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Maternal ambient air pollution, preterm birth, and markers of fetal growth in Rhode Island: Results of a hospital-based linkage study

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Abstract

Background—Maternal exposure to ambient air pollution has been associated with higher risk of preterm birth and reduced fetal growth, but heterogeneity among prior studies suggests additional studies are needed in diverse populations and settings. We examined the associations between maternal ambient air pollution levels, risk of preterm birth, and markers of fetal growth in an urban population with relatively low exposure to air pollution.

Methods—We linked 61,640 mother-infant pairs who delivered at a single hospital in Providence, Rhode Island from 2002–2012 to birth certificate and hospital discharge data. We used spatial-temporal models and stationary monitors to estimate exposure to fine particulate matter (PM$_{2.5}$) and black carbon (BC) during pregnancy. Using generalized linear models we evaluated the association between pollutant levels, risk of preterm birth, and markers of fetal growth.

Results—In adjusted models, an interquartile range (IQR; 2.5 µg/m$^3$) increase in pregnancy-average PM$_{2.5}$ was associated with odds ratios (ORs) of preterm birth of 1.04 (95% CI: 0.94, 1.15) and 0.86 (0.76, 0.98) when considering modeled and monitored PM$_{2.5}$, respectively. An IQR increase in modeled and monitored PM$_{2.5}$ was associated with a 12.1 g (95% CI:−24.2, −0.1) and...
15.9 g (95% CI: −31.6, −0.3) lower birthweight. Results for BC were highly sensitive to choice of exposure metric.

Conclusion—In a population with relatively low exposures to ambient air pollutants, PM$_{2.5}$ was associated with reduced birthweight but not with risk of preterm birth.

Keywords
air pollution; birthweight; preterm; fetal growth

Introduction
Preterm birth (<37 weeks gestation) is the leading cause of infant mortality and has been linked to increased risk of cognitive, behavioral, and medical disabilities in children and increased risk of adult diseases [1–4]. Globally, there are 15 million preterm births each year with the number increasing in many countries [5]. A growing literature suggests that maternal exposure to ambient fine particulate matter air pollution (PM$_{2.5}$) is associated with adverse birth outcomes [6–13], with the strongest evidence existing for reduced fetal growth [6–9].

Some prior studies have reported that PM$_{2.5}$ is associated with higher risk of preterm birth [6, 9, 14]. Under the assumption that these associations reflect widely generalizable causal effects, it has been estimated that in 2010 ambient PM$_{2.5}$ was responsible for 15,808 preterm births in the US (3.3% of all premature births) and a staggering 2.7 million (18% of all premature births) globally [15, 16]. However, a closer examination of the underlying evidence suggests the association between PM$_{2.5}$ and preterm birth remains incompletely understood. For example, two recent large studies (one in Canada including ∼2.5 million births and one in New York City including ∼258,000 births), as well as a few smaller studies using a novel matched design, and an analysis of detailed individual-level data from the European ESCAPE birth cohorts all failed to find an association between PM$_{2.5}$ and risk of preterm birth [17–22]. This heterogeneity in results across prior studies suggests the need for additional studies examining the potential impact of PM$_{2.5}$ on risk of preterm birth in diverse populations and settings.

Accordingly, we examined the associations between PM$_{2.5}$ and risk of preterm birth in Rhode Island (RI), a small coastal New England state where the majority of children are born at one hospital. We also estimated the association between black carbon (BC, a marker of traffic pollution) on preterm birth, an association that has been less extensively studied. Finally, we estimated the association between PM$_{2.5}$, BC, and markers of fetal growth.

Methods
Study Population
The study population consists of mother-infant pairs of women delivering at Women and Infants Hospital of Rhode Island (WIHRI) between 2001 and 2012. Located in Providence, RI, WIHRI is the 9th largest stand-alone obstetrical service in the United States with ∼8,500 deliveries annually [23]. Approximately 75% of pregnant women living in RI give birth at
and therefore this study sample is representative of the population living in this area. WIHRI provides hospital records, including individual-level demographic, clinical, and financial data on all deliveries to the National Perinatal Information Center (NPIC). We obtained these records from the NPIC Perinatal Center Data Base and merged them with birth certificate data from the RI Department of Health (RIDOH). The study protocol was approved by the Institutional Review Boards of Brown University, WIHRI, and RIDOH.

