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The effect of income growth and inequality on health inequality: Theory and empirical evidence from the European Panel[☆]

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ABSTRACT

Governments of EU countries have declared that they would like to couple income growth with reductions in social inequalities in income and health. We show that, theoretically, both aims can be reconciled only under very specific conditions concerning the type of growth and the income responsiveness of health. We investigate whether these conditions were met in Europe in the 1990s using panel data from the *European Community Household Panel*. We demonstrate that (i) in most countries, the income elasticity of health was positive and increases with income, and (ii) that income growth was not pro-rich in most EU countries, resulting in small or negligible reductions in income inequality. The combination of both findings explains the modest increases we observe in income-related health inequality in the majority of countries.

1. Introduction

A price to pay for income growth may be an increase in income-related disparities such as income-related health inequalities. Policy makers often aim to foster growth and simultaneously to reduce disparities, but are unaware of the conditions that make the combination attainable. For example, when the EU leaders met in the Lisbon European Council 2000, they claimed to strive to become "the most competitive and dynamic knowledge-based economy ... with more and better jobs and greater social cohesion"[1] (Atkinson et al., 2002). Social exclusion is broadly defined and includes wider social dimensions like housing, education and health and is monitored using a set of indicators like health status.[2]

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[1] There are 8 Lisbon strategic goals: 6 on economic performance and 2 on increasing social inclusion.

[2] Among the so-called Level 1 (Laeken) indicators – which consist of a restricted number of lead indicators covering the broad fields of social exclusion – are the ratio of equivalised income of the top and bottom quintile for income inequality (recommendation 15), and the same ratio for the proportion of the population classifying themselves in poor or very poor health (recommendation 23) in Atkinson et al. (2002).

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An important question, therefore, is to what extent – and under what conditions – the twin goals of income growth and reduction of social inequalities in health can be reconciled.[3] This paper focuses on the consequences of income growth, but – as will be shown – these cannot be analyzed independently from the effects of changing income inequalities on social inequalities in health. A second – no less important – question is how countries have fared with respect to the achievement of these goals. Building on earlier work by Contoyannis and Forster (1999) and Wagstaff et al. (2003) we develop a decomposition technique that points to the crucial role of the income elasticity of health. If this elasticity is increasing with income, then proportional income growth may – under certain conditions – lead to *higher* income-related health inequalities (hereafter denoted as IRHI). If this were the case, then Europe faces a trade-off between these two goals. If, on the other hand, growth goes hand in hand with a reduction in health inequality by income, then greater social inclusion derives as a windfall profit. It turns out that the degree to which income growth occurs disproportionately at higher or lower incomes, and the degree to which health responds to income changes at varying income levels, are both crucial elements in the relationship between income growth and inequality and the degree of income-related health inequality.

The paper then goes on to analyze the empirically observed trends in income (inequality) and health (inequality) in European countries using one of the monitoring tools the European Union created: a set of comparable longitudinal household level data across all member states: the *European Community Household Panel* (ECHP) survey. We estimate regression models of health and use our decomposition technique to relate trends in income growth and income inequality to changes in income-related inequalities in health as observed in 8 waves of data. There is also an epidemiological literature documenting trends in socioeconomic inequalities in self-assessed health (e.g. Dalstra et al., 2005; Kunst et al., 2005) but, we first theoretically identify the role of changes in the level and distribution of incomes on IRHI before empirically testing these relationships for a large set of European countries.

The paper is organised as follows. In Section 2 we develop a decomposition technique to analyze the consequences of income growth and income inequality for social inequalities in health. Section 3 describes the ECHP data and the empirical models used to implement the decomposition. Empirical results on income elasticities of health and on trends in income, health, inequality and their decomposition are presented in Section 4, while Section 5 provides a conclusion and discussion.

2. Decomposing the relation between the distributions of income and health

In this section, we present a decomposition technique to examine the impact of changes in the income distribution on the distribution of health. We focus on two aspects of the distribution of health, i.e. the evolution of mean health and of IRHI. Despite the analysis being far more complicated for income-related than for pure health inequalities, we did not consider the latter since policy makers are generally more concerned with the socioeconomic gradient in health. We extend the approach proposed by Wagstaff et al. (2003) by allowing for a non-linear relationship between income and health.

2.1. Decomposing IRHI in a linear framework

Wagstaff et al. (2003) have proposed a method for decomposing IRHI when health is a linear function of a set of determinants. Each individual i is characterised by her health level $h_i \geq 0$, her income level $y_i \geq 0$ and a vector of other characteristics x_i which could include demographics, etc. The linear relationship allows us – without loss of generality – to assume that x_i includes only one variable.

$$h_i = \alpha + \beta y_i + \gamma x_i \tag{2.1}$$

where α, β, γ are parameters.[4]

We are interested in the role of income growth on the evolution of both mean health and IRHI. First, the relation between mean health and mean income is straightforward:

$$H = \alpha + \beta Y + \gamma X \tag{2.2}$$

where capital letters denote means. β measures the impact of mean income on mean health. Consequently, other aspects of the distribution of income (e.g. income inequalities) do not matter.

Like Wagstaff et al. (1991), we measure IRHI using the concentration index $C(h_i|y_i)$ – which is widely used to measure relative IRHI (see e.g. van Doorslaer et al., 1997; van Doorslaer and Koolman, 2004; Bleichrodt and van Doorslaer, 2006). It can be written as:

$$C(h_i|y_i) = \frac{2\,\mathrm{cov}(h_i; R_i)}{H} = \frac{2\sum_{i=1}^{n} h_i R_i}{\sum_{i=1}^{n} h_i} - 1 \tag{2.3}$$

where $R_i = N^{-1}(i - 0.5)$ denotes the fractional rank of income. The first equality shows that the concentration index can be written as a simple function of mean health and the covariance of health and the fractional income rank. Wagstaff et al. (2003) and others[5] have shown that a factor decomposition in the spirit of Shorrocks (1982) can be obtained for the concentration index. Combining Eqs. (2.3) and (2.1), and exploiting the linearity of the covariance function gives:

$$C(h_i|y_i) = \beta\frac{2\,\mathrm{cov}(y_i, R_i)}{H} + \gamma\frac{2\,\mathrm{cov}(x_i, R_i)}{H} = \beta\frac{Y}{H}G(y_i) + \gamma\frac{X}{H}C(x_i|y_i) \tag{2.4}$$

where $\beta Y(H)^{-1}$ and $\gamma X(H)^{-1}$ can be interpreted as 'mean elasticities', $G(y_i) = C(y_i|y_i)$ is the Gini index, and $C(x_i|y_i)$ is the concentration index of x_i. Eq. (2.4) shows that IRHI are a linear function of the income-related inequalities of its determinants weighted by their respective

[3] Note that this paper does not fit into the literature on the role of income growth on poverty reduction (e.g. Kraay, 2006; Kalwij and Verschoor, 2007) and reduction of income inequality (e.g. Barro, 2000) since we study the joint income-health distribution.

[4] Note that the methodology in this section can cope with an error term by incorporating it in x_i, but for ease of exposition it is neglected here.

[5] See also Clarke et al. (2003) and Gravelle (2003).

'mean elasticity'. The advantage of their approach for our purposes is that it clearly demonstrates that IRHI is related to mean income through the income elasticity – i.e. IRHI are determined by $\beta Y(H)^{-1}$ – but it highlights that other aspects of the distribution of income also matter, in particular the Gini index – i.e. IRHI are influenced by income inequality as measured by $G(y_i)$ –, the effect of mean income on the elasticity of x_i – i.e. IRHI are influenced by mean income affecting $\gamma X(H)^{-1}$ (through the effect on H) –, *and* the income rank in $C(x_i|y_i)$ – i.e. income-related inequalities of the determinants affect IRHI.

2.2. Decomposing changes in IRHI in a non-linear setting

While the decomposition presented in the previous section has obvious intuitive appeal, it abstracts from the well documented non-linear relationship between income and health, i.e. that health shows diminishing returns to income (e.g. Smith, 1999; Deaton, 2003; Ecob and Smith, 1999; Gerdtham and Johannesson, 2000; Gravelle and Sutton, 2003; Mackenbach et al., 2005). A common approach has been to log-transform y_i (e.g. van Doorslaer and Koolman, 2004) or to use a power function of y_i (e.g. Gravelle and Sutton, 2003) in Eq. (2.1) to preserve linearity in the transformed variable and its decomposition in Eq. (2.4). These procedures are incapable of informing on the effect of changes in the Gini and in the 'mean income elasticity' as these only inform on the contribution of the elasticity and the Gini of the transformed variable (i.e. $\ln(y_i)$ or y_i^k). We propose to resolve this problem by re-computing the first term of Eq. (2.4) in a non-linear setting. We show that it is still possible to make (albeit weaker) inferences on the relative importance of the income elasticity versus income inequality.

While our basic interest lies in Eqs. (2.2) and (2.4), we are not interested in mean health and IRHI *per se*, but rather in their evolution over time. Therefore, we introduce a decomposition of (discrete) time differences in Section 2.3 to disentangle the effects of proportional income growth and income inequality allowing for a non-linear income effect by replacing βy_i in Eq. (2.1) by the non-linear function $f(y_{it};\beta)$. Our approach bears some resemblance to Eq. (8) in Wagstaff et al. (2003), which (i) considers a total differential[6] and (ii) is formulated in a continuous framework, but has the obvious disadvantage of being only an approximation – valid for very small changes – while our approach is exact. This is important since we will analyze rather large changes in mean income. In addition, it is easier to deal with non-linearities in our approach (see below). We further assume that Eq. (2.1) holds in each time period.

