The Incidence and Age Distribution of Death: Mortality by Caste, Gender, and Sector of Origin in India in the Mid-2010s

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THE INCIDENCE AND AGE DISTRIBUTION OF DEATH: MORTALITY BY CASTE, GENDER, AND SECTOR OF ORIGIN IN INDIA IN THE MID-2010s[1]

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Abstract

This paper considers the distribution of mortality across social groups classified by caste, gender and sector of origin in India in the mid-2010s. The analysis is carried out employing micro-data on the age-distributions of population and death-rates available in the National Family Health Survey of 2015-16 (NFHS-4). Mortality in the paper is measured in terms of the crude death rate, an indicator of 'inefficiency' in the age-distribution of deaths, and an 'age-adjusted' death rate which takes account of both the mean and the dispersion of a distribution. The last-mentioned indicator is taken to be the preferred measure of mortality. The analysis in the paper suggests that mortality outcomes across castes replicate the caste hierarchy and that there is a sharp rural-urban divide in the distribution of death. Mortality sex-ratios are found to be relatively more favourable for the lower than the higher castes. The results presented in the paper are not unexpected, but they provide quantitative confirmation of one's worst suspicions regarding the skewed distribution of mortality across social groups in India.

Key Words

Crude death rate, inefficiency indicator, 'age-adjusted' death rate, caste, gender, sector of origin

JEL Classification

D30, D63, I14, I31, J17, J19

[1] The author is greatly indebted to John Creedy for valuable comments and suggestions on an earlier draft of the paper. Detailed tabulated micro-data on the age-distributions of population and death-rates by caste and religion, for males and females and rural and urban areas, were made available, on request, by the International Institute of Population Sciences, Mumbai. Many thanks are owed to K. S. James (Director: IIPS) for this collegial and helpful sharing of a valuable research resource.

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1. Introduction

This paper is on a non-income dimension of the human condition—mortality. It addresses the theme of distribution in two senses. First, it appraises the role of the age-distribution of populations and deaths in measuring mortality: this is a theoretical, or measurement-related, aspect of distributional analysis. Second, it assesses the equality, or lack thereof, in the distribution of mortality across well-defined social groups in India: this is a social-empirical aspect of distributional analysis.

In the matter of the measurement of mortality in a society, the measure most frequently resorted to is the crude death rate (CDR), defined as the proportion of deaths in a population over a specified period of time (usually a calendar year). The CDR is a simple head-count ratio, essentially a measure of central tendency. As such, it does not take account of the precise age-specific distributions of population and deaths in reckoning the extent of mortality. A relatively young population, in which the death-intensive old-age cohorts account for a relatively small proportion of the total population, is a different proposition from a population in which the aged are thicker on the ground—a difference that is not picked up by the crude death rate. One way of taking account of age-distribution is through a device called the *mortality concentration curve*, and one called the *generalized mortality concentration curve*, which are useful in deriving an 'age-adjusted death rate' that takes account of the 'inefficiency' (or 'wastefulness' of early deaths) in the distribution of deaths across age cohorts. These devices and measures draw directly on the 'Lorenz-Gini framework' which is a very familiar feature of the literature on income distribution. They have been dealt with in detail in a couple of recent papers (Creedy and S. Subramanian, 2022a, 2022b). The issues involved are explicated in the text of this paper.

On the distribution of health, mortality, and life-expectancy outcomes across socio-economic groupings of the population, there has been a large volume of work undertaken in the 2000s, notably in the recent past. Some at least of the salient studies in this field of enquiry would include Sen (1985), Maharatna (2000), S. V. Subramanian et al (2006), Mohanty and Ram (2010), Borooah (2018), Saikia et al (2019), Kumari and Mohanty (2020), Yadav et al

(2020), Gupta and Sudharsanan (2022), and Vyas et al (2022). Many of these studies—like the present one—have been enabled by the data available in successive and immensely informative rounds of the *National Family Health Survey*. The present paper confirms the discouraging but not unexpected findings of earlier studies on the well-marked differentiation in mortality according to social groups, particularly with respect to the dimension of caste.

This paper is organized as follows. Section 2 deals with the mortality and generalized mortality concentration curves, and with a small set of measures of mortality: the crude death rate (D), a measure of 'inefficiency' in the age-distribution of deaths (I_M), and an 'age-adjusted' measure of mortality (D^*) which is obtained by combining D and I_M in a composite index. The data source employed in the study is briefly described in Section 3. Sections 4, 5, 6 and 7 deal with empirical applications of the measurement concerns of Section 2 in order to present a picture of aspects of inequality in the distribution of mortality across the social categories of caste, gender and sector of origin. Section 8 concludes.

