A weight-gain-for-gestational-age z score chart for the assessment of maternal weight gain in pregnancy1–3

Jennifer A Hutcheon, Robert W Platt, Barbara Abrams, Katherine P Himes, Hyagriv N Simhan, and Lisa M Bodnar

ABSTRACT

Background: To establish the unbiased relation between maternal weight gain in pregnancy and perinatal health, a classification for maternal weight gain is needed that is uncorrelated with gestational age.

Objective: The goal of this study was to create a weight-gain-for-gestational-age percentile and z score chart to describe the mean, SD, and selected percentiles of maternal weight gain throughout pregnancy in a contemporary cohort of US women.

Design: The study population was drawn from normal-weight women with uncomplicated, singleton pregnancies who delivered at the Magee-Womens Hospital in Pittsburgh, PA, 1998–2008. Analyses were based on a randomly selected subset of 648 women for whom serial prenatal weight measurements were available through medical chart record abstraction (6727 weight measurements).

Results: The pattern of maternal weight gain throughout gestation was estimated using a random-effects regression model. The estimates were used to create a chart with the smoothed means, percentiles, and SDs of gestational weight gain for each week of pregnancy.

Conclusion: This chart allows researchers to express total weight gain as an age-standardized z score, which can be used in epidemiologic analyses to study the association between pregnancy weight gain and adverse or physiologic pregnancy outcomes independent of gestational age. Am J Clin Nutr 2013;97:1062–7.

INTRODUCTION

Institute of Medicine guidelines for maternal weight gain in pregnancy attempt to balance the risks associated with excess gestational weight gain (such as unplanned cesarean delivery and fetal macrosomia) with the risks associated with inadequate gestational weight gain (including fetal growth restriction and preterm birth) (1). Establishing the optimal range of gestational weight gain, therefore, requires a solid evidence base on the association between total weight gain in pregnancy and a broad range of short- and longer-term maternal and child health outcomes.

Although several studies have examined the relation between gestational weight gain and pregnancy outcomes (1), recent work by our group suggests that the approaches used to classify total gestational weight gain in these studies may have serious flaws (2). It has long been recognized that linking total gestational weight with outcomes such as preterm birth or perinatal mortality is problematic because women who deliver at younger gestational ages have less time to gain weight than do women with longer pregnancy durations, which thus creates a spurious association between low gestational weight gain and preterm birth or adverse perinatal outcome (3). Investigators have attempted to overcome this potential bias by expressing total weight gain as an average rate of weight gain or as a ratio of the total weight gain recommended by Institute of Medicine guidelines (4, 5), but our work has shown that these classification methods are still problematic. We showed that the gestational weight gain measures remained correlated with gestational duration, and, because many adverse pregnancy outcomes are also strongly associated with gestational duration, their use introduces a nontrivial degree of bias to our understanding of the relation between weight gain and pregnancy outcomes (2).

To accurately understand the relation between gestational weight gain and maternal and child health, a method for classifying weight gain that is independent of gestational duration is needed. In the study of fetal, infant, and pediatric growth, weight-for-age percentiles or z scores are well-established tools for classifying growth during or after pregnancy (6–8). Weight percentiles or z scores provide an age-independent assessment of a fetus or child’s growth while accounting for the nonlinear shape of growth trajectories throughout gestation and childhood. Furthermore, percentiles and z scores may be preferable to ratios

1 From the Department of Obstetrics & Gynaecology, University of British Columbia, Vancouver, Canada (JAH); the Departments of Pediatrics and Epidemiology & Biostatistics, McGill University, Montreal, Canada (RWP); the Division of Epidemiology, School of Public Health, University of California, Berkeley, Berkeley, CA (BA); the Department of Obstetrics, Gynecology, and Reproductive Sciences, School of Medicine, University of Pittsburgh, Pittsburgh, PA (KPH, HNS, and LMB); and the Department of Epidemiology, Graduate School of Public Health, University of Pittsburgh, Pittsburgh, PA (LMB).

