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Demographic Estimates for Post-Independence India : A New Integration

Introduction

MANY statistical publications of Government of India contain a set of estimates of vital rates for the country, which are quoted freely in the press and taken uncritically in academic circles while commending or criticising or censuring our achievements since Independence. But have we got the facts correct? Are the official estimates consistent with the patterns of population growth recorded by the decennial censuses, and trace correctly the progress of demographic transition in the country? As pitfalls in our data are many, demographic scenarios can be very different depending on how one chooses to correct them. As new information may not always be consistent with the preceding one, demographic analysis in India must be a continuous activity of updating and integrating estimates. The celebration of the country's 50th year of Independence provides an opportune time to reflect and correct, if necessary, the historical record on the basis of information now become available.

The main objective of this paper is to put forth an integrated set of estimates of fertility and mortality for the country as a whole, that are based upon careful weighing of evidence from diverse sources, and employing methods that are least susceptible to typical forms of data errors. The period under consideration is from 1941 to 1991.1 have begun the review a few years a head of Independence to establish benchmarks for measuring the progress made since then, but have ended somewhat prematurely at 1991 owing to the want of census data to confirm the trends indicated by other sources.

While assessing demographic trends during the post-Independence era, one encounters a great data divide at around 1971. For the pre-1971 period, demographic estimation would have to rely almost exclusively on information on growth and age-sex structure of the population provided by the population censuses. The subsequent period is however inundated

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with information from many quarters: the Sample Registration System and other demographic sample surveys provide more recent information than the census, and the contents of census information itself has become richer owing to the inclusion of questions on fertility and child mortality. The problem has thus become one of choosing from among the many possible contenders.

The availability of overlapping sources of information for the post-1971 period has also implications for demographic estimation for the pre-1971 period. In particular, it contributes to the understanding of the nature of biases present in demographic estimates derived from the age-sex distribution of the population and in revising the rules of thumb devised to assist estimation in the presence of distortions in the age data. We are thus in an advantageous position than the analysts of the past in making estimates for this period too.

Mortality estimates are first taken up for review as they form inputs in estimating fertility measures from age-sex distributions of the population. As they often are derived from different data sources, estimates of child mortality and adult mortality are separately discussed before reviewing the estimates of life expectancy at birth.

Levels of Adult Mortality

Before the inception of the Sample Registration System, the sole bias for the measurement of mortality level was through the comparison of sizes of cohorts enumerated at different ages at periodic intervals. Although several methods were employed to achieve the same objective, all suffered from the basic flaws in the census data - changing enumeration completeness, age misreporting and migration - to varying degrees. As no independent information of reliable quality was available, there was no way of knowing the merits of the proposed estimates. The emergence of the SRS in the early 1970s provided for the first time a ready bias for making such an evaluation. Therefore I shall first consider the estimates of adult mortality for the decades 1971 -81 and 1981 -91.

Estimates for Periods after 1971

Aihong the many methods proposed for the estimation of mortality from intercensal analyses, theoretically most rigorous is the 'variable-r' methodology proposed by Preston and Bennett (1983). It makes no assumption of stability and use of the model life tables is kept to a minimum. It can further be improved by using a cohort-wise interpolation in place of the age-wise interpolation while estimating the intercensal person-years of exposure (Coale 1984). In populations where gross distortions are common in reported ages, the cohort interpolation has the additional advantage of providing some in-built smoothing to age distributions (see Bhat 1995). I have therefore applied this method to the data from the censuses of 1971, 1981 and 1991. The interpolation factors used in estimating the person-years lived were taken from Bhat (1987), and reproduced here in Table Al of the appendix[1].

[1] A programme in Lotus 1-2-3 that implements the method can be obtained from the author by sending Rs. 100/- by M. O. or demand draft if the floppy is to be mailed within India. For users outside India the price is U.S. $ 10, including postage.

To estimate a complete life table from this procedure information on the number of births in the intercensal period is necessary. A set of birth rate estimates presented later in this paper has been used to generate the whole life table. Here, however, we are concerned with the estimates for ages above 10 which are insensitive to the assumptions about the birth rate.

The estimates of life expectancies at various ages computed using the variable-r methodology are shown in Table A2 of the appendix. They make an interesting comparison with the estimates of life expectancies derived from the SRS data for the corresponding period[2]. Figure 1 shows the ratios of life expectancy estimates from the two sources for 1971-81 and 1981-91. If the estimates from the two sources are identical at all ages, the ratios should fall in a stright line. The plots shown in Figure 1 suggest a systematic deviation from the expect value of 1. The ratios of male life expectancies are close to 1 at ages under 10; thereafter the ratios show a rising trend up to age 25, then a fall until age 40 and a sharp rise after 45 years. Since the male ratios of 1971 -81 and 1981-91 show very similar patterns of deviation, they must be arising out of typical patterns of age misstatements in the census data. Nonetheless, as estimates at younger ages are close to one, the overall level of mortality suggested by the census analysis is almost identical to the SRS estimates. The slightly higher ratios of 1971 -81 compared to those of 1981 -91 suggest either (i) a somewhat higher level of underestimation of mortality in 1981 -91 compared to 1971 -81 in the SRS, or (ii) a higher level of underenumeration of males in 1971 and 1991 censuses compared with the 1981 census. This issue will be resolved shortly.

Since at most adult ages the census-based estimates of male life expectancy are higher than the SRS estimates—despite being nearly equal at younger ages - estimates for any arbitrarily chosen age range cannot be averaged to give an unbiased estimate of adult mortality level. However, the bias in the census-based estimate can be kept to a minimum by averaging the life expectancy estimates for ages 15 and 10 years, since for both the decades, the former is slightly higher than the SRS estimate while the latter is lower. In order to convert this average to a more commonly used index of adult mortality, such as life expectancy at age 5 *(e_s)*, I suggest, as Preston and Bennett do, taking the ratios of e_{10} and e_{15} to the corresponding values of a standard life table and multiplying the average of the ratios by the value of e_s of the standard table[3].

The life expectancy ratios of females presented in Figure 1 are far more puzzling. They do show a systematic deviation by age, similar to those of males, but the census estimates of life expectancies at all ages are higher than the SRS estimates for 1971-81 while being substantially lower for 1981-91. This could be suggesting either that death rates from the SRS had significant downward bias in 1981 -91 while being biassed in the opposite direction in 1971 -81, or that the underenumeration of females was significantly larger in the 1971 and 1991 censuses compared with the 1981 censuses.

[1] For this purpose, I have computed abridged life tables for 1971 -81 and 1981-91 from the annual SRS rates using an identical methodology. They differ only marginally from the official tables for the quinquennial periods. [3] However, to estimate e_{10} using the cohort interpolation method, it is necessary to have an estimate of live births in the intercensal period when the censuses are 10-year apart. Fortunately, the e_n estimate is not very sensitive to the assumed level of the crude birth rate. As such, only an approximate value is needed. Since the estimate of $e_{\%}$ from the method is sensitive to the choice of the crude birth rate, it is not used in arriving at an estimate of the adult mortality level.

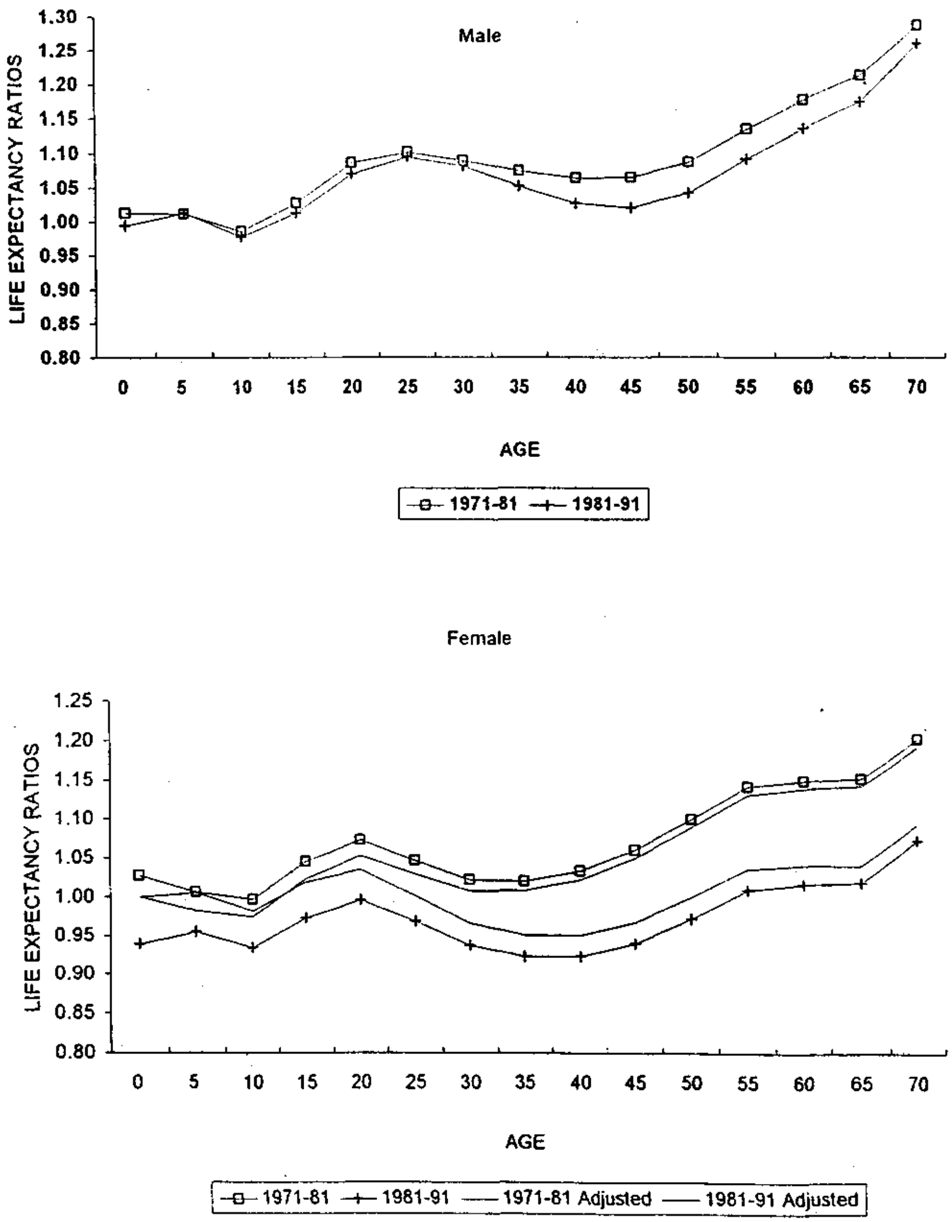


Fig. 1. Ratios of census-based estimates of life expectancy to that of SRS

Among the two propositions, the census underenumeration hypothesis appears far more credible for several reasons: First, a drastic change in the direction of SRS completeness between 1971 -81 and 1981-91 seems quite unlikely; if at all it changed, evidence favours an improvement in the SRS completeness during 1981-91 rather than its deterioration. Second, any changes in the SRS completeness would have affected both the sexes about equally. Ihird, the census-based estimates indicate a rise hi mortality levels of women between 1971-81 and 1981-91 while a large drop has been registered for the preceding period. For example, the census-based estimates of e_{l5} show an increase of 6.1 years between 1961-71 and 1971-81 and a drop of 0.7 years during the next ten years (see Table Al). Finally, there is independent evidence to suggest greater level of underenumeration in the 1971 census (Visarial971).

