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ORIGINAL ARTICLE

On age-specific variations in income-related inequalities in diabetes, hypertension and obesity

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Abstract
Objective This article examines whether the social health gradients in diabetes, hypertension and obesity for men and women vary significantly across different age groups.
Methods We use a pooled sample of German survey data from the years 2002 and 2006 with a total of 87,601 observations. We employ a varying Wagstaff index derived from the class of Gini-type concentration indices to estimate age-specific income-related health inequalities.
Results We find significant health disadvantages among poor women in mid-age, but no significant age-specific income-related health inequalities among men. Some leveling of inequalities in diabetes is observed.
Conclusions The results suggest that variations in age-specific inequalities are unlikely to be a purely artificial result of health-related selection into retirement or mortality.

Keywords Germany · Life course perspective · Age-specific inequality · Obesity · Hypertension · Diabetes

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Introduction

With the introduction of the health concentration index (Wagstaff et al. 1991), socioeconomic gradients in the distribution of health became an important research field in health economics (see e.g. van Doorslaer et al. 2004, 2006; Kakwani et al. 1997; McKinnon et al. 2011). Banks et al. (2007), Case and Deaton (2005) and van Kippersluis et al. (2010) point out that considering the life course perspective may be important for a better understanding of the origins of disparities in health.

The relation of low socioeconomic status and less healthy lifestyles, higher health risks and increased rates of premature mortality is well documented in the literature (Balia and Jones 2008; Case and Deaton 2005; Deaton and Paxson 1998; Kim and Durden 2007; Lynch 2003; Prus 2004). The two currently predominant explanations for variations in socioeconomic health gradients over the life course are the disadvantage accumulation hypothesis and the age as leveler hypothesis. The former states that social health gradients develop in early life and emerge over the complete life course through continual accumulation of health disadvantages among the deprived (Kim and Durden 2007; van Kippersluis et al. 2010; Lynch 2003; Prus 2004; Ross and Wu 1996). The age as leveler hypothesis, on the other hand, has long been considered as an antithesis (Dupre 2007). It is, however, rather an extension adding the assumption that the decline of health (being an inevitable part of aging) may outperform the accumulation of health disadvantages around mid-age (Case and Deaton 2005; Deaton and Paxson 1998; Kim and Miech 2009; Kunst and Mackenbach 1994). According to van Kippersluis et al. (2010), evidence for the age as leveler hypothesis is generally consistent with an accumulation process up to some point, where age begins acting as a leveler in later life.

Associating the accumulation process to the individual and leveling through age to the aggregate level, Dupre (2007) has shown that both mechanisms are by no means contradictory.

A growing body of literature analyzes the health status over the life course for distinct socioeconomic groups. van Kippersluis et al. (2010), for instance, use cross-sectional data from the Netherlands and compare the development of self-reported health, disability and mortality for distinct educational groups. Mirowsky and Ross (2005) investigate differences in the predicted health status between educational levels in the United States. Both found that health status declines faster among the less educated. By using data from 11 European countries, van Kippersluis et al. (2009) found a continual disadvantage accumulation for most countries; however, they discovered some leveling for France, Germany and the UK.

To our knowledge, no work has yet been done on the life course perspective in income-related inequalities of specific diseases. The aim of this paper is to highlight changes of the social gradients in three health outcomes across different age groups. Using German survey data, we investigate age-specific variations in income-related health inequalities in hypertension and diabetes. The risk of hypertension and diabetes is negatively correlated with socioeconomic status in developed countries and both are considered as avoidable through lifestyle and health behavior (Harati et al. 2010; Puska 2010). We also analyze obesity as it increases the risk of hypertension and diabetes considerably (Haffner 2006; WHO 2003). To measure age-specific income-related health inequalities, we use a semiparametric extension of the concentration index proposed by Siegel and Mosler (2010).

