | (0.690) |
| | 28-Day Readmission | 0.042 | 0.045 | (0.047)* | 0.047 | 0.046 | (0.522) |
| | 28-Day Readmission Charges | 852 | 938 | (0.376) | 1128 | 1287 | (0.346) |
| | 1-year Readmission Charges | 1763 | 1868 | (0.457) | 2189 | 2461 | (0.246) |
| | 28-Day Mortality | 0.0035 | 0.0036 | (0.836) | 0.0028 | 0.0030 | (0.829) |
| | 1-year Mortality | 0.0056 | 0.0053 | (0.687) | 0.0040 | 0.0042 | (0.711) |
| | Observations | 28898 | 29477 | | 16637 | 17283 | |

Calculations from the "40 minute" sample includes births from 11:37pm-11:55pm & 12:05am-12:24am. 28-day measures include outcomes 28 days from the midnight considered for each birth. "Before 1997 law change" includes births from January 1, 1991 to July 31, 1997; "After 1997 law change" includes births from September 1, 1997-December 31, 2002. The p-value is calculated from a test of the difference in means before and after midnight. ** = significant at 1%; * = significant at 5%.

## Table 3: Time of Birth and Length of Stay

**A. Before 1997 Law Change**

| Dependent Variable: | Additional Midnights | | | | |
|---|---|---|---|---|---|
| Model: | Local linear | OLS | OLS | OLS | OLS |
| | (1) | (2) | (3) | (4) | (5) |
| Birth After Midnight | 0.273 | 0.293 | 0.272 | 0.220 | 0.216 |
| | (0.036) | (0.033) | (0.032) | (0.0201) | (0.019) |
| Birth After Midnight * Minute from cutoff | | -0.001 | -0.001 | 0.001 | 0.001 |
| | | (0.002) | (0.00) | (0.0005) | (0.0004) |
| Birth Prior to Midnight* Minute from cutoff | | -0.004 | -0.003 | 0.001 | 0.001 |
| | | (0.002) | (0.002) | (0.0004) | (0.0004) |
| Full Controls | No | No | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2 hour | 2hour |
| Observations | 60398 | 58375 | 58375 | 162821 | 162821 |
| Mean of Dep. Variable Before Midnight | 1.00 | 0.99 | 0.99 | 0.97 | 0.97 |

**B. After 1997 Law Change**

| Dependent Variable: | Additional Midnights | | | | |
|---|---|---|---|---|---|
| Model: | Local linear | OLS | OLS | OLS | OLS |
| | (1) | (2) | (3) | (4) | (5) |
| Birth After Midnight | 0.255 | 0.238 | 0.230 | 0.252 | 0.246 |
| | (0.045) | (0.041) | (0.039) | (0.026) | (0.025) |
| Birth After Midnight * Minute from cutoff | | 0.006 | 0.006 | 0.001 | 0.001 |
| | | (0.003) | (0.002) | (0.0006) | (0.0005) |
| Birth Prior to Midnight* Minute from cutoff | | -0.001 | -0.001 | 0.001 | 0.001 |
| | | (0.0023) | (0.003) | (0.0006) | (0.0006) |
| Full Controls | No | No | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2 hour | 2hour |
| Observations | 35736 | 33920 | 33920 | 94879 | 94879 |
| Mean of Dep. Variable Before Midnight | 1.30 | 1.29 | 1.29 | 1.28 | 1.28 |

Columns (1)-(3) consider the "40 minute" sample from 11:37pm-11:55pm & 12:05am-12:24am. Columns (4) and (5) consider the "two hour" sample from 11:02pm-11:55pm & 12:05am to 12:59am. Column (1) reports the difference in local linear regression estimates just above and below the discontinuity using a triangle kernel and a bandwidth of 20 minutes. Asymptotic standard errors in parentheses. Columns (2)-(5) report OLS results using different samples close to the midnight cutoff, along with linear trends in the time of birth before and after midnight; robust standard errors reported in parentheses. Full controls include the controls listed in Appendix Table A1, as well as indicators for the mother's age, the father's age, the year of birth and the month of birth.

**Table 4: Time of Birth and Infant Readmissions**

**Dependent Variable: 28-Day Readmission**

**A. Before 1997 Law Change**

| Model: | Local linear | Probit | Probit | Probit | IV Probit | IV Probit |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Birth After Midnight | 0.0049 | 0.0059 | 0.0053 | 0.0027 | | |
| | (0.0036) | (0.0034) | (0.0033) | (0.0019) | | |
| Birth After Midnight * 100 Minutes from cutoff | | 0.0013 | 0.0031 | -0.0033 | | |
| | | (0.021) | (0.020) | (0.0043) | | |
| Birth Prior to Midnight* 100 Minutes from cutoff | | -0.027 | -0.024 | 0.0014 | | |
| | | (0.022) | (0.021) | (0.0044) | | |
| Length of Stay | | | | | 0.012 | 0.011 |
| | | | | | (0.0091) | (0.0089) |
| | | | | | [-0.463,3.35] | [-0.515, 3.25] |
| Full Controls | No | No | Yes | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2 hour | 2 hour | 2 hour |
| Observations | 60398 | 58375 | 58375 | 162821 | 162821 | 162821 |
| Mean of Dep. Variable Before Midnight | 0.042 | 0.042 | 0.042 | 0.042 | 0.042 | 0.042 |

**B. After1997 Law Change**

| Model: | Local linear | Probit | Probit | Probit | IV Probit | IV Probit |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Birth After Midnight | -0.0041 | -0.0038 | -0.0036 | -0.0012 | | |
| | (0.0048) | (0.0045) | (0.0043) | (0.0026) | | |
| Birth After Midnight * 100 Minutes from cutoff | | 0.020 | 0.017 | 0.0050 | | |
| | | (0.028) | (0.026) | (0.0058) | | |
| Birth Prior to Midnight* 100 Minutes from cutoff | | 0.0045 | 0.0034 | 0.0013 | | |
| | | (0.030) | (0.028) | (0.0059) | | |
| Length of Stay | | | | | -0.0046 | -0.0064 |
| | | | | | (0.011) | (0.011) |
| | | | | | [-2.21, 1.59] | [-2.38, 1.49] |
| Full Controls | No | No | Yes | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2hour | 2hour | 2hour |
| Observations | 35736 | 33920 | 33920 | 94879 | 94879 | 94879 |
| Mean of Dep. Variable Before Midnight | 0.046 | 0.047 | 0.047 | 0.046 | 0.046 | 0.046 |

40 minute sample includes births from 11:37pm-11:55pm & 12:05am-12:24am; 2 hour sample includes births from 11:02pm-11:55pm & 12:05am to 12:59am. Column (1) reports the difference in local linear regression estimates just above and below the discontinuity using a triangle kernel and a bandwidth of 20 minutes. Asymptotic standard errors in parentheses. Columns (2)-(4) report marginal effects evaluated at the mean of the covariates with robust standard errors reported in parentheses. Full controls include the controls listed in Appendix Table A1, as well as indicators for the mother's age, the father's age, the year of birth and the month of birth. Columns (5) and (6) report marginal effects from models that include a quartic in the residual from a regression of length of stay on the same variables that are included in this table. Robust standard errors in parentheses, 95% confidence interval of the bootstrapped t-statistics (500 replications) in brackets.

**Table 5: Time of Birth and 28-Day Mortality**

**Dependent Variable: 28-Day Mortality**

**A. Before 1997 Law Change**

| Model: | Local linear | Probit | Probit | Probit | IV Probit | IV Probit |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Birth After Midnight | -0.00035 | 0.00083 | 0.00001 | 0.00005 | | |
| | (0.0011) | (0.0010) | (0.00010) | (0.012) | | |
| Birth After Midnight * 100 Minutes from cutoff | | -0.0035 | 0.0001 | 0.00001 | | |
| | | (0.0060) | (0.0006) | (0.0003) | | |
| Birth Prior to Midnight* 100 Minutes from cutoff | | -0.0040 | -0.0003 | -0.0002 | | |
| | | (0.0069) | (0.0007) | (0.0003) | | |
| Length of Stay | | | | | -0.00036 | -0.00011 |
| | | | | | (0.0018) | (0.00049) |
| | | | | | [-1.61, 2.48] | [-1.91, 2.30] |
| Full Controls | No | No | Yes | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2 hour | 2 hour | 2 hour |
| Observations | 60398 | 58365 | 58365 | 162791 | 162791 | 162791 |
| Mean of Dep. Variable Before Midnight | 0.0036 | 0.0035 | 0.0035 | 0.0035 | 0.0035 | 0.0035 |

**B. After1997 Law Change**

| Model: | Local linear | Probit | Probit | Probit | IV Probit | IV Probit |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Birth After Midnight | 0.00043 | 0.00014 | 0.00270 | -0.00002 | | |
| | (0.0013) | (0.001207) | (0.0047) | (0.000071) | | |
| Birth After Midnight * 100 Minutes from cutoff | | -0.00007 | -0.0003 | 0.0001 | | |
| | | (0.000078) | (0.0003) | (0.0002) | | |
| Birth Prior to Midnight* 100 Minutes from cutoff | | 0.00007 | 0.00004 | -0.0002 | | |
| | | (0.000082) | (0.0003) | (0.0002) | | |
| Length of Stay | | | | | -0.0013 | -0.00034 |
| | | | | | (0.0014) | (0.00025) |
| | | | | | [-1.91, 2.09] | [-2.29, 1.69] |
| Full Controls | No | No | Yes | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2hour | 2hour | 2hour |
| Observations | 35736 | 31627 | 31627 | 94761 | 94761 | 94761 |
| Mean of Dep. Variable Before Midnight | 0.0030 | 0.0030 | 0.0030 | 0.0033 | 0.0033 | 0.0033 |

40 minute sample includes births from 11:37pm-11:55pm & 12:05am-12:24am; 2 hour sample includes births from 11:02pm-11:55pm & 12:05am to 12:59am. Column (1) reports the difference in local linear regression estimates just above and below the discontinuity using a triangle kernel and a bandwidth of 20 minutes. Asymptotic standard errors in parentheses. Columns (2)-(4) report marginal effects evaluated at the mean of the covariates with robust standard errors reported in parentheses. Full controls include the controls listed in Appendix Table A1, as well as indicators for the mother's age, the father's age, the year of birth and the month of birth. Columns (5) and (6) report marginal effects from models that include a quartic in the residual from a regression of length of stay on the same variables that are included in this table. Robust standard errors in parentheses, 95% confidence interval of the bootstrapped t-statistics (500 replications) in brackets.

Figure 1A: U.S. Vaginal Births:
Fewer Than 2 Days in the Hospital
1970-2004

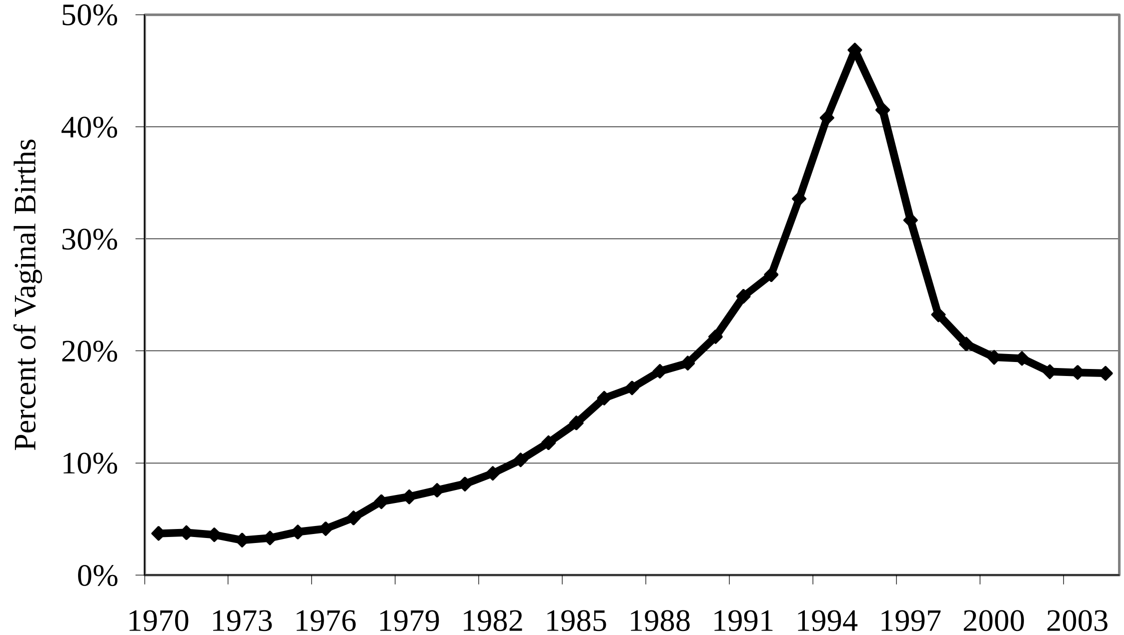
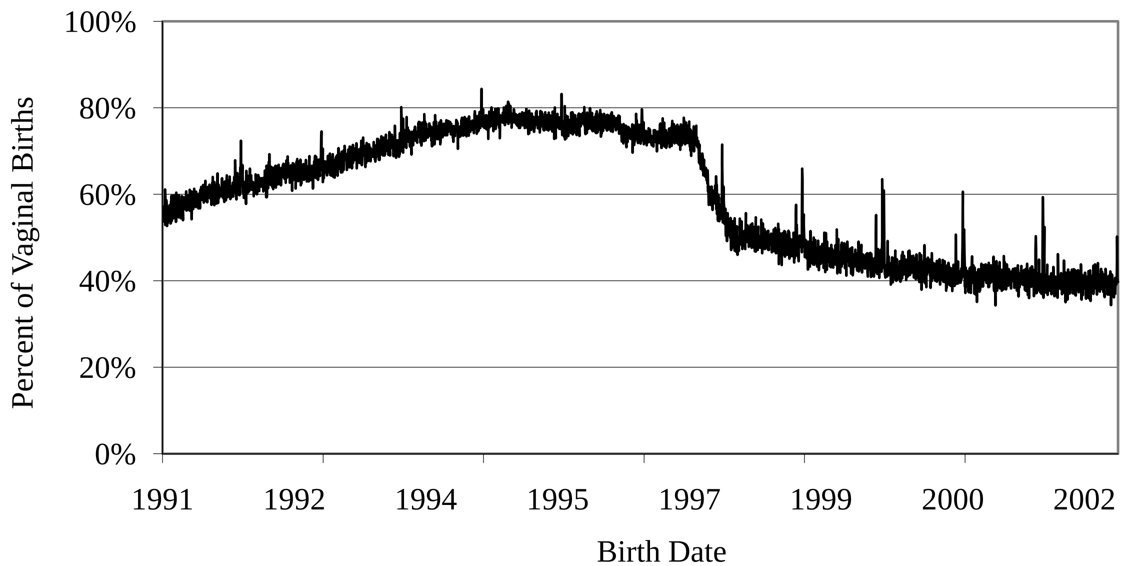



Figure 1B: California Vaginal Births:
Fewer than 2 Days in the Hospital
1991-2002




Sources: NCHS and California Linked Discharge-Birth Certificate-Death Certificate data.

