RESEARCH ARTICLE

Estimating the effect of Prenatal Care Coordination in Wisconsin: A sibling fixed effects analysis

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Abstract

Objective: To estimate Prenatal Care Coordination's (PNCC) effect on birth outcomes for Wisconsin Medicaid-covered deliveries.

Data Source: A longitudinal cohort of linked Wisconsin birth records (2008-2012), Medicaid claims, and state-administered social services.

Study Design: We defined PNCC treatment dichotomously (none vs. any) and by service level (none vs. assessment/care plan only vs. service uptake). Outcomes were birthweight (grams), low birthweight (<2500 g), gestational age (completed weeks), and preterm birth (<37 weeks). We estimated PNCC's effect on birth outcomes, adjusting for maternal characteristics, using inverse-probability of treatment weighted and sibling fixed effects regressions.

Data Collection/Extraction Methods: We identified 136 224 Medicaid-paid deliveries, of which 33 073 (24.3 percent) linked to any PNCC claim and 22 563 (16.6 percent) linked to claims for PNCC service uptake.

Principal Findings: Sibling fixed effects models—which best adjust for unobserved confounding and treatment selection—produced the largest estimates for all outcomes. For example, in these models, PNCC service uptake was associated with a 1.3 percentage point (14 percent) reduction and a 1.8 percentage point (17 percent) reduction in the probabilities of low birthweight and preterm birth, respectively (all $P < .05$).

Conclusions: PNCC's modest but significant improvement of birth outcomes should motivate stronger PNCC outreach and implementation of similar programs elsewhere.

KEYWORDS

care coordination, Medicaid, prenatal

1 | INTRODUCTION

Adverse birth outcomes remain a critical public health problem.¹⁻³ Low birthweight and preterm infants are at risk of neonatal and post-neonatal mortality and a host of morbidities,^{4,5} and scholars, policy makers, and practitioners argue that investment in prevention is essential.⁶⁻¹⁰

Following federal Medicaid expansions in the late 1980s, several states developed care coordination programs designed to supplement standard prenatal care and connect Medicaid-covered pregnant women to services supportive of maternal and child health.¹¹⁻¹³ One such initiative is Wisconsin's Prenatal Care Coordination (PNCC) program, a service that directs eligible pregnant Medicaid beneficiaries to medical, educational, and social services.¹⁴ On the whole,

evidence suggests that care coordination programs contribute to better birth outcomes.^{13,15-22} Prior studies have relied on observational data, however, and although recent investigations have used quasi-experimental methods to bolster internal validity,^{17-20,22,23} problems with estimating the effects of care coordination on birth outcomes remain.²⁴

The purpose of this study is to estimate the effects of PNCC utilization on four birth outcomes—birthweight, low birthweight, gestational age, and preterm birth—in Wisconsin during 2008-2012. We leverage a unique, statewide administrative dataset and use two analytic techniques to estimate these effects. We use inverse-probability of treatment weighting regression to adjust for observable differences in the characteristics of women who did and did not participate in PNCC, and sibling fixed effects regressions that adjust for stable unobserved differences among mothers who participated in PNCC during some pregnancies but not others. Considering the two sets of estimates together offers a more complete assessment of care coordination's effect on birth outcomes than has been possible in most prior work and, by extension, defines the likely population health effect of care coordination expansion.

2 | METHODS

2.1 | Prenatal Care Coordination

In Wisconsin, the purpose of PNCC is to improve maternal and infant health by connecting eligible Medicaid beneficiaries to medical, educational, and social services that supplement standard prenatal care.¹⁴ Potentially eligible women can seek PNCC assessment, or be referred by a health care provider or community agency, and receive services in clinical or community-based settings, such as a hospital or a local public health department, in their residential county. Institutions and agencies that provide PNCC services must submit a service outreach and care management plan to render PNCC assessments, care plans, and services. Participating women can receive a variety of services that may fit their needs, such as nutrition counseling, tobacco cessation, psychosocial therapy, and general health education. Qualified PNCC providers may be clinical (eg, nurse practitioners or clinical dietitians) or nonclinical (eg, social workers or occupational therapists), reflecting the broad range of potential services to address beneficiaries' needs.

PNCC's providers assess women for service eligibility with a brief questionnaire, and the eligibility criteria for service receipt are intentionally broad. To be eligible for service receipt, women must be receiving Medicaid and be younger than 18 years old upon assessment or report at least four risk factors on the questionnaire.¹⁴ Risk factors generally fall into one of four categories: socioeconomic or demographic background (eg, no high school education), pregnancy history (eg, prior preterm birth), health behavior (eg, tobacco smoking), or living conditions (eg, homelessness).²⁵ Eligible women who decide to use PNCC services meet with a care coordinator to develop an individually tailored care plan. While there are no strict

requirements regarding monitoring, care coordinators are encouraged to contact their assigned recipient at least once per 30 days. Wisconsin Medicaid covers PNCC services from program initiation up to 60 days postpartum.¹⁴ Although PNCC coverage is state-mandated, county health departments have autonomy in PNCC provision, so there is considerable county-level variation in the rates of PNCC assessment and service uptake.²⁴ Triaging by risk may contribute to this variation; some counties exhibit high PNCC assessment rates with relatively low service enrollment thereafter, whereas other counties appear to specifically target with high-risk pregnancies and enroll a large portion of those who were assessed.

2.2 | Data

Data come from Big Data for Little Kids (BD4LK), which links information from multiple Wisconsin administrative sources at the child level. Data for our analyses include the following: (a) vital records for all Wisconsin Medicaid-covered resident in-state births between 2008 and 2012; (b) Medicaid claims and encounters (henceforth "claims") associated with a paid obstetrical delivery between 2008 and 2012; and (c) social services utilization from the Multi-Sample Person File data system.²⁶ Building the dataset required using sensitive identifiable information and was managed by programmers who were not part of the analysis team.

