The changing contribution of socioeconomic deprivation to variance in age at death

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Abstract

Mortality inequalities demonstrate a double burden: the most deprived socioeconomic groups experience the lowest average age of death and the highest variation in age at death. No study has identified how variation in age of death is patterned by area-level socioeconomic deprivation despite the established literature demonstrating that area-level effects are an important part of the explanation for inequalities in risk of death. Two underlying processes drive variation in age at death: individual stochasticity (within-group variance) and heterogeneity (between-group inequality). Limited research has evaluated how these two components have changed over time. We address these research gaps by using population and mortality data for the entire population of Scotland stratified by a validated measure of area-level deprivation that covers the time period 1981-2011.

1 Background

The association between socioeconomic inequality and mortality is traditionally evidenced by life expectancy comparisons. The most deprived populations experience the lowest average age of death, and the least deprived populations experience the highest. Studies have further demonstrated that the most deprived populations also demonstrate the highest level of variation in age at death when measuring socioeconomic inequality by income, education, or occupation (Broennum-Hansen 2017, van Raalte et al. 2011, Sasson 2016, van Raalte et al. 2014). To our knowledge no study has identified how variation in age of death is patterned by area-level socioeconomic inequality even though it has long been established that area of residence is an important part of the explanation for socioeconomic differences in mortality outcomes (Carstairs and Morris 1989, Macintyre et al. 2002, Tunstall et al. 2011). Area-level measures aim to capture relative deprivation rather than absolute poverty. Relative deprivation is the idea that not having enough material cultural or social resources to participate in the accepted way of life in one's societal context is as important for health as the absolute minimum requirements for survival (Kearns et al. 2000, Carstairs and Morris 1989, Townsend 1987) .

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Two processes underly variance in age at death: Individual stochasticity (within-group variance) and heterogeneity (between-group inequality). Within-group variance tends to arise from differences due to random demographic processes. The lifetable assumes that every individual is subject to the same schedule of mortality rates, such that any variance in age of death can be interpreted as individual stochasticity. However, it can also be argued that within-group variance could be due to the lifetable being aggregated over heterogeneous populations. For example males and females experience different social and biological processes that mean they demonstrate different mortality schedules. Aggregating over both sexes would increase within-group variance because you are systematically aggregating over two groups that are known to be heterogeneous and that have age distributions which do not completely overlap. It is possible that a lifetable could be produced for populations that are heterogeneous but without knowing the reasons why the populations are disparate.

Between-group inequality can arise if individuals at the same age are subject to different mortality rates that may be due to unobserved social, economic or environmental contexts (Hartemink et al. 2017).

One existing study of 11 European countries found that the differences between educational groups accounted for between 1.2% and 10.9% of total variance in age at death (van Raalte et al. 2012). This study used aggregated data for the years 1990 to 2000, such that time trends could not be assessed. Education may also be a problematic socioeconomic measure for studying changes in components over time due to compositional changes (Hendi 2015). In addition, stratifying data by education or other acquired states forces the researcher to left-truncate data at some age, such as 35.

We contribute to previous findings by measuring how the within-group and between-group components of total variance in age at death are patterned by sex and area-level deprivation over time. We use decennial Census population estimates and mortality data stratified into population-weighted quintiles according to a validated area-level measure of relative deprivation. The data include the whole population of Scotland (ca 5 million persons) and cover the time period between 1981 and 2011. Using population-weighted quintiles according to an area-level measure of deprivation has two advantages. Firstly, it is applicable to the entire population meaning that age truncation is not necessary. Secondly, it is a relative measure of deprivation, meaning that there is always a notional 20% most deprived compared to a notional 20% least deprived even though absolute levels of deprivation may have changed over time.

2 Data and methods

2.1 Area level deprivation

Census population estimates and mortality data[1] by single year of age and sex for each part-postcode (zip code) sector in Scotland were obtained via a commissioned request to National Records of Scotland. There are around 1,012 part-postcode sectors in Scotland at each Census year each with an average population size of 5,000 individuals.

[1] Mortality data used came from 1980-1982, 1990-1991, 2000-2002 and 2010-2012 to increase the number of events centered around each census. 1990-1991 is just a two-year numerator sample due to geographical boundary changes occurring in 1990. Corresponding Census population estimates are adjusted accordingly.

