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Paul Allanson &
Dennis Petrie

Department of
Economic Studies,
University of
Dundee,
Dundee.
DD1 4HN

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On decomposing the causes of changes in income-related health inequality with longitudinal data

Paul Allanson[#] and Dennis Petrie
Economics Studies, University of Dundee, Scotland

Abstract

Regression-based decomposition procedures are used to both standardise the concentration index and to determine the contribution of inequalities in the individual health determinants to the overall value of the index. The main contribution of this paper is to develop analogous procedures to decompose the *income-related health mobility* and *health-related income mobility* indices first proposed in Allanson, Gerdtham and Petrie (2010) and subsequently extended in Petrie, Allanson and Gerdtham (2010) to account for deaths. The application of the procedures is illustrated by an empirical study that uses British Household Panel Survey (BHPS) data to analyse the performance of Scotland in tackling income-related health inequalities relative to England & Wales over the five year period 1999 to 2004.

[#] Corresponding author. Economics Studies, University of Dundee, Perth Road, Dundee DD1 4HN. Tel: +44 01382 384377 Fax: +44 01382 384691. Email: p.f.allanson@dundee.ac.uk. The authors acknowledge the financial support of the CSO under small grant CZG/2/451 “Development of tools to measure and explain changes in health inequalities in Scotland and benchmark performance”. All opinions expressed in this paper are solely the responsibility of the authors.

1. Introduction

The main measure of income-related health inequality within the health economics literature is the concentration index (Wagstaff and Van Doorslaer, 2000). However the value of this simple bivariate measure is determined not only by the relationship between health and income but also by other factors, such as age, gender and lifestyle choices, to the extent that they affect health and are correlated with income. This has led to the development of various standardisation techniques to control for the effects of factors deemed unamenable to policy intervention, such as age and gender, thereby yielding the (augmented) partial concentration index as a measure of 'policy relevant' income-related health inequality (see Gravelle, 2003). More generally, regression-based decomposition techniques have been employed to identify the contribution of inequalities in the individual determinants of health to the overall health concentration index at a specific point in time and thereby provide a basis for the analysis of changes in income-related health inequalities using repeated cross-sections (see, e.g., Wagstaff et al., 2003; and Gravelle and Sutton, 2003).

However, there are important aspects of income-related health inequality changes that are not revealed by examining changes in cross-sectional data over time. In particular, it is not possible to either measure or model the dynamic links between health, income, demographic factors and lifestyle choices in the absence of longitudinal or follow-up data on individuals. Thus, it is not possible to determine to what extent a fall in income-related health inequality over time might be due to a relative improvement in the health of those who were initially poor as opposed to an improvement of their position in the income distribution, where the former might be due to healthcare interventions targeted at the poor and the latter to broader changes in welfare/income provision and economic conditions.

In a recent paper, Allanson, Gerdtham and Petrie (2010; hereafter AGP) consider the characterisation and measurement of income-related health inequality using longitudinal data to track the experience of individuals. Their approach is based on the observation that any change in income-related health inequality over time must arise from some combination of changes in health outcomes (i.e. "health mobility") and changes in individuals' positions in the income distribution (i.e. "income mobility"). Accordingly, they decompose the change in the concentration index between two periods to provide an index of *income-related health mobility* (IRHM) that captures the effect on cross-sectional income-related health inequality of differences in relative health changes between individuals with different levels of initial income and an index of *health-related income mobility* (HRIM) that captures the effect of the reshuffling of individuals within the income distribution on cross-sectional socioeconomic inequalities in health. The main aim of this paper is to develop standardisation and decomposition techniques for these mobility indices akin to those available in the literature for the concentration index. In particular, we seek to develop a partial IRHM index that can provide a benchmark measure of income-related mobility in health after removing the effects of standardising variables which affect health changes, are correlated with initial income, but not amenable to policy. More generally, we seek to establish decomposition techniques that can account for the determinants of mobility and thereby provide appropriate indicators to monitor and evaluate performance in tackling health inequalities.

The paper is structured as follows. The following section develops standardisation and decomposition techniques for the AGP mobility indices akin to those available in the literature for the concentration index. Section 3 proposes analogous procedures for the extended mobility indices introduced in Petrie, Allanson and Gerdtham (2010; hereafter PAG), which takes account of mortality as well as morbidity, so as to enable the construction of partial mobility indices for the entire population in some initial period and not just for that part of the population who survive to the final period. These various measures are employed in Section 4 to investigate the dynamics of income and health using a measure of Quality Adjusted Life Years (QALYs) derived from the SF6D instrument (Brazier et al. 2002). Specifically, we use data from waves 9 and 14 of the British Household Panel Survey (BHPS) to analyse the performance of Scotland in tackling income-related health inequalities relative to England & Wales over the five year period 1999 to 2004. The final section summarises the contribution of the paper.

2. Decomposing health mobility indices: methods

Regression-based decomposition procedures have been used to both standardise the concentration index and to determine the contribution of inequalities in the individual determinants of health to the overall value of the concentration index. We briefly review these procedures before proposing analogous procedures for the AGP mobility indices.

Decomposition of the concentration index

Let CI_{tt} be the concentration index that is obtained when health outcomes in period t are ranked by income in period t. This health concentration index may be written as:

$$CI_{tt} = \frac{2}{\bar{h}_t}\mathrm{Cov}\left(h_t, R_t\right) \tag{1}$$

where h_t is a measure of health, $\bar{h}_t$ is the average health of the population; and $R_t \equiv F(y_t)$ denotes relative income rank, which is determined by the cumulative distribution function $F(\cdot)$ of income y_t. Following Gravelle (2003), suppose that the population health function is given by a linear regression model linking health to income and a set of *K* other health determinants x_{kt} (*k*=1, .. *K*):

$$h_t = \beta_0 + \beta_y y_t + \sum_{k=1}^{K} \beta_k x_{kt} + \varepsilon_t \tag{2}$$

where β_0, β_y, and the β_k's are coefficients, and ε_t is an error term. Accordingly, the concentration index may be re-written as:

$$\begin{aligned} CI_{tt} &= \beta_y \frac{2}{\bar{h}_t} \mathrm{Cov}(y_t, R_t) + \sum_{k=1}^{K} \beta_k \frac{2}{\bar{h}_t} \mathrm{Cov}(x_{kt}, R_t) + \frac{2}{\bar{h}_t} \mathrm{Cov}(\varepsilon_t, R_t) \\ &= \frac{\beta_y \bar{y}_t}{\bar{h}_t} G_{tt} + \sum_{k=1}^{K} \frac{\beta_k \bar{x}_{kt}}{\bar{h}_t} CI_{tt}^k = \eta_t^y G_{tt} + \sum_{k=1}^{K} \eta_t^k CI_{tt}^k \end{aligned} \tag{3}$$

where $\bar{y}_t$ and G_{tt} are the mean and concentration index (i.e. Gini coefficient) of income; $\bar{x}_t^k$ and CI_{tt}^k are the mean and concentration index of health determinant *k*; η_t^y and η_t^k are the elasticities of health with respect to income and health determinant *k* respectively, evaluated at population means; and $\mathrm{Cov}(\varepsilon_t, R_t) = 0$ by assumption. Thus the concentration index can be expressed as a weighted sum of the concentration indices of all the health determinants (including income), with the weight on each index equal to the share of health attributable to that determinant where this is given by the elasticity of health with respect to that determinant evaluated at the means. This decomposition is feasible so long as the health function is linear in the parameters with an additive error term.

Gravelle (2003) proceeds to consider standardisation procedures on the assumption that the set of *K* other health determinants can be partitioned between a sub-set of policy relevant variables and a sub-set of policy irrelevant or standardising variables, where the choice of partition will depend upon the policy context. Assuming that the first *J* ($0 \le J \le K$) other health determinants are deemed policy relevant then the effect of the remaining *K–J* standardising variables on the overall concentration index can be removed by deduction of the terms involving them from (3) to yield the augmented partial concentration index:

$$I_{tt}^{AUG} = \eta_t^y G_{tt} + \sum_{k=1}^{J \le K} \eta_t^k CI_{tt}^k \tag{4}$$

which reduces to the partial concentration index $I_{tt} = \eta_t^y G_{tt}$ if there are no policy relevant variables other than income. A consistent estimator of I_{tt}^{AUG} in (4) is provided by the following three step procedure (Gravelle, 2003). First, estimate the health function (2) to obtain:

$$h_t = \hat{\beta}_0 + \hat{\beta}_y y_t + \sum_{k=1}^{J} \hat{\beta}_k x_{kt} + \sum_{k=J+1}^{K} \hat{\beta}_k x_{kt} + \hat{\varepsilon}_t \; ; \tag{5}$$

where $\hat{\beta}_0$, $\hat{\beta}_y$, and the $\hat{\beta}_k$'s are estimates of β_0, β_y, and the β_k's from a regression of health on income and the other health determinants using cross-sectional data on a sample of individuals drawn from the population, and $\hat{\varepsilon}_t$ is the regression residual. Second, use (5) to generate predictions of 'directly standardised' health h^{AUG} in period *t*:

$$h_t^{AUG} = \hat{\beta}_0 + \hat{\beta}_y y_t + \sum_{k=1}^{J} \hat{\beta}_k x_{kt} + \sum_{k=J+1}^{K} \hat{\beta}_k \bar{x}_{kt} + \hat{\varepsilon}_t \; ; \tag{6}$$

where the values of the standardising variables are held fixed at their sample means, $\bar{x}_{kt}$ (*k=J+1, ...K*), across all individuals. Finally, the estimate of I_{tt}^{AUG} is obtained as the concentration index of h_t^{AUG}. Gravelle (2003) further notes that this method readily generalises to the case where the standardising variables are not additively separable from the other variables in the health function, though the value of the augmented partial concentration index will depend on the fixed values of the standardising variables in this case.

Gravelle (2003) observes that plotting the two components of I_{tt} in {elasticity, Gini} space can yield useful diagrams for showing the time path of inequality (see, for example, Figure 1 in Gravelle and Sutton, 2003). More formally, one can follow Wagstaff et al. (2003) and consider an Oaxaca-type decomposition of the change in the concentration index. Thus the change in the concentration index between an initial period *s* and a final period *f* may be written as either:

$$\Delta CI = \hat{\eta}_f^y \Delta G + G_{ss} \Delta \hat{\eta}^y + \sum_{k=1}^{K} \left(\hat{\eta}_f^k \Delta CI^k + CI_{ss}^k \Delta \hat{\eta}^k \right) + \left(\frac{GC_{ff}^{\hat{\varepsilon}}}{\bar{h}_f} - \frac{GC_{ss}^{\hat{\varepsilon}}}{\bar{h}_s} \right) \tag{7a}$$

or:

$$\Delta CI = \hat{\eta}_s^y \Delta G + G_{ff} \Delta \hat{\eta}^y + \sum_{k=1}^{K} \left(\hat{\eta}_s^k \Delta CI^k + CI_{ff}^k \Delta \hat{\eta}^k \right) + \left(\frac{GC_{ff}^{\hat{\varepsilon}}}{\bar{h}_f} - \frac{GC_{ss}^{\hat{\varepsilon}}}{\bar{h}_s} \right) \tag{7b}$$

where, for period t ($t=s,f$), $\hat{\eta}_t^y$ and $\hat{\eta}_t^k$ denote estimates of the health elasticities based on (5); $\Delta Z = Z_{ff} - Z_{ss}$ for $Z = CI, G, CI^k$, and $\Delta\hat{\theta} = \hat{\theta}_f - \hat{\theta}_s$ for $\hat{\theta}_t = \hat{\eta}_t^y, \hat{\eta}_t^k$; and $GC_{tt}^{\hat{\varepsilon}}$ is the generalised concentration index of the regression residual. This decomposition serves to show how far changes in inequality in health are attributable to changes in inequalities in the determinants of health rather than to changes in the other influences on health inequality. To further disentangle changes going on within the health elasticities, Wagstaff et al. (2003) also consider the total differential of the concentration index:

$$\begin{aligned} \mathrm{d}\,CI_{tt} &= \frac{\mathrm{d}\,CI_{tt}}{\mathrm{d}\,\hat{\beta}_0}\mathrm{d}\,\hat{\beta}_0 + \left(\frac{\mathrm{d}\,CI_{tt}}{\mathrm{d}\,\overline{y}_t}\mathrm{d}\,\overline{y}_t + \frac{\mathrm{d}\,CI_{tt}}{\mathrm{d}\,G_{tt}}\mathrm{d}\,G_{tt} + \frac{\mathrm{d}\,CI_{tt}}{\mathrm{d}\,\hat{\beta}_y}\mathrm{d}\,\hat{\beta}_y \right) \\ &\quad + \sum_{k=1}^{K} \left(\frac{\mathrm{d}\,CI_{tt}}{\mathrm{d}\,\overline{x}_{kt}}\mathrm{d}\,\overline{x}_{kt} + \frac{\mathrm{d}\,CI_{tt}}{\mathrm{d}\,CI_{tt}^k}\mathrm{d}\,CI_{tt}^k + \frac{\mathrm{d}\,CI_{tt}}{\mathrm{d}\,\hat{\beta}_k}\mathrm{d}\,\hat{\beta}_k \right) + \frac{\mathrm{d}\,CI_{tt}}{\mathrm{d}\left\{GC_{tt}^{\hat{\varepsilon}}/\overline{h}_t\right\}}\mathrm{d}\left\{GC_{tt}^{\hat{\varepsilon}}/\overline{h}_t\right\}. \end{aligned} \tag{8}$$

Using obvious notation, this yields the discrete approximation:

$$\begin{aligned} \Delta CI &\approx \frac{CI_{ss}}{\overline{h}_s}\Delta\hat{\beta}_0 + \left(\frac{\hat{\beta}_{ys}\left(G_{ss} - CI_{ss}\right)}{\overline{h}_s}\Delta\overline{y} + \frac{\hat{\beta}_{ys}\overline{y}_s}{\overline{h}_s}\Delta G + \frac{\overline{y}_s\left(G_{ss} - CI_{ss}\right)}{\overline{h}_s}\Delta\hat{\beta}_y \right) \\ &\quad + \sum_{k=1}^{K} \left(\frac{\hat{\beta}_{ks}\left(CI_{ss}^k - CI_{ss}\right)}{\overline{h}_s}\Delta\overline{x}_k + \frac{\hat{\beta}_{ks}\overline{x}_{ks}}{\overline{h}_s}\Delta CI + \frac{\overline{x}_{ks}\left(CI_{ss}^k - CI_{ss}\right)}{\overline{h}_s}\Delta\hat{\beta}_k \right) + \left(\frac{GC_{ff}^{\hat{\varepsilon}}}{\overline{h}_f} - \frac{GC_{ss}^{\hat{\varepsilon}}}{\overline{h}_s} \right) \end{aligned} \tag{9}$$

which will be accurate for small changes in the concentration index and reveals the contributions both of changes in the means and inequalities of the determinants of health, and of changes in the effects of the determinants on health.

