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RESEARCH ARTICLE

The Impact of Obstetric Unit Closures on Maternal and Infant Pregnancy Outcomes

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Objective. To define the association between large-scale obstetric unit closures and relative changes in maternal and neonatal outcomes.

Data Sources/Study Setting. Birth and death certificates were linked to maternal and neonatal hospital discharge records for all births between January 1, 1995 and June 30, 2005 in Philadelphia, which experienced the closure of 9 of 19 obstetric units between 1997 and 2005, and five surrounding counties and eight urban counties that did not experience a similar reduction in obstetric units.

Design. A before-and-after study design with an untreated control group compared changes in perinatal outcomes in Philadelphia to five surrounding control counties and eight urban control counties after controlling for case mix differences and secular trends (N = 3,140,782).

Results. Relative to the preclosure years, the difference in neonatal mortality (odds ratio (OR) 1.49, 95 percent CI 1.12–2.00) and all perinatal mortality (OR 1.53, 95 percent CI 1.14–2.04) increased for Philadelphia residents compared with both control groups between 1997 and 1999. After 2000, there was no statistically significant change in any outcome in Philadelphia county compared with the preclosure epoch.

Conclusions. Obstetric unit closures were initially associated with adverse changes in perinatal outcomes, but these outcomes ameliorated over time.

Key Words. Neonatal mortality, obstetric unit closures, bed supply

Between 1997 and 2005, 9 of 19 obstetric units in Philadelphia County closed, resulting in a 40 percent reduction in the number of obstetric beds (Maternity Care Coalition 2005). The literature on the reduction in hospital services largely studies potential reasons for hospital closures. These studies suggest that closures occur in highly competitive health care markets (Lee and Alexander 1999; Mullner et al. 1989), and the hospitals most at-risk for closing were smaller and for profit (Buchmueller, Jacobson, and Wold 2006; Ciliberto and Lindrooth 2007; Deily, McKay, and Dorner 2000; Lillie-Blanton et al. 1992; Lindrooth, Lo Sasso, and Bazzoli 2003). In Philadelphia, primary data suggest that increased fixed costs over the postclosure period, primarily increased malpractice premiums (Anonymous 2007; Burling 2007; George 2001, 2002, 2005, 2008a, 2009; McCullough 2003; Treaster 2002) and reduced reimbursement for obstetric care (Burling 2007; George 2008a,b, 2009; McCullough 2003) were the primary reason for closures. Most investigations into the effect of hospital closures are limited to the closure of rural hospitals (Rosenbach and Dayhoff 1995) or single hospitals within a large metropolitan area. These studies show conflicting results on the impact of hospital closures on patient outcomes. Even though obstetric units are the units most commonly eliminated by individual hospitals when reducing services (Kirby et al. 2006), the effects of a large urban obstetric service reduction on perinatal outcomes have not been systematically examined in the literature.

The goal of this study was to determine the effect of a large-scale reduction in obstetric bed supply on maternal and neonatal outcomes, using a before-and-after study design with an untreated control group to control for secular trends and changes in patient characteristics. Based on prior literature, we hypothesized that some outcomes, such as neonatal mortality, neonatal complications, and maternal complications would worsen as more units closed. These outcomes are most likely to change with changes in either access of hospital care or the timeliness of care delivery. Other outcomes, such as the delivery of infants between 23 and 32 weeks gestation, would not change because multiple studies suggest that medical management cannot reduce the risk of a delivery under 32 weeks gestation

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by mothers in premature labor (American College of Obstetricians and Gynecologists 2003), and thus timeliness of care may not affect these outcomes.

METHODS

Study Design and Choice of Control Groups

This study used a before-and-after study design with an untreated control group (Card 1990; Shadish, Cook, and Campbell 2002; Volpp et al. 2007a,b) to determine the association between obstetric unit closures in Philadelphia and changes in maternal and neonatal outcomes between a preclosure period (1995–1996) and three separate postclosure periods (1997–1999, 2000–2002, and 2003–2005). This study design answers questions where a randomized trial is not feasible, and the reasons for obstetric unit closures are not strongly associated with factors that both (1) changed more in Philadelphia, relative to the control group, between the preclosure and postclosure epoch, and (2) independently affect perinatal outcomes (Beasley and Case 2000). This study design controls for secular trends, stable differences between Philadelphia and the control counties over time, and differences in changes in each county's case mix.

