

MATERNAL MORTALITY IN INDIA: AN UPDATE

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Introduction

In spite of the growing concern about reproductive health, information on levels, trends and differentials in maternal mortality remains fragmentary in most developing countries. Policy initiatives often rest on judgements made on the basis of a small, selective cross-section of the population. For India, the National Family Health Survey of 1992-93 was the first to provide a national-level estimate of 437 maternal deaths per 100,000 births for the two-year period preceding the survey (International Institute for Population Sciences, 1995). But in spite of surveying nearly 90,000 households, it could not produce estimates at regional or state-levels owing to the smallness of the sample. Even at the national level, the sample inadequacies of the NFHS came into sharp focus when the second round of the survey in 1998-99 produced a maternal mortality estimate of 520, but failed to confirm statistically the possible rise in the level of maternal mortality (International Institute for Population Sciences and ORC-Macro, 2000).

To fill the data gap, in recent times, the potential of the Sample Registration System - a dual record system for collecting data on births and deaths - for estimating maternal mortality has also been explored. The source has recorded a maternal mortality rate of 408 and 407 for 1997 and 1998, respectively (India, Registrar General, 1999 and 2000). While the suggested level seems plausible at the all-India level, the state-level patterns indicated by this source appear highly improbable (such as very low estimates for Gujarat and Tamil Nadu and relatively high estimates for Kerala).

As direct investigations require a huge sample, several indirect methods have been proposed for the estimation of maternal mortality. One of these is the so called 'sisterhood method' developed by Graham and others (Graham *et al.*, 1989). This method makes use of the data collected from female respondents in a sample survey on the number of ever-married sisters they had, the number who were not currently alive, and the number who died while pregnant, during childbirth or within six weeks after delivery. The procedure of estimation can be made more direct by adding a few more questions on siblings, as done in some Demographic and Health Surveys (See Rutenberg and Sullivan, 1991). The sisterhood method cuts down the required sample size drastically because women generally have several sisters who could have been exposed to the risk of maternal mortality, each time when they were pregnant. The method has been used to estimate maternal mortality with some success in African populations. In India, opportunities for using the sisterhood method have been limited owing to paucity of data.

An indirect procedure of estimating maternal mortality from the sex differentials in mortality at reproductive ages has been developed by Bhat *et al.* (1995). Basically, it involves the regression of sex differentials in mortality rate on fertility rate, by age. The method can be applied to any source that provides estimates of mortality rates by age and sex, and fertility rates

by age. The main advantages of this method are that it requires no special questions to be canvassed, and provides estimates for more recent times than the indirect version of the sisterhood method. Comparisons made with the estimates from other sources indicate that this approach gives quite plausible levels of maternal mortality for India. However, there have been few applications of the method to data outside south Asia.

The objectives of this paper are two fold. First, it presents estimates of maternal mortality derived from the sisterhood questions asked in the Human Development Profile Survey (HDPS) conducted by the National Council of Applied Economic Research in 1994 (Shariff, 1999). As the relevant data have been collected from about 37,000 ever-married women, it is possible to derive from the survey data reasonably stable estimates of maternal mortality for broad geographical regions and by socio-economic characteristics.

The second objective of this paper is to update the estimates of maternal mortality derived using the indirect procedure developed by Bhat *et al.* (1995). The original paper carried estimates of maternal mortality for India and its major states for the period 1982-86 using the data from India's Sample Registration System. In this paper, estimates are presented using this method for the period 1987-96. An attempt is also made to compare the estimates derived from the two indirect methods to judge their relative performance.

Indirect Estimates from Sisterhood Method

The HDPS covered about 33,000 rural households spread over 1,765 villages and 195 districts in 16 major states of India. Table 1 shows the application of the sisterhood method to the all-India data for rural areas. Nearly 37,000 women 15-49 years at the time of the survey have reported on an average 1.9 married sisters. Among them, 2,671 sisters were not alive at the time of the survey (i.e., 4.3 percent). The data available for 2,102 dead sisters indicate that 32 percent died around the time of childbirth. From this information, following the procedure developed by Graham *et al.*, the lifetime probability of dying from maternal causes for a female baby has been calculated as 2.5 percent (see Table 1). This estimate of lifetime probability refers to, on an average, 11.8 years before the survey, or roughly to 1982. For computing the more conventional measure of maternal mortality rate, an estimate of average number of children born to women in their lifetime (i.e., total fertility rate) is required. The most convenient approximation to this is the average number of live births reported by the surveyed women of age 40-49 years. According to HDPS, the average number of children ever born to women of the age interval 40-49 was 4.4. But the reported sex ratio of 118 males for 100 female births suggests that women had underreported the number of daughters born to them. If a correction is made by assuming a sex ratio at birth of 105, the total lifetime births to an average woman rises to 4.7. This figure is close to the total fertility rate of 4.9 reported by the Sample Registration System for rural areas of India in 1982. With 4.7 as the estimate of lifetime births, the sisterhood data from the HDPS imply a maternal mortality rate of 544 per 100,000 live births in rural areas of India in 1982. With 4.9 as the estimate of TFR, the data imply a maternal mortality ratio of 518. As only the survey can provide estimates of lifetime births for various population groups, for the sake of comparability, the former could be adopted as the estimate of MMR for rural India from HDPS.

