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Accounting for the dead in the longitudinal analysis of income-related health inequalities

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Abstract

This paper develops an accounting framework to consider the effect of deaths on the longitudinal analysis of income-related health inequalities. Ignoring deaths or using inverse probability weights (IPWs) to re-weight the sample for mortality-related attrition can produce misleading results, since to do so would be to disregard the most extreme of all health outcomes. Incorporating deaths into the longitudinal analysis of income-related health inequalities provides a more complete picture in terms of the evaluation of health changes in respect to socioeconomic status. We illustrate our work by investigating health mobility in Quality Adjusted Life Years (QALYs) as measured by the SF6D from 1999 till 2004 using the British Household Panel Survey (BHPS). We show that for Scottish males explicitly accounting for the dead, rather than using IPWs to account for mortality-related attrition, changes the direction of the relationship between relative health changes and initial income position, while for other population groups it increases the strength of this relationship by up to 14 times. When deaths are explicitly incorporated into the analysis it is found that over this five year period for both Scotland and England & Wales the relative health changes were significantly regressive such that the poor experienced a larger share of the health losses relative to their initial share of health and a large amount of this was related to mortality.

Keywords: mortality, morbidity, income-related health inequality, mobility analysis, longitudinal data, inverse probability weights

JEL classifications: D39, D63, I18

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1. Introduction

A strong cross-sectional relationship between individuals' socioeconomic status and health has been documented in numerous studies (Benzeval and Judge, 2001; Deaton, 2003; Gerdtham and Johannesson, 2004). Significant income-related inequalities in health have persisted, and even increased, in some western countries over the last decade in spite of considerable improvements in average health status (Van Doorslaer and Koolman, 2004; Kunst et al., 2005). Thus, reducing socioeconomic inequalities in health has become a key policy objective for many European governments (Mackenbach and Bakker, 2002). As with any policy objective, it is important to be able to evaluate progress and understand reasons for progress in order to inform future policy (Exworthy, 2006).

Often the longitudinal analysis of income-related health inequalities focuses on how the cross-sectional relationship, between income (or some other socioeconomic status indicator) and the morbidity of those currently alive, evolves over time (Lahelma et al. 2002; Gravelle and Sutton, 2003; Kunst et al., 2005). However, in order to evaluate the performance of policies in reducing income-related health inequalities, a measurement framework is needed which simultaneously examines changes in inequality associated with both morbidity changes and mortality (Khang et al. 2004).

The main measure of income-related health inequality within the health economics literature is the concentration index (Wagstaff and Van Doorslaer, 2000). This captures the extent to which good health in any period is concentrated among the rich compared to the poor and is equal to twice the covariance between health and income rank normalised by average health.

Changes in the concentration index (*CI*) over time have been analysed in the manner of Gravelle and Sutton (2003) using repeated cross-sections, but this does not consider the impact of individuals dying and dropping out of the population between cross-sectional surveys. The changes in cross-sectional income-related health inequality are usually calculated based only on a sample of those in the population at each point in time. Holding all else equal, if those who are poor and sick in the initial period are more likely to die than the rich and sick then this will result in an improvement in the cross-sectional *CI* of those alive in the final period, even though such a transition is likely to be viewed as a policy failure rather than a success.

The longitudinal analyses of the concentration index have also been conducted using both balanced and unbalanced panel data on individuals where the dead are either excluded from the analysis is all periods or included only in periods where they are alive[1]. One recent longitudinal study, Allanson et al. (2010), tracks the performance of individuals over time by decomposing the change in the *CI* into "income-related health mobility", which measures the effect of the relationship between health changes and the initial income rank of the individuals on the change in the *CI*, and "health-related income mobility", which measures the effect of the relationship between income rank changes and the final health of the individuals on the change in the concentration index[2]. While this allows one to follow the performance of individuals over the period it again does not capture the impact of individuals who are alive in the initial period but dead by the final period, as it uses a balanced sample of only those alive in both periods. Taking

[1] In most cases individuals who die during the period are excluded from the sample when a longitudinal perspective is taken as in Wildman (2003), Jones and Lopez-Nicolas (2004) and Allanson et al. (2010). Islam et al. (2010) compare the results from an unbalanced sample with a balanced sample while investigating the extent to which income-related health inequalities change as the population ages.

[2] Note that in Allanson et al. (2010) we also outline an alternative decomposition which measures "income-related health mobility" and "health-related income mobility" from a different perspective.

mortality into account is important for the evaluation of policies which tackle health inequalities since a failure to do so would ignore perhaps the most important of all health outcomes.

One option used to deal with attrition in analysing the dynamics of health is to reweight the sample using inverse probability weights (IPWs) (Jones et al., 2006; Van Kippersluis et al., 2009). This involves placing extra weight on those individuals within the final sample who appear to have the same initial characteristics as those who drop out of the sample. However, in the current context, it seems unreasonable to assume that there are some individuals who stay within the sample (stay alive) who could represent the longitudinal experience of those that die, given that death is the most extreme health outcome possible[3]. In particular, if those that die between the initial and final period were likely to be in general sicker in the initial period, then by construction the sick in the initial period that stay alive between the initial and final period obviously had a better longitudinal experience in terms of their health. Therefore, simply placing more weight on the performance of these individuals would bias the result. In our empirical example we show that for Scottish males explicitly accounting for the dead rather than using IPWs to account for mortality-related attrition changes the direction of the relationship between relative health changes and initial income position, while for other population groups it increases the strength of this relationship by up to 14 times.

This paper aims to provide a unified framework for the longitudinal analysis of changes in income-related health inequality due to both morbidity changes and mortality, based on the assumption that the dead are assigned a health state of zero[4]. First, we provide an overview of the longitudinal methods employed by Allanson et al. (2010). Second, we extend these methods to

[3] Jones et al. (2006) do note that non-response associated with idiosyncratic morbidity shocks are likely problematic and that their Hausman test is unlikely to pick up this type of bias.
[4] We also show in our empirical example that, even when taking a more conservative assumption regarding the weight of mortality, using IPWs can create significant bias.

account for the impacts of mortality on income-related health inequalities. The paper then uses data from the BHPS (British Household Panel Survey) to perform an ex-ante evaluation of the extent to which relative health changes from 1999 to 2004 in England & Wales and in Scotland were progressive in the sense that they have favoured the initially poor. It compares the results when mortality is assumed to be a form of attrition and adjusted for using IPWs to when mortality is explicitly included in the analysis. Finally, the paper compares the performance of England & Wales versus Scotland in tackling income-related health inequalities over this period.

Methods

Review of the longitudinal analysis of income-related health inequalities

We provide an overview of the decomposition methods in Allanson et al. (2010) which are based on a balanced sample where all individuals considered are alive in both periods.

The approach is based on the simple observation that any change in income-related health inequality over time must arise from some combination of changes in health outcomes and income ranks. By decomposing the change in CI between two periods, an index of income-related health mobility is provided that captures the effect on short run income-related health inequality of differences in relative morbidity changes between individuals with different levels of initial income. Thus, the measure addresses the question of whether the pattern of morbidity changes is biased in favour of those with initially high or low incomes, providing a natural counterpart to measures of income-related health inequality that address the issue of whether those with better health tend to be the rich or poor. In addition, a health-related income mobility index that captures the effect of the reshuffling of individuals within the income distribution on cross-sectional income-related health inequalities is obtained.