Nevertheless, it remains the case that any change analysis based on cross-sectional data can capture the effects of movements of, but not within, the joint distribution of health and income due to individual dynamics in income and health. For example, the change in the cross-sectional concentration index will not pick up any effect from the joint permutation of health and income ranks between individuals. By extension, longitudinal data are required to distinguish between income-related health inequalities arising from chronic or persistent social disadvantage as opposed to those due to transitory episodes of poverty and sickness, where the former would call for policies to tackle the structural problems that trap some individuals in deprivation and ill-health while the latter might demand measures such as improvements in acute health services or temporary welfare assistance.

Decomposition of AGP mobility indices

To investigate mobility requires knowledge not only of the initial and final joint distributions of health and income, but also of the transition process linking the observations on these two joint distributions. AGP propose a decomposition of the change in the concentration index between two periods *s* and *f* into:[1]

$$\begin{aligned}\Delta CI &= \left(CI_{ff} - CI_{ss}\right) = \left(CI_{ff} - CI_{fs}\right) + \left(CI_{fs} - CI_{ss}\right) \\ &= \left(\frac{2}{\bar{h}_f}\operatorname{cov}(h_f, R_f) - \frac{2}{\bar{h}_f}\operatorname{cov}(h_f, R_s)\right) + \left(\frac{2}{\bar{h}_f}\operatorname{cov}(h_f, R_s) - \frac{2}{\bar{h}_s}\operatorname{cov}(h_s, R_s)\right) \\ &= M^R - M^H\end{aligned} \tag{10}$$

where CI_{fs} is the concentration index obtained when health outcomes in the final period are ranked by income in the initial period, and M^H and M^R are interpreted as *income-related health mobility* (IRHM) and *health-related income mobility* (HRIM) indices respectively.[2]

The IRHM index M^H in (10) captures the effect on cross-sectional income-related health inequality of differences in relative health changes between individuals with different levels of initial income. M^H will be positive (negative) if health changes are progressive (regressive) in the sense that the poorest individuals either enjoy a larger (smaller) share of total health gains or suffer a smaller (larger) share of total health losses compared to their initial share of health, and equals zero if relative health changes are independent of income or there are no health changes. M^H is in turn shown to depend on the progressivity and scale of health changes:

$$M^H = \left(\frac{2}{\bar{h}_s}\operatorname{cov}(h_s, R_s) - \frac{2}{\overline{\Delta h_f}}\operatorname{cov}(\Delta h_f, R_s)\right)\left(\frac{\overline{\Delta h_f}}{\bar{h}_f}\right) = \left(CI_{ss} - CI_{f-s,s}\right)\left(\frac{\overline{\Delta h_f}}{\bar{h}_f}\right) = Pq \tag{11}$$

where $CI_{f-s,s}$ is the concentration coefficient of health changes ranked by initial period income[3] and $\overline{\Delta h_f} = \bar{h}_f - \bar{h}_s$ is the average health change between the two periods; such that

[1] AGP also consider the alternative decomposition: $\Delta CI = \left(CI_{ff} - CI_{sf}\right) + \left(CI_{sf} - CI_{ss}\right) = M_{\#}^R - M_{\#}^H$. We choose to focus on (10) here though it is clear that similar procedures are readily applicable to this alternative.

[2] Allanson (2010) proposes an analogous decomposition of the change in the generalised concentration index, which provides a measure of absolute income-related health inequality. In this case the health mobility index is simply equal to minus the product of the concentration index of health changes ranked by initial income and the average health change. It is a trivial exercise to extend the standardisation and decomposition procedures developed here to this alternative measurement framework - indeed, the adoption of an absolute measure of inequality serves to greatly simplify the analysis.

[3] Note that $CI_{f-s,s}$ 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.

progressivity is captured by the Kakwani (1977)-type disproportionality index $P = \left(CI_{ss} - CI_{f-s,s}\right)$, and the scale of health changes $q = \overline{\Delta h} / \overline{h}^{f}$ is measured as the ratio of average health changes to average final period health. AGP argue that P can provide *inter alia* a useful measure of the performance of health improvement programmes in targeting the poor: a given reduction in income-related health inequality can be achieved either by a small-scale but highly targeted intervention to improve the health of the very poor or by a larger scale but broader health programme. For any given P, the gross impact of health changes on income-related health inequality is proportional to their average scale q.

The IRHM index M^{H} addresses the question of whether the pattern of relative health changes favour those with initially low or high incomes, providing a natural counterpart to measures of income-related health inequality that address the issue of whether the distribution of health favours the poor or the rich. However, M^{H} will provide a misleading measure of the extent to which relative health changes are directly attributable to initial income because health changes are determined not only by initial income but also by other factors that affect health mobility and are correlated with initial income. Thus, as with the concentration index, there is a need for procedures both to standardise IRHMs and to determine the contribution of inequalities in the individual determinants of health changes to the overall value of the IRHM.

For this purpose, we follow Hauck and Rice (2003) and Contoyannis et al. (2004) in considering a less restrictive, dynamic specification of the health function than (2), which allows for lagged as well as contemporaneous responses to changes in income and other health determinants. Specifically, we assume that the dynamic health function takes the form of an first-order autoregressive distributed lag (ARDL) model:[4]

$$h_{t+1} = \alpha_0 + \delta_y y_{t+1} + \sum_{k=1}^{K} \delta_k x_{k,t+1} + \alpha_y y_t + \sum_{k=1}^{K} \alpha_k x_{kt} + (1-\lambda) h_t + \varepsilon_{t+1} \;; \quad t=1, \ldots T\text{-}1 \tag{12}$$

which may be expressed as the error correction model (ECM):

$$\begin{aligned} \Delta h_{t+1} = \left(h_{t+1} - h_t\right) &= \delta_y \left(y_{t+1} - y_t\right) + \sum_{k=1}^{K} \delta_k \left(x_{k,t+1} - x_{kt}\right) + \lambda\left(\left(\beta_0 + \beta_y y_t + \sum_{k=1}^{K} \beta_k x_{kt}\right) - h_t\right) + \varepsilon_{t+1} \\ &= \delta_y \Delta y_{t+1} + \sum_{k=1}^{K} \delta_k \Delta x_{t+1}^{k} + \lambda\left(h_t^{*} - h_t\right) + \varepsilon_{t+1} \;; \end{aligned} \tag{13}$$

[4] Equation (12) can readily be extended to include higher-order lagged terms in y, the x_k's and h, leading to more complicated ECM representations in which the short-run dynamics are a function not only of current but also of lagged changes in the health determinants.

where $\beta_0 = \alpha_0/\lambda$, $\beta_y = \left(\alpha_y + \delta_y\right)/\lambda$ and $\beta_y = \left(\alpha_y + \delta_y\right)/\lambda$ may be interpreted as parameters of a long-run or equilibrium health function:

$$h_t^* = \beta_0 + \beta_y y_t + \sum_{k=1}^{K} \beta_k x_{kt} \tag{14}$$

such that $\left(h_t^* - h_t\right)$ corresponds to the 'equilibrium error' in the current period and $\lambda\left(0 \le \lambda \le 1\right)$ determines the rate of adjustment to equilibrium. Hence the ECM representation of the model states that the change in health over the next period depends on the effects of contemporaneous changes in income and the other health determinants, the extent of any disequilibrium in health in the current period and the size of the idiosyncratic health shock in the next period. For analytical purposes, the main attraction of this representation is the clear distinction between the short-run dynamics and the long-run equilibrium health relationship (14). In particular, it is straightforward using the ECM to explore the short-term impact on IRHI of policy interventions that impact on the determinants of health. Moreover, the concentration index of equilibrium health CI_{tt}^{h*} may be used to provide a measure of chronic or structural income-related health inequality, which may be further standardised or decomposed using the techniques set out at the beginning of this section. If $\lambda = 1$ then adjustment is instantaneous and the ECM collapses to the static model (2) with $h_t^* = h_t$ in all periods.

Accordingly, the IRHM index M^H, may be re-written from (11) as:

$$\begin{aligned} M^H &= \left(CI_{ss} - \frac{2}{\overline{\Delta h}_f}\operatorname{cov}\left(\delta_y \Delta y_f + \sum_{k=1}^{K}\delta_k \Delta x_{kf} + \lambda\left(\left(\beta_0 + \beta_y y_s + \sum_{k=1}^{K}\beta_k x_{ks}\right) - h_s\right) + \varepsilon_f, R_s\right)\right)\left(\frac{\overline{\Delta h}_f}{\overline{h}_f}\right) \\ &= \left\{CI_{ss} - \left(\frac{\delta_y \overline{\Delta y}_f}{\overline{\Delta h}_f} CI_{f-s,s}^{y} + \sum_{k=1}^{K}\frac{\delta_k \overline{\Delta x}_{kf}}{\overline{\Delta h}_f} CI_{f-s,s}^{k}\right)\right. \\ &\qquad \left. -\lambda\left(\frac{\beta_y \overline{y}_s}{\overline{\Delta h}_f} G_{ss} + \sum_{k=1}^{K}\frac{\beta_k \overline{x}_{ks}}{\overline{\Delta h}_f} CI_{ss}^{k} - \frac{\overline{h}_s}{\overline{\Delta h}_f} CI_{ss}\right)\right\}\left(\frac{\overline{\Delta h}_f}{\overline{h}_f}\right) \end{aligned} \tag{15}$$

where, additionally, $\overline{\Delta y}_f$ and $\overline{\Delta x}_{kf}$ are the average changes in income and health determinant k respectively between the two periods; $CI_{f-s,s}^{y}$ and $CI_{f-s,s}^{k}$ are the corresponding concentration indices of changes ranked by initial income; and $\operatorname{Cov}(\varepsilon_f, R_s) = 0$ by assumption (see Gravelle, 2003).

All other things equal, (11) shows that the IRHM index M^H will be more positive (or less negative) the more pro-poor is the distribution of health changes as reflected in the value of the concentration index of health changes, $CI_{f-s,s}$. $CI_{f-s,s}$ is in turn expressed in (15) as a weighted sum of the concentration indices of all the determinants (including initial health) of the dynamic health function (13), with weights equal to the share of the overall health change attributable to each health change determinant and defined as the product of the response coefficient on each determinant and the ratio of the average value of the determinant to the average health change. This decomposition will be feasible so long as the dynamic health function is linear in the parameters with an additive error term. In particular, we note that if health changes are a linear function not of income *y* itself but of some function of income *g*(*y*) then the results in this section will go through unchanged but with *g*(*y*) replacing *y* throughout.

Noting that M^H is a measure of relative not absolute health mobility, we proceed to define a partial mobility index conditional upon the actual distribution of health in the initial period as:

$$M_{PTL}^{H} = \left(CI_{ss} - \hat{\lambda}\left(\frac{\hat{\beta}_y \bar{y}_s}{\overline{\Delta h_f}} G_{ss} - \frac{\bar{h}_s}{\overline{\Delta h_f}} CI_{ss} \right) \right) \left(\frac{\overline{\Delta h_f}}{\bar{h}_f} \right) = \left(CI_{ss} - CI_{f-s,s}^{\Delta h^{PTL}} \right) \left(\frac{\overline{\Delta h_f}}{\bar{h}_f} \right) = P^{PTL} q \qquad (16)$$

where $CI_{f-s,s}^{\Delta h^D}$ is the concentration index of directly standardised health changes:

$$\Delta h_f^{PTL} = \hat{\delta}_y \overline{\Delta y_f} + \sum_{k=1}^{K} \hat{\delta}_k \overline{\Delta x_{kf}} + \hat{\lambda}\left(\left(\hat{\beta}_0 + \hat{\beta}_y y_f + \sum_{k=1}^{K} \hat{\beta}_k \bar{x}_{kf} \right) - h_s \right) + \hat{\varepsilon}_f \; ; \qquad (17)$$

and $\hat{\delta}_y$, $\hat{\delta}_k$'s, $\hat{\lambda}$, $\hat{\beta}_0$, $\hat{\beta}_y$, and $\hat{\beta}_k$'s are estimates of the corresponding parameters of the ECM (13) from a regression using balanced panel data on a sample of individuals drawn from the population, and $\hat{\varepsilon}_f$ is the regression residual. Equation (16) provides a benchmark measure of income-related mobility in health after the elimination of inequalities in all health change determinants other than initial income and health: we do not standardise for initial health in the definition of Δh_f^{PTL} in (17) as this would imply a counterfactual value for the concentration index of initial health CI_{ss} of zero in (16). P^{PTL} may in turn be interpreted as a partial progressivity index, being defined as the difference between the concentration indices of initial health and standardised health changes, and therefore of the same form as *P* in (11). All other things equal, P^{PTL} will be more positive (or less negative) the smaller the response of long-run or equilibrium health to changes in income as measured by the response coefficient $\hat{\beta}_y$; the higher the ratio of average initial income to health changes; the less unequal the initial distribution of income as measured by the income Gini coefficient G_{ss}; the lower the ratio of average initial health to health changes; the more unequal the initial distribution of health by income as measured by the income-related health concentration index CI_{ss}; and, if $\hat{\beta}_y \bar{y}_s G_{ss} > \bar{h}_s CI_{ss}$ ($\hat{\beta}_y \bar{y}_s G_{ss} < \bar{h}_s CI_{ss}$), the faster (slower) the rate of adjustment to equilibrium as measured by λ. For any given P^{PTL}, the gross impact on final period income-related health inequalities is again proportional to the scale of health changes q.

The conditional partial mobility index M^H_{PTL} in (16) may be augmented to evaluate the vertical stance implied by initial policy conditions, and hence the contribution of current policy to health mobility in the next period. Following Gravelle (2003), we assume that the set of K other health determinants can be partitioned between a sub-set of policy relevant variables and a sub-set of policy irrelevant or standardising variables, where the choice of partition will depend upon the policy context. Again assuming that the first J ($0 \le J \le K$) other health determinants are deemed policy relevant then the effect of the remaining $K–J$ standardising variables on the overall mobility index can be removed by deduction of the terms involving them from (15) to yield the partial mobility index:

$$\begin{aligned} M^H_{A(L)} &= \left(CI_{ss} - \hat{\lambda}\left(\frac{\hat{\beta}_y \bar{y}_s}{\overline{\Delta h_f}} G_{ss} + \sum_{k=1}^{J} \frac{\hat{\beta}_k \bar{x}_{ks}}{\overline{\Delta h_f}} CI^k_{ss} - CI_{ss} \right)\right)\left(\frac{\overline{\Delta h_f}}{\bar{h}_f} \right) \\ &= \left(CI_{ss} - CI^{\Delta h^{A(L)}}_{f-s,s} \right)\left(\frac{\overline{\Delta h_f}}{\bar{h}_f} \right) = P^{A(L)} q \end{aligned} \tag{18}$$

where standardised health changes are given in this case as:

$$\Delta h^{A(L)}_f = \hat{\delta}_y \overline{\Delta y}_f + \sum_{k=1}^{K} \hat{\delta}_k \overline{\Delta x_{kf}} + \hat{\lambda}\left(\left(\hat{\beta}_0 + \hat{\beta}_y y_f + \sum_{k=1}^{J} \hat{\beta}_k x_{kf} + \sum_{k=J+1}^{K} \hat{\beta}_k \bar{x}_{kf} \right) - h_s \right) + \hat{\varepsilon}_f \ ; \tag{19}$$

which reduces to the partial mobility index M^H_{PTL} if there are no policy relevant variables other than income, We note that the difference between $M^H_{A(L)}$ and M^H_{PTL} will provide an indicator of the contribution of existing policies to mobility in the following period, with $P^{A(L)}$ again interpretable as a measure of the progressivity of the existing policy stance in terms of the changes in health that it will induce in the following period. Moreover, as the length of this period is extended, such that $f–s \to \infty$, we note that $J^{A(L)}_{fs}$ will converge to the augmented partial long-run mobility index:

$$M^H_{A^*(L)} = \left(CI_{ss} - \left(\frac{\hat{\beta}_y \overline{y}_s}{\overline{\Delta h_f}} G_{ss} + \sum_{k=1}^{J} \frac{\hat{\beta}_k \overline{x}_{ks}}{\overline{\Delta h_f}} CI^k_{ss} - CI_{ss} \right) \right) \left(\frac{\overline{\Delta h_f}}{\overline{h}_f} \right)$$

$$= \left(CI_{ss} - CI^{\Delta h^{A(L)*}}_{f-s,s} \right) \left(\frac{\left(\overline{h}^*_s - \overline{h}_s \right)}{\overline{h}^*_s} \right) = P^{A(L)*} q \tag{20}$$

where $\overline{h}^*_s$ is mean equilibrium health, given the initial conditions, and standardised equilibrium health changes are given as:

$$\Delta h^{A^*(L)}_f = \left(\hat{\beta}_0 + \hat{\beta}_y y_f + \sum_{k=1}^{J} \hat{\beta}_k x_{kf} + \sum_{k=J+1}^{K} \hat{\beta}_k \overline{x}_{kf} \right) - h_s \ ; \tag{21}$$

such that $M^H_{A^*(L)}$ captures how much income-related health mobility there would be if there was full adjustment of health to the equilibrium levels implied by initial income levels and the current policy mix.

