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Confidence Intervals and Sample-size Calculations for the Sisterhood Method of Estimating Maternal Mortality

James A. Hanley, Catherine A. Hagen, and Tesfaye Shiferaw

The sisterhood method is an indirect method of estimating maternal mortality that has, in comparison with conventional direct methods, the dual advantages of ease of use in the field and smaller sample-size requirements. This report describes how to calculate a standard error to quantify the sampling variability for this method. This standard error can be used to construct confidence intervals and statistical tests and to plan the size of a sample survey that employs the sisterhood method. Statistical assumptions are discussed, particularly in relation to the effective sample size and to effects of extrabinomial variation. In a worked example of data from urban Pakistan, a maternal mortality ratio of 153 (95 percent confidence interval between 96 and 212) deaths per 100,000 live births is estimated. (STUDIES IN FAMILY PLANNING 1996; 27,4: 220–227)

The sisterhood method developed by Graham and colleagues (Graham et al., 1989) was an important advance in providing a method for calculating indexes of maternal mortality in countries or regions where data on vital events are not routinely and reliably collected. This indirect method uses the proportions of adult sisters dying during pregnancy, childbirth, or the puerperium reported by adults during a census or survey to derive an estimate of the lifetime risk of maternal mortality. From this estimate, the maternal mortality ratio (MMR, number of maternal deaths per 100,000 live births) can then be derived. The method has many practical advantages, such as ease and speed of application. Since its introduction, reports of the application of the sisterhood method in Africa and Asia have been published (Chiphangwi et al., 1992; David et al., 1991; de Groof et al., 1995; Oosterhuis, 1993; Shiferaw et al., 1993; Hagen, 1995; Walraven et al., 1994; Wirawan and Linnan, 1994), and other unpublished studies have been incorporated in national demographic surveys (AbouZahr and Royston, 1991) or mother and child health surveys.

The authors have recently used the sisterhood method to estimate the maternal mortality ratio in southwestern Ethiopia (Shiferaw and Tessema, 1993) and in urban Pakistan (Hagen, 1995). In order to compare these estimates with others, we wished to know whether the differences could simply be due to their being based on a sample survey and so subject to sampling variability.

This report addresses the need for a method of calculating sampling error for sisterhood estimates of maternal mortality. Because the derivations of the required inferential procedures are nonstandard, they are presented in detail in the hope that other users can now report their results not just as point estimates, but with interval estimates that convey the margin of sampling error. The formulas given also can be used to project the sampling variability associated with various sizes of sample surveys and thus to plan the size of surveys employing the sisterhood method.

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Preliminaries

The risk, Q, of lifetime mortality from maternal causes is estimated from the number of sisters' deaths reported by respondents (which Graham and colleagues call r) as a fraction of the number of sister units of exposure (which they call B), that is,

$$Q = r/B. \quad (1)$$

The lifetime probability of avoiding death from maternal causes is, therefore,

$$P = 1 - Q. \quad (2)$$

The maternal mortality ratio (MMR) can then be calculated from the approximation

$$MMR = 1 - (P)^{1/TFR}, \quad (3)$$

where TFR is the total fertility rate.[1] In fact, as is easily verified, even when Q and TFR are high, a good approximation of the MMR is simply

$$MMR \approx Q/TFR. \quad (3')$$

This simplified formula shows the relationship between the risk of death for each pregnancy (sometimes known as the obstetrical risk) and the number of pregnancies per woman (total fertility rate); the overall life risk of maternal death is a compound of both.