The change in the short run *CI* between any initial (or start) period *s* and any final period f of only those alive in both periods is decomposed into two parts:

$$
\begin{aligned}
CI_f^A - CI_s^A &= \frac{2}{\bar{h}_f^A}\operatorname{cov}(h_{if}^A, R_{if}^A) - \frac{2}{\bar{h}_s^A}\operatorname{cov}(h_{is}^A, R_{is}^A) \\
&= \left(\frac{2}{\bar{h}_f^A}\operatorname{cov}(h_{if}^A, R_{if}^A) - \frac{2}{\bar{h}_f^A}\operatorname{cov}(h_{if}^A, R_{is}^A)\right) + \left(\frac{2}{\bar{h}_f^A}\operatorname{cov}(h_{if}^A, R_{is}^A) - \frac{2}{\bar{h}_s^A}\operatorname{cov}(h_{is}^A, R_{is}^A)\right) \quad (1) \\
&= \left(CI_f^A - CI_{fs}^A\right) + \left(CI_{fs}^A - CI_s^A\right) \\
&= M^{RA} - M^{HA}
\end{aligned}
$$

where, CI_s^A and CI_f^A are the *CI*'s in periods *s* and *f* respectively of those individuals who are alive (*A*) in the final period, and CI_{fs}^A is the *CI* obtained when health outcomes in the final period are ranked by income in the initial period; $\bar{h}_f^A$ is the average final health of all those alive in the final period; $\bar{h}_s^A$ is the average initial health of all those still alive in the final period; h_{if}^A is the final health of individual *i* who is still alive in the final period; h_{is}^A is the initial health of individual *i* who is still alive in the final period; R_{if}^A is the final income rank of individual *i* in the final income distribution of all those alive in the final period, and R_{is}^A is the initial income rank of individual *i* in the initial income distribution of all those alive in the final period.

In (1), the index $M^{HA} = \left(CI_s^A - CI_{fs}^A\right)$ provides a measure of income-related health mobility, which captures the effect of differences in relative morbidity changes between individuals with different initial levels of income. M^{HA} is positive (negative) if morbidity changes are progressive (regressive) in the sense that the poorest individuals either enjoy a larger (smaller) share of total morbidity gains or suffer a smaller (larger) share of total morbidity losses compared to their initial share of health, and equals zero if relative morbidity changes are independent of income. M^{HA} in turn depends on the level of progressivity and scale of morbidity changes:

$$M^{HA} = \left(CI_s^A - CI_{fs}^A\right) = \left(\frac{2}{\overline{h}^s}\text{cov}(h_{is}^A, R_{is}^A) - \frac{2}{\overline{h}^f}\text{cov}(h_{if}^A, R_{is}^A)\right)$$
$$= \left(\frac{2}{\overline{h}_s}\text{cov}(h_{is}^A, R_{is}^A) - \frac{2}{\overline{\Delta h^A}}\text{cov}(h_{if}^A - h_{is}^A, R_{is}^A)\right)\left(\overline{\Delta h^A}\Big/\overline{h}_f{}^A\right) = \left(CI_s^A - CI_{\Delta s}^A\right)\left(\overline{\Delta h^A}\Big/\overline{h}_f{}^A\right) = P^A q^A \quad (2)$$

where $CI_{\Delta s}^A$ is the concentration coefficient of morbidity changes ranked by initial income[5] and $\overline{\Delta h^A} = \overline{h}_f^A - \overline{h}_s^A$ is the average morbidity change between the two periods of those who are still alive in the final period. Progressivity is captured by the disproportionality index $P^A = \left(CI_s^A - CI_{\Delta s}^A\right)$. Note that if the average morbidity change is negative, then P^A will be negative if health depreciation is progressive such that relative morbidity losses tend to be larger for rich individuals than poor ones. For any given P^A, the gross redistributive effect is proportional to the scale of morbidity changes $q^A = \overline{\Delta h^A}\Big/\overline{h}_f^A$, measured as the ratio of average morbidity changes to average final period health.

However, the income-related health mobility index M^{HA} will not exactly equal the change in income-related health inequality because it does not allow for the effect of changes in the ranking of individuals in the income distribution between the initial and final periods. This effect is captured by the health-related income mobility index $M^{RA} = CI_f^A - CI_{fs}^A$ which measures the extent to which the re-ranking of income is related to the final health status of individuals.

5 Note that $CI_{\Delta s}^A$ will be negative (positive) if individuals with low initial incomes experience a larger (smaller) share of total health gains or losses than those with high incomes, and will equal zero for a universal flat-rate gain or loss.

While this analysis does examine the performance of individuals over time it only involves a balance sample and thus excludes those that die between the two periods. Next we extend this analysis to consider the impact of mortality on the change in the *CI*.

Accounting for the dead

We now decompose the change in the concentration index across two periods where some of the initial population drops out of the population due to mortality. CI_f^A is again the *CI* for all those individuals who are alive in the final period whereas CI_s is the *CI* of all individuals in the initial period, including those who die between the two periods. The change in these cross-sectional concentration indices is decomposed into three main components, (a) the change in health ranked by initial income (which includes the people who die between the two periods who are assumed to have zero final health), (b) the change in income rank by final health (where both the initial and final income ranks only consider those people alive in both periods) and finally (c) the change in the concentration index due to the dead not being included in the final period concentration index.

$$\begin{aligned}
CI_f^A - CI_s &= \frac{2}{\bar{h}_f^A}\operatorname{cov}(h_{if}^A, R_{if}^A) - \frac{2}{\bar{h}_s}\operatorname{cov}(h_{is}, R_{is}) \\
&= \left(\frac{2}{\bar{h}_f^A}\operatorname{cov}(h_{if}^A, R_{if}^A) - \frac{2}{\bar{h}_f}\operatorname{cov}(h_{if}, R_{is})\right) + \left(\frac{2}{\bar{h}_f}\operatorname{cov}(h_{if}, R_{is}) - \frac{2}{\bar{h}_s}\operatorname{cov}(h_{is}, R_{is})\right) \quad (3) \\
&= \left(CI_f^A - CI_{fs}\right) + \left(CI_{fs} - CI_s\right) \\
&= M^R - M^H
\end{aligned}$$

where, $\bar{h}_s$ is the average initial period health of all individuals regardless of whether they are alive or dead in the final period; $\bar{h}_f$ is the average final period health of all individuals including the death who are assumed to have a health of zero; R_{is} is the income rank of an individual in the initial period (regardless of whether they are alive or dead in the final period); h_{is} is the health of an individual in the initial period regardless of whether they are alive or dead in the final period; h_{if} is the health of an individual in the final period where if the individual is dead they are assumed to have zero health; and CI_{fs} is the concentration index of final health ranked by initial income where they dead are also included in the calculation.

The income-related health mobility index $\left(M^{H} = CI_{fs} - CI_{s}\right)$, which captures the effect of differences in relative health changes (including both morbidity changes and mortality) between individuals with different initial levels of income, depends upon both the level of progressivity (P) of the health changes and scale of health changes (q).

$$M^{H} = \left(\frac{2}{\overline{h}_{f}} \operatorname{cov}(h_{if}, R_{is}) - \frac{2}{\overline{h}_{s}} \operatorname{cov}(h_{is}, R_{is}) \right)$$
$$= \left(\frac{2}{\overline{h}_{s}} \operatorname{cov}(h_{is}, R_{is}) - \frac{2}{\overline{\Delta h}} \operatorname{cov}(h_{if} - h_{is}, R_{is}) \right) \left(\overline{\Delta h} \big/ \overline{h}_{f} \right) = \left(CI_{s} - CI_{\Delta s} \right) \left(\overline{\Delta h} \big/ \overline{h}_{f} \right) = Pq \quad (4)$$

Where $\overline{\Delta h}$ is the average change in health (including both morbidity changes and mortality), and $CI_{\Delta s}$ is the concentration index of the change in health (included both morbidity changes and mortality) by initial income rank.