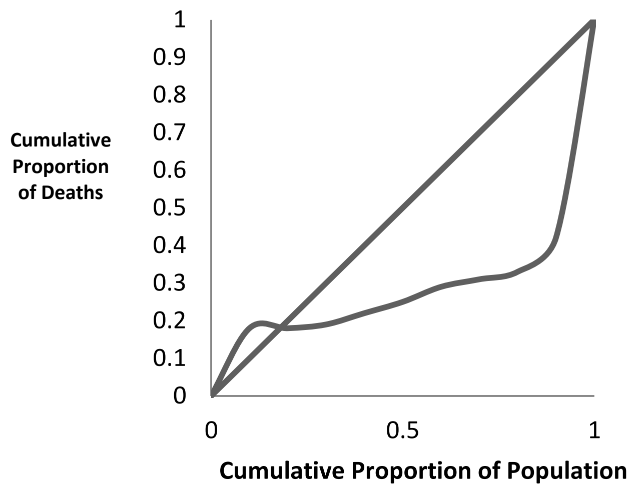
2. Mortality Concentration Curves and Mortality Measures

Consider a population of size n in which d deaths occur in the reference year under consideration. The crude death rate D is simply the proportion of all deaths in the total population: $D \equiv d/n$. Suppose the population is partitioned into K mutually exclusive and exhaustive age-groups, indexed, in ascending order of age, by $j = 1,...,K$. (The form in which data are typically available is in five-year age intervals, for example: 0-4, 5-9, 10-14,…,70-74, and 75+.) Let P_j be the cumulative proportion of the population, and Q_j the cumulative proportion of all deaths, accounted for by the *j*th youngest age-group, $j = 1,...,K$. Let $P_0 = Q_0 \equiv 0$; and, of course, $P_K = Q_K = 1$. The *mortality concentration curve*, or *M*-curve, is the graph obtained by plotting the points $(P_0, Q_0), (P_1, Q_1),...,(P_j, Q_j),...,(P_K, Q_K)$ in the unit square: it is the cumulative proportion of deaths as a function of the cumulative proportion of the population arranged in ascending order of age. The parallel with the Lorenz curve (the plot of cumulative income-share against cumulative population-share) should be obvious.

What can one say of the shape of the *M*-curve? It is clear that the curve is a non-decreasing one, going from (0,0) to (1,1) of the unit square. While the Lorenz-curve is strictly convex and never lies above the diagonal of the unit square, this need not be the case with the *M*-

curve because, unlike the case of the Lorenz curve, the *M*-curve is obtained from individuals being ranked by age rather than by age-specific death-rates: it is a sort of concentration curve. In a ‘demographically advanced’ regime, characterised by low rates of infant mortality, child mortality and fertility, one might expect the *M*-curve to lie below the diagonal of the unit square and to be uniformly convex. In relatively underdeveloped societies, one might expect the *M*-curve to have an initial concave ‘bulge’ above the diagonal, to intersect the latter, and thereafter to lie below it, so that the curve has a typically sigmoid shape, such as is featured in

Figure 1: A Possible *M*-Curve



If one subscribes to the value-judgement that a death acquires greater adverse significance the lower the age at which it occurs, then, for any given level of deaths, the ‘worst’, or ‘most inefficient’ age-distribution of deaths is one in which all the deaths are concentrated at the youngest age; and the ‘least bad’ or ‘most efficient’ distribution is one in which all the deaths occur at the oldest age. That is, the most inefficient distribution would be represented by an *M*-curve—call it the M^W-curve—which coincides with the upper-left right angle of the unit square. The most efficient distribution would be represented by an *M*-curve—call it the M^B-curve—which coincides with the lower-right right angle of the unit square (the super-script ‘*W*’ can be thought of as signifying ‘worst’, and ‘*B*’ as signifying ‘best’). The area under any actual *M*-curve is a ‘natural’ measure of a distribution’s divergence from the ‘best-case’ outcome represented by the M^B-curve. A normalized value of this divergence is given by the area under the *M*-curve as a proportion of the area enclosed by the M^W- and M^B-curves, which is simply unity. So a plausible measure of inefficiency of the age-distribution of deaths in any situation—call it I_M—is given just by the area under the *M*-curve: this area, in terms of the coordinates of the *M*-curve, can be derived by the usual 'trapezoidal approximation' method which is familiar from the calculation of the Gini coefficient of inequality from the Lorenz curve, as:

$$(1)\quad I_M = (1/2)\sum_{j=1}^{K}(P_j - P_{j-1})(Q_j + Q_{j-1}).$$

The index I_M is a measure of the 'sub-optimality' of the age-distribution of deaths, in terms of 'wasted life-years', and it lies in the interval [0,1].