Initially we matched 76,590 (79.8%) of the 95,948 hospital discharge records from WIHRI to state birth certificates, but upon closer examination few births were matched to birth certificate data in 2001 or between July 2004 and December 2005 as RIDOH transitioned to a new birth records system. Thus, we restricted our analyses to 74,165 deliveries occurring during time periods with high match rates (>88.6%, January 2002-June 2004 and January 2006-December 2012) and successfully matched to birth certificate data. We used ArcMap (ESRI, Redlands, CA) to geocode maternal residential addresses listed on birth certificates and linked 2010 census and geographic data to each address point. We excluded women aged <18 years at delivery (1,575) or missing maternal age (2,198); those who had multiple births (3,176); and those with addresses outside of RI (7,418) or missing (14). We also excluded women missing data for birthweight (6) or those with an implausible birthweight of <500 g or >5000 g (194). Our final sample included 61,640 mother-infant pairs.

**Exposure Assessment**

We estimated daily PM$_{2.5}$ levels at each maternal residential address using a hybrid of land-use regression and satellite remote sensing, as previously described [24]. Briefly, the model ($R^2=0.88$) uses a land-use regression model that includes spatial and temporal factors and satellite measurements of aerosol optical depth (AOD) on a 1 km grid and fits a daily calibration regression using ground-level PM$_{2.5}$ measurements. Differences between grid cell AOD and measured PM$_{2.5}$ are regressed against local land use features to generate estimates of local source (mostly automobile traffic) exposure on a finer scale (200 m × 200 m).

We also examined BC as a marker of traffic-related air pollution. We estimated daily BC levels at each address using an extended version of a validated spatial-temporal land-use regression model [25]. Briefly, this model incorporates daily average BC measurements from five RI monitors, meteorological data from nearly 2 dozen local weather stations, land use data, latitude and longitude, daily meteorological factors, and interaction terms between land use and daily meteorological factors. The model performed well in cold (November – April) and warm (May – October) seasons (10 fold cross-validated $R^2$ of 0.73 and 0.75, respectively).

To facilitate comparison with previous studies, we also obtained daily PM$_{2.5}$ and BC measurements from stationary monitors operated by the RI Department of Environmental Management. We obtained PM$_{2.5}$ data from six monitors in Providence County and BC data from two monitors in Providence and East Providence, RI and calculated daily averages of each pollutant.
Thus, modeled and monitored PM$_{2.5}$ estimates were available for deliveries taking place anytime during the study period (2002–2012), modeled BC estimates were available for 2004–2011, and monitored BC was available for 2005–2012. For both monitored and modeled pollutant levels, we averaged the daily exposure estimates to determine the average levels for the entire pregnancy as well as for the first trimester (weeks 1–12), second trimester (weeks 13–26), and third trimester (weeks 27 to birth), as in previous studies [26]. While our primary analyses were based on pregnancy-average pollutant levels, we additionally investigated trimester-specific exposures as vulnerability to air pollution may change over the course of pregnancy [9, 14, 27].

**Outcome Assessment**

We defined preterm birth as births before 37 completed weeks of gestation. Term birthweight, small for gestational age (SGA), and low birthweight (LBW) were used as markers of fetal growth among term births. Birthweight was obtained from birth certificates and was measured in grams. SGA was defined as the lowest 10$^{th}$ percentile for gestational age and sex based on 1999 and 2000 U.S. births [28–31]. We defined LBW as term births having a birthweight <2500 g.

**Covariates**

We considered the following potential confounders: maternal age, parity (0, 1, or ≥2), maternal race (Black, White, Other), maternal education (less than high school, high school, attended college but did not graduate, graduated from college, attended graduate school), marital status (married, single, other: divorced, separated, widowed, unknown), health insurance (private, public, other: Medicare, Champus, self-coverage, other coverage, coverage unknown), and tobacco use during pregnancy (yes, no). Public insurance (Medicaid or Medicaid/HMO) during this time was only provided to low income women. Gestational age was measured in weeks.

