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published in
European Journal of Obstetrics and Gynecology and Reproductive Biology
1993

DOI (link to publisher)
10.1016/0028-2243(93)90175-C

document version
Publisher's PDF, also known as Version of record

European Journal of Obstetrics & Gynecology and Reproductive Biology, 50 (1993) 115–122

EUROBS 01613

Maternal characteristics and expected birth weight

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Accepted for publication 13 April 1993

Summary

Fetal growth charts currently used aggregate birth weights of infants with various natural histories from 1931 until 1967. In order to modernize these charts, avoiding deviation from the natural history of fetal development, we report data from infants born after spontaneous onset of labour in 'normal' pregnancy from a gestational age of 267 to 295 days between 1972 and 1982 (n = 14 113). The relationship between birth weight and gestational age in days was studied by multiple regression analysis, containing dummy variables for parity and gender. The estimated proportion of the variance in the model, attributed to these characteristics, was 15%. This could be improved to 22% by supplementing the model with maternal characteristics such as age, height, mid-pregnancy weight and ethnic origin. According to this extended model, in the Dutch section of the population 511 (4.6%) babies had a birth weight below the 5th percentile, whereas 412 (3.7%) babies would be labeled as such according to the conventional birth weight tables. Moreover, 93 babies would be wrongly considered too small, corresponding with a sensitivity of 62.4%, and 192 babies would be wrongly considered normal, corresponding with a specificity of 99.3%. Integration of the four currently used tables into one, and adjustment for easily available maternal characteristics, could substantially improve classification methods.

Fetal growth charts; Ethnic origin; Maternal height; Mid-pregnancy weight; Percentiles

Introduction

Infants who are too small for their gestational age are at increased risk to perinatal death and subsequently to neuro-developmental abnormalities [1]. Therefore they should preferably be nursed in specialized neonatal wards. Although identification of abnormal fetal growth is a major objective of antenatal care in communities allowing deliveries at home, 40.6% of the children below the 2.3rd percentile, usually called small for gestational age (SGA), were not detected by midwives in a specific geographical area (Wormerveer). Despite the high neonatal referral rate of 59% of these SGA infants to neonatologists, perinatal mortality and morbidity were high [2].

The Amsterdam birth weight tables, currently used in the Netherlands consist of four tables. They summarize birth weights from 1931 until 1967, dichotomized by parity (primiparous and

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multiparous) and gender of the child, and the gestational age is expressed in weeks [3].

Endeavours to modernize these tables are biased by several selection phenomena. Newborns represent cross-sectional data of birth weights on gestational age and these data do not describe the chronological course of events. This may imply that fetal growth does not follow these so-called 'intra-uterine growth curves'. Birth weights of infants after spontaneous term labour in 'normal' pregnancies are part of the natural history of fetal development, in contrast to that after artificially induced labour. Boths sets of data should not be included together in the same reference tables. The figures relating to the lower and higher gestational ages are especially biased by the increasing proportion of births after induction of labour. Although perinatal vital statistics tend to describe gestational age in weeks, we might consider this grid to be too vague. In Cleveland, 75.7% of the deliveries took place between the 38th and the 41st pregnancy week, i.e. within 4 weeks [4]. Taking into consideration the weekly 'growth' of about 140 g, the mean weight within 1 week will not coincide with the mid-week weight.

Moreover, the confidence intervals (CI) of the percentiles in the Amsterdam birth weight tables are so wide that the 95% CI of the 2.3rd and the 5th percentile overlap [5]. Nowadays, these tables could be too inaccurate to provide sufficient guidelines on which to base, for instance, clinical decisions about how and where to nurse infants which are small for their gestational age. An additional problem is the increasing discrepancy between the existing (Caucasian) tables and the (mixed) ethnic characteristics of the newborns. It is not yet clear which tables are most appropriate for them; tables from the country of origin or those derived from the new social environment. Furthermore, Gordosi has recently drawn attention to misclassification resulting from the disregard of maternal characteristics [6]. We tried to develop a method with which it would be possible to study unbiased comparison of subgroups of newborn babies within a changing society, avoiding the inaccuracies mentioned above. This method of classification will not only improve the birth weight tables for clinical use, but can also be used to classify newborns accurately in clinical studies. For this purpose we chose a description of birth weights of infants born near term from apparently 'normal' pregnancies [5].