The analysis may be further extended to additionally take account of the contemporaneous effect of changes in income and other policy relevant variables on health mobility, yielding a second augmented partial mobility index:

$$M^H_{A(L,D)} = \left(CI_{ss} - \left(\frac{\hat{\delta}_y \overline{\Delta y}}{\overline{\Delta h_f}} CI^y_{f-s,s} + \sum_{k=1}^{K} \frac{\hat{\delta}_k \overline{\Delta x_{ks}}}{\overline{\Delta h_f}} CI^k_{f-ss} + \lambda \left(\frac{\hat{\beta}_y \overline{y}_s}{\overline{\Delta h_f}} G_{ss} + \sum_{k=1}^{J} \frac{\hat{\beta}_k \overline{x}_{ks}}{\overline{\Delta h_f}} CI^k_{ss} - CI_{ss} \right) \right) \right) \left(\frac{\overline{\Delta h_f}}{\overline{h}^f} \right)$$

$$= \left(CI_{ss} - CI^{\Delta h^{A(L,D)}}_{f-s,s} \right) \left(\frac{\overline{\Delta h_f}}{\overline{h}^f} \right) = P^{A(L,D)} q \tag{22}$$

where standardised health changes are given in this case as:

$$\Delta h^{A(L,D)}_f = \hat{\delta}_y \Delta y_f + \sum_{k=1}^{J} \hat{\delta}_k \Delta x_{kf} + \sum_{k=J+1}^{K} \hat{\delta}_k \overline{\Delta x_{kf}} + \hat{\lambda} \left(\left(\hat{\beta}_0 + \hat{\beta}_y y_f + \sum_{k=1}^{J} \hat{\beta}_k x_{kf} + \sum_{k=J+1}^{K} \hat{\beta}_k \overline{x}_{kf} \right) - h_s \right) + \hat{\varepsilon}_f \ ; \tag{23}$$

such that the difference between $M^H_{A(L,D)}$ and $M^H_{A(L)}$ will provide an indicator of the contribution of income and policy changes to mobility in the following period, with a comparison of $P^{A(L,\Delta)}$ and $P^{A(L)}$ providing an indication of whether the immediate effect of the policy changes is to increase or decrease the progressivity of the policy stance.

Equation (15) provides the basis for the standardisation procedures, but it is more instructive to base the full regression-based decomposition analysis on an alternative representation of the IRHM:

$$M^H = Pq = \Bigg\{ \left(CI_{ss} - CI_{f-s,s}^{y}\right)\frac{\hat{\delta}_y \overline{\Delta y}_f}{\overline{\Delta h}_f} + \sum_{k=1}^{K}\left(CI_{ss} - CI_{f-s,s}^{k}\right)\frac{\hat{\delta}_k \overline{\Delta x}_{kf}}{\overline{\Delta h}_f} + \left(CI_{ss} - C_{ss}^{\hat{\alpha}_0}\right)\frac{\hat{\alpha}_0}{\overline{\Delta h}_f}$$

$$+\left(CI_{ss} - G_{ss}\right)\frac{\hat{\lambda}\hat{\beta}_y \overline{y}_s}{\overline{\Delta h}_f} + \sum_{k=1}^{K}\left(CI_{ss} - CI_{ss}^{k}\right)\frac{\hat{\lambda}\hat{\beta}_k \overline{x}_{ks}}{\overline{\Delta h}_f} - \left(CI_{ss} - CI_{ss}\right)\frac{\hat{\lambda}\overline{h}_s}{\overline{\Delta h}_f} + \frac{\hat{\varepsilon}_f}{\overline{\Delta h}_f}\Bigg\}\left(\frac{\overline{\Delta h}_f}{\overline{h}_f}\right) \quad (24)$$

since:

$$\hat{\delta}_y \frac{\overline{\Delta y}_f}{\overline{\Delta h}_f} + \sum_{k=1}^{K}\hat{\delta}_k \frac{\overline{\Delta x}_{kf}}{\overline{\Delta h}_f} + \hat{\alpha}_0 \frac{1}{\overline{\Delta h}_f} + \hat{\lambda}\hat{\beta}_y \frac{\overline{y}_s}{\overline{\Delta h}_f} + \sum_{k=1}^{K}\hat{\lambda}\hat{\beta}_k \frac{\overline{x}_{ks}}{\overline{\Delta h}_f} - \hat{\lambda}\frac{\overline{h}_s}{\overline{\Delta h}_f} = 1;$$

and where $\hat{\alpha}_0 = \hat{\lambda}\hat{\beta}_0$. Hence (24) may be written as:

$$M^H = \Bigg\{ \left(CI_{ss} - CI_{f-s,s}^{y}\right)\frac{\hat{\delta}_y \overline{\Delta y}_f}{\overline{h}_f} + \sum_{k=1}^{K}\left(CI_{ss} - CI_{f-s,s}^{k}\right)\frac{\hat{\delta}_k \overline{\Delta x}_{kf}}{\overline{h}_f} + \left(CI_{ss} - C_{ss}^{\hat{\alpha}_0}\right)\frac{\hat{\alpha}_0}{\overline{h}_f}$$

$$+\left(CI_{ss} - G_{ss}\right)\frac{\hat{\lambda}\hat{\beta}_y \overline{y}_s}{\overline{h}_f} + \sum_{k=1}^{K}\left(CI_{ss} - CI_{ss}^{k}\right)\frac{\hat{\lambda}\hat{\beta}_k \overline{x}_{ks}}{\overline{h}_f} - \frac{GC_{fs}^{\hat{\varepsilon}}}{\overline{h}_f}\Bigg\} \quad (25)$$

$$= P_{\Delta y}q_{\Delta y} + \sum_{k=1}^{K}P_{\Delta x_k}q_{\Delta x_k} + P_{\hat{\alpha}_0}q_{\hat{\alpha}_0} + P_y q_y + \sum_{k=1}^{K}P_{x_k}q_{x_k} - \frac{GC_{fs}^{\hat{\varepsilon}}}{\overline{h}_f}$$

where $GC_{fs}^{\hat{\varepsilon}}$ is the generalised concentration index of the regression residual ranked by initial income, and $P_{\hat{\alpha}_0} = CI_{ss}$ since $C_{ss}^{\hat{\alpha}_0} = 0$ by definition.[5] Thus the contribution of each determinant of the dynamic health function to M^H can be expressed as the product of the progressivity and scale of health changes attributable to that determinant, plus a final term that is due to the regression residual and reflects the unpredictability of future health states. An individual health determinant with a positive (negative) scale factor - i.e. the contribution of the determinant to the overall health change is positive (negative) - will also have a positive impact on IRHM if it is distributed less (more) unequally than initial health and will have a negative impact on IRHM otherwise. Combining terms in (25), a second, alternative representation that will prove useful is given by:

$$M^H = P_{\Delta y}q_{\Delta y} + \sum_{k=1}^{K}P_{\Delta x_k}q_{\Delta x_k} + P_{\left(h_s^* - h_s\right)}q_{\left(h_s^* - h_s\right)} - \frac{GC_{fs}^{\hat{\varepsilon}}}{\overline{h}_f} \quad (26)$$

[5] Note that there is not a separate term in initial health h_s since $P_h = \left(CI_{ss} - CI_{ss}\right) = 0$ by definition. Hence the sum of the scale factors, $q_{\Delta y} + \sum_{k=1}^{K}q_{\Delta x_k} + q_{\hat{\alpha}_0} + q_y + \sum_{k=1}^{K}q_{x_k} = q - q_h$ not q, where $q_h = -\hat{\lambda}\,\overline{h}_s / \overline{h}_f$.

where, additionally, $P_{\left(h_s^*-h_s\right)} = \left(CI_{ss} - C_{ss}^{\left(h_s^*-h_s\right)}\right)$ and $q_{\left(h_s^*-h_s\right)} = \hat{\lambda}\left(\overline{h_s^* - h_s}\right) \Big/ \overline{h}_f$ are the disproportionality index and scale factor of the equilibrium error.

Returning to the HRIM index M^R in (10), this captures the effect of the reshuffling of individuals within the income distribution on cross-sectional socioeconomic inequalities in health. M^R will be positive (negative) if the effect of income re-ranking is to exacerbate (moderate) income-related health inequalities compared to what they would have been otherwise, implying that those who moved up the income ranking tended to be healthier (less healthy) in the final period compared to those who moved down. M^R may readily be re-written to give:

$$M^R = \left(CI_{ff} - CI_{fs}\right) = \left(\frac{2}{\overline{h}_f}\operatorname{cov}\left(h_f, R_f - R_s\right)\right) = \frac{2}{N\left(\overline{h}_s + \overline{\Delta h}_f\right)} \sum_i (h_{is} + \Delta h_{if})\left(R_{if} - R_{is}\right) \tag{27}$$

which allows for the construction of a set of partial HRIM indices, analogous to the partial IRHM indices defined above, by substitution of the various definitions of standardised health changes given by (17), (19), (21) and (23) for Δh_f:

$$M_{PTL}^R = \left(\frac{2}{\left(\overline{h}_s + \overline{\Delta h}_f\right)}\operatorname{cov}\left((h_s + \Delta h_f^{PTL}), R_f - R_s\right)\right) \tag{28}$$

$$M_{A(L)}^R = \left(\frac{2}{\left(\overline{h}_s + \overline{\Delta h}_f\right)}\operatorname{cov}\left((h_s + \Delta h_f^{A(L)}), R_f - R_s\right)\right) \tag{29}$$

$$M_{A^*(L)}^R = \left(\frac{2}{\left(\overline{h}_s + \overline{\Delta h}_f\right)}\operatorname{cov}\left((h_s + \Delta h_f^{A^*(L)}), R_f - R_s\right)\right) \tag{30}$$

and

$$M_{A(L,D)}^R = \left(\frac{2}{\left(\overline{h}_s + \overline{\Delta h}_f\right)}\operatorname{cov}\left((h_s + \Delta h_f^{A(L,D)}), R_f - R_s\right)\right) \tag{31}.$$

Thus M_{PTL}^R in (28) provides a benchmark measure of HRIM after removing the effects of all variables which affect health changes other than initial income and health. $M_{A(L)}^R$ in (29) reduces to the partial mobility index M_{PTL}^R if there are no policy relevant variables other than income, with the difference between the two providing an indicator of the contribution of existing policies to HRIM in the following period, and $M_{A^*(L)}^R$ in (30) captures how much

HRIM there would be if there was full adjustment of health to the equilibrium levels implied by initial income levels and the current policy mix. Finally, the difference between $M^{R}_{A(L,D)}$ in (31) and $M^{R}_{A(L)}$ measures the contribution of income and policy changes to HRIM in the following period. We note that (27) could be expanded to explicitly show the impact of particular health change determinants on HRIM, but it is not clear that this elaboration would be of particular interest. A complete analysis would further allow for a decomposition by the determinants of the change in income ranks, based on a joint model of the determination of health and income changes, but this lies beyond the scope of the current paper.

In conclusion, the various regression-based standardisation and decomposition techniques outlined in this section complement the methods (7)-(9) used in Wagstaff *et al.* (2003), with the main appeal of our approach being that the use of longitudinal data allows us to track individual outcomes and thereby characterise the redistributive effects of the co-evolution of health and incomes. In a recent paper, van Ourti et al. (2009) have proposed an alternative decomposition procedure based on the use of longitudinal data, but the focus of their analysis is on the consequences of changes in the income distribution for changes in the distribution of health by income. In particular, they explore how the effect of income changes on IRHI varies depending on the nature of income growth and the assumed form of the relationship between health and income, i.e. on the specification of the function $g(y)$. In contrast, the main focus of our analysis is on the consequences of changes in the health distribution for changes in IRHI, providing a measure of health change progressivity that may be further broken down into the individual contributions of specific health change determinants.

3. Accounting for deaths: extended decomposition techniques

The preceding section is based on the implicit assumption that the population is invariant over time, with all individuals presumed to be alive in both the initial and final periods. PAG argue that this is potentially misleading for the evaluation of policies designed to tackle health inequalities since it fails to take account of those who are alive in the initial period but die before the final period, thereby ignoring perhaps the most important of all health outcomes. Accordingly, they extend the AGP framework to additionally consider the effect of deaths on the longitudinal analysis of income-related health inequality. In this section we modify and extend our regression-based decomposition and standardisation procedures to deal with the resultant set of mobility indices.

Assigning the dead a health status of zero, PAG propose a decomposition of the change in the cross-sectional concentration index between the initial and final periods into:

$$CI_{ff}^{i\in A} - CI_{ss} = \left(CI_{ff}^{i\in A} - CI_{fs}\right) + \left(CI_{fs} - CI_{ss}\right) = \tilde{M}^{R} - \tilde{M}^{H} \tag{32}$$

where the superscript notation $i \in A$ denotes the sub-set of the population in the initial period who are still alive in the final period, so $CI_{ff}^{i\in A}$ is income-related health concentration index for the extant population in the final period, whereas all other concentration indices are now defined over the entire population alive in the initial period whether or not they survive to the final period. $\tilde{M}^{H}$ and $\tilde{M}^{R}$ in (32) are again interpreted as IRHM and HRIM indices respectively, but defined over different populations than in Section 2.