Table 1. Estimating maternal mortality ratio from sisterhood method, All India Rural, HDPS 1994

Age group	No. of Respondents	No. of ever-married sisters	No. of married sisters dead	Percent died of maternal causes *	No. of maternal deaths	Mean no. of ever-married sisters	Adjusted ever-married Sisters **	Adjusted Factor	Sister units of exposure to risk	Lifetime probability of dying	Reference time
15-19	2537	2975	58	40.2	23	1.17	4746	0.107	508	0.046	5.7
20-24	6763	9455	246	34.0	84	1.40	12653	0.206	2606	0.032	6.8
25-29	7684	12738	408	37.8	154	1.66	14376	0.343	4931	0.031	8.1
30-34	6429	11832	458	36.5	167	1.84	11832	0.503	5951	0.028	9.7
35-39	6110	11720	594	28.1	167	1.92	11720	0.664	7782	0.021	11.7
40-44	4181	7695	479	29.0	139	1.84	7695	0.802	6172	0.023	14.3
45-49	3289	6187	428	25.5	109	1.88	6187	0.900	5568	0.020	17.5
Total	36993	62602	2671	31.5	843	1.87	69209		33519	0.025	11.8

TFR = 4.7 MMR = 544

* As per data available for 2,102 dead sisters.

** Figures for ages 15-29 are adjusted on the assumption that the average number of sisters reaching the reproductive period would be the same as in the age groups 30+, i.e., 1.87.

As mentioned earlier, using a completely different approach, Bhat *et al.* (1995) had estimated a maternal mortality rate for India and its major states during 1982-86. Their estimate for rural India was 638 per 100,000 live births. Clearly, the sisterhood data from the HDPS implies lower levels of maternal mortality than the indirect procedure using the sex-differentials in mortality in reproductive ages. A recent review of the sisterhood method had indicated that it could generally be underestimating maternal mortality because of omission of events in the survey (Stanton *et al.*, 2000). On the other hand, some have claimed that the indirect version of the method could be overestimating maternal mortality owing to non-applicability of model assumptions (Garenne and Friedberg, 1997). A state or regional comparison could provide more clues to the source of this discrepancy.

Estimates for geographical zones

As one runs into problems of sample size in making state-specific estimates from the HDPS, estimates of maternal mortality were derived for five geographical zones (see Table 2). Maternal mortality levels were lowest in the north-western zone comprising of states of Punjab, Haryana and Himachal Pradesh (289), and in south India (383). On the other hand, maternal mortality was over 600 in the eastern zone (Assam, West Bengal and North-Eastern states) and north-central zone (Uttar Pradesh and Bihar). The accuracy of this geographical variation in maternal mortality can be checked by comparing with the estimates derived indirectly from sex-differentials in adult mortality. As the latter estimates were derived for the major states, they were pooled to produce the regional estimates (see Table 2).

Table 2. Indirect estimates of maternal mortality ratios from sisterhood method by geographical zones and their comparison with those derived from sex differentials in mortality

Zones **	Estimates for rural areas from HDPS, 1994						Indirect estimates of MMR by Bhat et. al for 1982-86 *
	Number of respondents aged 15-49	Lifetime Births for Women aged 40-49	Percentage of maternal deaths	Lifetime risk of maternal death	Time reference	Maternal mortality ratio	
North-West	4708	4.8	27.5	0.014	11.7	289	346
North-Central	7007	5.2	29.5	0.031	11.8	612	879
South-Central	9298	5.0	35.1	0.029	11.6	588	614
East	3034	5.0	36.7	0.032	11.4	636	709
West	4752	4.1	23.8	0.019	11.9	471	414
South	8193	4.0	29.5	0.015	12.1	383	379
All -India	36993	4.7	31.5	0.025	11.8	544	580 (638) @

* Pooled, state-specific estimates for the total area.