Material and Methods

Singletons were chosen from the 38 771 babies that were born alive in the Amsterdam maternity wards from 1972 until 1982 (Departments of Obstetrics at the Vrije Universiteit Academic Hospital, Onze Lieve Vrouwe Hospital, St. Lukas Hospital and Slotervaart Hospital/Training School for Midwives). Relevant details, ascertaining health and circumstances at birth were taken from the computerized Cooperative Obstetric Registration [7]. These routinely registered records, however, are often incomplete. Incomplete and insufficiently accurate cases were excluded. Term delivery was operationalized by an amenorrhoea from 267 to 295 days (40 ± 2 weeks). This very conservative definition was used for statistical convenience, and in order to ensure that only term-born babies were included in the study population. (This excludes children born by an amenorrhoea from 259 to 266 days (37th week), who according to current recommendations are classified as term.) The gestational age was considered informative for the study if women with a regular cycle had a cycle length of 28 ± 2 days, or if a body temperature chart was kept. Cases that did not comply with this operationalization were excluded (31.5% of the study population). The possibility of impaired intra-uterine growth was assumed to be present in congenital malformed infants and if the pregnancy was complicated by gross pathological problems such as hypertension and/or proteinuria (6.6%), maternal diabetes mellitus (0.9%) and blood-group immunization (0.4%). These cases were all excluded. Spontaneous onset of labour was operationalized by excluding cases of induced labour after amniotomia or oxcytocine employment (9.5%), and by excluding cases of caesarean section (1.2%).

The following maternal characteristics had to be available: ethnic origin, height, (mid-pregnancy) weight and age. Cases were often excluded for more than one reason. Excluding large numbers was inevitable, as accuracy and 'normality' were of paramount importance. In the resulting sample of 14 109 complete cases with accurate registration of the maternal characteristics, gestational age at birth and birth weight were considered to be the physiological endpoints of fetal growth after a 'normal' pregnancy.

The association between birth weight in grams and gestational age in days was examined by multiple linear regression analysis in the Dutch (Caucasian) section of the study population. Birth weight was the dependent variable. In addition to gestational age in days as an independent variable, categorical variables were taken for both gender and parity modalities. Interactions among the determinants were studied in order to assess the suitability of the use of one single model for the traditional four tables of the Amsterdam birth weight tables. Subsequently, continuous variables for additional maternal characteristics (age, height and weight) were introduced in the regression model. These factors were retained, if statistically significant. Additional categorization of the continuous variables allowed us to assess deviations of the model. Covariance analysis, with interaction terms between the ethnic group and the continuous variables, was used to examine the validity of the model for all ethnic groups.

The 5th percentile of the birth weight was estimated as the means minus 1.64 times the standard deviation (S.D.), assuming a normal distribution and similar variances throughout the selected part of the term period. These assumptions were verified by checking distributions and variances in the four parity/gender modalities, after classification of the gestational age in weeks. In addition to point estimates of the mean and the 5th percentile computed by the model, examples of non-parametric estimates with 95% confidence intervals are given. The confidence intervals were derived from binomial tables [8] after reducing the samples based on the classification of gestational age in weeks to 500 babies a week, similar to the approach adopted by Kloosterman constructing the Amsterdam birth weight tables [3]. Dutch babies with a birth weight below and above the 5th percentile, according to the Amsterdam birth weight tables, [3] were identified and compared with the results of the parametric models. The calculations were performed by Number Cruncher Statistical System (version 5.0 10/87) [9].

Results

The mean birth weight of the Dutch Caucasian section (11 154) of the study population (14 109) was 3407 g. Multiple linear regression analysis showed that, in addition to gestational age in days, parity and gender are statistically significant determinants of birth weight. In the simple model (Table II) these determinants account for 14.5% of the variance. The small range of the gestational age is partly debit to this low value. None of the interaction terms among parity and gender reached a level of statistical significance. The introduction of a centered quadratic term of gestational age did not improve precision. These results provide strong evidence that during the chosen term period, birth weight on gestational age curves are parallel for girls and boys, as well as for children of primiparous and multiparous women.

The distribution of the birth weight over the population was also studied for gender and parity groups, classified by gestational age in weeks. Twenty groups were available. The Martinez-

TABLE I

The expected mean birth weight and 5th percentile of first-born boys produced by: (a) the simple multiple linear regression model; (b) non-parametric percentiles of the study population; and (c) readings from the Amsterdam birth weight tables [3].