To estimate the relative levels of female underenumeration in 1971 and 1991, the respective census totals are to be raised relative to that of 1981 so that the female life expectancy ratios of 1971-81 and 1981-91 are, if not identical, nearest to each other. As the ratios of the two intercensal periods come nearest to the expected value of one (signifying that the SRS and census estimates are identical) at younger ages, it is reasonable to assume that, in a situation where the two censuses are equally complete, the mean of e_{lo} and e_{l5} values would give an unbiased estimate of the level adult mortality among females too. Accordingly, we raised the 1991 census population totals by age such that the mean of the implied e_{w} and e_{l5} values would be identical to the mean of the SRS estimates for the period 1981 -91. Similarly the 1971 census totals were raised to match the SRS estimates of 1971-81. In doing so, the levels of underenumeration was assumed to vary from age to age; for this purpose we made use of the age-pattern of underenumeration suggested by the post enumeration checks of the last three censuses[4].

These computations suggested an excessive undercount of 1.8 percent among females in 1991 census and 0.8 percent in 1971, both in relation to the level of undercount in the 1981 census. The number of females that went additionally missing in 1991 on account of underenumeration was put at 7 million, of which 1.2 million (17.5%) were in the age group 0-4 years. In 1971 the numbers were 2.1 million and 0.44 million (21 %), respectively. Similar computations were made for the males too; but the indicated changes in the level of underenumeration were marginal: 0.2% in 1991 and 0.3% in 1971.

It was further observed that even after raising tbe *female populations of 199) and)97)* ty the above margins, *life* expectancy ratios for the two periods remained far apart at older iges (see Figure 1). This may be suggesting that the excessive undercount at older ages was nuch larger than what is assumed in my computations. Altematively, there may have been m unusually large transfer of women to older ages in the 1981 census on account of age tnisstatements. A satisfactory explanation to this is yet to be found.

[4] As the three post-enumeration checks indicated very similar age patterns of underenumeration, they were averaged to obtain a standard pattern of net omission rates per 1000 of enumerated population :

Age Interval

	0-4	*5-9*	*10-14*	*15-19*	*20-24*	*25-29*	*30-34*	*35-39*	*40-44*	*45-49*	*50-54*	*55-59*	*60**
Males	23.5	15.8	18.4	15.4	24.6	18.3	14.1	12.4	10.8	9.4	9.1	7,7	10.4
Females	25.5	16.9	17.7	22.3	21.9	17.8	13.6	9.5	8.9	13.8	13.9	12,6	23,4

The review of mortality estimates for the decades 1971-81 and 1981-91 provides nevertheless an important rule of thumb: the average of the estimates life expectancy for ages 10 and 15 that are derived using cohort interpolation in the variable-r methodology gives a reasonably satisfactory estimate of adult mortality in situations where census completeness remains more or less stationary. Large changes in enumeration completeness can also be detected from the method by studying the ratios of life expectancy estimates. Armed with this information, we may now examine the estimates for earlier decades.

Estimates for Periods before 1971

Estimates of life expectancy derived using the variable-r methodology for the decades 1931-41 to 1961-71 are shown in Table A2 of the appendix. Owing to the suspected quality of census data for 1941,1 have also provided estimates for the decade 1931-41 for comparison. The estimates were made for both using the unadjusted census totals and adjusted figures for migration and suspected changes in enumeration completeness.[5] Further, adjusted estimates of life expectancy at age 5 were derived by accepting the level of adult mortality implied by the averages of the life expectancy estimates for ages 10 and 15 and converting them to e_s estimates using the SRS life table of 1981-91 as the standard. In Table 1, these adjusted estimates of $e_\$$ are compared with estimates from the official life tables and other sources.

For 1961-71 the preferred estimates are those derived after adjusting the 1971 census totals for underenumeration, as the deficiencies of this census has already been pointed out. The adjusted variable-r estimates of e_s are higher than the official life table estimate for the decade by 3 years among females and 1.8 years among males. However, the new estimates are lower than those contained in the report of the NAS Panel of India (Bhat, Preston and Dyson 1984), especially for males.

For 1951-61 the major cause for uncertainty is the required correction for migration. An official estimate put the net immigration to India (particularly to the eastern parts) during this period at 3.14 million. Others have estimated a somewhat lower level of immigration (e.g., Visaria 1969). But any migration correction in the intercensal analysis leads to unacceptably high estimates of mortality for this period. In particular, it implies extremely low gains in life expectancy between 1941-51 and 1951-61.1 suspect that the migration estimates for this decade do not adequately take account of the counter flows of migrants from India to Pakistan. Hence the unadjusted estimates are preferred, though later I will show that even these estimates may be biassed. Somewhat surprisingly the variable-r estimates of e_s that are unadjusted for migration compare favourably with the official estimates, which possibly were corrected for the assumed level of migration. Apparently there were other compensating errors in the method employed by the census actuary.

Although there were large territorial transfers during 1941-51, the available estimates suggest negligible impact of migration on census totals. But the alleged overenumeration of population in 1941 census and incomplete information on the age distribution of the population (see Mukharjee 1976) complicates the analysis. If an estimated overenumeration of 2 million

[5] For making migration correction, I have made use of an age distribution of migrants reporting a duration of residence of 0-9 years in 1971 (see Bhat 1987).

TABLE 1: ESTIMATES OF LIFE EXPECTANCY AT AGE 5 DERIVED FROM VARIABLE-*r* METHODOLOGY USING COHORT INTERPOLATION AND THEIR COMPARISON WITH THE ACTUARIAL AND OTHER ESTIMATES, INDIA, 1941 TO 1991

Decade/ Assumption	*Estimates derived from variable-r methodology*		*Actuarial estimates*		*Other estimates*		
	Male	*Female*	*Male*	*Female*	*Male*	*Female*	*Source*
1931-41							
Unadjusted census counts	44.0	44.9					
Adjusted for immigration			—	—	44.0	41.7	Davis (1951)
and overcount in 1941	42.9	44.1					
1941-51							
Unadjusted census counts	46.7	46.5					
Adjusted for immigration			40.9	40.9	45.6	45.2	Visaria (1969)
and overcount in 1941	47.1	46.6					
1951-61							
Unadjusted census counts	48.4	47.5					
Adjusted for immigration	47.2	46.8	48.7	47.0	48.2	47.9	Visaria (1969)
1961-71							
Unadjusted census counts	53.4	52.3	52.0	50.2	55.4	53.7	Bhat, Preston &
Adjusted for undercount	53.8	53.3					Dyson (1984)
in 1971 (0.3 & 0.8%)							
1971-81							
Unadjusted census counts	58.1	59.6					
Adjusted for undercount			-	-	57.9	59.0	Official SRS L.T.
in 1971 (0.3 & 0.8%)	57.6	58.3			57.7	58.4	Author's SRS based L.T.
1981-91							
Unadjusted census counts	59.8	58.9					
Adjusted for undercount			-	-	60.5	62.1	Official SRS L.T.
in 1991 (0.2 & 1.8%)	60.1	61.7			60.1	61.7	Author's SRS based L.T.

persons (India, Registrar General 1953) in 1941 is allowed for, large gains in life expectancy are implied between 1931-41 and 1941-51 and very little improvement is suggested for the period 1941 -51 to 1951 -61. The use of unadjusted totals for 1941 brings the suggested pattern of changes to a more acceptable level. I suspect that some over enumeration of the population in Punjab and Bengal was compensated by underenumeration in other parts of the country as the 1941 census was undertaken under disturbed conditions of World War DL

The variable-r estimates of e_5 computed from the unadjusted census totals are higher by about 6 years from the official estimates. Such a difference could not have occurred but from a use of different age distribution or an erroneous methodology. The variable-r estimates are however closer to the estimates made by Visaria (1969) from a stable population model.

Widowhood Estimates of Adult Mortality

The estimates of adult mortality reviewed so far are either based on the census data on age and growth rates of population, or directly derived from the SRS. It would particularly be valuable to have an independent check on the estimated trends in mortality during the pre-SRS period. Widowhood data from the censuses present one such possibility. These data have rarely been used because of the possible contamination from remarriages. However, as widow remarriages are not that common in India, information on women's current marital status may be a valuable source for estimating trends in mortality levels among men.

On this consideration, I have applied the widowhood method (United Nations, 1983) to the census data of 1951 to 1991. The estimates of adult survivorship and their approximate reference periods are shown in Table A3 of the appendix. Model life tables are used to convert the estimates of adult survivorship into a common index of adult mortality such as e_y Since the SRS life tables suggest that mortality rates of men between the ages 25 to 60 are more closely approximated by the West model life tables than the South model tables, I have used the former in deriving the estimates of e_s. These too are shown in Table A3.

The estimates based upon the widowhood data of women under the age of 40 suggest very low levels of male mortality possibly because of high rates of remarriage among young widows. Data for women of ages above 40 show more acceptable levels of mortality; hence in Figure 2 only the estimated trends in male adult mortality derived from the reports women aged 40 to 59 are shown. For comparison I have also shown the estimates of e_s derived from

ESTIMATES OF MALE LIFE EXPECTANCY AT AGE 5

Reference year	*Variable-r estimates*	*Interpolated Widowhood estimates (45-54 ages)*	*Ratio*
1936.2	44.0	50.1	1.14
1946.2	46.7	50.9	1.09
1956.2	48.4	54.9	1.13
1966.2	53.8	58.1	1.08
1976.2	57.7	61.5	1.07
1986.2	60.1	-	