Population-weighted quintiles (each 20% of the population) were created by aggregating the 1,012 part-postcode sectors according to Carstairs score of deprivation. The Carstairs score is a z-score for each part-postcode sector that is derived from four individual-level census variables: overcrowding, male unemployment, low social class, and no car ownership. The Carstairs Score (z-score) reflects the material resources that provide the means to access the goods, services, amenities, and physical environment seen as expected in society (Carstairs and Morris 1989). This means the Carstairs score is a method of capturing relative deprivation at the contextual level.

Deaths and census population denominators were used to construct complete lifetables for each deprivation quintile, at each Census year, for males and females seperately (40 lifetables in total). The Human Mortality Database protocol was used to extrapolate age specific mortality rates from ages 85 to 110 (Wilmoth et al. 2007).[2] From the complete lifetables we compute remaining life expectancy and the conditional standard deviation of the remaining lifespan distribution. Lifetable standard deviations are a common measure of the variability applied to the distribution of age at death (van Raalte and Caswell 2013). Details of the within-group and between-group component calculations are given in the following subsection.

2.2 Variance decomposition

Lifetable variance is calculated and decomposed into within- and between-group components using Markov chain methods (Caswell 2001)(Caswell 2009)(Caswell 2014). From the lifetable, we extract conditional single-age death probabilities, q_x, and take its complement, p_x. We then calculate the survival matrix for the i^{th} quintile, $\mathbf{U}_i$ as:

$$\mathbf{U}_i = \begin{bmatrix} 0 & \cdots & \cdots & \cdots & 0 \\ p_1 & & & & \vdots \\ 0 & \ddots & & & \vdots \\ \vdots & & \ddots & & 0 \\ 0 & \dots & 0 & p_{\omega-1} & p_\omega \end{bmatrix} \tag{1}$$

Conditional remaining survivorship is calculated as:

$$\mathbf{N}_i = (\mathbf{I} - \mathbf{U}_i)^{-1} \quad . \tag{2}$$

$\mathbf{N}_i$ ends up being 0s in the upper triangle, and conditional remaining survivorship in columns descending from the subdiagonal. The moments of longevity for individuals in group i are $\boldsymbol{\eta}_1^{(i)}$ and $\boldsymbol{\eta}_2^{(i)}$.

$$\boldsymbol{\eta}_1^{(i)} = (1^\intercal \boldsymbol{N}_i)^\intercal \tag{3}$$

The second moment is defined as:

$$\boldsymbol{\eta}_2^{(i)} = [1^\intercal \boldsymbol{N}_i(2\boldsymbol{N}_i - \boldsymbol{I})]^\intercal \tag{4}$$

The vectors with the means and variances, for group i, are

$$E(\boldsymbol{\eta}^{(i)}) = \boldsymbol{\eta}_1^{(i)} \tag{5}$$

$$V(\boldsymbol{\eta}^{(i)}) = \boldsymbol{\eta}_2^{(i)} - \left[\boldsymbol{\eta}_1^{(i)} \circ \boldsymbol{\eta}_1^{(i)}\right] \tag{6}$$

[2]Specifically, we apply equations (53) and (54) from the HMD protocol v5, modified to use information from ages 75+ rather than 80+.

To carry out calculations we procede by creating vectors that contain the combined age and stage specific values

$$E(\tilde{\boldsymbol{\eta}}) = \begin{pmatrix} E(\boldsymbol{\eta}^{(1)}) \\ \vdots \\ E(\boldsymbol{\eta}^{(g)}) \end{pmatrix} \tag{7}$$

and a similar vector for variances $V(\tilde{\boldsymbol{\eta}})$. The tilde indicates that these combine both age and quintile values, with length $= g\omega$.

The next step is to calculate the means and variances of remaining longevity, at each age x, within each group, as follows.

$$E(\boldsymbol{\eta}(x)) = (\mathbf{I}_g \otimes \mathbf{e}_x^{\mathsf{T}})\, E(\tilde{\boldsymbol{\eta}}) \qquad x = 1, \ldots, \omega \tag{8}$$

$$V(\boldsymbol{\eta}(x)) = (\mathbf{I}_g \otimes \mathbf{e}_x^{\mathsf{T}})\, V(\tilde{\boldsymbol{\eta}}) \qquad x = 1, \ldots, \omega \tag{9}$$

where $\mathbf{e}_x$ is a vector of length ω with a 1 in the xth position and zeros elsewhere. The resulting vectors here are of dimension $g \times 1$.