Figure 2A: 24-hour Frequency by Minute of Birth:
1991-2002

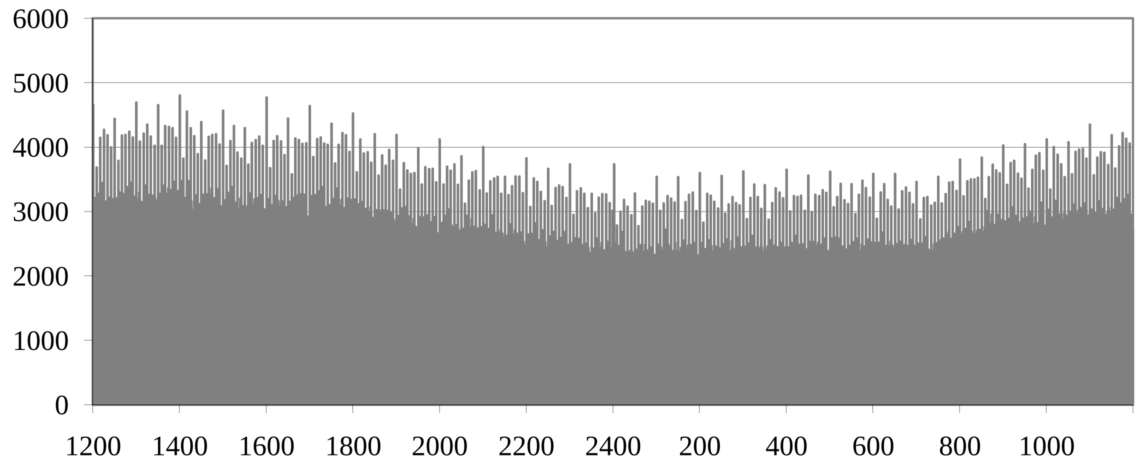
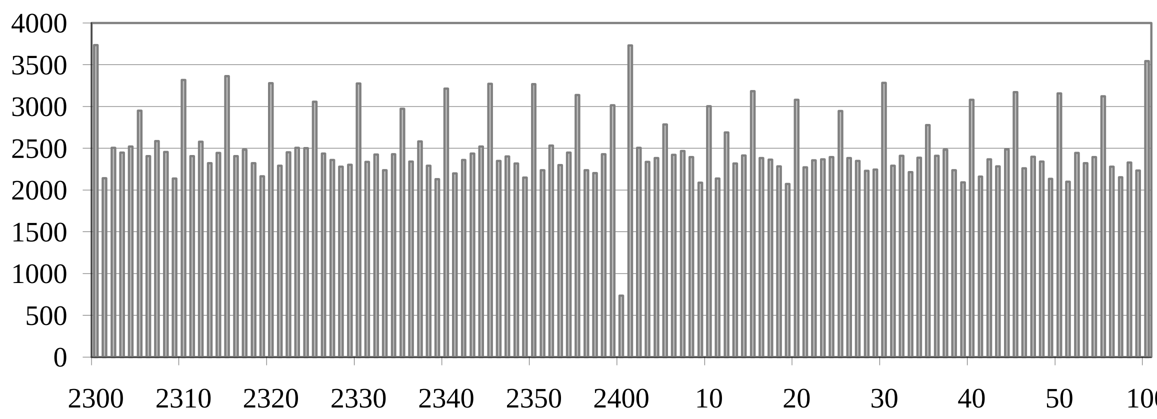
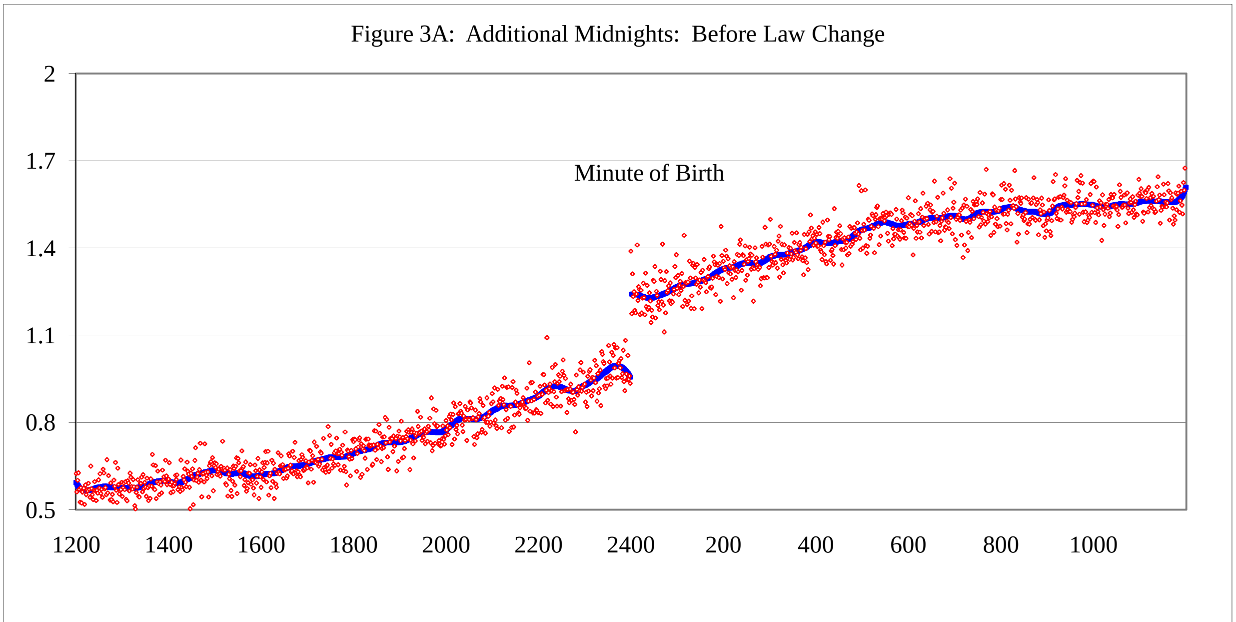
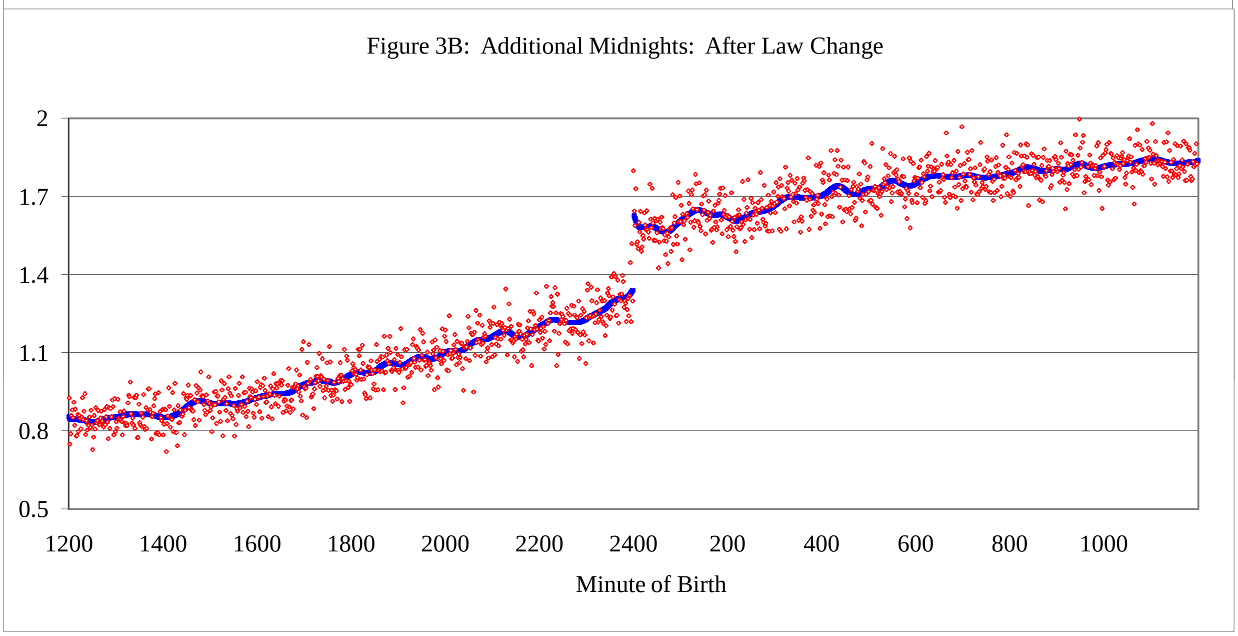



Figure 2B: Frequency by Minute of Birth (11pm-1am):
1991-2002




Source: California Linked Discharge-Birth Certificate-Death Certificate data. Figure 2B magnifies Figure 2A to examine births within 1 hour of midnight.




Number of additional midnights is the number of midnights for those born on or after midnight and before noon, while the number of additional midnights is measured as the number of midnights minus one for children born after noon and before midnight. Points represent means within 1-minute intervals from 12:00 noon to 11:59am. Lines represent local linear regressions, h=20

Figure 4A:  28-Day Readmission Rate:   Before Law Change

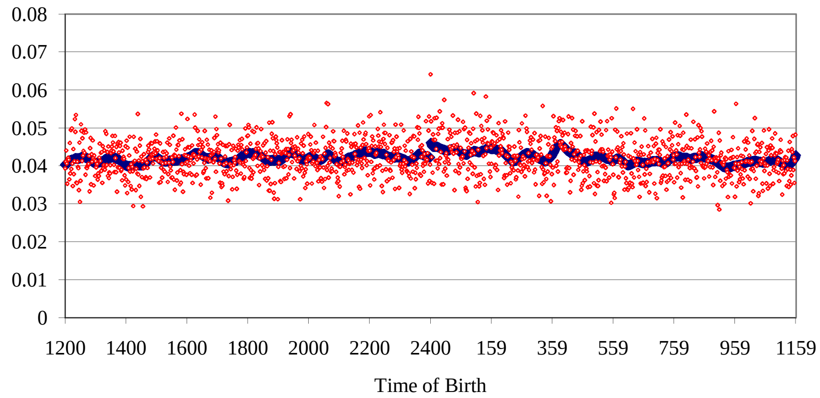



Figure 4B:  28-Day Readmission Rate:   After Law Change

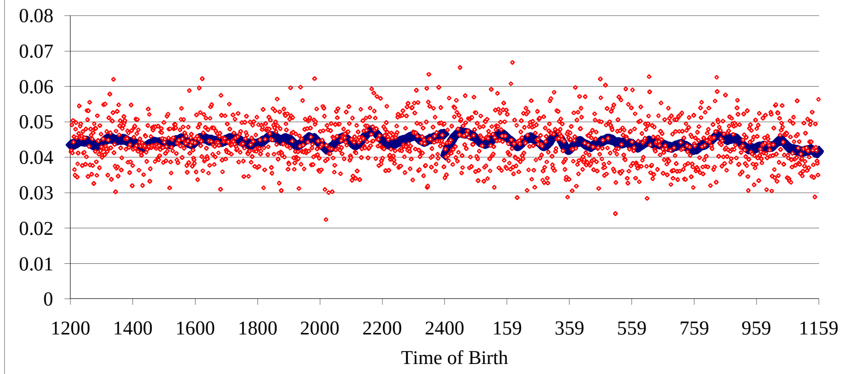



Figure 4C:  28-Day Mortality Rate:   Before Law Change

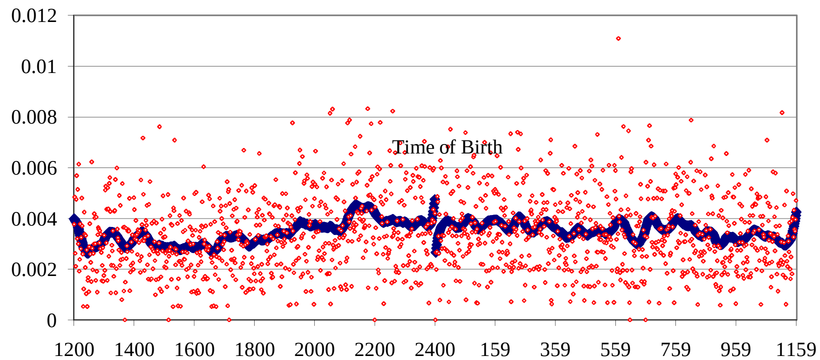
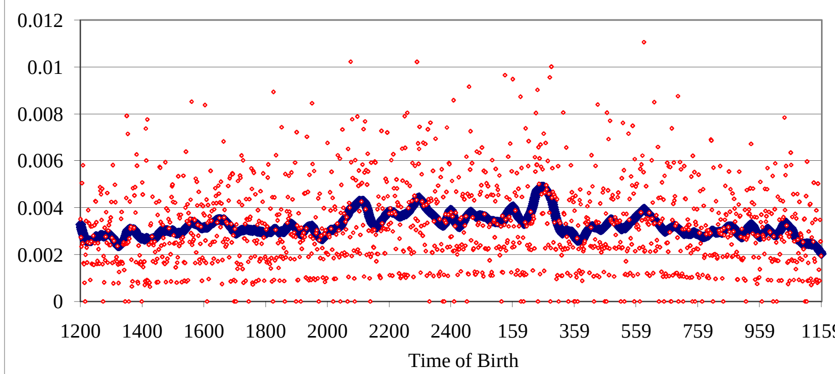



