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Are socioeconomic inequalities in mortality decreasing or increasing within some British regions? An observational study, 1990–1998

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Abstract

Background This paper evaluates claims in a recent study that inequalities in small area mortality rates have lessened. We examine the effect of differently estimated populations on time trends in age-specific mortality rates for Yorkshire and the Humber and East of England.

Methods Populations were estimated for wards using four methods that introduce increasing amounts of information. Age-specific mortality rates for age-groups 45–54, 55–64, 65–74 and 75–84 for both sexes were calculated for population-weighted deprivation quintiles. Inequality was tracked using ratios of rates in the most deprived quintile divided by those in the least.

Results When constant 1991 populations are used, rate ratios decrease for all age–sex groups, indicating shrinking inequality. When a method adjusting small area populations to official district estimates is used, both decreases and increases are observed in the mortality rate ratios. These results differ from Trent region findings of decreases in inequality. When small area populations are cohort-survived and adjusted to district populations, most differences in rate ratios indicate increasing inequality. When a method is used that includes information on migration and special populations, then seven out of eight age–sex groups exhibit increasing inequality.

Conclusions A judgement about trends in mortality inequality is highly dependent upon the denominator population used. Simpler estimation methods result in convergence of rate ratios, whereas more sophisticated methods result in increasing inequalities in most age–sex groups.

Keywords: population estimates, estimation methods, mortality, inequalities

Introduction

The health of the nation continues to improve: UK life expectancies increased at around 0.2 years per year during the 1990s (faster for males, slower for females). However, there is evidence[1] that inequalities across social classes in life expectancy have consistently increased since 1971. When the spatial pattern of the standardized mortality ratios is examined for areas grouped by deprivation rank, the trend is for increasing inequality for parliamentary constituencies in Great Britain,[2] a trend also observed for electoral wards in Yorkshire and the Humber.[3] The improvement in mortality experience of people living in the poorest areas has been lower than that of people in the richer areas, although the degree of difference has varied with the period studied.

Strong *et al.*[4] have challenged this view. They studied trends in small area age-specific mortality rates (ASMRs) for males and females aged 45–84 living in the Trent (health) region, classified by the Townsend deprivation rank (quintile), using the Enumeration District (ED) at the 1991 Census as the fixed geographical unit. They found evidence in seven out of eight age–sex groups studied of a narrowing gap between ASMRs in the least and most deprived quintiles. Those researchers, however, qualified their findings saying that: 'Possible artefactual explanations include changes in deprivation levels in EDs since 1991, population movements across EDs and increasing error in the denominator population estimation' (p. 122).[4]

This denominator error will relate to the method Strong *et al.* used to derive the populations at risk for their ASMRs. Base populations comprising age–sex counts (corrected for undercount) for 1991 Census EDs were scaled using health authority level age- and sex-specific mid-year estimates for other years in their study period to allow for population ageing, survival, births and migration each year at the health authority level. This approach imposes these demographic events in proportion to population size across the small area EDs that comprise the larger area health authority geography. However, this will be inappropriate as persons distributed throughout an area are not equally at risk of experiencing births, deaths and migration

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events. A concentration of population through migration into a new housing estate, for example, will erroneously be spread across all areas. The further away in time from a census this approach is used, the more dated the assumptions become about larger area change being applicable to constituent smaller areas and the more any error is compounded.

In this paper, we investigate the effects of differently estimated populations on measured inequality trends for the period 1990–1998 in two Government Office Regions (GORs), Yorkshire and the Humber[5] and the East of England.[6] These GORs are the focus of wider research by the workers who have devised population estimates using various approaches to evaluate the impact of estimation method on mortality and morbidity rates. For this paper, initially we replicate the Strong *et al.* analysis as far as possible and then, as it is expected that differently estimated denominators result in different ASMRs, extend it by considering the effect of using different estimation methods to derive ward populations (temporally harmonized to 1998 boundaries[7,8]).

Methods

For Yorkshire and the Humber region three estimates of ward populations by single years of age and sex at each mid-year from 1990 to 1998 were produced.[5] For the East of England estimates were generated using the same three methods together with a fourth, more complex method,[6] again for 1990–1998. A 1991 base population for both GORs and all methods was generated by the Estimating with Confidence Project.[9,10] Known as EwCPOP, these are ward age–sex mid-year estimates (MYEs) adjusted for underenumeration, which are available via the MIMAS service of the University of Manchester.[11] The estimation methods used to provide populations at risk are as follows:

Method 1. The 1991 EwCPOP 1991 MYEs for wards are held constant throughout the study period.