2 Supported by the Eunice Kennedy Shriver National Institute of Child Health and Human Development (R21 HD067851-01 to LMB and HS). JAH holds a Canadian Institutes of Health Research New Investigator Award and is a Career Scholar of the Michael Smith Foundation for Health Research. RWP holds a Chercheur-National from the Fonds de la Recherche en Santé du Quebec (FRSQ). The Montreal Children’s Hospital Research Institute receives core operating funds from the FRSQ.

3 Address correspondence and reprint requests to JA Hutcheon, Department of Obstetrics & Gynaecology, University of British Columbia, Shaughnessy Room E421A, BC Women’s Hospital & Health Centre, 4500 Oak Street, Vancouver, BC, Canada V6N 1H3. E-mail: jhutcheon@cfri.ca.

Received September 24, 2012. Accepted for publication January 16, 2013. First published online March 6, 2013; doi: 10.3945/ajcn.112.051706.
for use in epidemiologic analyses because their statistical characteristics make them less prone to bias and difficulty in interpretation due to nonlinearity (9). This well-accepted method for the assessment of offspring growth has the potential to be applied to the assessment of maternal weight gain to overcome the limitations of current classification methods. The objective of this study was to create a maternal weight-gain-for-gestational-age percentile and z score chart, describing the mean, SD, and selected percentiles of weight gain for each week of gestation in a cohort of normal-weight women in the United States with uncomplicated term pregnancies.

SUBJECTS AND METHODS

The study population was drawn from women who delivered a singleton birth at Magee-Womens Hospital in Pittsburgh, PA, from 1998 to 2008. Detailed demographic, medical, obstetrical, and neonatal data on all deliveries at Magee-Womens Hospital is contained in the electronic Magee Obstetrics, Medical, and Infant (MOMI) Database. As part of an ongoing case-cohort study investigating the relation between obesity, gestational weight gain, and stillbirth (NIH R21 HD067851), we used the MOMI Database as a sampling frame to perform a simple random sample of 812 normal-weight [prepregnancy BMI (in kg/m²) 18.5–24.9], 316 overweight (BMI 25–29.9), and 212 obese (BMI ≥30) women representative of the Magee-Womens Hospital obstetrical population. In the current study, we restricted our final population to women with ≥5 antenatal weight measurements who delivered live-born infants at term gestational ages (37+0 to 41+6 wk) and excluded pregnancies complicated by preexisting diabetes or hypertension, gestational diabetes, or preeclampsia. We excluded women with <5 prenatal visits because we hypothesized that this reflected suboptimal prenatal care, with an increased potential for missed diagnoses of hypertensive disorders or gestational diabetes, errors in recall of prepregnancy BMI, and errors in gestational age dating because of a lack of early ultrasonography. We focused on normal-weight women and provided provisional data for overweight and obese women. On the basis of preliminary data on the mean and SD of total gestational weight gain in pregnancies of normal-weight women documented in the MOMI (15.7 ± 5.7 kg), we estimated that a sample size of 598 women would be required to allow us to estimate the 16th and 84th percentiles of gestational weight gain (ie, 1 SD from the mean) to within 0.5 kg. Ethical approval was obtained from the University of Pittsburgh Institutional Review Board.

For each of the sampled women, serial maternal weight measurements obtained during antenatal clinic visits were abstracted from the medical records by a trained abstraction using an Access 2007 database (Microsoft Corporation) with programmed abstraction forms. The programmed forms help to eliminate data entry errors by disallowing implausible values, flagging questionable entries, and ensuring completion of all data elements. Self-reported prepregnancy weight and maternal height documented at the time of the first antenatal visit were also abstracted into the programmed forms. Gestational weight gain (kg) was calculated as the measured weight at the time of an antenatal visit or delivery admission less the self-reported prepregnancy weight ascertained at the first antenatal visit. The trajectories of women with extreme weight gains or weight losses [defined as a weight gain of ≥36.4 kg (80 lb) or weight loss of ≥−18.2 kg (40 lb)] were assessed for implausible values in either prepregnancy weight or weight at a specific antenatal visit. Gestational age was estimated by using the algorithm endorsed by the American College of Obstetricians and Gynecologists for clinical practice based on ultrasonography and estimates of the last menstrual period (10).