First, we allow for a non-linear income effect in Eq. (2.1) and add a time subscript:

$$h_{it} = \alpha + f(y_{it};\beta) + \gamma x_{it} \tag{2.5}$$

Note that we allow the function $f(\,)$ and its slope to vary with income. For notational brevity, we do not show the dependence upon β in the remainder of the paper, i.e. $f(y_{it};\beta) = f(y_{it})$. Second, by combining Eqs. (2.2) and (2.5) and introducing discrete time differences, we obtain:

$$H_t - H_1 = \frac{\sum_{i=1}^{n} f(y_{it}) - \sum_{i=1}^{n} f(y_{i1})}{n} + \gamma(X_t - X_1) \tag{2.6}$$

In contrast to Eq. (2.2), not only mean income, but also the non-linearity of the income profile now matters due to the aggregation of a non-linear income profile.[7] For example, in the special case of an increasing and concave second order polynomial, one can show that mean income increases mean health, whereas the variance of income (or income inequality) decreases mean health. Doing a similar exercise for IRHI, by introducing a non-linear income profile, changes Eq. (2.4) into:

$$C(h_{it}|y_{it}) - C(h_{i1}|y_{i1}) = \underbrace{\frac{\sum_{i=1}^{n} f(y_{it})}{nH_t} C[f(y_{it})|y_{it}] - \frac{\sum_{i=1}^{n} f(y_{i1})}{nH_1} C[f(y_{i1})|y_{i1}]}_{\text{term 1}} + \underbrace{\gamma \frac{X_t}{H_t} C(x_{it}|y_{it}) - \gamma \frac{X_1}{H_1} C(x_{i1}|y_{i1})}_{\text{term 2}} \tag{2.7}$$

Note that the introduction of $f(\,)$ removes the exact relationship between $C(h_{it}|y_{it})$ and the income elasticity and the Gini index, except if $f(y_{it}) = \beta y_{it}$. In addition, the effects of proportional income growth (i.e. a change in mean income) and the change in income inequality on IRHI are not easily inferred from Eq. (2.7). While it is still straightforward to decompose the effect of a change of x_{i1} to x_{it} on the change in IRHI into a mean elasticity and an inequality effect using the methods in Wagstaff et al. (2003), this is no longer the case for a change of y_{i1} to y_{it}.

2.3. Disentangling proportional income growth from the evolution of income inequality

Our approach consists of the introduction of two hypothetical health levels which allow us to put more analytic structure on the decompositions in Eqs. (2.6) and (2.7). These hypothetical health levels are h_{it}^{pg} (*pg* for proportional growth) and h_{it}^{ng} (*ng* for no growth), i.e.

$$h_{it}^{pg} = \alpha + f(y_{it}^{pg}) + \gamma x_{it} \tag{2.8}$$

$$h_{it}^{ng} = \alpha + f(y_{i1}) + \gamma x_{it} \tag{2.9}$$

where $y_{it}^{pg} = y_{i1} Y_t (Y_1)^{-1}$. Eq. (2.8) presents the hypothetical health level that individual i would have had in period t if her income growth had been equal to the *actual* mean growth. A similar intuition lies behind the introduction of h_{it}^{ng}, but in contrast to Eq. (2.8), the income

[6] Wagstaff et al. (2003) also allow for changes in α, β, γ. Although we keep β fixed, the nonlinear relationship of $f(\,)$ allows for different income-effects at different income levels, and consequently at different time periods. Fixing of α and γ is less important since we focus on the impact of the distribution of income on the distribution of health. Our approach could be generalised to allow for changes in α, β, γ, but consult Section 3.1 for our reasons not to do so.

[7] Our assumption does not rely on the literature investigating a direct negative effect of income inequality on individual health (see e.g. Wildman, 2003). Literature surveys (Wagstaff and van Doorslaer, 2000; Deaton, 2003) and recent contributions with similar data (Hildebrand and Van Kerm, 2005; Lorgelly and Lindley, 2008) did not find convincing evidence for a direct effect of income inequality on individual health.

distribution remains unchanged. Combining Eqs. (2.6), (2.7), (2.8) and (2.9), we now obtain:

$$H_t - H_1 = \underbrace{\frac{\sum_{i=1}^{n} f(y_{it}) - \sum_{i=1}^{n} f(y_{it}^{pg})}{n}}_{\text{term a}} + \underbrace{\frac{\sum_{i=1}^{n} f(y_{it}^{pg}) - \sum_{i=1}^{n} f(y_{i1})}{n}}_{\text{term b}} + \underbrace{\gamma(X_t - X_1)}_{\text{term c}} \tag{2.10}$$

$$\begin{aligned} C(h_{it}|y_{it}) - C(h_{i1}|y_{i1}) = & \underbrace{\frac{\sum_{i=1}^{n} f(y_{it})}{nH_t} C[f(y_{it})|y_{it}] - \frac{\sum_{i=1}^{n} f(y_{it}^{pg})}{nH_t^{pg}} C[f(y_{it}^{pg})|y_{i1}]}_{\text{term 1a}} + \underbrace{\frac{\sum_{i=1}^{n} f(y_{it}^{pg})}{nH_t^{pg}} C[f(y_{it}^{pg})|y_{i1}] - \frac{\sum_{i=1}^{n} f(y_{i1})}{nH_t^{ng}} C[f(y_{i1})|y_{i1}]}_{\text{term 1b}} \\ & \underbrace{+ C[f(y_{i1})|y_{i1}] \left[\frac{\sum_{i=1}^{n} f(y_{i1})}{nH_t^{ng}} - \frac{\sum_{i=1}^{n} f(y_{i1})}{nH_1} \right]}_{\text{term 1c}} + \underbrace{\gamma \frac{X_t}{H_t} C(x_{it}|y_{it}) - \gamma \frac{X_t}{H_t^{pg}} C(x_{it}|y_{i1})}_{\text{term 2a}} + \underbrace{C(x_{it}|y_{i1}) \left[\gamma \frac{X_t}{H_t^{pg}} - \gamma \frac{X_t}{H_t^{ng}} \right]}_{\text{term 2b}} \\ & \underbrace{+ \gamma \frac{X_t}{H_t^{ng}} C(x_{it}|y_{i1}) - \gamma \frac{X_1}{H_1} C(x_{i1}|y_{i1})}_{\text{term 2c}} \end{aligned} \tag{2.11}$$

where $C(\ldots|y_{it}^{pg}) \equiv C(\ldots|y_{i1})$. Eq. (2.10) clearly shows that the effects of proportional income growth (term b) and changes in income inequality (term a) are easily separated, and are unambiguous, leading to Proposition 1:

Proposition 1. *Mean health responds elastically/inelastically/unit elastically to proportional income growth if Eq. (2.5) is convex/concave/linear and increasing with income, while reductions of income inequality reduce/keep unchanged/increase mean health if Eq. (2.5) is convex/linear/concave and increasing with income.*

Proof. This result was already shown by Contoyannis and Forster (1999). □

With respect to Eq. (2.11) things are less straightforward. Terms 1a–c decompose the first term of Eq. (2.7), while terms 2a–c decompose the second term of Eq. (2.7). In the next subsections, we show that the a-terms are related to the evolution of income inequality, the b-terms to proportional income growth and the c-terms to the evolution of the other determinants of health.

2.3.1. Proportional income growth: term 1b and 2b

The influence of proportional income growth on IRHI is summarized by terms 1b and 2b in Eq. (2.11), i.e.

$$\frac{\sum_{i=1}^{n} f(y_{it}^{pg})}{nH_t^{pg}} C[f(y_{it}^{pg})|y_{i1}] - \frac{\sum_{i=1}^{n} f(y_{i1})}{nH_t^{ng}} C[f(y_{i1})|y_{i1}] \tag{2.12}$$

$$C(x_{it}|y_{i1}) \left[\gamma \frac{X_t}{H_t^{pg}} - \gamma \frac{X_t}{H_t^{ng}} \right] \tag{2.13}$$

Note that (2.12) and (2.13) are zero when there is no income growth. Moreover, by definition they are not influenced by changes in income inequality, nor by changes in x_{it}.

Proposition 2. *The direct effect of proportional income growth – as measured by Eq. (2.12) – is to increase/decrease IRHI if the income elasticity is rising/falling with income.*

Proof. The sign of (2.12) is easily obtained. After some algebra, (2.12) reduces to:

$$\frac{1}{n} \sum_{i=1}^{n} \left\{ (2R_{i1} - 1) \left[\frac{f(y_{it}^{pg})}{H_t^{pg}} - \frac{f(y_{i1})}{H_t^{ng}} \right] \right\} \tag{2.14}$$

One can show that (2.14) is positive/negative if the term between square brackets increases/decreases with income.[8] Taking the partial derivative of the term between square brackets; and multiplying by Y_1 gives[9]:

$$\frac{\partial[(f(y_{it}^{pg})/H_t^{pg}) - (f(y_{i1})/H_t^{ng})]}{\partial y_{i1}} Y_1 = \frac{\partial f(y_{it}^{pg})}{\partial y_{it}^{pg}} \frac{Y_T}{H_t^{pg}} - \frac{\partial f(y_{i1})}{\partial y_{i1}} \frac{Y_1}{H_t^{ng}} \quad \square \tag{2.15}$$

[8] Note that the first term of (2.14), $2R_{i1} - 1$, is negative/positive for incomes below/above the median. Consequently, the absolute value of its sum above and below the median is identical, i.e. $\left|\sum_{i=1}^{n/2}(2R_{i1} - 1)\right| = \left|\sum_{i=n/2}^{n}(2R_{i1} - 1)\right|$. Therefore, (2.14) is positive/negative if $\sum_{i=n/2}^{n}(2R_{i1} - 1)[\ldots] > / < \sum_{i=1}^{n/2}(2R_{i1} - 1)[\ldots]$. The latter condition is satisfied if the term between square brackets increases/decreases with income.