Methods

Data

To assure sufficiently large numbers of observations particularly in the oldest age groups, we pooled data from the 2002 and 2006 waves of the Healthcare Access Panel (HCAP) for our empirical analysis. The data were collected by TNS Healthcare, Munich (now Kantar Health) for commercial purposes. For each wave of this voluntary mail survey, ∼50,000 households were drawn randomly from a large database. In-field time was 6 weeks without active reminding and the response rates were ∼60 % in both waves. Potthoff et al. (2004) provide an extensive discussion on the 2002 wave of the HCAP showing that the underlying sampling procedure yields representative samples for the German population.

The pooled sample comprises 117,167 individuals in 48,574 households. 75,122 individuals in 29,421 households were included in 2002 and 60,555 individuals in 28,828 households in 2006. 18,510 individuals in 8,718 households participated in both surveys. We removed 28,390 individuals younger than 20 as chronic conditions rarely affect younger individuals and a meaningful interpretation of the body mass index (BMI) is problematic for children and adolescents. We dropped 1,176 observations owing to missing data on income where zero household income was also considered as nonresponse. The sample for our empirical analysis eventually comprises 87,601 individuals (45,889 women and 41,712 men) in 47,867 households.

We measure the social status using the monthly net equivalent household income. In the HCAP, individuals were asked to specify the sum of all incomes available to the household (e.g. from labor, capital, pensions or welfare benefits). We employ the modified OECD equivalence scale to compute net equivalent household income (see e.g. van Doorslaer et al. 2004; van Kippersluis et al. 2009). Benefits from the public pension scheme depend largely on incomes over the life course (Boersch-Supan and Wilke 2004) and German welfare policies are considered as highly status preserving (Brockmann et al. 2009). Disparities in the income distribution and the relative socioeconomic position are therefore likely to persist after retirement. Accordingly, we consider household income as a good measure of the social position before and after retirement.

The BMI is computed from self-reported anthropometric data as body weight in kilograms divided by the squared body height in meters, kg/m^2. We follow the standard WHO classification (WHO 2009) and define obesity as BMI over 30. The second health outcome is hypertension. Individuals were asked to indicate whether they had hypertension during the preceding 12 months. The third health outcome is diabetes in the preceding 12 months; however, the survey does not allow us to distinguish between type 1 and type 2 diabetes. Type 2 diabetes is agreed to be age-related and influenceable through lifestyles and health behaviors (see e.g. Harati et al. 2010; Puska 2010). We analyze diabetes irrespective of insulin dependency, because we consider type 1 diabetes as mainly genetic and therefore equally distributed across socioeconomic groups.

Econometric model

We measure age-specific income-related health inequalities using a semiparametric extension of Wagstaff's (2005) corrected concentration index. Concentration indices are derived from concentration curves where the cumulative share of health is compared with the cumulative share of the population ranked by income (Wagstaff et al. 1991). If no income-related inequality is present, the concentration curve coincides with the line of equality (45° line). Technically, the concentration index measures twice the area between the concentration curve and the line of equality. It is bound in the $(-1;\,1)$ interval and becomes positive (negative), if the variable of interest concentrates among the better off (worse off). The index equals zero if no income-related inequality is observed. Note that the concentration index is a bivariate extension of the Gini index; using income in place of health as outcome variable would yield the well-known Lorenz curve and Gini index. Konings et al. (2010) provide an intuitive introduction to the concept of Gini-type concentration indices.

We follow Siegel and Mosler (2010) to obtain the concentration index as a smooth function of some regressor z (i.e. age). They adapt the convenient regression approach (Kakwani et al. 1997) and propose to estimate

$$2\frac{\sigma_r^2(z)}{\mu_y(z)}y_i = \beta_0(z) + \beta_1(z)r_i(z) + \epsilon_i, \qquad (1)$$

using the varying smooth coefficient approach (Hastie and Tibshirani 1993; Li et al. 2002). Individuals $i = 1, \ldots, N$ must be sorted by income in ascending order and $r(z)$ denotes the local fractional rank. The coefficient $\beta_1(z)$ is the varying concentration index $C(z)$ and y is the dependent variable with local mean $\mu_y(z)$.