Figure 4D:  28-Day Mortality Rate:   After Law Change




28-day measures consider 28 days since the midnight prior to those born between midnight and noon, and 28 days since the following midnight for those born between noon and midnight. Points represent means within 1-minute intervals from 12:00 noon to 11:59am.  Lines represent local linear regressions, h=20

# Appendix Tables and Figures
(not for publication)

**Table A1: Table of Means**

| | | All Years | | | Before 1997 Law Change | | | After 1997 Law Change | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | Before Midnight | After Midnight | p-value | Before Midnight | After Midnight | p-value | Before Midnight | After Midnight | p-value |
| Pregnancy Characteristics | At least one pregnancy complication | 0.585 | 0.589 | (0.264) | 0.544 | 0.546 | (0.598) | 0.657 | 0.661 | (0.482) |
| | <9 prenatal visits | 0.199 | 0.204 | (0.052) | 0.225 | 0.232 | (0.058) | 0.153 | 0.156 | (0.393) |
| | 9-15 prenatal visits | 0.695 | 0.689 | (0.043)* | 0.675 | 0.669 | (0.105) | 0.729 | 0.723 | (0.222) |
| | >15 prenatal visits | 0.088 | 0.089 | (0.655) | 0.084 | 0.083 | (0.568) | 0.095 | 0.099 | (0.233) |
| | Prenatal visits missing | 0.019 | 0.019 | (0.850) | 0.017 | 0.018 | (0.338) | 0.022 | 0.021 | (0.413) |
| Mother's Characteristics | Born in California | 0.390 | 0.391 | (0.753) | 0.381 | 0.378 | (0.485) | 0.404 | 0.411 | (0.175) |
| | Born outside U.S. | 0.472 | 0.475 | (0.411) | 0.469 | 0.475 | (0.155) | 0.478 | 0.476 | (0.681) |
| | 1st Birth | 0.400 | 0.394 | (0.047)* | 0.396 | 0.394 | (0.673) | 0.407 | 0.393 | (0.009)** |
| | Age | 26.817 | 26.789 | (0.489) | 26.666 | 26.566 | (0.048)* | 27.085 | 27.168 | (0.222) |
| | High school drop out | 0.355 | 0.356 | (0.652) | 0.371 | 0.376 | (0.186) | 0.325 | 0.321 | (0.418) |
| | High school | 0.287 | 0.288 | (0.684) | 0.289 | 0.288 | (0.694) | 0.283 | 0.287 | (0.374) |
| | Some College | 0.184 | 0.181 | (0.250) | 0.183 | 0.18 | (0.303) | 0.187 | 0.185 | (0.672) |
| | College+ | 0.164 | 0.164 | (0.907) | 0.149 | 0.149 | (0.949) | 0.190 | 0.191 | (0.858) |
| | Missing education data | | | | | | | | | |
| Father's Characteristics | Age | 29.755 | 29.725 | (0.529) | 29.531 | 29.444 | (0.142) | 30.155 | 30.202 | (0.559) |
| | High school drop out | 0.302 | 0.303 | (0.848) | 0.316 | 0.321 | (0.222) | 0.277 | 0.271 | (0.229) |
| | High school | 0.287 | 0.285 | (0.465) | 0.292 | 0.285 | (0.056) | 0.278 | 0.284 | (0.231) |
| | Some College | 0.156 | 0.155 | (0.650) | 0.156 | 0.156 | (0.993) | 0.155 | 0.153 | (0.525) |
| | College+ | 0.181 | 0.181 | (0.898) | 0.172 | 0.171 | (0.774) | 0.197 | 0.197 | (0.909) |
| | Missing education data | 0.075 | 0.078 | (0.086) | 0.064 | 0.067 | (0.104) | 0.093 | 0.095 | (0.534) |
| Newborn Characteristics | Boy | 0.509 | 0.507 | (0.530) | 0.510 | 0.509 | (0.767) | 0.508 | 0.505 | (0.537) |
| | White | 0.544 | 0.537 | (0.029)* | 0.477 | 0.469 | (0.063) | 0.657 | 0.648 | (0.071) |
| | African American | 0.065 | 0.066 | (0.570) | 0.069 | 0.069 | (0.818) | 0.059 | 0.061 | (0.445) |
| | Hispanic | 0.186 | 0.188 | (0.443) | 0.295 | 0.3 | (0.173) | 0.000 | 0 | (.) |
| | Asian | 0.092 | 0.094 | (0.198) | 0.089 | 0.092 | (0.168) | 0.098 | 0.1 | (0.580) |
| Birth Characteristics | Birthweight bottom quartile | 0.197 | 0.196 | (0.803) | 0.199 | 0.197 | (0.459) | 0.193 | 0.196 | (0.499) |
| | Birthweight 2nd quintile | 0.212 | 0.22 | (0.003)** | 0.211 | 0.221 | (0.004)** | 0.212 | 0.217 | (0.271) |
| | Birthweight 3rd quintile | 0.216 | 0.217 | (0.740) | 0.214 | 0.216 | (0.502) | 0.221 | 0.219 | (0.685) |
| | Birthweight 4th quintile | 0.194 | 0.19 | (0.166) | 0.194 | 0.19 | (0.227) | 0.194 | 0.192 | (0.641) |
| | Birthweight Top quintile | 0.182 | 0.177 | (0.068) | 0.183 | 0.177 | (0.073) | 0.180 | 0.176 | (0.336) |
| | Gestational age < 37 weeks | 0.088 | 0.09 | (0.190) | 0.086 | 0.089 | (0.203) | 0.090 | 0.092 | (0.570) |
| | Gestational age 37 <= weeks < 40 | 0.424 | 0.434 | (0.003)** | 0.410 | 0.423 | (0.001)** | 0.449 | 0.452 | (0.616) |
| | Gestational age 40 <= weeks < 42 | 0.382 | 0.369 | (0.000)** | 0.390 | 0.376 | (0.001)** | 0.368 | 0.359 | (0.081) |
| | Gestational age weeks >=42 | 0.076 | 0.075 | (0.716) | 0.082 | 0.078 | (0.087) | 0.064 | 0.068 | (0.103) |
| | Gestational age missing | 0.031 | 0.032 | (0.298) | 0.032 | 0.034 | (0.217) | 0.029 | 0.029 | (0.912) |
| | Vaginal birth after C-section | 0.025 | 0.026 | (0.353) | 0.027 | 0.029 | (0.124) | 0.021 | 0.02 | (0.437) |
| | Forceps or vacuum | 0.099 | 0.095 | (0.022)* | 0.102 | 0.097 | (0.027)* | 0.093 | 0.091 | (0.493) |
| | Less than 3 hours | 0.018 | 0.019 | (0.228) | 0.019 | 0.019 | (0.843) | 0.015 | 0.018 | (0.051) |
| | More than 20 hours | 0.007 | 0.006 | (0.310) | 0.007 | 0.007 | (0.550) | 0.007 | 0.006 | (0.447) |
| | Labor stimulated | 0.119 | 0.115 | (0.053) | 0.106 | 0.104 | (0.375) | 0.141 | 0.133 | (0.036)* |
| | Labor induced | 0.099 | 0.095 | (0.065) | 0.092 | 0.089 | (0.289) | 0.111 | 0.105 | (0.072) |
| | Admitted on Sunday | 0.116 | 0.117 | (0.572) | 0.118 | 0.119 | (0.772) | 0.112 | 0.114 | (0.616) |
| | Admitted on Monday | 0.155 | 0.153 | (0.346) | 0.154 | 0.15 | (0.150) | 0.157 | 0.158 | (0.705) |
| | Admitted on Tuesday | 0.153 | 0.152 | (0.648) | 0.154 | 0.149 | (0.102) | 0.151 | 0.156 | (0.188) |
| | Admitted on Wednesday | 0.151 | 0.146 | (0.023)* | 0.148 | 0.145 | (0.257) | 0.155 | 0.146 | (0.026)* |
| | Admitted on Thursday | 0.152 | 0.15 | (0.427) | 0.151 | 0.15 | (0.703) | 0.152 | 0.149 | (0.435) |
| | Admitted on Friday | 0.147 | 0.146 | (0.735) | 0.147 | 0.147 | (0.916) | 0.147 | 0.145 | (0.534) |
| | Admitted on Saturday | 0.120 | 0.121 | (0.761) | 0.121 | 0.122 | (0.717) | 0.117 | 0.118 | (0.852) |
| | Missing admission day | 0.008 | 0.017 | (0.000)** | 0.007 | 0.019 | (0.000)** | 0.008 | 0.013 | (0.000)** |
| Primary Payer | Medicaid | 0.463 | 0.463 | (0.898) | 0.473 | 0.48 | (0.091) | 0.445 | 0.432 | (0.017)* |
| | Self pay/unknown | 0.041 | 0.039 | (0.182) | 0.046 | 0.046 | (0.862) | 0.033 | 0.029 | (0.056) |
| | Private | 0.480 | 0.482 | (0.458) | 0.463 | 0.457 | (0.161) | 0.509 | 0.525 | (0.004)** |
| | Other | 0.016 | 0.016 | (0.727) | 0.017 | 0.016 | (0.387) | 0.013 | 0.014 | (0.515) |
| | Government | 0.223 | 0.227 | (0.147) | 0.237 | 0.244 | (0.062) | 0.199 | 0.199 | (0.935) |
| Hospital Characteristics | Private nonprofit | 0.634 | 0.636 | (0.554) | 0.616 | 0.616 | (0.936) | 0.664 | 0.668 | (0.414) |
| | Private for-profit | 0.142 | 0.136 | (0.012)* | 0.147 | 0.14 | (0.018)* | 0.135 | 0.131 | (0.256) |
| | Observations | 45807 | 47046 | | 28898 | 29477 | | 16637 | 17283 | |

Estimates from the 40 minute sample, which includes births from 11:37pm-11:55pm & 12:05am-12:24am. Additional variables used the main analysis (not shown) include mother age indicators, father age indicators, year of birth indicators and month of birth indicators. ** = significant at 1%; * = significant at 5%.

**Table A2: Selected Covariates: 40 Minute Pooled Sample, 1991-2002**

| | Dependent Variable: | Additional Midnights | 28-Day Readmission |
|---|---|---|---|
| | | (1) | (2) |
| | Birth After Midnight | 0.255 | 0.0023 |
| | | (0.025)** | (0.0026) |
| | Birth After Midnight * Minute from cutoff | 0.001 | 0.0001 |
| | | (0.002) | (0.0002) |
| | Birth Prior to Midnight* Minute from cutoff | -0.002 | -0.0001 |
| | | (0.002) | (0.0002) |
| Pregnancy | At least one pregnancy complication | 0.021 | -0.0006 |
| Characterisitics | | (0.014) | (0.0014) |
| | 9-15 prenatal visits | -0.073 | -0.0022 |
| (<9 omitted) | | (0.020)** | (0.0017) |
| | >15 prenatal visits | 0.055 | 0.0019 |
| | | (0.029) | (0.0027) |
| | Prenatal visits missing | 0.060 | 0.0042 |
| | | (0.060) | (0.0051) |
| Mother's | Born in California | 0.023 | 0.0005 |
| Characteristics | | (0.021) | (0.0021) |
| | Born outside U.S. | 0.028 | -0.0047 |
| | | (0.023) | (0.0024)* |
| | 1st Birth | 0.232 | 0.0063 |
| | | (0.016)** | (0.0016)** |
| (Age<18 omitted) | Age = 20 | -0.057 | 0.0028 |
| | | (0.046) | (0.0047) |
| | Age = 30 | 0.219 | 0.0026 |
| | | (0.053)** | (0.0051) |
| | Age = 40 | 0.315 | 0.0119 |
| | | (0.071)** | (0.0075) |
| (College+ omitted) | High school drop out | 0.034 | 0.0011 |
| | | (0.027) | (0.0031) |
| | High school | 0.004 | -0.0005 |
| | | (0.023) | (0.0027) |
| | Some College | 0.009 | 0.0013 |
| | | (0.021) | (0.0026) |
| | Missing education data | 0.147 | 0.0016 |
| | | (0.091) | (0.0074) |
| Father's | Age = 20 | 0.080 | 0.0071 |
| Characterisitcs | | (0.076) | (0.0084) |
| | Age = 30 | 0.083 | 0.0041 |
| (Age<18 omitted) | | (0.074) | (0.0080) |
| | Age = 40 | 0.183 | 0.0055 |
| | | (0.075)* | (0.0080) |
| | Age Missing | 0.203 | 0.0044 |
| | | (0.075)** | (0.0076) |
| | High school drop out | 0.007 | -0.0026 |
| (College+ omitted) | | (0.026) | (0.0029) |
| | High school | 0.014 | -0.0030 |
| | | (0.023) | (0.0025) |
| | Some College | -0.038 | 0.0019 |
| | | (0.021) | (0.0026) |
| | Missing education data | 0.052 | 0.0014 |
| | | (0.044) | (0.0040) |
| | Mean of Dependent Variable | 1.10 | 0.044 |
| | Observations | 92853 | 92853 |

Additional characteristics included month, year, day of the week, mother's age, and father's age indicators. ** = significant at 1%; * = significant at 5%.