Furthermore M^{H} can be expanded to consider the separate vertical redistributive effects of health changes due to morbidity changes and mortality:

$$
\begin{aligned}
M^{H} &= \left(CI_{s} - CI_{\Delta s}\right)\left(\overline{\Delta h}/\bar{h}_{f}\right) = Pq \\
&= \left(\frac{2}{\bar{h}_{s}}\operatorname{cov}(h_{is}, R_{is}) - \frac{2}{\overline{\Delta h}}\operatorname{cov}(\Delta h_{i}, R_{is})\right)\left(\frac{\overline{\Delta h}}{\bar{h}_{f}}\right) \\
&= \left(\frac{2}{\bar{h}_{s}}\operatorname{cov}(h_{is}, R_{is}) - \frac{2}{\overline{\Delta h}}\operatorname{cov}(\Delta h_{i}^{MB} + \Delta h_{i}^{MT}, R_{is})\right)\left(\frac{\overline{\Delta h^{MB}} + \overline{\Delta h^{MT}}}{\bar{h}_{f}}\right) \\
&= \left(\begin{array}{l} \left(\dfrac{\overline{\Delta h^{MB}} + \overline{\Delta h^{MT}}}{\bar{h}_{f}}\right)\dfrac{2}{\bar{h}_{s}}\operatorname{cov}(h_{is}, R_{is}) \\ -\left(\dfrac{\overline{\Delta h^{MB}}}{\bar{h}_{f}}\dfrac{2}{\overline{\Delta h^{MB}}}\operatorname{cov}(\Delta h_{i}^{MB}, R_{is}) + \dfrac{\overline{\Delta h^{MT}}}{\bar{h}_{f}}\dfrac{2}{\overline{\Delta h^{MT}}}\operatorname{cov}(\Delta h_{i}^{MT}, R_{is})\right) \end{array}\right) \\
&= \left(\frac{2}{\bar{h}_{s}}\operatorname{cov}(h_{is}, R_{is}) - \frac{2}{\overline{\Delta h^{MB}}}\operatorname{cov}(\Delta h_{i}^{MB}, R_{is})\right)\frac{\overline{\Delta h^{MB}}}{\bar{h}_{f}} \\
&\quad + \left(\frac{2}{\bar{h}_{s}}\operatorname{cov}(h_{is}, R_{is}) - \frac{2}{\overline{\Delta h^{MT}}}\operatorname{cov}(\Delta h_{i}^{MT}, R_{is})\right)\frac{\overline{\Delta h^{MT}}}{\bar{h}_{f}} \\
&= \left(CI_{s} - CI_{\Delta s}^{MB}\right)\left(\frac{\overline{\Delta h^{MB}}}{\bar{h}_{f}}\right) + \left(CI_{s} - CI_{\Delta}^{MT}\right)\left(\frac{\overline{\Delta h^{MT}}}{\bar{h}_{f}}\right) = P^{MB}q^{MB} + P^{MT}q^{MT} \\
&= M^{HMB} + M^{HMT}
\end{aligned} \tag{9}
$$

where Δh_i^{MB} and Δh_i^{MT}, are health changes due to morbidity and mortality respectively and are defined so that only one of the measures can be non-zero for any individual. $\overline{\Delta h^{MB}}$ is the average morbidity changes for the whole sample including the dead whom by construction have no change in morbidity. The overall redistributive effect can thus be decomposed into the effects of the two sources of health changes, where each component can in turn be expressed in terms of the progressivity and scale of that type of health change relative to the initial distribution of health and income. By definition, $q = q^{MB} + q^{MT}$ where the scaling factor due to mortality will be higher than might be expected since the averages are determined by net rather than gross health changes. This is because and health changes due to mortality are all negative while health changes due to morbidity are both positive and negative and will therefore to some extent cancel each other out. It is also easily shown that the overall progressivity index is simply the weighted average of the component progressivity indices, $P=\left(q^{MB}/q\right)P^{MB}+\left(q^{MT}/q\right)P^{MT}$, with weights determined by the net health change shares.

We now expand out the income re-ranking term M^R into two key components. Some of the income re-ranking is due to the dead dropping out of the income distribution (and the remaining being re-ranked as a result) while some income re-ranking reflects the shuffling of those alive in both periods.

$$
\begin{aligned}
M^R &= \left(\frac{2}{\overline{h}^{Af}}\operatorname{cov}(h_{if}^A, R_{if}^A) - \frac{2}{\overline{h}^f}\operatorname{cov}(h_{if}, R_{is})\right) \\
&= \left(\frac{2}{N^A\overline{h}^{Af}}\sum_i^A (h_{if}^A)(R_{if}^A - 1/2) - \frac{2}{N\overline{h}^f}\sum_i^N (h_{if})(R_{is} - 1/2)\right)
\end{aligned} \quad (5)
$$

Noting that $N^A\overline{h}^{Af} = N\overline{h}^f$ given that those dead in the final period are given a health status of zero, (5) may be re-written as:

$$
M^R = \left(\frac{2}{N\overline{h}^f}\sum_i^A (h_{if}^A)(R_{if}^A - 1/2) - \frac{2}{N\overline{h}^f}\sum_i^A (h_{if}^A)(R_{is} - 1/2) - \frac{2}{N\overline{h}^f}\sum_{i=A+1}^N (h_{if}^D)(R_{is} - 1/2)\right) (6)
$$

where h_{if}^D is the health state of the dead which by assumption will be equal to zero. Thus, (6) becomes;

$$
M^R = \left(\frac{2}{N\overline{h}_f}\sum_i^A (h_{if}^A)(R_{if}^A - R_{is})\right) (7)
$$

We then use the income rank for the initial period income, excluding those dead in the final period, R_{is}^{A}, to separate the re-ranking effect due to reshuffling of those still alive compared to re-ranking effect due to the dead dropping out of the sample.

$$M^{R} = \frac{2}{N\bar{h}_{f}} \sum_{i}^{A} (h_{if}^{A})(R_{if}^{A} - R_{is}^{A}) + \frac{2}{N\bar{h}_{f}} \sum_{i}^{A} (h_{if}^{A})(R_{is}^{A} - R_{is}) \tag{8}$$
$$M^{R} = M^{RA} + M^{RMT}$$

The term $M^{RA} = \frac{2}{N\bar{h}_{f}} \sum_{i}^{A} (h_{if}^{A})(R_{if}^{A} - R_{is}^{A})$ measures the relationship between the change in income rank based on final health, only considering those living in both periods[6], and the term $M^{RMT} = \frac{2}{N\bar{h}_{f}} \sum_{i}^{A} (h_{if}^{A})(R_{is}^{A} - R_{is})$ reflects the change in the final concentration index due to the dead dropping out of the index.

For M^{RMT} if those who died were, in general, more likely to have been the poor in the initial period then $(R_{is}^{A} - R_{is})$ is likely to be negative as those who are still alive are likely to have been ranked higher in the initial period when those who died were included in the ranking rather than when they were excluded (e.g. if we remove the poorest person from the income distribution then all individual income ranks will fall as they become closer to being the poorest person). And thus, given that h_{if}^{A} must be positive then if those who died were, in general, poor in the initial period this second term will be negative and will make the change in concentration indices between the two periods look smaller and therefore make the final period look more equal than otherwise. If deaths were not related to initial income and were instead random then

[6] Note that as long as the sample weights for each observation are the same when including or excluding the dead then this is the same as the health-related income mobility when the dead are excluded.

$(R_{is}^{A} - R_{is})$ and M^{RMT} would be expected to be zero and there would be no effect on the final concentration index.

Decomposing changes in concentration indices for Scotland and England & Wales

We use both the original decomposition (Allanson et al. 2010) on the balanced sample of those alive in both periods and the new decomposition outlined above accounting for the dead to perform an ex-ante evaluation of the extent to which relative health changes from 1999 to 2004 in England & Wales and Scotland were progressive in the sense that they have favoured the initially poor.

Data

This paper employs data from the annual British Household Panel Survey (BHPS). The BHPS is a multi-purpose survey providing information on, inter alia, health, income and wealth, education, employment, housing, household composition, smoking, leisure activities and individual demographics. The causes of sample attrition between waves, including deaths, are recorded where known.

Health Measure

The health measure used is Quality Adjusted Life Years (QALYs) derived from the SF6D instrument (Brazier et al. 2002), which is available for 1999 and 2004 in the BHPS (i.e. Waves 9 and 14). Those who had been reported deceased by friends, family members or other contacts in or before 2004 were given a QALY weight of zero in 2004.

Income Measure

The income measure (*y*) used is equivalised household income, which takes into account the number of adults and children in the household using the McClements equivalence scale (Taylor, 1995). Because the analysis only involves the relative income rank at each point in time there is no need to convert incomes into real terms.