At age x the cohort consists of a mixture of the g different groups ($g = 5$ for quintiles, 10 for deciles) with mixing distribution $\boldsymbol{\pi}(x)$ generated by the differential survival of groups within the cohort.

The mixing distribution $\boldsymbol{\pi}$(x) at age x is a vector of dimension $g \times 1$, which sums to 1. It is obtained from the distribution of groups in an initial cohort. Since quintiles are by definition equally distributed, it would seem that the initial cohort should be evenly distributed. Some other distribution could be used if desired.

Let $\boldsymbol{\pi}(0)$ be the initial mixing distribution, and let $\boldsymbol{\eta}^{(i)}(0)$ be the initial cohort age distribution in group i. Then

$$\mathbf{n}^{(i)}(0) = \mathbf{e}_i \pi_i(0) \tag{10}$$

(i.e., a vector with $\pi_i(0)$ in the first entry and zeros elsewhere. For an evenly distributed cohort, $\pi_i(0) = 1/g$.)

Project each group with its appropriate survival matrix

$$\mathbf{n}^{(i)}(x) = \mathbf{U}_i^x \mathbf{n}^{(i)}(0) \qquad i = 1, \ldots, g \tag{11}$$

add up the entries

$$N^{(i)}(x) = \mathbf{1}_\omega^{\mathsf{T}} \mathbf{n}^{(i)}(x) \qquad i = 1, \ldots, g \tag{12}$$

and create $\boldsymbol{\pi}(x)$ by putting these into a vector and normalizing it to sum to 1

$$\boldsymbol{\pi}(x) = \begin{pmatrix} N^{(1)}(x) \\ \vdots \\ N^{(g)}(x) \end{pmatrix} \frac{1}{\sum_i N(i)(x)} \tag{13}$$

At age x remaining life expectancy for the total population is

$$E(\eta_x) = E_{\boldsymbol{\pi}(x)}\left[E(\boldsymbol{\eta}(x))\right] \tag{14}$$

$$= \boldsymbol{\pi}(x)^{\mathsf{T}} E(\boldsymbol{\eta}(x)) \tag{15}$$

$$= (\boldsymbol{\pi}(x)^{\mathsf{T}} \otimes \mathbf{e}_x)\, E(\tilde{\boldsymbol{\eta}}) \qquad x = 1, \ldots, \omega \tag{16}$$

Notice that η_x is a scalar. The remaining life expectancy at age x is a simple average weighted by the mixing distribution.

The variance in η_x is

$$V(\eta_x) = V_{\text{within}} + V_{\text{between}} \tag{17}$$

with

$$\begin{aligned} V_{\text{within}} &= E_{\boldsymbol{\pi}(x)}\Big[V\left(\boldsymbol{\eta}(x)\right)\Big] && (18)\\ &= \boldsymbol{\pi}(x)^{\mathsf{T}} V(\boldsymbol{\eta}(x)) && (19)\\ &= \left(\boldsymbol{\pi}(x)^{\mathsf{T}} \otimes \mathbf{e}_x^{\mathsf{T}}\right) V(\tilde{\boldsymbol{\eta}}(x)) && (20) \end{aligned}$$

and

$$\begin{aligned} V_{\text{between}} &= V_{\boldsymbol{\pi}(x)}\Big[E\left(\boldsymbol{\eta}(x)\right)\Big] && (21)\\ &= \boldsymbol{\pi}(x)^{\mathsf{T}}\Big[E(\boldsymbol{\eta}(x)) \circ E(\boldsymbol{\eta}(x))\Big] - \Big[\boldsymbol{\pi}(x)^{\mathsf{T}} E(\boldsymbol{\eta}(x)\Big]^2 && (22) \end{aligned}$$

Again, $V(\eta_x)$ is a scalar.

3 Preliminary results

Table 1 and Table 2 show the life expectancy and variation in age at death for males and females, respectively, in each deprivation quintile at each Census year. The same tables reporting life expectancy and variation in age at death conditional upon survival to age 35 are included in the appendices.

Table 1: Life expectancy and standard deviation for males, age 0.