The IRHM index $\tilde{M}^{H}$ in (32) is defined over the entire population alive in the initial period and captures the effect on income-related health inequality of differences in relative health changes between individuals with different levels of initial income, where health changes due to both morbidity and mortality are now taken into account. Specifically, PAG show that $\tilde{M}^{H}$ can be written as:

$$\begin{aligned}
\tilde{M}^{H} &= \left(CI_{ss} - CI_{fs}\right) = \left(CI_{ss} - CI_{f-s,s}\right)\left(\overline{\Delta h_f / \bar{h}_f}\right) = \tilde{P}\tilde{q} \\
&= \left(\tilde{P}^{MB}\left(\frac{\tilde{q}^{MB}}{\tilde{q}}\right) + \tilde{P}^{MT}\left(\frac{\tilde{q}^{MT}}{\tilde{q}}\right)\right)\left(\tilde{q}^{MB} + \tilde{q}^{MT}\right) = \tilde{P}^{MB}\tilde{q}^{MB} + \tilde{P}^{MT}\tilde{q}^{MT} \\
&= \left(CI_{ss} - CI_{f-s,s}^{MB}\right)\left(\frac{\overline{\Delta h_f^{MB}}}{\bar{h}_f}\right) + \left(CI_{ss} - CI_{f-s,s}^{MT}\right)\left(\frac{\overline{\Delta h_f^{MT}}}{\bar{h}_f}\right) = \tilde{M}^{HMB} + \tilde{M}^{HMT}
\end{aligned} \tag{33}$$

where Δh_f^{MB} and Δh_f^{MT} denote morbidity-related and mortality-related health changes respectively, which are defined so that only one of the measures can be non-zero for any individual, and averages are taken over all individuals. Thus $\tilde{M}^{H}$ may be used to address questions of whether the patterns of relative morbidity-related and mortality-related health changes favour those with initially low or high incomes. But, as with M^{H} in Section 2, there will be a need for decomposition and standardisation procedures to obtain an accurate picture of the contribution of inequalities in both income and other individual determinants of health changes to the overall level of income-related health mobility.

For this purpose, we incorporate the dynamic health function (13) into a Two-Part Model (TPM; see Leung and Yu (1996) and Puhani (2000) for discussion):

$$S_{t+1}^{*} = \gamma_0 + \gamma_y y_t + \sum_{k=1}^{K} \gamma_k x_{kt} + \gamma_h h_t + u_{t+1}\,; \quad u_{t+1} \sim N(0,1) \tag{34}$$

$$\Delta h_{t+1} = \begin{cases} \delta_y \Delta y_{t+1} + \sum_{k=1}^{K} \delta_k \Delta x_{k,t+1} + \lambda\left(\left(\beta_0 + \beta_y y_t + \sum_{k=1}^{K} \beta_k x_{kt}\right) - h_t\right) + \varepsilon_{t+1}\ ;\ if\ S_{t+1}^* \geq 0 & (35a) \\ -h_t\ ;\ if\ S_{t+1}^* < 0 & (35b) \end{cases}$$

where the first part (34) is assumed to take the form of a standard probit model with the link between the latent index variable S_{t+1}^* and observable survival status S_{t+1} following the rule that $S_{t+1} = 1$ if $S_{t+1}^* \geq 0$ and $S_{t+1} = 0$ otherwise. Thus (34) determines the survival status of individuals in the final period as a function of initial income y_t, initial health h_t and a set of other determinants x_t^k (l=1,….K), which it is plausible to assume are the same as the set of other determinants in the dynamic health function (13). For the second part (35), which determines the unconditional change in health outcome Δh_{t+1}, it is assumed that $E\left(\varepsilon_{t+1} \mid S_{t+1}^* \geq 0\right) = 0$ in (35a), but not necessarily that ε_{t+1} is normally distributed, and that the health status of those that do not survive in (35b) is identically equal to 0 in the final period.

Equations (34) and (35a) are conditionally independent, which serves to identify the latter in the absence of variables that might conceivably influence mortality but not morbidity. Accordingly, the health change of any individual alive in the initial period may be written as:

$$\begin{aligned} \Delta h_{t+1} &= E\left(\Delta h_{t+1}\right) + \upsilon_{t+1} \\ &= \left\{\text{Prob}\left(S_{t+1} = 1\right) E\left(\Delta h_{t+1} \mid S_{t+1} = 1\right)\right\} + \left\{\text{Prob}\left(S_{t+1} = 0\right)\left(0 - h_t\right)\right\} + \upsilon_{t+1} \\ &= \left\{\Phi\left(z_t\right)\left(\delta_y \Delta y_{t+1} + \sum_{k=1}^{K} \delta_k \Delta x_{k,t+1} + \lambda\left(h_t^* - h_t\right)\right)\right\} - \left\{\left(1 - \Phi\left(z_t\right)\right) h_t\right\} + \upsilon_{t+1} \\ &= E\left(\Delta h_{t+1}^{MB}\right) + E\left(\Delta h_{t+1}^{MT}\right) + \upsilon_{t+1} \end{aligned} \qquad (36)$$

where $z_t = \gamma_0 + \gamma_y y_t + \sum_{k=1}^{K} \gamma_k x_{kt} + \gamma_h h_t$, and $\Phi(\cdot)$ denotes the cumulative density function of the standard normal distribution. The decomposition of expected health changes into morbidity-related and mortality-related components parallels that in (33), though both $E\left(\Delta h_{t+1}^{MB}\right)$ and $E\left(\Delta h_{t+1}^{MT}\right)$ in (36) will typically be non-zero for all individuals given that matters of life and death are never certain.

Within this extended modelling framework, the standardised health changes given by (17), (19), (21) and (23) may simply be reinterpreted as standardised health changes conditional on survival to the final period. Using obvious notation, let the corresponding set of standardised survival probabilities be:

$$\text{Prob}\left(S_f^{PTL}=1\right)=\Phi\left(\hat{\gamma}_0+\hat{\gamma}_y y_s+\sum_{k=1}^{K}\hat{\gamma}_k\bar{x}_{ks}+\hat{\gamma}_h h_s\right) \tag{37}$$

$$\begin{aligned}\text{Prob}\left(S_f^{A(L)}=1\right)&=\text{Prob}\left(S_f^{A*(L)}=1\right)=\text{Prob}\left(S_f^{A(L,D)}=1\right)\\&=\Phi\left(\hat{\gamma}_0+\hat{\gamma}_y y_s+\sum_{k=1}^{J}\hat{\gamma}_k x_{ks}+\sum_{k=J+1}^{K}\hat{\gamma}_k\bar{x}_{ks}+\hat{\gamma}_h h_s\right)\end{aligned} \tag{38}$$

where $\hat{\gamma}_0$, $\hat{\gamma}_y$, $\hat{\gamma}_k$'s and $\hat{\gamma}_h$ are estimates of the corresponding parameters of the probit model (34).[6] We proceed to define a set of standardised unconditional health changes $\widetilde{\Delta h}_f^{PTL}$, $\widetilde{\Delta h}_f^{A(L)}$, $\widetilde{\Delta h}_f^{A*(L)}$ and $\widetilde{\Delta h}_f^{A(L,D)}$ equal to the corresponding standardised conditional health changes multiplied by the relevant standardised probability of survival, less the initial health level multiplied by one minus the relevant standardised probability of survival. Evaluation of these measures is unproblematic except in the case of individuals who actually die before the final period, whom we assume would have experienced both average changes in income and all other health determinants if they had gone on living and no health shock in the final period. But note that the averages of the standardised changes will not in general equal the observed unconditional mean health change due to the non-linearity of the probit survival model. Calculation of the IRHM index in (33) using these standardised unconditional health changes, rather than the observed changes, then yields partial mobility measures $\tilde{M}_{PTL}^H$, $\tilde{M}_{A(L)}^H$, $\tilde{M}_{A*(L)}^H$ and $\tilde{M}_{A(L,D)}^H$ analogous to those defined by (16), (18), (20) and (22) respectively, where these indices may be further decomposed into morbidity-related and mortality-related components.

Equation (36) provides a direct counterpart of (13) but does not similarly lend itself to a decomposition analysis of $\tilde{M}^H$ due to the inherent non-linearity of the TPM. To overcome this problem we adopt a hierarchical decomposition procedure in which we first break down $\tilde{M}^H$ on the basis of (36) into elements due to health changes resulting from expected morbidity changes, mortality and health shocks, and then further decompose the first two of these elements to determine the contributions of the individual health change determinants. The first stage of this procedure straightforwardly yields:

[6] Equation (38) provides a definition not only of $\text{Prob}\left(S_{t+1}^{A(L)}=1\right)$, but also of $\text{Prob}\left(S_{t+1}^{A*(L)}=1\right)$ and $\text{Prob}\left(S_{t+1}^{A(L,D)}=1\right)$, since the probit model (34) is a static function of initial conditions alone.

$$
\begin{aligned}
\tilde{M}^H &\approx \left(CI_{ss} - \frac{2}{\overline{\Delta h_f}} \operatorname{cov}\left(E\left(\Delta h_f^{MB}\right) + E\left(\Delta h_f^{MT}\right) + \hat{\upsilon}_f, R_s \right) \right) \left(\frac{\overline{\Delta h_f}}{\overline{h}_f} \right) \\
&= \left(CI_{ss} - CI_{f-s,s}^{E(MB)} \right) \frac{\overline{E\left(\Delta h_f^{MB}\right)}}{\overline{h}_f} + \left(CI_{ss} - CI_{f-s,s}^{E(MT)} \right) \frac{\overline{E\left(\Delta h_f^{MT}\right)}}{\overline{h}_f} + \left(CI_{ss} - CI_{fs}^{\hat{\upsilon}} \right) \frac{\overline{\hat{\upsilon}}_f}{\overline{h}_f} \\
&\equiv \tilde{P}_{E(MB)} \tilde{q}_{E(MB)} + \tilde{P}_{E(MT)} \tilde{q}_{E(MT)} + \tilde{P}_{\hat{\upsilon}} \tilde{q}_{\hat{\upsilon}} = \tilde{P} \tilde{q}
\end{aligned}
\tag{39}
$$

where $CI_{f-s,s}^{E(MB)}$, $CI_{f-s,s}^{E(MT)}$ and $CI_{fs}^{\hat{\upsilon}}$ are the concentration indices of expected morbidity changes, expected mortality and health shocks respectively, and $\overline{\hat{\upsilon}}_f$ is the sample average value of $\hat{\upsilon}_f$, which will not generally equal zero. In the second stage we make use of a first-order Taylor-series expansion about the sample means of the explanatory variables to obtain the following linear approximations:

$$
\begin{aligned}
E\left(\Delta h_f^{MB}\right) &\approx \omega_0^{MB} + \varpi_y^{MB} \Delta y_f + \sum_{k=1}^{K} \varpi_k^{MB} \Delta x_{kf} + \omega_y^{MB} y_t + \sum_{k=1}^{K} \omega_k^{MB} x_s^k - \omega_h^{MB} h_s \\
E\left(\Delta h_f^{MT}\right) &\approx \omega_0^{MT} + \omega_y^{MT} y_s + \sum_{k=1}^{K} \omega_k^{MT} x_s^k - \omega_h^{MT} h_s
\end{aligned}
\tag{40}
$$

where:

$$\omega_0^{MB} = \Phi\left(\overline{z}_s\right) \lambda \beta_0 - \left(\delta_y \overline{\Delta y}_f + \sum_{k=1}^{K} \delta_k \overline{\Delta x}_{kf} + \lambda \left(\overline{h_s^* - h_s} \right) \right) \phi\left(\overline{z}_s\right) \left(\gamma_y \overline{y}_s + \sum_{k=1}^{K} \gamma_k \overline{x}_{ks} + \gamma_h \overline{h}_s \right);$$

$$\varpi_y^{MB} = \Phi\left(\overline{z}_s\right) \delta_y;$$

$$\varpi_k^{MB} = \Phi\left(\overline{z}_s\right) \delta_k; \quad k = 1, \ldots K;$$

$$\omega_y^{MB} = \Phi\left(\overline{z}_s\right) \lambda \beta_y + \left(\delta_y \overline{\Delta y}_f + \sum_{k=1}^{K} \delta_k \overline{\Delta x}_{kf} + \lambda \left(\overline{h_s^* - h_s} \right) \right) \phi\left(\overline{z}_s\right) \gamma_y;$$

$$\omega_k^{MB} = \Phi\left(\overline{z}_s\right) \lambda \beta_k + \left(\delta_k \overline{\Delta y}_f + \sum_{k=1}^{K} \delta_k \overline{\Delta x}_{kf} + \lambda \left(\overline{h_s^* - h_s} \right) \right) \phi\left(\overline{z}_s\right) \gamma_k; k = 1, \ldots. K;$$

$$\omega_h^{MB} = -\Phi\left(\overline{z}_s\right) \lambda + \left(\delta_k \overline{\Delta y}_f + \sum_{k=1}^{K} \delta_k \overline{\Delta x}_{kf} + \lambda \left(\overline{h_s^* - h_s} \right) \right) \phi\left(\overline{z}_s\right) \gamma_h;$$

$$\omega_0^{MT} = -\overline{h}_s \, \phi\left(\overline{z}_s\right) \left(\gamma_y \overline{y}_s + \sum_{k=1}^{K} \gamma_k \overline{x}_{ks} + \gamma_h \overline{h}_s \right);$$

$$\omega_y^{MT} = \overline{h}_s \, \phi\left(\overline{z}_s\right) \gamma_y;$$

$$\omega_k^{MT} = \overline{h}_s \, \phi\left(\overline{z}_s\right) \gamma_k; k = 1, \ldots. K;$$

$$\omega_h^{MT} = -\Phi\left(-\overline{z}_s\right) + \overline{h}_s \, \phi\left(\overline{z}_s\right) \gamma_h;$$

with $\bar{z}_s = \gamma_1 + \gamma_y \bar{y}_s + \sum_{k=1}^{K} \gamma_k \bar{x}_{ks} + \gamma_h \bar{h}_s$, and $\phi(\cdot)$ denoting the probability density function of the standard normal distribution.[7] Hence $\tilde{P}_{E(MB)}\tilde{q}_{E(MB)}$ and $\tilde{P}_{E(MT)}\tilde{q}_{E(MT)}$ in (39) will be approximately equal to:

$$
\begin{aligned}
\tilde{P}_{E(MB)}\tilde{q}_{E(MB)} &\approx \left(CI_{ss} - CI_{f-s,s}^{y}\right)\frac{\hat{\varpi}_y^{MB}\overline{\Delta y}_f}{\bar{h}_f} + \sum_{k=1}^{K}\left(CI_{ss} - CI_{f-s,s}^{k}\right)\frac{\hat{\varpi}_k^{MB}\overline{\Delta x}_{kf}}{\bar{h}_f} \\
&\quad + \left(CI_{ss} - C_{ss}^{\hat{\omega}_o}\right)\frac{\hat{\omega}_0^{MB}}{\bar{h}_f} + \left(CI_{ss} - G_{ss}\right)\frac{\hat{\omega}_y^{MB}\bar{y}_s}{\bar{h}_f} + \sum_{k=1}^{K}\left(CI_{ss} - CI_{ss}^{k}\right)\frac{\hat{\omega}_k^{MB}\bar{x}_{ks}}{\bar{h}_f} \\
&\equiv \tilde{P}_{\Delta y}^{E(MB)}\tilde{q}_{\Delta y}^{E(MB)} + \sum_{k=1}^{K}\tilde{P}_{\Delta x_k}^{E(MB)}\tilde{q}_{\Delta x_k}^{E(MB)} + \tilde{P}_{\hat{\omega}_0}^{E(MB)}\tilde{q}_{\hat{\omega}_0}^{E(MB)} + \tilde{P}_{y}^{E(MB)}\tilde{q}_{y}^{E(MB)} + \sum_{k=1}^{K}\tilde{P}_{x_k}^{E(MB)}\tilde{q}_{x_k}^{E(MB)} \qquad (41) \\
\tilde{P}_{E(MT)}\tilde{q}_{E(MT)} &\approx \left(CI_{ss} - C_{ss}^{\hat{\omega}_o}\right)\frac{\hat{\omega}_0^{MT}}{\bar{h}_f} + \left(CI_{ss} - G_{ss}\right)\frac{\hat{\omega}_y^{MT}\bar{y}_s}{\bar{h}_f} + \sum_{k=1}^{K}\left(CI_{ss} - CI_{ss}^{k}\right)\frac{\hat{\omega}_k^{MT}\bar{x}_{ks}}{\bar{h}_f} \\
&\equiv \tilde{P}_{\hat{\omega}_0}^{E(MT)}\tilde{q}_{\hat{\omega}_0}^{E(MT)} + \tilde{P}_{y}^{E(MT)}\tilde{q}_{y}^{E(MT)} + \sum_{k=1}^{K}\tilde{P}_{x_k}^{E(MT)}\tilde{q}_{x_k}^{E(MT)}
\end{aligned}
$$

where $\hat{\varpi}_y^{MB}$, $\hat{\varpi}_k^{MB}$'s, $\hat{\omega}_0^{MB}$, $\hat{\omega}_y^{MB}$, $\hat{\omega}_k^{MB}$'s, $\hat{\omega}_h^{MB}$, $\hat{\omega}_0^{MT}$, $\hat{\omega}_y^{MT}$, $\hat{\omega}_k^{MT}$'s and $\hat{\omega}_h^{MT}$ are estimates of the corresponding parameters in (40). Finally, we calibrate the sets of scale factors in (41) to make the second-stage decompositions exact.

Returning to the HRIM index $\tilde{M}^R$ in (32), this may be re-defined, following PAG, over only that part of the population alive in both the initial and final periods:

$$
\tilde{M}^R = CI_{ff}^{i\in A} - CI_{fs} = \frac{2}{N^{i\in A}\left(\bar{h}_s^{i\in A} + \overline{\Delta h}_f^{i\in A}\right)}\sum_{i\in A}(h_{is} + \Delta h_{if})\left(R_{if}^{i\in A} - R_{is}\right) \qquad (42)
$$

where $N^{i\in A}$ is the number of those alive in both periods, $\bar{h}_s^{i\in A}$ is their average initial health and $\overline{\Delta h}_f^{i\in A}$ their average health change; and $R_{if}^{i\in A}$ is defined only over those alive in the final period whereas R_{is} is defined over the entire population in the initial period. Equation (42) is of the same form as (27) and can therefore in principle be similarly used to construct a

---

[7] In the empirical application, h_s enters the probit equation (34) as $\ln h_s$, and ω_0^{MB}, ω_0^{MT}, ω_h^{MB} and ω_h^{MT} are given by:

$$\omega_0^{MB} = \Phi(\bar{z}_s)\lambda\beta_0 - \left(\delta_y\overline{\Delta y}_f + \sum_{k=1}^{K}\delta_k\overline{\Delta x}_{kf} + \lambda\left(\overline{h_s^* - h_s}\right)\right)\phi(\bar{z}_s)\left(\gamma_y\bar{y}_s + \sum_{k=1}^{K}\gamma_k\bar{x}_{ks} + \gamma_h\right);$$

$$\omega_0^{MT} = -\bar{h}_s\,\phi(\bar{z}_s)\left(\gamma_y\bar{y}_s + \sum_{k=1}^{K}\gamma_k\bar{x}_{ks} + \gamma_h\right);$$

$$\omega_h^{MB} = -\Phi(\bar{z}_s)\lambda + \left(\delta_k\overline{\Delta y}_f + \sum_{k=1}^{K}\delta_k\overline{\Delta x}_{kf} + \lambda\left(\overline{h_s^* - h_s}\right)\right)\phi(\bar{z}_s)\gamma_h/\bar{h}_s;$$

$$\omega_h^{MT} = -\Phi(-\bar{z}_s) + \phi(\bar{z}_s)\gamma_h;$$

corresponding set of standardised HRIM indices. However, survival outcomes are assumed to be known with certainty in (42) whereas the nature of survival in the TPM is probabilistic, so in practice we employ the following generalisation of (42) in the standardisation procedure:

$$\tilde{M}^R = \frac{2}{\sum_i \text{Prob}\left(S_{if}=1\right)(h_{is}+\Delta h_{if})} \sum_i \text{Prob}\left(S_{if}=1\right)(h_{is}+\Delta h_{if})\left(\tilde{R}_{if}-R_{is}\right), \tag{43}$$

where $\tilde{R}_{if}$ is defined over the weighted population in the initial period, with weights equal to individuals' probability of survival until the final period $\text{Prob}\left(S_{if}=1\right)$,[8] and the final period incomes of those who did in fact die is imputed on the basis of their own initial income and the observed average income change so as to be consistent with the construction of the standardised health change measures. We then obtain the various standardised unconditional HRIM indices by substitution of the standardised survival probabilities $\text{Prob}\left(S_f^{PTl}=1\right)$, $\text{Prob}\left(S_f^{A(L)}=1\right)$, $\text{Prob}\left(S_f^{A*(L)}=1\right)$ and $\text{Prob}\left(S_f^{A(L,D)}=1\right)$ for $\text{Prob}\left(S_f=1\right)$, and of the standardised unconditional health changes $\widetilde{\Delta h}_f^{PTL}$, $\widetilde{\Delta h}_f^{A(L)}$, $\widetilde{\Delta h}_f^{A*(L)}$ and $\widetilde{\Delta h}_f^{A(L,D)}$ for Δh_f.

4. Empirical analysis

We employ the various standardisation and decomposition procedures developed in the preceding sections to investigate the dynamics of income and health in Great Britain and to evaluate the performance of Scotland in tackling income-related health inequalities relative to England & Wales over the five year period 1999 to 2004. Our empirical analysis employs data from the British Household Panel Survey (BHPS; University of Essex, Institute for Social and Economic Research, 2007)), which is a longitudinal survey of private households in Great Britain, based on an original, nationally representative sample of 5,500 households and 10,300 individuals in 1991. Specifically, we use data from waves 9 to 14 to construct a balanced panel consisting of observations on the sub-set of individuals in the BHPS for whom full information on health, income and a range of other socioeconomic variables was available in both 1999 and 2004 or for whom full information was available in 1999 and the individual was known to have died by 2004. The resultant sample comprises

[8] It is easy to check that (43) reduces to (42) if $\text{Prob}\left(S_{if}=1\right)$ is set equal to one for those that do in fact survive to the final period and to zero for those who die beforehand.

observations on 2136 individuals in Scotland, of whom 152 had died by 2004, and 7734 individuals in England & Wales, of whom 460 did not survive until 2004. Sample weights were used throughout the analysis with these being given by a set of adjusted BHPS cross-sectional weights for 1999, where the adjustments were made using inverse probability weights (see Wooldridge, 2002) to allow both for missing data in either 1999 or 2004 and for non-mortality related sample attrition between 1999 and 2004 (see PAG for further discussion). On this basis, 6.4% of the raised sample in Scotland and 6.2% in England & Wales did not survive until 2004.

Turning first to the dynamic health function given by (13), the dependent variable was specified as the change in health utility between 1999 and 2004, where our measure of health is expressed in terms of Quality Adjusted Life Years (QALYs) and derived from the responses to the SF-36 questionnaire using the SF-6D preference based algorithm (Brazier et al. 2002). This measure is bounded in the unit interval with full health corresponding to a value of one, the lowest possible health utility of anyone alive being equal to 0.319, and with death assigned a score of zero. Table 1 shows that the average QALY score fell among those who survived until 2004, both in Scotland and in England & Wales, with this morbidity-related decline being reinforced by health utility losses due to mortality. The income variable for each individual was defined as the natural logarithm of annual household income equivalised using the McClements scale (Taylor, 1995) to take account of household composition and deflated by the CPI to take account of inflation. Individuals with negative incomes were excluded from the sample, while zero incomes were set to an arbitrarily small positive value. The other determinants of equilibrium health included in the model were the number of cigarettes usually smoked each day, age, the square of age, gender, ethnicity and highest level of educational attainment. Changes in log income and smoking were also included in the specification to capture the possible impact effects of changes in these variables on health, with smoking treated as the sole policy-relevant, non-income health change determinant in the subsequent standardisation analysis. Of those alive in both years, average incomes rose by 11.7% from an average of £22451 in Scotland and by 11.2% from £23,390 in England & Wales, while smoking fell by 0.5 cigarettes per day from an average of 4.8 per day in Scotland and by 0.4 from 3.3 per day in England & Wales. Both average incomes and cigarette consumption were higher in 1999 among those who did survive until 2004 compared to those who did not.

Table 1. Variable definitions and panel attributes: (A) Panel without non-survivors (B) Panel including non-survivors

		SCOTLAND				ENGLAND & WALES			
Variable	Attribute	Mean	Std.Dev	Min	Max	Mean	Std.Dev	Min	Max
(A) Panel without non-survivors									
CHHEALTH	Change in health	-0.005	0.126	-0.605	0.492	-0.010	0.117	-0.699	0.495
HEALTH99	Health 1999	0.810	0.131	0.301	1	0.809	0.119	0.301	1
LNHEALTH99	Logarithm of health 1999	-0.227	0.186	-1.201	0	-0.225	0.167	-1.201	0
LNINCOME99	Logarithm of income 1999	2.796	1.122	-6.908	6.212	2.913	0.846	-7.419	6.341
CHLNINCOME	Change in log income	0.224	1.244	-9.914	10.471	0.146	0.925	-11.657	11.551
AGE99	Age 1999	44.615	16.999	16	93	46.109	17.747	16	93
AGESQ99	Age squared 1999	2279.202	1634.934	256	8649	2440.958	1739.298	256	8649
MALE	Gender (Male = 1)	0.482	0.500	0	1	0.476	0.499	0	1
NONWHITE	Race (Non White = 1)	0.044	0.204	0	1	0.183	0.387	0	1
ADVEDUC	At least Highers/A Levels 1999	0.467	0.499	0	1	0.364	0.481	0	1
STDEDONLY	Standards/CSEs only 1999	0.239	0.427	0	1	0.329	0.470	0	1
SMOKING99	Cigarettes smoked per day 1999	4.845	8.794	0	50	3.374	7.192	0	60
CHSMOKING	Change in smoking	-0.534	5.016	-40	30	-0.427	4.705	-40	30
(B) Panel including non-survivors									
CHHEALTH	Change in health	-0.047	0.207	-1	0.492	-0.051	0.200	-1	0.495
HEALTH99	Health 1999	0.800	0.138	0.301	1	0.801	0.127	0.301	1
LNHEALTH99	Logarithm of health 1999	-0.241	0.200	-1.201	0	-0.237	0.182	-1.201	0
LNINCOME99	Logarithm of income 1999	2.763	1.171	-6.908	6.212	2.888	0.843	-7.419	6.341
CHLNINCOME	Change in log income*	0.224	1.203	-9.914	10.471	0.146	0.896	-11.657	11.551
AGE99	Age 1999	46.441	18.129	16	93	47.854	18.744	16	96
AGESQ99	Age squared 1999	2485.197	1813.232	256	8649	2641.351	1902.892	256	9216
MALE	Gender (Male = 1)	0.481	0.500	0	1	0.476	0.499	0	1
NONWHITE	Race (Non White = 1)	0.043	0.203	0	1	0.174	0.379	0	1
ADVEDUC	At least Highers/A Levels 1999	0.449	0.498	0	1	0.349	0.477	0	1
STDEDONLY	Standards/CSEs only 1999	0.230	0.421	0	1	0.317	0.465	0	1
SMOKING99	Cigarettes smoked per day 1999	4.811	8.782	0	50	3.343	7.160	0	60
CHSMOKING	Change in smoking*	-0.534	4.853	-40	30	-0.427	4.556	-40	30
SURVIVAL	Survival status (Alive in 2004=1)	0.936	0.245	0	1	0.938	0.241	0	1

* Survivors only.

Table 2. Dynamic health functions conditional upon survival

CHHEALTH	SCOTLAND			ENGLAND & WALES		
	Coef.	Robust Std.Error	t	Coef.	Robust Std.Error	t
CHLNINCOME	0.004825	0.002745	1.76	0.009060	0.002509	3.61
CHSMOKING	-0.000317	0.000600	-0.53	-0.000203	0.000370	-0.55
LNINCOME99	0.010526	0.004541	2.32	0.014640	0.002504	5.85
SMOKING99	-0.001318	0.000392	-3.36	-0.000890	0.000226	-3.94
AGE99	0.002023	0.000948	2.13	0.000447	0.000429	1.04
AGESQ99	-0.000037	0.000010	-3.65	-0.000020	0.000004	-4.41
MALE	0.029019	0.005044	5.75	0.013474	0.002993	4.50
NONWHITE	-0.013798	0.013049	-1.06	-0.004129	0.003668	-1.13
ADVEDUC	0.007582	0.007160	1.06	0.011807	0.004611	2.56
STDEDONLY	0.010622	0.008256	1.29	0.014674	0.004534	3.24
HEALTH99	-0.530303	0.030895	-17.16	-0.509437	0.015108	-33.72
constant	0.374081	0.035900	10.42	0.373860	0.017481	21.39
№ obs.	1984			7274		
R-squared	0.3053			0.2599		
F	46.67			116.38		
Prob > F	0.0000			0.0000		
RMSE	0.10551			0.10095		

Robust standard errors allow for the sample design.

Table 2 reports the results from the estimation of the dynamic health functions conditional upon survival, with the dependent variable given by the change in health as in (13) and (35a). The first two coefficients show the short-run impact of changes in income and smoking on health. Thus increases in the logarithm of income led to contemporaneous improvements in health, consistent with other evidence that short run movements in individual health are related to transitory income shocks (see e.g. Benzeval and Judge, 2001). Conversely, increases in smoking are estimated to have an immediate negative impact on health, though this effect was not significantly different from zero in either set of results. The remainder of the function then serves to define the equilibrium error with the estimates of the coefficient on the initial health variable implying that just over half of any gap between individuals' actual and equilibrium health in 1999 was closed by 2004. Dividing through the coefficients on the other initial health determinants by this adjustment coefficient yields the parameters of the equilibrium health relationship (14). Thus the long-run effect of changes in the logarithm of income was between three and four times as large as the initial impact effect, while the long-term impact of smoking was between eight and ten times the short-term impact. The quadratic in age implies that equilibrium health levels peak before 30, with health declining at an increasing rate thereafter. Men had significantly higher equilibrium levels of health utility than women all other things equal. Non Whites had lower equilibrium levels of health utility though neither of the coefficient estimates was significantly different from zero. Finally, higher levels of educational attainment are associated with better long-term health than the omitted case of no educational qualifications. Overall, the similarity of the two sets of regression results presented in Table 2 provides grounds for confidence in the empirical robustness of the model specification, with expected signs on all coefficients significantly different from zero.