In BD4LK, there are 335 722 unique birth records for deliveries during 2008-2012. Of these records, 137 960 births (41.1 percent) linked to a Medicaid claim for delivery. Duplicate matches across included data sources occurred in relatively few cases ($n = 1136$, or <1 percent of births). In such cases, we randomly selected one match per birth for sample inclusion and dropped all others. We employed list-wise deletion to manage missing data on other variables (see "Variables" subsection), excluding 1736 birth records (1.3 percent) and yielding a final sample of 136 224 births (see Figure 1).

We identified siblings by flagging births associated with the same unique maternal identifier. In total, there were 111 599 unique mothers across the 136 224 observed births. We identified 89 825 sole births to a mother (65.9 percent of the sample) and identified 46 399 sibling births to mothers (34.1 percent of the sample). Among sibling births, there were 11 546 births (8.5 percent of the sample; 24.9 percent of sibling births) to 5860 mothers with varied PNCC exposure across pregnancies (ie, at least one pregnancy linked to PNCC services and at least one pregnancy did not link to PNCC services). We categorized births in this manner, as some analyses rely on sibling births only (detailed further in "Statistical Analyses" subsection). To compare these groups, we tabulated the baseline characteristics of all sampled births, sibling births, and sibling births with varied PNCC exposure (Table S1).

2.3 | Variables

Our primary exposure was PNCC receipt prior to delivery. We distinguished between any PNCC receipt (no PNCC; any PNCC) and PNCC service level (no PNCC; PNCC assessment/care plan only;

FIGURE 1 Multivariate regression estimates of associations between Prenatal Care Coordination receipt and birth outcomes among Medicaid-covered births, Wisconsin (2008-2012). *Notes:* IPTW models included birth-variant covariates (maternal age in years; maternal education at delivery; maternal marital status at delivery; Supplemental Nutrition Assistance Program receipt 0-6 mo prior to estimated conception; parity; plural birth; maternal tobacco use during pregnancy; maternal chronic hypertension; maternal prepregnancy diabetes; paternal information missing from birth record) and birth-invariant covariates (maternal race/ethnicity; maternal nativity; prior preterm birth). Fixed effects models included birth-invariant covariates only. Birth records supplied the clinical estimate of gestational age in completed weeks. We multiplied that value by 7 to calculate gestational age in completed days for ease of interpreting coefficients. ATE, average treatment effect; ATET, average treatment effect on treated; CI, confidence interval; IPTW, inverse-probability of treatment weighting; PNCC, Prenatal Care Coordination; pt., point; ref., referent

PNCC service uptake). We identified PNCC receipt in Medicaid claims using Healthcare Common Procedure Coding System values indicating care coordination assessment (H1000), care coordination plan (H1002, H1002 U2), education (H1003), follow-up home visit (H1004), and case management (T1016 TH). However, these codes do not identify the specific type of service rendered nor the specific type of provider of the service (eg, nutrition counseling with a dietitian). We coded mothers as having any PNCC if their deliveries linked to a PNCC claim. Likewise, we coded mothers as receiving PNCC assessment/care plan only if they linked to a claim for care coordination assessment and/or planning only, and we coded mothers as receiving PNCC service uptake if they linked to PNCC claims beyond care coordination assessment and planning. We used the PNCC service level variable in addition to the dichotomous PNCC receipt variable to distinguish the potential benefit of receiving PNCC core services from the benefit of only being assessed for services.

We drew birth outcomes from the birth record. We considered birthweight (grams), low birthweight (LBW; <2500 g), gestational age (weeks), and preterm birth (<37 weeks). We used gestational age in days (constructed as gestational age in weeks multiplied by 7) in our models for ease of interpretation. Additionally, we chose these outcomes because PNCC services are intended to intervene on maternal health factors of birthweight and gestational age.^{4,5,11,14}

Depending on model specification, we considered within-mother birth-invariant and birth-varying covariates. Birth-invariant covariates included mother's race/ethnicity and nativity; birth-varying

covariates included mother's age in years (which also accounts for interpregnancy interval), education, marital status, parity, plural birth, prior preterm birth, tobacco use during pregnancy, chronic hypertension, prepregnancy diabetes, whether father information was on the birth certificate, mother's county of residence, and Supplemental Nutrition Assistance Program receipt in the 6 months prior to the estimated conception date.

2.4 | Statistical analyses

2.4.1 | Main analyses

We first regressed birth outcomes on PNCC receipt using inverse-probability of treatment weighting (IPTW), a form of propensity score adjustment,²⁷ to account for mothers' selection into PNCC. We provide two weighted estimates. The average treatment effect (ATE) reflects the expected difference in the outcome for mothers in the population of women eligible to receive services who actually received PNCC compared to those who did not. The average treatment effect on the treated (ATET) reflects the expected difference in the outcome for those mothers who received PNCC relative to what would have been expected for *those same* mothers had they not received services.

If conditioning on observables does not sufficiently address confounding, however, the IPTW estimates remain biased. Sibling fixed effects models further improve on IPTW (as well as other

propensity score methods) for a subset of the PNCC target population because they account for unobserved characteristics of mothers that are stable across their pregnancies, thus allowing for within-mother (across sibling births) comparisons.²⁸ However, the models are only identified for mothers who had multiple births during the sample period and whose births varied on PNCC receipt. Our sibling fixed effects analysis included 46 399 sibling births (34.1 percent of all observed births); 11 546 sibling births (8.5 percent of all observed births and 24.9 percent of observed sibling births) with varied PNCC exposure contributed to the fixed effects estimator. While the remaining 34 853 sibling births did not influence the estimator value, their inclusion in the analysis improves the estimator's precision.