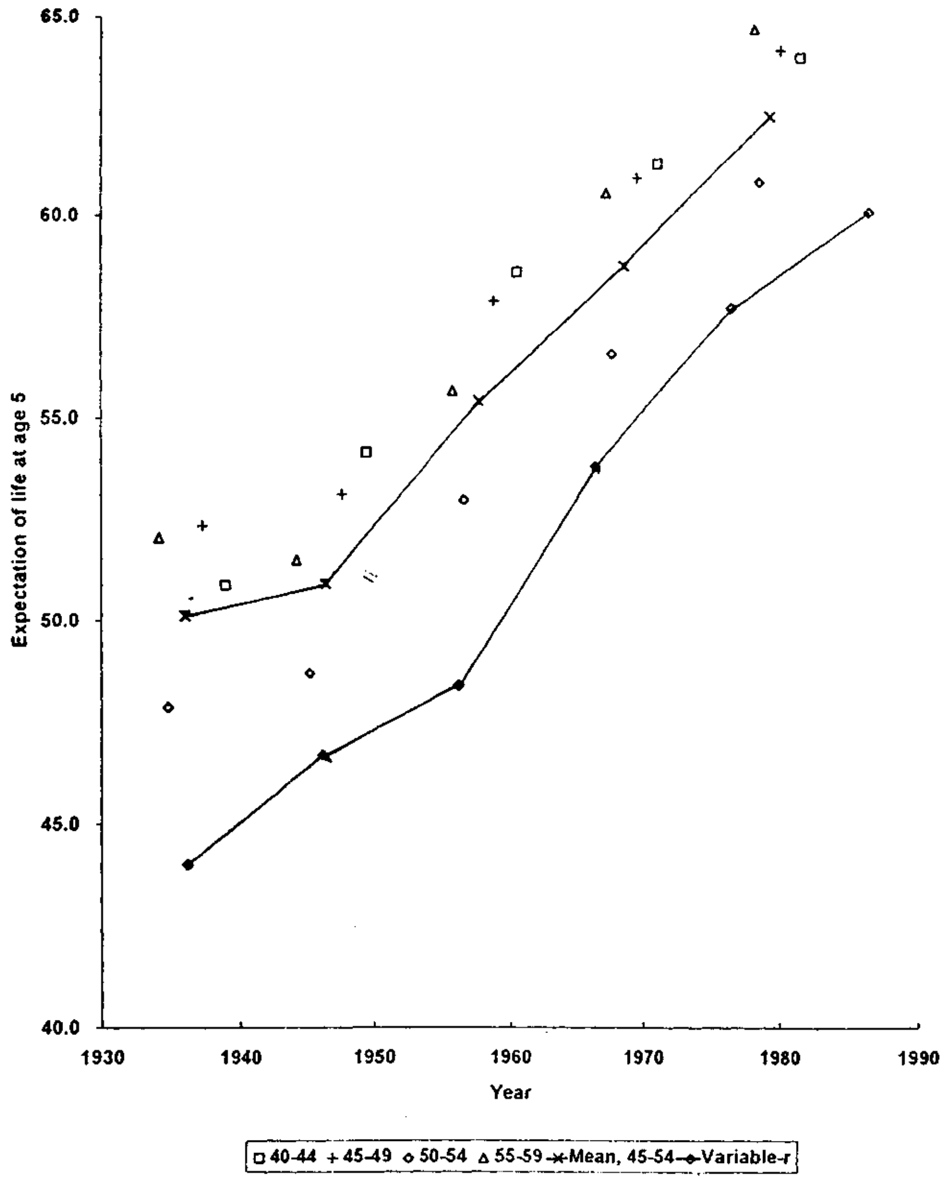


Fig. 2. Estimates of male life expectancy at age 5 from widowhood data, all India

the application of variable-r procedure. As can be seen from the figure, the widowhood estimates suggest very similar patterns of rise in male life expectancy as that of the variable-*r* estimates. However, not unexpectedly, all widowhood estimates of life expectancy are higher than those derived from the variable-r procedure. The widowhood estimates that come closest to the variable-r estimates are those based on the reports of women aged 50-54. This may be because of higher levels of digit preference in the age reports on widows. Therefore, the widowhood estimates of adult mortality based on women aged 45-49 and 50-54 may be averaged to study the trends in mortality:

Although both the series suggest rising levels of life expectancy, there are some subtle differences. The widowhood estimates indicate somewhat less rapid increase in life expectancy which may be because of the fall in widow remarriage rates as a result of a marriage squeeze against women (Bhat and Halli *in press)*. Nevertheless widowhood estimates suggest a sharper increase in male life expectancy between 1946 and 1956. In part this may be because of the problems associated with the 1941 census data. If the estimates for 1936 and 1946 are averaged, an upward bias of 12 percent is indicated in the widowhood-based estimates. The implied bias in the widowhood estimates shows a steadily fall except for an unexpected jump to 13 percent in 1956. This may be because of a downward bias in the variable-r estimate for the decade 1951-61. This bias could be corrected by assuming that the error in the widowhood estimate for 1956 should fall roughly in between those suggested for 1946 and 1966. Thus it is proposed that the true levels of male mortality in 1931-41 and 1941-51 were 12 percent lower than the widowhood estimates for the respective periods, and in 1951-61 they were lower by 10 percent. For the subsequent periods the estimates from the variable-r procedure could be relied upon without any correction. To obtain the corresponding estimates for females, the sex differential in adult mortality suggested by the variable-r estimates is used. The estimates of e_s corrected in this manner are shown in Table 3, which represent the final estimates of this section.

Estimates of Infant and Child Mortality

Assumed levels of infant and child mortality form the most contentious part of the demographic estimates proposed for the pre-SRS era. However, as with the estimates of adult mortality, it is convenient to review first the estimates of infant and child mortality for the more recent periods. When the SRS began to map yearly trends in the.early 1970s, it showed nearly unchanging levels of infant mortality until 1978. After this date however, it has registered a steady decline in infant mortality levels. A question arises as to how far the suggested constant levels of infant mortality for the early 1970s is correct?

When the trends in mortality levels at very early infancy are examined some strange patterns are revealed. While the post-neonatal mortality has registered a steady downward trend since 1972, neonatal mortality rates have shown a sharp increase from 70 to 80 per 1000 between 1973 and 1977, and a steady fall after 1978. If we .probe further near the event of birth, the early neonatal mortality rate (mortality in the first week of life) shows a strange pattern[6]. From a level of 33 per 1000 in 1973 it rose sharply to 49 per 1000 in 1976

[6] The estimate of early neonatal mortality rate is computable from the SRS data by subtracting the estimate of still birth ratio from the estimate of perinatal mortality rate.

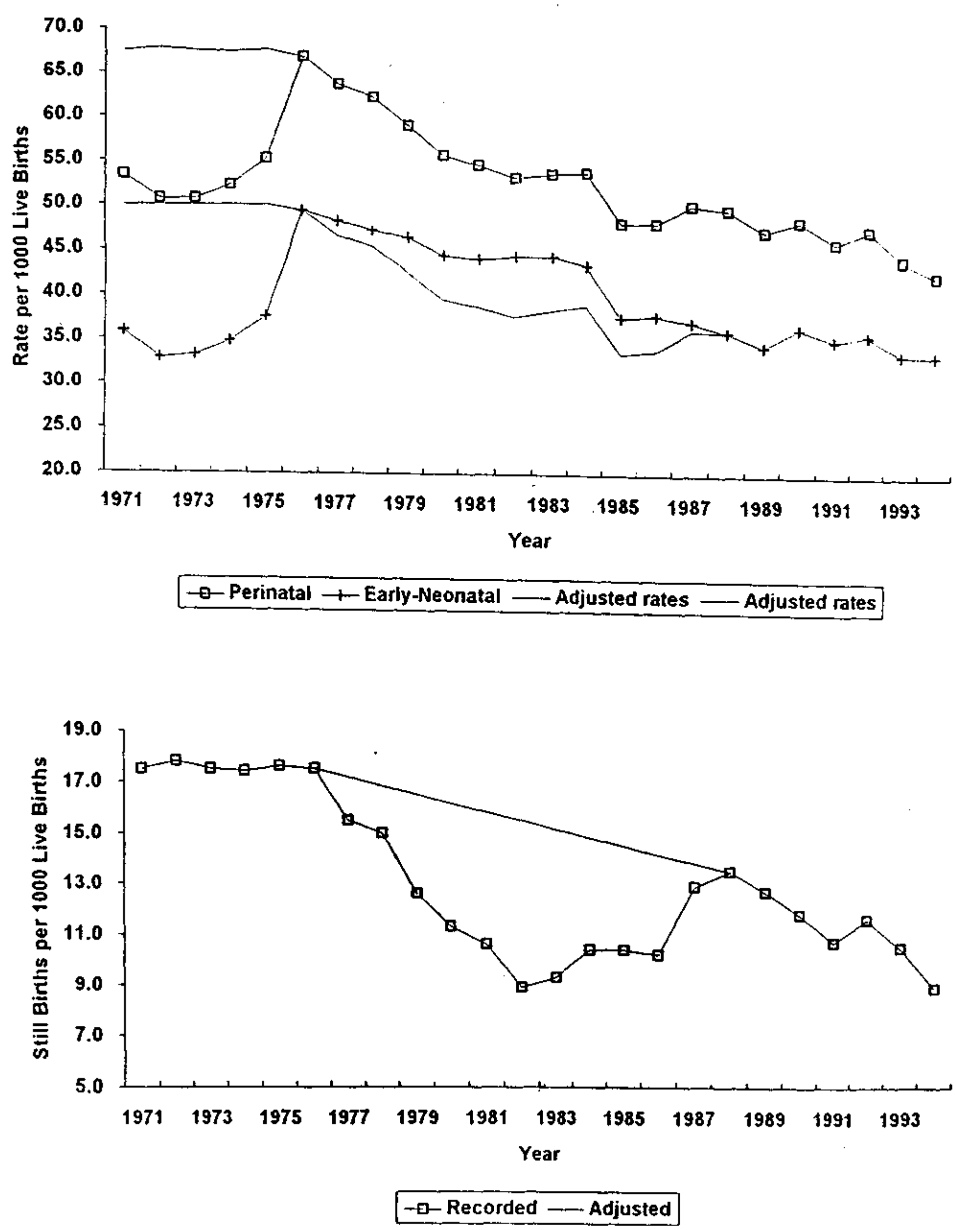


Fig. 3. Trends in perinatal mortality, early neontal mortality and still birth ratio from the SRS, 1971 to 1994

(see Figure 3). This increase was followed by a sharp decline until 1985 and a slow fall since then. But even in 1993-94, the reported levels of early neonatal mortality are higher than those reported for 1971 -73!

The suggested trend of the still birth ratio is even more enigmatic (see Figure 3). It remained unchanged at 17.5 per 1000 live births until 1976, and then began a sharp fall to 8 per 1000 in 1982; then the ratio began to climb and reached 13 per 1000 in 1988. Since then it has steadily fallen to reach 9 per 1000 in 1994. Thus, if one were to believe the SRS story, during the last 25 years still birth ratio has declined by about 100 percent while there has been no perceptible change in the mortality risk of the period immediately following the birth.

It appears quite likely that the SRS trends are influenced by the conceptual problems in distinguishing a still birth from an infant death. The large fall in still birth ratio between 1976 and 1982 and its subsequent rise until 1988 seems particularly spurious, and a result of the classification problem. By assuming a linear fall in the still birth rate from 17.5 in 1976 to 13 in 1988, it is possible to correct the SRS trends. However, the registered rise in the early nearly neonatal mortality during 1973-77 cannot be easily attributed to the classification problem as the still birth ratio has not shown any drastic changes during this period. Therefore the rise must have been due to an improvement in the netting of early infant deaths. The early neonatal mortality rates of the period 1971 -76 could be corrected by assuming that the rate was more or less constant at around 50 per 1000 (approximately its 1976 level). Thus the revised rates imply a fall of still birth ratio from 17.5 in the early 1970s to 9 per 1000 in the early 1990s, and a reduction of early neonatal mortality rate from 50 to 34 per 1000 during the same period - which appear far more plausible than the uncorrected SRS trends.

The suggested corrections for the estimates of early neonatal mortality rates have implications for the estimates of infant mortality rate and death rates of the age interval 0-4.

The infant mortality rate could easily be corrected by an addition, but correcting death rate of 0-4 ages is somewhat complicated. First an estimate of the probability of dying between birth and age 5 (also known as the under-5 mortality rate) is derived from the unadjusted rates using the formula:

$$q_5 = m_5 (5 - 2q_1) / (1 + 2.75\ m_5) \qquad (1)$$

where m_5 is the death rate of the age interval 0-4 and q_1 is the infant mortality rate. This formula assumes that infant deaths live for 3 months on an average and deaths of the age interval 1-4 live for 1.25 years after completing one year.