Model	a	b	c
Estimated weight at:		* 95% CI	*
273 days	3236	3280 (3220, 3320)	3290
280 days	3375	3435 (3390, 3480)	3410
287 days	3502	3545 (3480, 3590)	3500
5th percentile at:	#	* 95% CI	*
273 days	2581	2670 (2550, 2730)	2600
280 days	2718	2650 (2600, 2750)	2720
287 days	2854	2890 (2780, 2960)	2790

#Computed by: mean — 1.64 S.D.

*Values taken from 273 ± 3, 280 ± 3 and 287 ± 3 days.

TABLE II

The regression coefficients with standard error (SE) of the simple (only Dutch) and extended regression models, to estimate the expected birth weight. Also given is the mean birth weight and standard deviation (S.D.) for the ethnic groups

Caucasian:	Simple model	Dutch	Mediterranean
count	11 154	11 154	660
Constant	−2102 (156)	−4167 (181)	−1845 (679)
Height (cm)	—	9.185 (0.672)	8.112 (2.572)
Weight (kg)	—	9.902 (0.967)	9.403 (1.760)
Gest age (day)	19.524 (0.556)	18.694 (0.531)	12.139 (2.006)
Age (year)	—	4.963 (0.947)	−1.657 (2.844)*
Parity[a]	157.3 (7.7)	138.2 (7.8)	186.2 (31.0)
Gender[b]	−135.5 (7.7)	−133.0 (7.4)	−163.6 (28.6)
R^2	14.5%	22.3%	22.9%
Mean weight (S.D.)	3407 (406)	3407 (388)	3406 (367)
	Other European	African	Asiatic
count	312	531	697
Constant	−2614 (1112)	−2969 (773)	−1764 (644)
Height (cm)	8.415 (3.489)	9.523 (2.798)	3.360 (2.488)*
Weight (kg)	10.866 (2.732)	9.963 (1.707)	12.509 (1.870)
Gest age (day)	13.794 (3.408)	14.623 (2.259)	12.898 (1.925)
Age (year)	2.980 (5.804)*	−0.153 (4.009)*	7.284 (3.011)
Parity[a]	128.8 (47.4)	72.3 (37.8)#	71.61 (28.4)
Gender[b]	−121.9 (45.1)	−93.8 (33.4)	−151.4 (27.0)
R^2	18.8%	21.1%	20.0%
Mean weight (S.D.)	3384 (397)	3315 (383)	3215 (354)

*No significant predictor ($P > 0.05$).
#$P = 0.056$.
[a]nulliparous X = 0, parous X = 1.
[b]Boys X = 0, girls X = 1.

Iglewicz test [10] showed a departure from the normal distribution in 6 of the 20 groups. The distribution in these groups was long-tailed. On the other hand, the birth weights half way through the studied period (day 281), irrespective of gender and parity ($n = 1670$), did not significantly deviate from the normal distribution. In our opinion, the slight non-normality of the study population, as is often found in very large samples, gives no cause for concern.

The computed standard deviation (340–439 g) within the gender and parity groups did not significantly depend on the gestational age. (After extending this model with strata for the maternal characteristics, it could be shown that the standard deviations in the strata did not deviate. Although statistically insignificant, deviations from the model occurred in the more extreme values of the continuous variables).

Therefore, the use of one value for the standard deviation throughout the term-period seems to be justified. Table I gives the point estimates according to the modeled mean and the 5th percentile, together with the observed non-parametric percentiles and the 95% confidence intervals for first-born boys. These figures do not differ significantly from the Amsterdam birth weight tables. The estimated means meet the non-parametric medians of the Amsterdam birth weight tables within 0.15 S.D., the estimates of the 5th percentiles differ up to 0.30 S.D. In our series of 11 154 Dutch babies, according to the simple model, 519 (4.7%) babies had a birth weight below the empirically observed 5th percentile, whereas 412 (3.7%) babies would be labeled as such according to the conventional Amsterdam birth weight tables. Assuming that our simple model is the best, the use of the conventional Amsterdam birth weight tables would wrongly consider 20 babies to be too small, corresponding with a sensitivity of 75.5%, and 127 babies would be wrongly considered normal, corresponding with a specificity of 99.8%. Only 15% of the variance in the Dutch section of the study population is accounted for by the simple model, so the analysis was extended to include other maternal characteristics. Correlation coefficients between the maternal characteristics and birth weight suggest that maternal mid-pregnancy weight (r = 0.295) and gestational age (r = 0.292) are equally suitable to estimate birth weight. Lower correlations were found between birth weight and maternal height (r = 0.224) and birth weight and maternal age (r = 0.111).

