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Original article

Impact of Medicaid Expansion on Interpregnancy Interval

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A B S T R A C T

Objectives: Medicaid expansion under the Affordable Care Act (ACA) improved access to reproductive health care for low-income women and birthing people who were previously ineligible for Medicaid. We aimed to evaluate if the expansion affected the risk of having a short interpregnancy interval (IPI), a preventable risk factor for adverse pregnancy outcomes.

Methods: We evaluated parous singleton births to mothers aged 19 or older from U.S. birth certificate data 2009–2018. We estimated the effect of residing in a state that expanded Medicaid access (expansion status determined at 60 days after the prior live birth) on the risk of having a short IPI (<12 months) using difference-in-differences (DID) methods in linear probability models. We stratified the analyses by maternal characteristics and county-level reproductive health care access.

Results: Overall risk of short IPI was 14.9% in expansion states and 16.3% in non-expansion states. The expansion was not associated with a significant change in risk of having a short IPI (adjusted mean percentage point change 1.24 [–1.64, 4.12]). Stratified results also did not provide support for an association.

Conclusions: ACA Medicaid expansion did not have an impact on risk of short IPI. Preventing short IPI may require more comprehensive policy interventions in addition to health care access.

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Medicaid expansion under the Affordable Care Act (ACA) expanded health care access among people with low incomes (Miller & Wherry, 2017). Prior to the ACA, most states' Medicaid programs covered low-income pregnant women and birthing people until 60 days after giving birth, covering almost half of all births; on average across the states, low-income was defined as <133% of the federal poverty level (FPL). For non-pregnant women, Medicaid eligibility was much stricter; the average

income limit was <41% FPL for women with children, and even less for women without children (Wherry, 2018). In January 2014, 26 states and the District of Columbia expanded Medicaid eligibility to *all adults* with income <138% of FPL, regardless of pregnancy or postpartum status (Wherry, 2018). Thus, women with low incomes in these states would no longer lose coverage after 60 days postpartum, when their eligibility based on pregnancy status came to an end, and access to health care services was improved for many low-income women of reproductive age (Clapp, James, Kaimal, Sommers, & Daw, 2019).

Since not all states implemented the ACA Medicaid expansion, the policy provided a natural experiment to evaluate the hypothesis that improved access to reproductive health care could effectively reduce the incidence of adverse pregnancy

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outcomes. Previous studies confirmed that ACA Medicaid expansion increased health care access among women of reproductive age (Daw & Sommers, 2019; Johnston, Strahan, Joski, Dunlop, & Adams, 2018; Margerison, MacCallum, Chen, Zamani-Hank, & Kaestner, 2020). Regarding pregnancy outcomes, the expansion did not affect overall levels of neonatal outcomes such as infant mortality, preterm birth, or low birth-weight in all births, possibly because access to health care for women during pregnancy did not change for the majority of the birthing population. However, policies can affect health equity in addition to overall population health, and the expansion was found to reduce Black-White disparities in preterm birth, low birth weight, infant mortality, and maternal mortality (Bhatt & Beck-Sagué, 2018; Brown et al., 2019; Eliason, 2020).

Interpregnancy interval (IPI) is another outcome that could be affected by ACA expansion. IPI is the time from the last birth to the beginning of the subsequent pregnancy. Short IPI, commonly defined as having an IPI shorter than 6 or 12 months, is associated with adverse maternal and neonatal outcomes such as postpartum hemorrhage or preterm birth (Conde-Agudelo, Rosas-Bermudez, Castaño, & Norton, 2012), and is more likely among women with lower levels of education or income, of Hispanic or Black race/ethnicity, or with high parity (two or more previous live births) (Cheslack Postava & Winter, 2015).

Expanded access to inter-pregnancy reproductive health care, such as contraception or counseling regarding family planning, is a potential mechanism to reduce unplanned pregnancies or risk of short IPI (Ahrens et al., 2018). The ACA was estimated to reduce the uninsured rate of reproductive age women (who account for 1 out of 4 US women) by 7.4 percentage points (Daw & Sommers, 2019). Insurance can help women access contraceptive care and other forms of interpregnancy health care, and thereby improve pregnancy outcomes and reduce health inequities. Yet we are unaware of prior longitudinal studies examining the effect of expansion on the occurrence of short IPI in the entire nation. In addition, social determinants of health such as family relationships, socioeconomic context (Holt et al., 2020; Maness & Bui, 2016), and availability of health care providers (Lyu & Wehby, 2019; Wehby, Lyu, & Shane, 2019) may moderate the effectiveness of expanded access to reproductive health care.