Sampling Variability

Situations are considered now where r and B (and possibly TFR) are derived from sample surveys so that the estimates of Q (or equivalently, P) and MMR are subject to sampling variability. Although in theory B is random, for all practical applications it can be treated as fixed—in the same spirit that the n's in most studies are not really fixed, but are nevertheless treated as such for statistical inference. The implications of the various other components in our assumptions are discussed below, but we begin by assuming that the estimate r/B of Q can be treated for purposes of statistical inference as a binomial variable with parameter Q. (Later, because Q is low, r will be treated as a Poisson arising from a denominator of B, but here, because readers may be more familiar with the binomial distribution, it is treated as a binomial.) Thus, if r is 10 or higher, a confidence interval for Q can be calculated from the Gaussian approximation to the binomial as $r/B \pm z_{\alpha/2} SE(r/B)$, where $z_{\alpha/2}$ is the appropriate normal deviate corresponding to a two-sided confidence level of $100(1-\alpha)$ and the standard error, $SE(r/B)$, of r/B is calculated as $\sqrt{\{r/B\}\{1-r/B\}/B}$. Because they are used below, the lower and upper limits for Q are referred to as Q_L and Q_U, respectively. In most situations, r will be greater than the minimum of 5 or 10 suggested for using large-sample approximation; if it is not, tabulated exact confidence intervals can be resorted to for a binomial proportion, but in such a case, the confidence interval will be so wide as to be virtually useless.

If the TFR can be taken as a constant, that is, as without any sampling error, then the lower and upper limits for the MMR can be obtained by substituting Q_L and Q_U into equation *(2)* to obtain lower and upper limits for P and then by substituting these two limits into equation *(3)* or *(3')* to obtain the corresponding limits MMR_L and MMR_U.

If the TFR is itself an estimate that is subject to sampling error, then the lower and upper limits for MMR are more complex. The confidence interval (CI) can be written as $MMR \pm z_{\alpha/2} SE(MMR)$, where the SE is the standard error calculated from the laws governing the amalgamation of the sampling errors in the estimates of both P and the TFR. This SE, which is derived in Appendix 1, can be calculated by combining their sampling variances, as

$$SE[MMR] = \frac{1-MMR}{TFR}\sqrt{\frac{1-P}{B \cdot P} + [\log\{P\}]^2 Var[TFR]/TFR^2} \quad (4)$$

with P replaced by its estimate $1-r/B$. The two parts under the radical sign can be seen to represent the sampling uncertainties inherent in the estimates of P and the TFR respectively; if the estimate of the TFR can be taken as being without sampling error, then the second term becomes zero.

The Worked Example

In January 1994, a sisterhood survey was conducted in Karachi, Pakistan, as part of a maternal health and health services survey (Hagen, 1995). The sample of 2,897 households was chosen randomly from the catchment area of the three Aga Khan Health Services of Pakistan maternity homes in urban Karachi. Survey respondents were all ever-married women between the ages of 15 and 54 years; the survey response rate was 93 percent. The four sisterhood questions were used as suggested by Graham and colleagues (1989). Respondents were asked the number of ever-married sisters born to the same mother. The 2,651 respondents re-

ported that among their ever-married sisters, 146 deaths had occurred, of which $r = 27$ (18.5 percent) were reported as the result of maternal causes. The sisterhood data are presented in Table 1, showing that these 27 deaths arose from a denominator of B = 3,767 sister units at risk.

According to formula *(1)* above, the estimate of life risk of maternal death in this population was Q = r/B = 0.00717. Substituting this estimate of life risk of maternal death, *Q*, and an estimate for the total fertility rate of 4.69 (obtained directly from the survey data) into formula *(3)* yielded an estimate of the maternal mortality ratio of

$$MMR = 1 - (1 - 0.00717)^{1/4.69} = 0.00153$$

or 153 maternal deaths per 100,000 live births.

In the next step, 95 percent confidence limits are calculated for this estimate of the MMR. A 95 percent CI for *Q* is given by $0.00717 \pm 1.96\sqrt{\{0.00717\}\{0.99283\}/3767}$ or $Q_L = 0.00449$, $Q_U = 0.00989$. If, for the purposes of illustration, the total fertility rate of 4.69 is taken as being without sampling error, then the corresponding 95 percent limits for the estimated MMR can be found simply by substituting the 0.00449 and 0.00989 into equation *(3)* to obtain $MMR_L = 0.00096$ and $MMR_U = 0.00211$, that is, between 96 and 211 maternal deaths per 100,000 live births.