The fixed effects estimator also relies on variation in the outcome. In the case of continuous outcomes, this requirement is not problematic; we find that roughly 60 percent of the variance in birthweight and gestational age was among siblings, whereas the rest occurred between children born to different mothers. In the case of the nominal outcomes LBW and preterm birth, however, the strain on the data is greater, as fixed effects in nonlinear models can only be identified if a mother experienced, for example, a LBW and a non-LBW birth. To protect against sample loss, we estimated PNCC's association with nominal outcomes using linear probability models with sibling fixed effects, which do not require within-mother variation on the outcome.

Differences in estimates between fixed effects regressions and IPTW regressions may indicate that fixed effects estimates account for additional unobserved birth-invariant confounders. They may also reflect differences in sample composition between the two analyses.²⁹ We estimated IPTW ATE and ATET regressions using only the fixed effects subsample ($n = 46\,399$) to better determine the source of potentially disparate results across models.

PNCC's services can intervene on factors that may affect birthweight or gestational age, and assessment/care planning should not confer these benefits. If any PNCC is associated with improved birth outcomes, it should be explained by mothers who received services beyond assessment. Thus, we make the following hypothesis: PNCC assessment/care planning alone did not affect birth outcomes, but PNCC service uptake improved birth outcomes.

With the exception of maternal race/ethnicity, maternal nativity, and prior preterm birth, all covariates (see "Variables" subsection) were included in every model. Fixed effects models adjust for stable cluster-level factors, so conditioning on maternal race/ethnicity or maternal nativity is unnecessary.²⁸ Additionally, we did not adjust for prior preterm birth in the sibling fixed effects model as adjusting for an outcome can bias estimates (see "Sensitivity Analyses" subsection, below, for a detailed justification). The University of Wisconsin-Madison minimal risk institutional review board approved all study procedures, and we conducted all statistical analyses with Stata v.15.³⁰

2.4.2 | Sensitivity analyses

We consider four relevant sources of potential bias to our fixed effects estimates: (a) order of prenatal care receipt; (b) Women,

Infants, and Children Food and Nutrition Service (WIC) program receipt; (c) prior birth outcomes; and (d) plural births. Our sensitivity analyses investigate the extent to which these sources may bias our results.

The order of prenatal care receipt across pregnancies may confound the PNCC-birth outcome association, as prior exposure to PNCC may predict later outcomes via spillover; likewise, prior birth outcomes may predict later PNCC receipt.^{14,25} We stratified the sibling fixed effects regressions by order of PNCC receipt to investigate this potential confounder. First, among all sibling births, we kept the first two in-sample births for each sibling group (2847 births excluded). Second, among sibling pairs with varied PNCC exposure, we identified the order of PNCC receipt. Approximately 56 percent of these sibling pairs had PNCC in the first pregnancy only, and 44 percent had PNCC in the second pregnancy only. Third, we created two subsamples: (a) sibling pairs with constant PNCC exposure or with PNCC exposure in the first pregnancy only and (b) sibling births with constant PNCC exposure or with PNCC exposure in the second pregnancy only. We included siblings with constant PNCC exposure in both subsamples, as their inclusion improves the estimator's precision without affecting its value. Fourth, we replicated the fixed effects regressions in both subsamples and compared results.

WIC receipt may also be a confounder, as PNCC assessments can occur in WIC offices,¹⁴ and WIC receipt may increase infant birthweight.³¹ If WIC receipt systematically changed across births with and without PNCC receipt, then WIC receipt could confound the PNCC-birth outcome relation. We cannot adjust for WIC receipt in our overall models, as it was only captured on 2011-2012 birth records, after Wisconsin adopted the 2003 Revised Standard Birth Record. However, for 2011 and 2012, we investigated confounding by conducting chi-square tests of independence for PNCC and WIC receipt.

The relation between prior birth outcomes and subsequent PNCC use may also bias fixed effects estimates. Traditional approaches to sibling carryover analyses—including sibling fixed effects approaches—cannot recover unbiased estimators if an outcome for the older sibling predicts the exposure of the younger sibling.³² Prior birth outcomes have been shown to predict PNCC eligibility,^{14,24,25} which may bias our fixed effects estimates. However, whether that bias is consequential is uncertain. To assess the magnitude of this potential bias, we regressed *subsequent* PNCC use on *prior* birth outcomes using a subsample of only the first two births observed for sampled mothers. Specifically, we regressed the subsequent sibling's PNCC status on two of the younger sibling's birth outcomes: LBW and preterm birth. We used ordinary least-squares regression and adjusted for all previously listed covariates. The resulting estimates provide evidence of the magnitude of association of prior (adverse) birth outcomes on the likelihood of PNCC receipt for a future birth.

Lastly, plural births may also bias estimates, as plural delivery elevates the risk of adverse birth outcomes³³ which, in turn, may increase a mother's likelihood of engaging in PNCC services during subsequent pregnancies. Likewise, a mother may choose to delay or forgo future child birth as a result of her plural delivery. We

TABLE 1 Characteristics of Medicaid-covered births overall and by Prenatal Care Coordination receipt in Wisconsin, 2008-2012