**Table A2 (continued): Selected Covariates: 40 Minute Pooled Sample, 1991-2002**

| | Dependent Variable: | Additional Midnights | 28-Day Readmission |
|---|---|---|---|
| | | (1) | (2) |
| Newborn Characterisitcs | Boy | 0.095 | 0.0090 |
| | | (0.013)** | (0.0013)** |
| | White | -0.070 | -0.0018 |
| | | (0.022)** | (0.0022) |
| | African American | 0.148 | -0.0061 |
| | | (0.038)** | (0.0030)* |
| | Hispanic | -0.024 | -0.0032 |
| | | (0.029) | (0.0028) |
| | Asian | -0.058 | 0.0011 |
| | | (0.030) | (0.0031) |
| Birth Characterisitics | Birthweight 2nd quintile | -0.747 | -0.0090 |
| | | (0.023)** | (0.0018)** |
| | Birthweight 3rd quintile | -0.732 | -0.0115 |
| | | (0.023)** | (0.0018)** |
| | Birthweight 4th quintile | -0.718 | -0.0108 |
| | | (0.023)** | (0.0019)** |
| | Birthweight Top quintile | -0.636 | -0.0106 |
| | | (0.024)** | (0.0019)** |
| (Gestational Age <37 weeks omitted) | Gestational age 37 <= weeks < 40 | -1.511 | -0.0251 |
| | | (0.049)** | (0.0020)** |
| | Gestational age 40 <= weeks < 42 | -1.500 | -0.0303 |
| | | (0.048)** | (0.0020)** |
| | Gestational age weeks >=42 | -1.465 | -0.0259 |
| | | (0.051)** | (0.0018)** |
| | Gestational age missing | -1.472 | -0.0214 |
| | | (0.056)** | (0.0025)** |
| | Vaginal birth after C-section | 0.058 | 0.0068 |
| | | (0.040) | (0.0046) |
| | Forceps or vacuum | 0.109 | 0.0088 |
| | | (0.021)** | (0.0025)** |
| | Less than 3 hours | -0.018 | 0.0035 |
| | | (0.056) | (0.0051) |
| | More than 20 hours | 0.262 | 0.0062 |
| | | (0.086)** | (0.0084) |
| | Labor stimulated | 0.003 | -0.0007 |
| | | (0.020) | (0.0021) |
| | Labor induced | -0.029 | 0.0042 |
| | | (0.021) | (0.0024) |
| (Sat. omitted) | Admitted on Wednesday | 0.061 | 0.0010 |
| | | (0.025)* | (0.0026) |
| | Missing admission day | 0.148 | 0.0045 |
| | | (0.067)* | (0.0065) |
| Primary Payer (private omitted) | Medicaid | 0.143 | 0.0102 |
| | | (0.017)** | (0.0018)** |
| | Self pay/unknown | -0.250 | -0.0138 |
| | | (0.031)** | (0.0030)** |
| | Other | 0.315 | -0.0012 |
| | | (0.072)** | (0.0055) |
| Hospital Characterisitcs (for-profit omitted) | Government | 0.439 | -0.0018 |
| | | (0.020)** | (0.0022) |
| | Private nonprofit | 0.080 | -0.0041 |
| | | (0.016)** | (0.0020)* |
| | Mean of Dependent Variable | 1.10 | 0.044 |
| | Observations | 92853 | 92853 |

Additional characteristics included month, year, day of the week, mother's age, and father's age indicators. ** = significant at 1%; * = significant at 5%.

**Table A3: Time of Birth and Infant Readmissions**

**Dependent Variable: 7-Day Readmission**

**A. Before 1997 Law Change**

| Model: | Local linear | Probit | Probit | Probit | IV Probit | IV Probit |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Birth After Midnight | 0.00038 | 0.00188 | 0.00174 | 0.00134 | | |
| | (0.0028) | (0.0028) | (0.0025) | (0.0015) | | |
| Birth After Midnight * 100 Minutes from cutoff | | 0.018 | 0.018 | -0.00050 | | |
| | | (0.027) | (0.015) | (0.0033) | | |
| Birth Prior to Midnight* 100 Minutes from cutoff | | -0.026 | -0.023 | 0.0033 | | |
| | | (0.018) | (0.016) | (0.0033) | | |
| Length of Stay | | | | | 0.0017 | 0.0025 |
| | | | | | (0.0054) | (0.0050) |
| | | | | | [-1.12, 2.86] | [-1.15, 2.83] |
| Full Controls | No | No | Yes | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2 hour | 2 hour | 2 hour |
| Observations | 60398 | 58365 | 58365 | 162791 | 162821 | 162821 |
| Mean of Dep. Variable Before Midnight | 0.027 | 0.027 | 0.027 | 0.026 | 0.042 | 0.042 |

**B. After1997 Law Change**

| Model: | Local linear | Probit | Probit | Probit | IV Probit | IV Probit |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Birth After Midnight | 0.00040 | -0.00052 | -0.00035 | 0.00142 | | |
| | (0.0038) | (0.0036) | (0.0032) | (0.0020) | | |
| Birth After Midnight * 100 Minutes from cutoff | | 0.0023 | -0.00008 | -0.0009 | | |
| | | (0.022) | (0.019) | (0.0044) | | |
| Birth Prior to Midnight* 100 Minutes from cutoff | | -0.0042 | -0.0042 | -0.0038 | | |
| | | (0.023) | (0.021) | (0.0045) | | |
| Length of Stay | | | | | 0.0015 | 0.0015 |
| | | | | | (0.0054) | (0.0042) |
| | | | | | [-0.88, 2.75] | [-1.04, 2.69] |
| Full Controls | No | No | Yes | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2hour | 2hour | 2hour |
| Observations | 35736 | 33920 | 33920 | 94879 | 94879 | 94879 |
| Mean of Dep. Variable Before Midnight | 0.028 | 0.028 | 0.028 | 0.029 | 0.046 | 0.046 |

40 minute sample includes births from 11:37pm-11:55pm & 12:05am-12:24am; 2 hour sample includes births from 11:02pm-11:55pm & 12:05am to 12:59am. Column (1) reports the difference in local linear regression estimates just above and below the discontinuity using a triangle kernel and a bandwidth of 20 minutes. Asymptotic standard errors in parentheses. Columns (2)-(4) report marginal effects evaluated at the mean of the covariates with robust standard errors reported in parentheses. Full controls include the controls listed in Appendix Table A1, as well as indicators for the mother's age, the father's age, the year of birth and the month of birth. Columns (5) and (6) report marginal effects from models that include a quartic in the residual from a regression of length of stay on the same variables that are included in this table. Robust standard errors in parentheses, 95% confidence interval of the bootstrapped t-statistics (500 replications) in brackets.

**Table A4: Time of Birth and 1-Year Mortality**

**Dependent Variable: 1-Year Mortality**

**A. Before 1997 Law Change**

| Model: | Local linear | Probit | Probit | Probit | IV Probit | IV Probit |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Birth After Midnight | 0.000003 | 0.00127 | 0.00026 | 0.000013 | | |
| | (0.0013) | (0.00126) | (0.000465) | (0.00035) | | |
| Birth After Midnight * 100 Minutes from cutoff | | -0.0079 | -0.0013 | -0.0002 | | |
| | | (0.0076) | (0.0029) | (0.0008) | | |
| Birth Prior to Midnight* 100 Minutes from cutoff | | -0.0078 | -0.0029 | -0.0002 | | |
| | | (0.0082) | (0.0030) | (0.0008) | | |
| Length of Stay | | | | | -0.00059 | -0.00051 |
| | | | | | (0.0028) | (0.0015) |
| | | | | | [-2.28, 1.99] | [-2.31, 2.05] |
| Full Controls | No | No | Yes | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2 hour | 2 hour | 2 hour |
| Observations | 60398 | 58365 | 58365 | 162791 | 162791 | 162791 |
| Mean of Dep. Variable Before Midnight | 0.0058 | 0.0056 | 0.0056 | 0.0055 | 0.0055 | 0.0055 |

**B. After1997 Law Change**

| Model: | Local linear | Probit | Probit | Probit | IV Probit | IV Probit |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Birth After Midnight | 0.00083 | 0.00077 | 0.00028 | 0.000016 | | |
| | (0.0015) | (0.0014) | (0.00030) | (0.00029) | | |
| Birth After Midnight * 100 Minutes from cutoff | | -0.0019 | -0.0005 | 0.0005 | | |
| | | (0.0087) | (0.0018) | (0.0007) | | |
| Birth Prior to Midnight* 100 Minutes from cutoff | | -0.0033 | -0.0022 | -0.0013 | | |
| | | (0.0096) | (0.0019) | (0.0006) | | |
| Length of Stay | | | | | -0.00092 | 0.000033 |
| | | | | | (0.0025) | (0.0011) |
| | | | | | [-1.68, 2.21] | [-1.99, 2.13] |
| Full Controls | No | No | Yes | Yes | No | Yes |
| Sample | 40 minute | 40 minute | 40 minute | 2hour | 2hour | 2hour |
| Observations | 35736 | 33157 | 33157 | 94761 | 94761 | 94761 |
| Mean of Dep. Variable Before Midnight | 0.0043 | 0.0041 | 0.0041 | 0.0048 | 0.0048 | 0.0048 |

40 minute sample includes births from 11:37pm-11:55pm & 12:05am-12:24am; 2 hour sample includes births from 11:02pm-11:55pm & 12:05am to 12:59am. Column (1) reports the difference in local linear regression estimates just above and below the discontinuity using a triangle kernel and a bandwidth of 20 minutes. Asymptotic standard errors in parentheses. Columns (2)-(4) report marginal effects evaluated at the mean of the covariates with robust standard errors reported in parentheses. Full controls include the controls listed in Appendix Table A1, as well as indicators for the mother's age, the father's age, the year of birth and the month of birth. Columns (5) and (6) report marginal effects from models that include a quartic in the residual from a regression of length of stay on the same variables that are included in this table. Robust standard errors in parentheses, 95% confidence interval of the bootstrapped t-statistics (500 replications) in brackets.

Table A5: Characteristics of Compliers

| | | Before Law Change | | | After Law Change | | |
|---|---|---|---|---|---|---|---|
| | | 40-minute Sample | | 24-hour | 40-minute Sample | | 24-hour |
| | | Complier Mean | Overall Mean | Population Mean | Complier Mean | Overall Mean | Population Mean |
| Pregnancy | At least one pregnancy complication | 0.546 | 0.545 | 0.565 | 0.683 | 0.659 | 0.664 |
| Characteristics | <9 prenatal visits | 0.202 | 0.228 | 0.208 | 0.129 | 0.155 | 0.137 |
| | 9-15 prenatal visits | 0.699 | 0.672 | 0.685 | 0.696 | 0.726 | 0.734 |
| | >15 prenatal visits | 0.078 | 0.083 | 0.091 | 0.147 | 0.097 | 0.109 |
| | Prenatal visits missing | 0.021 | 0.017 | 0.016 | 0.028 | 0.021 | 0.021 |
| Mother's | Born in California | 0.380 | 0.380 | 0.391 | 0.396 | 0.407 | 0.414 |
| Characteristics | Born outside U.S. | 0.482 | 0.472 | 0.455 | 0.490 | 0.476 | 0.460 |
| | 1st Birth | 0.385 | 0.395 | 0.391 | 0.414 | 0.400 | 0.387 |
| | Age | 26.2 | 26.6 | 27.1 | 28.3 | 27.1 | 27.8 |
| | High school drop out | 0.387 | 0.374 | 0.346 | 0.297 | 0.323 | 0.298 |
| | High school | 0.274 | 0.289 | 0.291 | 0.237 | 0.285 | 0.282 |
| | Some College | 0.177 | 0.181 | 0.191 | 0.198 | 0.186 | 0.193 |
| | College+ | 0.156 | 0.149 | 0.165 | 0.252 | 0.191 | 0.213 |
| | Missing education data | 0.006 | 0.007 | 0.007 | 0.015 | 0.015 | 0.014 |
| Father's | Age | 28.9 | 29.5 | 29.9 | 31.4 | 30.2 | 30.7 |
| Characteristics | High school drop out | 0.312 | 0.319 | 0.299 | 0.232 | 0.274 | 0.256 |
| | High school | 0.259 | 0.289 | 0.290 | 0.253 | 0.281 | 0.276 |
| | Some College | 0.176 | 0.156 | 0.163 | 0.160 | 0.154 | 0.162 |
| | College+ | 0.169 | 0.171 | 0.187 | 0.254 | 0.198 | 0.220 |
| | Missing education data | 0.083 | 0.065 | 0.061 | 0.100 | 0.094 | 0.086 |
| Newborn | Boy | 0.496 | 0.509 | 0.510 | 0.482 | 0.506 | 0.510 |
| Characteristics | White | 0.466 | 0.473 | 0.508 | 0.615 | 0.652 | 0.681 |
| | African American | 0.073 | 0.069 | 0.069 | 0.075 | 0.060 | 0.060 |
| | Hispanic | 0.299 | 0.298 | 0.260 | . | . | |
| | Asian | 0.091 | 0.091 | 0.091 | 0.124 | 0.098 | 0.095 |
| Birth | Low birthweight (<2500g) | 0.005 | 0.048 | 0.059 | 0.029 | 0.047 | 0.061 |
| Characteristics | Full Term (>=37 weeks) | 0.927 | 0.909 | 0.898 | 0.915 | 0.906 | 0.896 |
| | Vaginal birth after C-section | 0.035 | 0.028 | 0.022 | 0.023 | 0.021 | 0.015 |
| | Forceps or vacuum | 0.077 | 0.099 | 0.083 | 0.102 | 0.092 | 0.072 |
| | Less than 3 hours | 0.019 | 0.020 | 0.014 | 0.016 | 0.016 | 0.011 |
| | More than 20 hours | 0.002 | 0.007 | 0.008 | 0.009 | 0.006 | 0.007 |
| | Labor stimulated | 0.085 | 0.105 | 0.097 | 0.124 | 0.137 | 0.116 |
| | Labor induced | 0.089 | 0.091 | 0.087 | 0.101 | 0.108 | 0.109 |
| | Admitted on a weekend | 0.266 | 0.243 | 0.228 | 0.261 | 0.233 | 0.216 |
| Primary Payer | Medicaid | 0.489 | 0.477 | 0.457 | 0.425 | 0.439 | 0.417 |
| | Self pay/unknown | 0.038 | 0.046 | 0.045 | 0.009 | 0.031 | 0.027 |
| | Private | 0.459 | 0.460 | 0.481 | 0.550 | 0.517 | 0.540 |
| | Other | 0.014 | 0.016 | 0.018 | 0.016 | 0.014 | 0.016 |
| | Government | 0.222 | 0.241 | 0.215 | 0.264 | 0.200 | 0.183 |
| Hospital | Private nonprofit | 0.644 | 0.616 | 0.630 | 0.585 | 0.666 | 0.666 |
| Characteristics | Private for-profit | 0.134 | 0.143 | 0.153 | 0.152 | 0.133 | 0.149 |
| Excluded Characteristics | Cesarean Section | . | . | 0.215 | . | . | 0.240 |
| | Observations | 58375 | | 3455661 | 33920 | | 2561629 |

Compliers are newborns who are induced into having a longer stay due to a post-midnight births, where longer stays are defined as at least one additional midnight prior to the law change and at least 2 additional midnights after the law change. The means are then calculated as described in the text. The 40-minute sample is the analysis sample used in Table 3. The 24-hour population includes all births in California hospitals. The estimated fraction of compliers in the 40-minute sample is 16% prior to the law change and 17% after the law change.