@ Estimate for rural India.

** North-West: Punjab, Haryana, Himachal Pradesh; North-Central: Uttar Pradesh and Bihar; South Central: Rajasthan, Madhya Pradesh and Orissa; East: Assam, West Bengal and North-eastern states; West: Maharashtra and Gujarat; South: Andhra Pradesh, Karnataka, Tamil Nadu and Kerala.

While making the comparison, it should also be noted that the indirect estimates derived from sex differentials in adult mortality are for all areas while the estimates from the HDPS are for only rural areas of the respective zones. Nonetheless, but for one zone, there is a reasonably good correspondence between the estimates from the two sources on the regional variations in maternal mortality in India. Both sources suggest that maternal mortality is lowest in the north-western zone and relatively high in the eastern zone. The discrepancy between the two sources is mainly in the north-central zone where the estimates from sisterhood data from HDPS is lower than the indirect estimate based on sex-differentials in mortality by as much as 30 percent (612 and 879, respectively). As this region is known to be one of the most backward in India, the estimate from the HDPS seems to be a bit too low.

Estimates by background characteristics of respondents

An advantage of the sisterhood data is that it can be used to study variation in maternal mortality among population subgroups. However, one should be cautious in interpreting the results because the personal characteristics used for the classification are not of the women who were exposed to the risk of maternal deaths, but those of their sisters. Nonetheless, this information could be of use because it is unlikely that the sisters would have come from very different socio-economic background. In Table 3 estimates of maternal mortality are presented according to respondent characteristics such as caste, religion education, level of poverty and developmental

level of the village.

Table 3. Indirect estimates of maternal mortality ratios from sisterhood method by background characteristics of respondents, Rural India, HDPS 1994

Background Characteristics	Number of respondents age 15-49	Lifetime births for women age 40-49	Percentage of maternal deaths	Lifetime risk of maternal death	Time reference	Maternal mortality ratio
Caste						
Scheduled caste	8368	4.9	32.1	0.029	11.8	584
Schedule tribe	4545	4.6	28.2	0.030	11.6	652
Others	24079	4.6	32.0	0.024	11.8	516
Religion						
Hindus	31088	4.6	31.7	0.026	11.8	573
Muslims	3654	5.8	32.0	0.022	11.7	384
Others	2250	3.9	25.9	0.017	12.0	428
Education						
Illiterate	24580	4.8	31.8	0.027	12.1	574
Primary or less	6694	4.3	30.4	0.021	11.6	492
Middle or more	5718	3.8	30.4	0.018	10.5	484
Poverty Group						
Poor	7334	5.3	31.3	0.029	11.6	555
Lower-middle	7823	5.1	31.2	0.022	11.5	439
Upper- middle	15688	4.5	32.7	0.027	11.8	611
Non poor	6147	4.2	28.9	0.020	12.3	484
Village development						
Low	10904	5.1	31.3	0.032	11.7	646
Medium	14640	4.7	30.4	0.023	11.8	501
High	11448	4.3	31.4	0.021	11.8	488
All India	36993	4.7	31.5	0.025	11.8	544

Not surprisingly, the maternal mortality ratio is estimated to be quite high among women of scheduled tribes (652) and scheduled castes (584) compared to women of other castes (516). Among the religious groups, Hindus have higher maternal mortality (573) than either Muslims (384) or women of other religions (428). The unusually low maternal mortality among Muslim women is a puzzle; it could partly be due to the fact that they generally live in larger villages having better access to emergency obstetric care.

As one would expect, maternal mortality is higher among illiterates (574) compared to those who had gone to primary school (492) or passed middle school (484). The educational differential in maternal mortality appears somewhat muted because some of the sisters of illiterate respondents could have gone to school, or the visa-versa.

Table 3 also shows the estimates of maternal mortality by poverty level of respondents. Using the 'head-count' method, the survey respondents have been classified as belonging to (i) lower segment below the poverty line, (ii) upper segment below the poverty line, (iii) lower

segment above the poverty line and (iv) upper segment above the poverty line (see Shariff, 1999). Surprisingly, the data do not disclose a strong relationship between poverty level of respondents and maternal mortality of their sisters. Those who are at the bottom most segment do have marginally higher maternal mortality than those at the top (555 and 484 respectively). But it is the two intermediary classes that show the lowest (439) and the highest (611) maternal mortality ratios. This could happen if the respondents and their sisters did not belong to the same poverty class.