In reality, the TFR of 4.69 does contain sampling variability. In Appendix 2, the method used to calculate a standard error for the TFR is shown, which is estimated to be $\sqrt{0.0245} = 0.156$. Thus, the 95 percent confidence limits for the TFR can be found as $4.69 \pm 1.96(0.156)$, or between 4.39 and 5.00. Therefore, the SE for the estimated MMR, taking into account sampling variability of both the estimate of *Q* and the estimate of the TFR, can be found from equation *(4)* as

$$\frac{0.99283}{4.69}\sqrt{\frac{0.00717}{3767\{0.99283\}} + \frac{[\log\{0.99283\}]^2\{0.0245\}}{4.69^2}}$$

or 0.000297, from which are derived the limits as $0.00153 \pm 1.96(0.000297)$ or {0.00095, 0.00211} or 95 to 211 maternal deaths per 100,000 live births.

Clearly, the limits here are only slightly different than they were before. Even if the second component were zero, the two methods use two slightly different approaches to forming a CI for the MMR: the first transforms symmetric uncertainty in *Q* into asymmetric uncertainty in the MMR, whereas the second assumes symmetric uncertainty directly in the MMR scale. When the relative contributions from the two sources of variation combined in equation *(4)* are examined, the variance stems predominantly from the *Q* component (where the SE relative to *Q* is 0.00137/0.00717 or 19 percent) rather than from that associated with the TFR (where the SE relative to the TFR is 0.156/4.69 or only 3.3 percent).

Table 1 For nine sisterhood studies, numbers of reported sister deaths in relation to maternal mortality rate (MMR), total fertility rate (TFR), and survey sample size, by study

Study/date	MMR[a]	TFR	Respondents N	Deaths r	r 1000N
Karachi (1994)	131	4.7	2,491	21	8
Bali (1991)	282	3.5	26,157	210	8
Zimbabwe (1991)	323	5.5	3,493	70	20
Tanzania (1990)	288	6.5	2,527	63	25
Malawi (1989)	373	6.7	3,333	88	26
Ethiopia (1991)	623	7.6	1,848	68	37
Djibouti (1989)	740	6.8	7,408	374	50
The Gambia (1987)	998	6.0	1,652	65	39
Niger (n.d.)	1,132	7.0	2,654	150	57

[a]All MMRs reported in this table are recalculated from published data, restricting respondent age to younger than 50 years.

Sources: Karachi—Hagen, 1995; Bali—Wirawàn et al., 1994; Zimbabwe—Oosterhuis, 1993; Tanzania—Walraven et al., 1994; Malawi—Chiphangwi et al., 1992; Ethiopia—Shiferaw and Tessema, 1993; Djibouti—David et al., 1991; The Gambia—Graham et al., 1989; Nigèr—de Groof et al., 1995.

Calculating Sample Size to Estimate the MMR

We were unable to find in the original work by Graham and colleagues (1989) or in subsequent literature any method for calculating sample-size requirements for surveys employing the sisterhood method to estimate maternal mortality. Graham and colleagues suggest that "for an estimate of the broad level of maternal mortality, interviews with 3,000–6,000 adults will be required, depending on the maternal mortality and the number of sisters per respondent that can be expected to have reached reproductive age." Ebràhim (1991) says that "a sample size of 2,500–3,000 is desirable to obtain reliable estimates." In this section, sample-size requirements are addressed in relation to the desired statistical precision of the survey estimate, the maternal mortality rate, and the upper age limit of the respondents to be surveyed.

The statistical precision of the MMR estimate depends on the size of the standard error, which in turn depends directly on the number of reported deaths, *r*. For practical purposes, and because it simplifies the form of standard deviations, *r* can be treated as having Poisson variation, with a large enough expectation that its distribution can be reasonably approximated by a Gaussian distribution with standard deviation = $\sqrt{\text{expectation}}$.