As proper sample weights are not available for the pooled data, the formula for the fractional rank presented by Siegel and Mosler (2010) simplifies to

$$r_i(z) = \left(\sum_{j=1}^{i} k_{h_z}(u_j) - \frac{k_{h_z}(u_i)}{2}\right)\mathcal{I}(|u_i| \le h_z) \qquad (2)$$

with $u_i = z_i - z$. $\mathcal{I}(|u_i| \le h_z)$ is an indicator function being 1 if $|u_i| \le h_z$ and 0 otherwise. Using the kernel weights $k_{h_z}(u_i)$ is important here to assure that the local estimates for the mean and variance of the local rank variable $r(z)$ equal their theoretical asymptotic values 0.5 and 1/12 throughout the support of z. We use a Nadaraya–Watson estimator with a quartic kernel. This kernel function assigns higher weights to observations closer to z, lower weights for observations being further away and zero weight to observations outside the bandwidth h_z. The kernel weights are $k_{h_z}(u_i) = \left(\sum_{j=1}^{N} K_{h_z}(u_j)\right)^{-1} K_{h_z}(u_i)$ with $K_{h_z}(u_i) = (1/h_z)K(u_i/h_z)$ and $\sum_{i=1}^{N} k_{h_z}(u_i) = 1$. The local bandwidth parameter h_z is chosen inversely to the local kernel density $f(z)$ owing to the tradeoff between estimation bias and uncertainty. The kernel density is the nonparametric estimate of the probability density function (pdf) of z.

Comparing concentration indices of binary variables with varying means may be misleading as the bounds then depend inversely on the mean μ, $|C|_{\max} = 1 - \mu$ (Wagstaff 2005, 2011). We follow Siegel and Mosler (2010) and adapt Wagstaff's (2005) approach such that $W(z) = C(z)/(1 - \mu(z))$ is our varying Wagstaff index. Similar to its homogeneous counterpart, $W(z)$ is always bound in the $(-1;\,1)$ interval irrespective of $\mu(z)$. This allows comparisons of inequalities throughout the support of z. We compute the varying Wagstaff indices separately for men and women. We report pointwise confidence intervals (for computational details see Appendix).

Results

The graphs for the estimated density $\hat{f}(z)$ with respect to age (the smoother or smoothing variable z) in Fig. 1 exhibit similar patterns for both sexes. Cohorts younger than 35 are smaller than those born before the 1970s. This corresponds with the population pyramid for Germany and supports existing evidence for an aging society (see e.g. von Weizsäcker 1996). Note that the observed kernel density $\hat{f}(z)$ implies that the bandwidth parameter h_z is minimal at age 33 for women ($h_z \approx 5.95$) and at 39 for men ($h_z \approx 6.22$). The bandwidth is largest for the 79 years old for both sexes ($h_z \approx 16.6$ for women and ≈ 16.9 for men).

The overall sample prevalence of obesity in Table 1 is higher among women than among men. Comparing the age-specific estimates in Fig. 1, we find that this holds in all age groups. The prevalence is highest among the 56-year-old women and 58-year-old men and lower in younger and older age groups. The overall sample prevalence of hypertension is higher among men. The shape of the age-specific prevalence curve in Fig. 1 suggests that prevalence rises almost monotonously with increasing age and flattens after age 70 for both sexes. It is highest among the 76-year-old women and 77-year-old men. The overall sample prevalence of diabetes in Table 1 is again somewhat higher for men. The age-specific estimates vary around 1 % being somewhat higher for women until the late thirties. The prevalence then rises faster among men, is highest at age 69 ($\approx$12 %) and varies around this value for the oldest. For women, it is highest at age 67 ($\approx$10 %) and decreases to $\sim$7 % for the oldest.