Appendix Table A6: Characteristics of Compliers for the Law Change

| | | January-August: 1997 vs. 1998 | | January-August: 1997 vs. 1999 | |
|---|---|---|---|---|---|
| | | Complier Mean | Overall Mean | Complier Mean | Overall Mean |
| Pregnancy | At least one pregnancy complication | 0.712 | 0.629 | 0.716 | 0.641 |
| Characte | <9 prenatal visits | 0.091 | 0.155 | 0.075 | 0.150 |
| | 9-15 prenatal visits | 0.797 | 0.720 | 0.785 | 0.722 |
| | >15 prenatal visits | 0.109 | 0.105 | 0.125 | 0.106 |
| | Prenatal visits missing | 0.003 | 0.019 | 0.015 | 0.022 |
| Mother's | Born in California | 0.475 | 0.418 | 0.454 | 0.418 |
| Characte | Born outside U.S. | 0.355 | 0.445 | 0.388 | 0.446 |
| | 1st Birth | 0.419 | 0.384 | 0.445 | 0.387 |
| | Age | 28.7 | 27.5 | 28.7 | 27.6 |
| | High school drop out | 0.167 | 0.312 | 0.188 | 0.309 |
| | High school | 0.261 | 0.285 | 0.262 | 0.285 |
| | Some College | 0.227 | 0.195 | 0.214 | 0.195 |
| | College+ | 0.321 | 0.194 | 0.323 | 0.199 |
| | Missing education data | 0.024 | 0.014 | 0.013 | 0.012 |
| Father's | Age | 31.6 | 30.4 | 31.6 | 30.4 |
| Characte | High school drop out | 0.142 | 0.258 | 0.162 | 0.256 |
| | High school | 0.256 | 0.278 | 0.261 | 0.280 |
| | Some College | 0.193 | 0.165 | 0.180 | 0.165 |
| | College+ | 0.339 | 0.208 | 0.331 | 0.211 |
| | Missing education data | 0.070 | 0.092 | 0.065 | 0.089 |
| Newborn | Boy | 0.509 | 0.510 | 0.518 | 0.510 |
| Characte | White | 0.722 | 0.692 | 0.717 | 0.692 |
| | African American | 0.058 | 0.064 | 0.057 | 0.063 |
| | Hispanic | . | | . | |
| | Asian | 0.124 | 0.089 | 0.125 | 0.090 |
| Birth | Low birthweight (<2500g) | 0.027 | 0.061 | 0.024 | 0.060 |
| Characte | Full Term (>=37 weeks) | 0.928 | 0.893 | 0.930 | 0.893 |
| | Vaginal birth after C-section | 0.025 | 0.021 | 0.017 | 0.020 |
| | Forceps or vacuum | 0.118 | 0.085 | 0.088 | 0.081 |
| | Less than 3 hours | 0.017 | 0.013 | 0.010 | 0.012 |
| | More than 20 hours | 0.008 | 0.008 | 0.007 | 0.007 |
| | Labor stimulated | 0.161 | 0.120 | 0.142 | 0.119 |
| | Labor induced | 0.161 | 0.115 | 0.140 | 0.113 |
| | Admitted on a weekend | 0.221 | 0.217 | 0.224 | 0.218 |
| Primary P | Medicaid | 0.193 | 0.415 | 0.266 | 0.415 |
| | Self pay/unknown | 0.016 | 0.034 | 0.014 | 0.032 |
| | Private | 0.776 | 0.539 | 0.696 | 0.539 |
| | Other | 0.015 | 0.012 | 0.024 | 0.015 |
| | Government | 0.152 | 0.181 | 0.153 | 0.180 |
| Hospital | Private nonprofit | 0.743 | 0.665 | 0.701 | 0.663 |
| Characte | Private for-profit | 0.103 | 0.152 | 0.145 | 0.156 |
| Excluded Characteristics | Cesarean Section | 0.117 | 0.216 | 0.137 | 0.221 |
| | Observations | 558620 | | 553139 | |

Law change ocurred in August 1997, but exempted (some) Medicaid patients. Medicaid patients were covered by a law

## Table A7: Maternal Length of Stay & Outcomes

| | Before 1997 Law Change | | After 1997 Law Change | |
|---|---|---|---|---|
| Dependent Variable: | Additional Midnights | 28-Day Readmission | Additional Midnights | 28-Day Readmission |
| | (1) | (2) | (3) | (4) |
| Birth After Midnight | 0.297 | -0.0006 | 0.229 | 0.0015 |
| | (0.086) | (0.0012) | (0.045) | (0.0015) |
| Birth After Midnight * Minute from cutoff | 0.0004 | -0.00002 | -0.0005 | -0.00004 |
| | (0.0077) | (0.00007) | (0.0020) | (0.0001) |
| Birth Prior to Midnight* Minute from cutoff | -0.0008 | 0.0000 | 0.0022 | -0.0002 |
| | (0.0025) | (0.00008) | (0.0018) | (0.00009) |
| Observations | 57599 | 57597 | 33560 | 32869 |
| Mean of Dep. Variable Before Midnight | 0.685 | 0.008 | 0.968 | 0.008 |

Analyses uses the "40 minute" sample, which includes births from 11:37pm-11:55pm & 12:05am to 12:24am. Columns (1) and (3) are estimated with by OLS, and Columns (2) and (4) report estimates that are marginal effects from a probit model, evaluated at the mean of the covariates. All models include full controls. Robust standard errors reported in parentheses.

Appendix Table A8: Infant Outcomes Across Patient Groups

**A. Before the Law Change**

| | Additional Midnights | | | 28-Day Readmission | | | 1-year Mortality | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Coeff. On After Midnight | S.E. | Mean of Dep. Var. | Marginal effect of After Midnight | S.E. | Mean of Dep. Var. | Marginal effect of After Midnight | S.E. | Mean of Dep. Var. | Obs. |
| **Subgroup:** | | | | | | | | | | |
| Medicaid Patient | 0.25831 | (0.02984)** | 1.10 | 0.00414 | (0.00294) | 0.047 | -0.00013 | (0.00054) | 0.0062 | 77272 |
| Unmarried | 0.25165 | (0.0360)** | 1.09 | 0.00146 | (.000332) | 0.044 | 0.00017 | (0.00063) | 0.0060 | 56440 |
| For-Profit Hospital | 0.23857 | (0.03859)** | 0.79 | -0.00138 | (0.00493) | 0.044 | 0.00009 | (0.00039) | 0.0041 | 19765 |
| Cesarean Section | 0.23909 | (0.06236)** | 2.87 | 0.00376 | (0.00428) | 0.049 | -0.00208 | (0.00089)* | 0.0094 | 34442 |
| Birthweight < 3000g | 0.22649 | (0.07206)** | 1.71 | 0.00654 | (0.00479) | 0.060 | -0.00024 | (0.00148) | 0.0183 | 34675 |
| High P(Readmission\|X) | 0.21937 | (0.03615)** | 1.25 | 0.00285 | (0.00318) | 0.055 | 0.00019 | (0.00060) | 0.00944 | 77012 |
| Low P(Readmission\|X) | 0.21634 | (0.01584)** | 0.72 | 0.00264 | (0.00237) | 0.031 | -0.00022 | (0.00039) | 0.00191 | 85319 |
| Low Maternal Education | 0.18903 | (0.03297)** | 1.07 | 0.00375 | (0.00317) | 0.043 | -0.00065 | (0.00063) | 0.0060 | 60403 |
| Kaiser Hospital | 0.18820 | (0.05150)** | 0.91 | 0.00768 | (0.00510) | 0.035 | 0.0005 | (0.00062) | 0.0062 | 19616 |
| Friday/Saturday Midnight | 0.186953 | (0.064019)** | 1.29361 | -0.001307 | (0.007888) | 0.04622 | -0.003232 | (0.002233) | 0.00479 | 11666 |
| Any Labor Complication | 0.18372 | (0.02364)** | 0.97 | 0.00549 | (0.00236)* | 0.042 | 0.00018 | (0.00039) | 0.0060 | 108265 |
| Scheduled Birth | 0.16825 | (0.05151)** | 0.75 | -0.00064 | (0.00608) | 0.049 | 0.00008 | (0.00021) | 0.0037 | 15225 |
| Any Pregnancy Complication | 0.16427 | (0.02678)** | 0.98 | 0.00455 | (0.00263) | 0.044 | 0.00030 | (0.00045) | 0.0063 | 88955 |
| All Data | 0.13561 | (0.05248)** | 1.81 | 0.00210 | (0.00163) | 0.043 | -0.00025 | (0.00033) | 0.0065 | 229554 |

**B. After the Law Change**

| | Additional Midnights | | | 28-Day Readmission | | | 1-year Mortality | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Coeff. On After Midnight | S.E. | Mean of Dep. Var. | Coeff. On After Midnight | S.E. | Mean of Dep. Var. | Coeff. On After Midnight | S.E. | Mean of Dep. Var. | Obs. |
| **Subgroup:** | | | | | | | | | | |
| Medicaid Patient | 0.26693 | (0.03925)** | 1.35 | 0.00464 | (0.00415) | 0.049 | 0.00027 | (0.00046) | 0.0051 | 41412 |
| Unmarried | 0.24456 | (0.07785)** | 1.28 | 0.00486 | (0.00735) | 0.050 | -0.00019 | (0.00032) | 0.0076 | 6374 |
| For-Profit Hospital | 0.31748 | (0.05329)** | 1.07 | -0.00175 | (0.00661) | 0.052 | -0.00021 | (0.00027) | 0.0086 | 5547 |
| Cesarean Section | 0.22570 | (0.08319)** | 3.14 | 0.00909 | (0.00570) | 0.052 | -0.00027 | (0.00083) | 0.0074 | 20777 |
| Birthweight < 3000g | 0.09485 | (0.09228) | 2.05 | 0.00836 | (0.00636) | 0.063 | -0.00026 | (0.00129) | 0.0168 | 19937 |
| High P(Readmission\|X) | 0.26061 | (0.04092)** | 1.49 | -0.00017 | (0.00397) | 0.057 | -0.00026 | (0.00037) | 0.00756 | 51520 |
| Low P(Readmission\|X) | 0.22471 | (0.02251)** | 1.01 | -0.00214 | (0.00332) | 0.033 | 0.00077 | (0.00052) | 0.00226 | 27239 |
| Low Maternal Education | 0.24245 | (0.04515)** | 1.29 | 0.00656 | (0.00459) | 0.045 | 0.00061 | (0.00042) | 0.0044 | 29431 |
| Kaiser Hospital | 0.08651 | (0.05824) | 1.15 | -0.00559 | (0.00538) | 0.034 | -0.000002 | (0.00014) | 0.0057 | 11176 |
| Any Labor Complication | 0.24057 | (0.02893)** | 1.27 | -0.00004 | (0.00305) | 0.046 | 0.00010 | (0.00031) | 0.0050 | 69767 |
| Scheduled Birth | 0.27764 | (0.04104)** | 1.21 | 0.00201 | (0.00419) | 0.041 | 0.00050 | (0.00044) | 0.0038 | 30303 |
| Any Pregnancy Complication | 0.23126 | (0.03095)** | 1.28 | -0.00033 | (0.00325) | 0.048 | 0.00001 | (0.00037) | 0.0055 | 62285 |
| All Data | 0.29971 | (0.06152)** | 2.10 | 0.00020 | (0.00196) | 0.045 | 0.00029 | (0.00031) | 0.0055 | 162427 |

Analyses use the 2 hour sample, which includes births from 11:02pm-11:55pm & 12:05am to 12:59am. The additional midnight models are estimated by OLS, and the readmission and mortality columns report marginal effects from a probit model, evaluated at the mean of the covariates. Models for unmarried do not include missing observations and only include 2 years in the post-law change period. Robust standard errors reported in parentheses. All models include full controls listed in Appendix Table A1, as well as mother age indicators, father age indicators, year of birth indicators and month of birth indicators. Number of observations listed is for the mortality model, which is a lower bound on the number of observations in that row, as some cells have zero deaths. ** = significant at 1%; * = significant at 5%.