Sample Attrition

Some individuals in our sample do not report health or income in the final period. A proportion of this is because individuals die between the two periods, but deaths are not the only cause of sample attrition and therefore to provide a more accurate picture of the ex-ante evaluation of changes in income-related health inequalities we also consider other reasons for sample attrition. We report the differences in initial health, income and age for all those that drop out of the sample in the final period against those who remain in the sample.

To control for attrition we use Inverse Probability Weights (IPWs) (as per Jones et al., 2006 and Wooldridge, 2002). This is done by using probit models to derive the probability with regards to the likelihood of non-response in the final period for each individual in the initial sample. These are then used to adjust the initial sample weights such that those who had a higher probability of non-response are given a higher weight, as they are underrepresented in the observed sample. We derive two different alternative final weightings, one which considers death as just another source of sample attrition and one which excludes deaths as a form of sample attrition. Note that while reweighting the sample to take account of death-related attrition is likely to produce misleading results, reweighting the sample for other attrition will have no effect on the progressivity and vertical redistribution indices if those that leave the sample have, on average, the same longitudinal profiles as those individuals with similar initial characteristics who remain in the sample. In all calculations we provide weighted indices where the weights are derived from a combination of the sample weights from the initial period given in the BHPS and adjustment due to sample attrition.

Sample and individual weights

Our initial sample includes all those who answered a full questionnaire in 1999[7] and the BHPS provides cross-sectional weights (W) for all those who have a full questionnaire in this wave. Those who were resident in Northern Ireland are excluded from the sample because they are all from the former European Community Household Panel (ECHP) sub-sample which stopped in 2001 (before our final period).

Some individuals who receive a full interview in 1999 have data missing on the required health variables in 1999[8]. In order to correct the sample weightings for this we use IPWs to re-weight the sample based on initial income (y_0), gender, age in 1999 (Age_0) and whether the individual was resident in Scotland (Sco_0) as opposed to England or Wales in 1999. The dependent variable $Full_i$ is equal to 1 if the individual has data on health in 1999 and ε_{1i} is assumed to be normally distributed.

[7] Only those who answered a full questionnaire have the data available to derive their QALY health status.
[8] 141 individuals did not have health data for 1999.

$$Full_i = 1 \quad if \quad \beta_0 + \beta_1 y_{0i} + \beta_2 Gender_i + \beta_3 Age_{0i} + \beta_4 Sco_{0i} + \varepsilon_{1i} > 0$$
$$= 0 \quad otherwise$$

The predicted probability for each individual $\hat{P}(Full_i)$ is then used to adjust the original cross-sectional weights W_i to derive the new weights (WN_{1i})[9];

$$WN_{1i} = W_i \big/ \hat{P}(Full_i)$$

The cross-sectional sample weights still include the ECHP sample, resident in England, Wales and Scotland, but who are excluded from the BHPS after 2001[10]. Therefore, in order to adjust the cross-sectional weights we use another set of IPWs to re-weight the sample based on initial income (y_0), initial health ($QALY_0$), gender, age in 1999 and whether the individual was resident in Scotland (Sco_0) in 1999, where $NECHP_i$ is equal to 1 if an individual was not in the ECHP sub-sample and ε_{2i} is assumed to be normally distributed.

$$NECHP_i = 1 \quad if \quad \alpha_0 + \alpha_1 y_{0i} + \alpha_2 QALY_{0i} + \alpha_3 Gender_i + \alpha_4 Age_{0i} + \alpha_5 Sco_{0i} + \varepsilon_{2i} > 0$$
$$= 0 \quad otherwise$$

The predicted probability for each individual $\hat{P}(NECHP_i)$ is then used to adjust the current cross-sectional weights WN_{1i} to derive the new weights (WN_{2i});

$$WN_{2i} = WN_{1i} \big/ \hat{P}(NECHP_i)$$

Next we address the problem of non-random attrition between the initial and final periods. We apply the same IPW procedure as above to take into account those individuals who did not receive a full interview or had missing data on health in the final period 2004. Here we

[9] Note that two individuals do not report their age and therefore are given their initial weights with no adjustment.
[10] The ECHP oversampled low-income households therefore it is important to re-weight the remaining observations to account for this.

derive two different possible weightings. First, one weighting assumes reported deaths are not attrition, such that, $NMissf_i$ is equal to 1 if the individual had data on health in the final period or had been recorded as having died before the final period.

$$\begin{aligned} NMissf_i &= 1 \quad if\ \phi_0 + \phi_1 y_{i0} + \phi_2 QALY_{i0} + \phi_3 Gender_i + \phi_4 Age_{i0} + \phi_5 Sco_{0i} + \varepsilon_{3i} > 0 \\ &= 0 \quad otherwise \end{aligned}$$

The random error term ε_{3i} is again assumed to be normally distributed as is the case with probit models. The predicted probability for each individual $\hat{P}(NMissf_i)$ is then used to adjust the current cross-sectional weights (WN_{2i}) to derive the new weights (WN_{3i});

$$WN_{3i} = WN_{2i} \Big/ \hat{P}(NMissf_i)$$

And the second final weighting which assumes death as just another form of attrition, such that, $NMissDf_i$ is equal to 1 if the individual had data on health in the final period but had not been recorded as having died before the final period.

$$\begin{aligned} NMissDf_i &= 1 \quad if\ \gamma_0 + \gamma_1 y_{0i} + \gamma_2 QALY_{0i} + \gamma_3 Gender_i + \gamma_4 Age_{0i} + \gamma_5 Sco_{0i} + \varepsilon_{4i} > 0 \\ &= 0 \quad otherwise \end{aligned}$$

The predicted probability for each individual $\hat{P}(NMissDf_i)$ is then used to adjust the current cross-sectional weights (WN_{2i}) to derive the new weights (WN_{4i});

$$WN_{4i} = WN_{2i} \Big/ \hat{P}(NMissDf_i)$$

Decompositions

We first apply the decomposition of Allanson et al. (2010) to only those alive in both periods (using weights assuming reported death as just some other form of attrition). Secondly we apply the new decomposition described in this paper to the change in the concentration index considering both those alive and dead in the final period (using weights assuming reported death is not a form of attrition). We apply the measures separately for males and females and conduct the analysis separately for those residing in England & Wales or in Scotland in 1999[11].

Robustness

In order to explore the statistical significance of the results and whether or not the results significantly differ across genders and countries we apply a bootstrap sampling procedure 2000 times, where the re-sampling occurs at the clustered level (Primary Sampling Unit)[12]. The mobility calculations are re-estimated for each bootstrapped sample to provide 95% confidence intervals around the concentration and mobility indices and provide significance levels.

To further illustrate that explicitly taking mortality into account rather than just using IPWs is important we also take a very conservative approach and estimate the income-related health mobility assuming death has an implied weight of 0.319 rather than zero, where 0.319 is the lowest QALY weight of anyone still alive in our population in the final period.[13]

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[11] Of those individuals whom resided in England in 1999, 16 (<1%) are found to reside in Scotland in 2004. Of those individuals who resided in Scotland in 1999, 26 (<1%) are found to reside in England or Wales in 2004.
[12] The individual weights are set from the original sample and the bootstrapping procedure does not include re-estimating the individual weights.
[13] We thank Tom Van Ourti for this suggestion.

Results

Deaths and Sample Attrition

Table 1 provides the percentages and reasons for sample attrition for both England & Wales and Scotland, plus some descriptive statistics relating to the initial mean health, mean income and mean age of each group. In England & Wales, 29.5% and 26.4% of the males and females respectively who had answered the full BHPS questionnaire in 1999 failed to answer the full questionnaire in 2004. These included, 4.9% and 4.3% of males and females respectively who were alive in England & Wales in the initial period but who had been reported deceased by the final period. This is compared to Scotland where 35.1% and 35.0% of males and females respectively failed to answer the full questionnaire in 2004 which included 5.0% and 4.9% of males and females respectively whom had died between the two periods. Between 1999 and 2004 health changes related to mortality made up 37% and 29% of all absolute health changes for English & Welsh males and females respectively. While for Scotland, mortality contributed 39% and 34% to all the absolute health changes for males and females respectively[14].