	1981		1991		2001		2011	
quintile	ex	sd	ex	sd	ex	sd	ex	sd
1 (least dep.)	71.6	15.4	74.5	14.4	77.6	13.7	80.2	13.6
2	69.9	16.1	72.9	14.9	75.3	15.1	78.5	14.6
3	69.1	16.2	72.1	15.4	73.9	15.4	77.0	15.1
4	68.2	16.2	70.4	16.1	72.2	16.1	75.3	15.6
5 (most dep.)	66.4	16.4	68.3	16.5	69.0	17.2	72.4	16.5
Total pop.	69.0	16.1	71.6	15.6	73.5	15.8	76.7	15.4

The most deprived quintile experiences the lowest life expectancy and the highest variation in age at death (highest standard deviation) at each year. For males there was an increase in variation in age at death between 1991 and 2001. Although there was some improvement between 2001 and 2011, variation in age at death was very similar to the level experienced 30 years earlier. Females from the most deprived quintile have experienced decreasing variation in age at death (decreasing standard deviation) but the decrease was greater for the least deprived. The change in standard deviation over time, across all ages is shown in Figure 1. Male total variation has increased somewhat, and female total variation has changed very little over the time period studied.

Figure 2 shows the proportion of the total difference in variation in age at death that is due to differences between area level deprivation. For males and females the proportion

Table 2: Life expectancy and standard deviation for females, age 0.

	1981		1991		2001		2011	
quintile	ex	sd	ex	sd	ex	sd	ex	sd
1 (least dep.)	77.1	14.9	79.1	14.1	81.3	13.0	83.4	12.8
2	75.8	15.3	78.6	13.9	80.4	13.6	82.1	13.5
3	75.1	15.6	77.7	14.7	78.9	14.5	80.9	13.9
4	74.4	15.7	76.6	14.8	78.1	14.6	80.0	14.0
5 (most dep.)	72.8	16.3	74.9	15.9	76.3	15.7	77.9	15.1
Total pop.	75.0	15.6	77.3	14.8	78.9	14.4	80.9	14.0

Figure 1: Standard deviation of remaining lifespan for total population by age, Census years 1981 until 2011.

(a) Males (b) Females

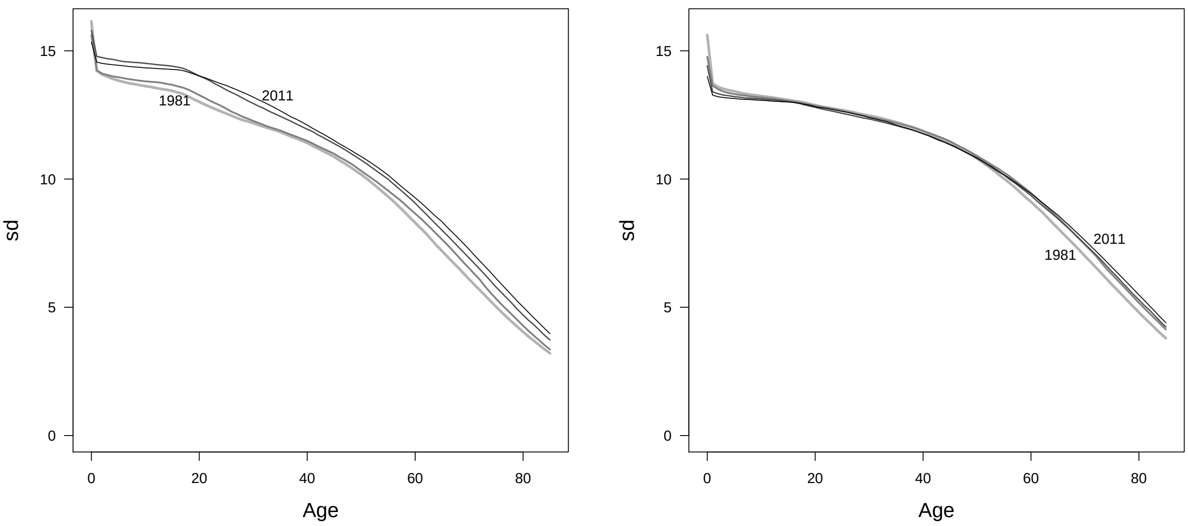
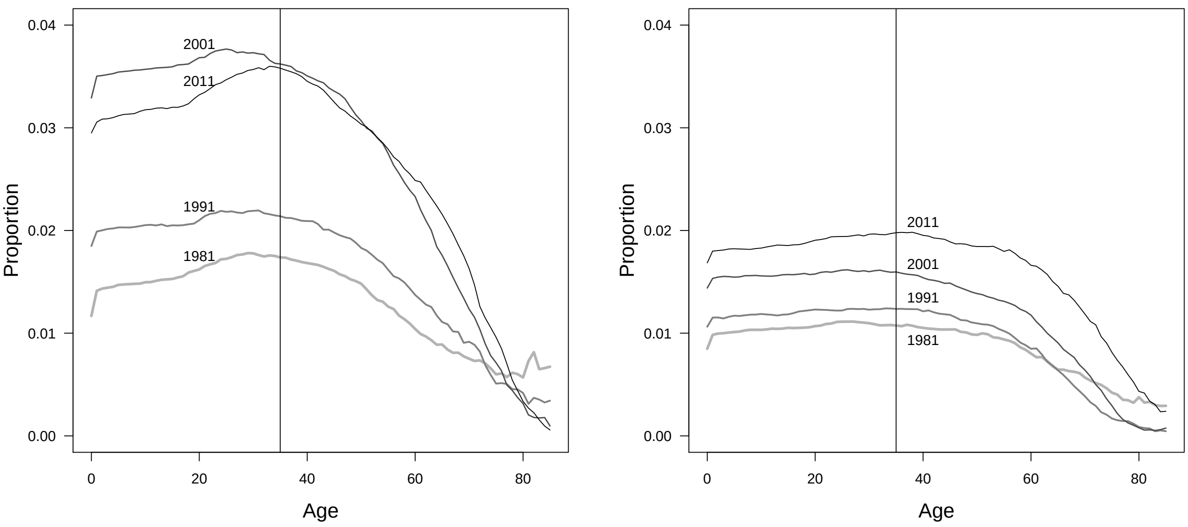