[9] Note that we treat the partial derivatives of H_t^{pg} and H_t^{ng} to y_{i1} as zero, which is justified since we only intend to investigate how the term changes if we move up in the 'unchanged' income distribution.

Proposition 2 highlights the crucial role of the income elasticity (evaluated at the mean values Y_t and Y_1).[10] This result points to the importance of using a flexible functional form of $f(\ )$ in our empirical exercise. For example, if we were to impose a linear income profile like in (2.1), proportional income growth would always lead to increasing IRHI since its income elasticity is always increasing with income (assuming a positive income effect). If we were to take the logarithm of income – a common approach in the literature (see Section 2.2) – proportional income growth would always give rise to decreasing IRHI since its income elasticity is always decreasing with income (assuming a positive income effect).

The combination of Propositions 1 and 2 is powerful. It implies that proportional income growth leads to a (welfare improving) mean health increase and – depending on the slope of the income elasticity – to a (welfare decreasing/increasing) increase/decrease in relative income-related health inequality. Because this result has only limited applicability, as it only refers to proportional income growth, we abstract from proportional income growth in the next section and focus on income inequality changes.

The above discussion illustrates that the income elasticity is a vital element to understand the evolution of IRHI. However, (2.12) assumes that only income affects health, whereas Eq. (2.11) allows for an additional determinant x_{it}. It follows – as can be inferred from (2.13) – that proportional income growth (through its effect on H) also *indirectly* affects the elasticity of x_{it}. Although H_t^{pg} is larger/smaller than H_t^{ng} if the income elasticity is positive/negative, we cannot derive the sign of (2.13), as we do not know the sign of $C(x_{it}|y_{i1})\gamma X_t$ *a priori*.

2.3.2. Changes in income inequality: term 1a and 2a

The effect of changes in income inequality on IRHI is summarized by the terms 1a and 2a in Eq. (2.11), i.e.

$$\frac{\sum_{i=1}^{n} f(y_{it})}{nH_t} C[f(y_{it})|y_{it}] - \frac{\sum_{i=1}^{n} f(y_{it}^{pg})}{nH_t^{pg}} C[f(y_{it}^{pg})|y_{i1}] \tag{2.16}$$

$$\gamma \frac{X_t}{H_t} C(x_{it}|y_{it}) - \gamma \frac{X_t}{H_t^{pg}} C(x_{it}|y_{i1}) \tag{2.17}$$

We start by noting that both expressions are by definition not influenced by proportional income growth, nor by changes in x_{it}. The only relevant determinant is the evolution of relative income inequality between period 1 and t. Clearly, without any change in income inequality, both terms are zero.

Proposition 3. *If the evolution of income inequality is on average pro-rich/poor and if the income elasticity is increasing/decreasing with income, the direct effect of the evolution of income inequality is to increase/decrease IRHI.*

Proof. After some algebra, Eq. (2.16) reduces to:

$$\frac{1}{n}\sum_{i=1}^{n}\left\{2\left[R_{it}\frac{f(y_{it})}{H_t} - R_{i1}\frac{f(y_{it}^{pg})}{H_t^{pg}}\right] - \left[\frac{f(y_{it})}{H_t} - \frac{f(y_{it}^{pg})}{H_t^{pg}}\right]\right\} \quad \square \tag{2.18}$$

Eq. (2.16) – which keeps mean income fixed – shows that a change in income inequality has an effect on IRHI through the R-terms and through the income elasticity. The intuition of Proposition 3 can be understood from a comparison between (2.14) and (2.18) which shows that in case of a pro-rich (pro-poor) change in income inequality and an elasticity that increases (decreases) with income, both effects have the same sign. These conditions are sufficient, but not necessary: if these effects are opposite, then one cannot *a priori* determine the sign of (2.16). Intuitively, the latter means that an increase/decrease in income inequality is offset by local changes in the income elasticity; which of the two effects dominates is then an empirical question.

Similar to Section 2.3.1, it is of vital importance to allow for a flexible functional form of $f(\ )$. For example, a linear income profile imposes that only relative income inequality (as measured by the Gini) matters for (2.16) *and* that the effect through the income elasticity is non-existent. The latter effect is relevant for the functional form with the logarithm of income (as is the effect of relative income inequality), but as we showed in Section 2.3.1, the income elasticity is always decreasing with income, which imposes that decreasing income inequalities always lead to decreasing IRHI.

It is important to add that (2.17) shows that a change in income inequality has two additional effects, i.e. (i) it affects the concentration indices of x_{it} through differences in the fractional rank, and (ii) it affects the 'mean elasticity of the x_{it}-determinant' through H. The sign of (2.17) cannot be determined *a priori*.

It is worth emphasizing that the results in this section imply that – contrary to a common belief[11] (e.g. Blakely and Wilson, 2006; Avendano et al., 2006; Dahlgren and Whitehead, 2006) – reductions in income inequality do *not necessarily* lead to lower IRHI – since there will be interactions between the evolution of income inequality and the 'slope of the income elasticity' as soon as income affects health in a non-linear way.

[10] A more restricted version of this condition was established by Contoyannis and Forster (1999). They imposed that the income elasticity has to rise/fall *monotonically* with income. Since the partial derivatives of $f(y_{it}^{pg})$ and $f(y_{i1})$ in (2.15) are at the individual level, it is sufficient if the income elasticity (*evaluated at* y_{i1}) increases/decreases 'on average' and not at each point of the income profile.

[11] Contoyannis and Forster (1999) is a notable exception.

2.3.3. The importance of other determinants: terms 1c and 2c

Terms c of Eq. (2.11) measure the effect of the change from x_{i1} to x_{it}. Term 1c measures the effect of this change through H while term 2c summarizes the effect of this change on the 'mean elasticity' of x_{it} and the effect that runs via changes in the concentration index of x_{it}.

$$C[f(y_{i1})|y_{i1}]\left[\frac{\sum_{i=1}^{n}f(y_{i1})}{nH_t^{ng}}-\frac{\sum_{i=1}^{n}f(y_{i1})}{nH_1}\right] \tag{2.19}$$

$$\gamma\frac{X_t}{H_t^{ng}}C(x_{it}|y_{i1})-\gamma\frac{X_1}{H_1}C(x_{i1}|y_{i1}) \tag{2.20}$$

Note that both expressions drop out of Eq. (2.11) if x_{it} is constant over time. However, we cannot derive the sign of (2.19) and (2.20) *a priori*. This is not a disadvantage since both expressions only enter Eq. (2.11) to correct the effect of changes in the income distribution on IRHI for the evolution of other health determinants. In other words, these are just control terms.

In summary, the above approach builds on Wagstaff et al. (2003), but explicitly accounts for a non-linear relationship between income and health, while allowing for large income changes. Depending on the restrictions imposed on the non-linear income profile, one can derive the sign of the effects of changes in mean income and changes in the variability of individual income on mean health. For instance, assuming a concave profile, one can show a positive effect from proportional income growth and a negative effect from rising income inequality. The picture is more complicated for the effect of changes in the income distribution on the evolution of IRHI. In order to disentangle the effect of proportional income growth from the impact of changes in income inequality, we introduced two hypothetical health levels, i.e. (i) the health level that would prevail in case of a non-changing income distribution and (ii) the health level that would prevail in case of proportional income growth. This enabled us (i) to isolate the effect of changes in the income distribution from changes in the other health determinants, and (ii) to isolate the effect of changes in income inequality from proportional income growth. In both instances there is a direct effect of the change in the income distribution on IRHI, but also an indirect effect through the other health determinants. Building on Contoyannis and Forster (1999) we showed that the *direct* effect of proportional income growth depends on the slope of the income elasticity. If this elasticity is rising/decreasing with income 'on average' (see Proposition 2), IRHI increase/decline. With respect to the *direct* effect of changes in income inequality, we found increasing/decreasing IRHI in case of a pro-rich/pro-poor change in income inequality in combination with an income elasticity that increases/decreases with income 'on average' (see Proposition 3). We find that pro-poor changes in income inequality do not always lead to reductions in IRHI if income inequality and the elasticity do not move together 'on average'. In the latter case, they have an opposite effect and the net effect is an empirical issue that cannot be determined *a priori*. The sign of the indirect effects (both for proportional income growth and income inequality) could not be inferred since these depend on the concentration indices and the elasticity of the other health determinants. Obviously, these are not known *a priori*.

Our empirical analysis has three objectives. First, as an introduction, we present estimates of the income elasticity of health and show how these vary with income in European countries. Because the income elasticity and its slope are crucial to determine the consequences of income growth, we use a flexible functional form in the estimation. Secondly, we describe the empirically observed trends of income growth, income inequality, mean health and IRHI in Europe in the 1990s. Third, we use our decomposition technique to isolate the effects of proportional income growth and changing income inequality.

3. Data and empirical model specification

The data used in this paper are taken from the full 8 annual waves (1994–2001) of the *European Community Household Panel User Database* (ECHP-UDB). The ECHP was designed and coordinated by EUROSTAT, and it contains socioeconomic, demographic and health variables, for a panel of households which only includes individuals aged 16 or older. We use all waves that are available for 13 EU member states: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain and the United Kingdom. We did not analyze the data for Luxembourg (small sample) and Sweden (no panel data). For Germany and the UK, we did not use the ECHP (which only ran from 1994 to 1997, i.e. waves 1–3) but instead used the corresponding waves from the *German Socio-economic Panel* (GSOEP) and the *British Household Panel Survey* (BHPS) that are provided in the ECHP-UDB. Austria joined the survey in 1995 (wave 2) and Finland in 1996 (wave 3).