We used U.S. vital records and a quasi-experimental design to examine the impact of ACA Medicaid expansion on short IPI, by comparing the changes in risk of having short IPI across states before and after the expansion occurred. To test underlying assumptions, we additionally used an event study approach to capture the pre- and post-expansion trends of short IPI and further examined whether the potential policy effect varied by maternal characteristics and availability of reproductive health care services.

Methods

Study Population

We obtained birth certificate data from the National Center for Health Statistics Vital Statistics Natality Birth Data files in 2009–2018, as compiled from data provided by the 57 vital statistics jurisdictions through the Vital Statistics Cooperative Program, for 39,528,323 live births; we refer to these births as the “current” births. We restricted the population to parous women aged 19 or older (being adult at the time of conception) who had a singleton live birth with known gestational age and year and month of the prior birth and IPI ≥ 2 months (because Medicaid

eligibility based on pregnancy ends after 60 days postpartum). We only utilized data from states of birth where the 2003 version Birth Certificate was fully implemented by 2011 (see Figure 1 and eTable 1 for further details). Non-singleton births were excluded because the data structure did not allow us to count one observation per multiple gestation birth (i.e., multiple gestation births from the same pregnancy were not linked to each other and would thus have been counted more than once). To allow consistent post-expansion observation time, we excluded births in 7 states with ACA Medicaid Expansion later than 1 Jan 2014 (eTable 1), leaving 15,155,756 births eligible for the main analysis. After excluding 281,761 births with missing covariates, the study population consisted of 14,873,995 births.

Outcome

We had access to date of birth of the current birth (reported as month and year; we imputed the day as the first day of the birth month), gestational age (in completed weeks), and months since last live birth. We estimated IPI (in days) by subtracting gestational age (in days) from the interval between the derived dates of last birth and the current birth. We defined short IPI as IPI shorter than 12 months (McKinney, House, Chen, Muglia, & DeFranco, 2017).

Exposure

Exposure to ACA Medicaid Expansion was assigned based on the expansion status of the maternal residence state at the 60th day after the estimated date of the prior birth (see ACA Medicaid expansion dates in eTable 1). Date of the prior birth was estimated by subtracting months since last live birth (transformed into days) from date of the current birth. If the state's Medicaid expansion happened before or at the 60th day after the prior birth, the exposure status was coded as 1; it was coded as 0 for all the other births, including all those for which the expansion occurred after the 60th day after the prior birth and all births in the non-expansion states. By this definition, we treated “partially exposed” women as non-exposed. For example, if a state expanded Medicaid 65 days after the prior birth, a woman residing in that state would have been covered for all but 5 days and was still defined as non-exposed.

Covariates

We examined sociodemographic factors from the birth certificate data, including maternal age (years), parity, race/ethnicity (non-Hispanic (NH) White, Hispanic, NH Black, NH Asian or Pacific Islander, NH American Indian & Alaska Native, and Other), education (less than high school, high school graduate, some college, undergraduate degree, graduate degree), and foreign-born status (U.S. born or foreign born). County-level confounders included annual unemployment and poverty rates. Because the effectiveness of expanded access to care may depend on availability of health care services, we also examined potential effect modifiers including primary health professional shortage area (HPSA), reproductive health care workforce supply, and contraceptive service supply. These annual county-level data were merged with the birth certificate data by county of residence, and values were assigned based on year of birth because of better data availability and quality than in the year of prior birth, which can be missing for births with longer IPIs. See eTable 2 for the county-level variable definitions and their