The age-specific mean net equivalent household income in Fig. 2 yields similar patterns for men and women. The graphs suggest that the youngest individuals live in households with the lowest incomes. We find a first peak around the age of 30 and the highest mean income shortly before the statutory retirement age of 65. The lower mean equivalent household incomes between 35 and 50 come from an increase in the equivalence weights owing to a

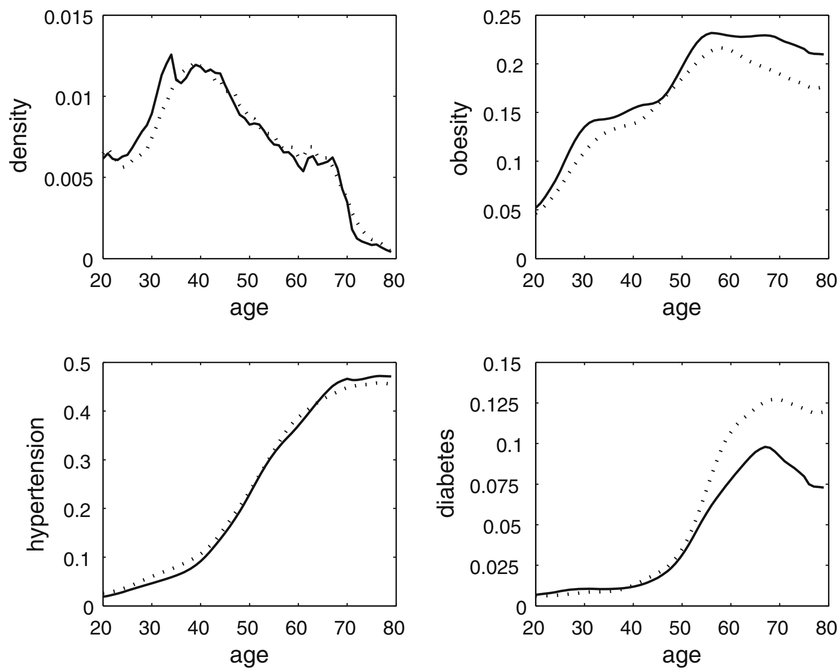


Fig. 1 Kernel density and age-specific prevalences. The kernel density estimate of age (*upper left*) and the age-specific prevalences of obesity (*upper right*), hypertension (*bottom left*) and diabetes (*bottom right*) for the female (*solid*) and male (*dotted*) sample (TNS Healthcare Access Panel; Germany 2002, 2006)

Table 1 Homogeneous (overall sample) estimates

	Female ($N = 45{,}889$)			Male ($N = 41{,}712$)		
	Prevalence	$\hat{W}$	$\hat{\sigma}_W$	Prevalence	$\hat{W}$	$\hat{\sigma}_W$
Income	1394.9[a]	0.2971**,[b]	0.0032[c]	1460.9[a]	0.2994**,[b]	0.0035[c]
Obesity	16.57 %	−0.1283**	0.0126	15.19 %	−0.0543**	0.0137
Hypertension	17.73 %	−0.0623**	0.0122	19.54 %	−0.0114	0.0123
Diabetes	3.14 %	−0.1283**	0.0273	4.15 %	−0.0535*	0.0246

The homogeneous (overal sample) estimates regardless of age

* Significant at the 5 % level

** Significant at the 1 % level

[a] Mean of net equivalent household income

[b] Gini index of net equivalent household income (without Wagstaff's correction)

[c] Standard error of the Gini index (TNS Healthcare Access Panel; Germany 2002, 2006)

peak in the average number of children per household in these age groups. In all age groups, men have a higher mean income than women. The age-specific Gini index in Fig. 2 is somewhat lower for women and men around 30 than for the youngest. After the age of 30, income inequality increases and is highest for women around 50 and men around 60. The Gini index drops back to ~0.27 for both the sexes during the common retirement period. Note that the statutory retirement age is 65, however, most people retire in the preceding 7 years (see Wingerter 2010). Although our results suggest that retirement leads to some leveling in household incomes, we find that considerable income inequalities persist after retirement.

The homogeneous Wagstaff indices W for obesity in Table 1 are both significantly negative suggesting a concentration of the disease among the worse off for both sexes. The varying Wagstaff index for men in Fig. 3 is negative but statistically significant for no age. The significantly negative index for women aged between 28 and 72 indicates a concentration among the worse off in these age groups. While the varying index for men shows no considerable changes over the life course, women exhibit a lower level of inequality in mid-age and a higher level among those over 55.