It is interesting to explore whether access to health care has an influence on the level of maternal mortality. The sample villages of the HDPS have been classified as less developed, moderately developed and well developed based on an index measuring the infrastructure and amenities in the village (see Shariff, 1999). Table 6 shows the estimates of maternal mortality for these three types of villages. It can be seen that the less developed villages have significantly higher maternal mortality (646) than either moderately or well-developed villages (501 and 488, respectively). Thus access to health care does appear to matter for the level of maternal mortality.

Indirect Estimates from Sex Differentials in Adult Mortality

The method

As the method of estimating maternal mortality from sex-differentials in adult mortality (SDAM) has not come into wide use, it is useful to briefly review its methodology. The SDAM method makes use of the following exact identity (for proof, see Bhat *et al.*, 1995):

$$\frac{d_{xf}}{d_{xm}} = g_x + \omega \frac{w_x f_x}{d_{xm}} \quad (1)$$

where d_{xf} and d_{xm} are death rates at age x for females and males, respectively; g_x is the sex ratio of death rates at age x in the absence of maternal mortality; f_x is the fertility rate at age x , w_x is the ratio of maternal mortality at age x relative to its level at age 20-24; ω is the maternal mortality ratio at age 20-24. Among the various terms in the equation, d_{xf} , d_{xm} and f_x are observed while g_x and w_x and ω are unknowns. However, if a mathematical form could be assumed for g_x , and the age pattern of maternal mortality (w_x) can be taken from another source, or 'borrowed' from another population with roughly similar levels of maternal mortality, the level of maternal mortality at age 20-24 (ω) can be estimated from eq. (1) by ordinary least squares regression of the data for age groups 10-14 to 50-54 years.

Once ω is estimated, the conventional estimate of maternal mortality ratio (MMR) for age 15-49 can be derived from the relation

$$MMR = \omega \frac{\sum_x w_x f_x N_x}{\sum_x f_x N_x} \quad (2)$$

where N_x is the number (or proportion) of females at age x . Bhat *et al.* have suggested that while using regression to estimate ω , the following alternative form for eq. (1) should also be tried out:

$$d_{xf} = g_x d_{xm} + \omega w_x f_x \quad (3)$$

They have further suggested that if OLS regressions of these two alternative forms give substantially different results, the forms assumed for g_x or w_x should be changed, or data for age groups at the beginning or end of the reproductive span should be discarded so that the two regressions give similar estimates of ω .

The functional form for g_x should be carefully chosen. Experience suggests that either a quadratic form, or a liner function with a dummy for the age interval 20-24 to capture the kink occasionally found at this age, is best suited. But if data on age-specific mortality and fertility rates are available for a large number of population subgroups (or time periods), the form of g_x can be estimated by a pooled, dummy-variable regression wherein g_x is specified as age effects. If such estimate of g_x is available, then the estimating equation for maternal mortality can be rewritten as

$$\frac{d_{xf}}{d_{xm}} - \hat{g}_x = C + \omega \frac{w_x f_x}{d_{xm}} \quad (4)$$

or,

$$d_{xf} - \hat{g}_x d_{xm} = C d_{xm} + \omega w_x f_x \quad (5)$$

where C is the estimated sex ratio of death rates at age 10-14 years. Using data for major states of India for 1982-86 from the Sample Registration System (SRS) Bhat *et al.* have estimated g_x values for India. The estimated g_x values vary slightly under the two alternate specifications of the regression model. Although using the same procedures it is possible to make separate estimates of g_x for the period after 1986, for the sake of maintaining comparability with maternal mortality estimates for earlier periods, I have used the same values of g_x as estimated by Bhat *et al.* from the data for 1982-86. For w_x , Bhat *et al.* had used the age pattern of maternal mortality implied by two field studies conducted in the Matlab area of Bangladesh (Chen *et al.*, 1974; Koenig *et al.*, 1988). The same pattern has been used in the present application too. The standard values of g_x and w_x employed in the application are reproduced below:

Function	10-14	15-19	20-24	25-29	30-34	35-39	40-44	45-49	50-54
g_x									
Set A	0.00	-0.05	-0.15	-0.24	-0.33	-0.36	-0.37	-0.44	-0.39
Set B	0.00	-0.03	-0.11	-0.22	-0.29	-0.33	-0.33	-0.41	-0.36
w_x	0.0	1.9	1.0	1.2	1.4	1.8	2.0	2.4	0.0