In general, for both countries it can be seen than those who died between the two periods where sicker, poorer and older in 1999 compared to those who survived. Those who did not respond in 2004 due to age, infirmary, disability or because they were institutionalized were also sicker, poorer and older in 1999 but these accounted for only a very small percentage of the total sample (<1%). In general, those in Scotland, compared to those in England & Wales, who answered the full questionnaire in 1999 were less likely to answer the full questionnaire again in 2004 mainly due to a higher refusal rate (including adamant refusals at previous waves), more people not being contactable and more telephone interviews taking place in Scotland for 2004.

[14] For both countries these figures are unweighted statistics.

Table 1: Reasons for sample attrition in 2004 and descriptive statistics for each group in 1999

Males

Interview status in 2004	Number in each Category Scotland		England & Wales		Mean Health 1999 Scotland	England & Wales	Mean Income (£,000s) Scotland	England & Wales	Mean Age 1999 Scotland	England & Wales
Full interview in 2004	892	(64.9%)	3391	(70.5%)	0.83 (0.12)	0.83 (0.11)	23.8 (24.1)	24.5 (20.5)	44.4 (16.6)	44.6 (16.9)
Dead	69	(5.0%)	234	(4.9%)	0.71 (0.17)	0.70 (0.16)	15.3 (11.2)	16.4 (10.6)	71.3 (12.5)	71.0 (14.0)
Full interview in 2004 (no health data)	12	(0.9%)	44	(0.9%)	0.81 (0.12)	0.80 (0.15)	29.9 (25.0)	23.8 (18.6)	47.3 (16.8)	57.5 (19.7)
Proxy interview	10	(0.7%)	62	(1.3%)	0.86 (0.12)	0.79 (0.16)	28.7 (17.7)	22.3 (21.3)	40.8 (11.6)	44.6 (19.3)
Telephone interview	74	(5.4%)	169	(3.5%)	0.81 (0.14)	0.84 (0.11)	23.3 (13.7)	23.8 (14.6)	43.7 (14.4)	42.0 (17.1)
Refusal	86	(6.3%)	212	(4.4%)	0.84 (0.11)	0.82 (0.13)	21.4 (13.7)	20.0 (12.5)	43.0 (18.5)	41.0 (17.9)
Other non-interview	12	(0.9%)	41	(0.9%)	0.87 (0.06)	0.84 (0.13)	24.4 (16.6)	16.7 (9.1)	28.8 (15.2)	31.8 (13.6)
Age, infirmity or disability	5	(0.4%)	13	(0.3%)	0.61 (0.19)	0.67 (0.13)	13.4 (4.1)	13.2 (10.8)	69.4 (9.9)	72.0 (15.4)
Non-contact	100	(7.3%)	242	(5.0%)	0.81 (0.13)	0.82 (0.14)	17.2 (13.3)	21.0 (17.2)	31.1 (13.2)	32.7 (12.9)
Out-of-scope	34	(2.5%)	83	(1.7%)	0.84 (0.10)	0.86 (0.10)	25.3 (16.8)	26.8 (16.2)	35.8 (13.6)	35.9 (13.4)
Institutionalised	3	(0.2%)	10	(0.2%)	0.77 (0.19)	0.67 (0.11)	14.5 (5.0)	14.3 (12.3)	57.0 (31.2)	73.2 (8.2)
Isolated Temporary Sample Member	14	(1.0%)	155	(3.2%)	0.84 (0.11)	0.82 (0.12)	20.4 (19.8)	31.2 (39.4)	34.6 (10.6)	34.3 (13.9)
Adamant Refusal at Previous Wave	61	(4.4%)	145	(3.0%)	0.85 (0.11)	0.81 (0.15)	24.1 (13.4)	23.6 (18.1)	44.7 (16.5)	43.9 (18.6)
Long-term untraced or withdrawn	3	(0.2%)	8	(0.2%)	0.88 (0.08)	0.86 (0.03)	12.9 (9.5)	29.6 (15.7)	42.0 (12.5)	31.6 (7.3)
Total with health data in 1999	1375		4809							

Females

Interview status in 2004	Number in each Category Scotland		England & Wales		Mean Health 1999 Scotland	England & Wales	Mean Income 1999 (£,000s) Scotland	England & Wales	Mean Age 1999 Scotland	England & Wales
Full interview in 2004	1129	(65.0%)	4122	(73.6%)	0.80 (0.13)	0.79 (0.13)	22.1 (23.9)	22.5 (21.8)	44.7 (17.1)	45.5 (17.4)
Dead	85	(4.9%)	241	(4.3%)	0.63 (0.16)	0.64 (0.16)	14.0 (12.0)	14.1 (9.8)	74.2 (12.5)	75.5 (12.4)
Full interview in 2004 (no health data)	6	(0.3%)	49	(0.9%)	0.82 (0.09)	0.75 (0.16)	16.2 (13.0)	22.4 (12.5)	61.3 (13.0)	52.1 (15.9)
Proxy interview	4	(0.2%)	26	(0.5%)	0.77 (0.10)	0.77 (0.16)	24.7 (12.2)	17.6 (11.3)	41.8 (26.0)	53.1 (22.9)
Telephone interview	79	(4.6%)	211	(3.8%)	0.79 (0.13)	0.81 (0.12)	20.0 (12.3)	22.9 (21.5)	46.8 (15.2)	42.5 (15.6)
Refusal	110	(6.3%)	215	(3.8%)	0.80 (0.14)	0.79 (0.14)	19.3 (12.3)	19.2 (13.8)	44.1 (18.8)	42.1 (19.6)
Other non-interview	23	(1.3%)	39	(0.7%)	0.79 (0.13)	0.78 (0.13)	23.9 (21.2)	18.6 (13.1)	38.0 (19.7)	43.1 (20.8)
Age, infirmity or disability	8	(0.5%)	40	(0.7%)	0.67 (0.18)	0.66 (0.16)	8.4 (4.4)	12.1 (8.2)	70.6 (12.0)	79.4 (9.4)
Non-contact	154	(8.9%)	189	(3.4%)	0.79 (0.13)	0.79 (0.14)	17.0 (13.6)	17.8 (15.8)	29.5 (10.8)	33.0 (14.8)
Out-of-scope	39	(2.2%)	107	(1.9%)	0.84 (0.14)	0.83 (0.11)	22.5 (15.4)	26.4 (16.8)	36.4 (15.2)	34.8 (15.0)
Institutionalised	2	(0.1%)	10	(0.2%)	0.75 (0.12)	0.72 (0.13)	9.2 (0.25)	15.0 (13.3)	77.5 (3.5)	73.8 (19.7)
Isolated Temporary Sample Member	10	(0.6%)	159	(2.8%)	0.78 (0.13)	0.79 (0.12)	15.9 (9.3)	24.3 (23.1)	32.5 (12.8)	33.8 (16.2)
Adamant Refusal at Previous Wave	84	(4.8%)	183	(3.3%)	0.82 (0.13)	0.80 (0.14)	24.6 (16.1)	22.8 (21.0)	44.2 (17.4)	48.4 (19.3)
Long-term untraced or withdrawn	3	(0.2%)	7	(0.1%)	0.88 (0.02)	0.73 (0.14)	11.4 (10.9)	24.4 (20.8)	39.3 (8.4)	33 (14.7)
Total with health data in 1999	1736		5598							

Un-weighted statistics. Standard deviations in brackets for health, income and age.

Sample weights and Inverse Probability Weightings (IPWs)

Appendix A, Table A1, provides the results for each of the probit models which are used to derive the inverse probability weightings (IPW) and thus re-weight the sample for; missing initial health data in 1999, ECHP sub-sample exclusion, sample attrition and missing health data in 2004 including death and the sample attrition and missing health data in 2004 excluding death. The results suggest that older individuals who answered the full questionnaire in 1999 were significantly less likely to have their health variable available in 2004 than younger individuals. As expected, those in the ECHP sample were more likely to be poorer, older and sicker in 1999 than those in the regular BHPS sample.

In both cases when deaths were included and excluded as a form of attrition, non-response was significantly related to among other things initial health and income. Those who did not give a full interview (apart from those recorded as dead) or did not answer the necessary health questions in 2004 were significantly more likely to be male, younger, poorer, sicker and from Scotland in 1999 than those who either answered the health questions in the full interview for 2004 or were recorded as dead by 2004.