The homogeneous Wagstaff index for hypertension among women in Table 1 is again highly significant. The varying Wagstaff index in Fig. 4 is statistically significant for those between 50 and 69 in the female sample. For all age groups, the index varies around −0.1. The varying index for men is negative for those aged 23 or older. The

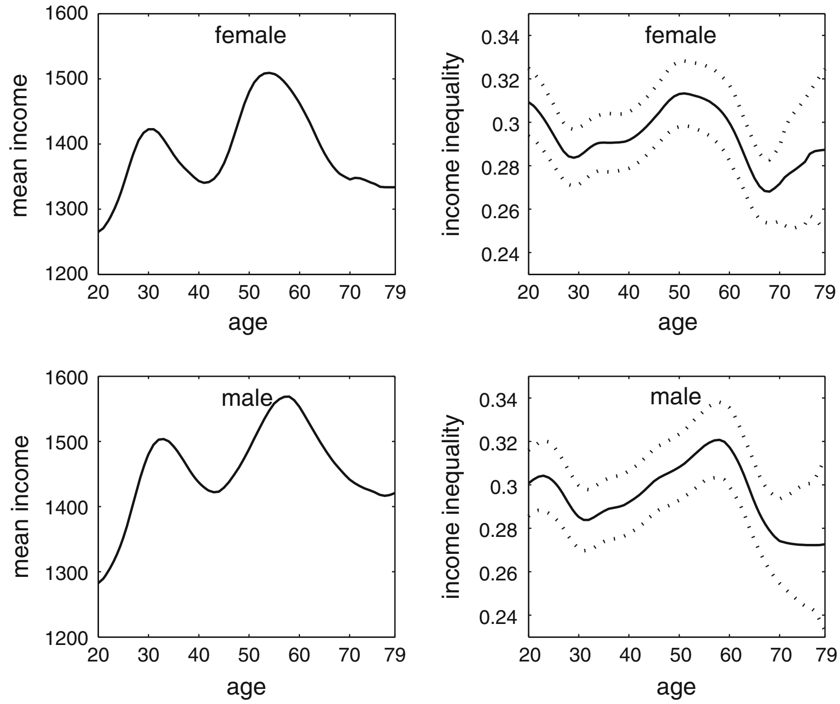


Fig. 2 Age-specific income and income inequality. The age-specific mean net equivalent household incomes (*left*) and income inequality as Gini index (*right*) with 95 % confidence bands (*dotted lines*) for the female (*upper*) and male (*bottom*) samples (TNS Healthcare Access Panel; Germany 2002, 2006)