3.1. Estimating the elasticity of health with respect to income

The two key variables for this study are health and income. The ECHP income measure is self-reported disposable (i.e. after-tax) household income, which is all net monetary income received by the household members during the *previous* year. It includes income from work (employment and self-employment), private income (from investments and property and private transfers to the household), pensions and other direct social transfers received. No account has been taken of indirect social transfers (e.g. reimbursement of medical expenses), receipts in kind and imputed rent from owner-occupied accommodation. The income variable is (i) converted in Euros by yearly PPPs (see EUROSTAT, 2003) to allow for comparability across countries, and (ii) expressed in constant (1996) prices, i.e. deflated by the harmonised index of consumer prices (HICP), to allow for comparability across waves (European Central Bank, 2000, 2003). Income was then divided by the OECD modified equivalence scale in order to account for household size and composition (giving a weight of 1.0 to the first adult, 0.5 to the second and each subsequent person aged 14 and over, and 0.3 to each child aged under 14 in the household).

Self-assessed health (SAH) is measured as the response to an ordered 5-point scale (ranging from very good to very poor) on the question "How is your health in general?" In addition to the mere language differences, the question wording was slightly different in 3 of the 13 countries. For France and Germany, a 6-point and 10-point (health satisfaction) scale respectively was recoded into the common 5-point scale by EUROSTAT. In the UK, the question wording adds a reference to people of the same age (except for wave 6). Reporting heterogeneity in self-assessed health across cultures and populations is a notable concern (Lindeboom and van Doorslaer, 2004) but as our paper is basically about health trends *within* countries this is less of a concern here. We have adopted the scaling methods based on interval regression proposed by van Doorslaer and Jones (2003) and used by van Doorslaer and Koolman (2004) on the ECHP data.[12,13] This approach assumes that there is a stable mapping from the Health Utility Index (HUI) (see e.g. Feeny et al., 2002) to the (latent) variable that determines reported *SAH*. The interval regression results in individual predicted health utility scores that have cardinal measurement properties. While the internal validity of this approach was confirmed in Canadian data (van Doorslaer and Jones, 2003), it is not possible to test the external validity on the European data. However, sensitivity analysis using other boundaries has shown that the results are almost identical when the imposed thresholds were derived from other (European) data and generic measures like the Euroqol (Lauridsen et al., 2004; Lecluyse and Cleemput, 2005).

Because we intend to investigate the impact of proportional income growth and changing income inequality on mean health and IRHI, we need an estimate of the income effect in Eq. (2.5) that can be interpreted as a causal effect. One could apply a simultaneous structural estimation technique to estimate a Grossman type model (e.g. Wagstaff, 1993), but we have opted for a single equation approach because of its transparency, and because we are only interested in the overall income effect here, not its pathways to influence income. In our health equation, we have included as covariates – besides income – only demographics like age and gender:

$$h_{it}^* = \alpha + f(y_{it}) + x_{it}'\gamma + \varepsilon_{it} \tag{3.1}$$

where h_{it}^* is the latent health outcome, $f(y_{it})=f(y_{it};\beta)$ is a non-linear function of income (see below), α, β and γ are parameters to be estimated, and $\varepsilon_{it} \sim N(0, \sigma_\varepsilon^2)$. We do not observe h_{it}^*, but we do observe *SAH* and can impose its interval boundaries derived from HUI scores. As a result, its predictions are contained in the [0,1] interval and can be interpreted as (predicted) health utilities on the HUI scale. The vector of covariates (x_{it}) includes age dummies (categories: 16–29; 30–44; 45–59; 60–69; 70+) for both sexes. We limit the specification to only these covariates on the grounds that these can safely be assumed to be exogenous and that we are mainly interested here in an estimate of the overall income elasticity of health (utility), not in the effects of endogenous variables (like life style or labour choices) that may channel the effect of income on health.[14]

Despite the exogeneity of the covariates and the fact that our income measure refers to disposable household income of the previous year – which makes it less prone to reverse causation bias compared to current income – our approach does not necessarily solve the endogeneity problem between income and health. Therefore, we have also estimated a dynamic version of Eq. (3.1) which includes *SAH* for the previous and the first wave (see e.g. Hurd and Kapteyn (2003), Contoyannis et al. (2004) and Jones et al. (2006) for similar approaches). As the latter specification captures state dependence, removes any correlation between income and initial health from the estimate of β, and models income effects on health transitions, it is far less likely to reflect reverse causation. A comparison of the latter estimates with estimates based on Eq. (3.1) revealed no major differences except for a smaller income effect. The latter finding is obvious given that the dynamic approach effectively models health transitions. We also repeated all other analyses in Section 4 and found no important differences. Since the dynamic approach did not alter the qualitative interpretations of any of our analyses and since the approach based on Eq. (3.1) is less complex and saves an additional wave, we decided to present the latter. Results from the dynamic approach are obtainable from the authors upon request.

We have run pooled[15] models on the balanced panel of 8 waves. We considered, but do not present unbalanced panels for two reasons. First, Jones et al. (2006) have shown that health-related non-response in the ECHP hardly affects estimates of income effects in health equations. Second, although our restriction to a balanced panel means that the results apply to a cohort of individuals, we have repeated all analyses in Section 4 using an unbalanced panel – such that our results are no longer cohort-specific – and found no important differences. In Section 4, we provide some additional discussion of this issue.

We did not include any time dummies in (3.1) on the grounds that these may capture some mean income changes and moreover, in 9 countries the set of time dummies was jointly not statistically significant. We have also kept the β's fixed across time (see also footnote 6). For 7 countries we could not reject the null that interactions between time dummies and the non-linear income profile are statistically irrelevant. Moreover, in those cases where the null was rejected, almost none of the individual interactions were statistically significant. We have also experimented with models that allow for interactions between the income polynomial and the age dummies for both sexes, but did not prefer these estimates as the resulting income elasticities hardly differed from those resulting from the more parsimonious specification in Eq. (3.1).[16] The cross-country differences in health, income and demographics are documented in Table 3.1 which presents unweighted means of all variables for the pooled balanced samples.

In view of the literature on a non-linear, concave relationship between income and individual health, the discussion on the relevance of Eq. (2.5) in Section 2.2 and the importance of non-linearity for the current paper, we allow for a flexible functional form by implementing polynomial transformations of income. This allows for the income elasticity of health to decrease with income in some income ranges and increase in others. The order of the polynomial was determined by first, estimating each model with a fifth order polynomial and then reducing the order of the polynomial until a likelihood ratio test (1% significance level) rejected the 'reduced order' against the higher order polynomial.

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[12] Thresholds are 0, 0.428, 0.756, 0.897, 0.947, and 1.

[13] We have not considered the more flexible ordered probit model to scale the SAH responses, since van Doorslaer and Jones (2003) have shown it is outperformed by the interval regression approach. In addition, we would loose the neat interpretation and advantages of the HUI scale.

[14] For adults, education would also be an obvious exogenous candidate to include and the ECHP records information on the highest level of general or higher education completed. Nevertheless, we had to exclude education from our regression model as EUROSTAT (2003) notes classification problems related to this variable.

[15] We also experimented with panel models. Since a fixed effects interval regression model suffers from the incidental parameters problem (Wooldridge, 2002), we estimated a random effects interval regression model in which we parameterised the individual effects as a function of the means of time-varying variables (Chamberlain, 1980). We also estimated a traditional random effects interval regression model. Although the estimated β's were in both cases very similar to those of the pooled model, we prefer the pooled specification (in which we correct the statistical inference for clustering at the individual level) as it imposes less stringent exogeneity assumptions (see e.g. chapter 15 in Wooldridge, 2002).

[16] The differences were slightly more pronounced for Portugal (and to a lesser extent for Italy and Ireland), but the resulting income elasticities still followed the same pattern as those based on the more parsimonious model in Eq. (3.1).

Table 3.1
Summary statistics by country: (unweighted) means.

	Austria	Belgium	Denmark	Finland	France	Germany	Greece	Ireland	Italy	Netherlands	Portugal	Spain	UK
M1629	0.089	0.052	0.059	0.062	0.064	0.069	0.045	0.069	0.096	0.041	0.080	0.077	0.066
M3044	0.141	0.173	0.163	0.147	0.148	0.167	0.126	0.142	0.141	0.167	0.121	0.128	0.146
M4559	0.129	0.120	0.150	0.139	0.132	0.140	0.128	0.143	0.133	0.143	0.118	0.113	0.124
M6069	0.072	0.059	0.062	0.069	0.064	0.068	0.090	0.072	0.070	0.060	0.078	0.078	0.058
M70+	0.049	0.050	0.047	0.031	0.053	0.032	0.068	0.055	0.046	0.048	0.062	0.067	0.053
F1629	0.076	0.061	0.066	0.075	0.077	0.079	0.069	0.071	0.094	0.058	0.073	0.082	0.078
F3044	0.152	0.203	0.171	0.172	0.170	0.176	0.142	0.157	0.149	0.198	0.136	0.143	0.181
F4559	0.138	0.131	0.162	0.182	0.142	0.140	0.137	0.151	0.137	0.146	0.142	0.131	0.152
F6069	0.080	0.077	0.061	0.072	0.079	0.074	0.099	0.075	0.073	0.075	0.097	0.091	0.065
F70+	0.074	0.073	0.058	0.051	0.070	0.055	0.096	0.064	0.061	0.064	0.091	0.091	0.077
income	14,631	16,298	15,718	12,948	14,516	15,030	7836	11,876	10,811	14,264	7293	9611	14,962
sahverybad	0.013	0.006	0.010	0.006	0.034	0.032	0.021	0.005	0.016	0.005	0.039	0.017	0.019
sahbad	0.058	0.035	0.040	0.054	0.038	0.138	0.074	0.022	0.092	0.037	0.194	0.104	0.076
sahfair	0.216	0.216	0.170	0.309	0.338	0.348	0.185	0.165	0.299	0.232	0.343	0.245	0.223
sahgood	0.429	0.531	0.328	0.461	0.473	0.406	0.270	0.367	0.437	0.561	0.398	0.485	0.466
sahverygood	0.285	0.212	0.451	0.170	0.118	0.076	0.450	0.441	0.157	0.165	0.026	0.150	0.216
N	27,769	24,200	20,352	19,314	54,688	60,160	49,072	22,976	72,288	36,448	56,776	58,456	47,400