Figure 2: Proportion of variance due to differences between deprivation quintiles by age, Census years 1981 until 2011.

(a) Males (b) Females



of variation explained by between group inequality has increased over time. The increase in this component was greater in magnitude for males than for females.

3.1 Sensitivity analysis

We tested the sensitivity of our results to the size of deprivation group by replicating the analysis using deciles of deprivation, each representing 10% of the population. The conclusions were the same for males and for females. However, the increase in the between group component over time was greater in magnitude when using deciles. We chose to report results for quintiles of deprivation as they are the preferred analytical grouping for routine reporting of health measures in Scotland (NHS Public Health & Intelligence).

4 Discussion and conclusion

4.1 Summary of main findings

Deprivation differences in age at death were evident at all Census years when measuring socioeconomic inequality by area-level. Those living in the most deprived areas can expect to live the shortest lives and experience the greatest variation in age at death: a double burden of mortality inequality. The difference between deprivation groups was larger for males than for females. Males from the most deprived quintile experienced increasing variation in age at death between 1991 and 2001 so that the level of variation in 2011 was the same as that experienced 30 years earlier. The between group component of inequality also increased over the time period. This may indicate that area-level deprivation is becoming a more important determinant of mortality. This finding has relevance for governments who are deciding whether to distribute resources between social policies seeking to intervene on neighborhoods and communities versus social policies seeking to intervene on individuals.

4.2 Strengths and limitations

The data used for this study includes the most robust population estimates and mortality data for the entire population of Scotland. Using a validated area-level measure of socioeconomic inequality meant that complete lifetables could be constructed and no ages were truncated from the analysis. However, it is important to acknowledge the reasons why studies interested in the social distribution mechanisms of adult mortality may consider restricting analysis to older ages. Smits and Monden (2009) suggest only looking at ages 15+ because these are the ages where 80% of deaths in developed countries now occur. Looking only at adult mortality may better reflect the causes of death driving mortality change in more recent time periods: infectious disease and effective medical intervention historically reduced infant and childhood deaths rapidly but reductions in adult mortality are influenced by more complex mechanisms that change slowly (Smits and Monden 2009, Vallin and Mesl 2004). Our results indicate that the age at which the difference in variation in age at death is greatest is around 35 years old. This provides some reassurance for studies that are forced to truncate out younger age groups: the peak of variation in age at death (at least in developed countries) is likely to be captured.