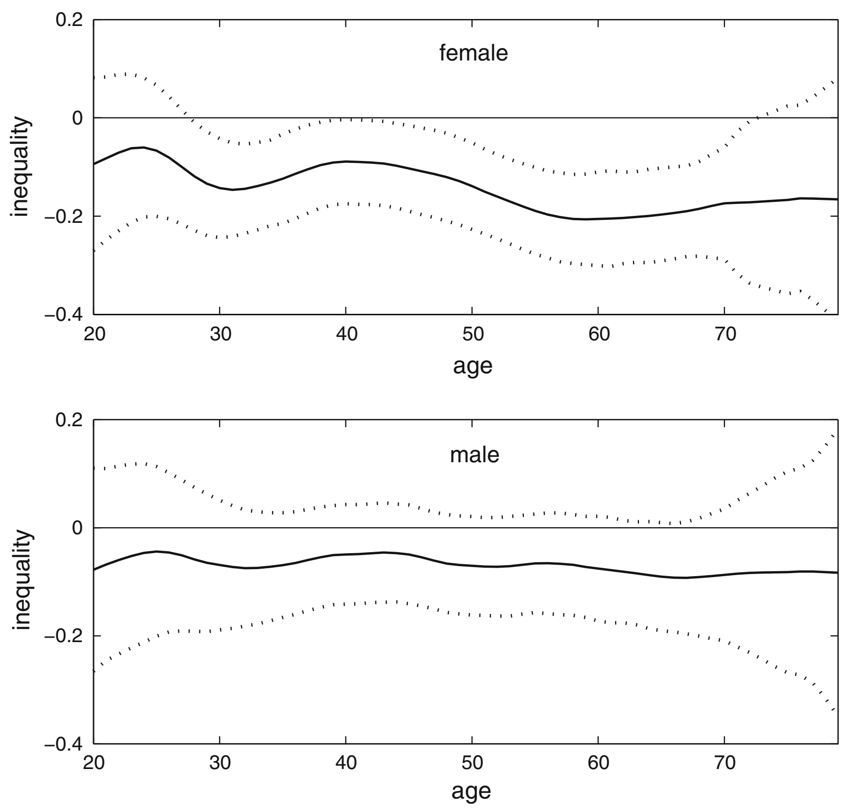


Fig. 3 Age-specific inequality of obesity. The age-specific varying Wagstaff indices (*solid lines*) with 95 % confidence bands (*dotted lines*) for women (*upper graph*) and men (*bottom graph*) (TNS Healthcare Access Panel; Germany 2002, 2006)

homogeneous and varying Wagstaff indices for the male sample are both insignificant.

The homogeneous Wagstaff indices for diabetes in Table 1 are significantly negative for women and men, again suggesting a concentration of the disease among the worse off. The varying Wagstaff index for the female sample in Fig. 5 is negative for all age groups. It is, however, only significant for women aged between 51 and 62. We find no significant age-specific income-related gradient in diabetes in the male sample.

Fig. 4 Age-specific inequality of hypertension. The age-specific varying Wagstaff indices (*solid lines*) with 95 % confidence bands (*dotted lines*) for women (*upper graph*) and men (*bottom graph*) (TNS Healthcare Access Panel; Germany 2002, 2006)

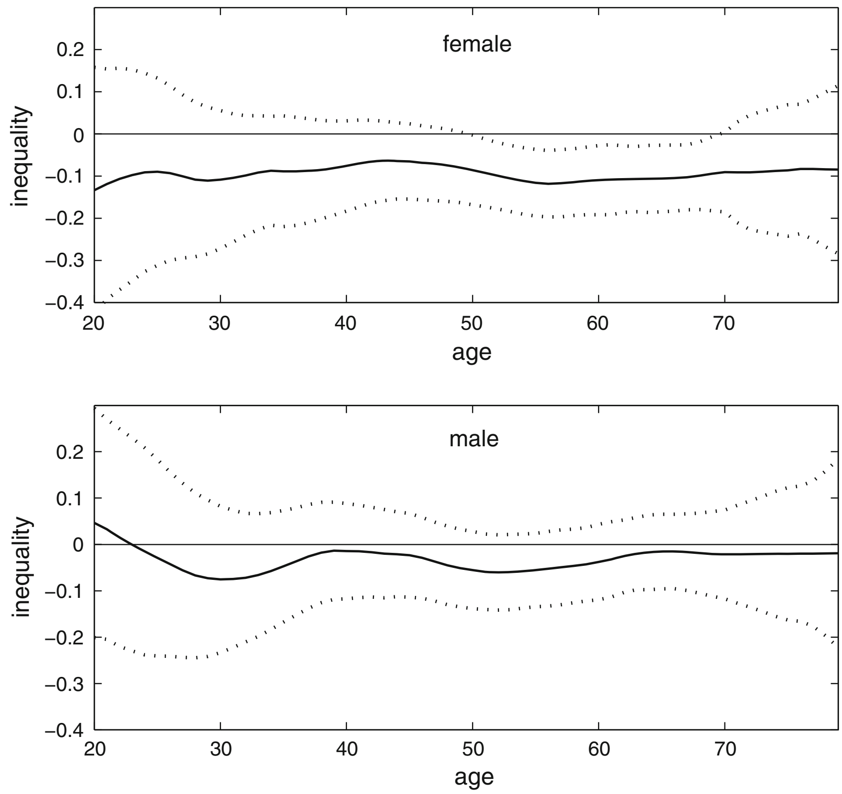
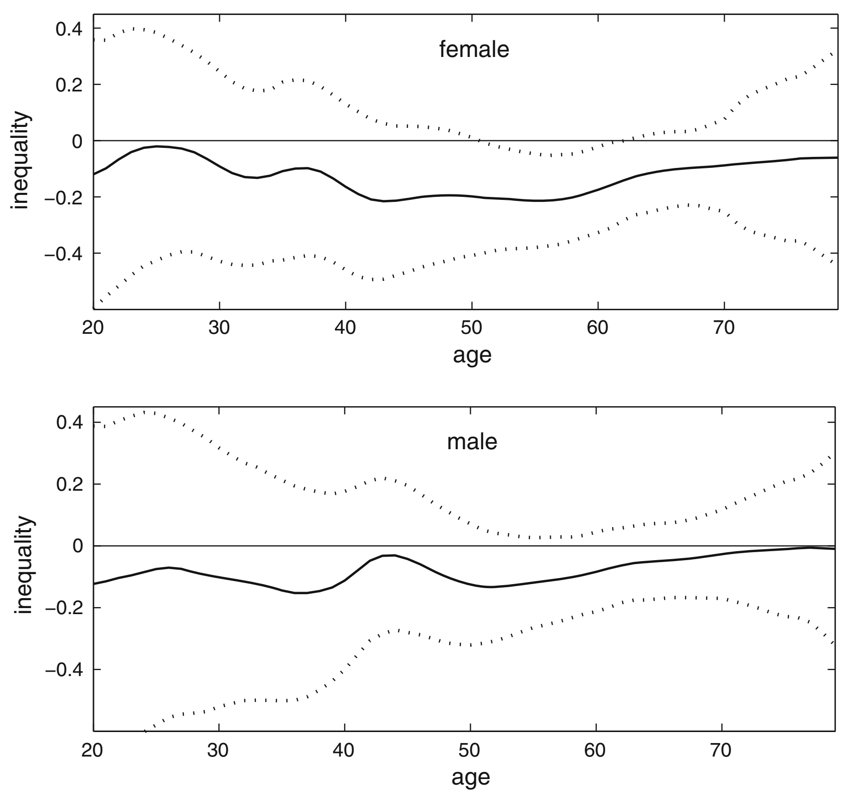