We did not use the EUROSTAT-provided cross-sectional individual sampling weights to estimate Eq. (3.1), but we did for the estimation of the elasticity of health with respect to income (see e.g. chapter 24 in Cameron and Trivedi, 2005):

$$\hat{\varepsilon}^{y} = \frac{\partial f(y_{it};\hat{\beta})}{\partial y_{it}} \frac{y_{it}}{\hat{h}_{it}} \tag{3.2}$$

where ^ denotes an estimate and $\hat{h}_{it}$ is the predicted value of Eq. (3.1). We computed Eq. (3.2) for each individual and calculate the (weighted) mean to obtain the mean elasticity over *all time periods*. In order to verify whether the elasticity is increasing/decreasing with income, this procedure was repeated for each income decile for all time periods.[17]

3.2. Descriptive statistics for health, income and their distributions

With respect to the distribution of (predicted) health (utility), we calculate $\hat{H}_t$ and $C(\hat{h}_{it}|y_{it})$ for each t using the balanced panel. The concentration index of health $C(\hat{h}_{it}|y_{it})$ is computed using a separate OLS-regression for each wave t. Kakwani et al. (1997) have shown that the point estimate of $\hat{\lambda}_t$ in the following equation equals $C(\hat{h}_{it}|y_{it})$.

$$2\hat{\sigma}^2_{R_t} \frac{\hat{h}_{it}}{\hat{H}_t} \sqrt{w_i} = \phi_t \sqrt{w_i} + \lambda_t \sqrt{w_i}\hat{R}_{it} + \zeta_{it} \tag{3.3}$$

where $\hat{h}_{it}$ is the predicted value of Eq. (3.1), and thus the resulting $C(\hat{h}_{it}|y_{it})$ can be interpreted as income-related inequality in predicted health utility. $\hat{H}_t$ is the weighted mean of $\hat{h}_{it}$ in wave t, w_i is the sampling weight[18] of individual i in wave 1, ϕ_t and λ_t are parameters to be estimated, and ζ_{it} is an error term with zero mean. $\hat{R}_{it}$ is the estimated weighted fractional rank of income in wave t and $\hat{\sigma}^2_{R_t}$ is the estimated weighted variance of $\hat{R}_{it}$ in wave t.

We estimate two characteristics of the income distribution using the balanced panel. First, we calculate the weighted mean income in each period t. Second, we calculate the Gini index of income in each wave t using Eq. (3.3) where $\hat{h}_{it}$ and $\hat{H}_t$ have been replaced by respectively y_{it} and $\hat{Y}_t$.

3.3. The role of proportional income growth and changing income inequality

We estimate (weighted) versions of (2.12), (2.13), (2.16), (2.17), (2.19) and (2.20) for each country in each wave t, except the first. To compute the decomposition terms, we need estimates of H_t^{pg} and/or H_t^{ng} which are the weighted means of $\hat{h}_{it}^{pg}$ and $\hat{h}_{it}^{ng}$. The latter are obtained by substituting y_{it} by respectively $\hat{y}_{it}^{pg}$ and y_{i1} in Eq. (3.1), and calculating the predicted value of health while keeping the coefficients fixed and the other variables at their actual value. For $\hat{h}_{it}^{pg}$ we need to generate an estimate for $\hat{y}_{it}^{pg}$, i.e. $y_{i1}(Y_t/Y_1)$. The latter estimate allows proportional income growth to differ between each period t and the first period. The sums of (2.12)–(2.13), (2.16)–(2.17) and (2.19)–(2.20) provide an indication of the *total* effect of respectively proportional income growth, income inequality, and the other determinants.

3.4. Statistical inference

For statistical inference on the point estimates of the income elasticity at the various deciles, the trends in mean income, the income Gini, mean health, the health concentration index and its decomposition, we use the bootstrap procedure of Mills and Zandvakili (1997).

[17] We also calculated elasticities for each decile in each time period since these are more relevant to get inferences on Eq. (2.12). However since the results are confirmed by the elasticities for all time periods and since it would overload the paper with additional tables, we decided not to present these results.
[18] Due to the restriction to a balanced panel we applied the first period weights to all subsequent periods.

We draw 100 bootstrap samples[19] on the level of the individual (i.e. if an individual is drawn in one time period, he is included in all time periods) which corrects the statistical inference for the dependence between time periods. For each bootstrap sample, we repeat all calculations, and compute 95% normal confidence intervals for the elasticities, the trends and all terms of the decomposition in Section 2.3.

4. Empirical results

All analyses presented in this section are based on a balanced panel, i.e. illustrating how changes in the income distribution within a single cohort affect mean health and IRHI. In principle, our analysis does not have to be restricted to a cohort and can similarly be used to track changes in the income distribution on the health distribution of a general population. We have repeated all analyses in this section using an unbalanced panel and found that these generally mimic our findings using the balanced panel. A slight difference occurred in two cases (discussed further in this section), but this did not affect the conclusions of our decomposition exercise. We therefore conclude that the general findings in this section are applicable beyond the single cohort and of broader relevance. All additional results are obtainable from the authors upon request.

We first present results on one of the most crucial elements in our decomposition, i.e. the income elasticity of health. Next, we present the trends in mean income and health, and income inequalities and IRHI, and only in the last subsection we discuss the decomposition itself.

Due to space limitations, we only summarize the estimates of the regression model, but full results are available in Appendix A. The coefficients for the age–gender dummies show the expected signs and magnitudes, i.e. younger and male persons have better health than older and female respondents. Income coefficients showed a concave and highly non-linear pattern and were (jointly) significant at the 1% level in all specifications and for all countries.

4.1. Income elasticity of health

A summary of the income elasticity estimates (averages over all time periods) is presented in Table 4.1. All elasticities are below one – implying a concave income profile – but positive and increasing with income, except for Austria where the confidence intervals of all deciles overlap. In all other countries, the point estimate of the elasticity starts decreasing only at the highest deciles, but the decline is only statistically relevant for Greece and Ireland.

In general, we can fairly safely conclude that the income elasticity is positive and non-decreasing with income over most of the income range. The elasticities are rather low; for example, a doubling of income in Austria, results on average in a 2.7% increase in health. This may be related to the fact that (good) health has an upper bound while income is unbounded. Nevertheless, the elasticity differences across deciles are highly relevant. Countries with particularly large differences between higher and lower deciles are Belgium, Greece, Italy, Portugal and Spain.

4.2. Trends in real incomes and income inequality, and trends in mean health and income-related health inequalities

First of all, it is obvious from Table 4.2 (compare y_{2001} with y_{1994}) and Panel a of Fig. 4.1 that income growth has been unequal across European countries. There have been ups and downs, but over the entire period (1994–2001), mean incomes have seen statistically significant growth in *all* countries in real terms. In percentage terms, mean annual real income growth has been particularly strong in Portugal (4%), Spain (3.6%), Ireland (3%), Greece (3%), UK (2.8%) and the Netherlands (2%), while it was below 2% in the other countries.

Second, the Gini trends in Table 4.2 (compare G_{2001} with G_{1994}) and Panel b of Fig. 4.1 indicate that very few countries have experienced a sustained increase in income inequality over the period 1994–2001. While the trends are by no means monotonic, it is clear from the graph that, on the whole, most countries have experienced either pro-poor income growth (Austria, France, Germany, Greece, Italy, Portugal and Spain) or income inequality has remained fairly stable (Belgium, Denmark, Ireland, the Netherlands and the UK). The sole exception is Finland which shows a statistically significant positive trend: its Gini index was about 10% higher in 2001 than in 1996.

While these findings may be seen as somewhat surprising in view of the often reported rising relative income inequality over this period in the OECD context (see e.g. Smeeding, 2002; Kenworthy and Pontusson, 2005), they are consistent with earlier findings reported by Garcia et al. (2004) and Hildebrand and Van Kerm (2005) on the same data and with the series of cross sections compared in Atkinson (2003) and Moran (2006a,b). For example, for the same period, also Atkinson (2003) reports stable Gini indices for the Netherlands, Italy and the UK, a modest rise in Germany and a strong increase only in Finland. In addition, the strong increase for Finland is in line with the observed reduction of progressivity of the Finnish tax system (Jäntti, 2005).[20] Also noteworthy is that the estimates of income growth and income inequality based on an unbalanced panel are largely confirmed by those resulting from the balanced panel. This suggests that attrition is not the main reason of our findings.

Third, we discuss the trends in mean health utility (compare h_{2001} with h_{1994} in Table 4.2 and consult Panel c in Fig. 4.1). Since reporting heterogeneity in self-assessed health might invalidate comparisons between countries, but probably not within countries, we only compare relative *changes* in mean health utility, not absolute values. As expected, mean health of the ageing cohort decreases but at a slow *and* statistically insignificant rate in all countries except in the UK where it remained stable.[21] This is one of the two findings that differ if an unbalanced panel is used where births and deaths are included. In the latter case, mean health remains stable or is slightly increasing (mostly not significantly). However, it turns out that our decomposition is not affected by this finding (see also further): the difference between a balanced and unbalanced panel is in our empirical example captured by the control term in Eq. (2.10).