Differences in results when the dead are explicitly included in the analysis

As can be seen in Table 2, for Scottish males, when the dead are excluded from the analysis, income-related health mobility is positive suggesting that the relative health changes from 1999 to 2004 appear to have been progressive (though not significant at conventional levels) such that those with initially low incomes experienced a larger relative share of the health gains compared to their initial share of health. However, once deaths are explicitly incorporated into the picture the result completely reverses with health changes now shown to have been significantly regressive with the poor experiencing a greater relative share of the health losses compared to their initial share of health. In addition, for Scottish females, English & Welsh males and English & Welsh females, the income-related health mobility goes from having a mild regressive effect when mortality is accounted for by using IPWs to be being 14.5, 6.7 and 4.9 times larger respectively when mortality is explicitly included in the analysis (Tables 2 & 3).

The breakdown in terms of the mortality and morbidity contributions to the income-related health mobility are also provided Tables 2 & 3. In all cases the mortality related income-related health mobility provided the greatest contribution, however, this may not be surprising given that the poor were sicker to start with and those who are sicker tend to have a higher chance of death. For both Scottish and English & Welsh males when deaths were explicitly included in the analysis the progressivity index, or how concentrated the relative health changes are to the poor, increases. However, while the Scottish females' relative health losses are more concentrated in the poor when deaths are taken into account, for English & Welsh females the progressivity index falls when deaths are explicitly taken into account. This does not suggest that mortality was progressive but only that the mortality related health losses were potentially less concentrated among the poor compared to the morbidity losses for this population group.

For all cases, when deaths are excluded from the analysis, the health-related income mobility (M^{RA}) is positive and results in a larger final *CI*. This is because those with lower final health are more likely to have moved down the income rank between 1999 and 2004. When deaths are explicitly included in the analysis the effect of the income re-ranking of individuals alive in both periods (M^{RA}) on the final *CI* is again positive with those alive, but with worse health in 2004, more likely to have moved down the income rank compared to those with better final health. However, this effect on the CI is more than offset by the effect of the income re-ranking as a result of the dead dropping out of the final population (M^{RMT}). This is because the dead, who were mostly the initially poor, drop out of the income distribution resulting in a lower income rank for all those still alive in the final period. This produces a smaller final concentration index compared to if the dead were included in the final concentration index but assumed to have kept their same initial income rank. Thus, while this process may show up as an improvement in cross-sectional CI of the population alive at each point in time it is obviously a consequence of the undesirable situation where the initially poor are more likely to die between the two periods.

From these results we can see that mortality can play a large role in explaining the evolution of income-related health inequalities over time and the small changes in the cross-sectional *CI*'s between 1994 and 2004 which are insignificant[15] for both sexes and regions hide the large changes which have taken place at the individual level.

Sensitivity to mortality assumption

Even when we take a conservative approach and assume mortality to be equal to the lowest health level achieved of those still living in the final period (0.319) instead of zero, we find large differences in income-related health mobility compared with using IPWs methods to account for the mortality-related attrition. Given this conservative assumption, Scottish males' income-related health mobility is still regressive (M^H = -0.00873) compared to being progressive in the IPW case. While for Scottish females, English & Welsh males and English & Welsh females, the income-related health mobility goes from having a mild regressive effect when mortality is

[15] At all conventional levels of significance.

accounted for by using IPWs to be being 7.4, 3.9 and 2.8 times larger respectively when mortality is explicitly included in the analysis and given a QALY weight of 0.319.

Comparing the results across regions

When mortality is explicitly included both males and females in England & Wales appear to have higher levels of income-related health mobility than Scotland, though the differences are not significant at conventional levels. Examining the progressivity index and the scale factors allows us to explore the extent to which this is a result of either larger relative health changes in England & Wales (a scale factor) or health changes which are more concentrated in the poor in England & Wales (progressivity index), compared to Scotland. For males, the relative health losses in England & Wales are significantly larger (10% level) than for Scotland and this appears to drive the higher income-related health mobility. While for females the opposite appears true with a the higher progressivity index in England & Wales, which is significant at the 5% level, driving the differences and suggesting that the relative health losses in English & Welsh females were more concentrated within the poor compared to Scotland.

Table 2. Males: Concentration and Mobility Indices for Scotland and England & Wales

		Scotland (excluding dead)	England & Wales (excluding dead)	Scotland (including the dead)	England & Wales (including the dead)	Difference (Scotland-E&W) (including the dead)
Mean health 1999	$\bar{h}_s^A / \bar{h}_s$	0.816*** (0.806, 0.827)	0.825*** (0.820, 0.829)	0.817*** (0.806, 0.827)	0.822*** (0.817, 0.827)	-0.00559 (-0.0167, 0.00602)
Mean health 2004 (included dead)	$\bar{h}_f$	-	-	0.778*** (0.760, 0.794)	0.767*** (0.758, 0.776)	0.0105 (-0.00832, 0.0291)
Mean health 2004 (excluding the dead)	$\bar{h}_f^A$	0.823*** (0.813, 0.832)	0.815*** (0.810, 0.820)	0.829*** (0.820, 0.837)	0.820*** (0.815, 0.825)	0.00882* (-0.00143, 0.0183)
Mean income 1999	$\bar{y}_s^A / \bar{y}_s$	22.6*** (20.9, 24.6)	23.9*** (23.2, 24.7)	22.6*** (20.8, 24.6)	23.8*** (23.0, 24.6)	-1.19 (-3.05, 0.095)
Mean income 2004 (excluding the dead)	$\bar{y}_f^A$	27.7*** (25.9, 29.5)	28.3*** (27.5, 29.2)	28.3*** (26.5, 30.1)	28.8*** (27.9, 29.7)	-0.513 (-2.44, 1.50)
Concentration index 1999	CI_s^A / CI_s	0.0198*** (0.0122, 0.0281)	0.0153*** (0.0121, 0.0185)	0.0198*** (0.0131, 0.0271)	0.0175*** (0.0144, 0.0205)	0.00237 (-0.00503, 0.01040)
Concentration index 2004	CI_f^A	0.0243*** (0.0169, 0.0316)	0.0232*** (0.0197, 0.0268)	0.0227*** (0.0162, 0.0294)	0.0216*** (0.0183, 0.0250)	0.00011 (-0.00614, 0.00831)
Change in concentration index	ΔCI	0.00445 (-0.00462, 0.0125)	0.00792*** (0.00427, 0.01147)	0.00284 (-0.00500, 0.00099)	0.00413*** (0.00065, 0.00774)	-0.00129 (-0.00991, 0.00680)
Income-related health mobility	M^H	0.00155 (-0.00507, 0.00908)	-0.00382*** (-0.00679, -0.00093)	-0.0179*** (-0.0303, -0.0054)	-0.0257*** (-0.0322,-0.0198)	0.00784 (-0.00633, 0.0226)
Income-related morbidity mobility	M^{HMB}	-	-	0.000833 (-0.00524, 0.00775)	-0.00325*** (-0.00603, -0.000553)	0.00408 (-0.00264, 0.0115)
Income-related mortality mobility	M^{HMT}	-	-	-0.0187*** (-0.0290, -0.00978)	-0.0225*** (-0.0280, -0.0170)	0.00376 (-0.00795, 0.0147)
Progressivity index	P	0.192 (-2.93, 2.73)	0.310*** (0.0775, 0.6832)	0.355*** (0.120, 0.588)	0.357*** (0.282, 0.438)	0.00178 (-0.252, 0.241)
Morbidity progressivity	P^{MB}	-	-	0.173 (-3.659, 5.360)	0.249*** (0.044, 0.560)	-0.0763 (-3.85, 4.94)
Mortality progressivity	P^{MT}	-	-	0.339*** (0.193, 0.484)	0.381*** (0.300, 0.457)	-0.0415 (-0.0390, 0.127)
Scale factor	q	0.00807 (-0.00283, 0.0191)	-0.0123*** (-0.0179, -0.0065)	-0.0504*** (-0.0703, -0.0322)	-0.0721*** (-0.0830, -0.0614)	0.0217* (-0.00046, 0.0430)
Morbidity scale factor	q^{MB}	-	-	0.00458 (-0.00516, 0.0141)	-0.0121*** (-0.0169,-0.0070)	0.0167*** (0.00585, 0.0277)
Mortality scale factor	q^{MT}	-	-	-0.0525*** (-0.0670, -0.0391)	-0.0551*** (-0.0629, -0.0472)	0.00254 (-0.0141, 0.0187)
Health-related income mobility	M^R	0.00600*** (0.0009, 0.0111)	0.00410*** (0.00122, 0.00700)	-0.0151** (-0.0289, -0.0026)	-0.0216*** (-0.0284, -0.0151)	0.00655 (-0.00956, 0.0208)
Due to income re-ranking of those still alive	M^{RA}	-	-	0.00687*** (0.0021, 0.0117)	0.00442*** (0.00164, 0.00725)	0.00244 (-0.00316, 0.00809)
Due to income re-ranking as the dead drop-out	M^{RMT}	-	-	-0.0219*** (-0.0338, -0.0109)	-0.0260*** (-0.0323, -0.0199)	0.00410 (-0.00962, 0.0168)