From the unadjusted estimates of q_1 and q_5 the estimate of probability of dying between the ages 1 and 5 years is estimated using the formula:

$$_1q_4 = (q_5 - q_1) / (1 - q_1) \qquad (2)$$

Since the mortality rate in the age interval 1-4 were unaffected by the errors m the estimate of early neonatal mortality, an adjusted estimate of q_s can be derived by plugging in the adjusted estimate of infant mortality rate in equation (2) along with the estimate of $_1q_4$ derived using the unadjusted rates. Further, the adjusted rates of both q_s and infant mortality rate can be plugged back in equation (1) to get a revised estimate of the death rate of ages 0-4.

The adjusted SRS estimates of infant and child mortality rates are shown in Table A3 of the appendix. The corrected estimates show that infant mortality was indeed falling in the early 1970s; from about 150 per 1000 in 1971-72 it decreased to 125 in 1978. The decline has continued during the 1980s to reach a level of 75 in 1993-94. The under-5 mortality rate has shown an even more impressive fall: from about 240 per 1000 live births in 1971-72, it declined to 200 in 1978; except for an abrupt increase during 1983-84, it has decreased steadily to record in 1993-94 a value less than half of the level it started with in 1971-72.[7] It is possible to check the validity of this pattern of change from the indirect estimates of child mortality derived from census/survey data on children ever-born and children surviving.

Indirect Estimates of Child Mortality

Besides the 1981 and 1991 censuses, a number of national surveys have collected information required for the indirect estimation of child mortality since the mid 1960s. In Table 2 I have shown the indirect estimates of child mortality derived from the two recent population censuses and five country-wide surveys. As the information supplied had slightly varied from source to source, it has not been possible to apply the same set of multipliers in each case; but this should not significantly affect their comparability. For the sake of brevity, I will restrict discussion to the graduated estimates of q_s derived from the responses of women aged 20-24 and 25-29 years at the time of census/survey. Figure 4 shows a comparison of the indirect estimates of q_s with those computed from the adjusted SRS rates. As can be seen, but for the estimate from the National Family Health Survey, all indirect estimates of under-5 mortality are significantly lower than the direct estimates from the SRS. The child mortality estimate from the National Family Health Survey for 1989 (130) is quite close to SRS estimate for the same year.[8] Baring the NFHS-based estimate, other indirect estimates of child mortality show an average downward bias of 25 percent. This suggests that unless a complete birth history is probed, even the younger women tend to underreport child deaths in a survey, and more so in a census.

This doesn't however imply that the indirect estimates of child mortality are worthless. When the estimates from several sources are taken together, they show a pace of child mortality decline very similar to that registered by the SRS (see Figure 4). But for some minor deviations - possibly reflecting the relative completeness of survey data - child mortality is shown to have fallen almost linearly from the early 1960s to the late 1980s. If a straight line is fitted to the data excluding the NFHS estimate, a drop of 31 per 1000 in under-5 mortality is indicated for every five years. If a similar line is fitted to the SRS data from 1971 to 1994, it too suggests a decline of 30 per 1000 in child mortality for every 5 years. Such a consistency is more than what one would ask for in any data.

[7] The abrupt rise in child mortality in 1983-94 may be because of the switching of the SRS sampling units which affects the population reported at ages 0-4 years.
[8] The child mortality estimates from the NFHS birth history data are however lower than the corresponding SRS estimates, which points to the possibility of date misstatement in the NFHS birth history.

TABLE 2: INDIRECT ESTIMATES OF CHILD MORTALITY DERIVED FROM DATA ON CHILDREN EVER-BORN AND CHILDREN SURVIVING FROM VARIOUS CENSUSES AND SURVEYS, ALL INDIA

Source	*Year of survey/ census*	q^2	q^3	q^5	q^{10}	*Graduated* q^5 *@@*		*Excess female mortality*
NSS 20th round *	1965-66	0.211	0.240	0.268	0.309		0.251	-
ORG Survey	1970	0.212	0.214	0.234	0.242	0.269	0.238	-
SRS Special Survey @	1972	0.174	0.190	0.206	0.220		0.204	-
SRS Special Survey @	1979	0.162	0.152	0.160	0.173	0.184	0.177	-
Census of India	1981	0.123	0.132	0.152	0.173	0.192	0.144	10.1
Census of India	1991	0.085	0.086	0.095	0.107	0.121	0.097	17.1
NFHS -I	1992-93	0.123	0.107	0.117	0.143	0.163	0.130	6.0
Reference period **		2.2	4.2	6.5	9.1	12.2	3.2	

* Derived from data classified by duration of marrige after applying some correction to those reported for widows, see Bhat (1987). @ Weighted averages of state-specific estimates, see Bhat (1987).
** Years before the survey/census; as they showed little variation, estimates from several sources were averaged. @@ Graduated estimates of q^2 & q^3 with a South model life table.

TABLE 3: LEVELS OF INFANT AND CHILD MORTALITY IN MODEL LIFE TABLES CORRESPONDING TO ESTIMATES OF LIFE EXPECTANCY AT AGE 5, AND EXTRAPOLATED ESTIMATES OF INFANT & CHILD MORTALITY FROM THE SRS

	*Estimates for both the sexes combined**											
Decade	*Preferred estimates of* e_s		*Corresponding to West model*		*Corresponding to South model*		*Weighted average of West & South*		*Extrapolated SRS trends*		*Actuarial life-tables*	
	Male	*Female*	q^5	*I MR*	q^5	*IMR*	q^5	*I MR*	q^5	*IMR*	q^5	*IMR*
1931-41	45.1	46.0	0.348	234	0.431	243	0.348	234				
1941-51	45.8	45.6	0.344	232	0.427	241	0.344	232	0.395	243	0.309	185
1951-61	49.9	49.0	0.276	184	0.356	202	0.303	184	0.335	207	0.224	146
1961-71	53.8	53.3	0.208	140	0.284	166	0.259	153	0.275	171	0.192	129
1971-81	57.7	58.4	0.139	97	0.213	133	0.215	135	0.215	135		
1981-91	60.1	61.7	0.102	73	0.173	114	0.154	95	0.154	95		

* Sex-specifc estimates were combined using a sex ratio at birth of 106.

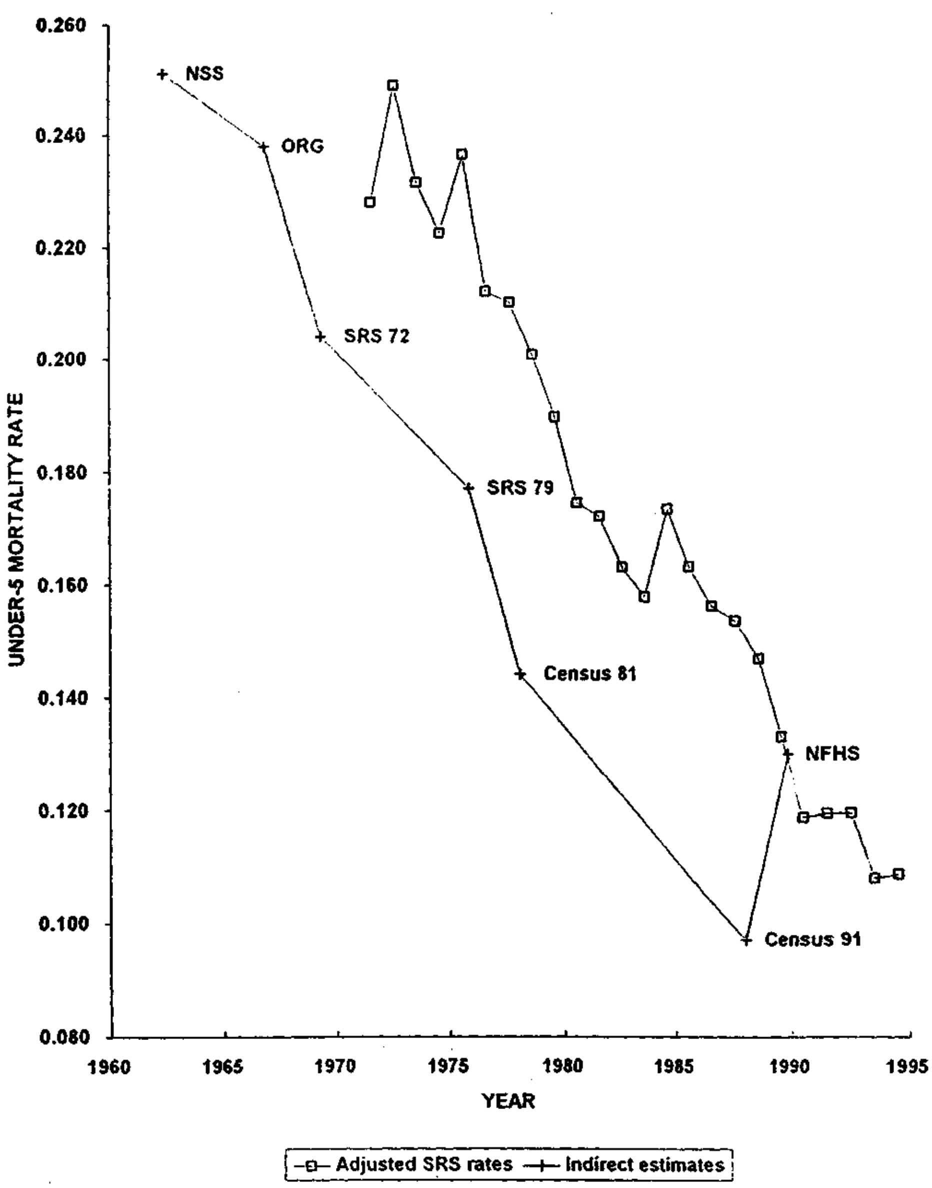
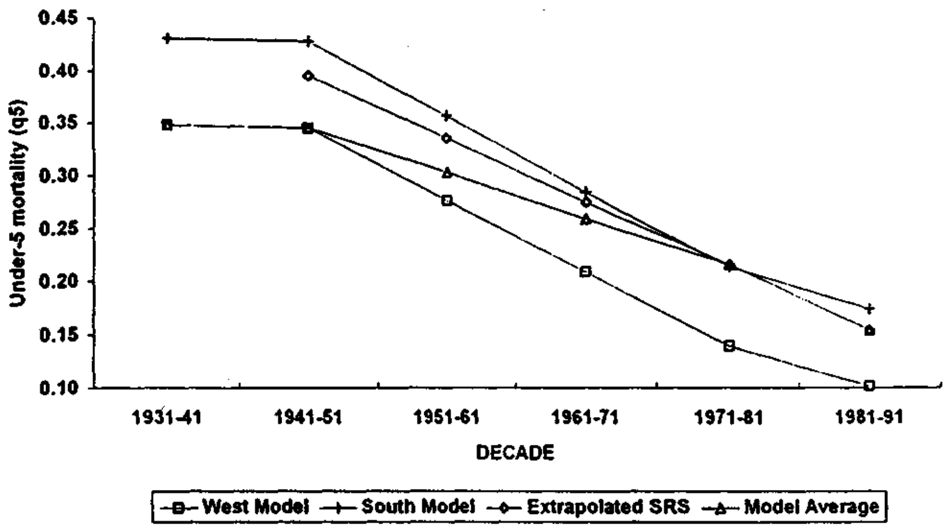