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[19] We did not consider more replications due to the time-intensity of our procedure and since the null hypothesis of a normal distribution could not be rejected (at 5% level) for 92% of our bootstrap samples in our application.

[20] We thank Unto Häkkinen for this suggestion.

[21] It is quite likely that the diverging finding for the UK is influenced by the health assessment 'compared to your own age' (see also Section 3.1).

Table 4.1
Summary of income elasticity estimates.

	Austria			Belgium			Denmark			France			Finland			Germany			Greece		
	95−	point	95+	95−	point	95+	95−	point	95+	95−	point	95+	95−	point	95+	95−	point	95+	95−	point	95+
average	0.023	0.027	0.031	0.009	0.014	0.018	0.022	0.028	0.034	0.026	0.030	0.034	0.016	0.021	0.025	0.024	0.030	0.035	0.018	0.021	0.024
decile 1	0.015	0.021	0.026	0.003	0.006	0.010	0.013	0.017	0.021	0.014	0.016	0.018	0.010	0.013	0.016	0.013	0.018	0.023	0.008	0.009	0.011
decile 2	0.020	0.027	0.033	0.005	0.009	0.014	0.017	0.022	0.027	0.019	0.023	0.026	0.013	0.017	0.021	0.020	0.026	0.033	0.013	0.015	0.017
decile 3	0.022	0.029	0.035	0.006	0.011	0.016	0.019	0.025	0.030	0.022	0.026	0.030	0.014	0.019	0.023	0.022	0.029	0.035	0.016	0.018	0.021
decile 4	0.024	0.030	0.036	0.007	0.012	0.017	0.021	0.027	0.032	0.025	0.029	0.033	0.015	0.020	0.025	0.023	0.030	0.037	0.018	0.021	0.024
decile 5	0.025	0.030	0.036	0.008	0.013	0.019	0.022	0.028	0.034	0.027	0.031	0.035	0.016	0.021	0.026	0.025	0.032	0.039	0.020	0.023	0.026
decile 6	0.026	0.031	0.036	0.009	0.015	0.020	0.024	0.030	0.036	0.029	0.033	0.037	0.017	0.022	0.027	0.026	0.033	0.040	0.022	0.025	0.028
decile 7	0.026	0.030	0.035	0.011	0.016	0.021	0.025	0.031	0.038	0.031	0.035	0.040	0.018	0.023	0.028	0.027	0.034	0.040	0.023	0.026	0.030
decile 8	0.025	0.029	0.034	0.012	0.017	0.023	0.026	0.033	0.040	0.033	0.037	0.042	0.019	0.024	0.029	0.028	0.034	0.041	0.024	0.028	0.031
decile 9	0.021	0.027	0.032	0.014	0.019	0.025	0.028	0.035	0.042	0.034	0.039	0.044	0.020	0.025	0.030	0.026	0.034	0.041	0.025	0.028	0.031
decile 10	0.003	0.015	0.027	0.010	0.017	0.025	0.022	0.030	0.038	0.027	0.032	0.038	0.016	0.022	0.028	0.011	0.026	0.041	0.013	0.017	0.022

	Ireland			Italy			Netherlands			Portugal			Spain			UK		
	95−	point	95+	95−	point	95+	95−	point	95+	95−	point	95+	95−	point	95+	95−	point	95+
average	0.018	0.021	0.025	0.016	0.018	0.020	0.012	0.015	0.018	0.034	0.037	0.040	0.022	0.025	0.027	0.025	0.028	0.032
decile 1	0.013	0.016	0.019	0.006	0.007	0.008	0.007	0.009	0.011	0.013	0.016	0.018	0.010	0.011	0.013	0.011	0.014	0.017
decile 2	0.016	0.020	0.024	0.010	0.012	0.014	0.009	0.012	0.015	0.022	0.026	0.029	0.015	0.018	0.020	0.017	0.021	0.025
decile 3	0.017	0.022	0.026	0.013	0.015	0.016	0.010	0.013	0.016	0.027	0.031	0.035	0.018	0.021	0.023	0.020	0.024	0.028
decile 4	0.019	0.023	0.028	0.015	0.017	0.019	0.011	0.014	0.017	0.031	0.035	0.039	0.020	0.023	0.026	0.023	0.027	0.031
decile 5	0.020	0.025	0.029	0.016	0.019	0.021	0.012	0.015	0.019	0.034	0.039	0.043	0.023	0.025	0.028	0.025	0.030	0.034
decile 6	0.022	0.026	0.030	0.018	0.020	0.023	0.013	0.016	0.020	0.038	0.042	0.046	0.025	0.028	0.030	0.028	0.032	0.036
decile 7	0.022	0.026	0.030	0.020	0.022	0.024	0.014	0.017	0.021	0.041	0.045	0.050	0.027	0.030	0.032	0.030	0.034	0.038
decile 8	0.021	0.025	0.029	0.021	0.024	0.026	0.015	0.018	0.022	0.044	0.049	0.053	0.029	0.031	0.034	0.032	0.036	0.040
decile 9	0.018	0.022	0.027	0.022	0.025	0.028	0.015	0.019	0.023	0.047	0.052	0.057	0.030	0.033	0.036	0.033	0.038	0.042
decile 10	0.001	0.008	0.015	0.017	0.023	0.028	0.014	0.018	0.022	0.022	0.036	0.050	0.020	0.026	0.032	0.019	0.028	0.036

Note: Elasticities are derived from an interval regression model (see Section 3.1 and Eq. (3.2)). The column ‘point’ refers to point estimate, and the columns ‘95−’ and ‘95+’ to the 95% confidence interval. The confidence intervals are derived from the bootstrap procedure described in Section 3.4.

Table 4.2
Decomposition results, 1994–2001.

	Austria	Belgium	Denmark	Finland	France	Germany	Greece	Ireland	Italy	Netherlands	Portugal	Spain	UK
CI_{1994}	0,0092	0,0058	0,0067	0,0057	0,0081	0,0088	0,0140	0,0073	0,0053	0,0041	0,0201	0,0094	0,0091
h_{1994}	0,9042	0,9090	0,9205	0,8909	0,8745	0,8439	0,9092	0,9263	0,8791	0,9056	0,8370	0,8829	0,8860
y_{1994}	14546	15754	14539	11596	14177	15431	6817	10089	10020	13000	7183	8911	13625
G_{1994}	0,2630	0,2908	0,2158	0,2246	0,3335	0,2832	0,3723	0,3168	0,3331	0,2550	0,3669	0,3380	0,3004
CI_{2001}	0,0112	0,0076	0,0103	0,0099	0,0102	0,0091	0,0166	0,0099	0,0076	0,0043	0,0237	0,0127	0,0111
h_{2001}	0,8946	0,9028	0,9145	0,8831	0,8678	0,8294	0,8962	0,9246	0,8646	0,9008	0,8245	0,8752	0,8859
y_{2001}	15711	17118	16313	12660	15889	16819	8426	12509	11534	15081	9527	11483	16652
G_{2001}	0,2487	0,2951	0,2294	0,2495	0,2712	0,2525	0,3337	0,3101	0,2936	0,2399	0,3602	0,3176	0,3014
N	3967	3025	2544	3219	6836	7520	6134	2872	9036	4556	7097	7307	5925
CI_{2001}-CI_{1994}	0,00205	0,00175	0,00358	0,00421	0,00206	0,00029	0,00266	0,00252	0,00228	0,00016	0,00355	0,00330	0,00199
ineq_direct	-0,00030	-0,00026	0,00042	0,00037	-0,00040	-0,00113	-0,00024	0,00017	-0,00068	-0,00033	-0,00093	-0,00062	-0,00059
elas_direct	-0,00004	0,00024	0,00031	0,00015	0,00038	0,00015	0,00030	-0,00012	0,00043	0,00026	0,00146	0,00081	0,00075
age_direct	0,00010	0,00003	0,00006	0,00005	0,00012	0,00017	0,00015	0,00004	0,00012	0,00003	0,00035	0,00013	0,00005
ineq_indirect	0,00258	0,00179	0,00244	0,00362	0,00197	0,00122	0,00214	0,00254	0,00261	0,00039	0,00249	0,00314	0,00168
elas_indirect	0,00000	0,00000	0,00000	0,00000	0,00000	0,00000	-0,00003	0,00000	0,00000	0,00000	-0,00007	-0,00001	0,00000
age_indirect	-0,00028	-0,00005	0,00037	0,00002	-0,00001	-0,00012	0,00034	-0,00010	-0,00020	-0,00018	0,00026	-0,00015	0,00011
ineq	0,00228	0,00153	0,00286	0,00399	0,00157	0,00009	0,00190	0,00271	0,00193	0,00006	0,00156	0,00252	0,00109
elas	-0,00004	0,00024	0,00030	0,00015	0,00038	0,00014	0,00027	-0,00013	0,00043	0,00026	0,00139	0,00080	0,00074
age	-0,00018	-0,00002	0,00043	0,00007	0,00010	0,00005	0,00050	-0,00006	-0,00009	-0,00015	0,00060	-0,00002	0,00016

Note: CI = $C(h_{it}|y_{it})$ = concentration index of health; h = H_t = mean health; y = Y_t = mean income; G = $G(y_{it})$ = Gini of income; N: number of observations per year; $CI_t - CI_1 = C(h_{it}|y_{it}) - C(h_{i1}|y_{i1})$; ineq_direct: Eq. (2.16); elas_direct: Eq. (2.12); age_direct: Eq. (2.19); ineq_indirect: Eq. (2.17); elas_indirect: Eq. (2.13); age_indirect: Eq. (2.20); ineq: sum of all ineq_-terms; elas: sum of all elas_-terms; age: sum of all age_-terms; shaded: statistically significantly different from zero at 5% level, based on a bootstrap procedure: see Section 3.4. For Austria and Finland, 2001 is compared with 1995 and 1996 respectively.