Multiple regression analysis showed that, in addition to gestational age, parity and fetal gender, the three maternal characteristics (maternal age, t = 5.2; maternal height, t = 13.7; mid-pregnancy weight, t = 10.2) are statistically significant determinants of birth weight in the Dutch section of the study population. Despite the high correlation between maternal height and weight (r = 0.455), both contribute significantly to the extended model. The extended model (Table II) accounts for 22.3% of the variance. The introduction of a centred quadratic term for the gestational age, or quadratic terms for the other continuous determinants causes no substantial change. The results of the introduction of additional categorization of the continuous variables are not given. The stratified analyses did not modify the results to any extent. By and large the deviations from the extended model are limited, and were of no statistical significance. In general, the regression approach seems to be correct. The introduction of additional determinants for the ethnic groups was less successful. By covariance analysis it could be shown that the combined estimates of birth weight by maternal height, mid-pregnancy weight and fetal gender do not differ among the maternal ethnic origins. Statistically significant interaction could be shown by the gestational age on the Mediterranean (t = −2.81) and Asiatic (t = −2.68) birth weights and by the parity modality on African (t = −3.20) birth weights. Discrepancy among the sizes of the ethnic groups makes interpretation difficult. We choose, for the time being, to present extended models separately for each ethnic group (see Table II).

Classification by the 5th percentile limits of the Amsterdam birth weight tables was also compared with that by the extended model in the series of 11 154 Dutch babies. Again, and now with less reserve, it was assumed that our model is the best. According to the extended model, in the Dutch section of the population 511 (4.6%) babies had a birth weight below the 5th percentile, whereas 412 (3.7%) babies would be labeled as such according to the Amsterdam birth weight tables. Moreover, 93 babies would be wrongly considered to be too small, corresponding with a sensitivity of 62.4%, and 192 babies would be wrongly considered normal, corresponding with a specificity of 99.3%.

Discussion

There is no generally accepted method available with which it is possible to describe reference birth weights. The birth weight tables used in various countries are constructed by adopting their own method of population selection or grouping data. The point estimates of the percentiles in the conventional Amsterdam birth weight tables are derived by non-parametric methods in weekly samples of less than 500 babies. This procedure of constructing percentiles results in a large random variation of the given figures.

This was recognized by Kloosterman [3], who softened his crooked lines by drawing smooth lines through the original data, and extracted from these lines the Amsterdam birth weight tables. In a former study [5] we described the same Caucasian part of the study population, irrespective of the availability of maternal characteristics (n = 15 877) and with an amenorrhoea from 36 to 43 weeks in a non-parametrical manner, complying with the methods of Kloosterman. The correlation between these data and the Amsterdam birth weight tables was very strong (r = 0.983). Because our study population is defined by leaving out all cases in which doubt could arise about the natural course of events, it approximates the population which described the natural history. By and large our selection of the study population corresponds to the Cleveland study [4]. In our report birth weights of 'normal' pregnancies are described with the purpose of improving accuracy of means and percentiles. The reasons for which cases were-excluded can be assumed to be independent of the result of interest (birth weight) and will thus not bias the estimates. This could possibly not be valid with regard to iatrogenic induced labour. However, in this case the infants would be mostly either too small (and premature), or too large (and serotine). This potential minor bias was softened by leaving out the minor adjustment by the quadratic term. We drew straight lines through our data, restricted to (a part of) the term period, using an accepted statistical method. Our statistical procedure could not justify four different shaped curves for each parity/gender combination, because we found that weight differences between newborn boys and girls were not modified by maternal parity. The advantage of being able to fit all data into one model is an improvement on precision.

The gender-related birth weight difference, after adjustment for parity and other factors calculated by Gardosi et al. [6] (120 g), agrees very well with ours (133 g). The weight differences among babies born after an intra-uterine sojourn of varying duration, 19.5 g/day (Table II, simple model), do not deviate much from the estimates that can be obtained from the Amsterdam birth weight tables (20 g/day). The diet-related differences Olsen et al. [11] describe, gestational age (283.3 vs. 279.4 days) and birth weight (3571 vs. 3445 g), after adjustment for gender and parity, are also 19.5 g/day. Therefore, the parameter estimates of the gestational age in our models comply with those from other population samples.