The choice of control counties is important for this study design. We adopted the framework of Rosenbaum and Campbell (Campbell 1969; Rosenbaum 1987) in choosing two separate control groups for this study. The results from an observational study are strengthened if we find similar results after choosing two control groups with different similarities to Philadelphia. The first control group was a combination of five counties surrounding Philadelphia (Bucks, Montgomery, Delaware, Chester, and Berks). These counties operated under similar state and insurance regulations to Philadelphia during the 11-year time frame of the study, and these counties have a similar medical culture to Philadelphia county because of their geographic proximity. However, these counties have different obstetric systems to Philadelphia, such as large rural communities without an obstetric unit or many smaller urban areas with only a single obstetric unit. To address these differences, we selected a second control group made up of control counties that are similar to Philadelphia in terms of urban location, educational achievement, insurance status, and obstetric network (Table 1). The eight counties that made up this urban control group were Alameda, CA; Allegheny, PA; Lehigh, PA; Los Angeles, CA; Sacramento, CA; San Diego, CA;

Table 1: Demographics of Study Population

	Philadelphia County	*Five Surrounding Control Counties*	*Eight Urban Control Counties*
No. of deliveries	155,261	269,570	2,985,521
% Deliveries <1,500 g	2.32%	1.28%	1.19%
% Deliveries ≤ 32 weeks GA	3.56%	2.07%	2.27%
% Deliveries via C-section	24.27%	25.53%	25.67%
Race			
White	30.02%	75.75%	32.41%
Black	42.51%	7.47%	8.07%
Asian	3.50%	2.01%	11.49%
Hispanic	9.72%	6.49%	44.36%
Other	14.25%	8.28%	3.66%
Insurance			
Fee for service	7.03%	19.34%	3.99%
HMO	35.58%	53.07%	49.46%
Medicaid/Medicare	42.46%	13.68%	42.51%
Other	13.03%	12.43%	1.02%
Uninsured	1.09%	0.99%	2.99%
Missing	0.80%	0.49%	0.03%
Mother's education level			
At least high school diploma	72.76%	91.05%	69.34%
No high school diploma	23.49%	7.11%	28.90%
Missing	3.75%	1.85%	1.76%
Cormorbidities			
Chronic hypertension	1.40%	0.91%	0.59%
Diabetes mellitus	0.78%	0.54%	0.76%
Gestational diabetes	3.19%	3.54%	4.36%
Renal disease	0.17%	0.11%	0.09%

Note. GA, gestational age.

San Francisco, CA; and San Jose, CA. These counties, though, are in either a different part of Pennsylvania or in another state. Thus, the secular trends in perinatal outcomes may differ in these counties. By comparing the change in outcomes in Philadelphia county to two dissimilar control populations, we learn more from the observational study than by comparing with a single control population.

Data Population and Sources

We obtained live birth and fetal death certificates from all in-hospital deliveries occurring in Pennsylvania and California between January 1, 1995 and June 30, 2005. Each state's department of health linked these birth certificates to death certificates using name and date of birth, and then deidentified the records. We then matched over 98 percent of live birth or fetal death certificates to maternal and newborn hospital records using prior methods (Herrchen, Gould, and Nesbitt 1997; Phibbs et al. 2007). Over 80 percent of the unmatched live birth or fetal death certificate records were missing the delivery hospital, suggesting a birth at home or a birthing center. The unmatched records had similar gestational age and racial/ethnic distributions to the matched records. The departments of health in California and Pennsylvania and the IRB at the Children's Hospital of Philadelphia approved this study.

Mothers residing in Philadelphia, the five surrounding control counties, and the eight urban control counties were identified using the maternal county of residence on the birth certificate, and validated with the maternal zip code from the hospital record. Birth records were excluded if they had a gestational age less than 23 weeks or greater than 44 weeks, a birth weight less than 400 g or greater than 8,000 g, or if the birth weight was more than 5 standard deviations from the mean birth weight for the recorded gestational age in the cohort (Parker and Schoendorf 2002). Initially, 3,195,062 birth records were identified; 54,280 met the exclusion criteria, leaving 3,140,782 births in the final cohort.