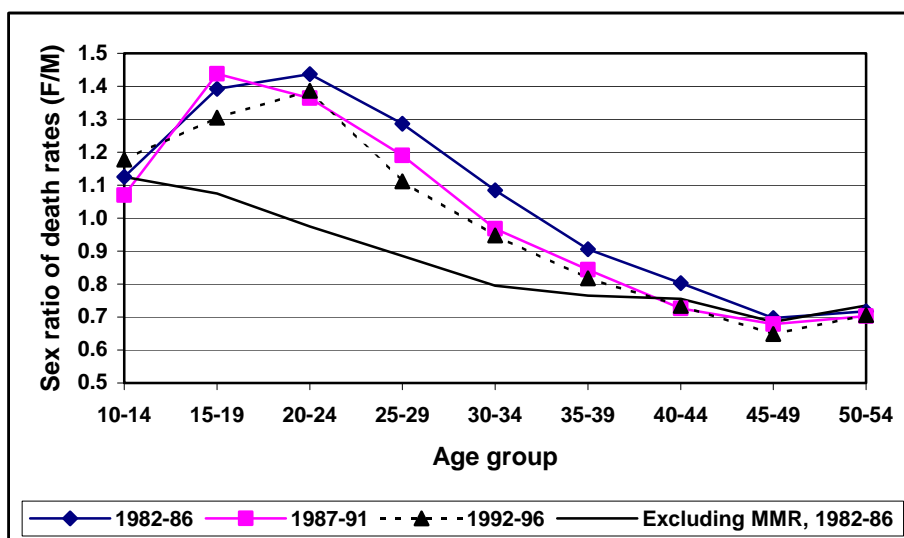
Note: Set A is used with eq.(4), and set B is used with eq. (5).

Maternal mortality during 1987-96

For India and its major states the Sample Registration System provides annual estimates of age-specific fertility and mortality rates. Indirect assessments suggest that about 90-95 percent of

the vital events get registered in the system (see Bhat 2000). Figure 1 shows the ratios of female-male mortality rates by age for all-India during 1982-86, 1987-91 and 1992-96. The sex ratios of mortality rates at reproductive ages show a clear fall overtime, suggesting declines in levels of maternal mortality. For establishing the levels and changes in maternal mortality more precisely, the SDAM method has been applied to the SRS data for 1987-91 and 1992-96 in the case of all-India, and for 1987-96 in the case of major states of India. As sampling errors in the SRS age-specific rates could be quite large (see, India, Registrar General 1990), a five-year time interval has been used in the case of all-India while a ten-year time interval has been adopted in the case of major states.

Figure 1. Ratio of death rate of females to that of males at reproductive ages, All India, 1982-86, 1987-91 and 1992-96



Source: Annual reports of the Sample Registration System.

The SDAM method was first applied to all-India data for 1987-91 and 1992-96 by assuming a quadratic form for g_x . Table 4 shows the parameters of the model estimated from OLS regressions. In all the regressions, the coefficient of the square of age variable (x^2) is statistically significant, showing that g_x is non-linear with age. The implied levels of maternal mortality (w) are slightly different under the two models, but they fall within the confidence interval for either of the estimates. If estimates under the two model specifications are averaged, the results imply that maternal mortality ratio was 641 during 1987-91 and 511 during 1992-96.

The SDAM method was also applied using the g_x function estimated for India by Bhat et al. from their analysis of state-level data for 1982-86. Table 5 shows the results of these regressions for India by rural-urban residence, and also for 15 major states. The slope coefficient, giving the estimate of maternal mortality at age 20-24 (w) is less than zero in the case of Punjab and Kerala. This may be because maternal mortality in these states has reached such low levels that its effects on sex differentials in mortality can no more be separated from other 'noises' in the data. The two model specifications give more or less similar estimates of maternal mortality in most cases. But in Haryana and Uttar Pradesh there were some significant differences. In these cases, the estimates from the two model specifications were brought closer

by dropping data for the age groups 45-49 and 50-54. The standard errors of maternal mortality estimates show that errors in state-specific estimates are 3-4 times larger than in the estimate for all-India.