We recognize that our results are vulnerable to the ecological fallacy : it is possible that the association found at the area-level may differ from the association found at the individual (Diez Roux 2002). The consistency of our findings with the existing literature on socioeconomic inequalities in variation in age at death (Broennum-Hansen 2017, van Raalte et al. 2014) indicates that the findings by area-level deprivation are not an artefact. This does not mean that area-level measures and individual level measures are substitutes for one another. Area-level measures capture characteristics of populations and individual level measures capture characteristics of individuals (Leyland et al.

2007b). An example helps to illustrate the contentions. GPs aiming to reduce inequalities between individuals by providing preventative screening programmes may rely on area-level indicators to target those who are most deprived but may target an individual in a deprived area who is actually well-off. So relying on an area-level measure to reduce health inequalities between individuals can be problematic if there is an assumption that the underlying characteristics of the population are socially homogenous (Fischbacher 2014). We acknowledge that deprived individuals do not exclusively reside in deprived areas and affluent individuals do not exclusively reside in affluent areas (Leyland et al. 2007b).

The Carstairs score has been the focus of further criticisms. For example, the meaning of car ownership is fundamentally different for individuals in rural contexts compared to urban contexts. It is also acknowledged that overcrowding may occur out of choice and for cultural reasons rather than simply being a marker of deprivation (Fischbacher 2014). Therefore it has been suggested that the Carstairs score may be an out of date measure of socioeconomic deprivation (Schofield et al. 2016, Tunstall et al. 2011) because the relevance of the variables used for capturing the meaning of deprivation varies across contexts and over time (Norman 2010). In response, it was demonstrated that the scores for each postcode sector at each Census year are highly correlated despite changes to the formal definitions of the variables. This is interpreted as evidence that the underlying information the variables aim to capture is similar or that deprivation has remained stable over time (Leyland et al. 2007b).

4.3 Next steps

Further sensitivity analysis will be carried out using population weighted quintiles derived from an alternative measure of area-level deprivation called the Scottish index of multiple deprivation (SIMD). The SIMD includes indicators from six deprivation domains (employment,income,health,education,access to services, crime and housing). Limitations of the SIMD are that it is only recommended for trend analysis for data beginning in 1996 and that it includes indicators of health and mortality meaning that the full SIMD can not be used for health inequalities research. Instead health inequalities research tends to use the income domain only (Leyland et al. 2007a). The income domain is highly correlated with the full SIMD and is one of the most heavily weighted domains (NHS Public Health & Intelligence, The Scottish Government 2016).

5 Conclusion

Area-level measures of deprivation are an invaluable tool for population health research. Perhaps more importantly, they have pragmatic advantages for governments seeking to identify how resources should be distributed across societies and are actively used by policies which intervene on neighborhoods or communities rather than individuals (Allik et al. 2016, Diez Roux 2001, Robert 1999, The Scottish Government 2016). For these reasons, it is important that more countries evaluate inequalities in variation in age at death by an area-level measure of socioeconomic inequality. This will help to identify if increasing contributions from between group inequality to total variation in age at death is a finding that is dependent upon, and thus amenable to, country specific contexts.

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6 Appendices

Table 3: Life expectancy and standard deviation for males, age 35.

	1981		1991		2001		2011	
quintile	ex	sd	ex	sd	ex	sd	ex	sd
1 (least dep.)	38.4	11.4	40.9	11.2	43.7	11.3	46.3	11.1
2	37.1	11.6	39.5	11.6	41.9	11.8	44.7	12.0
3	36.4	11.8	38.8	11.8	40.5	12.2	43.4	12.4
4	35.5	11.8	37.6	11.9	39.1	12.5	41.7	13.0
5 (most dep.)	33.8	12.1	35.8	12.3	36.6	13.2	39.2	13.5
Total pop.	36.2	11.8	38.5	11.8	40.3	12.4	43.1	12.6

Table 4: Life expectancy and standard deviation for females, age 35.

	1981		1991		2001		2011	
quintile	ex	sd	ex	sd	ex	sd	ex	sd
1 (least dep.)	43.4	11.7	45.1	11.5	47.0	11.1	49.0	49.0
2	42.2	12.0	44.4	11.7	46.1	11.5	47.9	47.9
3	41.6	12.2	43.8	12.1	45.0	12.0	46.7	46.7
4	41.0	12.1	42.7	12.3	44.1	12.2	45.7	45.7
5 (most dep.)	39.6	12.7	41.4	12.8	42.6	13.1	43.9	43.9
Total pop.	41.5	12.2	43.4	12.2	44.9	12.1	46.7	46.7