We built a multilevel (random-effects) model that described the longitudinal maternal weight gain measurements as a function of gestational age. We modeled gestational age using a restricted cubic spline to allow the weight gain curves to vary across gestation in a smooth manner (11). Random effects were specified for the model intercept and the linear spline term for gestational age, with an unstructured covariance structure. Formulas for the gestational age–specific means and SDs of weight gain are provided in Appendix A. Weight gain measurements were log transformed (natural log) to ensure that the model’s assumptions of homoscedasticity of residual errors and normality of distribution of the response variable were not violated. A constant of 9.18 was added to all values to shift the minimum value of the distribution to 1 (ie, to ensure nonnegative observations) (12). Because exponentiating the predicted mean of a log-transformed response variable will produce an unbiased estimate of the median and geometric mean, but not the arithmetic mean (13, 14), we presented the medians and selected percentiles back-transformed to the original scale (kg) and the means and SDs on the log scale for use in calculating z scores. This means that to calculate a given woman’s weight gain z score, her observed total weight gain must first be log transformed before being substituted into the formula $z = (\text{observed weight gain} - \text{mean})/\text{SD}$. We additionally built models for gestational weight gain in overweight and in obese women; however, because of the relatively small number of women in these BMI categories, we presented these results as provisional values only.

As recommended (15), we assessed the goodness-of-fit of our estimated means and SDs of gestational weight gain using a scatter diagram that superimposed the fitted means and SDs on the raw data. We further checked that the appropriate proportion of observations fell below and above the estimated thresholds for ±1 SD and confirmed that the z scores followed a normal distribution. The analyses were performed by using STATA SE version 11 (StataCorp). STATA code to produce the percentiles and z scores is available from the corresponding author on request.

RESULTS

The exclusion of 84 preterm deliveries, 2 stillbirths, 39 women with comorbidities, and 39 women with <5 antenatal weight measurements left a total of 648 normal-weight women. The 39 (5.7%) women excluded because they had <5 prenatal visits were less likely to be white (71% compared with 85%), were more likely to receive Medicaid (60.5% compared with 24.2%), were younger (27.4 y compared with 29.8 y), were more likely to be parous (81.6% compared with 54%), and had their first prenatal visit at a later gestational age (28.5 wk compared with 10.4 wk). One woman had a weight gain at 35 wk of −43 kg; this observation was excluded as implausible given her weight gain trajectory (which otherwise increased steadily from 0 kg at 8 wk to 21 kg at 39 wk), and her remaining 13 weight gain observations were used for analysis. There were a total of 6727 weight gain measurements (average of 10.4 per woman). The median gestational age at the first antenatal visit was 9.2 wk
The mean white, 11% non-Hispanic black, and 4% of other race-
ethnicity. The study population was 85% non-Hispanic
receiving Medicaid assistance.

The crude weight gain trajectories of 50 randomly selected
normal-weight women are shown in Figure 1. As expected, the
rate of weight gain in the first trimester was considerably lower
than that in the second and third trimesters.