Definition of Study Outcomes

Six outcome measures were used in this study that either (1) could be affected by changes in the delivery hospital or (2) are common outcomes measures used by public health departments to monitor perinatal care. Mortality is a common perinatal outcomes measure. This was measured in two ways. *Neonatal deaths* were defined as any death during the initial birth hospitalization and were determined from death certificate records. In addition, we examined *fetal deaths* in two ways. First, we used death certificate data to count all fetal deaths in each county with either a gestational age of at least 23 weeks or a birth weight of 400 g or more. Second, we counted only those fetal deaths that met a prior definition of a preventable fetal death that may be influenced by care delivered at the hospital (Phibbs et al. 2007). These deaths were identified by a previously constructed algorithm of ICD-9CM codes and birth certificate data. Complications around the time of delivery have also been used as adjunct outcome measures to mortality, as mortality rates are typically very low. We constructed a composite of *neonatal complications* that may be affected by a delay in delivery as infants with any one of the following complications: asphyxia (ICD-9CM codes 768.5, 768.6, 768.9), birth trauma (ICD-9CM codes 767.2, 767.4–767.9), or seizures (ICD-9CM codes 779.0, 780.3x). We constructed a composite of *maternal complications* that may be increased with a delay in delivery or a change in the quality of care at the remaining hospitals. This composite included women with any one of the following complications: wound infection (ICD-9CM codes 674.1x, 674.2x, 674.3x), postdelivery hemorrhage (ICD-9CM codes 641.3x, 641.8x, 641.9x, 660.0x, 660.1x, 660.2x, 660.3x, 667.1x), and blood transfusion (ICD-9CM procedure code 99.0x), as a proxy for severe postdelivery hemorrhage. Finally, there are common measures of perinatal care that may or may not be influenced by either the closure of the obstetric hospital or even obstetrical care. However, they are commonly used by public health departments to monitor obstetric care. These measures include *preterm deliveries*, defined as a delivery between 23 and 32 weeks, 23 and 37 weeks, or 34–37 weeks gestational age on the birth certificate; and deliveries via *Cesarean section*, identified from an ICD-9CM code of 669.7x in the maternal delivery record or a notation of a Cesarean section delivery in the birth certificate. For birth certificate records missing gestational age, we multiply imputed gestational age when it was the dependent variable in a multivariable model.

Definition of Covariate Variables

We included specific covariate variables in a risk-adjustment model based on their association with one or more study outcomes. The final models included birth weight, grouped into 250–500 g strata, because lower birth weight infants have an increased risk of death or complications; maternal sociodemographic factors, such as race, age, education, and insurance status; maternal comorbid conditions listed in the appendix, which may increase the risk of each outcome measure listed above; an interaction term between gender and birth weight, as female infants tend to have lower birth weights than male infants; and 49 congenital anomalies grouped by affected organ system, listed in the appendix, which have previously been identified as being associated with higher risk of neonatal and fetal death (Phibbs et al. 2007).

Data Analysis

For each outcome measure, we constructed two separate regression models: (1) unadjusted models that included county, treatment epoch, and an interaction term between each postclosure treatment epoch and Philadelphia county, and (2) multivariable logistic regression models that adjusted for patient-level factors, secular trends, and differences between the counties during the preclosure time period. The coefficients of interest were the interaction term between each postclosure treatment epoch and the Philadelphia County indicator variable. This interaction term measured the relative effect of obstetric unit closures on each outcome, after controlling for patient-level factors, secular trends, and differences between Philadelphia and the control counties that existed before the obstetric closures began. To adjust for hospital-level clustering of each outcome, we calculated standard errors using the robust methods of Huber-White (Angrist and Pischke 2008; Hansen 2007; Huber 1967; Localio et al. 2001; White 1980). All statistical analyses reported two-tailed P values with a statistical significance level of 5 percent.

Based on a total population of 3.14 million deliveries, and 155,261 in Philadelphia, we had 90 percent power to detect an odds ratio of 1.27 if the outcome occurred 0.5 percent of the time in the control group in both the before and after period and in the treatment group in the before time period (Demidenko 2008). This detectable odds ratio decreased to 1.20 if the outcome occurred 1 percent of the time in the control group. We assumed that the effect of clustering by county would increase the variance by 10 percent.

With two separate control groups we conducted a series of analyses. First, the multivariable models compared Philadelphia county to the control groups combined into one group. Statistically significant interaction terms were then further analyzed by comparing Philadelphia county to each of the two control groups separately. Only those outcomes and treatment epochs that were statistically significant for both the control groups combined into one group and for each control group examined separately were reported as statistically significant. Given the large size of Los Angeles county in this analysis, we also performed a sensitivity analysis where we removed this county from the urban control population and repeated the analyses.