Fig. 4. Estimates of under-5 mortality from indirect methods and Sample Registration System



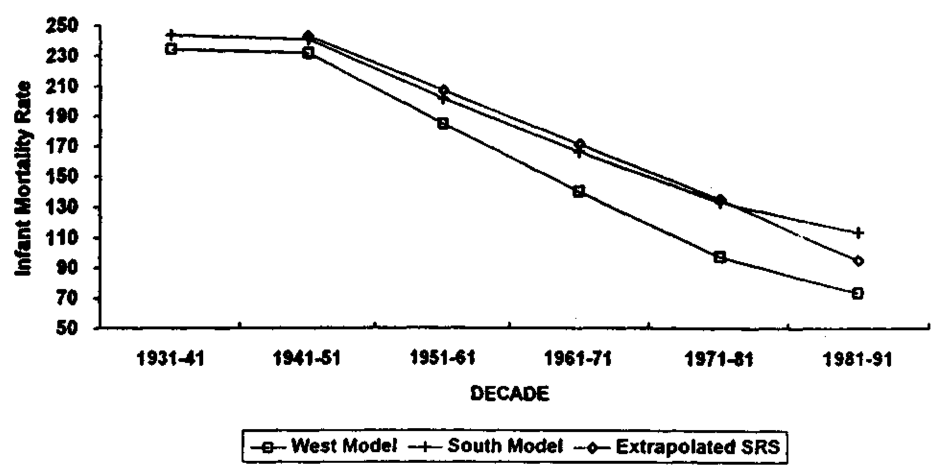


Fig. 5. Estimates of infant & child mortality through model and SRS extrapolation

Child mortality estimates by extrapolation

While the all-India series of infant and child mortality rates from the SRS begins at 1971 only, by using the Brass-type indirect estimates of child mortality it is possible to go back only up to the beginning of 1960s. For making estimates for earlier periods it is necessary to resort to some form of extrapolation. One such method is the extrapolation of adult mortality levels to childhood ages with the aid of model life tables. In Table 3,1 have shown the model life table estimates of infant mortality and under-5 mortality rate corresponding to the adult mortality estimates made in the preceding section. The estimates corresponding to West and South model life tables have only been shown as they are the main contenders to represent the age pattern of mortality in India. There is quite a bit of variability in the levels of under-5 mortality indicated by the two model families but the difference in the suggested levels of infant mortality is quite small (see Figure 5). For example, for 1941 -51, the level of under-5 child mortality (for both sexes combined) corresponding to the estimates of adult mortality for the decade is about 350 per 1000 under the West model and 430 under the South model. But the suggested range for the infant mortality level is only between 230 to 240. This is because the major cause for uncertainty is the level of mortality in the age interval 1-4 years.

Attention may be drawn to another interesting aspect of Figure 5. The changes in adult mortality levels in India was such that a linear fall in the levels of infant and child mortality could be expected between 1941-51 and 1971-81. Such a trend is predicted by both the models. Between 1971-81 and 1981-91 however, a slowing down of the decline is predicted (even though this did not actually happen). It has already been noted that both the adjusted SRS rates and indirect child mortality estimates have indicated a linear fall in child mortality since the 1960s. Therefore the adjusted SRS rates for the 1970s could be projected backwards in time assuming a linear change in mortality. The trends in infant and child mortality suggested by such an exercise have also been shown in Table 3 and Figure 5.

In the case of infant mortality rates, the back-projected SRS rates are very close to the levels suggested by the South model life tables. For the past however, there is not much difference in the estimates of infant mortality predicted by the two models. For 1981-91, where the difference begins to widen, the SRS rates show a tendency to move towards the West model life tables.

In the case of under-5 mortality rate, the SRS rates are almost identical to the south model values in 1971-81, But as one goes back in the past, the back-projected SRS rates move closer to the West model values. By 1941-51 the extrapolated SRS under-5 mortality rates were roughly halfway between the West and the South model values.

What this exercise shows is that during the course of mortality transition in India, for an unit change in the adult mortality level, child mortality levels have declined slower than what is implicit in the model tables. As such, the age pattern of mortality has undergone a shift from a pattern closer to West model tables to that of South model tables. This transition was completed in 1971-81. Since then, declines in adult mortality levels have tapered off (especially among males), while declines in infant and child mortality have continued - thanks perhaps to immunization and child nutrition programmes- and the age pattern of mortality is tilting back towards the West model life tables.

Keeping this in mind, I propose two alternative scenarios for the pre-1971 levels of child mortality. Under the first scenario, child mortality levels during 1941 to 1971 are assumed to be the same as those obtained from back-projecting the adjusted SRS trends. Under the second scenario, the West model age pattern of mortality is assumed to apply for 1941-51, and the South model pattern for 1971-81. For the intervening period, infant and child mortality rates are computed by linear interpolation. It may be noted that under both the scenarios, the implied levels of infant and child mortality are higher than those contained in the official Indian life tables (see Table 3). Implications of the two mortality scenarios for the levels of fertility and patterns of population growth will be gauged later in this paper.

Sex Differentials in Child Mortality

Discussion on this issue has been avoided so far because the model life tables do not adequately describe the patterns observed in India. Fortunately, the estimates of sex differentials in child mortality can be obtained directly from the application of variable-r procedure to the census data on age-sex structure of the population. This estimate is totally insensitive to the assumed level of crude birth rate, but requires an assumption on the sex ratio at birth. The estimate is also sensitive to the changes in enumeration completeness and differences in age misstatements patterns by sex.

TABLE 4: SEX DIFFERENTIALS IN CHILD MORTALITY IMPLIED BY THE VARIABLE-/-ANALYSIS AND IN THE SRS LIFE TABLES, ALL INDIA, 1941 TO 1991

Decade	*Estimate of q^{15}*		*Female-male mortality* ratio
	Males	*Females*	
1931-41	0.449	0.472	1.051
1941-51	0.444	0.440	0.991
1951-61	0.365	0.380	1.041
	Without underenumeration correction		
1961-71	0.322	0.354	1.099
1971-81	0.240	0.264	1.100
1981-91	0.183	0.223	1.219
	With underenumeration correction		
1961-71	0.319	0.344	1.078
1971-81	0.245	0.273	1.114
1981-91	0.183	0.199	1.087
	Life tables constructed from SRS data		
1971-81	0.229	0.254	1.109
1981-91	0.167	0.183	1.096

Table 4 shows the estimates of probability of dying before attaining the age of 15 years derived from applications of the variable-r methodology for both males and females. The age range up to 15 years is chosen to minimize the affect of age misstatements on the estimated sex-differentials in mortality. The absolute level of mortality indicated by the estimates is to be ignored since it is affected by the assumption on the level of birth rate. The sex-specific estimates of child mortality show that the excess mortality of girls was no more than 4 percent up to 1961. It has shown a tendency to increase in recent periods. If no correction is made for the differential underenumeration of females, a sharp increase in the female disadvantage from 10 percent in 1971-81 to 22 percent in 1971-81 is indicated. This is supported by the sex-specific indirect estimates of child mortality derived from the 1991 census data (see Table 2). However, this is not borne out by either the SRS data or the NFHS data which show an excess female mortality of less than 10 percent in the late 1980s. As there are grounds to assume that the 1991 census data on child mortality were of poor quality than that of 1981 census, its evidence may be set aside. Also to be noted is that in my calculations I have used throughout a sex ratio at birth of 106 males for 100 females. If, as some claim, the sex ratio at birth too has gone up, it would have the effect of dampening the estimated rise in the sex ratio of child mortality. It seems to me that the evidence is slightly in favour of the underenumeration hypothesis, as there certainly was some undercount of women at older ages in 1991. If a correction is made for this, one actually sees some reduction in the sex differentials in mortality between 1971-81 and 1981-91 (see Table 4).

Estimates of Life Expectancy at Birth

By combining the estimates of adult mortality and child mortality, it is possible to arrive at the estimates of life expectancy at birth.[9] As two estimates of child mortality are proposed for the pre-1971 period, there will be two estimates of expectation of life at birth:

Decade	*Official estimates*		*New estimates*	
	Male	*Female*	*Male*	*Female*
1931-41	32.1	31.4	30.9-33.2	31.1-33.5
1941-51	32.4	31.7	31.4-33.8,	30.9-33.4
1951-61	41.9	40.6	37.3-39.0	35.9-37.6
1961-71	46.4	44.7	43.6-44.5	41.9-42.9
1971-81	51.5*	50.5*	50.0	49.3
1981-91	56.5*	56.9*	55.6	56.2

* As per the official SRS life tables.

In spite of the large differences in the estimates of adult motility and child mortality levels, my estimates of expectation of life at birth are not substantially different from the official estimates. This is because while the new estimates of child mortality are substantially

[9] Estimates of life expectancy at birth is derived using the relation:

$$e_0 = (e_5 + 5 - a_0)(\backslash - q_5) + a_0$$

The value of a_n is assumed to be 1.1.

higher than the official estimates, the suggested levels of adult mortality are lower. Thus the presence of compensating errors in the official estimates reduced the discrepancies in the estimates of life expectancy at birth. Nevertheless some differences can be noted. For all the periods after 1951, the estimates now being proposed are lower than the official estimates. This is particularly so for the decades 1951-61 and 1961 -71, where the official estimates of expectation of life at birth are higher by about 2 to 4 years. The ability of the new estimates of life expectancies to correctly describe the intercensal patterns of population growth will be checked later in this paper through a simulation test.

Estimates of Fertility Level

I propose to use a uniform methodology to estimate fertility levels for all the decades from 1941 to 1991. Although for the last two decades estimates from the SRS appear to provide credible alternatives, it will be shown that the system significantly underestimates fertility levels, even if, somewhat paradoxically, estimates of age-specific mortality it generates can be trusted.

I shall use the census data on child-woman ratios to estimate the fertility level. Among the available methods to estimate fertility from the age-structure of the population, methods that make use of child-woman ratios are least affected by assumptions on mortality level. As there is considerable uncertainty on levels of mortality prevailed in the past, it would make sense to use a method least affected by it.

The late Professor Rele (1967,1987) had proposed a method of estimating fertility from the child-woman ratios. His method has an advantage over others in that it can be used to build a time series of fertility estimates using data from censuses taken 10 years apart. Rele, however, had ignored possible variations in the age pattern of mortality and in the mean age at maternity that could influence any fertility estimate derived from the child-woman ratios. He had also ignored the possible variation in sex differentials in mortality. Further, the multipliers he had developed are derived from a limited set of stable populations (36).

On these considerations, we have developed a new set of multipliers that are specific to model life table patterns and explicitly account for the mean age at maternity.[10] In estimating gross reproduction rate, we suggest using female children only in the numerator of the child-woman ratios. Further, rather than using an estimate of life expectancy at birth, the use of an estimate of the under-5 mortality is recommended to minimize the impact of the assumed model age pattern of mortality. When direct information on the mean age of the fertility schedule (m_) is lacking we suggest estimating it from the singulate mean age at marriage of females and the estimated level of GRR. For this purpose we make use of the following empirically derived equation:

$$\overline{m} = 0.9SMAM + \frac{GRR(40 - SMAM)}{7.3} + 4.5$$

This equation is used iteratively with the multipliers to estimate both GRR and m

[10] The paper is under preparation for publication.