We compared models with gestational age modeled as a linear
term, a restricted cubic spline with 4 knots, and a restricted cubic
spline with 5 knots. Whereas a linear term for gestational age
resulted in a model with a more parsimonious fit [ie, with a lower
Akaike Information Criterion (AIC)], we opted to use the spline
for gestational age because examination of residuals showed a
better fit at the extremes of gestational age (particular for the
first trimester). A restricted cubic spline with 4 knots was selected
because it resulted in a model with a lower AIC than did that with
5 knots. Our final model therefore included gestational age as a
restricted cubic spline with 4 knots at 13, 23, 33, and 39 wk. Knot
locations were based on default positions, with the ex-
ception of the knot at 13 wk. This knot was shifted from its initial
default position of 9 wk because of available evidence suggesting
that the end of the first trimester is a biologically meaningful
threshold (1) and was supported by our finding that shifting the
knot to 13 wk resulted in a model with a lower AIC. Our final
model was estimated as follows: E[ln(GWG)] = 2.04 + 0.032
(GA0) + 0.004(GA1) − 0.035(GA2), where GWG is gestational
weight gain and the 3 coefficients for gestational age (GA) are
the cubic spline regression coefficients (ie, the linear term, first
cubic term, and second cubic term), and random-effects pa-
rameters of var βw = 0.00005, var βGA1 = 0.082, and cov (βw,
βGA1) = −0.0015. The smoothed means and SDs of maternal
weight gain, estimated by the final model superimposed on the
crude observations, are shown in Figure 2.

Seventy-two percent of the weight measurements fell within
the smoothed 1-SD threshold, which was more than the 68%
expected. However, the absolute magnitudes of the differences
between the observed and smoothed 1-SD threshold values
within each week were modest (<1 kg). The values of the
smoothed week-specific means, SDs, and selected percentiles
are provided in Table 1, and shown graphically in Figure 3.

TABLE 1
Week-specific means, SDs, and selected percentiles of GWG among
normal-weight women who delivered at Magee-Womens Hospital,
Pittsburgh, PA, 1998–2008, estimated by using a multilevel linear
regression model.

<table>
<thead>
<tr>
<th>Gestational age</th>
<th>Mean ln(GWG)</th>
<th>SD ln(GWG)</th>
</tr>
</thead>
<tbody>
<tr>
<td>6 wk</td>
<td>2.47</td>
<td>0.26</td>
</tr>
<tr>
<td>7 wk</td>
<td>2.40</td>
<td>0.29</td>
</tr>
<tr>
<td>8 wk</td>
<td>2.31</td>
<td>0.25</td>
</tr>
<tr>
<td>9 wk</td>
<td>2.34</td>
<td>0.24</td>
</tr>
<tr>
<td>10 wk</td>
<td>2.37</td>
<td>0.24</td>
</tr>
<tr>
<td>11 wk</td>
<td>2.40</td>
<td>0.23</td>
</tr>
<tr>
<td>12 wk</td>
<td>2.44</td>
<td>0.23</td>
</tr>
<tr>
<td>13 wk</td>
<td>2.47</td>
<td>0.23</td>
</tr>
<tr>
<td>14 wk</td>
<td>2.50</td>
<td>0.23</td>
</tr>
<tr>
<td>15 wk</td>
<td>2.53</td>
<td>0.23</td>
</tr>
<tr>
<td>16 wk</td>
<td>2.55</td>
<td>0.22</td>
</tr>
<tr>
<td>17 wk</td>
<td>2.60</td>
<td>0.22</td>
</tr>
<tr>
<td>18 wk</td>
<td>2.63</td>
<td>0.22</td>
</tr>
<tr>
<td>19 wk</td>
<td>2.66</td>
<td>0.21</td>
</tr>
<tr>
<td>20 wk</td>
<td>2.70</td>
<td>0.21</td>
</tr>
<tr>
<td>21 wk</td>
<td>2.73</td>
<td>0.21</td>
</tr>
<tr>
<td>22 wk</td>
<td>2.76</td>
<td>0.21</td>
</tr>
<tr>
<td>23 wk</td>
<td>2.80</td>
<td>0.21</td>
</tr>
<tr>
<td>24 wk</td>
<td>2.83</td>
<td>0.21</td>
</tr>
<tr>
<td>25 wk</td>
<td>2.87</td>
<td>0.20</td>
</tr>
<tr>
<td>26 wk</td>
<td>2.90</td>
<td>0.20</td>
</tr>
<tr>
<td>27 wk</td>
<td>2.94</td>
<td>0.20</td>
</tr>
<tr>
<td>28 wk</td>
<td>2.97</td>
<td>0.20</td>
</tr>
<tr>
<td>29 wk</td>
<td>3.00</td>
<td>0.20</td>
</tr>
<tr>
<td>30 wk</td>
<td>3.03</td>
<td>0.20</td>
</tr>
<tr>
<td>31 wk</td>
<td>3.06</td>
<td>0.21</td>
</tr>
<tr>
<td>32 wk</td>
<td>3.09</td>
<td>0.21</td>
</tr>
<tr>
<td>33 wk</td>
<td>3.13</td>
<td>0.21</td>
</tr>
<tr>
<td>34 wk</td>
<td>3.15</td>
<td>0.21</td>
</tr>
<tr>
<td>35 wk</td>
<td>3.16</td>
<td>0.21</td>
</tr>
<tr>
<td>36 wk</td>
<td>3.17</td>
<td>0.21</td>
</tr>
<tr>
<td>37 wk</td>
<td>3.19</td>
<td>0.20</td>
</tr>
<tr>
<td>38 wk</td>
<td>3.20</td>
<td>0.20</td>
</tr>
<tr>
<td>39 wk</td>
<td>3.22</td>
<td>0.22</td>
</tr>
<tr>
<td>40 wk</td>
<td>3.22</td>
<td>0.22</td>
</tr>
</tbody>
</table>