	Full sample (n = 136 224)	No PNCC (n = 103 151)	PNCC assessment/care plan only ^a (n = 10 510)	PNCC service uptake ^a (n = 22 563)
Trimester of PNCC initiation, n (%)				
First trimester	11 293 (8.3)	-	3382 (32.2)	7911 (35.1)
Second trimester	14 122 (10.4)	-	4070 (38.7)	10 052 (44.6)
Third trimester	7650 (5.6)	-	3057 (29.1)	4593 (20.4)
None	103 151 (75.7)	103 151 (100)	-	-
Birthweight in grams, mean (SD)	3257 (590.5)	3262 (594.5)	3279 (560.2)	3219 (584.8)
Low birthweight (<2500 g), n (%)	11 180 (8.2)	8422 (8.2)	722 (6.9)	2036 (9.0)
Gestational age in weeks, mean (SD)	38.7 (2.1)	38.6 (2.2)	38.7 (1.9)	38.7 (2.1)
Gestational age in days, mean (SD)	270.9 (14.7)	270.2 (15.4)	270.9 (13.3)	270.9 (14.7)
Preterm birth (<37 wks), n (%)	12 441 (9.1)	9469 (9.2)	898 (8.5)	2074 (9.2)
Maternal age in years, mean (SD)	25.3 (5.5)	25.7 (5.5)	25.1 (5.3)	23.5 (5.5)
Maternal education, n (%)				
No HS	6108 (4.5)	4174 (4.1)	486 (4.6)	1448 (6.4)
Some HS	26 050 (19.1)	17 711 (17.2)	2040 (19.4)	6299 (27.9)
HS grad/GED	56 701 (41.6)	42 464 (41.2)	4577 (43.6)	9660 (42.8)
1-3 years college	37 913 (27.8)	30 528 (29.6)	2845 (27.1)	4540 (20.1)
4 years college	7806 (5.7)	6838 (6.6)	471 (4.5)	497 (2.2)
5 + years college	1646 (1.2)	1436 (1.4)	91 (0.9)	119 (0.5)
Maternal race/ethnicity, n (%)				
White NH	78 531 (57.7)	61 615 (59.7)	6534 (62.2)	10 382 (46.0)
Black NH	25 593 (18.8)	18 500 (17.9)	1268 (12.1)	5825 (25.8)
Native American NH	3324 (2.4)	2506 (2.4)	252 (2.4)	566 (2.5)
Laotian/Hmong NH	2327 (1.7)	1859 (1.8)	170 (1.6)	298 (1.3)
Asian NH	1823 (1.3)	1550 (1.5)	107 (1.0)	166 (0.7)
Hispanic	21 120 (15.5)	14 407 (14.0)	1902 (18.1)	4811 (21.3)
Other NH	3506 (2.6)	2714 (2.6)	277 (2.6)	515 (2.3)
Maternal marital status, n (%)				
Unmarried	91 225 (67.0)	66 054 (64.0)	7209 (68.6)	17 962 (79.6)
Married	44 999 (33.0)	37 097 (36.0)	3301 (31.4)	4601 (20.4)
Maternal nativity, n (%)				
Foreign-born	18 590 (13.7)	13 594 (13.2)	1392 (13.2)	3604 (16.0)
Native-born	117 634 (86.4)	89 557 (86.8)	9118 (86.8)	18 959 (84.0)
SNAP receipt, 0-6 mo before conception, n (%)	112 182 (82.4)	83 220 (80.7)	9175 (87.3)	19 787 (87.7)
Previous preterm birth, n (%)				
Yes	8718 (6.4)	6872 (6.7)	585 (5.6)	1261 (5.6)
No (non-first birth)	78 014 (57.3)	62 669 (60.8)	6229 (59.9)	9046 (40.1)
No (first birth)	49 492 (36.3)	33 610 (32.6)	3626 (34.5)	12 256 (54.3)
Parity				
First birth	49 492 (36.3)	33 610 (32.6)	3626 (34.5)	12 256 (54.3)
Second birth	39 919 (29.3)	31 659 (30.7)	3363 (32.0)	4897 (21.7)
Third birth	24 700 (18.1)	19 859 (19.3)	2002 (19.1)	2839 (12.6)
Fourth or subsequent birth	22 113 (16.2)	18 020 (17.5)	1519 (14.5)	2571 (11.4)
Plural birth, n (%)	3621 (2.7)	2706 (2.6)	247 (2.4)	668 (3.0)

(Continues)

TABLE 1 (Continued)

	Full sample (n = 136 224)	No PNCC (n = 103 151)	PNCC assessment/care plan only ^a (n = 10 510)	PNCC service uptake ^a (n = 22 563)
Tobacco use during pregnancy, n (%)	37 332 (27.4)	27 611 (26.8)	3175 (30.2)	6546 (29.0)
Chronic hypertension, n (%)	2453 (1.8)	1845 (1.8)	163 (1.6)	445 (2.0)
Prepregnancy diabetes, n (%)	1265 (0.9)	963 (0.9)	65 (0.6)	237 (1.1)
Father information missing on birth record, n (%)	18 159 (13.3)	13 410 (13.0)	1181 (11.2)	3568 (15.8)

Abbreviations: GED, General Equivalency Diploma; HS, high school; NH, non-Hispanic; PNCC, Prenatal Care Coordination; SNAP, Supplemental Nutrition Assistance Program.

^a“PNCC Assessment/Care Plan Only” indicates that a mother linked to a Medicaid claim for PNCC assessment and/or care plan only. “PNCC Service Uptake” indicates that a mother linked to a Medicaid claim for PNCC services beyond development of a care plan.

replicated the main analysis after excluding plural births from our sample (n = 3620 plural births).

3 | RESULTS

3.1 | Baseline characteristics

Table 1 shows characteristics for the full sample of 136 224 births and by PNCC receipt. Since the characteristics are of births and not mothers, some mothers appear in the data multiple times. Overall, PNCC services were provided in association with approximately 25 percent of Medicaid-covered live births (33 073 of 136 224 births). Among pregnancies assessed for PNCC, nearly 70 percent took up services beyond the care planning stage. Most assessed pregnancies received PNCC evaluation during the second trimester, although trimester of PNCC initiation was somewhat earlier during pregnancies that eventually took up services.