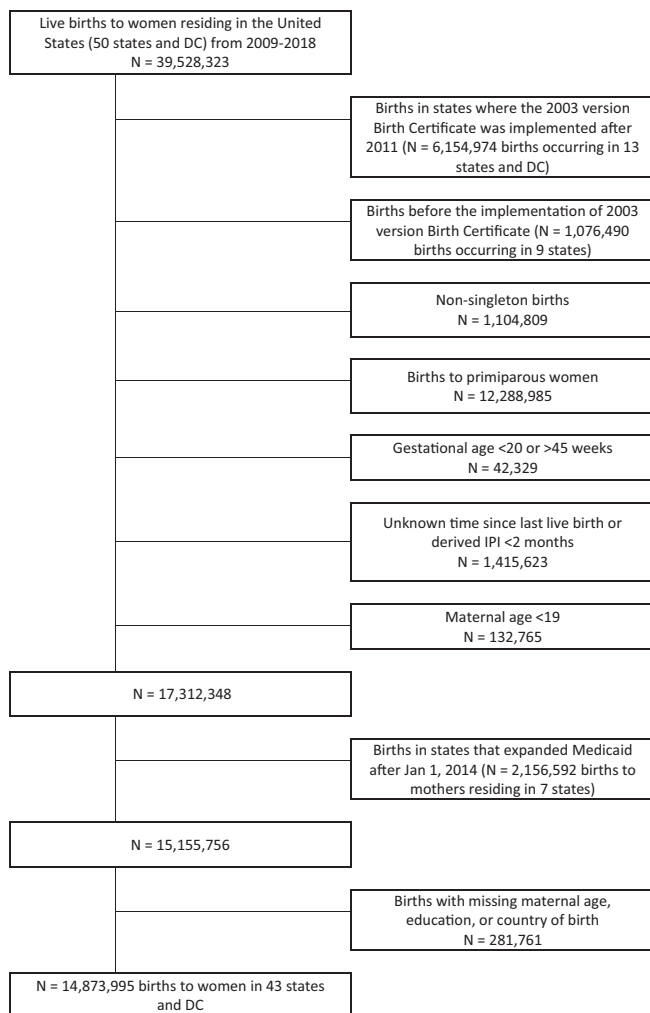


Figure 1. Flowchart of study population.

corresponding sources. Except for HPSA, we categorized the county-level effect modifiers by quartiles within each state in the whole study period, to better distinguish the high- and low-supply counties within the state.

Statistical Analysis

We used difference-in-differences (DID) analysis to estimate the effect of ACA Medicaid expansion on the risk of having a short IPI, by fitting linear probability models with state and year-month (at the 60th day after the last birth) fixed effects, adjusting for confounders: maternal number of previous live births, age in quadratic form, education, race/ethnicity, foreign-born status, and county unemployment rate and poverty rate. Standard errors were estimated using state-cluster robust estimates. To understand the policy effect on expansion states over time and to test the parallel trend assumption, we also used event study models to visualize the changes in relative risk of having a short IPI over time in the expansion states and tested for pre-policy linear trends. See Appendix A for model specifications.

Subgroup analyses

We stratified the analysis by payer, maternal age, education, race/ethnicity, and foreign-born status (foreign-born women were more likely to be ineligible for Medicaid because of residency status), county-level HPSA status, reproductive health care workforce supply, and contraceptive service supply. We additionally performed event study models for the strata where there was a significant association.

Sensitivity analyses

In an attempt to exclude partially exposed women, we conducted a sensitivity analysis in which we estimated the effect of the expansion after excluding all women who had the prior birth before 2014 and the current birth in or after 2014 (N = 5,025,638).

Some expansion states had decreased eligibility from January 2013 to January 2014 (Illinois, Minnesota, New Jersey, New York, Rhode Island, and Vermont) and some non-expansion states had increased eligibility (Kansas, South Dakota, Utah, Virginia, and Wyoming; see eTable 3 for information). The non-expansion states Wisconsin and Maine dropped eligibility limits for parents from 200% to 100% and from 200% to 105% FPL from January 2013 to January 2014. We conducted a sensitivity analysis excluding these states with contradicting eligibility change or large reduction of eligibility.

To evaluate differential selection into observation, we performed additional analyses of the policy effect on county female population sizes and the nulliparous and parous birth rates (Appendix B).

To test the assumption of having stable composition in the expansion and non-expansion states, we examined associations of the expansion with probability of having Medicaid payment at birth hospitalization (vs. private insurance), high school graduation or less education (vs. college or more education), maternal age <26 years (vs. ≥ 26 years), race/ethnicity (Hispanic, NH Black, or NH Asian or Pacific Islander vs. NH White, respectively), and foreign-born status (vs. U.S. born), adjusting for county unemployment rate and poverty rate. Even if the expansion resulted in compositional changes based on any of these measured factors, that would not bias our estimates given that we control for these factors in the main models.

Results

Table 1 shows that overall, births in the expansion states had a lower prevalence of short IPI (14.9% vs. 16.3%) and tended to have more mothers with older age, higher education, or foreign-born nativity, and fewer mothers with NH Black race/ethnicity. Mothers in the expansion states were more likely to reside in counties with lower poverty rates and higher mean numbers of reproductive health care specialists and contraceptive clinics per 1000 women aged 15–45.