Therefore, if the variation in the estimate of the TFR is much smaller than that of r, and if a 100(1–α) percent confidence level is used, the percentage margin of error (%ME) in the estimate of the MMR is close to $\frac{100z_{\alpha/2}\sqrt{r}}{r}$. Thus, the sample size should be large enough that the desired $\%ME \geq 100z_{\alpha/2}/\sqrt{r}$, or

$$r \geq z_{\alpha/2}^2 \{100/\%ME\}^2.$$

As an example, if we wished to estimate the MMR with a margin of error of 25 percent (95 percent CI), we should plan on a survey that would be large enough to yield $r \geq 1.96^2 \{100/25\}^2 = 62$ deaths. For other margins of error, but with the same 95 percent confidence, the requirements are:

Margin of error (%)	Number of deaths (r)
±50	16
±40	24
±30	43
±25	62
±20	97
±15	171
±10	385

Unfortunately, in practical terms, translating the desired number of reported deaths, r, into sample size directly is difficult, because r varies with the magnitude of both the MMR and the TFR, and with the age structure of the respondent population. Although r can clearly be seen to be higher in populations with higher maternal mortality and higher fertility, the relationship with age structure is more complex. This greater complexity is due to the necessity for age-adjustment factors used in the sisterhood method to generate sister units of risk. Simply put, in a given population, the required number of reported deaths, r, could be obtained by sampling a small number of older respondents or a larger number of younger respondents. Young respondents contribute fewer sister units of risk to the sample total because of the adjustment factors necessary to the method. Yet samples that have a higher proportion of young respondents estimate maternal mortality for a more recent time period. Therefore, a trade-off is inherent in sisterhood-method sample-size calculations: For any given sample size, the older the average age of the respondents, the smaller the margin of error, but the greater the lag between the date of data collection and the time-reference period of the study. For this reason, we would reiterate Graham's suggestion that studies employing the sisterhood method restrict the age of respondents to younger than 50 years, which has the effect of restricting the time-reference period for the MMR estimate to a date that is consistently 11 to 12 years prior to data collection. (Graham et al. describe the method for calculating time location.) If no restriction is placed on respondent age (Shiferaw and Tessema, 1993), time-reference periods can be unacceptably long (18 years or more).

If a decision is made to limit the respondent age to younger than 50 years, the sample-size requirement to generate a given number of reported deaths can be studied empirically. In Table 1, the yield (defined as the number of reported sister deaths per thousand respondents) in nine sisterhood surveys was summarized. In the table, yield was shown to be related to both the maternal mortality rate and the total fertility rate. Roughly speaking, the data suggest that in areas with maternal mortality in the range of 250 deaths per 100,000 live births, surveying 1,000 respondents will generate approximately 20 reported sister deaths; in areas where the MMR is 500, 30 sister deaths; and in areas where the MMR is 750, 40 sister deaths.

In Table 2, the yield from published studies is used to anticipate sample-size requirements according to MMR and the desired margin of error. The investigator is required to make an educated guess as to the prevailing MMR in the study and to choose an acceptable margin of error. The table then provides the sample sizes. For example, in an area with an MMR of approximately 500 deaths per 100,000 live births, where a study is being planned to provide an estimate of the MMR with a 20 percent margin of error, the sample size would be 3,200 respondents who are younger than 50 years of age.

Alternative, conventional methods of estimating maternal mortality (which use direct ratios of reported maternal deaths to live births in each household over a fixed recall period) are costly in terms of sample size. For example, conventional sample-size calculations (assuming the same MMR, TFR, and a five-year fixed recall period) for the population of urban Pakistan presented above would suggest the need to survey more than 17,000 households to produce an MMR estimate

Table 2 Approximate sample sizes (numbers of respondents) according to level of maternal mortality ratio (MMR) and desired margin of error

		Margin of error		
MMR	Yield[a]	±30% (r > 43)[b]	±20% (r > 97)	±10% (r > 385)
250	15	3,000	6,400	25,000
500	30	1,500	3,200	13,000
750	45	1,000	2,100	8,000

[a]Yield = Approximate number of reported deaths per 1,000 respondents, obtained from Table 1. [b]r = Total number of deaths.

with a comparable margin of error to the sisterhood study that used fewer than 3,000 households.