Method 2. The 1991 EwCPOP ward populations are adjusted to Office for National Statistics (ONS) Mid-Year Estimates for local authority (LA) district populations in the relevant year. This method is equivalent to the approach of Strong *et al.* to update their ED populations that is described above.

Method 3. A partial cohort-component model is used based on survival rates of ward-specific mortality, allowing for births in each ward and then adjusting through Iterative Proportional Fitting (IPF)[12–14] to electorate-based total ward populations and ONS district MYEs by age-group and sex.

Method 4. A full cohort-component model is employed with survival, births and migration components for single years of age with separate treatment for special populations of students, armed forces and communal establishments together with IPF adjustment to electorate-based total ward populations and ONS MYEs.

The enhancement offered by Methods 3 and 4 over the simpler approaches of Methods 1 and 2 is that increasing detail about ward-level demographic events is included. The utilization of vital statistics on births and deaths in estimates for each ward is more appropriate than apportioning equivalent larger area information amongst all constituent small areas. Electoral registers, although imperfect, can be used as indicators of ward-level change (through migration, for example), relative to other wards within each district. Scaling small area estimates to ONS MYEs ensures consistency with official data but, as noted above, district-level age–sex estimates should not necessarily be imposed across all constituent wards. IPF is a mathematical procedure that ensures a table of data is scaled so that the table's row and column totals agree with constraining row and column totals obtained from alternative sources. For Method 3, initial age–sex counts are estimated using a partial cohort-component model including ward-level births and deaths data, and in Method 4 initial counts are estimated using a full cohort-component model with further information on ward-level migration and special population sub-groups. These counts are then made consistent using IPF with electorate-derived ward total populations and LA district age–sex counts. IPF places equal importance on both the ward totals and district information and so has the advantage that neither data source is dominant and consistency with ONS MYEs is retained.

Using these methods, three (Yorkshire and Humberside) or four (East of England) estimates of age–sex-specific populations were computed and aggregated along with geographically harmonized all-cause deaths into deprivation quintile counts from which ASMRs were computed for 10 year age-groups from 45–54 to 75–84.

Results

The ratios (and 95 per cent confidence intervals) of the most deprived quintile 5 (Q5) to least deprived quintile 1 (Q1) rates are reported in the Table for Yorkshire and the Humber GOR and East of England GOR, along with the differences in the ratios between the two time periods 1990–1992 and 1996–1998. When Method 1 with constant populations is used the changes in rate ratios in all but one of the age–sex groups are negative, indicating decreasing inequality. When Method 2 is used, both decreases and increases are observed in the mortality rate ratios. These results differ from those for Trent health region, where Strong *et al.* reported decreases in inequality in seven out of eight groups. Method 3 explicitly allows for ageing, births and survival at small area scale and indirectly includes a contribution from migration through the adjustment to ward electoral populations.[7] Most of the resulting differences in rate ratios are positive: six out of eight in Yorkshire and the Humber and seven out of eight in the East of England. When Method 4 is used in the East of England analysis then seven out of eight age–sex groups exhibit increasing inequality.

The difference in rate ratios in 1990–1992 compared with 1996–1998 observed in both GORs does not differ greatly from zero. Figure 1 illustrates the ASMRs by 10-year age-groups for deprivation Q1 and Q5 with 95 per cent confidence intervals indicated. Graphs relating to Methods 2 and 4 are given for East of England GOR, but the same pattern of results is found in Yorkshire and the Humber and for all estimation methods. For each period 1990–1992 and 1996–1998, because there are non-overlapping confidence intervals, there is a significant difference in the ASMRs between the least and most deprived locations for all age-groups except females aged 75–84 in East of England. In Yorkshire and the Humber there is a greater disparity between