Table 2 shows the mean risk differences of having a short IPI in the expansion states compared to the non-expansion states, overall and stratified by the potential effect modifiers at the individual- or county-level. After adjusting for confounders, there was no significant association between short IPI and Medicaid expansion (percentage point 1.24 [−1.64, 4.12]). The result was consistent after excluding partially exposed women (percentage point 1.75 [−1.86, 5.37]). To understand which covariate drove the shifted estimation from the unadjusted to the adjusted model, we additionally performed piecewise adjusted models. The shift of the point estimation and enlargement of confidence

Table 1
Distribution of Variables among all Eligible Births in 43 States and DC by State's Expansion Status (N = 14,873,995)

	All		Non-expansion State		Expansion State	
	No.	Col%	No.	Col%	No.	Col%
Total	14,873,995	100	7,039,813	100	7,834,182	100
Interpregnancy interval (months) in categories						
<6	688,747	4.6	356,103	5.1	332,644	4.3
6–11	1,623,225	10.9	787,821	11.2	835,404	10.7
12–17	2,002,274	13.5	936,303	13.3	1,065,971	13.6
18–23	1,745,850	11.7	813,605	11.6	932,245	11.9
24–29	1,432,170	9.6	669,160	9.5	763,010	9.7
30–35	1,137,528	7.7	534,598	7.6	602,930	7.7
36–41	945,788	6.4	447,543	6.4	498,245	6.4
42–47	779,095	5.2	370,168	5.3	408,927	5.2
48–53	659,537	4.4	312,933	4.5	346,604	4.4
54–59	551,774	3.7	261,293	3.7	290,481	3.7
≥60	3,308,007	22.2	1,550,286	22.0	1,757,721	22.4
Maternal age (years)						
19–24	2,853,243	19.2	1,523,249	21.6	1,329,994	17.0
25–29	4,464,383	30.0	2,212,051	31.4	2,252,332	28.8
30–34	4,558,901	30.7	2,070,480	29.4	2,488,421	31.8
35–39	2,434,748	16.4	1,011,405	14.4	1,423,343	18.2
40–44	533,371	3.6	211,916	3.0	321,455	4.1
45–50	29,349	0.2	10,712	0.2	18,637	0.2
Number of previous live births						
1	7,762,637	52.2	3,575,616	50.8	4,187,021	53.5
2	4,125,700	27.7	1,991,981	28.3	2,133,719	27.2
3	1,758,797	11.8	870,291	12.4	888,506	11.3
4 or more	1,226,861	8.3	601,925	8.6	624,936	8.0
Maternal education						
Less than high school	2,691,186	18.1	1,256,286	17.9	1,434,900	18.3
High school graduate	3,886,576	26.1	1,936,243	27.5	1,950,333	24.9
Some college	4,347,299	29.2	2,115,218	30.1	2,232,081	28.5
Undergraduate degree	2,570,736	17.3	1,203,970	17.1	1,366,766	17.5
Graduate degree	1,378,198	9.3	528,096	7.5	850,102	10.9
Maternal race/ethnicity						
Hispanic	4,426,493	29.8	1,941,949	27.6	2,484,544	31.7
NH White	7,307,336	49.1	3,548,411	50.4	3,758,925	48.0
NH Black	1,981,427	13.3	1,184,075	16.8	797,352	10.2
NH Asian or Pacific Islander	889,445	6.0	235,362	3.3	654,083	8.4
NH American Indian & Alaska Native	124,919	0.8	68,086	1.0	56,833	0.7
Other	144,375	1.0	61,930	0.9	82,445	1.1
Maternal nativity						
U.S.	10,876,533	73.1	5,426,722	77.1	5,449,811	69.6
Foreign	3,997,462	26.9	1,613,091	22.9	2,384,371	30.4
Sex of infant						
Female	7,269,079	48.9	3,443,259	48.9	3,825,820	48.8
Male	7,604,916	51.1	3,596,554	51.1	4,008,362	51.2
Payment source						
Medicaid	6,875,065	46.2	3,259,169	46.3	3,615,896	46.2
Private	6,405,087	43.1	2,835,320	40.3	3,569,767	45.6
Self Pay	717,345	4.8	480,641	6.8	236,704	3.0
Other	728,233	4.9	394,079	5.6	334,154	4.3
Unknown	148,265	1.0	70,604	1.0	77,661	1.0
County unemployment rate quartile*						
Mean, Std.dev			7.1	3.0	4.7	1.6
County poverty rate quartile*						
Mean, Std.dev			15.7	5.3	13.3	4.7
County HPSA designation [†]						
Not designated as HPSA	1,034,357	7.0	613,529	8.7	420,828	5.4
Part of the county as HPSA	1,735,848	11.7	699,162	9.9	1,036,686	13.2
Whole county as HPSA	12,101,774	81.4	5,725,106	81.4	6,376,668	81.4
County reproductive health care specialists per 1000 women aged 15–45						
Mean, Std.dev			0.5	0.3	0.5	0.3
County public-funded clinics providing contraceptive service per 1000 women aged 15–45						
Mean, Std.dev			0.1	0.2	0.2	0.2