Table 4. Estimates of model parameters and maternal mortality ratios under two alternate model specifications that use quadratic form for g_x , India, 1987-91 and 1992-96

Year	Model	Coefficients of g_x function			MMR for age 20-24 (ω)	R ²	N	MMR for age 15-49
		Constant	x	X ²				
1987-91	Eq. (1)	1.087	-0.150 (0.018)	0.013 (0.002)	502 (42)	0.992	9	663
	Eq. (3) *	1.107	-0.145 (0.015)	0.012 (0.001)	468 (42)	0.999	9	618
1992-96	Eq. (1)	1.176	-0.158 (0.021)	0.012 (0.003)	387 (48)	0.988	9	500
	Eq. (3) *	1.179	-0.167 (0.032)	0.013 (0.003)	405 (84)	0.997	9	523

Note: Figures in parentheses are standard errors of the estimates.

* The regression model has no intercept, but the constant of the g_x function is the coefficient of death rates of males.

For all- India, the use of the standard g_x function results in substantially lower estimate of maternal mortality than using a quadratic function. For example, for 1987-91, with the sex ratio of mortality as dependant variable, the estimated maternal mortality ratio at age 20-24 is 502 if the quadratic form is used, whereas it is 410 if the Indian standard g_x function is employed. This is because, though the standard g_x function is non-linear, it is not exactly quadratic. As the true form of g_x is unknown, it is better to use a set that has been estimated from the data themselves. As the Indian standard g_x function is empirically derived, the estimates of maternal mortality derived from it are to be preferred. To some extent, regression results also point towards the same, as the standard error of ω is lower when the Indian standard is used instead of the quadratic function. However, the India standard is unlikely to have wider geographical applicability.

Table 6 shows the conventional estimates of maternal mortality ratio and maternal death rate (i.e., maternal mortality rate) derived from the regression results presented in Table 5 for all-India by rural-urban residence. The estimates from the two model specifications have been averaged. The results indicate that at the all-India level, the maternal mortality ratio has fallen from 580 per 100,000 live births around 1984 to 519 in 1989, and further to 440 in 1994. In the meanwhile, in the maternal death rate has fallen from 82 per 100,000 women age 15-49 in 1984 to 66 in 1989 and 51 in 1994. Maternal mortality has declined in both rural and urban areas. In rural areas the maternal mortality ratio has declined from 638 in 1984 to 498 in 1994. In urban areas, the estimated trend is somewhat uneven owing to larger sampling errors in the data; the maternal mortality ratio appears to have fallen from a level just below 400 in 1984 to near 300 in 1994.

Table 5. Model parameter estimates and regression statistics by state and residence, under two alternate model specifications that use the Indian standard gx function, 1987-96

State/Area	Time Period	Model A (eq. 4)			Model B (eq. 5)		
		MMR for age 20-24 (ω)	Standard error of ω	Model R^2	MMR for age 20-24 (ω)	Standard error of ω	Model R^2 @
Andhra Pradesh	1987-96	221	43	0.793	205	43	1.000
Assam	1987-96	810	103	0.898	711	90	0.999
Bihar	1987-96	386	52	0.888	362	53	0.999
Gujarat	1987-96	492	24	0.984	483	36	1.000
Haryana *	1987-96	382	122	0.661	373	87	0.994
Karnataka	1987-96	387	90	0.727	363	90	0.998
Kerala	1987-96	-48	83	0.047	-23	56	0.999
Madhya Pradesh	1987-96	532	35	0.971	524	33	1.000
Maharashtra	1987-96	313	84	0.665	290	79	0.998
Orissa	1987-96	465	70	0.863	467	64	0.999
Punjab	1987-96	-279	163	0.294	-448	231	0.983
Rajasthan	1987-96	399	88	0.746	439	80	0.997
Tamil Nadu	1987-96	181	83	0.406	135	83	0.999
Uttar Pradesh**	1987-96	579	57	0.946	517	65	0.998
West Bengal	1987-96	360	82	0.735	334	69	0.999
India, Total	1987-91	410	24	0.977	376	32	1.000
	1992-96	341	38	0.922	340	30	1.000
India, Rural	1987-91	431	18	0.988	408	28	1.000
	1992-96	385	35	0.946	374	24	1.000
India, Urban	1987-91	261	62	0.719	208	82	0.998
	1992-96	243	104	0.440	269	95	0.998

@ R^2 of regression through the origin, computed without the mean correction.

* Model fitted to data for age intervals 10-14 to 40-44 years.