TABLE 5: ESTIMATING TOTAL, FERTILITY RATE FROM FEMALE CHILD-WOMAN RATIO ALL INDIA- 1941 TO 1991

Decade	*Female child-woman ratio **	*Female singulate mean age at marriage*	*Assumed level of under-5 mortality*		*Preliminary TFR estimates (Scenario I)*			*Smoothed estimates*			
								Scenario I		*Scenario II*	
			Scenario I	*Scenario II*	*West model*	*South model*	*Weighted series*	*TFR*	*m*	*TFR*	*m*
1941-46	0.291	a	0.405	0.355	6.28	6.70	6.53	5.77	29.5	5.06	28.7
1946-51	0.281	a	0.395	0.345	5.40	5.43	5.42	6-19	29.5	5.63	28.8
1951-56	0.347	15.4	0.360	0.320	7.09	7.53	7.40	6-53	29.6	6.13	28.8
1956-61	0.328	15.8	0.330	0.300	5.89	5.93	5.92	6-72	29.7	6.44	29.1
1961-66	0.366	16.3	0.300	0.280	7.30	7.71	7.63	6-72	29.9	6.49	29.4
1966-71	0.324	16.9	0.280	0.260	5.67	5.71	5.70	6-36	29.5	6.22	29.3
1971-76	0.337	17.5	0.250	0.250	6.16	6.41	6.41	5.72	28.8	5.69	28.8
1976-81	0.269	18.0	0.205	0.205	4.32	4.36	4.36	5.05	28.1	5.05	28.1
1981-86	0.306	18.4	0.170	0.170	4.97	5.09	5.07	4-60	27.7	4.60	27.7
1986-91	0.252	18.8	0.150	0.150	3.88	3.91	3.90	4,52	28.0	4.52	28.0

* Child-woman ratios shown agianst the first half of a census decade is the ratio based on 5-9 year old children and that against the second half is the ratio based on 0-4 year children.

a No SMAM value is used; estimates of *m* (raw) for the period 1951-61 have been repeated.

Table 5 shows the input values used in estimating fertility from the child-woman ratios recorded in the 1951 census onwards. The two input values of under-5 mortality correspond to the two scenarios discussed in the previous section. The preliminary estimates of GRR are derived using both West and South model multipliers. In Table 5, I have shown the preliminary estimates of total fertility rate computed using a sex ratio at birth of 106. As can be seen from the table, the estimates of TFR derived from using the ratios of 0-4 year children to adult women aged 15-49 (estimate for the later half of the intercensal period) are not significantly different under the two model multipliers. However, the estimates of TFR computed using the ratios of 5-9 year children with South model multipliers are significantly higher than those that used the West model multipliers. This clearly shows the need to take account of possible variations in the age pattern of mortality in Rele's method.[11]

Our estimates of child mortality in the two mortality scenarios do not exactly fit either the South or the West model patterns. But their exact deviations from the model patterns are known for each time period (see Table), which may be used to compute a set of weights that can be used for the TFR estimates corresponding to the West and South model tables to derive a more appropriate series for India.

But the weighted series too is not without errors. In particular, the estimate of TFR level for the first half of an intercensal period are always higher than the estimate for the second half of the same intercensal period or that of the preceding one. Such a pattern in the preliminary estimates of TFR is almost certainly due to age misreporting and under enumeration. I therefore suggest smoothing the TFR estimates by using the following moving average formula:

$$TFR_t = 0.25\, TFR_{r}\, 5 + 0.5\, TFR_t + 0.25\, TFR_t + 5$$

In the moving average formula, the TFR for any given 5-year period gets twice the weight of the estimates for the two neighbouring time intervals to compensate for the additive impact of biasses in the estimates of the adjacent time periods (see Figure 6). "The estimates for the beginning period of the series (1941-45) and for the last period (1986-90) could be obtained by assuming the 10-year rate of change implied in the preliminary TFRs: the smoothed TFR for 1941-45 is computed by multiplying the smoothed TFR of 1951-55 by the ratio of preliminary TFR of 1941 -45 to that of 1951 -55; and smoothed TFR for 1986-90 is computed by multiplying the smoothed TFR of 1976-80 by the ratio of preliminary TFRs of the two periods.

The smoothed estimates of TFR so derived under the two mortality scenarios are shown in Table 5. Figure 5 shows a plot of the estimated trends in TFR. The estimated TFR for the period 1941 -45 is 5.8 under Scenario I which uses the extrapolated estimates of child morality from the SRS, and 5.1 under Scenario II which uses child mortality estimates that were derived by assuming a West model pattern. The difference in the TFR estimates diminish with time and the two series converge in 1971-81. The most intriguing aspect of the TFR estimates is, however, the suggested rise in fertility between 1941 and 1965. As per the first scenario, the rise is by about one child in TFR, and by about 1.4 under the second scenario.

[11] The use of the original multipliers of Rele gives estimates of TFR closer to the West model estimates shown here.

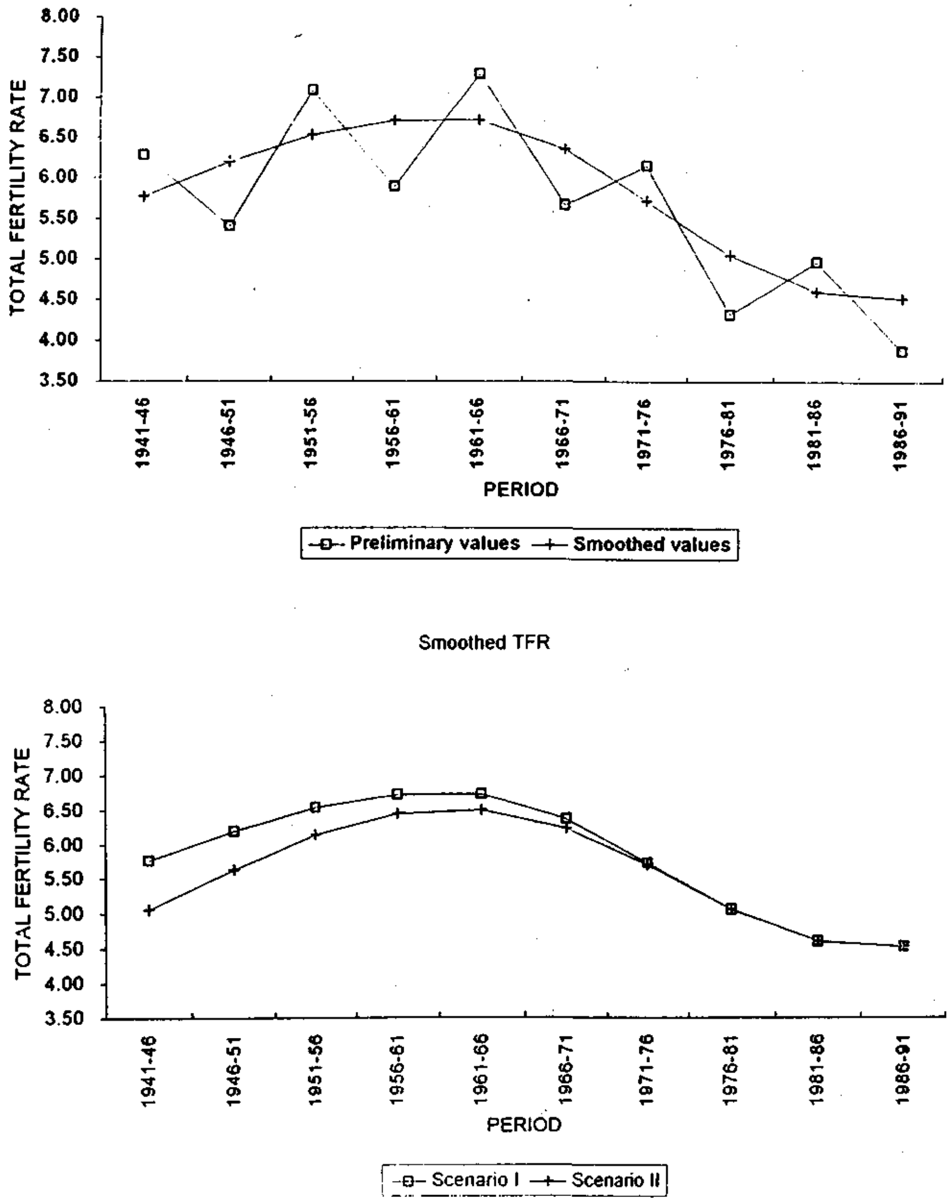


Fig. 6. Estimates of TFR derived from child-woman ratios
Mortality Scenario I

The maximum level of fertility was attained during 1956-65 when the TFR was between 6.4 to 6.7 births per woman. These are the highest ever estimates of fertility made for India.

But why did fertility levels increase? The most probable reason is the declines in mortality itself. Dyson and Murphy (1985) have documented the rise in fertility in many Third World countries following malaria eradication campaigns. The selected analyses of civil registration data have also disclosed a similar rise in fertility in several parts of India (e.g. Dyson, 1989; Dasgupta, 1995). But it is the magnitude of the suggested rise that intrigues. The second scenario that assumes slower declines in child mortality suggests larger rise in fertility. One is therefore inclined to trust the first scenario which suggests larger increase in child morality and relatively moderate rise in fertility.

The new estimates of TFR have significant revisions to offer for the estimated trends in fertility of more recent times. Below I show a comparison of the estimates of TFR derived from census child woman ratios with those from the SRS for the period 1971 to 1991:

Period	*New estimates*	*SRS estimates*		*Ratios*	
	(1)	*Raw (2)*	*Smoothed (3)*	*(2)/(l)*	*(3)/(l)*
1971-75	5.72	5.00	-	0.87	-
1976-80	5.05	4.51	4.62	0.89	0.91
1981-85	4.60	4.47	4.36	0.97	0.95
1 986-90	4.52	3.98	4.08	0.88	0.90

As the new estimates are smoothed, it is more appropriate to compare the SRS estimates that are similarly smoothed. As can be seen from the SRS figures^ the smoothing doesn't substantially change the underlying trend. The comparison shows that birth incompleteness in the SRS was of the order of 10 percent in the 1970s; there perhaps was only a marginal improvement in the coverage between 1971-75 and 1976-80. There, however, was substantial improvement in the first half of the 1980s when the underregistration fell to about 5 percent. But in the late 1980s incompleteness had shot back to 10 percent.

Simulation Test

The analyses of the foregoing sections have suggested that (i) for the most part, age-specific mortality rates from the SRS do not require significant correction, and (ii) fertility rates of the SRS have a downward bias of about 10 percent. Taken together these facts would suggest that the growth rates of the population recorded in the SRS are erroneous. But how is it that the source recorded approximately the same growth rate of the population as the 1981 and 1991 censuses? To resolve this paradox, it is necessary to check the growth implications of the estimates of fertility and mortality levels set forth in this paper. Therefore a projection exercise was carried out using the two mortality variants and their corresponding fertility rates.