Table 1 Ratios of age-specific mortality rates 1990–1992 and 1996–1998 calculated using different population estimate methods for wards in the most deprived quintile to those in the least deprived quintile

		m 45–54	m 55–64	m 65–74	m 75–84	f 45–54	f 55–64	f 65–74	f 75–84
Yorkshire and the Humber									
Method 1: constant 1991 EwCPOP estimates									
1990–1992	Rate ratio	2.11	1.69	1.37	1.17	1.86	1.69	1.42	1.12
	CI LL–UL	(1.84–2.38)	(1.55–1.82)	(1.29–1.45)	(1.10–1.24)	(1.54–2.19)	(1.51–1.86)	(1.32–1.52)	(1.05–1.18)
1996–1998	Rate ratio	1.90	1.64	1.32	1.14	1.59	1.54	1.39	1.14
	CI LL–UL	(1.65–2.15)	(1.49–1.79)	(1.24–1.40)	(1.07–1.21)	(1.28–1.90)	(1.33–1.74)	(1.27–1.51)	(1.07–1.22)
	Difference	–0.20	–0.04	–0.05	–0.03	–0.27	–0.14	–0.03	0.02
Method 2: 1991 EwCPOP estimates constrained to ONS MYE populations for other years									
1990–1992	Rate ratio	2.14	1.69	1.38	1.18	1.88	1.69	1.43	1.12
	CI LL–UL	(1.87–2.41)	(1.56–1.83)	(1.30–1.46)	(1.11–1.25)	(1.56–2.20)	(1.51–1.87)	(1.33–1.53)	(1.06–1.19)
1996–1998	Rate ratio	1.98	1.72	1.40	1.14	1.62	1.63	1.46	1.18
	CI LL–UL	(1.73–2.24)	(1.57–1.87)	(1.31–1.49)	(1.07–1.21)	(1.30–1.94)	(1.41–1.84)	(1.34–1.58)	(1.10–1.25)
	Difference	–0.15	0.03	0.02	–0.04	–0.25	–0.06	0.03	0.06
Method 3: partial cohort-component with IPF to electorate-based and ONS MYE populations for other years									
1990–1992	Rate ratio	2.13	1.71	1.39	1.18	1.89	1.71	1.44	1.13
	CI LL–UL	(1.86–2.40)	(1.57–1.85)	(1.31–1.47)	(1.11–1.25)	(1.57–2.22)	(1.54–1.89)	(1.34–1.54)	(1.06–1.19)
1996–1998	Rate ratio	2.15	2.01	1.53	1.17	1.66	1.87	1.65	1.28
	CI LL–UL	(1.88–2.42)	(1.84–2.18)	(1.44–1.62)	(1.10–1.24)	(1.34–1.98)	(1.64–2.10)	(1.52–1.78)	(1.20–1.36)
	Difference	0.02	0.30	0.14	–0.01	–0.23	0.16	0.21	0.15
East of England									
Method 1: constant 1991 EwCPOP estimates									
1990–1992	Rate ratio	2.08	1.61	1.33	1.13	1.50	1.37	1.26	1.08