** Model fitted to data for age intervals 10-14 to 45-49 years.

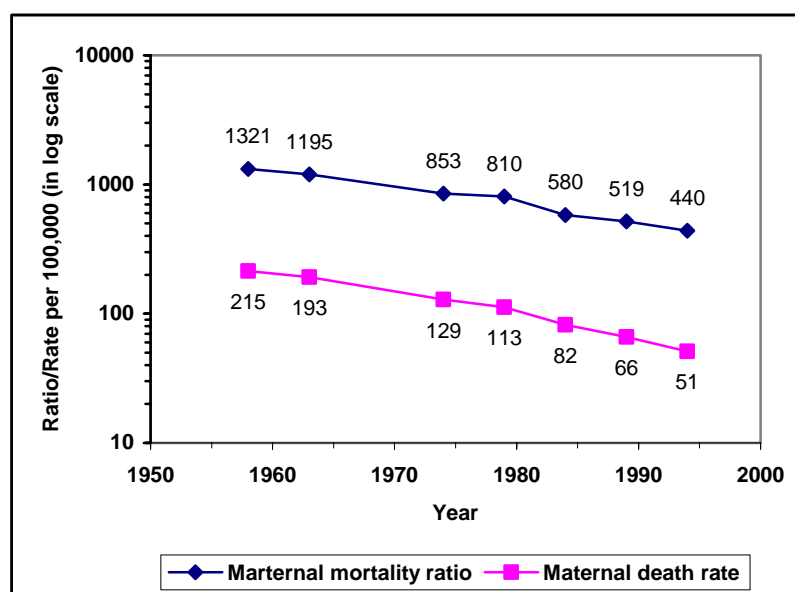
Table 6. Indirect estimates of maternal mortality for India derived from sex differentials in adult mortality, by rural-urban residence, 1982-86, 1987-91 and 1992-96

Area	Time period	Maternal mortality ratio (per 100,000 live births)	Maternal death rate (per 100,000 women)	Percentage of maternal deaths at age 15-49 years
India, Total	1982-86	580	82	21
	1987-91	519	66	19
	1992-96	440	51	17
India, Rural	1982-86	638	96	22
	1987-91	563	76	20
	1992-96	498	62	18
India, Urban	1982-86	389	44	17
	1987-91	299	29	12
	1992-96	321	27	13

Source : Bhat *et al.* (1995).

Figure 2 shows a more long-term trend in maternal mortality for India. All the estimates shown are derived from the sex differentials in mortality using the data from the National Sample Surveys and the Sample Registration System (see Bhat *et al.*, 1995). As the graph shows, maternal mortality ratio has been steadily falling in India from a level of near 1,300 in the late 1950s. It was between 800 and 900 in the 1970s, 500-600 in the 1980s, and 400-500 in the 1990s. The maternal death rate has declined in the meanwhile from a level slightly over 200 in the late 1950s, to around 120 in the 1970s, 75 in the 1980s and 50 in the 1990s. The figure suggests acceleration in the decline of maternal death rate since the 1970s, which could be related to the declines in fertility.

Figure 2. Estimated trends in measures of maternal mortality for India, 1955-95



Source: Bhat *et al.*, (1995).

According to the methodology adopted by WHO and UNICEF (1996), maternal mortality in India was 570 in 1990. On the other hand, a figure near 500 would be consistent with the estimates proposed here. It should be noted that the WHO-UNICEF methodology is based on an international regression line relating the maternal mortality with proportion of births assisted by trained persons and general fertility rate. From such a proximate-determinant one expects to provide only an approximate estimate in a given case. Although regression is also the basis of the present estimates, the regression model is based on an exact identity and, in each application, its parameters are estimated from the data. But the estimates from this method could be sensitive to the form assumed for g_x . If a quadratic form is assumed, the resulting estimates of maternal mortality are higher, and close to the WHO-UNICEF estimate. But it is theoretically more appropriate to use empirically derived values of this function. The estimates of maternal mortality so derived also have smaller standard errors.

Table 7. Indirect estimates of maternal mortality derived from sex differentials in adult mortality, by state, 1982-86 and 1987-96

State	1987-96			1982-86			
	Maternal mortality rate *	Maternal death rate **	Percentage of maternal deaths @	Maternal mortality rate *	Maternal death rate **	Percentage of maternal deaths @	
Andhra Pradesh	283	29	9	394	50	13	
Assam	984	119	27	1068	145	27	
Bihar	513	76	18	813	139	27	
Gujarat	596	67	23	373	51	15	
Haryana	472	65	25	494	79	25	
Karnataka	480	50	19	439	53	17	
Kerala	@@	@@	@@	247	21	15	
Madhya Pradesh	700	104	26	507	84	21	
Maharashtra	380	40	16	439	54	19	
Orissa	597	67	17	844	113	24	
Punjab	@@	@@	@@	207	25	11	
Rajasthan	580	86	28	627	110	29	
Tamil Nadu	195	15	5	372	38	10	
Uttar Pradesh	737	118	30	920	160	32	
West Bengal	458	49	17	561	73	20	
India	Total	479	58	18	580	82	21
	Rural	528	68	19	638	96	22
	Urban	311	28	13	389	44	17

* Rate per 100,000 live births.