The age-sex structure of the population in 1941 was first constructed using the variable-r-relations. This required life tables for males and females representing mortality levels of the period. As per our estimates, in the case of both males and females, life expectancy at birth in 1941 was about 31 years under the first scenario and 33.5 years under the second scenario. While the age pattern of mortality in the 1940s was similar to the West model pattern under the second scenario, under the first, a pattern in between the West and the South model life tables was applicable. Accordingly, for the first scenario, life tables were constructed by taking mortality rates for ages above 5 years from West model life tables with e5 values of 45.5 for males and 45.8 for females (the average estimates for 1931-51) but under-5 mortality rates were assumed to be 390 for males and 400 for females. Under the second scenario, male and female life tables with life expectancy at birth of 33.5 years were directly chosen from the West model family. In both the cases, the age-specific growth rates employed were those recorded for 1931-51.

The two estimated age-sex distributions for 1941 were projected forward to 1991 using their respective mortality and fertility rates. The time interval used in the projection was five years which required interpolating the decadal estimates of life expectancies. The age-specific mortality rates corresponding to these life expectancies were compute on the assumption that the age pattern of mortality progressively shifted from 1941 to that hi the SRS life table of 1971 -75. For the subsequent periods, life tables constructed from the adjusted SRS rates were employed. The age-specific fertility rates used in the simulation were those corresponding to the estimates of TFR and m derived from the child-woman ratios. The age pattern of fertility in 1941 was approximated by the average SRS rates for rural areas of$_t$ Uttar Pradesh, Madhya Pradesh and Rajasthan in 1971-72. For other projection intervals, age-specific fertility rates were computed by interpolating for the required values of m and adjusting them to sum to the TFR estimate.

Table 6 shows the projected values of the rate of natural increase and sex ratio of the population for every 10-year interval from 1941. For validation, the simulation results are to be compared with the unadjusted census figures up to 1961, and since 1971, with the estimates adjusted for underenumeration. The growth rates of population and sex ratio trends implied by the two projection runs are very similar, and as such the results are not helpful in discriminating the two variants. For males, the projected growth rates for every decade come very close to the recorded or the adjusted census figures, as the case may be. A slightly higher growth rate found for 1951 -61 was not unexpected as the variable-r estimate for this decade was substituted by a corrected widowhood estimator of e_v

The projected growth rates of female population are close to the adjusted census estimates for 1971-81 and 1981-91. However for other decades the projected figures are somewhat lower than the census estimates. Partly for this reason, the simulations show sharper reductions in the sex ratio of population (females per 1000 males) than that described by the adjusted census figures. This indicates that the levels of female life expectancies in the past may have been slightly higher than what my estimates suggest. Perhaps the rule of thumb employed in the variable-r procedure did not perform so satisfactorily in the case of females. I, therefore, suggest jacking up slightly the estimates of female life expectancy at birth, (see p. 50):

In effect, the expectation of life at birth was assumed to be the same for males and females in 1941-51;but it was assumed to be higher for males by half ayear in 1951-61 and by one year in 1961-71. The estimates for 1971-81 and 1981-91 were left unaltered as they were able to reproduce the adjusted census growth rates.

TABLE 6. ESTIMATES OF GROWTH RATE OF POPULATION, SEX RATIO, CRUDE BIRTH RATE AND CRUDE DEATH RATE IMPLIED BY THE PROJECTION OF 1941 POPULATION BY AGE AND SEX UNDER TWO ALTERNATE SCENARIOS OF MORTALITY AND FERTILITY LEVELS, ALL INDIA
(ALL FIGURE ARE PER 1000)

Population parameters	*Decade*				
	1941-51	*1951-61*	*1961-71*	*1971-81*	*1981-91*
Receded/adjusted values*					
Growth rate:					
Male	12.5	19.9	22.5 (22.8)	22.1 (21.8)	21.7 (22.0)
Female	12.6	19.3	21.4 (22.1)	22.5 (21.7)	20.9 (22.7)
Sex ratio (F/M)	946	941	930 (934)	934	927 (942)
Estimates from Scenario I					
Growth rate:					
Male	12.4	20.3	22.6	21.8	22.2
Female	11.7	19.1	21.7	21.7	22.7
Sex ratio (F/M)	938	928	920	918	923
Estimates from Scenario II					
Growth rate :					
Male	11.8	20.6	22.4	21.9	22.3
Female	11.3	19.4	21.6	21.8	22.8
Sex ratio (F/M)	941	930	922	921	925
With Adjusted Estimates of Female $e_{(t}$ and Sex Ratio at Birth of 105					
Estimates from Scenario I					
Growth rate:					
Male	12.3	20.3	22.8	22.2	22.5
Female	12.3	19.8	22.1	21.9	22.9
Sex ratio (F/M)	945	940	934	931	935
Crude birth rate	43.9	47.5	44.3	38.1	34.7
Crude death rate	31.6	27.4	21.9	16.1	12.0
Estimates from Scenario II					
Growth rate:					
Male	11.7	20.7	22.6	22.2	22.6
Female	11.7	20.3	22.0	21.9	23.0
Sex ratio (F/M)	945	942	937	934	937
Crude birth rate	40.1	46.0	43.1	38.2	34.9
Crude death rate	28.5	25.5	20.8	16.1	12.1

Note: Figures of sex ratio of population refer to the end of the decade.
* Recorded values adjusted for assumed levels of enumeration completeness are shown in parantheses.

Decade	Adjusted female e_0		Male-female difference in e_0 (same for two scenarios)	
	Scenario I	*Scenario II*	*Before adjustment*	*After adjustment*
1941-51	31.5	33.8	0.5	0.0
1951-61	36.8	38.5	1.4	0.5
1961-71	32.5	43.5	1.6	1.0
1971-81	49.3	49.3	0.7	0.7
1981-91	56.2	56.2	-0.6	-0.6

When the projections were rerun using the adjusted female life expectancies, the implied growth rates of female population shored up to match the corresponding census estimates for 1941-51 and 1961-71. For 1951-61, the projected growth rates were now somewhat higher than the census figures, as in the case of males. This is consistent with the belief that there were some unaccounted emigration from the country during this decade. Without this the implied gains in life expectancy between 1941-51 and 1951-61 would be too small to trigger the suggested rise in fertility.

However, the proposed corrections to the estimates of female life expectancies were found to make only a marginal difference to the projected trends in the sex ratio of population. The remaining part of the discrepancy between the projected and actual trends was traced to the assumed level of sex ratio at birth of 106 males for 100 females, as it had a cumulative effect on the calculated sex ratio of population of recent years. The basis for using this sex ratio was a large scale study of hospital births during 1949-58 (Ramachandran and Deshpande 1964). In retrospect, we find that this estimate may have had a slight upward bias, as the use of a more conventional value of 105 males for every 100 female births brings the projected trends in the population sex ratios very cbse to those seen in the census figures (see Table 6). Some discrepancies do remain in the estimated ratios for the years 1981 and 1991, which could be suggesting a slight upward bias in the SRS-based estimates of male life expectancies for 1971 - 81 and 1981-91. As these differences are quite small, and their source less obvious, no further corrections are proposed for the life expectancy estimates.

Table 6 also shows the decadal estimates of crude birth and death rates implied by the simulation exercise that employed the adjusted estimates of female life expectancies and a sex ratio at birth of 105. The sharp rise hi fertility indicated by the estimates of TFRs for the 1940s and 1950s is retained in the estimates of crude birth rate. According to the first scenario, the rise in fertility accounted for nearly half of the rise in the growth rate of population between 1941-51 and 1951-61; under the second scenario, the growth acceleration attributed to the birth rate increase is as much as two-thirds. Thus the estimates of the first scenario seem more realistic, but they imply that the level of crude birth rate in India in 1951-61 was 47.5 per 1000, in comparison to the official estimate of 41.7 and other independent estimates in the neighbourhood of 45 per 1000 (e.g., Visaria 1969; Rele 1987).

Although the estimated fall in total fertility was negligible between 1951-61 and 1961-71, projection results under both the scenarios show a decline in the crude birth rate of the order of 3 points per 1000. This is attributable to the rejuvenation of population's age structure from past declines in mortality and rise in fertility levels. This could be confirmed from the

TABLE 7. COMPARISON OF OFFICIAL ESTIMATES OF SOME KEY DEMOGRAPHIC INDICATORS WITH THE INTEGRATED SERIES BEING PROPOSED NOW, ALL-INDIA, 1941 TO 1991

Demographic indicator	*Series* *	*1941-51*	*1951-61*	7967-77	7977-57	7957-97
Fertility Measures						
TFR	O I	5.3-6.0	6.3-6.6	6.4-6.6	(4.8) 5.4	(4.2) 4.6
CBR	O I	39.9 40-44	41.7 46-48	41.2 43-44	37.2 38	(32.5) 35
Mortality Measures						
Male e_0	O I	32.4 31.5-34	41.9 37-39	46.4 43.5-44.5	(51.5) 50	(56.5) 55.5
Female e_0	O I	31.7 31.5-34	40.6 37-38.5	44.7 42.5-43.5	(50.5) 49	(56.9) 56
CDR	O I	27.4 28-32	22.8 26-28	19 21-22	15 16	(11-4) 12
Under-5 mortality (q^5)	O I	309 340-400	224 300-335	192 260-275	(201) 215	(147) 154
IMR	O I	185 230-240	146 180-210	129 150-170	(129) 135	(98) 95

Note: Figures in parentheses are from the SRS, which are treated as quasi-official. * O - Official series; I - Integrated series now proposed.

census figures (as well as from projection results) which show a rise in the proportion of population under age 15 from 38 percent in 1951 to 41 percent in 1961 and 42 percent in 1971. There also was some rise in the mean at maternity during this period which could have contributed to the fall in the crude birth rate.

The comparison of the crude birth rate estimates from the projections with those of the SRS shows a 9 percent under registration in 1971 -81 and 7 percent underreporting of births in 1981-91 by the SRS. This is consistent with the level of incompleteness arrived at from comparing TFR estimates. What however is more significant is the implied level of death under registration. The crude death rate of female population from the projection exercise is higher than the corresponding SRS estimate by 9 percent in 1971-81 and by 5 percent in 1981 -91. This has the implication that even though the age-specific mortality risks are correctly estimated by the SRS, the system underestimates total deaths as well as births in the population. What could explain this anomaly? Perhaps the births that go unreported are those that occur to usual residents outside the sample units. If the reporting of these births were not contingent upon the survival of the child, estimates of infant mortality and other indices from the system would still be reliable.

Conclusions

For recapitulation, in Table 71 bring together the integrated set of estimates of fertility and mortality proposed at several sections of this paper and compare them with the official estimates. Among the two variants of the integrated set, a slight preference may be given to the combination that implies higher levels fertility and mortality in the past. As it is self evident in the table, it is not necessary to dwell at length on differences between the official and present estimates. Instead attention may be focused on some salient features of the estimates that are proposed here.

It appears that fertility rates rose substantially during the 1950s as mortality levels began to fall. According to my estimates, total fertility rose at least by one birth per woman, and the resulting increase in the crude birth rate accounted for about half of the rise in the growth rate of population between 1941-51 and 1951-61. It is possible that some regions in India experienced this rise in the level of natural fertility more than others did, and became laggards in the process of fertility transition. In explaining the regional differences in fertility in India, greater attention ought to be given to levels that prevailed before the onset of fertility decline.