Recall that a principal motivation underlying these measurement exercises is to find a way of 'adjusting' the crude death rate *D* for age-distributional considerations. This would require us to enhance the information provided by *D* on the average rate of mortality with information on the sub-optimality of the age-distribution of deaths as furnished by I_M. A simple means to this end is an index—call it $D*$--that is given by:

$$(2)\quad D^* = D(1 + I_M).$$

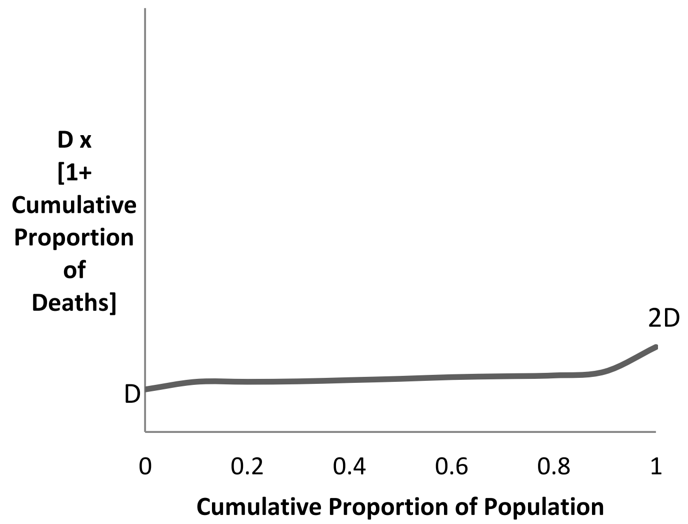
It is no accident that the measure $D*$ bears a distinct resemblance to Sen's (1976) welfare index, in the context of income distributions, given by $W = \mu(1-G)$, where μ is mean income and G is the Gini coefficient of inequality. Sen.'s welfare index (or index of 'real national income', as he called it) is also Atkinson's (1970) 'equally distributed equivalent income', derived for a situation in which the underlying social welfare function is the Borda rank-order-weighted sum of incomes. Analogously, the index $D*$ may be regarded as an 'optimally distributed equivalent death rate', with suitable contextual adaptation. In the Atkinson case, the underlying value judgement is that one is willing to trade a lower mean income for a lower level of inequality; in the case of the mortality measure, the underlying value judgement is that one is willing to trade a higher crude death rate for lower inefficiency in its distribution across age-groups.

Specifically, the welfare index *W* is a function of the *size* of the distribution (as captured by μ) and the extent of *inequality* in the distribution (as captured by G). *W* is increasing in μ, which is a 'good', and declining in G (which is a 'bad'), whence *W* = μ times (1 *minus G*).

Analogously, one could regard D^* as an abbreviated *illfare* or *loss* function: its two arguments are the size of the distribution (as captured by the mean or crude death rate *D*) and the sub-optimality of the age-specific distribution of deaths (as captured by the inefficiency coefficient I_M). D^*, being an index of illfare, ought to be an increasing function of *each* of its arguments, whence D^* = *D* times (1 *plus D*). So all that needs to be remembered is that *W* is a *welfare* function, μ is a 'good', and *G* is a 'bad'; and that D^* is an *illfare function*, with both *D* and I_M being 'bads'.

These considerations lead to the quest for a device which is an analogue of Shorrocks' (1983) *Generalized Lorenz Curve* (*GL* curve). This contrivance enables one to compare distributions in terms of their aggregate welfare content, which depends on both the size and inequality of the distributions in question. The Generalized Lorenz curve is just the Lorenz curve scaled by the mean income μ. It is well-known that the area under the Lorenz curve is (1-*G*)/2. Therefore, the area under the *GL* curve, which is just the Lorenz curve scaled by μ, should be $\mu(1-G)/2$. That is to say, Sen's welfare index is simply twice the area under the *GL* curve.

Figure 2: A Possible *GM* Curve



The preceding discussion suggests the case for a *Generalized Mortality* (*GM*) curve corresponding to the Generalized Lorenz curve, such that the area under the *GM* curve can simply be read off as the value of the abbreviated illfare function D^*. Specifically, one can derive the *GM* curve from the *M*-curve as follows: first, shift the *M*-curve up by the crude death rate *D*; then, scale the *M*-curve by multiplying by *D* (a more detailed account is available in Creedy and Subramanian, 2022a). This is shown in the accompanying Figure 2, which portrays a *GM* curve. The curve is obtained by plotting the points $\{(P_j, D(1+Q_j))\}_{j=0,\ldots,K}$. It is a non-decreasing curve which goes from the point $(0, D)$ to the point $(1, 2D)$. If one imagines a horizontal line at *D* in Figure 2, then the area enclosed by the *GM* curve and this line is DI_M, while the area below the horizontal line is just *D*. Hence, the area under the *GM* curve, as drawn in Figure 2, is a sum of two areas, given by $DI_M + D = D(1+I_M) \equiv D*$.