Table 3 presents the results of an analysis of the equilibrium levels of income-related health inequality implied by the ECM estimates reported in Table 2 for the sub-set of the population in 1999 who were still alive in 2004. Bootstrap estimates of standard errors are reported for all measures, where the bootstrapping procedure reflect the sample design with re-sampling at the cluster (Primary Sampling Unit) rather than the individual level. In the case of the equilibrium results, these estimates reflect not only the inherent sampling variability of the measures but also the precision of the dynamic health function estimates employed in their calculation.[9] A comparison of the values of the estimated equilibrium and observed concentration indices for 1999 reveals that levels of structural or chronic inequality conditional on survival were greater than would be inferred from the cross-sectional measures in both Scotland and England & Wales. This finding that income-related health inequality was worse in the long-run than in the short-run is consistent with the negative estimates of the Jones and Lopez Nicholas (2004) "health-related income mobility" index reported in the literature, which AGP argue result from a stronger association between permanent income and health than between short run changes in income and health. The decomposition of the equilibrium concentration indices further reveals that roughly one half of chronic income-related health inequality in 1999 was attributable to income inequality *per se*. A further third was attributable to income-related inequalities in age, with the old more likely to be both poorer and in worse health, all other things equal. Finally, smoking, education and gender also contributed significantly to overall levels of equilibrium income-related health inequality, with smokers more likely to be poor and in bad health while well-qualified individuals and men were more likely to be both better off and have better underlying levels of health.

[9] The bootstrapping procedure does not include re-estimation of the individual weights which are constructed from the original sample.

Table 3. Decomposition of equilibrium income-related health inequality in 1999 by heath determinant

		SCOTLAND				ENGLAND & WALES			
		Conc. Index	Contribution	Total		Conc. Index	Contribution	Total	
		CI	η	Value	Share	*CI*	η	Value	Share
Income-related health inequality in 1999 (panel without non-survivors)		-	-	0.01669	-	-	-	0.01674	-
				0.00249				*0.00122*	
Equilibrium income-related health inequality (conditional on survival)		-	-	0.02303	100.0%	-	-	0.02684	100.0%
				0.00527				*0.00222*	
of which due to:	LNINCOME99	0.16133	0.06955	0.01122	48.7%	0.13304	0.10642	0.01416	52.7%
		0.01057	*0.03119*	*0.00522*		*0.00361*	*0.01775*	*0.00232*	
	SMOKING99	-0.12104	-0.01509	0.00183	7.9%	-0.08293	-0.00749	0.00062	2.3%
		0.02896	*0.00413*	*0.00059*		*0.02126*	*0.00198*	*0.00023*	
	AGE99	-0.02483	0.21328	-0.00530	-23.0%	-0.03969	0.05145	-0.00204	-7.6%
		0.00755	*0.10390*	*0.00279*		*0.00382*	*0.05068*	*0.00203*	
	AGESQ99	-0.06479	-0.19775	0.01281	55.6%	-0.09063	-0.12008	0.01088	40.5%
		0.01384	*0.05706*	*0.00411*		*0.00702*	*0.02842*	*0.00285*	
	MALE	0.03785	0.03303	0.00125	5.4%	0.05340	0.01599	0.00085	3.2%
		0.01172	*0.00626*	*0.00046*		*0.00619*	*0.00335*	*0.00021*	
	NONWHITE	-0.05918	-0.00142	0.00008	0.4%	0.03425	-0.00188	-0.00006	-0.2%
		0.10795	*0.00145*	*0.00024*		*0.01985*	*0.00179*	*0.00007*	
	ADVEDUC	0.16259	0.00837	0.00136	5.9%	0.22439	0.01074	0.00241	9.0%
		0.01689	*0.00739*	*0.00119*		*0.01218*	*0.00387*	*0.00088*	
	STDEDONLY	-0.03792	0.00600	-0.00023	-1.0%	0.00207	0.01206	0.00003	0.1%
		0.02718	*0.00420*	*0.00024*		*0.01061*	*0.00333*	*0.00013*	
	constant	0	0.88402	0	0%	0%	0.93281	0	0%
		-	*0.05592*	-	-	-	*0.03071*	-	-

** Bootstrap errors based on 200 replications in italics*

Table 4 presents the results from the AGP decomposition of the change in income-related health inequality for the panel without non-survivors, where bootstrap estimates of standard errors are again reported for all measures. The results in column (1) show that average health declined over the period in both Scotland and in England & Wales, which is to be expected given the balanced nature of the panels, but that income-related health inequality, as measured by the concentration index, increased in both countries. The decomposition of these increases in health inequality reveals three common points of interest. First, the negative values of the index of income-related health mobility M_H imply that health depreciation in the two panels had the effect of increasing health inequalities, though this effect was not significant in Scotland, with the positive values of the disproportionality index P indicating that relative health losses were concentrated among the worse-off. Second, the positive values of the health-related income mobility index M_R imply that income-related health inequalities were exacerbated by income re-ranking: by implication, those who moved up the income ranking tended to be healthier in 2004 than those who moved down the income ranking. Third, the disequalising effects of health changes and income re-ranking reinforced each other to determine the overall increases in income-related health inequality between 1999 and 2004. Overall, the AGP decomposition results for Scotland and for England & Wales are broadly similar, with none of the differences between the results for the two countries proving to be significant at conventional levels on the basis of the test procedure set out in Zandvakili (2008).

The remaining columns in Table 4 present the results from the various standardisations considered in Section 2, where it may be noted that the average standardised health changes are equal by construction to the observed average changes reported in column (1) with the exception of the predicted equilibrium health changes in (4). Bootstrap estimates again reflect not only the inherent sampling variability of the measures but also the precision of the dynamic health function estimates employed in their calculation. Turning first to the results in column (2) then for both Scotland and England & Wales these show a decrease, rather than an increase, in income-related health inequality over the period after standardising health changes to eliminate the effects of

Table 4. AGP decomposition of unstandardised and standardised changes in income-related health inequality (survivors only).

	Unstandardised results			Standardised results							
	(1)			(2)		(3)		(4)		(5)	
Health change measure:	Δh_f			Δh_f^{PTL}		$\Delta h_f^{A(L)}$		$\Delta h_f^{A*(L)}$		$\Delta h_f^{A(L,D)}$	
	SCO	E&W	*p-value*	SCO	E&W	SCO	E&W	SCO	E&W	SCO	E&W
Average health 1999	0.80973	0.80907	0.465	0.80973	0.80907	0.80973	0.80907	0.80973	0.80907	0.80973	0.80907
$\bar{h}_s$	*0.00437*	*0.00191*		*0.00437*	*0.00191*	*0.00437*	*0.00191*	*0.00437*	*0.00191*	*0.00437*	*0.00191*
Average health change	–0.00499	–0.00997	0.115	–0.00499	–0.00997	–0.00499	–0.00997	–0.01177	–0.02234	–0.00499	–0.00997
$\overline{\Delta h}_f$	*0.00345*	*0.00183*		*0.00345*	*0.00183*	*0.00345*	*0.00183*	*0.00701*	*0.00392*	*0.00345*	*0.00183*
Conc. Index of health changes	–0.03964	–0.33420	0.240	0.71064	0.00237	0.55595	–0.02260	–0.35960	0.03197	0.55663	–0.02366
$CI_{f-s,s}$	1.70523	0.10839		6.61482	0.12211	5.33223	0.12307	53.54738	0.10012	5.32138	0.12314
Concentration Index 1999	0.01669	0.01674	0.460	0.01669	0.01674	0.01669	0.01674	0.01669	0.01674	0.01669	0.01674
CI_{ss}	*0.00249*	*0.00122*		*0.00249*	*0.00122*	*0.00249*	*0.00122*	*0.00249*	*0.00122*	*0.00249*	*0.00122*
Concentration Index 2004	0.02337	0.02529	0.235	0.01378	0.01480	0.01486	0.01504	0.00999	0.01001	0.01480	0.01506
CI_{ff}	*0.00251*	*0.00128*		*0.00229*	*0.00102*	*0.00232*	*0.00103*	*0.00284*	*0.00130*	*0.00232*	*0.00103*
Change in inequality	0.00668	0.00855	0.220	–0.00291	–0.00194	–0.00183	–0.00170	–0.00670	–0.00673	–0.00189	–0.00168
$CI_{ff}-CI_{ss}=M_R-M_H$	*0.00257*	*0.00115*		*0.00269*	*0.00109*	*0.00270*	*0.00109*	*0.00366*	*0.00158*	*0.00269*	*0.00109*
Income-related health mobility	–0.00035	–0.00438	0.055	0.00431	–0.00018	0.00335	–0.00049	0.00506	0.00043	0.00335	–0.00050
M_H	*0.00260*	*0.00108*		*0.00330*	*0.00137*	*0.00329*	*0.00137*	*0.00576*	*0.00255*	*0.00329*	*0.00137*
- *Disproportionality Index*	0.05632	0.35094	0.245	–0.69395	0.01437	–0.53926	0.03934	–0.34291	–0.01523	–0.53994	0.04039
P	*1.70489*	*0.10799*		*6.61477*	*0.12177*	*5.33214*	*0.12273*	*53.5473*	*0.09976*	*5.32129*	*0.12280*
- *Scale factor*	–0.00621	–0.01248	0.105	–0.00621	–0.01248	–0.00621	–0.01248	–0.01476	–0.02840	–0.00621	–0.01248
q	*0.00430*	*0.00231*		*0.00430*	*0.00231*	*0.00430*	*0.00231*	*0.00889*	*0.00512*	*0.00430*	*0.00231*
Health-related income mobility	0.00633	0.00417	0.210	0.00140	–0.00212	0.00152	–0.00219	–0.00164	–0.00630	0.00146	–0.00219
M_R	*0.00243*	*0.00104*		*0.00202*	*0.00107*	*0.00200*	*0.00108*	*0.00293*	*0.00147*	*0.00202*	*0.00108*

Bootstrapped standard errors in italics based on 200 replications.

p-value is the probability value from a one-sided test of the equality of the measures for Scotland and England & Wales.

inequalities in all health change determinants other than income and health in 1999.[1] This decrease is driven in Scotland by the positive value of the index of income-related health mobility M_H^{PTL} due to the pattern of relative standardised health losses favouring the poor not the rich, and in England & Wales by the negative value of the health-related income mobility index M_R^{PTL} due to the effect of re-ranking being equalising not disequalising when calculated with the directly standardised final health measure. Comparing the results in column (2) with those in (3) reveals that the concentration of smoking among the poor in 1999 had an adverse impact on the subsequent evolution of income-related health inequalities due to the negative health consequences. However the redistributive effects of income-related inequalities in smoking were small and full adjustment of health levels to individuals' income and smoking behaviour in 1999 would result in the even lower equilibrium levels of income-related health inequality reported in (4). Finally, a comparison of the results in columns (3) and (5) suggests that the short-run impact of changes in income and smoking behaviour on income-related health inequalities in 2004 was negligible with a slight reduction in Scotland due to a reduction in the adverse consequences of re-ranking and a slight increase in England & Wales despite a marginal increase in the pro-poor bias of relative health losses.

Table 5 extends the AGP decomposition analysis by reporting the results from the further decomposition of the income-related health mobility index by health change determinant, where the standard errors are obtained from the bootstrapping procedure as before. The results reveal that the net effect of the adjustment of health towards the equilibrium levels implied by the set of initial conditions was to moderately exacerbate income-related health inequalities, as reflected by the contribution of the equilibrium error term to income-related health mobility, with the disequalising effects of health changes due to income, smoking, age, gender and education in 1999 almost entirely offset by the equalising effect of the positive constant term (a uniform improvement in health will reduce relative health inequality). In particular, the contribution of income-related inequalities in age to income-related health mobility was strongly disequalising due to the combination of a positive association between age and poverty (which gives rise to the positive P values for *AGE99* and *AGESQ99* since the corresponding

[1] Note that $P^{PTL} = P_y + \left(1-\left(q_y + q_h\right)\right) CI_{ss} - GC_{fs}^{\hat{\varepsilon}} \big/ \overline{\Delta h_f}$, not P_y, since both uniform health changes (due to the combined effect of the standardising non-income health determinants) and the regression residual also have an impact on relative income-related health inequalities. See Table 5 for an estimate of P_y.

Table 5. Decomposition of AGP index of income-related health mobility by heath change determinant

			SCOTLAND				ENGLAND & WALES			
			Disproportionality *P*	Scale *q*	Mobility M_H	Share	Disproportionality *P*	Scale *q*	Mobility M_H	Share
Income-related health mobility			0.05632	– 0.00621	– 0.00035	100.00%	0.35094	– 0.01248	– 0.00438	100.0%
			1.70489	*0.00430*	*0.00260*		*0.10799*	*0.00231*	*0.00108*	
of which due to:	CHLNINCOME		1.17212	0.00134	0.00157	-449.8%	1.33328	0.00165	0.00220	-50.3%
			0.19785	*0.00113*	*0.00115*		*0.12859*	*0.00049*	*0.00057*	
	CHSMOKING		0.03663	0.00021	0.00001	-2.2%	-0.10460	0.00011	-0.00001	0.3%
			0.14309	*0.00043*	*0.00006*		*0.09498*	*0.00019*	*0.00003*	
	Equilibrium error		0.42999	-0.00776	-0.00334	954.1%	0.35584	-0.01424	-0.00507	115.7%
			14.96583	*0.00449*	*0.00281*		*0.08886*	*0.00240*	*0.00106*	
	of which due to:	LNINCOME99	-0.14464	0.03657	-0.00529	1512.8%	-0.11630	0.05337	-0.00621	141.7%
			0.01050	*0.01605*	*0.00241*		*0.00375*	*0.00905*	*0.00103*	
		SMOKING99	0.13773	-0.00793	-0.00109	312.4%	0.09967	-0.00376	-0.00037	8.6%
			0.02925	*0.00228*	*0.00035*		*0.02134*	*0.00100*	*0.00013*	
		AGE99	0.04152	0.11215	0.00466	-1331.5%	0.05643	0.02580	0.00146	-33.2%
			0.00812	*0.05444*	*0.00229*		*0.00427*	*0.02534*	*0.00143*	
		AGESQ99	0.08148	-0.10398	-0.00847	2423.0%	0.10737	-0.06023	-0.00647	147.6%
			0.01418	*0.02942*	*0.00250*		*0.00741*	*0.01409*	*0.00165*	
		MALE	-0.02116	0.01737	-0.00037	105.1%	-0.03666	0.00802	-0.00029	6.7%
			0.01173	*0.00319*	*0.00022*		*0.00612*	*0.00172*	*0.00008*	
		NONWHITE	0.07587	-0.00075	-0.00006	16.2%	-0.01751	-0.00095	0.00002	-0.4%
			0.10766	*0.00075*	*0.00012*		*0.01994*	*0.00090*	*0.00003*	
		ADVEDUC	-0.14590	0.00440	-0.00064	183.7%	-0.20765	0.00538	-0.00112	25.5%
			0.01648	*0.00396*	*0.00056*		*0.01199*	*0.00197*	*0.00041*	
		STDEDONLY	0.05461	0.00316	0.00017	-49.3%	0.01467	0.00605	0.00009	-2.0%
			0.02735	*0.00224*	*0.00015*		*0.01054*	*0.00169*	*0.00007*	
		HEALTH99	0	-0.53359	0	-	0	-0.51580	0	-
			-	*0.03126*	-		-	*0.01543*	-	
		constant	0.01669	0.46485	0.00776	-2218.4%	0.*01674*	0.46785	0.00783	-178.8%
			0.00249	*0.04513*	*0.00132*		*0.00122*	*0.02091*	*0.00061*	
	Residual		-	-	0.00141	-402.1%	0.00120	-0.01424	-0.00150	34.3%
			-	-	*0.00128*		*0.00040*	*0.00240*	*0.00050*	