Fig. 5 Age-specific inequality of diabetes. The age-specific varying Wagstaff indices (*solid lines*) with 95 % confidence bands (*dotted lines*) for women (*upper graph*) and men (*bottom graph*) (TNS Healthcare Access Panel; Germany 2002, 2006)



Discussion

Using a varying inequality index recently introduced by Siegel and Mosler (2010), this article addressed age- and sex-specific variations in income-related health inequalities. One may consider computing batteries of concentration indices for distinct age groups (as e.g. in van Kippersluis et al. 2009) and smoothing over these as an alternative. Siegel and Mosler (2010), however, point out that this would involve double uncertainty and a considerable bias. The authors argue that first, particularly, the oldest age groups may become rather small which would involve high uncertainty and large confidence intervals. Second, smoothing over such estimates would add its own uncertainty. In addition, the estimates close to the age group limits would likely be biased which is avoided in the kernel smoothing approach.

We were able to show that socioeconomic gradients in obesity, hypertension and diabetes vary across age groups in Germany. Significant age-specific inequalities to the detriment of the deprived in hypertension and diabetes are only observed for women in midlife. The age-specific gradient for obesity is significant for almost all age groups in the female sample. Surprisingly, we find no significant age-specific income-related inequalities among men. Although we find some leveling for diabetes among both sexes and hypertension among men, one may agree that our results are neither a clear support for the accumulation nor for the age as leveler hypothesis.

Using the varying index has some advantages compared with computing homogeneous concentration or Wagstaff indices. Using the latter may invoke what Islam et al. (2010) referred to as student and pension effects. According to the former, young people in the beginning of their career, in job trainings or universities are usually in good health conditions but have low incomes. This may result in an underestimation of health disadvantages for lower income households. Conversely, the pension effect describes the notion that older people have comparably bad health conditions due to their age and lower incomes after retirement. This, again, may lead to an overestimation of health disadvantages for the poor. The homogeneous indices actually underestimate the extent of inequality for most age groups in hypertension and obesity compared with the age-specific results.