On average, pregnancies with PNCC service uptake had infants with lower birthweight (3219 g) relative to pregnancies with no PNCC (3262 g) and to pregnancies with PNCC assessment/care plan only (3279 g). Similarly, PNCC service uptake pregnancies had a higher incidence of LBW (9.0 percent) compared to non-PNCC pregnancies (8.2 percent) and PNCC assessment/care plan only pregnancies (6.9 percent). Average gestational age was nearly identical across PNCC service levels, although the incidence of preterm birth was somewhat lower among PNCC assessment/care plan only pregnancies (8.5 percent) relative to non-PNCC or PNCC service uptake pregnancies (9.2 percent). Pregnancies with PNCC service uptake were more likely for mothers who were younger, less educated, nonwhite, unmarried, had first-time deliveries, or used tobacco during pregnancy.

Table 2 shows the characteristics among births in the sibling fixed effects sample (sibling births with varied PNCC exposure), both overall and by PNCC service level. There were 11 546 births in this subsample, of which 5596 (48.5 percent) received any PNCC. Of observed PNCC-assessed pregnancies in this subsample, approximately 66 percent took up services beyond a care plan. Pregnancies with no PNCC or PNCC service uptake had similar average birthweights (~3200 g) and incidences of LBW (~9.5 to

10 percent) and preterm birth (~10 to 11 percent), whereas PNCC assessment/care plan only pregnancies had a higher average birthweight (3288 g) and lower incidences of LBW (6.9 percent) and preterm birth (8.8 percent). In general, PNCC service uptake births were more likely among mothers who were younger, less educated, nonwhite, unmarried, had first-time deliveries, or used tobacco during pregnancy.

3.2 | Regression results

Table 3 presents PNCC coefficients for all models, and for ease of comparison, Figure 1 plots the coefficient estimates and their 95 percent confidence intervals. Generally, the models show that PNCC is associated with better birth outcomes. IPTW models showed that any PNCC was associated with improved birth outcomes, and ATET estimates were consistently greater than ATE estimates. For example, any PNCC was associated with a 0.6 percentage point reduction (95% CI 0.2-0.9 percentage points) in the probability of LBW in the IPTW ATE model, whereas any PNCC was associated with a 0.8 percentage point reduction (95% CI 0.5-1.2 percentage points) in the probability of LBW in the IPTW ATET model. These corresponded to a 7 percent and a 9 percent reduction in the risk of LBW in the ATE and ATET models, respectively. We observed similar patterns for the probability of preterm birth and for the continuous outcomes of birthweight and gestational age. However, results from these models suggested little or no benefit of the program when we considered PNCC service level. PNCC service uptake was not significantly associated with any birth outcome in the IPTW models, and PNCC assessment/care plan only was associated with a 12.7 g heavier birthweight (95% CI 1.7-23.6 g; ATET estimate) and with a 0.9 percentage point reduction in the probability of LBW (both ATE and ATET estimates).

In contrast to the IPTW models, the sibling fixed effects results suggested that PNCC service uptake improved birth outcomes. For example, any PNCC was associated with 0.9 percentage point reduction (95% CI 0.0-1.8 percentage points) and a 1.2 percentage point reduction (95% CI 0.3-2.1 percentage points) in the probabilities of LBW and preterm birth, respectively. However, PNCC

TABLE 2 Characteristics of Medicaid-covered births to mothers with multiple in-sample births who varied on Prenatal Care Coordination receipt across births, Wisconsin, 2008-2012

	Full sample (n = 11 546)	No PNCC (n = 5950)	PNCC assessment/care plan only ^a (n = 1890)	PNCC service uptake ^a (n = 3706)
Trimester of PNCC initiation, n (%)				
First trimester	1787 (15.5)	-	586 (31.0)	1201 (32.4)
Second trimester	2389 (20.7)	-	758 (40.1)	1631 (44.0)
Third trimester	1414 (12.3)	-	546 (28.9)	868 (23.4)
None	5950 (51.5)	5950 (100)	-	-
Birthweight in grams, mean (SD)	3214 (604.5)	3201 (627.0)	3288 (545.6)	3197 (593.7)
Low birthweight (<2500 g), n (%)	1055 (9.1)	557 (9.4)	130 (6.9)	368 (9.9)
Gestational age in weeks, mean (SD)	38.5 (2.3)	38.4 (2.5)	38.7 (1.8)	38.6 (2.2)
Gestational age in days, mean (SD)	269.5 (16.1)	268.8 (17.5)	270.9 (12.6)	270.2 (15.4)
Preterm birth (<37 wks), n (%)	1199 (10.4)	654 (11.0)	166 (8.8)	379 (10.2)
Maternal age in years, mean (SD)	23.6 (4.8)	23.7 (4.7)	24.4 (4.8)	22.9 (4.8)
Maternal education, n (%)				
No HS	576 (5.0)	291 (4.9)	81 (4.3)	204 (5.5)
Some HS	3189 (27.6)	1608 (27.0)	412 (21.8)	1169 (31.5)
HS grad/GED	5095 (44.1)	2609 (43.9)	850 (45.0)	1636 (44.1)
1-3 years college	2321 (20.1)	1252 (21.0)	441 (23.3)	628 (17.0)
4 years college	303 (2.6)	155 (2.6)	90 (4.8)	58 (1.6)
5 + years college	62 (0.5)	35 (0.6)	16 (0.9)	11 (0.3)
Maternal race/ethnicity, n (%)				
White NH	5266 (45.6)	2697 (45.3)	1101 (58.3)	1468 (39.6)
Black NH	3241 (28.1)	1680 (28.2)	308 (16.3)	1253 (33.8)
Native American NH	431 (3.7)	221 (3.7)	70 (3.7)	140 (3.8)
Laotian/Hmong NH	259 (2.2)	112 (1.9)	43 (2.3)	104 (2.8)
Asian NH	81 (0.7)	33 (0.6)	14 (0.7)	34 (0.9)
Hispanic	1930 (16.7)	991 (16.7)	299 (15.8)	640 (17.3)
Other NH	338 (2.9)	216 (3.6)	55 (2.9)	67 (1.8)
Maternal marital status, n (%)				
Unmarried	8399 (72.7)	4270 (71.8)	1217 (64.4)	2912 (78.6)
Married	3147 (27.3)	1680 (28.2)	673 (35.6)	794 (21.4)
Maternal nativity, n (%)				
Foreign-born	1437 (12.5)	738 (12.4)	215 (11.4)	484 (13.1)
Native-born	10 109 (87.6)	5212 (87.6)	1675 (88.6)	3222 (86.9)
SNAP receipt, 0-6 mo before conception, n (%)	10 828 (93.8)	5581 (93.8)	1713 (90.6)	3534 (95.4)
Previous preterm birth, n (%)				
Yes	984 (8.5)	527 (8.9)	146 (7.7)	311 (8.4)
No (non-first birth)	7619 (66.0)	4278 (71.9)	1330 (70.4)	2011 (54.3)
No (first birth)	2943 (25.5)	1145 (19.2)	414 (21.9)	1384 (37.3)
Parity				
First birth	2943 (25.5)	1145 (19.2)	414 (21.9)	1384 (37.3)
Second birth	4101 (35.5)	2427 (40.8)	678 (35.8)	996 (26.9)
Third birth	2311 (20.0)	1225 (20.6)	420 (22.2)	666 (18.0)
Fourth or subsequent birth	2193 (19.0)	1153 (19.4)	378 (20.0)	660 (17.8)
Plural birth, n (%)	293 (2.5)	139 (2.3)	36 (1.9)	118 (3.2)