One can think of a situation in which for two populations 1 and 2 with crude death rates of D_1 and D_2 respectively, it is the case that $D_1 < D_2$, but when it comes to their 'age-adjusted' death rates D_1^* and D_2^* respectively, it turns out that $D_1^* > D_2^*$. The rank-reversal could happen because the inefficiency index I_M is sufficiently higher in population 1 to negate its advantage over population 2 in terms of the crude death rate. The crude death rate, by failing to take account of the precise age-distribution of population and deaths, is thus a potentially misleading indicator of comparative mortality, hence the need for a distribution-sensitive index such as $D*$, which is the mortality measure employed in this paper.

3. Data on Mortality by Social Groups.

Over the last three decades, the National Family Health Survey (NFHS) has been a major source of information for policy-makers and researchers on crucial development indicators at the household level in India, extending to the country as a whole and its constituent States. NFHS is a large-scale household sample survey that provides data on various aspects of family welfare related to demography, health and gender. The surveys are conducted under the auspices of the Ministry of Health and Family Welfare and under the overall guidance of the International Institute of Population Sciences (Mumbai), with technical assistance from ICF (USA), and funding from various international agencies. There have been five rounds of the NFHS, conducted in 1992-93, 1998-99, 2005-06, 2015-16 and 2019-20. The present study is based on NFHS-4 (2015-16).

The specific data employed in this paper are on the age-specific distributions of populations and death-rates available across different combinations of social and geographical categories. These include classifications by caste, the four caste-groups employed being the Scheduled Castes (so designated in Article 341 of the Constitution in reference to the historically marginalized and oppressed communities—erstwhile 'untouchables'—excluded from the traditional Hindu hierarchy of caste), the Scheduled Tribes (so designated in Article 342 of the Constitution in reference to specified tribal communities), the Other Backward Classes (so described in the Constitution in reference to socially and educationally backward groups of people), and the 'Others', or 'the General Category' (principally the 'upper castes' in the caste hierarchy). Distributions are also available by gender (male/female) and by sector of origin (rural/urban). Effectively, the four caste groupings are constituted by the set {Scheduled Caste, Scheduled Tribe, Other Backward Classes, Others}, the three gender groupings by the set {Female, Male, Both Sexes}, and the three sector-of-origin groupings by the set {Rural, Urban, Rural-plus-Urban}, making for a total of 4 x 3 x 3 = 36 possible combinations of caste, gender, and sector of origin sub-groups. This is a rich collection of disaggregated age-distributional data. NFHS-4 (2015-16) is a survey conducted across a household population of 601,509 persons. Some results from an analysis of selected distributions are presented in what follows.

4. Mortality and Caste

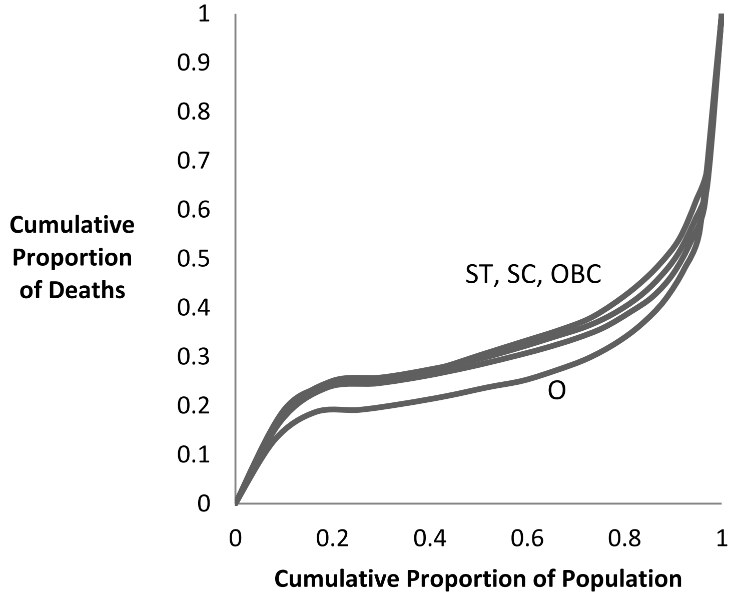
For each of the four caste-groups considered in this study, Table 1 provides information on the three mortality-related indicators mentioned earlier: the crude death rate (D), the mortality-inefficiency index (I_M), and the age-adjusted mortality measure (D^*). In everything that follows, SC stands for 'Scheduled Caste', 'ST for 'Scheduled Tribe', and OBC for 'Other Backward Classes'. As discussed earlier, the crude death rate D is not a wholly reliable measure of mortality because of its failure to take account of the age-distributions of population and deaths, and for this reason, the preferred measure is the index D^*. A case in point, as Table 1 reveals, relates to the ST population, which has the second lowest D-value of 8.3 deaths per one thousand population—lower than the OBC population's D-value of 8.5. However, it also turns out that the inefficiency index I_M for the ST population, at 0.3322, is the highest among all caste-groups. Hence, in terms of the age-adjusted mortality measure $D*$, the ST and OBC populations end up tied in second place ($D* = 11.1$). Figure 3, which features the *M*-curves for the four caste-groups, suggests an inefficiency ranking (from largest to smallest), of ST, SC, OBC and Others, in that order: this is verified by the respective I_M values. The gap between the worst- and best-performing groups is also larger for the measure $D*$ than for the measure *D*: the ratio of the *D*-values for the SC and the 'Others' groups is 1.14 (= 9.1/8.0), while the ratio of the respective $D*$-values is higher, at 1.18 (= 12.0/10.2), indicating that the SC group's relatively poor performance in terms of *D* is reinforced by its relatively poor performance in terms of I_M. The caste hierarchy is clearly reflected in the age-adjusted death-rate figures: the SC group heads the list with a $D*$-value of 12.0, followed by the ST and OBC groups tying at 11.1, and the 'Others' scoring the lowest value of 10.2. The inequities of caste are preserved in death, as in life.