Sample Size for a Comparison of Two MMRs

To make a comparison of two MMRs, each based on a survey using the sisterhood method, the statistical power to detect a percentage difference in rates will ultimately depend on the expected numbers of "events" in each sample (Smith and Morrow, 1991). Thus, for a given type I error α, and power 1–ß to detect a Δ percent increase or decrease in the MMR relative to that in the reference population, the number of deaths b in the sample from the reference population should be large enough that

$$r \geq 100 \left\{ \frac{z_{\alpha/2}\sqrt{200} + z_{\beta}\sqrt{200 + \%\Delta}}{\%\Delta} \right\}^2.$$

As an example, for a comparison of two MMRs, with a two-sided test with $\alpha = 0.05$ and power $1-ß = 0.80$ to detect a 50 percent decrease ($\%\Delta = -50$) in the MMR from that in the reference population, the number of deaths r in the sample from the reference group should be large enough that

$$r \geq 100 \left\{ \frac{1.96\sqrt{200} + 0.84\sqrt{200 - 50}}{50} \right\}^2 = 58.$$

Appropriateness of the Binomial Formula

In treating r/B, the estimate of Q, as a binomial proportion with the conventional standard error based on a sample size of B, the denominator (and thus the numerator) is implicitly assumed to involve sisters, each of whom is reported on once only. As Trussell and Rodriguez (1990) point out, the multiple counting that occurs when there are two or more respondents from the same family does not create any bias in the estimate of Q. However, in judging the possible sampling error in the estimate, any multiple counting means that the effective sample size (B') is smaller than B; therefore, strictly speaking, B should be replaced by B' in the two formulas shown above for calculating SE(Q) and SE(MMR); otherwise, the calculated SEs and CIs will be artificially narrow.

In the hypothetical example described by Trussell and Rodriguez, Q was estimated using denominators and numerators that were each inflated by a factor of 1.5 (if every sister in the family is included as a respondent, the inflation factor is of the order of (1– Q) times the average number of sisters in a family who reach childbearing age). In practice, two factors will tend to make the inflation in sample size that results from multiple reporting negligible (that is, they will make B only slightly higher than B'). First, unless the survey is carried out in a wide area and unless sisters live very near each other, the survey is unlikely to include as respondents all living members of families. Second, even if it is carried out in a wide area, it is likely to involve sampling; a sampling fraction smaller than one in four will reduce dramatically the multiple appearances of data on the same sisters. However, cluster sampling will lead to overcounting if, in rural areas, marriages tend to be indigenous.

Trussell and Rodriguez closed their note by emphasizing, as did Graham and colleagues, that "the key assumption of the sisterhood method is independence of the mortality experiences of adult sisters." By using the binomial formula to calculate a standard error, this assumption is made here also. If, however, maternal mortality does segregate by families, that is, if the variation in mortality from family to family is more than that predicted by the binomial distribution (in technical terms, if the variation is extrabinomial), SEs based on the binomial would be too small. However, given that Q is typically no more than 5 percent, and the number k of sisters reported on per respondent is fewer than three on average, the segregation of rates would have to be considerable before the extrabinomial variation would contribute substantially to the variability of the estimate based on the sisterhood method. For example, if Q is 2 percent on average in 50 percent of families and 8 percent in the other 50 percent, the SE of the estimate of Q calculated from the binomial formula should be raised by a factor of only 1 percent if k is 2 and of 1.9 percent if k is 3; even if the rates are as heterogeneous as 1 percent and 9 percent, the SEs would have to be raised by only 1.7 percent and 3.3 percent, respectively. In practice then, any extrabinomial variation in risk (that is, clustering of mortality within families) can be ignored in the calculation of standard errors.