(Continues)

TABLE 2 (Continued)

	Full sample (n = 11 546)	No PNCC (n = 5950)	PNCC assessment/care plan only ^a (n = 1890)	PNCC service uptake ^a (n = 3706)
Tobacco use during pregnancy, n (%)	3399 (29.4)	1779 (29.9)	543 (28.7)	1077 (29.1)
Chronic hypertension, n (%)	231 (2.0)	115 (1.9)	30 (1.6)	86 (2.3)
Prepregnancy diabetes, n (%)	90 (0.8)	50 (0.8)	9 (0.5)	31 (0.8)
Father information missing on birth record, n (%)	1934 (16.8)	971 (16.3)	258 (13.7)	705 (19.0)

Abbreviations: GED, General Equivalency Diploma; HS, high school; NH, non-Hispanic; PNCC, Prenatal Care Coordination; SNAP, Supplemental Nutrition Assistance Program.

^a“PNCC Assessment/Care Plan Only” indicates that a mother linked to a Medicaid claim for PNCC assessment and/or care plan only. “PNCC Service Uptake” indicates that a mother linked to a Medicaid claim for PNCC services beyond development of a care plan.

service uptake was associated with a 1.3 percentage point reduction (95% CI 0.3-2.4 percentage points) and a 1.8 percentage point reduction (95% CI 0.8-2.9 percentage points) in the probabilities of LBW and preterm birth, respectively. These corresponded to a 14 percent lower relative risk of LBW and a 17 percent lower relative risk of preterm birth. Similar patterns emerged for the continuous outcomes—PNCC service uptake was associated with a 42.9-gram increase in birthweight (95% CI 24.9-60.8 g) and a 1.5-day longer gestation (95% CI 0.9-2.0 days). Moreover, PNCC assessment/care plan only was not associated with any birth outcome in fixed effects models.

Results from the sibling fixed effects models potentially indicate that the IPTW estimates are downwardly biased by unobserved birth-invariant confounders. To explore whether the model’s functional form or the sibling sample drive these differences, we estimated IPTW ATE and ATET regressions using only the fixed effects model subsample (Table S2). Overall, IPTW estimates from the fixed effects subsample are somewhat larger than those from the full sample. However, the patterns of our IPTW models are consistent regardless of sample composition: Any PNCC is associated with improved birth outcomes with the dichotomous exposure but using the categorical PNCC exposure nullifies these associations. The model with continuous birthweight is the exception; assessment/care plan only is associated with a ~23-g increase in birthweight in both ATE and ATET models. Nonetheless, these findings confirm that the fixed effects model’s sample does not fully explain its estimates; rather, it appears that adjusting for stable maternal characteristics is the driving difference.

3.3 | Sensitivity analyses results

We evaluated the robustness of our estimates with three sensitivity analyses as described in the “Sensitivity Analyses” subsection (results not shown). First, we stratified the fixed effects regressions by order of PNCC receipt (prior or subsequent birth) to determine whether birth order confounded those estimators. We found that PNCC was associated with healthier birth outcomes regardless of the order of PNCC receipt. Estimates were stronger for mothers who received PNCC during the second pregnancy (53.6 g and 1.4 days) than for

mothers who received it only during the first pregnancy (32.1 g and 1.1 days, respectively). Nonetheless, care coordination was measurably associated with birth outcomes. This is notable given the fact that mothers who received care coordination during the first in-sample birth were at most risk of a carryover effect (a prior experience of care coordination could diminish the contrast between the outcome of a pregnancy with care coordination services and the outcome of a pregnancy without), and that mothers may be more likely to receive care coordination services when they are at higher risk of an adverse birth outcome.