Table 1: Mortality Measures Across Caste-Groups

	SC	ST	OBC	Others
D	9.1	8.3	8.5	8.0
I_M	0.3240	0.3322	0.3110	0.2733
$D*$	12.0	11.1	11.1	10.2

Source: Computed from NFHS-4 (2015-16) micro-data.

Figure 3: *M*-Curves: SC, ST, OBC and Others



Note: The figure suggests that the 'Others (O)' *M*-curve dominates the OBC *M*-curve, which dominates the SC *M*-curve, which dominates the ST *M*-curve.
Source: Based on computations from NFHS-4 (2015-16) micro-data.

5. Mortality and Gender

5.1. The Aggregate Picture

Table 2 summarises information on mortality-related indicators separately for males and females. The record for females is better than for males with respect to each of the indices *D*, I_M and *D**, and this is confirmed, in terms of the male and female *M*-curves, by Figure 4. Any direct comparison of male and female death rates as a guide to, or symptom of, 'unfair distribution' is rendered problematic by the possibility that females have a natural, or biologically determined, edge over males in the matter of survival chances. For example, Desjardins (2004) says: 'The genetic advantage of females is evident…the genetic difference between the sexes is associated with a better resistance to biological ageing.' Thus, the observation that the sex-ratio of deaths is in favour of females is not in itself a remarkable piece of social commentary.

What could, however, be a matter of interest is in exploring if, and why, there are differences across populations in the sex ratio of mortality. This is suggested by an observation of

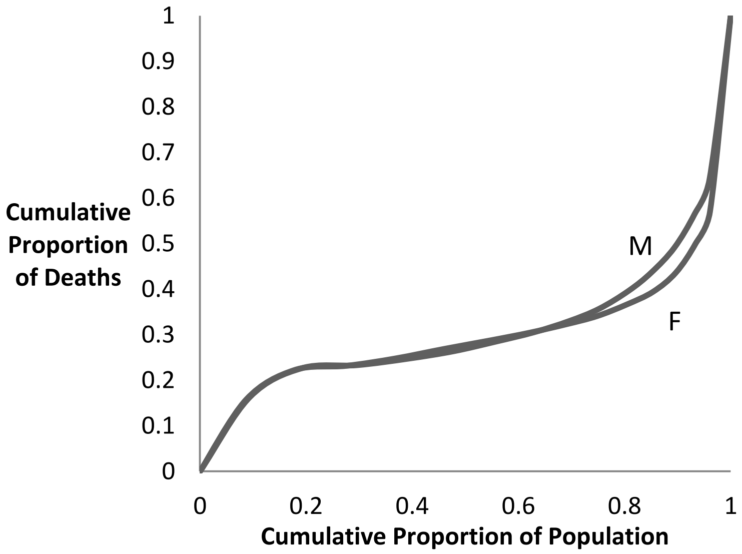
Desjardins' (2004): 'Even though many biological and genetic factors have been identified, their overall effect is impossible to measure, *especially given the influence of social factors on mortality*' [emphasis added]. In an exploration of sex differentials in life expectancy across a set of industrialized economies, Travolta and Lalu (1996) have found a narrowing of the differential (in favour of females) from the early 1970s to 1990s, suggesting, among other things, the possibility of faster gains in life expectancy accruing to males. In general, demographically and economically 'advanced' countries might be expected to have higher female-to-male mortality ratios. This is reflected, for instance, in the fact (as Table 2 reveals), that the ratio of female-to-male D*-values for India in 2015-16 was 0.78 (= 9.7/12.4), while this ratio for New Zealand in 2019—see Creedy and Subramanian (2022b)—was of the order of 0.92 (= 7.6/8.4). In what follows, the variability of mortality sex ratios is explored further in the context of caste in India.