** Bootstrap errors based on 200 replications in italic* concentration indices are both negative) and a negative long-run relationship between age and health over most of the population (as indicated by the negative q values). Income inequalities *per se* also played an important disequalising role in income-related health mobility, contrary to the impression given by the standardisation results, due to the conjunction of a greater concentration of (the logarithm of) income than of health among the rich in 1999 (which gives rise to the negative P values) and the positive long-run contribution of income to health (as indicated by the positive q values). In contrast, the contribution of income growth to income-related health mobility was equalising because the poor received a large share of relative income growth between 1999 and 2004 than of health utility in 1999. Indeed, in both Scotland and England & Wales, the rise in income-related health inequality due to morbidity changes would have been substantially larger had it not been for the positive redistributive impact of income growth on health. Reductions in smoking over the period reduced health inequalities in Scotland but not in England and Wales, though the effects were very small. Finally, the contribution of the residual shows that the effect of idiosyncratic health shocks was equalising in Scotland but disequalising in England & Wales, with the difference sufficiently large to account for most of the observed disparity in income-related health mobility between the two countries.

Table 6 presents the unstandardised results from the PAG decomposition of the change in income-related health inequality, where bootstrap estimates of standard errors are again reported for all measures. These results are based on unstandardised health changes for the entire population alive in 1999, whether or not they survived to 2004, and may be directly compared with the unstandardised AGP results, excluding non-survivors, in column (1) of Table 4. Thus health was both lower on average and slightly more concentrated among the rich in the full population alive in 1999 compared to the sub-population who survived until 2004, consistent with the estimates of the effects of initial health and income on survival from the probit model (see Table 7). Average health losses including deaths were much larger and, given that the concentration of health changes due to mortality was greater than that due to morbidity, so was the concentration of those losses among the poor with this difference particularly marked in Scotland. Taken together, these two factors resulted in the more disequalising effects of health changes as reflected in the larger negative values of the income-related health mobility index $\tilde{M}_H$. Nevertheless, the increases in cross-sectional income-related health

Table 6. PAG decomposition of income-related health inequality changes (all alive in 1999)

		Unstandardised results		
Health change measure:		Δh_f		
		SCO	E&W	*p-value*
Average health 1999	$\bar{h}_s$	0.80042	0.80055	0.450
		0.00489	0.00202	
Average health change	$\overline{\Delta h_f}$	–0.04727	–0.05112	0.260
		0.00535	0.00322	
Conc. Index of health changes	$CI_{f-s,s}$	–0.28312	–0.35789	0.125
		0.06278	0.03644	
Concentration Index 1999	CI_{ss}	0.01959	0.02067	0.400
		0.00267	0.00134	
Concentration Index 2004	$CI_{ff}^{i\in A}$	0.02337	0.02529	0.210
		0.00250	0.00130	
Change in inequality $CI_{ff} - CI_{ss} = \tilde{M}_R - \tilde{M}_H$		0.00378	0.00462	0.315
		0.00274	0.00136	
Income-related health mobility	$\tilde{M}_H$	–0.01900	–0.02582	0.065
		0.00391	0.00254	
of which due to:	*morbidity*	-0.00024	-0.00410	0.050
		0.00260	*0.00109*	
	mortality	-0.01876	-0.02172	0.265
		0.00371	*0.00213*	
- *Disproportionality Index*	$\tilde{P}$	0.30272	0.37856	0.115
		0.06316	0.03620	
of which due to:	*morbidity*	0.03879	0.32859	0.215
		1.73558	*0.12641*	
	mortality	0.33168	0.38975	0.170
		0.05041	*0.03030*	
- *Scale factor*	$\tilde{q}$	–0.06277	–0.06821	0.260
		0.00758	0.00459	
of which due to:	*morbidity*	-0.00621	-0.01248	0.105
		0.00430	*0.00237*	
	mortality	-0.05656	-0.05573	0.470
		0.00665	*0.00333*	
Health-related income mobility	$\tilde{M}_R$	–0.01522	–0.02120	0.125
		0.00504	0.00258	

Bootstrapped standard errors in italics based on 200 replications.

p-value is the probability value from a one-sided test of the equality of the measures for Scotland and England & Wales.

inequality reported in Table 6 are less than those in Table 4 as the concentration index in 2004 is defined over the extant population only, with the balancing item given by the health-related income index $\tilde{M}_R$, the value of which is dominated by the effect of the dead dropping out of the index (see PAG for further discussion). Overall, the unstandardised PAG decomposition results for Scotland and for England & Wales are again broadly similar, with only the difference between the morbidity-related IRHM indices marginally significant at conventional levels.

Table 7 reports the results from the estimation of the probit survival model (34) over the full sample of individuals alive in 1999, with the dependent variable given by survival status. The model specification includes the same set of health determinants as the equilibrium health function (14), together with the logarithm of health in 1999 to control for individuals' initial state of health. Higher levels of income and education improved survival chances, though the effects were not generally significant, while smoking had the opposite effect. The quadratic in age implies that individuals in their early twenties had the highest probability of survival over the five year period, with survival chances declining at an increasing rate thereafter. Men were less likely to survive than women all other things equal, while the effects of ethnicity were mixed with only the Scottish coefficient significantly

Table 7. Probit model of survival

	SCOTLAND			ENGLAND & WALES		
SURVIVAL	Coef.	Robust Std.Error	t	Coef.	Robust Std.Error	t
LNINCOME99	0.00695	0.03069	0.23	0.01747	0.04325	0.40
SMOKING99	-0.00384	0.00642	-0.60	-0.01621	0.00441	-3.68
AGE99	0.02490	0.02328	1.07	0.03263	0.01341	2.43
AGESQ99	-0.00067	0.00020	-3.36	-0.00067	0.00011	-5.94
MALE	-0.23357	0.12504	-1.87	-0.34078	0.06554	-5.20
NONWHITE	-0.66800	0.24143	-2.77	0.15690	0.15560	1.01
ADVEDUC	0.07912	0.13541	0.58	0.24595	0.10569	2.33
STDEDONLY	0.21853	0.17784	1.23	0.13375	0.09289	1.44
LNHEALTH99	1.57014	0.22220	7.07	1.38384	0.16771	8.25
constant	3.12219	0.67869	4.60	2.66592	0.43543	6.12
№ obs.	2136			7734		
PsuedoR^2	0.4054			0.3689		
Wald χ^2	307.83			534.23		
Prob > χ^2	0.0000			0.0000		

Robust standard errors allow for the sample design.

different from zero. Finally, the survival chances of those who were in better health in 1999 were, as would be expected, significantly better than those who were less well.

Table 8 repeats the basic output from the unstandardised PAG decomposition analysis together with the corresponding results from the various standardisations considered in Section 3, where the bootstrap estimates of the standard errors additionally reflect the precision of the probit survival model in the case of the standardised results. It should be noted that none of the average standardised health changes in columns (2) to (5) equal the observed mean health changes reported in (1) due to the non-linearity of the probit survival model. Specifically, the standardised health losses were consistently less on average, often by a substantial amount, than the observed changes since the survival probabilities based on mean values of the standardising variables were higher on average than the mean survival probabilities based on observed values.

Nevertheless, taking deaths into account affects the detail rather than the broad picture emerging from the standardisation analysis. Thus the results in column (2) again show that income-related health inequality in both Scotland and England & Wales decreased, rather than increased, after standardising health changes to eliminate the effects of inequalities in all health change determinants other than income and health in 1999. However, with deaths taken into account, these results were driven entirely by the less inequitable distribution of relative standardised health losses due to the higher standardised survival probability, which resulted in the lower scale and pro-rich bias of the standardised health changes, with the consequences of income re-ranking more adverse when calculated with the standardised final health measure. And the results in column (3) once more reveal that the concentration of smoking among the poor in 1999 had an adverse impact on the subsequent evolution of income-related health inequalities due to the negative consequences for both morbidity and mortality, but that these effects were relatively small so full adjustment of health levels to individuals' income and smoking behaviour in 1999 continued to result in the even lower equilibrium levels of income-related health inequality reported in (4). Finally, the similarity of the results in columns (3) and (5) again suggest that the short-run impact of changes in income and smoking behaviour on income-related health inequalities in 2004 was negligible.

Table 9 reports the parameters from the linearization of the Two-Part Model (TPM) representation of health changes (36), which are employed in the derivation of the results reported in Table 10 from the hierarchical decomposition of the PAG income-related health mobility index. Bootstrap estimates of standard errors are again reported for all measures,

Table 8. PAG decomposition of unstandardised and standardised changes in income-related health inequality (all alive in 1999)

	Unstandardised results		Standardised results							
	(1)		(2)		(3)		(4)		(5)	
Health change measure:	Δh_f		Δh_f^{PTL}		$\Delta h_f^{A(L)}$		$\Delta h_f^{A^*(L)}$		$\Delta h_f^{A(L,D)}$	
	SCO	E&W	SCO	E&W	SCO	E&W	SCO	E&W	SCO	E&W
Average health 1999	0.80042	0.80055	0.80042	0.80055	0.80042	0.80055	0.80042	0.80055	0.80042	0.80055
$\bar{h}_s$	*0.00489*	0.00202	*0.00487*	*0.00202*	*0.00487*	*0.00202*	0.00487	0.00202	0.00487	0.00202
Average health change	–0.04727	–0.05112	–0.01870	–0.02427	–0.01880	–0.02484	–0.02631	–0.03747	–0.01880	–0.02484
$\overline{\Delta h}_f$	*0.00535*	0.00322	*0.00479*	*0.00292*	*0.00479*	*0.00291*	0.00794	0.00460	0.00479	0.00291
Conc. Index of health changes	–0.28312	–0.35789	0.16387	0.00939	0.12022	–0.00573	0.16608	0.05334	0.12037	–0.00613
$CI_{f-s,s}$	*0.06278*	0.03644	*0.21173*	*0.05497*	*0.20103*	*0.05444*	0.31738	0.06056	0.20054	0.05431
Concentration Index 1999	0.01959	0.02067	0.01959	0.02067	0.01959	0.02067	0.01959	0.02067	0.01959	0.02067
CI_{ss}	*0.00267*	0.00134	*0.00267*	*0.00134*	*0.00267*	*0.00134*	0.00267	0.00134	0.00267	0.00134
Concentration Index 2004	0.02337	0.02529	0.01668	0.01837	0.01776	0.01866	0.01265	0.01259	0.01770	0.01868
$CI_{ff}^{i \in A}$	*0.00250*	0.00130	*0.00278*	*0.00128*	*0.00280*	*0.00128*	0.00337	0.00161	0.00279	0.00127
Change in inequality	0.00378	0.00462	–0.00292	–0.00230	–0.00183	–0.00201	–0.00694	–0.00808	–0.00189	–0.00199
$CI_{ff} - CI_{ss} = \tilde{M}_R - \tilde{M}_H$	*0.00274*	0.00136	*0.00287*	*0.00120*	*0.00289*	*0.00121*	0.00395	0.00175	0.00287	0.00121
Income-related health mobility	–0.01900	–0.02582	0.00345	–0.00035	0.00242	–0.00085	0.00498	0.00160	0.00242	–0.00086
$\tilde{M}_H$	*0.00391*	0.00254	*0.00343*	*0.00163*	*0.00339*	*0.00164*	0.00593	0.00281	0.00339	0.00164
- *Disproportionality Index*	0.30272	0.37856	–0.14428	0.01128	–0.10063	0.02640	–0.14649	–0.03267	–0.10077	0.02680
$\tilde{P}$	*0.06316*	0.03620	*0.21101*	*0.05471*	*0.20029*	*0.05415*	0.31663	0.06022	0.19981	0.05402
- *Scale factor*	–0.06277	–0.06821	–0.02392	–0.03126	–0.02405	–0.03202	–0.03399	–0.04910	–0.02405	–0.03202
$\tilde{q}$	*0.00758*	0.00459	*0.00623*	*0.00387*	*0.00623*	*0.00385*	0.01053	0.00630	0.00623	0.00385
Health-related income mobility	–0.01522	–0.02120	0.00054	–0.00266	0.00059	–0.00285	–0.00196	–0.00648	0.00053	–0.00285
$\tilde{M}_R$	*0.00504*	0.00258	*0.00193*	*0.00110*	*0.00191*	*0.00111*	0.00275	0.00142	0.00193	0.00111

Bootstrapped standard errors in italics based on 200 replications.

Table 9. Linearization of Two-Part Model
Panel A – Expected morbidity changes. Panel B – Expected mortality.