* County level unemployment rate and poverty rates are categorized by quartiles in the whole country in 2009–2018

[†] HPSA = primary health professional shortage area

intervals occurred after adjustment for county-level covariates (eTable 4). Figure 2 shows the mean risk differences in short IPI in the expansion versus the non-expansion states, over calendar

quarters before and after the expansion. Test for differential linear trend supported parallel trend in the expansion states compared to the non-expansion states before the expansion

Table 2

Difference-In-Differences Analysis of ACA Medicaid Expansion and Short IPI, Overall and Stratified by Maternal and County-Level Characteristics (N = 14,873,995)

	Difference-In-Differences	
	Unadjusted [95% CI]	Adjusted [95% CI] [†]
	Percentage points [*]	Percentage points [*]
All	–1.43 [–2.49, –0.38]	1.24 [–1.64, 4.12]
Payment at delivery		
Medicaid	–1.27 [–2.54, 0.01]	0.95 [–2.34, 4.25]
Private insurance	–0.96 [–0.21, 0.19]	1.68 [–0.92, 4.27]
Maternal age		
19–25	–0.24 [–1.86, 1.37]	3.27 [–2.17, 8.70]
≥26	–1.24 [§] [–2.17, –0.30]	1.06 [–1.47, 3.59]
Maternal education		
High school graduation or less	–1.02 [–2.34, 0.29]	1.41 [–1.90, 4.73]
Some college education or more	–1.29 [§] [–2.34, –0.24]	1.25 [–1.42, 3.92]
Maternal race		
Hispanic	–0.21 [–1.08, 0.67]	3.37 [§] [1.43, 5.31]
NH White	–1.45 [§] [–2.81, –0.09]	0.65 [–2.29, 3.58]
NH Black	–0.85 [–2.69, 0.99]	–0.16 [–3.96, 3.63]
NH Asian or Pacific Islander	–1.71 [–3.61, 0.19]	2.63 [–0.38, 5.64]
NH AI/AN	–1.83 [–3.79, 0.13]	0.25 [–4.74, 5.23]
NH Other	–2.90 [–4.65, –1.14]	–1.01 [–2.39, 0.36]
Maternal nativity		
U.S.	–1.59 [§] [–2.81, –0.37]	1.08 [–2.21, 4.37]
Foreign	–1.01 [§] [–1.99, –0.03]	1.84 [–0.24, 3.92]
County-level HPSA designation [‡]		
Not designated as HPSA	–4.47 [§] [–7.83, –1.10]	–2.10 [–8.20, 4.00]
Part of the county as HPSA	–2.50 [–7.35, 2.35]	–1.63 [–6.81, 3.56]
Whole county as HPSA	–0.16 [–1.72, 1.40]	1.55 [–1.15, 4.25]
County-level reproductive health care specialists per 1000 women aged 15–45 [‡]		
Lowest quartile	–0.88 [–3.22, 1.50]	1.62 [–1.71, 4.96]
Second	–1.43 [–3.49, 0.62]	0.65 [–2.95, 4.25]
Third	–0.64 [–3.44, 2.16]	1.13 [–2.88, 5.15]
Highest quartile	–0.39 [–3.96, 3.17]	1.26 [–2.60, 5.12]
County-level publicly funded contraceptive clinics per 1000 women aged 15–45 [‡]		
Lowest quartile	–1.72 [–4.15, 0.71]	2.33 [–1.22, 5.89]
Second	2.56 [–2.16, 7.27]	5.12 [–2.37, 12.62]
Third	–3.02 [–6.54, 0.50]	–0.89 [–5.15, 3.37]
Highest quartile	–1.29 [–2.89, 0.32]	0.15 [–1.73, 2.03]

* The stated percentage points reflect the difference in the percentage of births with short IPI among women exposed to Medicaid expansion compared to those not exposed to expansion (starting within 60 days after the birth of their prior infant).

[†] Model adjusted for maternal number of previous live births, age in quadratic form, education, race/ethnicity, foreign-born status, and county unemployment rate and poverty rate. For stratified analysis, the model did not adjust for the moderator.

[‡] HPSA = primary health professional shortage area. County level unemployment rate and poverty rates are categorized by quartiles in the whole country in 2009–2018. County level reproductive health care specialists and public funded clinics are categorized by quartiles in the whole state in 2009–2018 for stratified analyses.