RESULTS

Demographic information for all pregnancies included in the study is shown in Table 1. In 1995, there were 19 hospitals in Philadelphia with obstetric units. Between 1997 and 2005, Philadelphia lost nine obstetric units that delivered 30.75 percent of the county's births in 1995 and 1996. These hospitals were primarily private and localized to the Northwest and Northeast areas of the city. Closed hospitals typically had between 500 and 1,000 deliveries per year, delivered prenatal care to women in their local communities, and few had neonatal intensive care units. The control counties experienced the closure of 1–2 maternity wards that delivered 1.43–4.78 percent of the births in 1995 and 1996. In the year preceding their closure, only 1 of the 9 closed obstetric units in Philadelphia had a delivery volume less than 80 percent of its 1995–1996 average volume, whereas four of the nine units had a volume over 100 percent of its 1995–1996 average volume. Between 1995–1996 and 2005, the average number of Philadelphia county residents delivering at a Philadelphia obstetrics unit that remained open throughout the study period increased by 37 percent, whereas the percentage of Philadelphia county patients delivering at a hospital outside of the county increased by 23 percent. During the preclosure years, we did not detect a statistically significant difference in risk-adjusted rates of mortality, neonatal complications, or maternal complications between those hospitals that later closed between 1997 and 2005 and those that remained open.

Univariable Analyses

The unadjusted pregnancy outcomes in Philadelphia and the combination of the two control groups between 1995 and 2005 are shown in Table 2. Mortality rates, primarily neonatal deaths, were increased in Philadelphia by 18.1 percent during the first postclosure epoch compared with 1995–1996, the preclosure epoch. This increase was in contrast to the 6.1 percent decline in neonatal mortality seen in the combined control counties ($p = .04$). Cesarean section rates took longer to increase in Philadelphia compared with the combined control counties, with a 0.2 percent decrease in Philadelphia over the first postclosure epoch compared to a 6.8 percent increase in the combined control counties ($p = .02$). After 1999, the rate of Cesarean sections increased at a similar rate in both Philadelphia and the combined control counties. The change in 34–37 weeks gestational age births was lower in Philadelphia compared with the combined control counties in the 1997–1999 epoch (5.9 percent decline compared to a 3.5 percent increase in the combined control counties, $p = .006$). There were no significant changes in the rate of fetal death, neonatal complications, or maternal complications between Philadelphia and the combined control groups in any epoch.

Multivariable Analyses

Combined Control Groups. In multivariable analyses (Table 3), the relative difference in neonatal (odds ratio (OR) 1.49, 95 percent CI 1.12–2.00) and

Table 2: Relative Change in the Rate of Specific Perinatal Outcomes in Philadelphia and Combined Control Counties, 1995–2005, per 1,000 Births*

	1995–1996		*1997–1999*		*2000–2002*		*2003–2005*	
	Philadelphia	*Control Counties*	*Philadelphia*	*Control Counties*	*Philadelphia*	*Control Counties*	*Philadelphia*	*Control Counties*
Newborn mortality								
Neonatal deaths	5.63	4.42	**6.65 (18.1%)**	**4.15 (−6.1%)**	4.76 (−15.5%)	3.85 (−12.9%)	5.15 (−8.5%)	3.65 (−17.4%)
p-value			**.04**		.79		.48	
All fetal deaths	2.67	3.81	3.24 (21.3%)	3.51 (−7.9%)	2.82 (5.6%)	3.17 (−16.8%)	3.16 (18.4%)	2.91 (−23.6%)
p-value			.10		.23		.14	
All fetal deaths + neonatal deaths	8.29	8.21	**9.87 (19.1%)**	**7.64 (−6.9%)**	7.57 (−8.7%)	7.00 (−14.7%)	8.30 (0.1%)	6.56 (−20.1%)
p-value			**.02**		.59		.18	
Possibly preventable fetal deaths†	0.75	1.10	0.95 (26.7%)	1.09 (−0.9%)	0.83 (10.7%)	0.94 (−14.5%)	0.83 (10.7%)	0.90 (−18.2%)
p-value			.37		.35		.31	
Possibly preventable fetal + neonatal deaths†	6.37	5.51	7.59 (19.2%)	5.23 (−5.1%)	5.59 (−12.2%)	4.78 (−13.2%)	5.97 (−6.3%)	4.55 (−17.4%)
p-value			.06		.97		.40	
Cesarean sections								
All deliveries	211	218	**210.4 (−0.2%)**	**232.6 (6.8%)**	256.2 (21.5%)	265.3 (21.8%)	290.2 (37.6%)	307.6 (41.2%)
p-value			**.03**		.73		.28	