If the assumption of declining fetal growth before spontaneous onset of labour [12] is correct, the intra-uterine weight increase of a foetus just before spontaneous onset of labour would be minimal. Consequently, at the time of labour fetuses of mothers with induced labour are growing faster than spontaneous starters at the same gestational age, unless growth has stopped. These arguments justify the restriction of the analysis to babies born after a spontaneous onset of labour.

Other authors [e.g. 13–15] also defend the use of a linear relation between gestational age and birth weight. The mean birth weights at 40 completed weeks reported by Car Hill and Pritchard [16] (first-born boys 3421 g) and Altman and Coles [17] (first-born boys: 3414 g, S.D. 400) are very similar to ours. The mean birth weight given by Campbell and Newman [18], calculated from infants born after a 'normal' pregnancy (3400 g, S.D. 410), approximates our values, as do the figures given by Olsen [11]. The summary [19] of the percentiles during the term-period composed from five well-known birth weight tables, confirms our results, and the standard deviations are essentially the same. Caucasian newborns thus appear to be remarkably similar. Differences among birth weights among different Caucasian populations could possibly be mainly attributed to methods of selection, accuracy and analysis. It could be expected that labour starts haphazardly in any of the suitable pregnancies on a particular term-day. However, intra-uterine growth throughout pregnancy of heavier newborns is stronger than that of lighter newborns [18]. Random births in the set of growing fetuses, with a large range of growth velocity, would result in an increase in variance of the fetuses waiting to be born. That was not the case. Therefore, we postulate that slow-growing fetuses will be born at a lower gestational age than fast-growing foetuses.

Although the description of a population excludes variance, we handle population data as if it concerns samples. In fact, the data of standard growth curves are used to assess expected birth weights in other, chronologically later samples from the same population. Even large birth weight samples show a large annual variation. In our study population the median could differ as much as 120 g, and the non-parametric 5th percentile up to 240 g in tables corresponding with one parity and gender stratum [5]. Birth weight standardized for gestational age does not seem to be a very stable population characteristic. The parametric approach at least makes clear this limited importance of gestational age in assessing the expected birth weight. Our method of analysis was similar to the one adopted by Dougherty and Jones [20]. They restricted their study population to gestational ages up to 37 weeks and, in addition to physical determinants, also included social characteristics. The variance this model accounted for was 26%, which is similar to ours (22%). The inclusion of variables for diet composition [11,21] and smoking habits [22] could possibly improve the model. The model presented by Gardosi et al. [6] is more complicated than ours. They superimpose term birth weights, adjusted for population characteristics on weight estimates from published mathematical formulae. However, their population is not screened for spontaneous onset of labour, and pathological cases are not excluded. Nevertheless, their mean birth weight is only 101 g (0.25 S.D.) lower than ours. They found misclassification at the 10th percentile similar to that which we found at the 5th percentile.

Migrants are a specific subgroup of the population because characteristics of their ethnic origin and environment are intermingled with those of the new country. The number of migrant cases in our study population seems to be too small to suggest a definite relationship between population characteristics and birth weight. Nevertheless, differences among ethnic groups seem mainly to be directly related to population characteristics such as maternal height and mid-pregnancy weight. Doornbos and Nordbeck [23] analyzed the influence of ethnic origin on birth weight in a migrant population. Their results are difficult to interpret, as they combined the sexes to increase numbers. We conclude that birth weight percentiles for members of migrant populations can be improved if adjustments are made for population characteristics.

How much the birth weight deviates from the population mean can only be determined if accurate expected birth weights are available. The Amsterdam birth weight tables summarize properties of growth, with a minimum of medical intervention and a variety of natural characteristics. Nowadays these tables provide only very rough guidelines on which to base, for instance, clinical decisions about how and where to nurse infants who are small for their gestational age. Updating these tables is difficult in view of improved medical skills and interventions. We construed a reference set of data derived from infants born after spontaneous onset of labour after an apparently 'normal' pregnancy.

Our conclusion is that integration of the four currently used tables into one, and adjustment for easily available maternal characteristics, could substantially improve classification methods.