Discussion

Our experience with using the sisterhood method in multipurpose household surveys confirms its ease of use in the field. However, the method has some important limitations. The most important of these, as demonstrated by Graham and colleagues, is the lag time between the period of data collection and the time-reference period of the study. If respondents up to and in-

cluding the age group older than 60 years are included, the time-reference period will be centered as much as 18 years prior to data collection. By restricting respondents to those younger than 50, this lag can be reduced to approximately 12 years; further restrictions will continue to decrease the lag, but will increase the number of households required to complete the sample. Even if only the youngest respondents are included (15–19 years), however, the time-reference period will be centered, on average, 5.7 years prior to data collection; the sisterhood method cannot be used to produce more recent estimates. Whereas the time lag inherent in the sisterhood method may be unimportant where a simple descriptive estimate of the maternal mortality ratio is required, the method may be inappropriate for use in areas where maternal mortality has been changing quickly.

The sisterhood method is efficient in terms of sample size when compared with conventional methods of estimating maternal mortality. Part of the efficiency is the result of the method whereby each respondent reports concerning all of his or her sisters; part is the result of using as the unit of recall the life experience of each sister.

Conclusions

Where a regional or local estimate of maternal mortality is required and the age of the estimate (five versus 12 years) is not critical, the sisterhood method is efficient and practical. It has the great advantage of simplicity of use; the four questions necessary can be appended to a multipurpose survey. Where current estimates of maternal mortality are required, alternate, direct, methods of data collection using networking (Boerma and Mati, 1989) and a combination of official and unofficial sources (Kumar et al., 1989; Bhatia, 1990) may be better.

Estimates of sampling variability for the sisterhood estimate can be used a priori to calculate sample size requirements and a posteriori to provide confidence limits for estimates of the maternal mortality ratio. The use of a confidence interval is beneficial in quantifying the degree of uncertainty in the estimate that is the result of sampling alone. Besides the sampling variability, many potential sources of bias exist in retrospective estimates based on cross-sectional surveys. Several uncertainties, in the form of assumptions and approximations, are built into estimates derived from the sisterhood method. Whereas sampling uncertainties can be quantified, the other, less quantifiable components that can lead to an inaccurate (but possibly reproducible) estimate cannot be ignored. Although the purpose of this report is to quantify estimation errors whose magnitudes are a function of the number of sister units of risk (B) and the number of deaths (r), continued efforts should be made to minimize the effects of potential sources of bias that, unfortunately, will not be reduced simply by increasing the number of sister units of risk.

Appendix 1: Derivation of the Standard Error of an Estimate of the MMR

We calculate $P = 1-Q = 1 - r/B$. Then,

$$MMR = 1 - P^{1/TFR},$$

so

$$1 - MMR = P^{1/TFR},$$

or

$$\log\{1 - MMR\} = \log\{P^{1/TFR}\} = \log\{P\} / TFR.$$

Using the formula for the variance of the log of a statistic in relation to the variance of the statistic itself (see equation 3.17, page 92, in Armitage and Berry, 1987), the formula for the variance of a ratio of two uncorrelated statistics (equation 3.15 from page 91), and the formula for the variance of a binomial statistic (equation 3.4, page 85), we calculate

$$\begin{aligned}
\text{Var}[MMR] &= \text{Var}[\{1 - MMR\}] \\
&= \{1 - MMR\}^2 \text{Var}[\{\log\{1 - MMR\}] \\
&= \{1 - MMR\}^2 \text{Var}[\log\{P\} / TFR] \\
&= \{1 - MMR\}^2 \{\text{Var}[\log\{P\}] / TFR^2 + (\log\{P\})^2 \text{Var}[TFR] / TFR^4\} \\
&= \{1 - MMR\}^2 \{\{\text{Var}[P] / P^2\} / TFR^2 + (\log\{P\})^2 \\
&\quad \text{Var}[TFR] / TFR^4\} \\
&= \{1 - MMR\}^2 \{\{P(1-P) / B\} / P^2\} / TFR^2 + \\
&\quad (\log\{P\})^2 \text{Var}[TFR] / TFR^4\},
\end{aligned}$$

so that the standard error of the estimated MMR simplifies to

$$SE[MMR] = \frac{1 - MMR}{TFR} \sqrt{\frac{1-P}{B \cdot P} + \frac{[\log\{P\}]^2 \text{Var}[TFR]}{TFR^2}}.$$