Table 2: Mortality Measures Across The Sexes

	Male	Female
D	9.5	7.5
I_M	0.3070	0.2975
$D*$	12.4	9.7

Source: Computed from NFHS-4 (2015-16) Report: Tables 2.12 and 12.9

Figure 4: *M*-Curves: Males and Females



Source: Based on computations from NFHS-4 (2015-16) Report: Tables 2.12 and 12.9

5.2 Mortality Sex Ratios Across Caste Groups

Table 3 has results on the age-adjusted mortality index *D** for each combination of caste-group and sex, as also on the female-to-male ratio of *D**. The relevant mortality sex ratios for the SC and ST groups, at 0.73 and 0.71 respectively, are substantially lower than for the OBCs and 'Others', at 0.80 and 0.84 respectively.

Table 3: *D Measures and Mortality Sex Ratios Across Caste Groups**

	SC	ST	OBC	Others
Male	13.9	12.9	12.3	11.1
Female	10.2	9.1	9.8	9.3
D* Sex Ratio	0.73	0.71	0.80	0.84

Source: Computed from NFHS-4 (2015-16) micro-data.

There could be at least three reasons for the observed pattern of sex ratios across caste groups. The first has to do with cultural factors. The tradition of gender equality has for long been a feature among the Scheduled Tribes: on this, see in particular Maharatna (2000; p. 1333): '…gender relations among Indian tribes have historically been more balanced and egalitarian…'. To the extent that 'son-preference' as an indicator of anti-female bias is reflected in sex-specific mortality, the Scheduled Tribes would again seem to be an exception to the rule among higher castes: Yadav et al (2020; p. 1142), in an analysis of NFHS-4 data, suggest that '…In caste terms, Scheduled Tribe women have a lower desire for sons as compared to women from other castes. This may be due to the higher gender egalitarian norms among tribal communities in India.' Pande and Astone (2007; pp. 5-6) also see caste as playing a role in son-preference: 'Caste may also be associated with cultural practices that influence women's roles, and thus son preference, such that one may expect less son preference among lower castes and tribals than among high castes. Compared to lower castes, higher castes have more rigid gender stratification systems, with strictly enforced rules of seclusion… for women…and greater use of dowry. Lower caste and tribal women may have fewer restrictions placed on their movement or employment outside the home…'[2]

A second reason could have to do with a relatively larger incidence of male deaths from other than natural causes among lower caste populations. In this context, the reader is referred to Lewnard et al (2022; p.471) : 'Unintentional injuries are the leading single cause of death among Indian adults younger than 40 years, particularly among men, and occur disproportionately in communities of lower socioeconomic status.'[3]

A third reason has to do with the possibility that as general reductions in mortality occur, the gains accrue disproportionately to males, which is one of the theses explored by Travolta and Lalu (1996) in exploring an over-time reduction in relative female life expectancy advantage in some advanced industrialised countries. Sen (1987 [1985]; p.61) makes a similar diagnosis about the secular decline, over much of the 20th century, in the sex (female-to-male) ratio of life expectancy in India: 'Indeed, with economic and social progress, as the *absolute*

[2] These observations are—or should be—a salutary deterrent to uninformed and casual claims to cultural superiority by the higher castes.

[3] A case in point is death due to poisoning in sewage pits among sanitation workers who are drawn almost entirely from the Scheduled Caste community. This is a casualty of the brutal caste-based occupation of manual scavenging (now outlawed in the statute books), though deaths from this cause continue to occur.

positions of both Indian men and Indian women have improved, the relative position of Indian women seems to have fallen behind. If we judge well-being by the capability to live long, women's well-being has fallen vis-à-vis men's, even though absolutely both have increased substantially.'

These observations on time-series comparisons have similar implications for cross-section comparisons. Tables 1 and 2 reveal that as we move up the caste hierarchy from the Scheduled Castes to the OBCs, there is a 7.5 per cent decline in the adjusted mortality measure D^*, but while the decline for males is 11.5 per cent, that for females is only 3.9 per cent. The difference between the Scheduled Castes and 'Others' is reflected in an overall D^* measure for the 'Others' which is lower by 15 per cent, but while the reduction is as much as 20.1 per cent for males, it is only 8.8 per cent for females. As we move up the caste hierarchy, both males and females perform better on the mortality front, but males fare proportionately better than females.