Figure A1A: Low Birthweight vs. Minute of Birth:
Before the Law Change

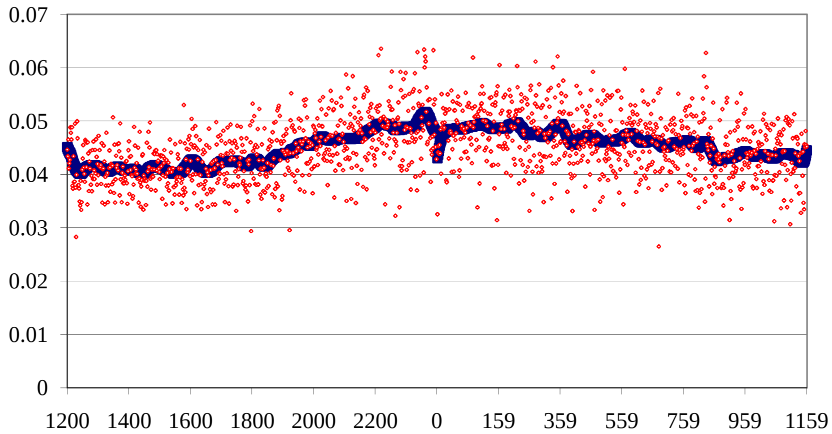



Figure A1B: P(28-Day Readmission|X) vs. Minute of Birth:
Before the Law Change

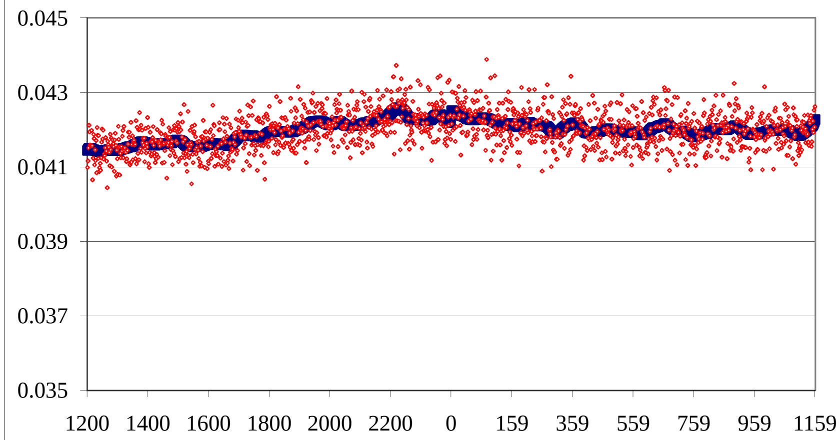



Figure A1C: Low Birthweight vs. Minute of Birth:
After the Law Change

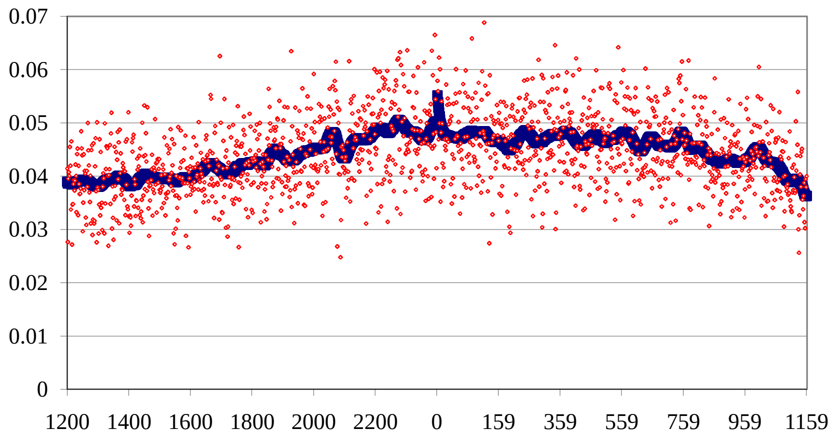
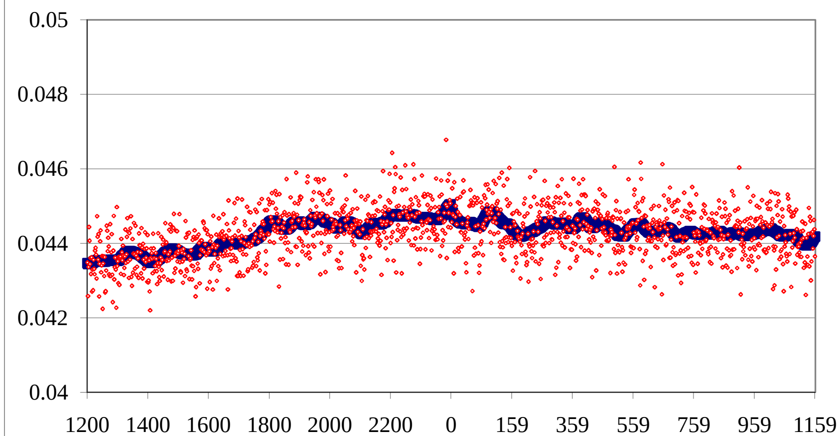



Figure A1D: P(28-Day Readmission|X) vs. Minute of Birth:
After the Law Change




Cell minute means are shown in red; local linear regression estimates (with a bandwidth of 20) are shown in blue. Predicted P(28-day readmission) used a probit model and full controls.

Figure A2A: Raw Length of Stay: Before 1997 Law Change

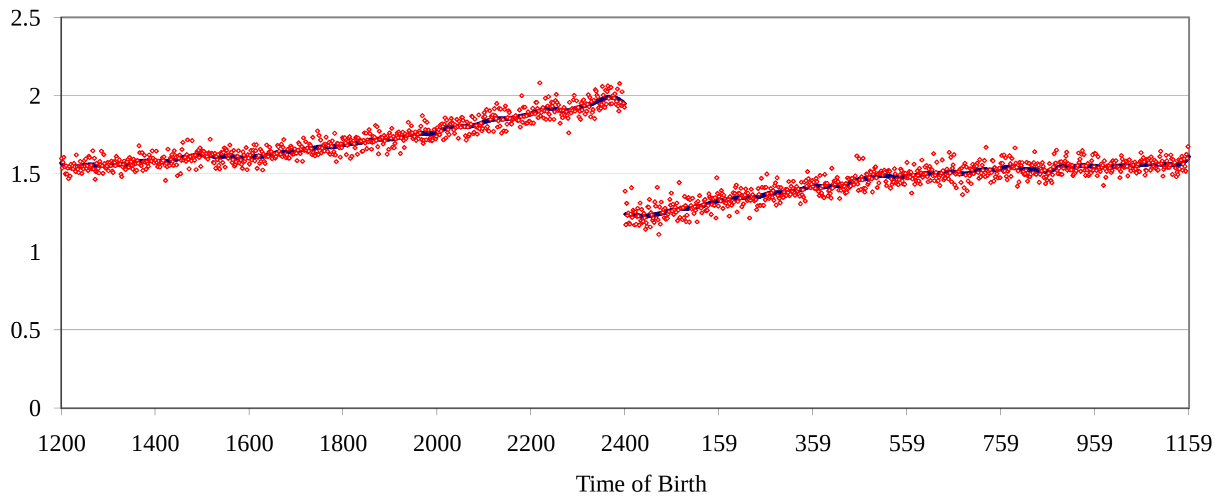
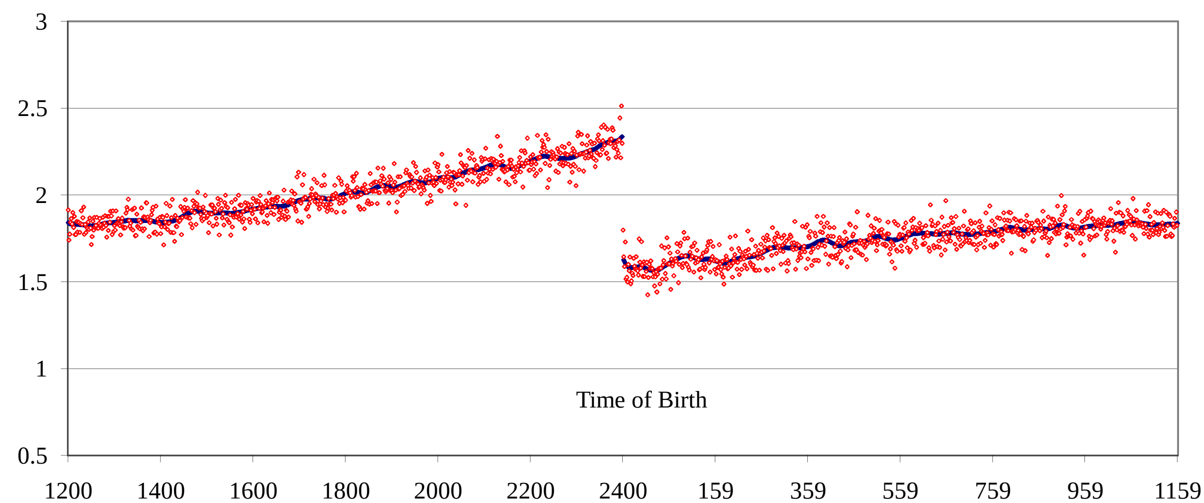



Figure A2B: Raw Length of Stay: After 1997 Law Change




Raw length of stay is the number of midnights in care. Points represent means within 1-minute intervals from 12:00 noon to 11:59am. Lines represent local linear regressions, h=20. The intercepts between Figures A2A and A2B differ so that they have the same scale.