Another important implication of the present analysis is that births and deaths in the sample registration system are incomplete by about 5-10 percent. The improvement in completeness it achieved in the early 1980s has now been reversed. As such, acceleration observed in fertility decline in the late 1980s may have been largely spurious. However the estimates of age-specific mortality risks from the system may still be flawless as birth reporting is probably not contingent on child survival.

The level of underenumeration in the 1991 census appears to have been higher than in 1981, particularly among older women. The unadjusted census totals of 1981 and 1991 imply a substantial increase in the mortality risks of older women. This is contradicted by the SRS data which shows rising levels of life expectancies among them. If the SRS mortality trends were taken to be correct, female population of the 1991 census would have to be raised by about two percent for excessive underenumeration. As a result, the sex ratio of population \tv

1991 would have been around 940 females per 1000 males instead of the recorded figure of 927 in 1991 and 934 in 1981.

Although the indirect estimates of child mortality and widowhood estimators of adult mortality have found to contain significant downward biases, they appear to give correct indications of the changes in mortality level. Therefore the data needed for the indirect estimation of mortality should continued to be collected as they would contribute for a better understanding of mortality differentials and trends at national and subnational levels.

These observations on the demographic trends in India over a half century are by no means last words on the subject. To the extent that I have been able to reduce the range of uncertainty in the past levels of fertility and mortality, it is primarily due to the availability of overlapping sources of information for the more recent times. Until earnest efforts are made to improve the completeness of civil registration system, it becomes imperative to maintain the costly investments on multiple sources of demographic data.

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TABLE 1. WEIGHTING FACTORS FOR ESTIMATING THE AGE-SPECIFIC PERSON-YEARS LIVED DURING A TEN-YEAR INTERCENSAL PERIOD

Intercensal age interval	*Cohort 's age in first census*			*Cohort's age in second census*		
x to x+5	*x-10 to x-5*	*x-5 to x*	*x to x+S*	*x to x+5*	*x+5 to x+10*	*x+10 to x+15*
0-4	0.5 @	1.1 @	1.6	1.6	3.9	3.3
5-9	0.05*	1.1	2.1875	2.1875	3.9	0.3125
10-14	0.01	2.5	2.1875	2.1875	2.5	0.3125
15-19	0.3125	2.5	2.1875	2.1875	2.5	0.3125
20-24	"	"	"		"	"
	„	„	„		„	„
w to w-5	„				"	0.3125 #
w-5 +	"	-2.1875	k	2.1875	-0.3125 #	10-k

Note: If the census interval exceeds (or is less than) ten years by a small margin, age-specific growth rates may be used to adjust the census totals by age.

w The age at which the second census data is truncated.

k This is the weight to be applied to the cohort aged w-10 and over in the the first census; its value depends on the value of w; when w is 80, k is 4.35; for other values, see Bhat (1987).

@ Weight for births during 0-4 years before the survey.

* Weight for births during 5-9 years before the survey.

\# Weight for estimated survivors of the cohort aged w-10 to w-5 in the first census derived by applying the survival rate of the cohort aged 5 years younger.

TABLE 2: ESTIMATES OF EXPECTATION OF LIFE AT VARIOUS AGES DERIVED FROM VARIABLE-r METHODOLOGY WITH COHORT INTERPOLATION, INDIA, 1931-41 TO 1971 TO 1981

Decade	*0*	5	*10*	*15*	*20*	*25*	*30*	*35*	*40*	*45*	*50*	55	*60*	*65*	*70*	*Mean e,*	*CBR level **
						Age											
							Males with unadjusted census counts										
1931-41	31.2	42.0	39.3	39.1	38.6	34.4	29.4	25.6	22.6	19.6	17.0	-				44.0	46, 46
1941-51	32.3	45.0	42.3	41.0	39.3	34.4	29.0	25.0	21.8	18.7	15.8	-				46.7	43, 46
1951-61	37.8	47.9	44.1	42.3	40.3	35.7	30.5	26.4	23.1	19.9	17.1	-	-	-	-	48.4	47, 48
1961-71	44.9	52.6	48.3	46.9	45.3	41.1	36.1	31.9	27.9	24.0	20.5	17.6	14.8	12.1	10.1	53.4	46, 44
1971-81	50.6	58.3	53.0	50.7	48.7	44.5	39.2	34.0	29.2	24.9	21.3	18.3	15.3	12.4	10.1	58.1	40, 37
1981-91	55.3	60.9	54.7	52.0	50.1	46.4	41.0	35.2	29.9	25.5	21.9	19.0	16.0	13.1	10.8	59.8	36, 35
Males with adjusted census counts																	
1961-71	45.3	53.0	48.7	47.2	45.6	41.3	36.3	32.0	28.0	24.1	20.6	17.7	14.8	12.2	10.1	53.8	46, 44
1971-81	50.0	57.8	52.6	50.3	48.3	44.2	39.0	33.8	29.1	24.8	21.2	18.2	15.2	12.4	10.1	57.6	40, 37
1981-91	55.5	61.2	55.0	52.3	50.3	46.5	41.1	35.3	30.0	25.5	21.9	19.0	16.0	13.1	10.8	60.1	36, 35
							Females with unadjusted census counts										
1931-41	30.7	41.8	40.5	40.2	37.9	32.7	28.3	25.6	23.1	20.6	18.1	-	-	-	-	44.9	46, 46
1941-51	32.5	44.8	42.7	40.9	37.6	32.1	27.5	24.6	22.1	19.6	17.0	-	-	-	-	46.5	43, 46
1951-61	36.7	46.3	43.6	41.8	38.6	33.4	29.0	25.9	23.3	20.7	18.1	-	-	-	-	47.5	47, 48
1961-71	42.8	50.5	47.5	46.5	43.7	38.4	33.9	30.5	27.4	24.1	21.1	18.2	15.1	12.2	10.1	52.3	46, 44
1971-81	50.6	58.8	54.6	52.6	49.5	44.0	38.8	34.5	30.6	26.9	23.6	20.2	16.5	13.1	10.5	59.6	40, 37
1981-91	52.8	58.9	54.1	51.9	48.8	43.4	38.0	33.4	29.4	25.8	22.7	19.6	16.1	12.9	10.5	58.9	36, 35
Females with adjusted census counts																	
1961-71	44.0	51.6	48.5	47.4	44.4	39.0	34.3	30.9	27.7	24.4	21.3	18.4	15.2	12.3	10.2	53.3	46, 44
1971-81	49.3	57.4	53.4	51.5	48.5	43.3	38.2	34.1	30.3	26.7	23.3	20.0	16.3	13.0	10.5	58.3	40, 37
1981-91	56.2	62.1	56.8	54.3	50.7	44.9	39.2	34.4	30.3	26.6	23.3	20.1	16.5	13.1	10.7	61.7	36, 35

* Figures shown are the assumed birth rate levels for the first and the second parts of the intercensal period.

TABLE 3: ESTIMATES OF MORTALITY LEVELS OF MEN DERIVED FROM PROPORTION OF WIDOWS AMONG EVER-MARRIED WOMEN, ALL-INDIA, 1951 TO 1991

Census year	*Estimated parameter*	*Women's Age Interval* 20-24 x = 25	25-29 x = *30*	*30-34 x = 40*	*35-39 x = 45*	*40-44 x = 35*	*45-49 x = 50*	*50-54 x = 55*	55-59 x = *60*
	IJ l_n	0.993	0.990	0.981	0.967	0.932	0.904	0.806	0.805
!991	West gj 1989.6	66.1 Ref. year	67.9 1987.4	67.0 1985.1	66.1 1983.0	64.0 1981.2	64.2 1979.7	60.8 1978.2	64.7 1977.7
	U1$_{20}$	0.993	0.989	0.976	0.955	0.903	0.861	0.732	0.728
981	West e_s 1979.3	66.5 Ref. year	67.5 1977.0	65.8 1974.8	64.3 1972.7	61.3 1970.8	60.9 1969.2	56.6 1967.4	60.6 1966.9
	ljl$_u$	0.993	0.988	0.970	0.943	0.872	0.816	0.667	0.633
i971	West e_s 1968.9	66.5 Ref. year	67.1 1966.6	64.4 1964.3	62.7 1962.2	58.6 1960.3	57.9 1958.6	53.0 1956.5	55.7 1955.7
	l/l to*	o,992	0.982	0.951	0.906	0.814	0.741	0.586	0.552
1961	West *e}* 1958.2	65.4 Ref. year	65.0 1955.9	61.1 1953.6	58.5 1951.5	54.1 1949.4	53.1 1947.6	48.7 1945.2	51.5 1944.2
	ljl$_u$	0.975	0.962	0.909	0.879	0.770	0.729	0.570	0.563
1951*	West *e}* 1947.9	56.3 Ref. year	59.5 1945.7	54.9 1943.3	55.7 1941.1	50.9 1938.9	52.3 1937.3	47.9 1934.8	52.0 1934.1

Required interpolation of marital status data classified by 10 year **age-intervals,** 15-24, 25-34, 35-44, etc.

TABLE 4: ADJUSTED ESTIMATES OF SOME INDICATORS OF INFANT AND CHILD MORTALITY FROM THE SAMPLE REGISTRATION SYSTEM, ALL-INDIA, 1971-1994

Year	*Still birth ratio*	*Early neonatal mortality*	*Perinatal mortality rate*	*Neonatal mortality rate*	*Infant mortality rate*	*0-4 year death rate*	*Under-5 mortality q^5*
1971	17.5	50.0	67.5	89.3	144	55.8	0.228
1972	17.8	50.0	67.8	88.7	157	62.2	0.249
1973	17.5	50.0	67.5	85.0	151	57.0	0.232
1974	17.4	50.0	67.4	85.3	141	54.2	0.222
1975	17.6	50.0	67.6	90.7	153	58.5	0.237
1976	17.5	49.3	66.8	77.0	129	51.0	0.212
1977	17.2	46.5	63.7	78.5	128	50.4	0.210
1978	16.8	45.4	62.2	75.6	125	47.8	0.201
1979	16.5	42.5	59.0	67.8	116	44.7	0.190
1980	16.2	39.5	55.7	64.4	109	40.5	0.174
1981	15.8	38.8	54.6	64.7	105	39.9	0.172
1982	15.5	37.7	53.2	60.1	98	37.4	0.163
1983	15.2	38.4	53.6	61.3	99	36.1	0.158
1984	14.8	39.0	53.8	61.4	100	40.1	0.173
1985	14.5	33.6	48.1	56.0	93	37.4	0.163
1986	14.2	33.9	48.1	55.8	92	35.6	0.156
1987	13.8	36.3	50.1	56.8	94	35.0	0.153
1988	13.5	36.1	49.6	56.8	95	33.3	0.147
1989	12.7	34.5	47.2	56.4	91	29.9	0.133
1990	11.8	36.6	48.4	52.5	80	26.3	0.119
1991	10.7	35.3	46.0	51.1	80	26.5	0.120
1992	11.6	35.9	47.5	50.0	79	26.5	0.120
1993	10.5	33.7	44.2	47.1	73	23.7	0.108
1994	8.9	33.6	42.5	47.7	74	23.9	0.109