[§] $p < .05$.

^{||} $p < .01$.

($p = .161$). There was no obvious disruption of the trend around the time of the expansion (the first quarter of 2014) in either model.

Subgroup Analyses

Confidence intervals for most of the covariate-adjusted stratum-specific differences included the null. For example, women under 26, who potentially also had expanded access to private insurance due to the ACA, had a 3.27 [–2.17, 8.70] percentage points change, in contrast to older women who had a 1.06 [–1.47, 3.59] percentage points change. The point estimates were small, and confidence intervals for both estimates were wide and included the null. Hispanic women showed an increased risk of short IPI in the expansion states vs in the non-expansion states (3.37 [1.43, 5.31] percentage points change). This appeared to be part of a long-term trend of relative increase (eFigure 1), as there was a nonparallel trend between the expansion and non-expansion states before the expansion ($p = .001$).

Sensitivity Analyses

The association between Medicaid expansion and short IPI was similar after excluding the states having contradicting eligibility change or large reduction of eligibility (percentage point 0.97 [–2.29, 4.22]).

The expansion was not associated with changes in birth rates (See Appendix B.). Over time, there was a lower proportion of births with socioeconomic or demographic characteristics that are associated with short IPI in the expansion states relative to the non-expansion states, including having Medicaid payment at birth hospitalization (adjusted percentage point difference –3.12 [–5.32, –0.92]), high school graduation or less education (adjusted percentage point difference –2.76 [–5.59, 0.07]), maternal age <26 years (adjusted percentage point difference –0.87 [–2.82, 1.08]), Hispanic ethnicity (adjusted percentage point difference –3.97 [–6.94, –1.00]), NH Black race (adjusted percentage point difference –2.19 [–5.06, 0.68]), NH Asian or Pacific Islander race (adjusted percentage point difference –1.56

ACA Medicaid Expansion on risk of short IPI by time (quarter)