Figure A3A: Bandwidth of 10:
Additional Midnights Before Law Change

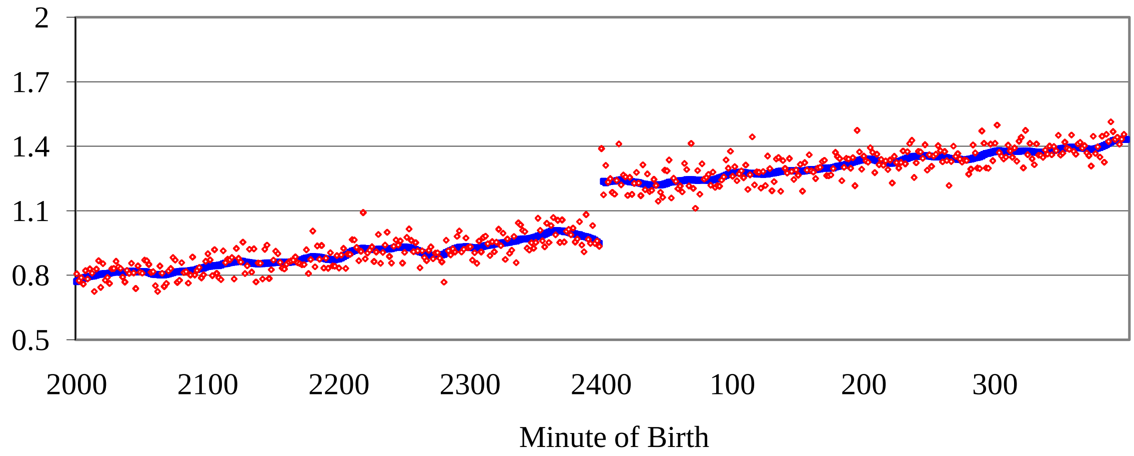



Figure A3B: Bandwidth of 10:
Additional Midnights After Law Change

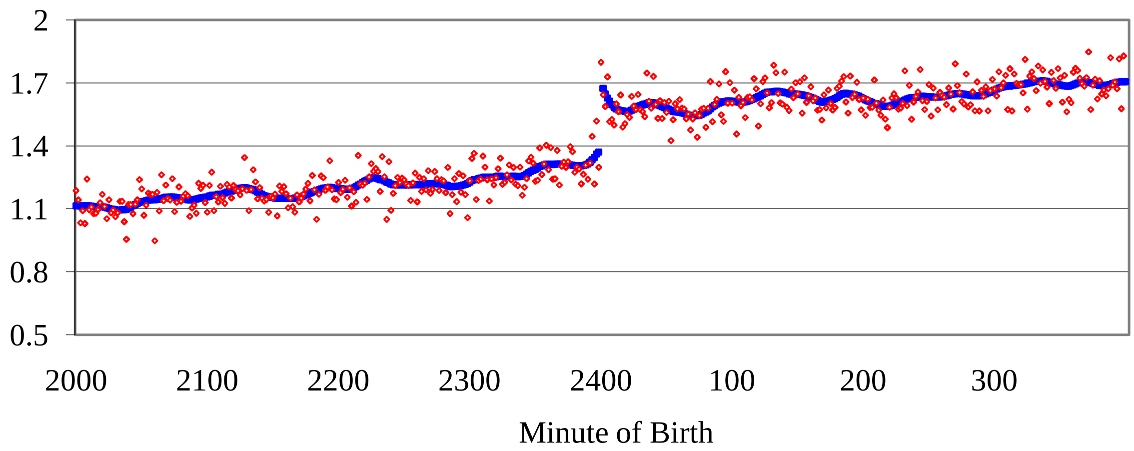


Figure A4A: 1 or More Additional Midnights: Before Law Change

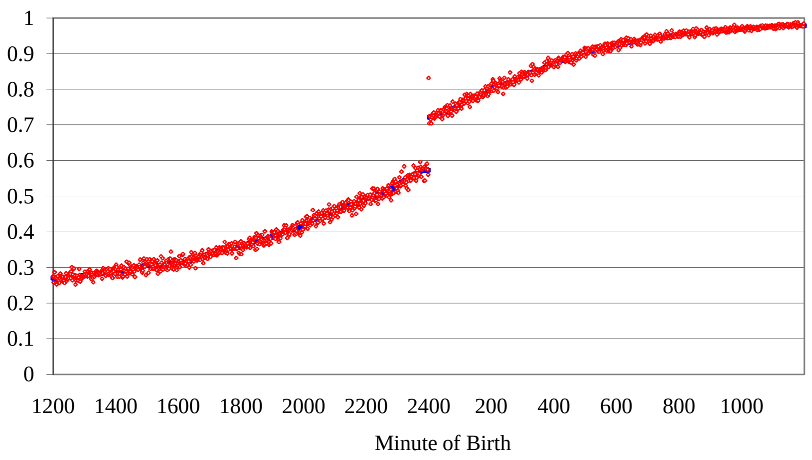



Figure A4B: 2 or More Additional Midnights: Before Law Change

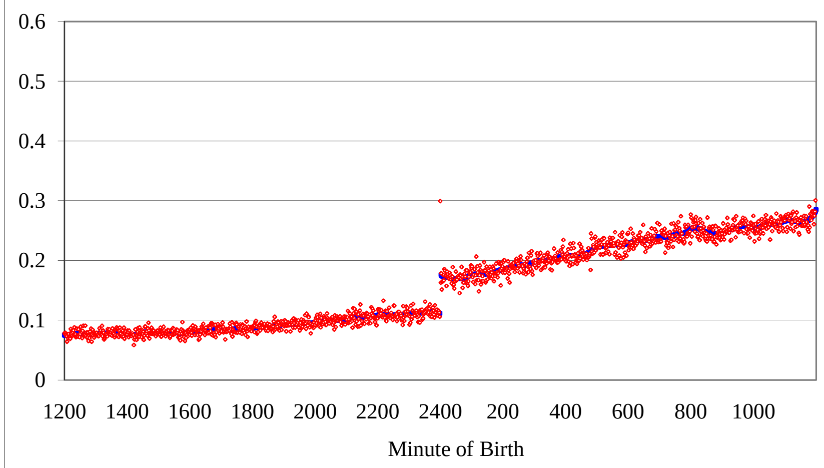



Figure A4C: 1 or More Additional Midnights: After Law Change

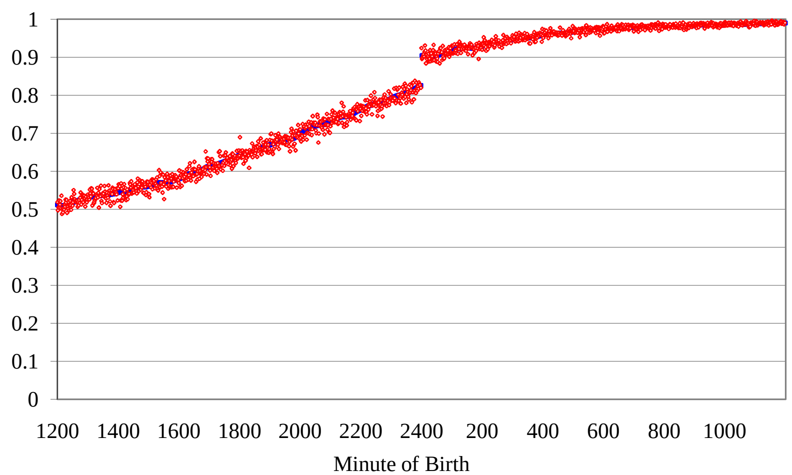
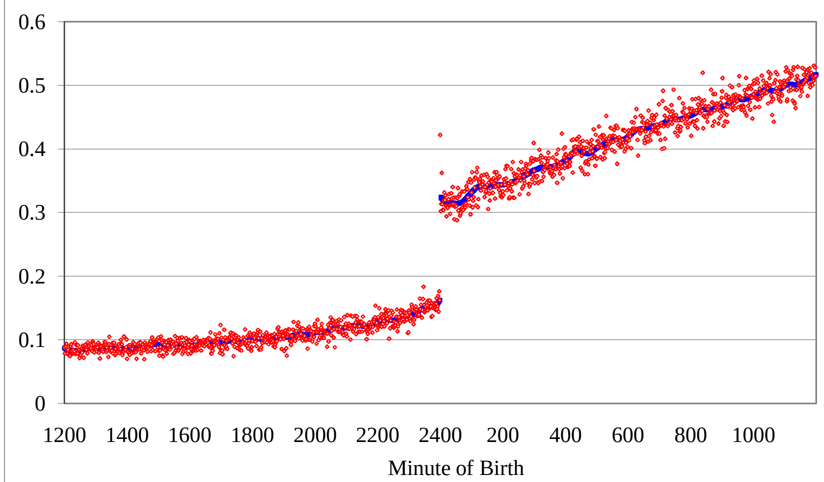



Figure A4D: 2 or More Additional Midnights: After Law Change




Number of additional midnights is the number of midnights for those born on or after midnight and before noon, while the number of additional midnights is measured as the number of midnights minus one for children born after noon and before midnight. Points represent means within 1-minute intervals from 12:00 noon to 11:59am. Lines represent local linear regressions, h=20

Figure A5A: 7-Day Readmission Rate: Before Law Change

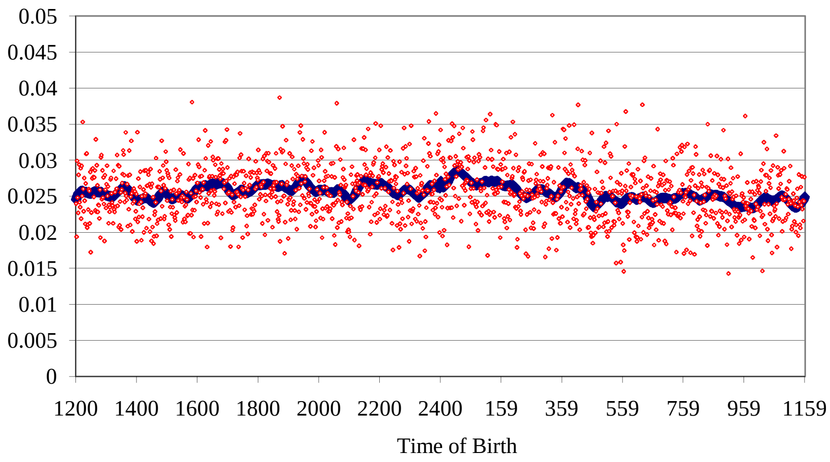



Figure A5B: Infant Mortality Rate: Before Law Change

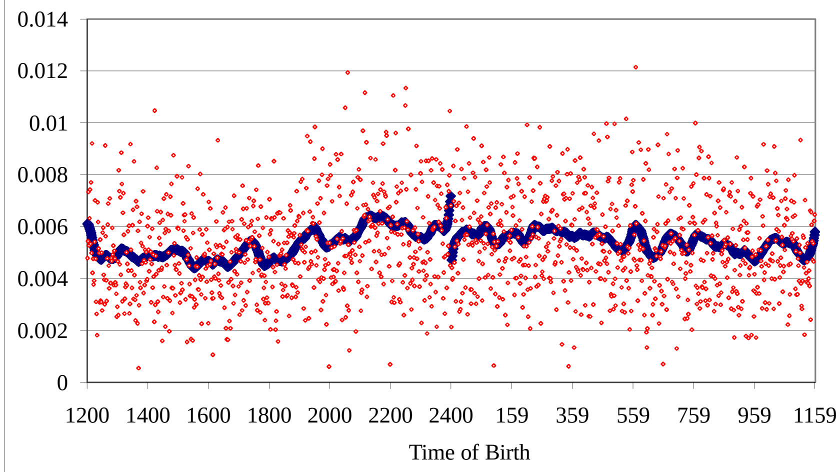



Figure A5C: 7-Day Readmission Rate: After Law Change

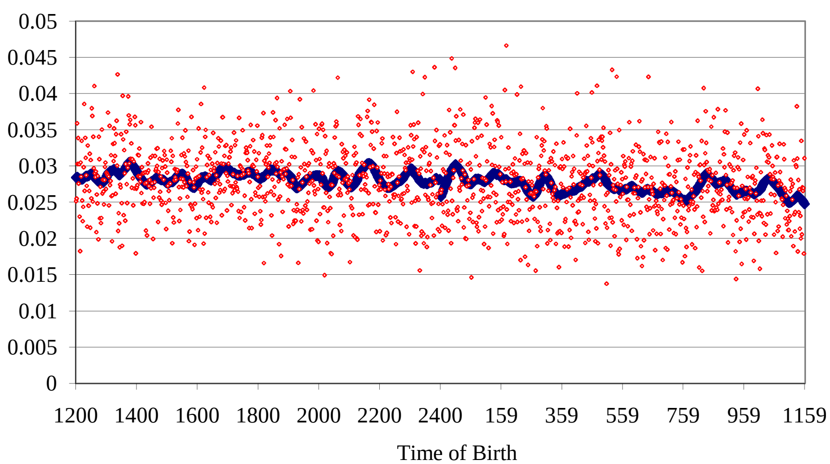
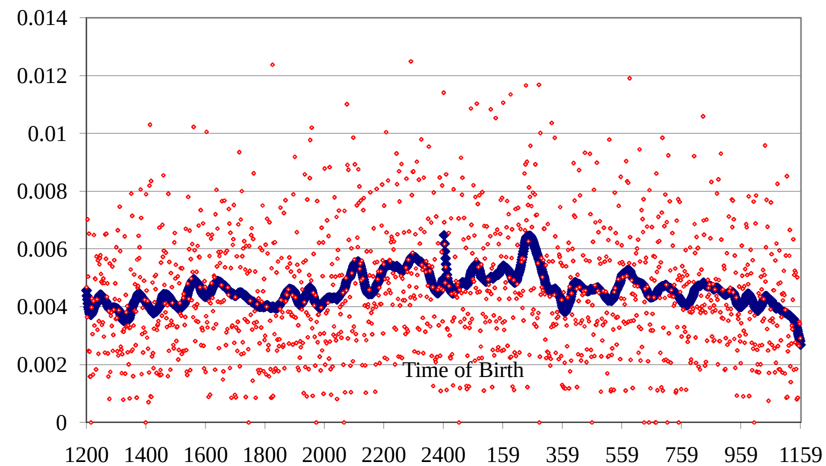



Figure A5D: Infant Mortality Rate: After Law Change




7-day measures consider 7 days since the midnight prior to those born between midnight and noon, and 7 days since the following midnight for those born between noon and midnight.
Points represent means within 1-minute intervals from 12:00 noon to 11:59am. Lines represent local linear regressions, h=20

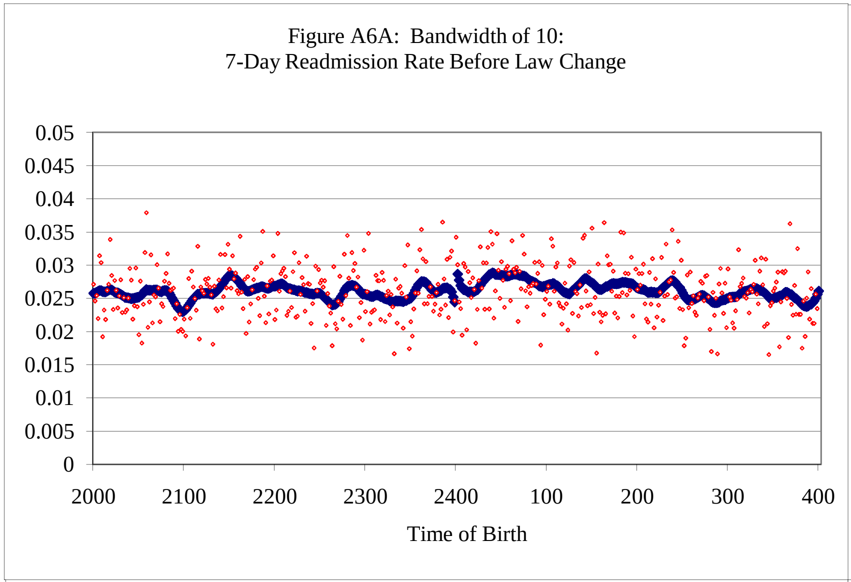


Figure A6A: Bandwidth of 10:
7-Day Readmission Rate Before Law Change
0.05
0.045
0.04
0.035
0.03
0.025
0.02
0.015
0.01
0.005
0
2000
2100
2200
2300
2400
100
200
300
400
Time of Birth

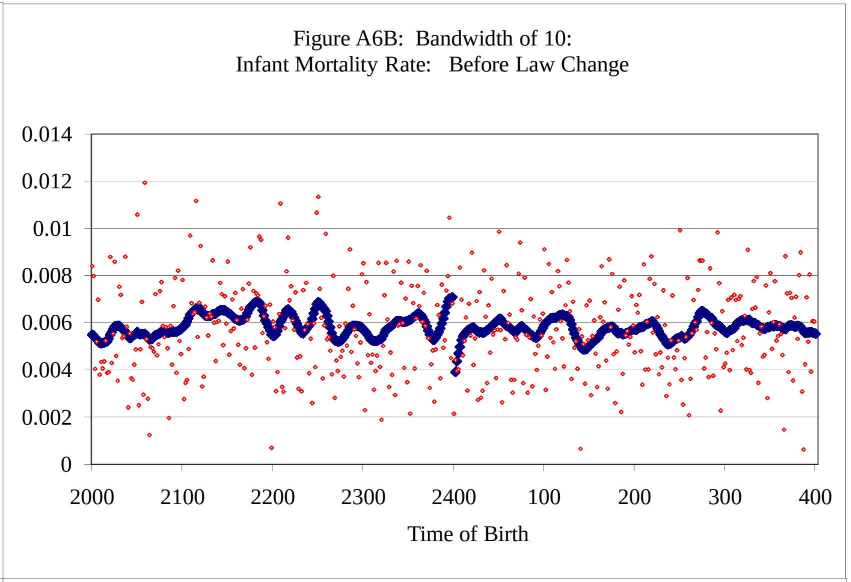


Figure A6B: Bandwidth of 10:
Infant Mortality Rate: Before Law Change
0.014
0.012
0.01
0.008
0.006
0.004
0.002
0
2000
2100
2200
2300
2400
100
200
300
400
Time of Birth