All are weighted statistics. The "excluding deaths" statistics use the weights which are derived assuming deaths are just another form of attrition and this therefore explains the differences in mean health 2004. The parentheses are the lower and upper 95% bootstrapped percentiles from 2000 replications. ***,**, * represent significance at the 1, 5 and 10% levels respectively. Income is measured in thousands of pounds. Note that when the deaths are excluded and just treated as attrition this places greater weight on those remaining individuals who were poor and sick in 1999 compared to when deaths are explicitly included in the analysis.

Table 3. Females: Concentration and Mobility Indices for Scotland and England & Wales

		Scotland (excluding dead)	England & Wales (excluding dead)	Scotland (including the dead)	England & Wales (including the dead)	Difference (Scotland-E&W) (including the dead)
Mean health 1999	$\bar{h}_s^A / \bar{h}_s$	0.786***	0.782***	0.785***	0.780***	0.00440
		(0.775, 0.797)	(0.777, 0.787)	(0.774, 0.796)	(0.775, 0.785)	(-0.00754, 0.0160)
Mean health 2004 (included dead)	$\bar{h}_f$	-	-	0.730***	0.732***	-0.00231
				(0.712, 0.747)	(0.723, 0.742)	(-0.0229, 0.0162)
Mean health 2004 (excluding the dead)	$\bar{h}_f^A$	0.774***	0.774***	0.782***	0.780***	0.00183
		(0.762, 0.786)	(0.768, 0.780)	(0.771, 0.793)	(0.775, 0.786	(-0.0103, 0.0141)
Mean income 1999	$\bar{y}_s^A / \bar{y}_s$	21.3***	22.1***	21.3***	22.0***	-0.685
		(19.9, 23.0)	(21.3, 22.9)	(19.9, 23.0)	(21.2, 22.8)	(-2.29, 1.11)
Mean income 2004 (excluding the dead)	$\bar{y}_f^A$	24.8***	26.0***	25.3***	26.5***	-1.23
		(23.6, 26.1)	(25.2, 26.9)	(24.1, 26.7)	(25.7, 27.4)	(-2.75, 0.376)
Concentration index 1999	CI_s^A / CI_s	0.0181***	0.0185***	0.0186***	0.0205***	-0.00186
		(0.0112, 0.0258)	(0.0146, 0.0222)	(0.0118, 0.0258)	(0.0168, 0.0240)	(-0.00944, 0.00590)
Concentration index 2004	CI_f^A	0.0240***	0.0276***	0.0220***	0.0261***	-0.00417
		(0.0173, 0.0309)	(0.0238, 0.0313)	(0.0157, 0.0286)	(0.0225, 0.0296)	(-0.0113, 0.00324)
Change in concentration index	ΔCI	0.00594	0.00908***	0.00336	0.00568***	-0.0230
		(-0.00192, 0.01333)	(0.00508, 0.0130)	(-0.00477,0.0108)	(0.00157, 0.00976)	(-0.0111, 0.00621)
Income-related health mobility	M^H	-0.00137	-0.00558***	-0.0198***	-0.0273***	0.00756
		(-0.00824, 0.00615)	(-0.00873, -0.00226)	(-0.0286, -0.0109)	(-0.0340, -0.0212)	(-0.00317, 0.0186)
Income-related morbidity mobility	M^{HMB}	-	-	-0.000444	-0.00534***	0.00490
				(-0.00681, 0.00664)	(-0.00835, -0.00217)	(-0.00231, 0.0128)
Income-related mortality mobility	M^{HMT}	-	-	-0.0193***	-0.0220***	0.00266
				(-0.0286, -0.0109)	(-0.0281, -0.0166)	(-0.00842, 0.0122)
Progressivity index	P	0.0897	0.547***	0.265***	0.419***	-0.154**
		(-0.5629, 0.7763)	(0.228, 1.379)	(0.150, 0.391)	(0.335, 0.510)	(-0.303, -0.00586)
Morbidity progressivity	P^{MB}	-	-	0.0271	0.480***	-0.453*
				(-0.514, 0.571)	(0.200, 1.094)	(-1.284, 0.159)
Mortality progressivity	P^{MT}	-	-	0.332***	0.406***	-0.0745
				(0.218, 0.443)	(0.331, 0.482)	(-0.203, 0.0524)
Scale factor	q	-0.0153***	-0.0102***	-0.0747***	-0.0653***	-0.00940
		(-0.0270, -0.0037)	(-0.0162, -0.00408)	(-0.0952, -0.0564)	(-0.0774, -0.0537)	(-0.0327, 0.0112)
Morbidity scale factor	q^{MB}	-	-	-0.0152***	-0.0104***	-0.00480
				(-0.0252, -0.00506)	(-0.0159, -0.00487)	(-0.0162, 0.0068)
Mortality scale factor	q^{MT}	-	-	-0.0543***	-0.0508***	-0.00341
				(-0.0701, -0.0402)	(-0.0593, -0.0429)	(-0.0208, 0.0130)
Health-related income mobility	M^R	0.00457	0.00351**	-0.0164***	-0.0217***	0.00525
		(-0.00237, 0.0124)	(0.000530, 0.00638)	(-0.0273, -0.00455)	(-0.0288, -0.00152)	(-0.00773, 0.0185)
Due to income re-ranking of those still alive	M^{RA}	-	-	0.00516	0.00395***	0.00122
				(-0.00153, 0.0127)	(0.00107, 0.00674)	(-0.00576, 0.00945)
Due to income re-ranking as the dead drop-out	M^{RMT}	-	-	-0.0216**	-0.0256***	0.00403
				(-0.0316, -0.0121)	(-0.0323, -0.0197)	(-0.00788, 0.0151)

All are weighted statistics. The "excluding dead" statistics use the weights which are derived assuming deaths are just another form of attrition and this therefore explains the differences in mean health 2004. The parentheses are the lower and upper 95% bootstrapped percentiles from 2000 replications. ***,**, * represent significance at the 1, 5 and 10% levels respectively. Income is measured in thousands of pounds. Note that when the deaths are excluded and just treated as attrition this places greater weight on those remaining individuals who were poor and sick in 1999 compared to when deaths are explicitly included in the analysis.

Conclusion

This paper extends Allanson et al. (2010) by outlining a decomposition method in order to account for the dead in the longitudinal analysis of income-related health inequalities. Excluding deaths from the analysis gives a misleading picture of the performance in tackling income-related health inequalities. A consistent or even increasing cross-sectional concentration index of the population over time may not be a bad outcome if it is a result of efforts to keep the poor and sick alive for longer than previously may have been the case. Reweighting the sample to account for mortality-related attrition does not solve the problem as those who die between the initial and final periods experience the most extreme health changes possible and are thus significantly different from those with the similar initial characteristics but who stay alive.