We assessed the following six census tract characteristics to address potential confounding by neighborhood SES: median household income; percent of households with interests, dividends, or rent income; percent of residents with high school diploma; percent with college degree; percent with professional occupation; and median value of owner-occupied housing units. We calculated a z-score for each variable and summed the scores to create a z-sum score [32].

**Statistical Analyses**

We used linear regression to quantify the association and 95% confidence intervals (CI) between pollutants and birthweight and used logistic regression models to obtain odds ratios and 95% CIs of the association between pollutants and odds of preterm birth, SGA, and LBW. Analyses examining birthweight, SGA, and LBW were restricted to term births only. We used causal diagrams to identify potential confounders requiring adjustment in each model [33]. All models were adjusted for maternal age (modeled as a natural spline with 3 degrees of freedom), tobacco use during pregnancy, parity, education, race, insurance, marital status, neighborhood SES z-sum score, and year of last menstrual period modeled as
a factor to account for secular trends in pollutant levels (Supplemental Figure 1). Models for
birthweight and LBW were additionally adjusted for gestational age.

Our primary definition of preterm birth based on gestational age at delivery does not
distinguish between spontaneous and induced preterm deliveries. We hypothesize that air
pollution would be more strongly associated with spontaneous than medically-induced
preterm birth. Thus, we conducted exploratory analyses to examine the association between
each exposure and preterm birth identified by ICD9 code 644.21, which defines preterm
birth as the onset of spontaneous delivery before 37 completed weeks of gestation [34].

We conducted all analyses using R (v3.2.0) and imputed missing covariate data using the
Multivariate Imputation by Chained Equations (MICE) package [35]. Missing covariate data
included: tobacco (13.7% missing), parity (2.4%), maternal education (3.7%), maternal race
(8.4%), maternal insurance (0.6%), and marital status (1.0%). We also conducted a
complete-case analysis examining the associations between pregnancy-average air pollution
levels and birthweight, SGA, and LBW adjusted for the same covariates as the primary
analyses.

Results

Most of the women in this study were white (64.7%), married (62.6%), and did not use
tobacco during pregnancy (93.4%) (Table 1). The mean maternal age was 29 (standard
deviation, SD = 5.9) years. The mean gestational age at delivery was 39 (SD = 2.0) weeks
with about 8% of births delivered preterm. Among term births, we found that gestational
age, parity, maternal race, maternal education, marital status, health insurance, and tobacco
use during pregnancy were associated with birthweight (Supplemental Table 1). Modeled
pregnancy-average PM$_{2.5}$ levels had a mean of 9.5 µg/m$^3$ (SD=1.5), which is below the
current National Ambient Air Quality Standard for annual PM$_{2.5}$ of 12 µg/m$^3$ [36]. Levels of
modeled PM$_{2.5}$ and BC were similar to monitored levels (Supplemental Table 2).

In fully adjusted models, modeled pregnancy-average PM$_{2.5}$ was not associated with risk of
preterm birth (Table 2). In secondary analyses considering trimester-specific PM$_{2.5}$, only
first-trimester PM$_{2.5}$ was associated with a lower risk of preterm birth (OR=0.93; 95% CI:
0.88, 0.98). Results were qualitatively similar, when instead considering measured PM$_{2.5}$
from stationary monitors and when considering risk of spontaneous preterm birth identified
from discharge diagnosis codes rather than gestational age (Supplemental Table 3).

Pregnancy-average PM$_{2.5}$ was associated with a 12.1 g lower (95% CI: −24.2, −0.1 g)
birthweight per interquartile range (IQR; 2.5 µg/m$^3$) increase in model-estimated PM$_{2.5}$
(Table 3). Results were qualitatively similar when considering instead PM$_{2.5}$ measured at
stationary monitors. Pregnancy-average PM$_{2.5}$ was associated with higher risk of being born
SGA, but this only reached statistical significance for monitor-estimated PM$_{2.5}$ (Table 4;
OR=1.15; 95% CI: 1.00, 1.31). PM$_{2.5}$ was not associated with risk of being born LBW. The
complete-case analysis results were similar to the main results with imputation, but have
wider 95% CIs (Supplemental Table 4).
We also considered the association between BC and birth outcomes. Modeled BC was not associated with risk of preterm birth (Table 2). Monitored BC levels were positively associated with preterm birth when averaged over the entire pregnancy but negatively associated in the third trimester. Model-estimated BC was not associated with birthweight and results were more extreme with wider CIs when considering monitored BC (Table 3). BC was not associated with risk of SGA and LBW, except for monitored second trimester BC (OR=0.88; 95% CI: 0.79, 0.98). Complete-case analysis results were similar to the main analysis with multiple imputation, but have wider 95% CIs (Supplemental Table 4).