continued

Table 2. *Continued*

	1995–1996		1997–1999		2000–2002		2003–2005	
	Philadelphia	*Control Counties*	*Philadelphia*	*Control Counties*	*Philadelphia*	*Control Counties*	*Philadelphia*	*Control Counties*
All deliveries with GA 37–42 weeks	188.4	199.1	**188.0 (−0.2%)**	**212.1 (6.6%)**	231.3 (22.8%)	243.9 (22.5%)	263.2 (39.7%)	284.4 (42.9%)
p-value			**.04**		.83		.33	
Premature delivery								
23–32 weeks preterm	35.0	20.8	33.8 (−3.4%)	21.2 (1.9%)	32.7 (−6.6%)	20.7 (−0.5%)	35.1 (0.3%)	21.7 (4.3%)
p-value			.40		.43		.73	
23–37 weeks preterm	212.8	177.8	**201.4 (−5.4%)**	**183.6 (3.3%)**	217.5 (2.2%)	184.9 (4.0%)	219.3 (3.1%)	194.6 (9.4%)
p-value			**.01**		.65		.39	
34–37 weeks preterm	168.2	148.6	**158.3 (−5.9%)**	**153.8 (3.5%)**	174.1 (3.5%)	155.7 (4.8%)	174.0 (3.4%)	163.4 (10.0%)
p-value			**.006**		.72		.33	
Complications								
Any infant complication	23.1	18.0	25.0 (8.2%)	16.1 (−10.4%)	16.9 (−26.8%)	13.2 (−26.6%)	18.2 (−21.3%)	12.1 (−32.4%)
p-value			.28		.99		.43	
Any maternal complication	33.6	29.4	36.1 (7.2%)	34.1 (16.2%)	41.1 (22.0%)	34.2 (16.5%)	45.1 (34.0%)	35.4 (20.6%)
p-value			.45		.41		.23	

Note. All differences, statistically significant at a $p < .05$ level, are shown in bold.
*Percentages denote the change in rate of each outcome from the 1995–1996 baseline.
†As per the coding methodology of (Phibbs et al. 2007).
GA, gestational age.

Table 3: Relative Change in Pregnancy Outcomes in Philadelphia County Compared with Combined Control Counties in Postclosure Years

	1997–1999	*2000–2002*	*2003–2005*
Newborn mortality			
Neonatal deaths	**1.49 (1.12, 2.00)**	1.14 (0.85, 1.53)	1.05 (0.81, 1.36)
All fetal deaths	1.61 (0.95, 2.74)	1.52 (0.98, 2.37)	0.85 (0.31, 2.37)
All fetal deaths + neonatal deaths	**1.53 (1.14, 2.04)**	1.22 (0.94, 1.58)	0.93 (0.59, 1.46)
Possibly preventable fetal deaths*	1.33 (0.58, 3.02)	1.39 (0.71, 2.71)	0.61 (0.15, 2.43)
Possibly preventable fetal + neonatal deaths*	**1.45 (1.09, 1.94)**	1.15 (0.87, 1.52)	0.94 (0.66, 1.35)
Cesarean sections			
All deliveries	0.90 (0.79, 1.03)	0.96 (0.83, 1.12)	**0.85 (0.73, 0.98)**
All deliveries with GA 37–42 weeks	0.89 (0.79, 1.01)	0.94 (0.81, 1.10)	**0.82 (0.70, 0.96)**
Premature delivery			
23–32 weeks preterm	1.10 (0.95, 1.28)	1.08 (0.85, 1.37)	0.91 (0.69, 1.19)
23–37 weeks preterm	0.95 (0.89, 1.02)	1.01 (0.87, 1.17)	0.84 (0.69, 1.03)
34–37 weeks preterm	0.94 (0.88, 1.01)	1.00 (0.89, 1.12)	0.88 (0.76, 1.02)
Complications			
Any infant complication	1.18 (0.84, 1.64)	0.97 (0.72, 1.31)	1.04 (0.73, 1.48)
Any maternal complication	0.95 (0.78, 1.14)	1.04 (0.88, 1.22)	1.09 (0.88, 1.35)

Note. All values reported as odds ratios with 95% confidence intervals. All statistically significant differences at a $p < .05$ level are shown in bold. All results control for secular time trends, stable differences in county outcomes, maternal comorbid conditions, presence of a congenital anomaly, gestational age, and birth weight.
*As per the coding methodology (Phibbs et al. 2007).
GA, gestational age.

neonatal + fetal mortality (OR 1.53, 95 percent CI 1.14–2.14) increased for Philadelphia residents compared with the combined control counties in the first postclosure epoch. There were no statistically significant changes in the neonatal or maternal complication rates, Cesarean section rates, or preterm delivery rates in Philadelphia between 1997 and 1999 compared with combined control counties. After 2000, there was no statistically significant change in any perinatal outcome in Philadelphia compared with the combined control counties after adjusting for confounding factors except for a significantly lower Cesarean section rate in 2003–2005 in Philadelphia.