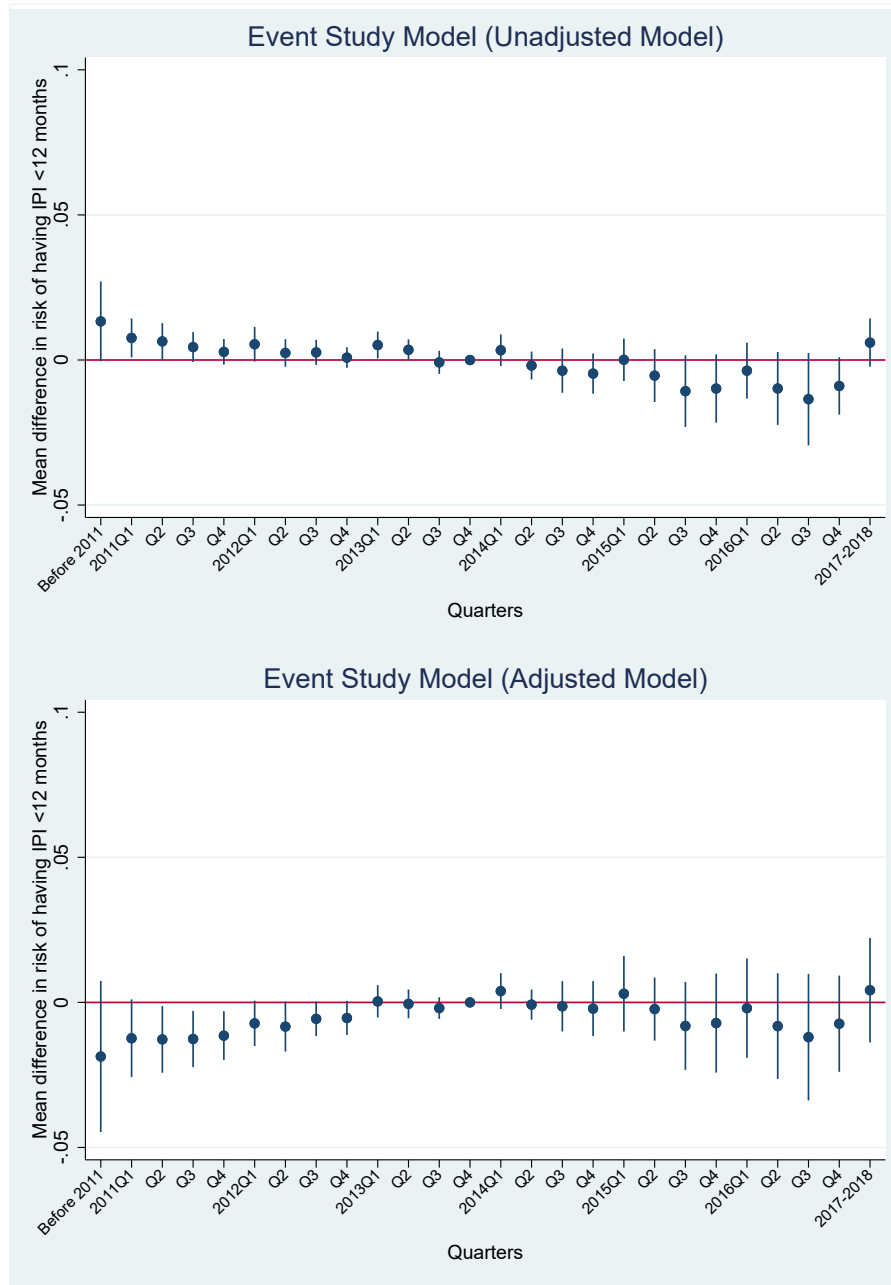


Figure 2. Mean risk difference of having short IPI in expansion states compared to non-expansion states by calendar quarters of the date of the prior birth plus 60 days. Above: unadjusted fixed effect model; bottom: adjusted model, adjusting for number of previous live births, maternal age in quadratic form, maternal education, race/ethnicity, and maternal foreign-born status, and county unemployment rate and poverty rate.

[−3.46, 0.34]), and foreign-born status (vs. US-born) (adjusted percentage point difference −2.95 [−7.08, 1.19]). Corresponding event time analyses mostly showed continuous trends for these composition changes, which existed before the expansion (eFigures 4–10).

Discussion

We found no statistically significant impact of ACA Medicaid expansion on the risk of short IPI using a difference-in-differences

model. The policy effects in analyses stratified by sociodemographic factors and healthcare availability were also found insignificant.

Using national birth certificate data from 2009 to 2018, we observed a crude reduction of short IPI in the expansion states compared to the non-expansion states, a finding similar to that in the recent work by Patel et al. using data from 2006 and 2012 (Patel, Bakk, Pensak, & DeFranco, 2021). Rather than using only two years of data, we used data of every year from 2009 to 2018 and fixed effect DID model to adjust for common trends shared

between the expansion and non-expansion states. Furthermore, we adjusted for state-level socioeconomic variables. In contrast to Patel et al., the results of our study did not suggest an impact of the ACA on parous women's risk of having a short IPI after adjusting for covariates. The absence of an ACA impact on short IPI may be attributed to the fact that the ACA primarily affects women with incomes <138% of FPL, especially women without children (Johnston et al., 2018; Margerison et al., 2020), who were not included in this study. Additionally, immigrants, especially the undocumented migrant women who accounted for about 7% of the births recorded in the U.S. (Passel & Cohn, 2016), were generally not covered by traditional or pregnancy-related Medicaid during pregnancy (Swartz, Hainmueller, Lawrence, & Rodriguez, 2017), but still contributed to the observed births in this study. On the other hand, the policy may have successfully prevented unintentional pregnancies by increasing the effective contraception use in the population (Darney et al., 2020), thereby reducing short IPIs. However, successfully prevented unintentional pregnancies could not be observed, thus this impact cannot be included in the policy effect estimation of this study.

The expanded eligibility to access reproductive health care, and family planning services in particular, may not translate to actual increased use. Factors associated with postpartum contraception use in general remain understudied, and more research is needed to verify if ACA expansion was associated with increases in postpartum contraception (Myerson, Crawford, & Wherry, 2020). Among all women who requested insurance preapproval for long-acting reversible contraception (LARC), women with Medicaid payment had lower probability of LARC insertion than women with private insurance, possibly due to their socioeconomic disadvantages and barriers in making another hospital visit for the insertion (Higgins, Dougherty, Badger, & Heil, 2018). On the suppliers' side, family planning facilities may have difficulties meeting the increasing needs of women newly covered by Medicaid, due to funding or other challenges (Boudreaux, Choi, Xie, & Marthey, 2019).

Health care access, sociodemographic, and socioeconomic factors may also have a complicated interplay in affecting individual reproductive outcomes. As our stratified analyses suggested, individual socioeconomic or sociodemographic factors may interact with the socioeconomic environment and affect the timing of next birth. The intention to have another child shortly after the last birth may depend on maternal age and socioeconomic circumstances (Cheslack Postava & Winter, 2015). For example, women who start childbearing at a more advanced age may intentionally shorten their IPI to achieve desired family size. We are not certain if some aspects of the ACA could have factored into a family's decision to have another child, such as reduced financial burden of child health care (Wisk, Peltz, & Galbraith, 2020). Stratified analyses showed generally null associations among race/ethnicity groups, except among Hispanic people where the effect estimation suggested that ACA expansion was associated with increased risk of short IPI. Yet the parallel trend assumption for the analyses was not fulfilled and we could not draw a conclusion on whether the policy affected the risk of short IPI among the Hispanic people.