Discussion

Among more than 60,000 deliveries, we found no evidence to suggest that PM$_{2.5}$ is associated with higher risk of preterm birth in the current study population in RI. Indeed, the only statistically significant associations between PM$_{2.5}$ and risk of preterm birth were negative rather than positive. Results were qualitatively similar when considering PM$_{2.5}$ levels estimated by a spatial-temporal model versus measured values from stationary monitors. On the other hand, we did observe the expected association between PM$_{2.5}$ and lower birthweight.

The lack of association between PM$_{2.5}$ and preterm birth stands in contrast to a number of prior studies. Pooled estimates of the relative risk for preterm birth from four recent meta-analyses ranged from 1.05 (95% CI: 0.98, 1.13) to 1.15 (95% CI: 1.14, 1.16) per 10 µg/m$^3$ change in PM$_{2.5}$ [6, 9, 14, 37]. Comparatively, our results for pregnancy-average PM$_{2.5}$ rescaled to a 10 µg/m$^3$ increase would yield an estimate of 1.17 (95% CI: 0.89, 1.75), which is within the CIs of three of these estimates. Thus, our results could be interpreted as consistent with the prior evidence, but lacking statistical significance due to our smaller sample size. However, our results are closer to those from several studies that have either failed to find an association between PM$_{2.5}$ and preterm birth or have reported a negative association [17–22]. These findings suggest that there is little evidence of an increased risk of preterm birth associated with PM$_{2.5}$ levels in this and at least some other populations.

Our results for PM$_{2.5}$ and birthweight are similar in direction and magnitude to prior studies. For example, in New York City Savitz, et al. [26] found that a 10 µg/m$^3$ increase in PM$_{2.5}$ was associated with a 48.4 g (95% CI: −62.3, −34.5) lower birthweight. For comparison, extrapolating our results to a 10 µg/m$^3$ increase in PM$_{2.5}$ yields a 48.5 g (95% CI: −96.6, −0.44) difference in birthweight. In neighboring Massachusetts, Kloog, et al. [38] applied the same spatial-temporal model as our study to estimate PM$_{2.5}$ and found a more modest 13.8 g (95% CI: −21.10, −6.05) lower birthweight per 10 µg/m$^3$ increase in PM$_{2.5}$. More broadly, in comparison to estimates reported in recent meta-analyses [7, 8, 37] our estimates of the association between PM$_{2.5}$ and birthweight are comparable to some and larger than others. However, our estimates have wider CIs that include the null hypothesis of no association when considering trimester-specific exposures.

Consistent with our findings for birthweight, pregnancy-average PM$_{2.5}$ was associated with higher risk of SGA, although these results only reached statistical significance when considering monitored PM$_{2.5}$ levels. These associations were in the same direction, but

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larger in magnitude compared to estimates from recent meta-analyses. In particular, our finding of relative risks of SGA of 1.41 (95% CI: 0.92, 2.16) per 10 µg/m$^3$ modeled PM$_{2.5}$ are somewhat larger than the pooled estimate of 1.15 (95% CI: 1.10, 1.20) reported previously for SGA [7–9, 14, 27, 37]. However, our CIs are wide and include the null, suggesting that our study may have been underpowered to detect associations of this magnitude.