	CI LL–UL	(1.73–2.42)	(1.43–1.79)	(1.23–1.43)	(1.05–1.21)	(1.15–1.85)	(1.16–1.57)	(1.13–1.38)	(1.00–1.16)
1996–1998	Rate ratio	2.06	1.55	1.26	1.12	1.57	1.28	1.26	1.00
	CI LL–UL	(1.71–2.40)	(1.36–1.73)	(1.15–1.36)	(1.04–1.20)	(1.24–1.90)	(1.06–1.50)	(1.13–1.38)	(0.92–1.08)
	Difference	–0.01	–0.06	–0.07	–0.00	0.067	–0.08	–0.00	–0.08
Method 2: 1991 EwCPOP estimates constrained to ONS MYE populations for other years									
1990–1992	Rate ratio	2.08	1.61	1.33	1.13	1.50	1.36	1.26	1.08
	CI LL–UL	(1.73–2.42)	(1.43–1.79)	(1.23–1.44)	(1.05–1.21)	(1.15–1.85)	(1.16–1.57)	(1.14–1.38)	(1.01–1.16)
1996–1998	Rate ratio	2.11	1.63	1.32	1.16	1.62	1.37	1.30	1.03
	CI LL–UL	(1.76–2.46)	(1.44–1.83)	(1.21–1.42)	(1.08–1.24)	(1.29–1.96)	(1.14–1.60)	(1.17–1.43)	(0.96–1.11)
	Difference	0.03	0.02	–0.01	0.03	0.12	0.01	0.04	–0.05
Method 3: partial cohort-component with IPF to electorate-based and ONS MYE populations for other years									
1990–1992	Rate ratio	2.11	1.66	1.36	1.15	1.53	1.40	1.28	1.10
	CI LL–UL	(1.76–2.46)	(1.48–1.84)	(1.26–1.47)	(1.07–1.23)	(1.18–1.89)	(1.19–1.61)	(1.16–1.41)	(1.02–1.18)
1996–1998	Rate ratio	2.21	1.92	1.57	1.26	1.78	1.63	1.51	1.10
	CI LL–UL	(1.84–2.57)	(1.69–2.14)	(1.45–1.69)	(1.17–1.34)	(1.42–2.14)	(1.36–1.89)	(1.37–1.66)	(1.02–1.18)
	Difference	0.10	0.26	0.21	0.11	0.25	0.23	0.23	–0.00
Method 4: full cohort-component with special populations and IPF to electorate-based and ONS MYE populations									
1990–1992	Rate ratio	2.10	1.64	1.35	1.14	1.53	1.40	1.27	1.09
	CI LL–UL	(1.75–2.45)	(1.47–1.82)	(1.25–1.45)	(1.06–1.22)	(1.17–1.89)	(1.19–1.61)	(1.15–1.40)	(1.01–1.17)
1996–1998	Rate ratio	2.15	1.82	1.49	1.19	1.76	1.59	1.44	1.03
	CI LL–UL	(1.79–2.51)	(1.61–2.03)	(1.37–1.61)	(1.11–1.27)	(1.40–2.12)	(1.34–1.85)	(1.31–1.58)	(0.96–1.11)
	Difference	0.05	0.18	0.14	0.05	0.23	0.19	0.17	–0.06