Separate Control Groups. The change in neonatal and neonatal + fetal mortality rates in Philadelphia county over the first postclosure epoch remained statistically higher when compared with both the five-county surrounding control group and the eight-county urban control group separately (Table 4).

Table 4: Relative Change in Pregnancy Outcomes in Philadelphia County Compared with Nearby Control Counties and Urban Control Counties Separately

Outcome and Epoch	*Philadelphia vs .Nearby*	*Philadelphia vs. Urban*
Neonatal death, 1997–1999	**1.59 (1.07, 2.37)**	**1.40 (1.10, 1.79)**
All fetal deaths + neonatal deaths, 1997–1999	**1.60 (1.11, 2.32)**	**1.45 (1.11, 1.89)**
Possibly preventable fetal + neonatal deaths*, 1997–1999	**1.52 (1.05, 2.20)**	**1.39 (1.07, 1.80)**
Cesarean sections, all deliveries, 2003–2005	**0.79 (0.67, 0.92)**	0.91 (0.77, 1.07)
Cesarean sections, all deliveries with GA 37–42 weeks, 2003–2005	**0.76 (0.64, 0.89)**	0.89 (0.75, 1.06)

Note. All values reported as odds ratios with 95% confidence intervals. All statistically significant differences at a $p < .05$ level are shown in bold. All results control for secular time trends, stable differences in county outcomes, maternal comorbid conditions, presence of a congenital anomaly, gestational age, and birth weight.
*As per the coding methodology of (Phibbs et al. 2007).

Philadelphia experienced a 59 percent larger increase in the odds of neonatal death compared with the surrounding control group (95 percent CI 1.07–2.37) and a 40 percent larger increase in the odds of neonatal death compared with the urban control group (95 percent CI 1.10–1.79). This increase in odds for Philadelphia was similar for neonatal + fetal deaths and neonatal + preventable fetal deaths. For Cesarean sections between 2003 and 2005, though, Philadelphia county only had a statistically significant change in odds compared with the surrounding control group (OR 0.79, 95 percent CI 0.67–0.91) but not the urban control group (OR 0.91, 95 percent CI 0.77–1.07). Rerunning the regression models after omitting Los Angeles County showed similar results to those of the full model.

1997–1999 Epoch. After stratifying the population by birth weight, we found that most groups had increased mortality in Philadelphia compared with the combined control groups (Table 5). Neonatal mortality and neonatal + all fetal mortality reached statistical significance for infants with a birth weight under 1,000 g (neonatal mortality OR 1.91, 95 percent CI 1.25–2.93; neonatal + fetal mortality OR 2.09, 95 percent CI 1.40–3.12). This finding persisted when we compared Philadelphia county to the surrounding county and urban county control groups separately.

We next examined the changes in risk-adjusted mortality rates at the remaining open hospitals in Philadelphia county in 1997–1999 and

Table 5: Relative Change in Mortality in Philadelphia Compared with Combined Control Counties, 1997–1999, Stratified by Birth Weight

Birth Weight Category	*Neonatal Deaths Odds Ratio*	*95% CI*
<1,000 g	1.91	(1.25, 2.93)
1,001–1,500 g	0.97	(0.47, 2.01)
1,501–2,000 g	1.66	(0.76, 3.66)
2,001–2,500 g	1.36	(0.65, 2.86)
2,501–3,000 g	1.32	(0.68, 2.58)
3,001–3,500 g	1.15	(0.64, 2.08)
3,501–4,000 g	1.37	(0.58, 3.28)
>4,000 g	3.03	(0.58, 15.8)
Birth Weight Category	*Neonatal + All Fetal Deaths Odds Ratio*	*95% CI*
<1,000 g	2.09	(1.40, 3.13)
1,001–1,500 g	0.90	(0.49, 1.62)
1,501–2,000 g	1.58	(0.85, 2.94)
2,001–2,500 g	1.58	(0.87, 2.85)
2,501–3,000 g	1.42	(0.82, 2.47)
3,001–3,500 g	1.35	(0.80, 2.28)
3,501–4,000 g	1.34	(0.60, 2.95)
>4,000 g	2.53	(0.63, 10.15)

Note. All values reported as odds ratios with 95% confidence intervals. All results control for secular time trends, stable differences in county outcomes, maternal comorbid conditions, presence of a congenital anomaly, gestational age, and birth weight.

the changes in delivery volume at these hospitals. More than 50 percent of the deliveries in Philadelphia county occurred at hospitals with a higher risk-adjusted mortality rate between 1997 and 1999 compared with 1995 and 1996, whereas only 20 percent of the deliveries occurred at hospitals with lower risk-adjusted mortality rates. There was only a minor correlation between change in volume and change in risk-adjustment mortality rate (Pearson $r = .07$). By 2003–2005, risk-adjusted mortality rates returned to the 1995–1996 baseline in the majority of hospitals that remained open.