** Rate per 100,000 females in the age interval 15-49 years.

@ In the age interval 15-49.

@@ Maternal mortality is so low that it is not possible to estimate its level from sex differentials in mortality at reproductive ages.

Turning to state-level variations, Table 7 shows the estimates of maternal mortality for 15 major states of India during 1987-96. For comparison, the table also shows the estimates derived by Bhat *et al.* for 1982-86. The suggested variation in maternal mortality by state during 1987-96 is very similar to that implied by the estimates made earlier for 1982-86. As in the earlier period, among the major states, Assam has the highest maternal mortality ratio (984). It is followed by Uttar Pradesh (737) and Madhya Pradesh (700). The lowest level of maternal mortality is estimated for Tamil Nadu (195). But it is perhaps even lower in Punjab and Kerala, where it has become difficult to discern the exact levels from the data on sex-differentials in adult mortality for a sample population. All but three states show sizeable reductions in maternal mortality between 1982-86 and 1987-96. In Madhya Pradesh, Gujarat and Karnataka the estimates suggest an increase in the level of maternal mortality. Among the three, the suggested increase is quite large in Madhya Pradesh and Gujarat. Perhaps in these two states, the maternal mortality estimates for 1982-86 were underestimates as suggested by a comparison with their levels of infant mortality (see Bhat *et al.*, 1986: 226). Given the sampling errors in the state-level estimates, maternal mortality estimates for 1987-96 should generally be more reliable than for 1982-86 as the former is based on data averaged for 10-years rather than for 5-years. One exception to this could be Bihar where the quality of the SRS data may have deteriorated in the late 1980s.

Summary and Conclusions

The paper has presented estimates of maternal mortality for India from two indirect procedures, the sisterhood method and regression method involving sex differentials in adult mortality (SDAM). The application of sisterhood method to the data from the Human Development Profile Survey of 1994 yields an estimate of maternal mortality of 544 deaths per 100,000 births in rural India for a period roughly 12 years before the survey. It also shows that maternal mortality ratio was more than 600 in east and north-central India, while it was between 300 to 400 in north-western and southern India. The survey results also show that maternal mortality levels were high among scheduled tribes and scheduled castes, and surprisingly low among Muslims. The level of maternal mortality is also strongly related to amenities and infrastructure available in the village. However, its relationships with poverty and educational levels of respondents are found to be weak perhaps because the characteristics of respondents were not ideal surrogates for their sister's attributes.

For India, the level of maternal mortality implied by the sisterhood method is found to be lower than the estimate derived from the sex differentials in adult mortality. If the sisterhood data implied a maternal mortality ratio of 544 for rural India in the early 1980s, the SDAM method suggested its level as 638. A regional comparison showed that the sisterhood estimate is particularly low for north central zone, comprising states of Uttar Pradesh and Bihar. Otherwise the two sources suggest very similar regional variations in maternal mortality.

The SDAM method can also be applied to more recent data. The application of this method to data from the Sample Registration System for 1987-91 and 1991-96 shows that maternal mortality ratio for India as a whole was 519 around 1989 and 440 around 1994. These estimates are not inconsistent with the findings of the National Family Health Survey, if the estimates from the two rounds are averaged (479), or with the direct estimates from the SRS for 1997-98 (407-408). But they are lower than the WHO-UNICEF estimate of 570 for 1990. As sisterhood estimates of maternal mortality are even lower than those of the SDAM method, the WHO-UNICEF estimate appears a bit too high.

The sisterhood method has been criticised for its lack of precision essential for a trend analysis (Stanton *et. al.*, 2000). In the case of India, by applying the SDAM method to the past data, it has been possible to make reasonably accurate estimate of the trend in maternal mortality. The results suggest that maternal mortality was steadily falling in India since the late 1950s. It has continued to fall during the 1980s and 1990s. But the method's continued applicability in future is under some cloud. The rising incidence of HIV/AIDS could drastically change the sex differentials in mortality in reproductive ages. As the application of the method for Punjab and Kerala has demonstrated, when maternal mortality falls to very low levels, it may not be possible to isolate its effects on sex differentials in mortality from other disturbances.

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