		SCOTLAND			ENGLAND & WALES		
PANEL A: $E\left(\Delta h_f^{MB}\right)$		Coef.	*Bootstrap Std.Error*	*z*	Coef.	*Bootstrap Std.Error*	*z*
CHLNINCOME	ϖ_Y	0.00476	*0.00329*	*1.45*	0.00892	*0.00252*	*3.54*
CHSMOKING	$\varpi_{SMOKING}$	-0.00031	*0.00056*	*-0.55*	-0.00020	*0.00038*	*-0.53*
LNINCOME99	ω_Y	0.01038	*0.00464*	*2.24*	0.01440	*0.00253*	*5.7*
SMOKING99	$\omega_{SMOKING}$	-0.00130	*0.00037*	*-3.53*	-0.00087	*0.00022*	*-3.88*
AGE99	ω_{AGE}	0.00199	*0.00094*	*2.13*	0.00043	*0.00040*	*1.07*
AGESQ99	ω_{AGESQ}	-0.00004	*0.00001*	*-3.72*	-0.00002	*0.00000*	*-4.75*
MALE	ω_{MALE}	0.02867	*0.00513*	*5.59*	0.01339	*0.00284*	*4.71*
NONWHITE	$\omega_{NONWHITE}$	-0.01351	*0.01328*	*-1.02*	-0.00412	*0.00397*	*-1.04*
ADVEDUC	$\omega_{ADVEDUC}$	0.00747	*0.00667*	*1.12*	0.01153	*0.00441*	*2.62*
STDEDONLY	$\omega_{STEDONLY}$	0.01045	*0.00740*	*1.41*	0.01439	*0.00442*	*3.26*
HEALTH99	ω_H	-0.52353	*0.03127*	*-16.74*	-0.50208	*0.01706*	*-29.43*
constant	ω_0	0.36925	*0.03645*	*10.13*	0.36843	*0.01952*	*18.87*
PANEL B: $E\left(\Delta h_f^{MT}\right)$							
CHLNINCOME	ϖ_Y	-	-	-	-	-	-
CHSMOKING	$\varpi_{SMOKING}$	-	-	-	-	-	-
LNINCOME99	ω_Y	0.00019	*0.00096*	*0.2*	0.00055	*0.00144*	*0.38*
SMOKING99	$\omega_{SMOKING}$	-0.00011	*0.00016*	*-0.66*	-0.00051	*0.00016*	*-3.21*
AGE99	ω_{AGE}	0.00068	*0.00069*	*0.99*	0.00103	*0.00052*	*1.96*
AGESQ99	ω_{AGESQ}	-0.00002	*0.00001*	*-2.39*	-0.00002	*0.00001*	*-3.75*
MALE	ω_{MALE}	-0.00640	*0.00343*	*-1.86*	-0.01075	*0.00217*	*-4.95*
NONWHITE	$\omega_{NONWHITE}$	-0.01829	*0.00620*	*-2.95*	0.00495	*0.00491*	*1.01*
ADVEDUC	$\omega_{ADVEDUC}$	0.00217	*0.00379*	*0.57*	0.00776	*0.00322*	*2.41*
STDEDONLY	$\omega_{STEDONLY}$	0.00598	*0.00505*	*1.19*	0.00422	*0.00285*	*1.48*
HEALTH99	ω_H	0.04039	*0.00975*	*4.14*	0.03881	*0.00723*	*5.37*
constant	ω_0	-0.02777	*0.01925*	*-1.44*	-0.03714	*0.01503*	*-2.47*

Bootstrap standard errors based on 200 replications.

with these additionally reflecting the accuracy of the linear approximation in the case of the second-stage decomposition results.

The first-stage decomposition results in Table 10 reveal two main points of interest. First, health changes due to expected morbidity changes, expected mortality and health shocks were all estimated to have had the effect of increasing income-related health inequality over the period in both Scotland and England & Wales, with the disequalising effects of expected mortality having been the dominant factor. Thus expected mortality-related health changes accounted for nearly 95% of overall income-related health mobility in Scotland, and nearly 80% of that in England and Wales, as a result of both the scale of expected health losses due to death and their concentration among the poor. These findings further point to the importance of taking deaths into account in the evaluation of policies designed to tackle health inequalities. Second, the estimated disproportionality, scale and mobility effects of expected morbidity changes and expected mortality closely matched the corresponding measures of the actual effects reported in Table 7. In particular, the TPM decomposition successfully captures the observed difference in morbidity-related health mobility between Scotland and England & Wales, which was not satisfactorily accounted for in the decomposition analysis based on the ECM alone (cf. the discussion of Table 5). Moreover, the TPM decomposition estimates also mirror the empirical finding that mortality-related health mobility was less disequalising in Scotland as a result of the slightly larger scale of health losses due to death being more than offset by their less regressive distribution in the population. As a result, the TPM explains the bulk of the observed income-related health mobility, with the residual terms consequently making minor contributions in both countries.

Finally the second-stage decomposition results suggest that the principal contributor to mortality-related health mobility was income-related inequalities in age, with the old in 1999 more likely both to be poor and to die in the following five year period. Income inequality *per se* also contributed to the disequalising effects of expected deaths as the better off were less likely to die, while income-related inequalities in educational attainment similarly contributed because the highly educated were both more likely to be better off and less likely to die, and inequalities in smoking had the same effect but for the opposite reason. The mortality contributions of these factors to income-related health inequalities thus reinforced those due to morbidity changes, as previously discussed with reference to the results of the AGP decomposition analysis and confirmed by the break down of the morbidity change term in Table 10. However these second-stage decomposition results need to be treated with some caution as the linearised version of the TPM does not yield accurate predictions of the overall levels of income-related health mobility due to expected mortality and morbidity changes in either Scotland or England & Wales. We conclude that further work is required to obtain consistent estimates of the effects of inequalities in individual health change determinants on income-related health mobility once deaths are taken into account.

Table 10. Decomposition of PAG index of income-related health mobility

SCOTLAND		Disproportionality *P*	Scale *q*	Mobility *M*	Share
Income-related health mobility		**0.30272**	**-0.06277**	**-0.01900**	**100.00%**
		0.06316	*0.00758*	*0.00391*	
***of which due to:* Expected morbidity changes**		**0.04759**	**-0.00582**	**-0.00028**	**1.5%**
		3.59633	*0.00427*	*0.00243*	
of which due to:	CHLNINCOME	1.11105	0.00018	0.00020	-1.1%
		0.17114	*0.01894*	*0.02133*	
	CHSMOKING	0.03621	0.00003	0.00000	0.0%
		0.13218	*0.00517*	*0.00029*	
	LNINCOME99	-0.14779	0.00495	-0.00073	3.8%
		0.01280	*0.62248*	*0.09017*	
	SMOKING99	0.13427	-0.00108	-0.00014	0.7%
		0.02873	*0.13482*	*0.01668*	
	AGE99	0.05559	0.01595	0.00089	-4.7%
		0.00879	*1.68183*	*0.09842*	
	AGESQ99	0.10440	-0.01548	-0.00162	8.5%
		0.01459	*1.85232*	*0.19849*	
	MALE	-0.01839	0.00238	-0.00004	0.2%
		0.01058	*0.30532*	*0.00507*	
	NONWHITE	0.08423	-0.00010	-0.00001	0.1%
		0.10666	*0.01871*	*0.00273*	
	ADVEDUC	-0.15862	0.00058	-0.00009	0.5%
		0.01657	*0.07732*	*0.01128*	
	STDEDONLY	0.04634	0.00041	0.00002	-0.1%
		0.02684	*0.03947*	*0.00207*	
	HEALTH99	0	-0.07225	0	0.0%
		-	*10.10233*	-	
	constant	0.01959	0.06367	0.00125	-6.6%
		0.00267	*9.54495*	*0.19220*	
Expected mortality		**0.31803**	**-0.05662**	**-0.01801**	**94.8%**
		0.04256	*0.00648*	*0.00293*	
of which due to:	LNINCOME99	-0.14779	0.00251	-0.00037	1.9%
		0.01280	*0.01479*	*0.00214*	
	SMOKING99	0.13427	-0.00242	-0.00032	1.7%
		0.02873	*0.00412*	*0.00056*	
	AGE99	0.05559	0.15127	0.00841	-44.3%
		0.00879	*0.17151*	*0.00985*	
	AGESQ99	0.10440	-0.21688	-0.02264	119.2%
		0.01459	*0.08131*	*0.00908*	
	MALE	-0.01839	-0.01470	0.00027	-1.4%
		0.01058	*0.00839*	*0.00025*	
	NONWHITE	0.08423	-0.00375	-0.00032	1.7%
		0.10666	*0.00163*	*0.00046*	
	ADVEDUC	-0.15862	0.00464	-0.00074	3.9%
		0.01657	*0.00821*	*0.00129*	
	STDEDONLY	0.04634	0.00656	0.00030	-1.6%
		0.02684	*0.00581*	*0.00035*	
	HEALTH99	0	0.15442	0	0.0%
		-	*0.03207*	-	
	constant	0.01959	-0.13263	-0.00260	13.7%
		0.00267	*0.09440*	*0.00187*	
Residual		**2.21467**	**-0.00032**	**-0.00072**	**3.8%**
		97.51672	*0.00036*	*0.00177*	

Bootstrap errors based on 200 replications in italics

……continued

Table 10 (continued)

ENGLAND & WALES		Disproportionality *P*	Scale *q*	Mobility *M*	Share
Income-related health mobility		**0.37856**	**-0.06821**	**-0.02582**	**100.00%**
		0.03620	*0.00459*	*0.00254*	
***of which due to:* Expected morbidity changes**		**0.32399**	**-0.01216**	**-0.00394**	**15.3%**
		0.11392	*0.00220*	*0.00119*	
of which due to:	CHLNINCOME	1.25952	0.00243	0.00306	-11.9%
		0.11855	*0.00714*	*0.00952*	
	CHSMOKING	-0.09127	0.00016	-0.00001	0.0%
		0.08591	*0.00178*	*0.00033*	
	LNINCOME99	-0.11427	0.07790	-0.00890	34.5%
		0.00365	*0.46607*	*0.05161*	
	SMOKING99	0.09199	-0.00545	-0.00050	1.9%
		0.01881	*0.07577*	*0.00628*	
	AGE99	0.07250	0.03836	0.00278	-10.8%
		0.00412	*0.19613*	*0.01454*	
	AGESQ99	0.13303	-0.09477	-0.01261	48.8%
		0.00686	*1.02864*	*0.13856*	
	MALE	-0.03271	0.01195	-0.00039	1.5%
		0.00567	*0.13681*	*0.00527*	
	NONWHITE	-0.03792	-0.00134	0.00005	-0.2%
		0.01863	*0.03207*	*0.00113*	
	ADVEDUC	-0.22353	0.00753	-0.00168	6.5%
		0.01114	*0.13594*	*0.02895*	
	STDEDONLY	-0.00020	0.00855	0.00000	0.0%
		0.01130	*0.10168*	*0.00128*	
	HEALTH99	0	-0.75278	0	0.0%
		-	*9.39212*	-	
	constant	0.02067	0.69001	0.01426	-55.2%
		0.00140	*9.37285*	*0.20476*	
Expected mortality		**0.36184**	**-0.05590**	**-0.02023**	**78.3%**
		0.02515	*0.00354*	*0.00173*	
of which due to:	LNINCOME99	-0.11427	0.00598	-0.00068	2.6%
		0.00365	*0.01619*	*0.00184*	
	SMOKING99	0.09199	-0.00643	-0.00059	2.3%
		0.01881	*0.00191*	*0.00020*	
	AGE99	0.07250	0.18521	0.01343	-52.0%
		0.00412	*0.08933*	*0.00650*	
	AGESQ99	0.13303	-0.20842	-0.02773	107.4%
		0.00686	*0.04658*	*0.00630*	
	MALE	-0.03271	-0.01926	0.00063	-2.4%
		0.00567	*0.00401*	*0.00016*	
	NONWHITE	-0.03792	0.00323	-0.00012	0.5%
		0.01863	*0.00349*	*0.00015*	
	ADVEDUC	-0.22353	0.01018	-0.00227	8.8%
		0.01114	*0.00433*	*0.00097*	
	STDEDONLY	-0.00020	0.00503	0.00000	0.0%
		0.01130	*0.00353*	*0.00007*	
	HEALTH99	0	0.11685	0	0.0%
		-	*0.02202*	-	
	constant	0.02067	-0.13967	-0.00289	11.2%
		0.00140	*0.04959*	*0.00103*	
Residual		**11.06800**	**-0.00015**	**-0.00165**	**6.4%**
		699.96210	*0.00018*	*0.00074*	

** Bootstrap errors based on 200 replications in italics.*

5. Conclusion

Regression-based decomposition procedures are used both to standardise the concentration index and to determine the contribution of inequalities in the individual health determinants to the overall value of the index. The main contribution of this paper is to develop analogous procedures to decompose the income-related health mobility (IRHM) and health-related income mobility indices (HRIM) first proposed in Allanson, Gerdtham and Petrie (2010) and subsequently extended in Petrie, Allanson and Gerdtham (2010) to account for deaths. These procedures are based on the specification and estimation of a Two-Part Model consisting of a dynamic health function conditional upon survival, which allows for lagged as well as contemporaneous responses to changes in income and other health determinants, and a probit model of survival. More specifically, we employ an error correction model (ECM) of conditional health changes in order to clearly distinguish between the short-run and long-run effects of changes in health determinants on income-related health inequality. Our empirical estimates imply that income-related health inequality was greater in the long-run than in the short-run, suggesting that priority should be given to policies that tackle the structural problems that trap some individuals in deprivation and ill-health.

The various procedures developed in the paper are applied to the investigation of the dynamics of income and health in Great Britain over the five year period 1999 to 2004, using a measure of Quality Adjusted Life Years (QALYs) derived from the British Household Panel Survey. Our empirical results show that the output from the standardisation procedures is in general difficult to interpret because controlling for the effects of standardising variables in the usual way does not serve to entirely eliminate their influence on the results of the analysis since uniform health changes have an impact on (relative) income-related health inequalities (cf. Wagstaff et al., 2003). Nevertheless, the standardisation procedures may be useful to explore the implications of alternative counter-factual policy scenarios for both IRHM and HRIM, and hence for the overall evolution of income-related health inequality.

In contrast, the output from the proposed decomposition procedures is readily intelligible, identifying the separate contributions of individual health change determinants to overall IRHM as the product of the progressivity and scale of the health changes attributable to each determinant. Health changes due to expected morbidity changes, mortality and health shocks are all found to have had a disequalising effect over the period in both Scotland and England & Wales, with the overall effect dominated by mortality-related health losses. We further find that the major driver of the disequalising effects of health losses due to both expected morbidity and mortality was the positive association between (old) age and poverty given that ageing was predicted to have led to both poorer health and an increased risk of dying for the majority of the population represented by the BHPS sample in 1999.

The decomposition procedures may also be used to evaluate the short-run impact of contemporaneous changes in health determinants on IRHM and thereby facilitate the evaluation of policy interventions. We find that, although income inequality in 1999 contributed to rising income-related health inequalities over the subsequent five year period in both Scotland and England & Wales, the overall increase due to morbidity-related health changes was substantially moderated by the poor enjoying a larger share of real income gains between 1999 and 2004 than of health in 1999. This period largely coincided with New Labour's second term in office, which was characterised by income growth rates that were highest at the very bottom of the income distribution (Joyce *et al.*, 2010: 25) as a result of range of factors including low unemployment, the introduction of the minimum wage and an assortment of new and enhanced social security benefits. The pattern of income growth is shown by (Joyce *et al.*, 2010: 25) to have led to reductions in both income inequality and relative poverty over the period, with our study further demonstrating that it also helped to moderate the disequalising effects of IRHM among the population resident in 1999. More generally, the findings may be taken to indicate the potential to tackle income-related health inequalities by reducing income inequality *per se* through welfare and other policies, given that roughly half of the equilibrium level of income-related health inequality in 1999 was attributable to income inequality. In contrast, we were unable to demonstrate any significant effects of the decline in smoking over the period on the level of income-related health inequality.

Overall, the decomposition results for Scotland and for England & Wales are broadly similar, providing grounds for confidence in the empirical robustness of the procedures, with only the difference in morbidity-related IRHM indices marginally significant at conventional levels. The first-stage decomposition results from our Two-Part Model successfully captures this difference but further work is required to fully disentangle the separate contributions of the regression coefficients, the distributions of health change determinants and residuals in a consistent manner. In future research we intend to use the re-weighting procedure of Lemieux (2002) for this purpose.

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