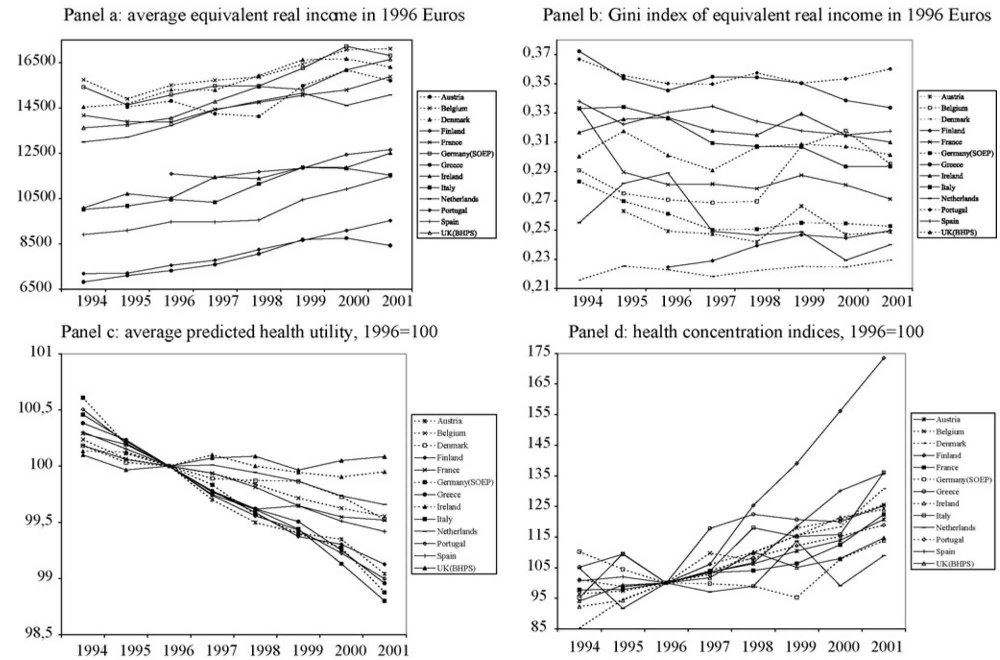

Fig. 4.1. Trends for a single cohort in 13 EU countries, 1994–2001. Panel a: average equivalent real income in 1996 Euros. Panel b: Gini index of equivalent real income in 1996 Euros. Panel c: average predicted health utility, 1996 = 100. Panel d: health concentration indices, 1996 = 100.

Fourth, we focus on the trends of health concentration indices within countries (consult Panel d in Fig. 4.1 and compare CI_{2001} with CI_{1994} in Table 4.2) and do not compare between countries. The positive and increasing concentration indices indicate that relative IRHI favoured the rich in all countries and increased between 1994 and 2001 in all countries, except Germany and the Netherlands. But they have clearly risen much faster in Finland than in any other European country in this period.[22] The marked difference between Finland and the remainder of Europe also emerges when we use the unbalanced panel, but the relative change of the concentration indices is in general somewhat lower in all countries. This is probably due to the fact that the correlation between income and health is higher within a cohort than across cohorts. However, and reassuringly, as was the case for the evolution of mean health utility, the difference between the balanced and unbalanced panel is in this empirical exercise again captured by the control term in Eq. (2.11).

4.3. Mean health and IRHI: the role of income growth and income inequality

Calculation of Eq. (2.10) was redundant since one can determine *a priori* the sign of the effect of income growth and income inequality on mean health trends. Since the income elasticities reported in Section 4.1 are all between zero and one, the income profile is concave. The income elasticities are also increasing over most of the income range. Consequently, mean predicted health utility will respond inelastically to proportional income growth and rising/decreasing income inequalities will have a negative/positive effect on mean health. Since all countries (except Finland) experienced stable or decreasing income inequalities, the combined health effect of income growth and income inequality was positive between 1994 and 2001. Note that this does not contradict the (insignificant) decreases in mean health levels and increases in mean income levels reported in the previous section. As explained, the decrease in mean health levels can be traced down to the restriction to a balanced panel. This implies a negative contribution of the x_{it} explanatory variables to mean health, as captured by the control term in Eq. (2.10), i.e. the effects of income growth and income inequality are similar for the unbalanced panel, but the contribution of the x_{it} variables differs.

With respect to trends in IRHI, the mechanics of our decomposition prove somewhat more complex. We present the results for the decomposition of the difference between the concentration indices of 2001 and 1994 in Table 4.2.

The term "elas_direct" is an estimate of Eq. (2.12). It measures the direct effect of proportional income growth on IRHI, i.e. the effect that runs via the income elasticity. Since we found positive income growth between period 1 and t, and an elasticity that increases with income in the previous section, one would expect a positive effect on IRHI. It is found that Eq. (2.12) is positive in all periods and countries where there was positive income growth, except for Austria, Germany and Ireland where the effects are not statistically significant. The negative signs for Austria and Ireland are easily explained. First, the increase of the income elasticity with income was not significant for Austria, while the decrease of the elasticity at the higher income deciles is statistically significant for Ireland (see Section 4.1). Overall, these findings imply that without proportional income growth (*and* even in the absence of changes in income inequality), the direct effect would be to find smaller IRHI, except in Austria, Germany and Ireland.

The second term for which we were able to derive the sign of the effect theoretically is Eq. (2.16), i.e. the direct effect of a change in income inequality. Since the income elasticities were found to be increasing with income, only in case of a pro-rich change in income

[22] It is worth mentioning that for our bounded health utility variable (range [0,1]) the bounds of the concentration index have some relationship with the mean health utility level: the higher the mean, the lower the bounds of the concentration index (see e.g. Wagstaff (2005) and Erreygers (2008)). Using either the Wagstaff (2005) or the Erreygers (2008) normalisation, we found that this 'predetermined' relationship only explains a minor fraction of the observed increase in IRHI.

inequality – a condition only found for Finland – we can *a priori* derive that IRHI will increase. With opposite configurations of income inequality and the elasticity, offsetting effects will occur and it is an empirical question which of the two dominates. Only for Finland and Denmark (even though the trend in income inequality is not statistically significant in Denmark), the configurations reinforce each other, i.e. ineq_direct is positive and significant when the Gini becomes larger.[23] In the majority of countries, a decreasing Gini has a negative effect on IRHI, which implies that the direct effect of income inequality is more important than the direct offsetting effect through the income elasticity (i.e. IRHI would have been higher without the concurrent change in income inequality). For some countries both effects seem to cancel out, i.e. Belgium, Greece and Ireland have an insignificant estimate for ineq_direct. We thus conclude that for the time period considered, income inequality has evolved in the same direction as IRHI (although there are some exceptions) and that only in the case of Finland (and Denmark) we observe increasing health inequalities due to the combination of growing pro-rich income inequality and an elasticity that increases with income. In the majority of countries, opposite effects from proportional income growth and decreasing income inequalities are observed.

The term "age_direct" estimates Eq. (2.19), i.e. the effect of ageing through changes in the age structure on IRHI (through the effect of health on the income elasticity). In general, this can be considered as a 'control term' measuring the impact of changes in 'other variables. It is of specific interest here as it measures the 'direct' impact of ageing (in a single cohort). Obviously the 'direct' effect of ageing is to *increase* IRHI as it increases the income elasticity through a reduced mean health.

Next, we discuss the term "elas_indirect" (cf. Eq. (2.13)), i.e. the indirect effect of proportional income growth on the elasticities of the age–sex structure. The effect runs through the impact of income growth on mean health which affects these elasticities. We could not determine *a priori* the sign of these terms, but find that they all have a very low contribution (in most cases smaller than 0.00000) and are in many cases statistically insignificant.

The term "ineq_indirect" (cf. Eq. (2.17)) summarizes the indirect effect of changes in income inequality. There is an indirect effect on the elasticities of the age–sex dummies (through mean health) and an indirect effect through reranking on the concentration indices of the same variables. Again, since the sign of the effect could not be derived theoretically, this is an empirical question. We observe in most cases a positive sign, although the effect is not significant for Germany and the Netherlands. Its interpretation is complicated since it is determined by the reranking *and* the effect on the elasticities.[24] From more detailed analyses where we decomposed the difference in concentration indices between all years and 1994, it seemed that the effect increased over time, which is not surprising as reranking (and the change in elasticities) becomes a more important phenomenon when a longer period is considered.

Finally, the term "age_indirect" (cf. Eq. (2.20)) summarizes the effect of changes in the age–sex dummies on the elasticities (both through the mean of the dummies and mean health) and on the concentration indices of these variables. Again, this term is only of limited interest since it enters the decomposition as a "control term". The term is insignificant for all countries, except the Netherlands.

Table 4.2 also shows the sum of all ineq, elas and age terms. These are interesting since they reflect the sum of the direct *and* indirect effects of respectively income inequality, proportional income growth, and 'ageing'. In the case of opposite indirect and direct effects, they indicate which effects dominate.