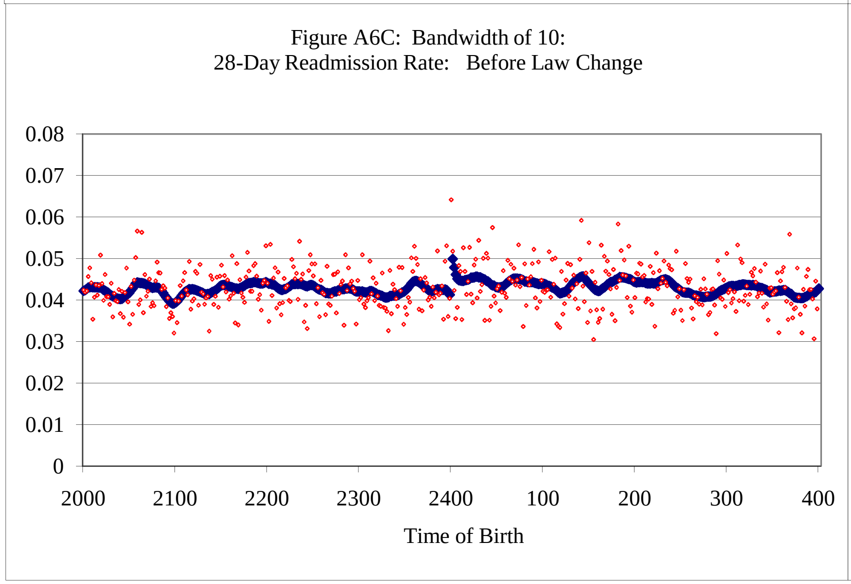


Figure A6C: Bandwidth of 10:
28-Day Readmission Rate: Before Law Change
0.08
0.07
0.06
0.05
0.04
0.03
0.02
0.01
0
2000
2100
2200
2300
2400
100
200
300
400
Time of Birth

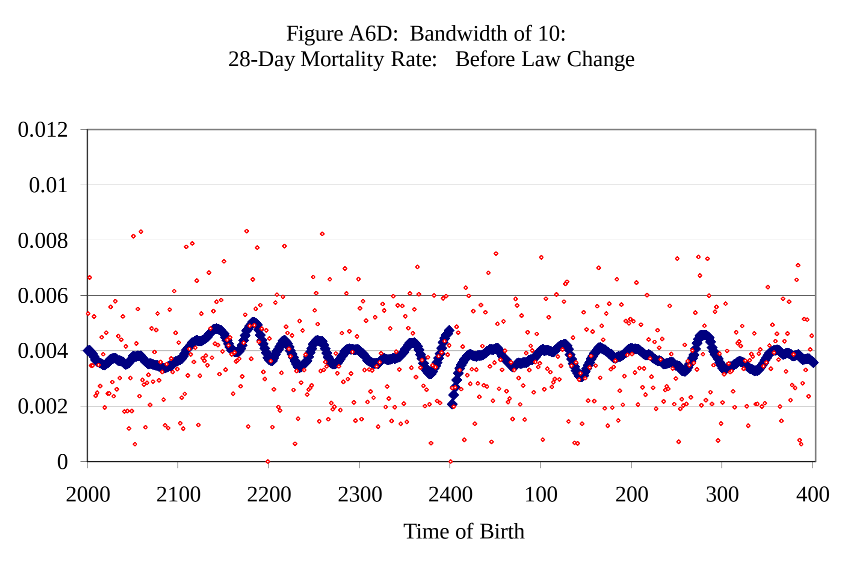


Figure A6D: Bandwidth of 10:
28-Day Mortality Rate: Before Law Change
0.012
0.01
0.008
0.006
0.004
0.002
0
2000
2100
2200
2300
2400
100
200
300
400
Time of Birth

Figure A7A: Bandwidth of 10:
7-Day Readmission Rate: After Law Change

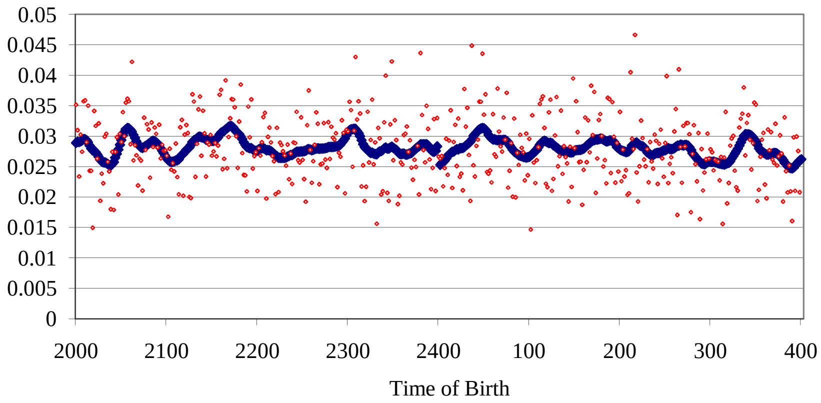



Figure A7B: Bandwidth of 10:
Infant Mortality Rate: After Law Change

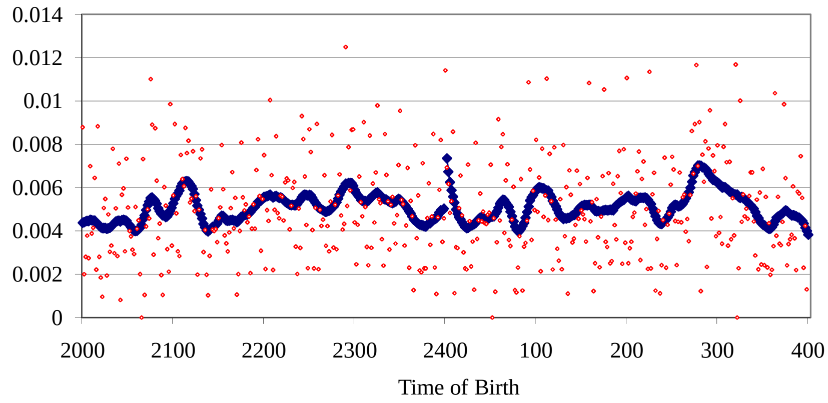



Figure A7C: Bandwidth of 10:
28-Day Readmission Rate: After Law Change

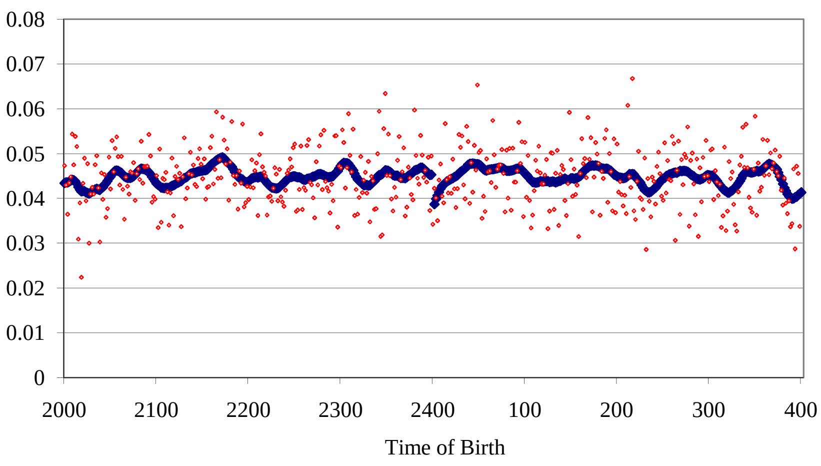
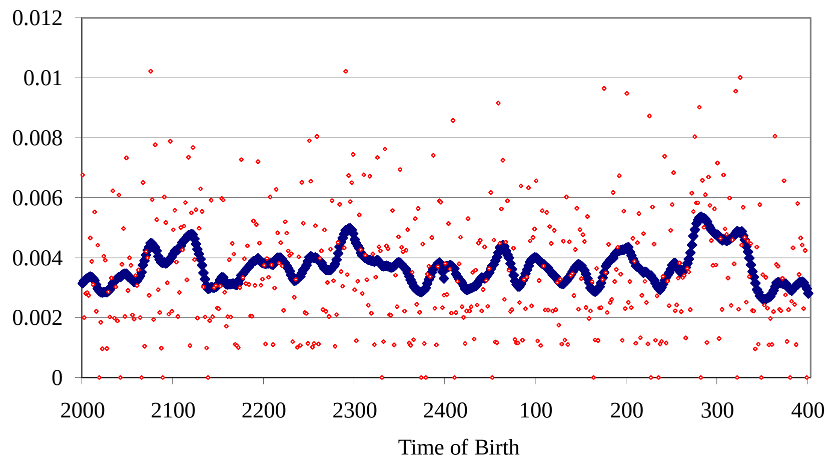



Figure A7D: Bandwidth of 10:
28-Day Mortality Rate: After Law Change




28-day measures consider 28 days since the midnight being compared. Points represent means within 1-minute intervals, lines represent local linear regressions, h=10

Figure A8A: 3-28 Day Readmissions vs. Minute of Birth:
Before the Law Change

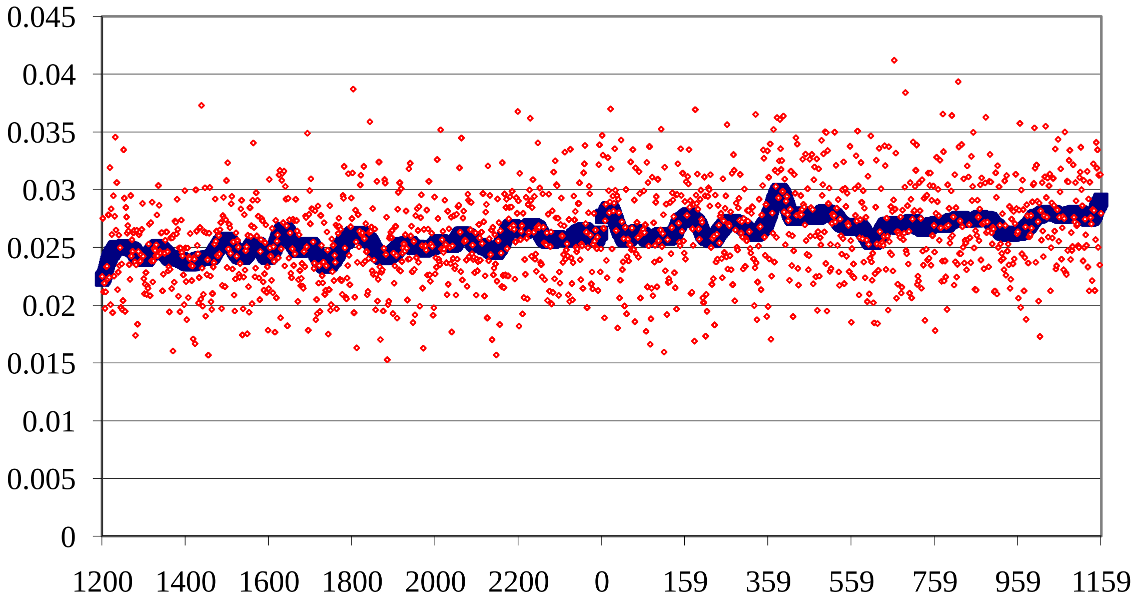
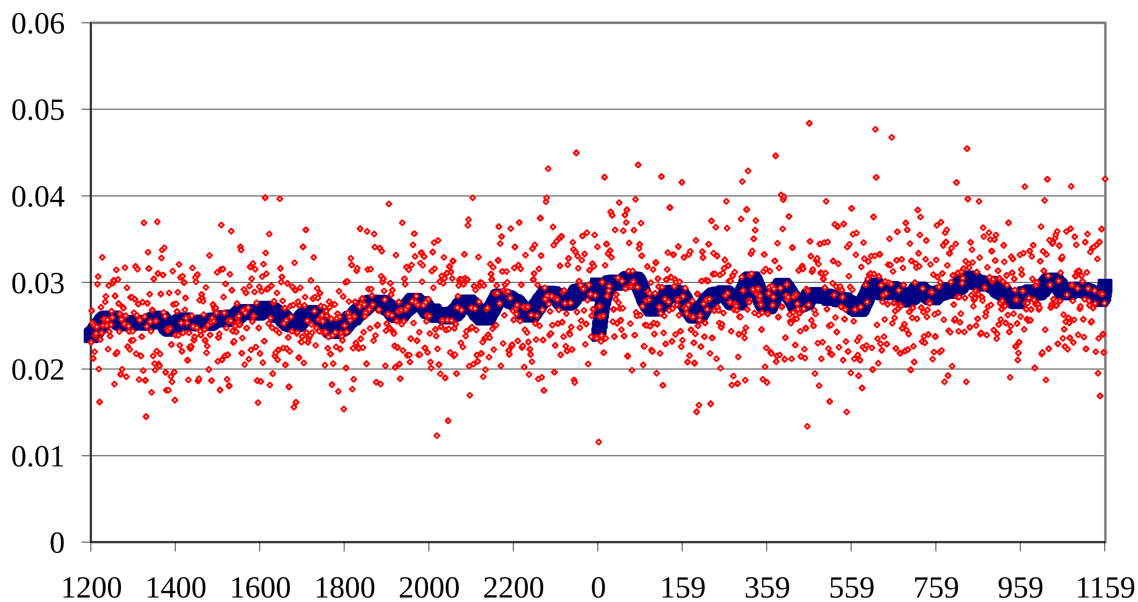



Figure A8B: 3-28 Day Readmissions vs. Minute of Birth:
After the Law Change




28-day measures consider 28 days since the midnight being compared. Points represent means within 1-minute intervals, lines represent local linear regressions, h=20