Source: computations based on mortality and population data reported in Refs 5 and 6.

Deprivation in Yorkshire and the Humber is measured using the Townsend Index and in East of England using the Carstairs Index. The confidence intervals (CIs) lower limit (LL) to upper limit (UL) for the ratio of Q5 to Q1 ASMRs were computed as follows. Standard errors (SE) of the differences between Q5 and Q1 rates were calculated using the conventional formula (Ref. 15, p. 130). The UL of the ratio is the Q1 rate plus the difference plus 1.96 times the SE of the difference divided by the Q1 rate. The LL of the ratio is the Q1 rate plus the difference minus 1.96 times the SE of the difference divided by the Q1 rate.

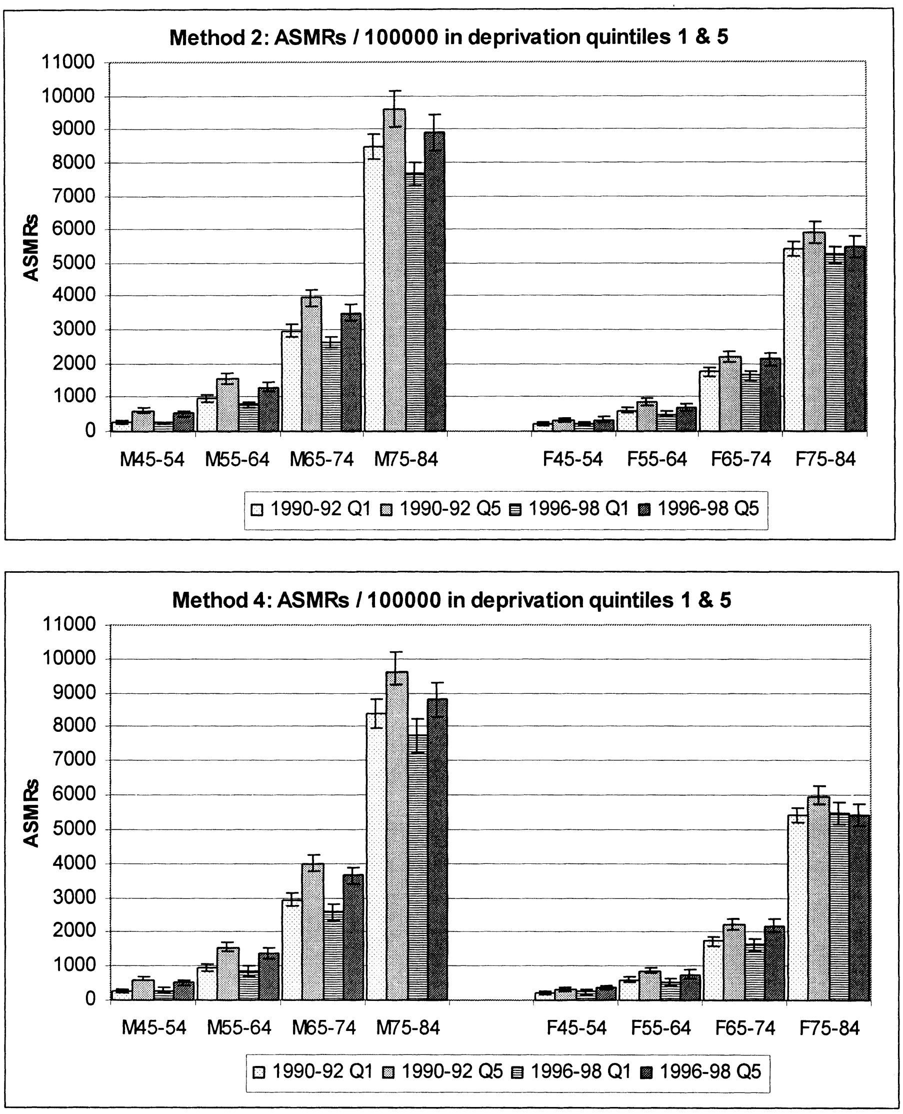

Figure 1 ASMRs in deprivation quintiles 1 and 5 in East of England GOR. Deprivation is measured using the Carstairs Index. Error bars are 95 per cent CIs. Source: computations based on mortality and population data reported in Ref. 6.

the ASMRs in deprivation Q1 and Q5. For all estimation methods, over the 9 year period, reductions in mortality are observed within both deprived and non-deprived locations with a slightly greater improvement in Q1.

It is useful to decompose elements of the ASMR calculations that influence the variations in rate ratios; example data from East of England GOR will be used. First the outputs of the various estimation methods are compared. Method 1 is a constant 1991 population; the outputs of Methods 2–4 are for 1998, as differences are likely to appear further away from the base year of the estimates time-series. Figure 2 illustrates ratios of age–sex population counts from estimation Methods 1, 3 and 4 to Method 2, equivalent to the approach of Strong *et al.* Method 1 is the population in 1991 and, with a net reduction of population by 1998, the ratios with Method 2 are largely less than 1.00. The apportionment of district-level change amongst constituent wards in Method 2 results in little relative change in the age–sex structure of the wards between each deprivation quintile. The ratios of Method 3 and 4 to Method 2 indicate a moderate change in the age–sex structures of quintiles 2–4 and reasonable consistency between the 45–54 and 75–84 age-groups. However, in Q1 a relatively larger population aged 65–74 is estimated by Methods 3 and 4 and a relatively smaller population aged 55–64 in Q5. Overall, Methods 3 and 4 are suggesting a shift in population during the 1990s from the more to the less deprived wards.

So does a differently estimated population lead to a different ASMR? In Figure 3 for males aged 65–74 in deprivation quintile 1 a higher rate for Method 1 than 2 is shown with lower rates when Methods 3 and 4 are used for the population at risk. These lower rates result because a larger population is estimated. In Q5