6. The Rural-Urban Divide

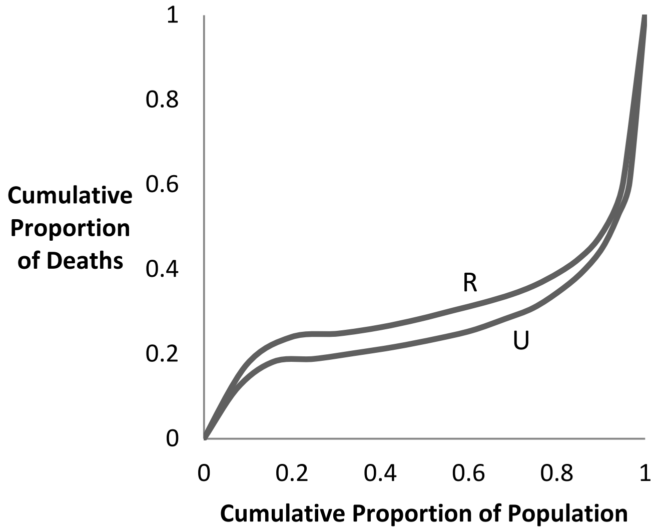
Results at the aggregate level on the three mortality indicators, separately for the rural and urban areas of the country, are presented in Table 5. Figure 5 features *M*-curves for the two sectors of origin. The picture is one of a substantially worse rural than urban record: the relatively poor rural performance with respect to the crude death rate is reinforced by its relatively poor performance with respect to the inefficiency indicator, resulting in a widening of the rural-urban gap in terms of the D^* measure vis-à-vis the D measure.

Table 5: Mortality Measures Across Rural and Urban Areas

	Rural	Urban
D	9.0	7.4
I_M	0.3138	0.2685
D^*	11.9	9.4

Source: Computed from NFHS-4 (2015-16) Report: Tables 2.12 and 12.9

Figure 5: *M*-Curves: Rural and Urban



Source: Based on computations from NFHS-4 (2015-16) Report: Tables 2.12 and 12.9

A more disaggregated picture of the rural-urban divide is available in Table 6, which presents information on the age-adjusted mortality measure *D** for each of eight social groups combining caste, gender and sector-of-origin. Reading the data in Table 6 column-wise enables one to see that for each pair of groupings in which only the sector of origin varies, the rural group in question displays a higher value of the mortality measure than the urban group. The rural-urban divide is systematic and pervasive across the relevant partitions of the population, providing confirmation, in the instant case, for Lipton's (1977) thesis of 'urban bias'.

Table 6: The Rural-Urban Divide in the *D Measure for Different Caste-Gender Groups**

	(SC,M)	(ST,M)	(OBC,M)	(Others,M)	(SC,F)	(ST,F)	(OBC,F)	(Others,F)
R	15.2	13.0	13.1	12.1	10.7	9.4	10.7	10.5
U	11.7	11.6	10.8	9.8	8.8	7.4	8.0	7.7

Source: Computed from NFHS-4 (2015-16) micro-data.

7. The Ratchet Effect of Reinforcing Disadvantages

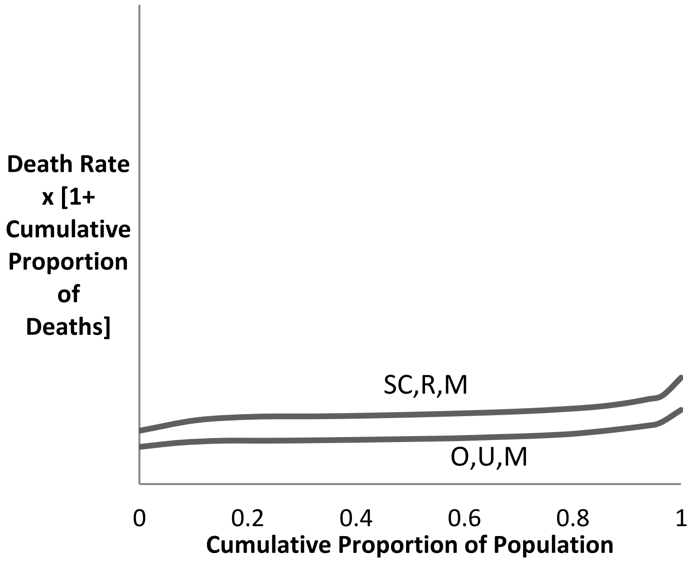
The preceding sections offer a picture of how, other things equal, caste, gender and sector of origin can and do affect mortality outcomes. Other things equal, mortality outcomes are worse for lower caste populations, as they are for populations of rural origin. (As stressed in Section 5, a direct male-female comparison is not meaningful because it is difficult to unscramble biological factors assisting women and social factors discriminating against women.) When caste and sector of origin are combined, one obtains a view of disparity across population groups which highlights the double disadvantage of being rural *and* low caste in contrast to the double advantage of being urban *and* high caste. These ratchet effects of criss-crossing identity compound the relative disadvantage or advantage arising from a person's separate identities associated with each of a number of the person's group affiliations. Such an 'intersectional' view of the experience of death in India is available from a pair of figures in Table 6, those relating to the D^* measure of 15.2 for the rural, male, Scheduled Caste population, and of 9.8 for the urban, male, 'Others' population. The ratio of the one to the other is 1.55. The mortality divide between the two groups is visually captured in the gap between their respective Generalized Mortality curves, as featured in Figure 6. The two populations might well be inhabiting different universes. But there is a certain inevitability to this outcome when one considers that the experience of death is but a logical culmination of the experience of life. The latter is reflected in some summary statistics on group-related indicators of human development, as provided in Table 7.