Our findings for BC are difficult to interpret since we unexpectedly found discrepant results depending on whether we used BC estimated from a land-use regression model or from stationary monitors, and the time period being considered. For example, we observed that monitored pregnancy-average BC was associated with a pronounced higher risk of preterm birth while third trimester BC was associated with a significantly lower risk of preterm birth. Similarly contradictory results were observed between monitored BC and birthweight, suggesting that estimates of associations with monitored BC are unstable and should be interpreted with caution. On the other hand, modeled BC was not associated with either risk of preterm birth or birthweight. Interestingly, our results for BC are consistent with findings by Brauer, et al. [39] who estimated residential BC from a land-use regression model and found no association between BC and preterm birth (defined as birth <30 weeks of gestation; OR=0.99, 95% CI: 0.87, 1.13) in Vancouver, Canada. Similarly, in the Dutch PIAMA study, Gehring, et al. [40] did not find a statistically significant association between soot (a component of BC, estimated with a land-use regression model) and preterm birth (OR=1.08, 95% CI: 0.88, 1.34). Thus, the effects of BC and other markers of traffic pollution on risk of preterm birth require further study.

Our study has important limitations. First, as is common to studies based on large-scale administrative data, we expect exposure misclassification due to lack of information on residential history throughout pregnancy, amount of time spent at home, and housing characteristics such as air conditioning and air ventilation. However, we used sophisticated spatial-temporal models to estimate PM$_{2.5}$ and BC levels at the residence representing an improvement over some prior studies and resulting in lower exposure misclassification compared to using stationary monitors alone. Second, our results may not be generalizable to pregnant women living in metropolitan regions outside of the study area. Third, there may be residual confounding due to misclassification of tobacco use during pregnancy and unmeasured data for diet, exposure to environmental tobacco smoke, substance abuse, and alcohol consumption during pregnancy. Also, information on tobacco use during pregnancy was missing in 13.4% of women, which potentially may not be missing at random, even when multiply imputed conditional on other available covariates. However, results of a complete-case analysis were similar to the main analyses with multiple imputation. Fourth, average pollutant levels in the study area were relatively low with limited variation over space or time, limiting our statistical power to detect associations of the expected magnitude. The contradictory and sometimes unexpected results related to monitored BC in particular suggest that we may have had limited statistical power for some of these analyses. On the other hand, strengths of this study include use of spatial-temporal models to estimate PM$_{2.5}$ and BC exposure at each address, adjustment for several potentially important confounders, multiple imputation to reduce the risk of selection bias from missing data, and investigation
within the context of births from a single hospital that delivers the great majority of babies in this state.

In conclusion, we did not find evidence of a positive association between PM$_{2.5}$ and risk of preterm birth, but did find the expected association between PM$_{2.5}$ and lower birthweight. These findings suggest that the etiologic relationship between PM$_{2.5}$ and risk of preterm birth remains incompletely understood and that further research is needed before attributing a large proportion of global preterm births to PM$_{2.5}$. Our results also add further support to the mounting evidence that PM$_{2.5}$ is associated with reduced fetal growth and extend previous findings to a new location.

**Supplementary Material**

Refer to Web version on PubMed Central for supplementary material.

**Acknowledgments**

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**References**


23. Hospital WL. About Us. 2017


J Epidemiol Community Health. Author manuscript; available in PMC 2018 April 02.


36. EPA. National Ambient Air Quality Standards for Particulate Matter. EPA. 2013


Summary Box

What is already known on this subject?

- Maternal exposure to ambient air pollution during fetal development has been associated with reduced fetal growth, with some studies also suggesting an increased risk of preterm birth.
- Heterogeneity among prior studies suggests the need to evaluate these relationships in diverse populations and settings.

What this study adds?

- We used spatial-temporal models to estimate PM$_{2.5}$ and black carbon (a marker of traffic-related pollution) levels at each residential address to capture both spatial and temporal variation in exposure.
- In this population with relatively low exposures to ambient air pollutants, PM$_{2.5}$ was not associated with risk of preterm birth, but was associated with lower birthweight.
- Further research is needed before attributing a large proportion of global preterm births to PM$_{2.5}$.
### Table 1