Second, we tested the association between WIC receipt and PNCC receipt in the subsample of 2011-2012 births for which WIC participation data were available from the birth records. As anticipated, mothers receiving WIC were more than four times as likely to receive any care coordination. This difference was entirely attributable to their greater likelihood of being assessed for services; conditional on assessment, approximately 70 percent of mothers receiving WIC took up services, compared to 71 percent of mothers who were not receiving WIC. Since our results show that PNCC’s effect on birth outcomes depended on service uptake, we suspect minimal confounding of WIC receipt.

Third, we used a subsample consisting of only the first two observed siblings in the fixed effects sample and tested the association between the first sibling’s birth outcomes and the second sibling’s exposure to PNCC. Preterm birth for the first sibling was not associated with the second sibling’s exposure to PNCC, and LBW for the first sibling only increased the probability of subsequent PNCC exposure by 2.5 percent. Moreover, several maternal characteristics—age, education, marital status, SNAP receipt, and chronic hypertension—were more strongly associated with the second sibling’s exposure to PNCC. These results suggest that prior birth outcomes have, at most, only modest influence on future PNCC use for mothers. Therefore, it is unlikely that our sibling fixed effects estimates are substantially biased as a result of selection into PNCC participation based on prior birth outcomes.

Lastly, we replicated the main IPTW and sibling fixed effects analysis after excluding plural births from our sample. All results from this analysis were nearly identical to those from our full sample (Table S3).

TABLE 3 Multivariate regression estimates of associations between Prenatal Care Coordination receipt and selected birth outcomes among Medicaid-covered births using Inverse-Probability of Treatment Weighting (IPTW) and sibling fixed effects models, Wisconsin (2008-2012)

	IPTW, average treatment effect (n = 136 224) ^a	IPTW, average treatment effect on treated (n = 33 073) ^a	Sibling fixed effects (n = 46 399) ^b
Birthweight, change in grams (95% CI)			
Dichotomous PNCC exposure			
None	Reference	Reference	Reference
Any	10.6 (3.2, 18.0)**	19.3 (12.0, 26.5)***	34.7 (19.2, 50.2)***
Categorical PNCC exposure			
None	Reference	Reference	Reference
Assessment/care plan only ^c	11.3 (−0.3, 22.9)	12.7 (1.7, 23.6)*	19.9 (−4.2, 42.3)
Service uptake	−0.7 (−10.2, 8.9)	1.0 (−8.7, 10.7)	42.9 (24.9, 60.8)***
Low birthweight (<2500 g), change in percentage points (95% CI)			
Dichotomous PNCC exposure			
None	Reference	Reference	Reference
Any	−0.6 (−0.9, −0.2)**	−0.8 (−1.2, −0.5)***	−0.9 (−1.8, −0.0)*
Categorical PNCC exposure			
None	Reference	Reference	Reference
Assessment/care plan only ^c	−0.9 (−1.5, −0.4)**	−0.9 (−1.4, −0.4)***	−0.1 (−1.4, 1.3)
Service uptake	−0.1 (−0.5, 0.4)	0.1 (−0.3, 0.5)	−1.3 (−2.4, −0.3)*
Gestational age, change in days (95% CI) ^d			
Dichotomous PNCC exposure			
None	Reference	Reference	Reference
Any	0.2 (0.1, 0.4)**	0.4 (0.2, 0.6)***	1.1 (0.7, 1.6)***
Categorical PNCC exposure			
None	Reference	Reference	Reference
Assessment/care plan only ^c	0.2 (−0.1, 0.5)	0.1 (−0.1, 0.4)	0.4 (−0.2, 1.1)
Service uptake	0.1 (−0.1, 0.3)	0.0 (−0.2, 0.3)	1.5 (0.9, 2.0)***
Preterm birth (<37 wks), change in percentage points (95% CI)			
Dichotomous PNCC exposure			
None	Reference	Reference	Reference
Any	−0.4 (−0.8, −0.0)*	−0.7 (−1.0, −0.3)***	−1.2 (−2.1, −0.3)*
Categorical PNCC exposure			
None	Reference	Reference	Reference
Assessment/care plan only ^c	−0.3 (−0.9, 0.3)	−0.2 (−0.8, 0.3)	0.1 (−1.3, 1.4)
Service uptake	−0.1 (−0.6, 0.3)	0.0 (−0.4, 0.5)	−1.8 (−2.9, −0.8)**

Abbreviations: GED, General Equivalency Diploma; HS, high school; NH, non-Hispanic; PNCC, Prenatal Care Coordination; SNAP, Supplemental Nutrition Assistance Program.

^aIPTW models included birth-variant covariates (maternal age in years; maternal education at delivery; maternal marital status at delivery; Supplemental Nutrition Assistance Program receipt 0-6 mo prior to estimated conception; parity; plural birth; maternal tobacco use during pregnancy; maternal chronic hypertension; maternal prepregnancy diabetes; paternal information missing from birth record) and birth-invariant covariates (maternal race/ethnicity; maternal nativity; prior preterm birth).

^bSibling fixed effects models included birth-variant covariates, as these models automatically control for birth-invariant confounders without their manual inclusion in the model.

^c“Assessment/care plan only” indicates that a mother linked to a Medicaid claim for PNCC assessment and/or care plan only. “Service uptake” indicates that a mother linked to a Medicaid claim for PNCC services beyond development of a care plan.

^dBirth records supplied the clinical estimate of gestational age in completed weeks. We multiplied that value by 7 to calculate gestational age in completed days for ease of interpreting coefficients.

*P-value <.05; **P-value <.01; ***P-value <.001.