Under the ACA, women with younger age (<26 years) who were not eligible for Medicaid may have obtained private health insurance coverage through their parents (Monaghan, 2014). Given this broader expansion of access to health care, we expected to see a greater policy impact in this younger age group than women aged 26 or older, but this was not supported by the data.

Strengths and Limitations

This study's use of national birth certificates is a strength, as most births in the US were included. The study also used a quasi-experimental design, which is a rigorous approach to examine if the risk of having a short IPI can be affected by policy intervention on reproductive health care access.

Longitudinal data are ideal to study the interval between births, because the process between births is affected by many time-varying factors that lead to event or censoring. Nonetheless, we only have national cross-sectional birth data and did not have prospective information on the entire population at risk, i.e., we only observed women who had a multiparous birth and not those who had a "previous birth" but did not give birth in the study period. Thus, the estimated policy effect did not account for any reduction of short IPI attributable to successfully preventing unwanted pregnancies. We did compare the multiparous birth rates, to gauge whether the probability of being observed may have changed in response to the expansion and thus impact our results; however, the multiparous birth rate did not appear to be associated with the expansion, as previous studies showed (Geiger, Sommers, Hawkins, & Cohen, 2021; Palmer, 2020).

Without longitudinal data linking the sequential births, we could not rule out differential selection into the study population in the expansion vs. non-expansion states (i.e., the policy effect that prevented unwanted pregnancy may result in different compositions of the birth population in the expansion vs non-expansion states). Over time, the expansion states' birthing populations were gradually less at risk of having short IPI. Nonetheless, the analyses did not suggest that the change resulted from the expansion, so the impact of this potential bias was limited.

We categorized women who were "partially exposed" as non-exposed, which could have diluted the difference between the groups, although sensitivity analyses excluding the partially exposed did not show different results. To examine a longitudinal health outcome such as interpregnancy interval, future studies preferably should use a cohort design to "assign" treatment strategy at the prior birth and follow women at risk of having another child.

As the policy impact lasts over time and is dependent on the women's state of residence, the data and statistical analyses also need to accommodate when exposure status changes over time, e.g., when women move. We assumed that mothers lived in the same state and county when they had their prior birth. This may not be the case for some women. Thus, their exposure to Medicaid expansion, as well as covariates measured at the county level, may have been misclassified for an unknown number of women. In addition, moving between expansion and non-expansion states could affect the insurance coverage and in turn the mothers' access to reproductive health care, i.e., insurance churn (Sommers, Gourevitch, Maylone, Blendon, & Epstein, 2016). We could not evaluate the impact of these types of exposure misclassifications on our results with the available data, but we do not expect them to be systematically different by women's exposure status.

Irrespective of expansion status, changes in Medicaid pregnancy coverage in 2014 could also make some women more likely to be covered, which in turn could affect immediate postpartum contraception (Clapp, James, Kaimal, Sommers, & Daw, 2019) and dilute the estimated policy effect. Furthermore, changes in some states' income limitation for eligibility of parents contradicted the state's expansion status, although these

changes tended to be modest in magnitude. Yet there were only a few states that were included in our main analyses for which the contradiction between expansion status and change in eligibility (from 2009 to 2018) seemed substantial (Illinois, Vermont, and Wisconsin; see eTable 3). After excluding the states with contradictory income eligibility changes, we found results consistent with those from the main analyses.

Implications for Practice and/or Policy

Our finding on the lack of policy impact on IPI may suggest that the time between births is a more complicated health outcome that requires more comprehensive policy interventions than insurance coverage. For example, in addition to the public health and health care policy, education and labor policies may also impact intention and timing of having another child. The limitations of the current study also suggest that future evaluations of the policy impact on postpartum health outcomes, including IPI, will benefit from having more comprehensive longitudinal data.

Conclusion

This study did not find evidence that ACA Medicaid expansion impacted short IPI. To fully understand the policy impact on IPI, we may need longitudinal data, and to take into consideration other social, structural, demographic, and health-related factors that may affect birth spacing.

Supplementary Data

Supplementary data related to this article can be found at <https://doi.org/10.1016/j.whi.2021.12.004>.

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