Characteristics of study population

<table>
<thead>
<tr>
<th>Characteristics</th>
<th>All births (n= 61,640)</th>
<th>Term births (n=56,635)</th>
<th>Preterm births (n=5,007)</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Maternal Characteristics</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age (years)</td>
<td>29 ± 5.9</td>
<td>29 ± 5.8</td>
<td>29 ± 6.2</td>
</tr>
<tr>
<td>Parity</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
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<td>25805 (42.9)</td>
<td>23603 (42.7)</td>
<td>2202 (45.3)</td>
</tr>
<tr>
<td>1</td>
<td>20209 (33.6)</td>
<td>18797 (34.0)</td>
<td>1412 (29.1)</td>
</tr>
<tr>
<td>≥2</td>
<td>14142 (23.5)</td>
<td>12898 (23.3)</td>
<td>1244 (25.6)</td>
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<tr>
<td>Race</td>
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<td>33657 (64.9)</td>
<td>2853 (62.2)</td>
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<td>Black</td>
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<td>4228 (8.2)</td>
<td>478 (10.4)</td>
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<tr>
<td>Other</td>
<td>15226 (27.0)</td>
<td>13967 (26.9)</td>
<td>1259 (27.4)</td>
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<td>Education</td>
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<td></td>
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<td>Less than high school</td>
<td>7609 (12.8)</td>
<td>6861 (12.6)</td>
<td>748 (15.8)</td>
</tr>
<tr>
<td>High School</td>
<td>15932 (26.8)</td>
<td>14609 (26.7)</td>
<td>1323 (27.9)</td>
</tr>
<tr>
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<td>10149 (18.6)</td>
<td>878 (18.5)</td>
</tr>
<tr>
<td>College</td>
<td>14785 (24.9)</td>
<td>13695 (25.1)</td>
<td>1090 (23.0)</td>
</tr>
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<td>Any Graduate School</td>
<td>10010 (16.7)</td>
<td>9303 (17.0)</td>
<td>707 (14.9)</td>
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<tr>
<td>Marital status</td>
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<tr>
<td>Married</td>
<td>38347 (62.6)</td>
<td>35440 (63.0)</td>
<td>2907 (58.3)</td>
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<td>21139 (34.5)</td>
<td>19249 (34.2)</td>
<td>1890 (37.9)</td>
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<td>Other</td>
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<td>1607 (2.9)</td>
<td>186 (3.7)</td>
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<td>Health Insurance</td>
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<tr>
<td>Private</td>
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<td>31628 (56.2)</td>
<td>2592 (52.1)</td>
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<td>24843 (40.6)</td>
<td>22668 (40.3)</td>
<td>2175 (43.7)</td>
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<td>Other</td>
<td>2185 (3.6)</td>
<td>1973 (3.5)</td>
<td>212 (4.2)</td>
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<td>Tobacco use in pregnancy</td>
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<td></td>
</tr>
<tr>
<td>No</td>
<td>49678 (93.4)</td>
<td>45837 (93.7)</td>
<td>3841 (89.9)</td>
</tr>
<tr>
<td>Yes</td>
<td>3529 (6.6)</td>
<td>3097 (6.3)</td>
<td>432 (10.1)</td>
</tr>
<tr>
<td>Neighborhood Socioeconomic Status, sum of z-score</td>
<td>−0.2 ± 4.4</td>
<td>−0.2 ± 4.4</td>
<td>−0.5 ± 4.3</td>
</tr>
<tr>
<td>Modeled ambient air pollution</td>
<td>PM$_{2.5}$ (µg/m$^3$)</td>
<td>9.5 ± 1.5</td>
<td>9.5 ± 1.5</td>
</tr>
<tr>
<td>BC (µg/m$^3$)</td>
<td>0.52 ± 0.1</td>
<td>0.52 ± 0.1</td>
<td>0.52 ± 0.1</td>
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<tr>
<td>Infant Characteristics</td>
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<td></td>
</tr>
<tr>
<td>Gestational Age (weeks)</td>
<td>39 ± 2.0</td>
<td>39 ± 1.1</td>
<td>34 ± 2.9</td>
</tr>
<tr>
<td>Preterm birth</td>
<td>5007 (8.1)</td>
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<td></td>
</tr>
<tr>
<td>Birthweight (g)</td>
<td>3341 ± 563.2</td>
<td>3426 ± 459.3</td>
<td>2380 ± 716.0</td>
</tr>
<tr>
<td>Characteristics</td>
<td>All births (n= 61,640)</td>
<td>Term births (n=56,633)</td>
<td>Preterm births (n=5,007)</td>
</tr>
<tr>
<td>---------------------------------</td>
<td>------------------------</td>
<td>------------------------</td>
<td>--------------------------</td>
</tr>
<tr>
<td>Low birthweight (&lt;2500 g)</td>
<td>3758 (6.1)</td>
<td>1145 (2.0)</td>
<td>2613 (52.2)</td>
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<tr>
<td>Small for gestational age</td>
<td>5442 (8.8)</td>
<td>4810 (8.5)</td>
<td>632 (12.6)</td>
</tr>
</tbody>
</table>