4 | DISCUSSION

We found strong evidence that PNCC service uptake among pregnant Medicaid beneficiaries in Wisconsin decreased their risks of LBW and preterm birth, increased birthweight, and lengthened gestation. Additionally, PNCC assessment/care plan only was not associated with improved birth outcomes, confirming our hypothesis that actual service receipt drives PNCC's potential benefit. These findings align with the results of recent quasi-experimental analyses of care coordination programs in Iowa,²³ which found a dosage effect of participation on preterm birth and LBW; North Carolina,¹⁷ which recorded a nearly 2 percentage point lower rate of preterm birth among participating mothers; and Michigan,²² which found lower odds of LBW and preterm birth among mothers who received care coordination, when compared to the outcomes of matched controls. Our results also parallel those obtained in a previous investigation of Wisconsin's PNCC program, which found that care coordination receipt was associated with lower odds of LBW and preterm birth.¹³

We also found that the sibling fixed effects models, which most rigorously adjusted for observed and unobserved birth-invariant confounders and selection into care coordination services, and which have not been used in prior work, produced the largest estimated effects of PNCC. This suggests that prior studies adjusting only for observable attributes may have underestimated the program's effects. Alternately, it may be that the subpopulation of Medicaid-covered mothers with multiple births in the observed five-year interval was, for some unobserved reason, more sensitive to PNCC's effects. If we assume, however, that the period of observation is random with respect to birth outcomes—a reasonable assumption—then the fixed effects estimates are arguably applicable to eligible Medicaid-covered mothers with more than one child whose birth spacing is 60 months or less.

In sum, our work shows that sibling fixed effects regressions best measure the effect of care coordination relative to other conventional regression techniques, and that care coordination dosage matters. Absent an experiment, we cannot confirm that our estimates are truly unbiased. Nonetheless, our fixed effects estimators are highly protective against bias from stable unobserved confounding and treatment selection factors within mothers. Prior quasi-experimental studies used propensity score matching to measure care coordination's impact on birth outcomes,^{17-20,22,23} but this approach cannot control for any unobserved confounding. For this reason, we believe that our fixed effects estimators are among the most internally valid and minimally biased estimates of care coordination on birth outcomes produced to date. Our findings may also support the clinical relevance of PNCC. For example, maternal tobacco use is a modifiable cause of adverse birth outcomes,³⁴ and smoking cessation during pregnancy may reduce the risk of LBW by 17 percent.³⁵ We found that PNCC service uptake reduced the risk of LBW by 14 percent—a potential benefit that is comparable to that of maternal smoking cessation. Moreover, the estimated effects of PNCC likely depend on program fidelity. Among mothers who received PNCC services, some may have initiated services late into the pregnancy or

may have not adhered to their care plans. Such factors could attenuate PNCC's benefit.

It is fitting that mothers who appeared to be at elevated risk of adverse birth outcomes were more likely to participate in PNCC (see Table 1), as it aligns with the program's intent. However, only 25 percent of Medicaid-covered women were assessed for PNCC, which means that three out of four Medicaid-covered pregnant women in Wisconsin during our sample period had no opportunity to take up or benefit from services. The possible underutilization of PNCC has been highlighted in prior research.²⁴ These findings underscore the untapped potential of PNCC: There is robust evidence that services are associated with clinically improved birth outcomes, yet the program is widely underutilized by otherwise eligible Medicaid beneficiaries. County- or provider-level variation in service outreach may partially explain this. With approval from the Wisconsin Department of Health Services, counties and individual PNCC providers may restrict (but not widen) eligibility for PNCC service receipt to target specific beneficiary populations, such as first-time mothers.¹⁴ Other reasons for PNCC underutilization may include a lack of awareness about PNCC, competing resources for beneficiaries such as work, and little coordination between prenatal care providers and other PNCC service providers. Expanding the reach of PNCC's services could potentially improve maternal and infant health at a population level.

4.1 | Limitations

Despite its strengths, our work has limitations that merit consideration. Estimates from the IPTW regressions cannot account for any unobserved confounder. These regressions allow for estimating PNCC's impact in a comprehensive population of Medicaid-covered mothers in Wisconsin, yet the estimates are almost certainly biased due to omitted variables. The sibling fixed effects models protect against mother-invariant confounders, but this advantage is not without its trade-offs. First, the smaller sample size for mothers with multiple births necessarily reduced estimate precision. Second, the fixed effects subsample represented disproportionately higher-risk profile births relative to the full population of Medicaid-covered births. Findings from this subsample may have less generalizability to all Medicaid-covered births. Moreover, we cannot entirely rule out that selection into treatment accounted for at least some of PNCC's observed benefits. This may be indicative of effect heterogeneity by risk-level—the benefit of PNCC increases alongside the risk of the pregnancy. Third, sibling fixed effects estimates may be biased because the birth outcomes for one pregnancy may predict PNCC uptake for a subsequent pregnancy, although our sensitivity analysis suggests that this bias is likely minimal. Additionally, fixed effects estimates were subject to bias from unobserved confounders that vary across mothers' births. Fourth, PNCC in Wisconsin is primarily managed at the county level, which allows service delivery to vary in idiosyncratic ways. We are unable to assert that our estimates apply consistently to mothers across the state; rather, they

represent the average statewide effect. Moreover, other states have implemented care coordination programming, but states vary in design and implementation so our estimates may not apply elsewhere. Finally, PNCC billing codes neither indicate the specific service that a mother received nor the type of provider who rendered the service, so we cannot identify the precise mechanism of PNCC's potential benefit.

5 | CONCLUSIONS

Wisconsin's PNCC program may have prevented LBW and preterm birth, and the program's benefits were most pronounced among mothers who used services beyond care planning. This finding, in conjunction with supporting literature, should inform stakeholders and policy makers interested in reforming and enhancing program outreach and implementation. Expanding PNCC services may improve population health in a modest but significant way among Medicaid-covered pregnant women in Wisconsin and likely in other states.

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SUPPORTING INFORMATION

Additional supporting information may be found online in the Supporting Information section.

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