Acknowledgments

The authors are very grateful to everyone who helped with the deliveries and to those who scrupulously registered the data. We are indebted to Dr. A. van Enk and Dr. W.J. Honnebier for their permission to make use of the data from their departments. We thank F. Maddever for correcting the English. We gratefully acknowledge the late Professor Dr. L.A.M. Stolte's interest and enduring discussion about fetal growth. Grant: Praeventiefonds, 281204.

References

1 Callan NA, Witter FR. Intrauterine growth retardation: characteristics, risk factors and gestational age. Int J Gynecol Obstet 1990;33:215–220.
2 Eskes M. The Wormerveer study. Thesis. University of Amsterdam, Amsterdam. 1989;Chapter VIIIA:155–164.
3 Kloosterman GJ. On intrauterine growth. Int J Gynaecol Obstet 1970;8:895–912.
4 Brenner WE, Edelman DA, Hendricks CH. A standard of fetal growth for United States of America. Am J Obstet Gynecol 1976;126:555–564.
5 Voorhorst FJ, Puyenbroek JI, Robertson EA, Bezemer PD, Kurver PHJ. Verschillen de geboortegewichten van vroeger en nu? Ned Tijdschr Geneeskd 1990;134:998–1002.
6 Gordosi J, Chang A, Kalyan, Sahota D and Syonds EM. Customised antenatal growth charts. Lancet 1992;339:283–287.
7 Van Hemel OJS. An obstetrical data-base. Human factors, design and reliability. Thesis. Amsterdam: Vrije Universiteit, 1977.
8 Documenta Geigy. Wissenschaftliche Tabellen. 7. Auflage. In: Diem K, Lentner C, eds. Basel: Geigy Pharma, 1968;102.
9 Hintze JL. Number Cruncher Statistical System. Version 5.01. Kaysville, Utah 84037, 1987.
10 Martinez J, Iglewicz B. A test for departure from normality based on a biweight estimator scale. Biometrika 1981;68:331–333.
11 Olsen SF, Sorensen JD, Secher NJ, Hedegaard M,

Henriksen TB, Hansen HS, Grant A. Randomised controlled trial of effect of fish-oil supplementation on pregnancy duration. Lancet 1992;339:1003–1007.

12 Gruenwald P. Pathology of the deprived fetus and its supply line. In: Size at birth. Ciba foundation Symposium 27. Amsterdam: Elsevier-Exerpta Medica-North-Holland Associated Publishers, 1974:3–19.

13 Secher NJ, Kern Hansen PL, Lenstrup C, Pedersen-Bjergaard L, Sindberg Eriksen P, Thomsom BL and Keiding N. Birth weight for gestational age charts based on early ultrasound estimation of gestational age. Br J Obstet Gynaecol 1986;93:128–134.

14 Dunn PM. The Gestogram — A New Standard Perinatal Growth Chart. In: Saling E, Schulte FJ, eds. Perinatal Medizin: Band 2. Proceedings of the 3rd German Congress of Perinatal Medicine. Thieme Verlag: Stuttgart, 1970:242–247.

15 Bjerkedal T, Bakketeig L, Lehman EH. Percentiles of birth weights of single, live births at different gestational periods. Acta Paediatrica Scand 1973;62:449–457.

16 Carr Hill RA, Pritchard CW. Reviewing birth weight standards. Br J Obstet Gynaecol 1983;90:718–725.

17 Altman DG, Coles EC. Assessing birth weight for dates on a continuous scale. Ann Hum Biol 1980;7:35–44.

18 Campbell S, Newman GB. Growth of the biparietal diameter during normal pregnancy. J Obstet Gynaecol Br Commonw 1971;78:513–519.

19 Dunn PM. A Perinatal Growth Chart for International Reference. Acta Paediatr Scand, Suppl 1985;319:180–191.

20 Dougherty CRS, Jones AD. The determinants of birth weight. Am J Obstet Gynecol 1982;144:190–200.

21 Kardjati S, Kusin JA, De With C. Energy supplementation in the last trimester of pregnancy in East Java: I. Effect on birthweight. Br J Obstet Gynaecol 95:1988;783–794.

22 De Haas JH. Review article. Parental smoking. Its effect on fetus and child health. Eur J Obstet Gynecol Reprod Biol 1975;5:283–296.

23 Doornbos JPR, en Nordbeck HJ. Birth weight. In: Perinal Mortality. Obstetric factors in a community of mixed origin in Amsterdam. Amsterdam: Universiteit van Amsterdam, 1985. Thesis. Dordrecht: ICG Printing, 1985:156–162.