\(^a\) Frequencies may not sum to full sample size due to missing data

\(^b\) Includes American Indian, Filipino, Hispanic, Other Asian, and Other (not specified)

\(^c\) Includes divorced, separated, widowed, and unknown

\(^d\) Private: Blue Cross/Blue Shield, Commercial insurance, and HMO coverage; Public: Medicaid or Medicaid/HMO; Other: Medicare, Champus, self-coverage, other coverage, and coverage unknown
Table 2

Odds ratio (95% confidence intervals) for preterm birth per interquartile range (IQR)\(^a\) increase in PM\(_{2.5}\) and black carbon, overall and by trimester among 61,640 deliveries

<table>
<thead>
<tr>
<th></th>
<th>Entire Pregnancy</th>
<th>First Trimester</th>
<th>Second Trimester</th>
<th>Third Trimester</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>PM(_{2.5})</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Modeled</td>
<td>1.04 (0.94, 1.15)</td>
<td>0.93 (0.88, 0.98)(^e)</td>
<td>1.03 (0.98, 1.09)</td>
<td>1.02 (0.97, 1.07)</td>
</tr>
<tr>
<td>Monitored</td>
<td>0.86 (0.76, 0.98)(^e)</td>
<td>0.90 (0.85, 0.94)(^e)</td>
<td>1.02 (0.97, 1.07)</td>
<td>1.00 (0.96, 1.05)</td>
</tr>
<tr>
<td><strong>Black Carbon</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Modeled</td>
<td>1.00 (0.97, 1.05)</td>
<td>0.99 (0.96, 1.03)</td>
<td>0.99 (0.95, 1.02)</td>
<td>0.99 (0.95, 1.02)</td>
</tr>
<tr>
<td>Monitored</td>
<td>1.21 (1.05, 1.39)(^e)</td>
<td>1.03 (0.97, 1.09)</td>
<td>1.04 (0.99, 1.09)</td>
<td>0.95 (0.90, 0.99)(^e)</td>
</tr>
</tbody>
</table>

All models are adjusted for maternal age, tobacco, parity, education, race, health insurance, marital status, last menstrual period, and neighborhood socioeconomic status. We used multiple imputation to account for covariates with missing data: tobacco (13.7% missing), parity (2.4% missing), maternal education (3.7% missing), maternal race (8.4% missing), maternal insurance (0.6% missing), and marital status (1.0%).

\(^a\)IQR is 2.5 µg/m\(^3\) for PM\(_{2.5}\) and 0.11 µg/m\(^3\) for black carbon.

\(^e\)p<0.05
Table 3
Gram change in birthweight (95% CI) per interquartile range (IQR)\(^a\) increase in PM\(_{2.5}\) and black carbon among 56,633 term births, overall and by trimester of pregnancy

<table>
<thead>
<tr>
<th></th>
<th>Entire Pregnancy</th>
<th>First Trimester</th>
<th>Second Trimester</th>
<th>Third Trimester</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>PM(_{2.5})</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Modeled</td>
<td>−12.1 (−24.2, −0.1) (^*)</td>
<td>−4.2 (−10.6, 2.3)</td>
<td>−0.3 (−6.8, 6.2)</td>
<td>−5.3 (−11.0, 0.4)</td>
</tr>
<tr>
<td>Monitored</td>
<td>−15.9 (−31.6, −0.3) (^*)</td>
<td>−3.0 (−9.1, 3.1)</td>
<td>−0.14 (−6.4, 6.1)</td>
<td>−4.3 (−9.6, 1.1)</td>
</tr>
<tr>
<td><strong>Black Carbon</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Modeled</td>
<td>0.0 (−4.7, 4.7)</td>
<td>1.3 (−2.9, 5.5)</td>
<td>1.2 (−3.0, 5.4)</td>
<td>−3.6 (−7.8, 0.7)</td>
</tr>
<tr>
<td>Monitored</td>
<td>5.7 (−10.9, 22.3)</td>
<td>−1.4 (−8.1, 5.4)</td>
<td>9.2 (3.3, 15.0) (^*)</td>
<td>−6.6 (−12.0, −1.2) (^*)</td>
</tr>
</tbody>
</table>