DISCUSSION

We observed an adjusted 49 percent increase in the odds of neonatal mortality during the first 3-year epoch after obstetric units began to close in Philadelphia county, compared with both surrounding counties that experienced similar state and insurance regulations to Philadelphia and other large urban counties with similar obstetric networks. Unlike our initial hypotheses, after the year 2000, the relative difference in mortality rates returned to the preclosure levels, but it did not improve beyond these baseline levels. The resolution of these observed changes in risk-adjusted mortality rates may be related to increased cooperation between obstetric departments in the city, increased monitoring of the obstetric care delivery system by the public health department, or adaptive changes within the remaining open hospitals to cope with the increased volume of deliveries. Relative differences in complication rates during the postclosure epochs remained statistically similar to preclosure rates after controlling for differences in case mix and delivery hospital. These results suggest that even small-scale closures of obstetric units may be associated with adverse effects on perinatal outcomes, but these outcomes could recover.

These results should be placed in context of three other studies of the health impact of a reduction in hospital health care services. Hemmelgarn, Ghali, and Quan (2001) studied the impact of a 1996 hospital closure in Calgary on the outcomes of coronary revascularization procedures. This study found shorter lengths of stay and no change in mortality in the county in the 2 years succeeding the closure. However, this study did not include a control group or control for secular trends in their outcomes. Rosenbach and Dayhoff (1995) used a multiple time-series design to examine the impact of rural hospital closures on health care utilization and mortality. These rural hospital closures were associated with higher admissions to urban hospitals with no change in all-cause mortality rates. While this study included a control group of patients who lived in rural counties without a hospital closure, the study minimally controlled for case mix differences in the different counties.

Buchmueller, Jacobson, and Wold (2006) examined the closure of 15 small hospitals in Los Angeles between 1997 and 2001. For each one-mile increase in distance between a zip code and the nearest hospital, patients admitted with acute myocardial infarction had a 6.5 percent increase in mortality and patients admitted with an unintentional injury had an 11–20 percent increase in their mortality rates. The observed effect of the hospital closures was partially reduced by transport to higher quality hospitals. As this prior study is the only examination of multiple closures across a single urban area, its results are more representative to the situation observed in obstetric care closures.

Examining the relative change in mortality rates suggests potential consequences of obstetric closures. Our results suggest that the effect resulted from changes to risk-adjusted mortality rate at several hospitals after closures initially began, affecting over 50 percent of the deliveries in Philadelphia county. First, changes in delivery volume at the remaining hospitals could reduce the ability of health care providers to provide patient-specific management that would prevent a specific complication from occurring, resulting in higher neonatal mortality. However, we only found a minimal correlation between changes in delivery volume and changes in risk-adjusted mortality rates at these hospitals. This measure, though, does not account for potential surges in the number of deliveries, which have been identified as issues in previous studies of emergency room closures in Los Angeles (Sun et al. 2006), French obstetric units (Pilkington et al. 2008), pediatric hospitalizations (Lorch et al. 2008), and urban hospital closures (Lindrooth, Lo Sasso, and Bazzoli 2003).

Alternatively, the closure of obstetric units could have impacted patients' access to health care, either during the pregnancy or during the immediate delivery period. Many obstetric units also provide antepartum services to women, especially to those residing close to the hospital (Phibbs et al. 1993). This feature is particularly important in Philadelphia, as almost all obstetric hospitals provide some sort of prenatal care. Thus, in 2008, a survey of Philadelphia women found reduced access to early and adequate prenatal care, particularly in areas most affected by the closure of obstetric units (Knauer 2008). Also, the remaining open hospitals reported more patients delivering at their hospital without previously receiving care at an affiliated health care site (Bishop 2006). While historic studies fail to show a significant effect of prenatal care on maternal or neonatal outcomes (Fiscella 1995; Villar and Berqsio 1997), the importance of early and adequate prenatal care is a stated national health goal (U.S. Department of Health and Human Services 2010). These changes in patient access may change the types of services provided by obstetric units, especially concerning social services and access to health care. The relative difference in mortality at both the county level and at many affected hospitals returned to the preclosure baseline by 2000, suggesting that the health care system and individual hospitals compensated for these changes.