Results are based on 648 women with 6727 weight-gain observations
(average of 10.4 per woman). GWG, gestational weight gain.
Estimates are presented for gestational ages (wk) at which a minimum of 30 weight observations were available (6–40 wk). Note that because the means and SDs are presented on the log scale, the women’s gestational weight gains had to be converted to the log scale for the purpose of calculating $z$ scores. For example, the $z$ score for a woman delivering at 35 wk with a total weight gain of 18 kg would be calculated as follows: $[\ln(18 + c) - 3.156)/0.216$, where $c$ is the constant of 9.18, and 3.156 and 0.216 are the log-transformed mean and SD of weight gain at 35 wk, respectively, in normal-weight women. Provisional values for overweight and obese women, estimated by using methods similar to those used for normal-weight women, are provided elsewhere (see Supplemental Tables S1 and S2 under “Supplemental data” in the online issue).

**DISCUSSION**

In this study, we created smoothed reference values for maternal weight gain in pregnancy from a contemporary cohort of US women with healthy pregnancy outcomes. Our weight gain chart allows a woman’s weight gain for gestational age to be expressed as a standardized $z$ score that is independent of pregnancy duration, which can in turn be used in epidemiologic studies to obtain an unbiased estimate of the association between total weight gain and maternal and offspring health.

Although the chart should be useful for epidemiologic analyses, it is not yet appropriate for use in the clinical setting because it remains to be established which $z$ score range is associated with optimal short- and long-term maternal and offspring health. Rather than arbitrarily selecting certain statistical values as the thresholds for high or low risk (eg, the 10th percentile or 2 SD), risk thresholds should instead be established following future work linking maternal weight $z$ scores with a broad range of short- and longer-term maternal and offspring health outcomes, including infant mortality, preterm delivery, postpartum weight retention, and childhood obesity. Indeed, because many women with uncomplicated pregnancies gain more weight than is recommended by the Institute of Medicine (as seen in our cohort and in other contemporary US cohorts) (1, 16), we hypothesized that the $z$ score range that optimizes maternal and child health (in particular, prevention of long-term maternal weight retention and childhood obesity) may be shifted to $z$ score values systematically $<0$ (ie, lower than the population average observed in contemporary cohorts).