All models adjusted for maternal age, gestational age, tobacco, parity, maternal education, maternal race, maternal insurance, marital status, last menstrual period, and neighborhood socioeconomic status. We used multiple imputation to account for covariates with missing data: tobacco (13.7% missing), parity (2.4% missing), maternal education (3.7% missing), maternal race (8.4% missing), maternal insurance (0.6% missing), and marital status (1.0%).

\(^a\)IQR is 2.5 µg/m\(^3\) for PM\(_{2.5}\) and 0.11 µg/m\(^3\) for black carbon.

\(^*\) p<0.05
### Table 4
Odds ratios and 95% CIs for small for gestational age (SGA) and low birthweight (LBW) per interquartile range (IQR) increase in PM$_{2.5}$ and black carbon among 56,633 term births, overall and by trimester of pregnancy

<table>
<thead>
<tr>
<th></th>
<th>Entire Pregnancy</th>
<th>First Trimester</th>
<th>Second Trimester</th>
<th>Third Trimester</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Small for Gestational Age</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>PM$_{2.5}$</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Modeled</td>
<td>1.09 (0.98, 1.21)</td>
<td>1.00 (0.95, 1.06)</td>
<td>1.04 (0.99, 1.1)</td>
<td>1.02 (0.97, 1.08)</td>
</tr>
<tr>
<td>Monitored</td>
<td>1.15 (1.00, 1.31)*</td>
<td>1.00 (0.95, 1.06)</td>
<td>1.05 (0.99, 1.11)</td>
<td>1.01 (0.97, 1.06)</td>
</tr>
<tr>
<td>Black Carbon</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Modeled</td>
<td>1.04 (0.99, 1.08)</td>
<td>1.03 (0.99, 1.06)</td>
<td>1.03 (0.99, 1.07)</td>
<td>1.03 (0.99, 1.07)</td>
</tr>
<tr>
<td>Monitored</td>
<td>0.96 (0.83, 1.11)</td>
<td>0.99 (0.93, 1.05)</td>
<td>0.97 (0.92, 1.02)</td>
<td>1.03 (0.98, 1.08)</td>
</tr>
<tr>
<td><strong>Low Birthweight</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>PM$_{2.5}$</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Modeled</td>
<td>1.05 (0.84, 1.29)</td>
<td>0.96 (0.86, 1.08)</td>
<td>1.03 (0.92, 1.16)</td>
<td>1.05 (0.96, 1.16)</td>
</tr>
<tr>
<td>Monitored</td>
<td>0.92 (0.71, 1.19)</td>
<td>0.94 (0.84, 1.04)</td>
<td>1.01 (0.91, 1.13)</td>
<td>1.03 (0.95, 1.13)</td>
</tr>
<tr>
<td>Black Carbon</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Modeled</td>
<td>1.01 (0.93, 1.10)</td>
<td>1.01 (0.94, 1.09)</td>
<td>1.02 (0.95, 1.09)</td>
<td>1.01 (0.94, 1.09)</td>
</tr>
<tr>
<td>Monitored</td>
<td>0.81 (0.60, 1.08)</td>
<td>0.83 (0.91, 1.17)</td>
<td>0.88 (0.79, 0.98)*</td>
<td>1.03 (0.94, 1.13)</td>
</tr>
</tbody>
</table>

All models adjusted for maternal age, gestational age, tobacco, parity, maternal education, maternal race, maternal insurance, marital status, last menstrual period, and neighborhood socioeconomic status. Models of LBW are additionally adjusted for gestational age. We used multiple imputation to account for covariates with missing data: tobacco (13.7% missing), parity (2.4% missing), maternal education (3.7% missing), maternal race (8.4% missing), maternal insurance (0.6% missing), and marital status (1.0%).

* IQR is 2.5 µg/m$^3$ for PM$_{2.5}$ and 0.11 µg/m$^3$ for black carbon.

* p<0.05