The decomposition method outlined in the current paper provides a comprehensive picture of the extent to which both morbidity changes and mortality are related to socioeconomic status and how these impact on income-related health inequalities. Using the performance of England & Wales and Scotland from 1999 till 2004 as an example it was found that explicitly accounting for the dead is important and can lead to very different results in evaluating the performance of reducing income-related health inequalities compared to when the sample is just reweighted assuming mortality is another form of attrition.

Including deaths in the analysis suggests that the relative health changes for both regions and genders between 1999 and 2004 were significantly regressive such that the initially poor experienced a greater share of the health losses compared to their initial share of health. This was mostly because those who were initially poor were more likely to die between the two periods than the rich. However, as the dead drop out of the population this also contributes to a lower final cross-sectional CI of those still alive in 2004 for both sexes and regions compared to if those you had died between 1999 and 2004 were included but given their initial income rank. If the health system in these regions had done a better job at keeping the poor and sick alive then this may have made the 2004 CI appear worse but would have show up in our new decompositions by making the income-related health mobility less regressive.

When deaths are explicitly taken into account all of the mobility indices are significant for each region and gender, however, there are very few significant differences between regions for the indices. The significant differences that are found suggest that the relative health losses for females in England & Wales are more concentrated in the poor compared to relative health losses for females in Scotland. While for males there is some evidence to suggest that those in England & Wales experienced a larger total relative health loss over the period than the Scottish but there was very little difference in terms of how concentrated these relative health losses were among the poor.

In the current analysis we have only considered the longitudinal analysis of those that reported their health and income levels in the initial period and thus have excluded from the analysis some individuals who were alive in the initial period but for some reason may not have completed the questionnaire[16]. As such the analysis may not be representative of the population as a whole if those that did not answer the initial questionnaire experienced different longitudinal outcomes than those that did. Also we have assumed that those who did not provide responses to the questionnaire in the final period due to reasons other than mortality experienced similar patterns in terms of health and income rank changes between the initial and final periods as their counterparts with similar characteristics in the initial period and thus controlled non-mortality

[16] The longitudinal data will also not capture the experience of those who enter the population during the period such as recent migrants or those born during the period.

related attrition by using IPWs. Further research is needed to explore the extent to which this fails to hold and how it impacts on the results.

It is also worth noting that the comparison of mobility indices across countries and genders may be misleading to the extent that socioeconomic health differentials may be expected to have changed over the period simply due to both the ageing of the sample (Kiula and Mieszkowski, 2007) and changes in other determinants of health which may differ across populations groups. Further research is needed to develop standardized measures of mobility to account for those factors which are outside the control of policy.

While the decomposition in this paper has been applied to the change in the *CI* over time and therefore measures changes in relative income-related health inequalities it can also be easily extended to consider the decomposition of changes in absolute and other income-related health inequality measurement tools (see Allanson, 2010).

References

Allanson, P., 2010. Longitudinal analysis of income-related health inequality: welfare foundations and alternative measures. *Dundee Discussion Papers in Economics*, No 240.

Allanson, P. Gerdtham, U-G. and Petrie, D., 2010. Longitudinal analysis of income-related health inequality. *Journal of Health Economics*, 29, 78-86.

Brazier, J., Roberts, J. and Deverill, M., 2002. The estimation of a preference-based measure of health from the SF-36. *Journal of Health Economics*, 21, 271–292.

Deaton, A.S., 2003. Health, inequality, and economic development. *Journal of Economic Literature* XLI, 113–158.

Exworthy, M., Bindman, A., Davies, H. and Washington, Eugene, 2006. Evidence into Policy and Practice? Measuring the Progress of U.S. and U.K. Policies to Tackle Disparities and Inequalities in U.S.and U.K. Health and Health Care. *Milbank Quarterly*, 84, 76-109.

Gravelle, H. and Sutton, M., 2003. Income related inequalities in self assessed health in Britain: 1979–1995. *Journal of Epidemiology and Community Health*, 57, 125-129.

Islam, M., Gerdtham, U., Clarke, P. and Burstrom, K., 2010. Does income-related health inequality change as the population ages? Evidence from Swedish panel data. *Health Economics*, 19(3), 334-349.

Jones, A. M. and Lopez Nicolas, A., 2004. Measurement and explanation of socioeconomic inequality in health with longitudinal data. *Health Economics*, 13, 1015-1030.

Jones, A.M., Koolman, X., Rice, N., 2006. Health-related non-response in the British Household Panel Survey and European Community Household Panel: using inverse-probability-weighted estimators in non-linear models. *Journal of the Royal Statistical Society*, Series A 169, 543–569.

Khang Y-H, Lynch JW, Yun S, Lee S-I. 2004. Trends in socioeconomic health inequalities in Korea: use of mortality and morbidity measures, *Journal of Epidemiology Community Health*, 58, 308-314.

Kiula, O. and Mieszkowski, P., 2007. The effects of income, education and age on health. *Health Economics*, 16, 781–798.

Kunst A, Bos V, Lahelma E, et al., 2005. Trends in socio-economic inequalities in self-assessed health in 10 European countries. *International Journal of Epidemiology* , 34, 295–305.

Lahelma E, Kivela K, Roos E, et al. 2002. Analysing changes of health inequalities in the Nordic welfare states. *Social Science & Medicine*, 55, 609–625.

Mackenbach, J.P., Bakker, M., 2002. Reducing Inequalities in Health: A European Perspective. London: Routledge.

Strategic review of health inequalities in England post 2010, 2010. Fair Society, Healthy Lives (The Marmot Review). University College London: London.

Taylor, M.F., 1995. In: British Household Panel Survey User Manual Volume A: Introduction, Technical Report and Appendices, University of Essex, Colchester.

Van Kippersluis, H., Van Ourti, T., O'Donnell, O., and Van Doorslaer, E., 2009, Health and income across the life cycle and generations in Europe. *Journal of Health Economics*, 28, 818-830.

Wagstaff, A. and Van Doorslaer, E., 2000. Equity in health care finance and delivery. In: Culyer, A.J., Newhouse, J.P. (Eds.) Handbook of Health Economics. Amsterdam, Netherlands: North Holland, p. 1803–1862.

Wildman J. 2003. Income related inequalities in mental health in Great Britain: analyzing the causes of health inequality over time. *Journal of Health Economics*, 22, 295–312.

Wooldridge, J.M., 2002. Econometric Analysis of Cross-section and Panel Data. MIT Press, Cambridge, MA.

Appendix A

Table A1: Probit models for adjusting sample weights

Dependent variable	Health Data available for 1999	Not in ECHP sample	Health data available for 2004 and not reported dead	Health data available for 2004 or reported dead
Explanatory variable	Coefficient (std error)	Coefficient (std error)	Coefficient (std error)	Coefficient (std error)
Constant	2.89*** (0.112)	0.656*** (0.105)	-0.127 (0.0858)	0.00882 (0.0903)
Age	-0.0113*** (0.00168)	-0.0021** (0.00081)	0.000809 (0.000641)	0.0116*** (0.000687)
Income(1999)	0.000240 (0.00154)	0.0106*** (0.00098)	0.00369*** (0.000663)	0.00231*** (0.000645)
Health(1999)	-	0.649*** (0.110)	0.811*** (0.0901)	0.232** (0.0953)
Male	0.0351 (0.0645)	-0.0011 (0.0300)	-0.106*** (0.0232)	-0.0624*** (0.0240)
Scotland	0.0705 (0.0806)	0.361*** (0.040)	-0.207*** (0.0267)	-0.201*** (0.0275)
Sample size	14986	14845	13516	13516
Pseudo R^2	0.0313	0.0343	0.0123	0.0244

All dependent variables are equal to 1 when they are still included in the sample and 0 when they are to be excluded from the sample due to missing data. ***, **, * refers to significance at the 1, 5 and 10% level respectively. Age refers to the age in 1999. Scotland refers to the fact that the individual was recorded as resident in Scotland in 1999 and are given a value of zero if they are resident in England or Wales in 1999. Note that two individuals do not report their age and therefore these are individual were given their original cross-sectional weights with no adjustments. Income is equivalised annual income measured in thousands of pounds.