Our maternal weight gain chart was created by using a random sample of women who delivered at Magee-Womens Hospital in Pittsburgh, PA, and reflects a mix of non-Hispanic white and black women from diverse socioeconomic backgrounds. Because serial prenatal weight measurements are rarely available in large population or clinical databases (only total weight gain is typically available), our abstracted serial prenatal weight data provided a unique opportunity to model patterns of weight gain in a large contemporary US cohort with uncomplicated term pregnancies. As shown in Table 2, the total gestational weight gains at this institution are very similar to those observed in several other large, contemporary American cohorts (1, 16), which suggests that the reference values we presented are likely suitable for use in many general obstetrical populations in the United States. The Magee-Womens Hospital population has also previously been used in many studies that examined maternal weight gain and pregnancy outcomes (4, 17, 18), which further supports the generalizability of our results. The 10th, 50th, and 90th percentiles for first-trimester weight in our cohort (−0.4, 2.7, and 6.9 kg at 13 wk, respectively) were higher than those previously reported for a cohort of normal-weight women with good pregnancy outcomes (−2.21, 2.20, and 6.59 kg, respectively) (19), although this may be explained by temporal

**TABLE 2**

Comparison of total GWG among women who delivered at Magee-Womens Hospital, Pittsburgh, PA, 1998–2008, with weight gain of women in other contemporary US cohorts

<table>
<thead>
<tr>
<th>Prepregnancy BMI group</th>
<th>Magee-Womens Hospital</th>
<th>PRAMS$^2$</th>
<th>PIN$^3$</th>
<th>Project Viva$^4$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Underweight, &lt;18.5 kg/m$^2$</td>
<td>15.7 ± 4.6$^5$</td>
<td>14.8</td>
<td>15.4 ± 4.4</td>
<td>15.7 ± 4.5</td>
</tr>
<tr>
<td>Normal weight, 18.5–24.9 kg/m$^2$</td>
<td>15.5 ± 5.3</td>
<td>15.0</td>
<td>16.6 ± 5.3</td>
<td>16.4 ± 4.8</td>
</tr>
<tr>
<td>Overweight, 25–29.9 kg/m$^2$</td>
<td>15.1 ± 6.2</td>
<td>13.9</td>
<td>15.5 ± 6.2</td>
<td>15.3 ± 5.3</td>
</tr>
<tr>
<td>Obese, ≥30 kg/m$^2$</td>
<td>9.9 ± 6.5</td>
<td>11.2</td>
<td>12.0 ± 7.1</td>
<td>13.1 ± 7.1</td>
</tr>
</tbody>
</table>

$^1$Data derived from references 1 and 16. GWG, gestational weight gain.


$^4$Project Viva (1999–2002, Massachusetts). Note that underweight was defined as $<19.8$ kg/m$^2$, normal weight as 19.8–26 kg/m$^2$, overweight as 26.1–29.0 kg/m$^2$, and obese as $>29.0$ kg/m$^2$.

$^5$Mean ± SD (all such values).
increases in gestational weight gain among American women, with the latter study reflecting the weight gain of women who delivered between 1980 and 1990. Whereas our results should be generalizable to women in a large number of obstetrical care settings in the United States, we hope that our work will pave the way for the future creation of gestational weight gain charts based on the serial antenatal weight gain measurements of a nationally representative sample of women.

Although our weight-gain-for-gestational-age charts are novel to the study of maternal weight gain in pregnancy, they are based on methods that are well established in the study of fetal and infant growth. Since the publication of the Lubchenco birth weight-for-gestational-age charts in 1963 (20), birth weight charts have been published for a broad range of infant populations and subgroups (6, 21–23). Reference charts of longitudinal estimated fetal weight measurements have also been created by using statistical multilevel modeling techniques similar to those used in this study to appropriately account for the repeated weight measurements from each individual (8). By requiring a minimum of 5 weights per woman, we helped to ensure that our chart reflected longitudinal trajectories of weight gain rather than cross-sectional weights of different women at different gestational ages. Birth weight and estimated fetal weight percentiles and z scores have been used extensively to improve our understanding of the determinants and consequences of abnormal fetal growth (24–26); we believe that applying this methodology to the study of maternal weight gain will be equally valuable.