Table 7: Select Human Development Indicators for Rural, Urban, Scheduled Caste and 'Others' Populations : 2015-16

% Distribution of households by source of drinking water	*Piped into dwelling/yard/plot*	*Unprotected dug well*	
Rural	18.4	6.0	
Urban	52.1	1.0	
% Distribution of households by access to electricity	*With Access*	*Without Access*	
Rural	83.2	16.8	
Urban	97.5	2.5	
% Distribution of households by type of house	*Katcha*	*Pucca*	
Rural	8.1	41.2	
Urban	0.9	84.5	
% Distribution of households by type of cooking fuel	*LPG/Natural Gas*	*Dung Cakes*	
Rural	23.0	10.2	
Urban	78.3	1.5	
% Distribution of households by type of toilet facility	*Improved, non-shared facility*	*No facility*	
Rural	36.7	54.1	
Urban	70.3	10.5	
% Distribution of households possessing various goods	*Mattress*	*Pressure Cooker*	*Any Television*
Rural	58.4	42.2	53.5
Urban	82.3	83.6	81.0
% Distribution of de facto household population ≥ age 6 years by highest number of years of school completed	*No Schooling*	*At least 12 Years*	
Scheduled Caste	36.2	9.6	
Others	21.5	20.6	
Rural	36.8	8.7	
Urban	19.2	23.8	
% Distribution of de jure population by wealth quintiles	*Lowest Quintile*	*Highest Quintile*	
Scheduled Caste	25.9	11.3	
Others	9.4	34.0	

Source: Various Tables of National Family Health Survey (NFHS-4), 2015-16

Figure 6: Generalized *M*-curves for Scheduled Caste Rural Male and 'Other' Urban Male Populations



Source: Based on computations from NFHS-4 (2015-16) micro-data.

8. Conclusions

This paper has been concerned with quantifying the distribution of mortality in India in the mid-2010s across the social categories of caste, gender and sector-of-origin. Mortality has been measured in terms of the crude death rate *D* and an 'age-adjusted' death rate *D** which takes account of the age-distribution of population and deaths. This is done by combining the crude death rate with a measure of 'inefficiency' I_M in the age-distribution of deaths that relies on the value-judgement of relative aversion to young-age deaths. By virtue of its taking account of both central tendency and dispersion considerations, *D** is the preferred measure of mortality.

The data employed in the paper are the detailed caste, gender and sector-of-origin age-distributions of population and death rates available in the Report of the National Family Health Survey (2017) (NFHS-4).

In terms of the age-adjusted mortality measure *D**, the Scheduled Castes fare worst, followed by the Scheduled Tribes and Other Backward Castes, while the higher castes constituting the 'Others' display the lowest mortality level. A direct comparison of male and female mortality is rendered somewhat meaningless by the fact that women have a genetic advantage in the matter of survival over men. However, it is interesting to examine variations in the mortality sex-ratio (female-to-male). This ratio, in terms of the D^*-measure, is found to be generally lower for the backward compared to the forward caste groups. The phenomenon could arise for three reasons: (a) a culture of lower anti-female bias among the Scheduled Castes and Tribes compared with the OBCs and Others; (b) a greater incidence of under-40 male mortality from accidental injuries among the relatively backward socioeconomic groups; and (c) the tendency for males to benefit disproportionately from the gains of general improvement on the mortality front (as on other fronts): as one goes up the caste hierarchy, mortality rates decline, but proportionately more for males than for females.

Lipton's (1977) thesis of 'urban bias' is borne out, in the matter of mortality, by a systematic and pervasive record of worse mortality outcomes in the rural compared with the urban areas of the country. This is the case both at the aggregate level and at the disaggregated level of groups differentiated by caste and gender.

Additionally, the paper notes the stark disparity between the mortality outcomes for the rural male Scheduled Caste group and the urban male Others group: this reflects the ratchet effect of mutually re-enforcing disadvantages arising from particular combinations of group affiliation. Such outcomes in the space of mortality are found to faithfully replicate the caste and sector of origin distributions of the benefits and burdens of human development.

Finally, the essay also offers a pair of pictorial devices—the mortality concentration curve and the generalized mortality concentration curve—along with numerical measures which enable more reliable mortality comparisons, ones that have the advantage of lending themselves to easy visual interpretation.

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