First, for all countries, except Denmark, the total effect of 8 years of ageing is insignificant. While the direct effect of ageing was to increase the income elasticity of health, the combination of direct and indirect effects is no longer statistically significant. We may conclude that IRHI is hardly influenced by 8 years of ageing in these adult cohorts. An alternative interpretation is to consider the term 'age' as a control term that includes age- and sex-related reporting heterogeneity of self-assessed health. Whichever of the two explanations chosen, we repeat that in our empirical example the sum of the 'ageing' terms captures the majority of the differences between the evolution of IRHI in a balanced and an unbalanced panel. Consequently, the results for the other terms on income inequality and proportional income growth in Table 4.2 are hardly affected by using an unbalanced panel, and are thus of relevance beyond the cohort we considered.

Second, "elas" is in most cases almost identical to elas_direct, since the effect of proportional income growth on the elasticity of the other variables is almost negligible. This means that positive income growth adds in all countries, except Austria, Germany and Ireland, to an increase in IRHI, even without any changes in income inequality. This is the most striking finding of our empirical analysis.

Third, the sign of the total effect of income inequality "ineq" is positive, except for Germany and the Netherlands. This implies that, despite decreasing or stable income inequality, the overall effect has been an increase in IRHI. In other words, the indirect effects seem to dominate the effect of income inequality. This may not be so surprising given that this term absorbs most of the effects of income reranking. Note that for the effect of income growth, the direct effect was most important.

5. Conclusion and discussion

This paper set out to try and answer the question to what extent the twin goals of income growth and reduction of social inequalities – such as formulated in the Lisbon EU strategic goals – are compatible, both on theoretical grounds and empirically. In particular, we concentrated on the consequences of changes in the income distribution for changes in the distribution of health by income. We developed a decomposition technique to analyze the role of changes in (proportional) income growth and income inequality in explaining expected trends in income-related health inequality. The decomposition was then applied to the empirical analysis of these trends in 13 European countries using 8 waves of European panel data.

The theoretical model indicates that – when the relationship between income and health is concave – proportional income growth increases mean health and rising pro-rich income inequality with constant mean income reduces mean health. For trends in IRHI, it is more difficult to isolate the role of income growth and income inequality. Our solution was to introduce two hypothetical health levels.

[23] For two countries (Belgium and the UK) the Gini also increases (although not in a statistically significant sense), but without being overall 'pro-rich', i.e. the changes in the income distribution occur mainly among the rich or among the poor. As can be seen from Eq. (2.18), overall pro-richness is needed for triggering increasing IRHI.

[24] Given that γX_t is negative in our specification, this term can only be positive if $C(x_{it}|y_{it})/H_t < C(x_{it}|y_{i1})/H_t^{pg}$. If income inequality decreases, one would expect $H_t^{pg} < H_t$ (due to the positive, but decreasing with income, income profile). It follows that the required inequality always arises if $C(x_{it}|y_{it}) \leq C(x_{it}|y_{i1})$, i.e. if reranking does not increase income-related inequalities in the other determinants.

Using this method, we found that income growth and income inequality have a direct and indirect effect on IRHI. The sign of the direct effects can be derived *a priori*, but not of the indirect effects. Building on Contoyannis and Forster (1999), we showed that the expected *direct* effect of proportional income growth depends crucially on the slope of the income elasticity. If this elasticity is rising/decreasing with income 'on average', IRHI will increase/decline. With respect to the *direct* effect of changes in income inequality, we find increasing/decreasing IRHI in case of 'on average' pro-rich/pro-poor evolving income inequality in combination with an income elasticity that increases/decreases with income 'on average'. The signs of the indirect effects (both for proportional income growth and income inequality) could not be inferred from our decomposition technique and therefore remain essentially empirical questions.

In the empirical analysis, we first examined how estimates of the income elasticity of health varied with income since it is an important determinant of the consequences of income growth and income inequality. We found that in all countries, the marginal effect of income on health is positive and *decreasing* with income. In other words, the income–health relationship is concave, as expected. But more importantly, the income elasticity of health was found to be *rising* with income ('on average').

Second, we presented trends on income growth, income inequality, mean health and IRHI for a single cohort. While between 1994 and 2001, all European countries were found to have experienced real income growth, in most countries this growth was not equally distributed by income level. In all EU countries, income growth was found to be either pro-poor or equally distributed, with one exception: only Finland experienced a clear pro-rich growth. Given this combination of income elasticities rising with income and mostly pro-poor growth, no clear pattern of the impact of income inequality upon income-related health inequality could be derived *a priori*. Only for Finland, a steady rise in the concentration index of health could be anticipated. We also presented evidence on the changing distribution of health. Since we analyzed a cohort of individuals, mean health deteriorated over time, while IRHI increased.

Third, in order to clarify the role and quantify the contribution of proportional income growth and income inequality in the evolution of mean health and IRHI, we used our decomposition technique to disentangle direct from indirect effects. We concluded that proportional income growth leads to better mean health, and that this is true *a fortiori* when simultaneously income inequality is falling. So economic growth coupled with reduced income inequality is good for mean health levels. However, both the direct and indirect effects of proportional income growth were found to increase IRHI, a rather striking result. On the other hand, the direct effect of pro-poor changes in income inequality in Europe (with a few exceptions) has reduced IRHI, while its indirect effect has increased IRHI. Finally, the model also allows for an analysis of the direct and indirect effects of ageing but we found that the 8 years of ageing in our cohort hardly affected IRHI.

So can the twin goals of economic growth and greater social (health) cohesion in the European Union be reconciled? Our analysis suggests that there may be a problem and that Wagstaff's (2002b) hypothesis that developing countries are "swimming against the tide?" by trying to couple growth with a reduction in relative inequalities may similarly apply to high-income economies like the European. Given the universal observation that everywhere in the 'old' European Union the income elasticity of health rises with income, attaining both goals proves to be difficult. It means that then even proportional growth, which leaves income inequality unchanged, will lead to greater IRHI. *A fortiori*, the same applies when countries' income inequality rises. Therefore, if European countries had not managed to lower or stabilize their income inequality in the 1990s, then more countries than only Finland would have experienced sharp increases in socioeconomic inequalities in health, simply as a result of the economic growth.

It is clear that an evaluation of the overall welfare implications of improved mean health coupled with rising relative IRHI depends on the relative weight given to improvements in the mean versus the distribution. This trade-off can be made more explicit by using social welfare type functions (see e.g. Wagstaff, 2002a) but the results crucially depend on the prevailing degree and type of societal aversion to health inequality. Little is known on this empirically and more evidence is needed before such measures can usefully be applied to a welfare analysis of health trends.

Appendix A. Health equation estimates

	Austria	Belgium	Denmark	Finland	France	Germany	Greece	Italy	Ireland	Netherlands	Portugal	Spain	UK
M3044	−0.018**	−0.011**	−0.016**	−0.019**	−0.018**	−0.030**	−0.009**	−0.023**	−0.010**	−0.012**	−0.022**	−0.018**	−0.013**
M4559	−0.060**	−0.027**	−0.037**	−0.063**	−0.046**	−0.079**	−0.038**	−0.055**	−0.026**	−0.027**	−0.067**	−0.047**	−0.027**
M6069	−0.078**	−0.040**	−0.052**	−0.087**	−0.069**	−0.103**	−0.087**	−0.090**	−0.045**	−0.044**	−0.120**	−0.087**	−0.041**
M70+	−0.109**	−0.061**	−0.076**	−0.098**	−0.093**	−0.125**	−0.136**	−0.144**	−0.053**	−0.051**	−0.163**	−0.100**	−0.039**
F1629	−0,002	−0,004	−0.007*	−0.007*	−0,002	−0.008*	0.005*	−0.003+	−0,001	−0.008*	0.001	−0,001	−0.011**
F3044	−0.017**	−0.019**	−0.023**	−0.022**	−0.025**	−0.035**	−0.012**	−0.030**	−0.011**	−0.021**	−0.032**	−0.020**	−0.019**
F4559	−0.057**	−0.038**	−0.055**	−0.061**	−0.056**	−0.090**	−0.051**	−0.071**	−0.021**	−0.038**	−0.095**	−0.065**	−0.041**
F6069	−0.089**	−0.052**	−0.062**	−0.080**	−0.079**	−0.112**	−0.109**	−0.116**	−0.037**	−0.053**	−0.156**	−0.114**	−0.032**
F70+	−0.127**	−0.070**	−0.092**	−0.120**	−0.105**	−0.148**	−0.151**	−0.170**	−0.057**	−0.070**	−0.196**	−0.134**	−0.058**
(eqinc/10000)	0.042**	0.011**	0.025**	0.027**	0.031**	0.032**	0.050**	0.024**	0.047**	0.016**	0.069**	0.040**	0.030**
(eqinc/10000)ˆ2	−0.010**	−0.001**	−0.002**	−0.005**	−0.004**	−0.005**	−0.014**	−0.003**	−0.014**	−0.002**	−0.015**	−0.008**	−0.004**
(eqinc/10000)ˆ3	0.001**	0.000**		0.000**	0.000**	0.000**	0.001**	0.000**	0.002**	0.000**	0.001**	0.001**	0.000**
(eqinc/10000)ˆ4	−0.000*	−0.000**			−0.000**		−0.000**		−0.000**			−0.000**	−0.000**
(eqinc/10000)ˆ5									0.000**				
constant	0.907**	0.920**	0.92**	0.913**	0.880**	0.870**	0.929**	0.911**	0.914**	0.914**	0.861**	0.902**	0.879**
σ_{ε}^{2}	0.079**	0.064**	0.078**	0.068**	0.105**	0.114**	0.094**	0.086**	0.060**	0.063**	0.114**	0.090**	0.098**
Observations	27,769	24,200	20,352	19,314	54,688	60,160	49,072	72,288	22,976	36,448	56,776	58,456	47,400

Notes: +: significant at 10%; *: significant at 5%; **: significant at 1%. Statistical inference is based on a robust covariance matrix that allows for clustering at the level of the individual.

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