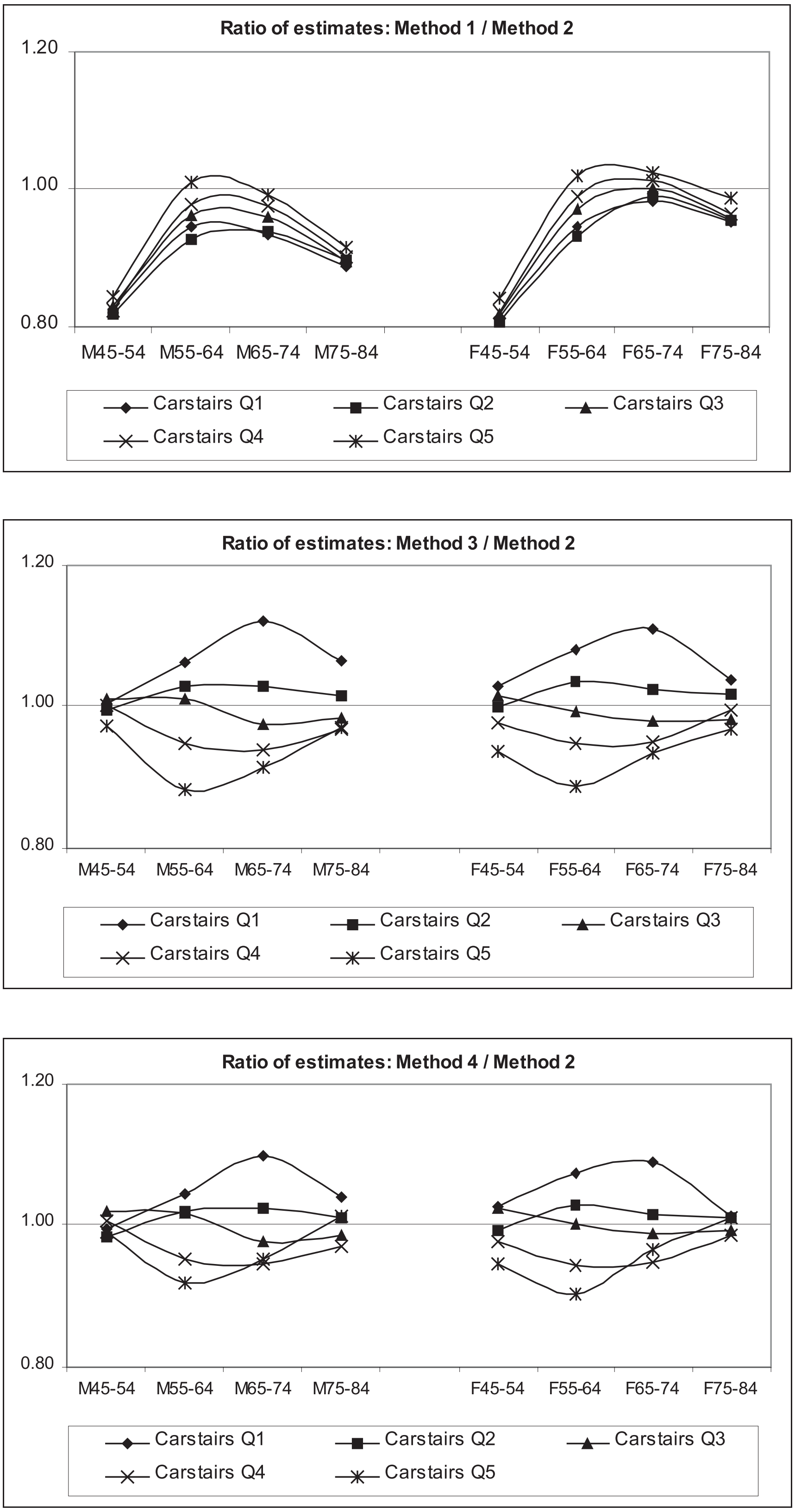

Figure 2 Comparisons of age–sex population counts estimated using various methods for East of England GOR. Deprivation is measured using the Carstairs Index. Source: computations based on mortality and population data reported in Ref. 6.

for males aged 55–64 similar ASMRs result for both Methods 1 and 2 but higher rates for Methods 3 and 4 are due to the smaller estimated populations. The 95 per cent confidence intervals show some overlap for males aged 55–64, suggesting the differences between the calculated rates may not be significant, despite the different denominator populations. For males aged 65–74 in Q1 there is almost no overlap between Methods 1 and 2 and Methods 3 and 4, and in these cases there are significant differences in the rates. The same situations exist for females in the same age-groups and quintiles, but in other age–sex groups, although there are variations in the ASMRs, the differences may not be significant.

Discussion

The analysis shows that a judgement about whether inequality is worsening or improving at small area scale is highly dependent on the method for estimating small area populations. We emulated the approach of Strong *et al.* in calculating ASMRs but did not find the tendency for decreasing inequality they identified in Trent. Our analysis, however, has not matched exactly the work of Strong *et al*. They worked with EDs whereas we worked with wards and in East of England deprivation was measured using the Carstairs Index. Some of the differences could, of course, result from different conditions in the study regions. Where more sophisticated methods are used, except for the elderly, the evi-