Appendix 2: Method for Calculating the TFR and Its Sampling Uncertainty

This method is illustrated using the data in the first three columns of Table A2.

Data: For each age category, the number (y) of births over a period of T years, and the number (n) of women in the age category.

Table A2 Worked example (based on data from Hagen, 1995)

Maternal age group (each w = 5)	Births 1979–83 (T = 5) y	Women (N) n	Births per women-year $p = \frac{y}{n \cdot T}$	Subtotal: Births over w years $= w \cdot p$	Var[p] $= \frac{p \cdot [1-p]}{n \cdot T}$	$w^2 \cdot Var[p]$
15–19	300	790	0.076	0.380	1.8×10^{-5}	4.4×10^{-4}
20–24	618	552	0.224	1.120	6.3×10^{-5}	1.6×10^{-3}
25–29	559	478	0.234	1.169	7.5×10^{-5}	1.9×10^{-3}
30–34	302	319	0.189	0.947	9.6×10^{-5}	2.4×10^{-3}
35–39	100	248	0.081	0.403	6.0×10^{-5}	1.5×10^{-3}
40–44	19	37	0.103	0.514	5.0×10^{-4}	1.2×10^{-2}
45–49	6	37	0.032	0.162	1.7×10^{-4}	4.2×10^{-3}
Total				4.694		2.45×10^{-2}
				$\hat{TFR}$		Var $\hat{TFR}$

$SE(\hat{TFR}) = \sqrt{0.0245} = 0.156$. 95% confidence interval: 4.694±1.96(0.156) or (4.39,5.00)

Note: The formula for the variance of the TFR is similar to that for cumulative rates given in equation 2.2 on page 59 of Breslow and Day (1987). Since the events they are concerned with are less common, they ignore the factor (1–p) in each subvariance.

Step 1—Subtotal and total fertility rates: (1) Calculate category-specific fertility rates (p) by dividing the number (y) of live births by the number of woman-years ($n \cdot T$). Each such rate, called p, is in units of live births per woman-year. (2) Obtain the subtotal fertility for the age span by multiplying the category-specific fertility rate (p) by the width (w) of the age category. Coincidentally, in the example above w = T. (3) Sum the category-specific fertility subtotals ($w \cdot p$) to give the TFR.

Step 2—Variances of subtotal fertility rates: If the sampling of women within an age category is close to simple random sampling, the variance of each p can be approximated by a binomial formula with proportion p and denominator = $n \cdot T$, namely $Var[p] = p(1-p)/(n \cdot T)$. (If the sampling is based on stratified or cluster sampling, the binomial-based variance may have to be adjusted downward or upward by an appropriate factor.) The variance of the subtotal fertility for the age range is then calculated as $var[w \cdot p] = w^2 \cdot Var[p]$.

Step 3—Variance and standard error (SE) of total fertility rate: The variance of the estimate of the TFR is obtained by summing the variances of the fertility subtotals. (If the samples for the different age categories are from the same sample of geographic clusters, we could add any positive covariances induced by this.) The SE is then calculated as the square root of the total variance.

Note

1 Elsewhere (for example, see Haub, 1988) other authors have used 1.2TFR in this formula rather than 1.0TFR; presumably they have done so to reflect the fact that a woman can die from any pregnancy (not just from one that ends in a live birth) and that an estimated 0.2 pregnancies do not produce a live birth for every one that does.

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