There were multiple outcome measures that differed between Philadelphia and the two control groups in univariable analysis but did not reach statistical significance in multivariable analysis. There were several likely reasons for these differences. First, as shown in Table 1, there were differences in the sociodemographic makeup between Philadelphia and each of the two control groups. The most striking was the larger percentage of white, insured patients in the surrounding control group compared with Philadelphia. This difference may have resulted in both differences in access to health care as these changes were occurring, or better navigation of the health care system in the surrounding counties compared with Philadelphia. Also, there were differences in the rate of several comorbid conditions between Philadelphia and each of the two control populations. These differences, which are controlled for in the multivariable analysis, may have explained the differences between univariable and multivariable analyses.

One major advantage to our study is the use of multiple comparison groups. The choice of control populations is important for this study design. Without a control group, our results would be confounded by secular trends in many outcomes. For example, Cesarean section rates increased throughout the study population by 20 percent by 2002 and 40 percent by 2005 compared with the 1995–1996 baseline. However, no single control group ideally reflects the racial/ethnic and sociodemographic breakdown in Philadelphia nor the practice environment experienced by patients in Philadelphia county. Nearby counties experienced similar practice and regulatory environments, but there are likely unmeasurable differences between the largely suburban population of these five counties and the urban population of Philadelphia beyond those measured by racial/ethnic and insurance background. Other urban areas have a more similar case mix to Philadelphia county, but they have a different practice environment. In these cases, studies suggest that the use of multiple control groups may strengthen the evidence from an observational study (Campbell 1969; Meyer 1995; Rosenbaum 1987). As these two control groups differ from each other, the fact that we found a statistically significant increase in mortality in the first postclosure epoch, whether we compared Philadelphia county to either both groups separately or combined, suggests that the choice of control group does not explain the results we found. Changes in Cesarean section rates in the third postclosure epoch, though, were only detected between Philadelphia and the nearby control counties—not the urban control counties. Thus, we cannot be certain whether the difference was related to the progressive closure of obstetric units in Philadelphia or unmeasured secular differences in Cesarean section rates.

There are several potential limitations to this study. First, we did not examine medical error rates or other patient safety indicators. Examining these outcomes through a careful chart review could help further quantify the impact of the obstetric unit closures. Second, the study was not designed to determine the reason for the high level of obstetric unit closures. Additional studies are needed to answer this question. Third, there is also potential unmeasured confounding because we did not have specific clinical data to further improve our risk-adjustment models, such as prenatal labs and ultrasound results. Other work, though, has used similar lists of comorbid conditions to control for differences in case mix (Phibbs et al. 2007). As with other population-based studies, we relied on ICD-9CM codes to determine some of our outcomes. However, by comparing the results of Philadelphia versus either other urban counties or nearby counties to each other during the preclosure and three postclosure epochs, we reduced these potential biases from both of these issues as well as the effect of secular trends. Finally, there could be a change in payments for obstetric care, the supply of obstetricians in Philadelphia, or another omitted variable (Meyer 1995) that coincided with the closures of obstetric units but did not occur in the other urban areas. Primary data, though, suggest that increased fixed costs over the postclosure period, primarily increased malpractice premiums (George 2001, 2002, 2005, 2008a, 2009; Treaster 2002; Anonymous 2007; Burling 2007; McCullough 2003), and reduced reimbursement for obstetric care (McCullough 2003; Burling 2007; George 2008a,b, 2009) were the primary reason for closures. While economic changes in the area could have resulted in both increased neonatal mortality rates and increased closure of obstetric units, no indices changed solely in Philadelphia compared to other counties between 1997 and 1999 to support this potential explanation.

In conclusion, obstetric unit closures in Philadelphia were associated with potentially adverse changes in the maternal and neonatal outcomes. These changes are similar to the rise in mortality seen in the closure of smaller, adult hospitals in Los Angeles (Buchmueller, Jacobson, and Wold 2006). State and county public health agencies should carefully monitor and prepare for changes to the health of communities experiencing large-scale reductions in obstetric services.

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Supporting Information

Additional supporting information may be found in the online version of this article:

Appendix SA1: Author Matrix.

Appendix SA2: ICD-9CM Codes for Maternal Comorbid Conditions and Infant Congenital Anomalies.

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