Several limitations of our research should be noted. Gestational weight gain measurements were based on self-reported prepregnancy weights. It is well documented that women of childbearing age systematically underestimate their weight on self-report (27), so the mean gestational age–specific weight gains presented herein may overestimate true pregnancy weight gain (ie, the observed weight gain reflects the true total weight gain plus the amount of underreporting in prepregnancy BMI). Nevertheless, because measured prepregnancy weight is rarely available in routine antenatal care, our reference values are a pragmatic tool that reflects the information available to obstetrical care providers in practice. We also did not have a sufficient number of women to produce a definite reference chart for underweight, overweight, or obese women. Although we produced provisional values for overweight and obese women, future work to establish more definitive reference values would be valuable.

The authors’ responsibilities were as follows—IAH, RWP, BA, and LMB: designed the research; LMB, KPH, and HNS: collected the data and provided essential materials; and JAH: performed the statistical analysis, wrote the manuscript, and had primary responsibility for the final content. All authors read and approved the final manuscript. None of the authors declared a conflict of interest.

REFERENCES
APPENDIX A

Regression formulas

We modeled serial antenatal weight gain measurements using the random-effects model:

\[ GWG_{ij} = \beta_{0i} + \beta_{GAi}X_{ij} + e_{ij} \] (A1)

where: \( i \) is the \( i \)th woman, \( j \) is the \( j \)th measurement occasion, GWG is the log-transformed response variable of gestational weight gain, \( e \) is the within-woman variability in weight-gain measurements.

- \( X \) is the independent variable of gestational age, modeled as a restricted cubic spline with random effect on the linear spline basis term.

- \( \beta_{0i} = \beta_0 + u_i \) is the random effect (latent variable) at the level of the woman to allow each woman to have her own intercept.

- \( \beta_{GA(i)} = \beta_{GA} + u_{GA(i)} \) is the random effect at the level of the woman, here, allowing each woman’s rate of gain to have its own slope (i.e., taking into account that women will gain weight at different rates during pregnancy).

\[ \text{var} \beta_{0i} = \sigma_{\beta_0}^2 \]
\[ \text{var} \beta_{GAi} = \sigma_{\beta_{GA}}^2 \]
\[ \text{cov}(\beta_{0i}, \beta_{GAi}) = \sigma_{\beta_0\beta_{GA}} \]

The variance of GWG at gestational age \( X \) was estimated as follows:

\[ \text{var}(GWG) = \sigma_{e}^2 + \sigma_{\beta_{GA}}^2X^2 + 2X\sigma_{\beta_0\beta_{GA}} = \sigma_{e}^2 \] (A2)
Erratum


In the first full sentence of the right column on page 1064, the values for the random-effects parameters are inverted. The corrected sentence in the Results section should read as follows: “Our final model was estimated as follows: E[\ln(GWG)] = 2.04 + 0.032(GA_{ij}) + 0.004(GA_{ij})^2 - 0.035(GA_{ij})^3, where GWG is gestational weight gain and the 3 coefficients for gestational age (GA) are the cubic spline regression coefficients (ie, the linear term, first cubic term, and second cubic term), and random-effects parameters of var(\beta_{GA}) = 0.00005, var(\beta_0) = 0.082, and cov(\beta_{GA}, \beta_0) = -0.0015.”

In equation A2 of Appendix A, the final equals sign should be a plus sign instead. The equation should read as follows:

\[ \text{var}(GWG) = \sigma_0^2 + \sigma_{GA}^2 X^2 + 2X\sigma_0\beta_{GA} + \sigma_e^2. \]