When considering self-reported data on specific diseases, one may argue that certain diseases have to be diagnosed by a physician. Individual awareness may therefore, to some extent, depend on health care utilization and communication between patients and physicians. However, $\sim$90 % of the German population contact a physician within a year and, more importantly, Germany is known for a fairly equitable access to health care (see e.g. van Doorslaer et al. 2004, 2006). Considerable reporting biases owing to inequalities in health care utilization are hence rather unlikely. The potential of biases owing to social distances between physicians and less educated or lower income patients, however, remains. Considering the results found in Kelly-Irving et al. (2011), one may speculate that such biases may lead to a pro-rich bias, i.e. an underestimation of inequalities to the detriment of the deprived. Concerning the results for obesity, it should be mentioned that self-reported anthropometric data may involve some measurement or reporting bias, which may lead to an underestimation of the prevalence and socioeconomic gradient of obesity.

Using cross-sectional data involves some limitations. First, they do not allow us to fully distinguish between life course effects and cohort effects. However, van Kippersluis et al. (2009) found no consistent cohort effects for Germany. Although we cannot fully exclude cohort effects, we assume that these play a minor role and the observed variations to be mostly an age effect. Second, the data do not allow us to distinguish between income effects on health and health effects on income or to test the causal relationship between then. We assume possible losses of income through adverse health-related selection into both early old age and reduced earning capacity retirement to be of minor importance (Brockmann et al. 2009). Only 1.5 and 1.6 % of women and men among the 45–50 years old in the sample are retired. The shares are somewhat higher among the 50- to 55-year olds, 2.8 and 4.3 %, respectively. Further, German welfare policies are seen as rather status preserving (Brockmann et al. 2009). Pensions largely depend on compulsory contributions over the life course (Boersch-Supan and Wilke 2004). The relative socioeconomic position within a birth cohort will therefore change only marginally with retirement. Finally, Dupre (2007) and Prus (2004) argue that leveling may be an artificial result owing to mortality selection. However, Beckett (2000) has shown that this needs not necessarily be true for self-reported health. The age-specific mortality rates did not exceed 0.5 (1) % for women younger than 60 (65) in the respective years in Germany (Human Mortality Database 2011). The rates were about double for men. We therefore reject the notion that the observed variations in income-related inequalities could be solely caused by mortality selection.

Conclusions

We believe that our results are of particular interest for researchers and health policy makers alike, since the health outcomes under consideration are among the major risk factors for cardiovascular diseases and premature mortality in developed countries. In addition, this paper adds an important contribution to the field of age-specific analyses of income-related health inequalities. Given the demographic developments in most industrialized countries as demonstrated by the evidence for an aging German society found in the data, such analyses will likely become an increasingly important tool for health policy makers to maintain an efficient allocation of both preventive and curative health services.

Appendix: Statistical inference

The local standard error $\sigma_C(z)$ of the varying concentration index can be approximated by estimating β_0^* and β_1^* from Eq. (1) with the untransformed health variable y in place of $\left(2\sigma_r^2(z)/\mu_y(z)\right)y$. One may then take advantage of the fact that $\beta_0^*(z) + \mu_r(z)\beta_1^*(z) = \mu_y(z)$ and apply the δ method to $C(z) = 2\sigma_r^2(z)\beta_1^*(z)\left[\beta_0^*(z) + \mu_r(z)\beta_1^*(z)\right]^{-1}$ (Kakwani et al. 1997; Siegel and Mosler 2010; Wildman 2003). The standard error $\sigma_W(z)$ for $W(z)$ can be estimated analogously (Siegel and Mosler 2010). Note that $\mu_r(z)$ and $\sigma_r^2(z)$ need not be considered as stochastic as they are sample independent. The error term may likely be heteroscedastic and autocorrelated (Kakwani et al. 1997; Wildman 2003; McKinnon et al. 2011); we therefore follow Siegel and Mosler (2010) and estimate local Newey-West type standard errors.

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