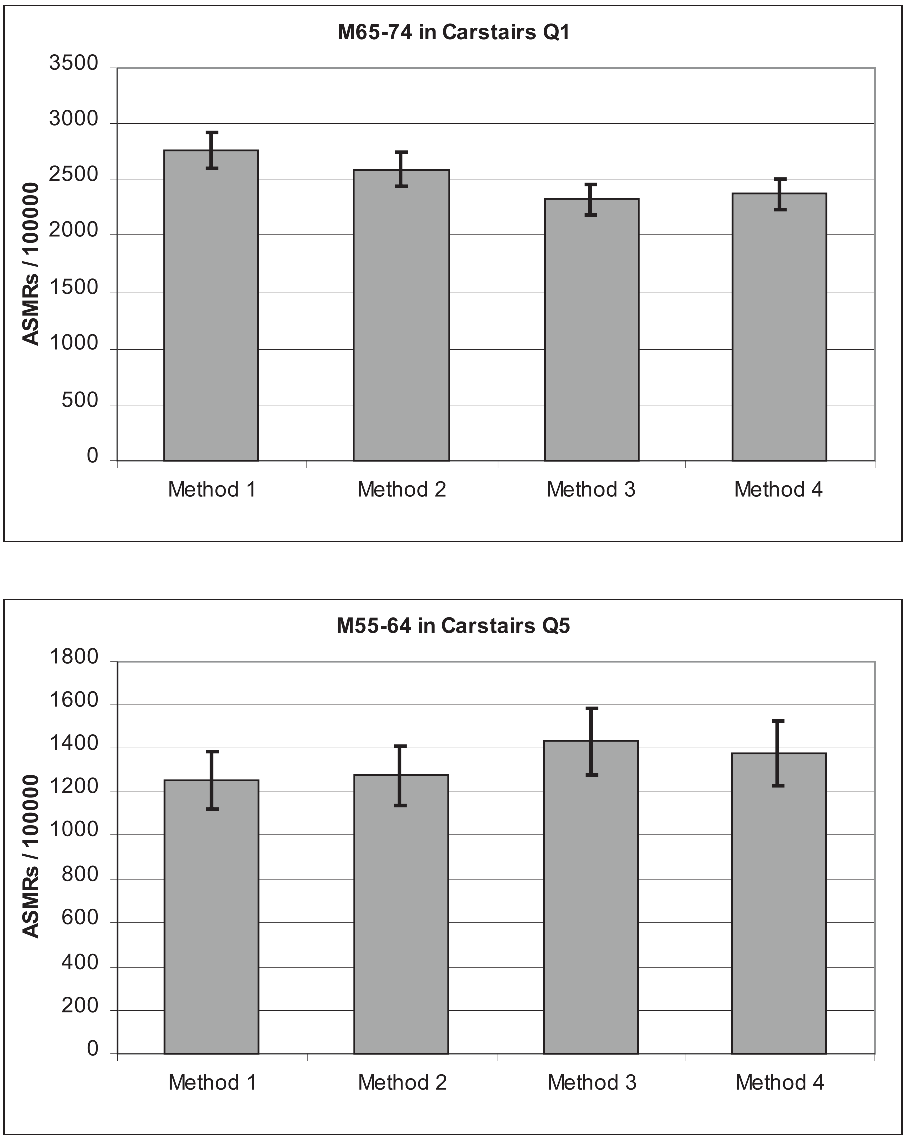

Figure 3 Comparison of ASMRs calculated with age–sex counts estimated using various methods for East of England GOR. Deprivation is measured using the Carstairs Index. Error bars are 95 per cent CIs. Source: computations based on ONS mortality and population data reported in Ref. 6.

dence for our two regions is that inequality increased; a result that supports previous research.[2,3] It is clear, however, that in looking at trends in mortality at small area scale, careful attention must be paid to the estimation of denominator populations.

Finally, does it matter whether the trend is towards increasing or decreasing inequalities in health when the differences in mortality rates over a few years are so slight that the direction of the trends is dependent on the population estimates used? We would argue that the direction of these trends is extremely important as, apart from the unjust deaths they directly represent, the trends are indicative of whether life chances in society are becoming more or less unfairly distributed overall. Inequalities in mortality are the tip of the iceberg of inequalities more generally. Given that, it is imperative that we can determine accurately whether life in Britain is becoming more or less fair.

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Contributors

D.B. and P.N. carried out the analysis. P.R. wrote the first draft, D.D contributed to the literature review and with P.R. redrafted the paper. P.N. assembled the tables, devised the figures and revised the report following referees' comments. P.R. is the guarantor for the study.

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