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# The Shape of the Relationship Between Mortality and Income in France

**Florence Jusot***

**ABSTRACT.** – Using a case-control study constructed with two fiscal databases, this paper investigates the shape of the relationship between income and the probability of death in France. The results show that the risk of mortality is strongly correlated with the level of income, independent from the occupational status. This relationship holds across the whole range of income distribution. Specifically the protective effect of highest incomes casts some doubt on the hypothesis of the concavity of the income-health relationship.

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## La forme de la relation existant entre mortalité et revenu en France

**RÉSUMÉ.** – Cette recherche explore la forme de relation existante entre probabilité de décès et revenu en France, sur la base d'une étude cas-témoins constituée à partir de deux bases de données fiscales. Les résultats montrent que le risque de décès est fortement corrélé au niveau de revenu, après contrôle par la profession. Cette relation existe tout au long de la distribution des revenus. En particulier, l'effet protecteur des plus hauts revenus remet en cause l'hypothèse de concavité de la relation revenu-santé.

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* F. JUSOT : IRDES (Institut de Recherche et Documentation en Economie de la Santé), 10 rue Vauvenargues, 75018 Paris, France. E-mail : jusot@irdes.fr, téléphone : 01.53.93.43.16, fax : 01.53.93.43.50.

# 1 Introduction

In France, as in many other countries, strong social inequalities exist in mortality. In the 1990s, a 35 year-old manager in France could expect to live 7 more years compared with a manual worker (a life expectancy of 46 years against 39 years) (Monteil and Robert-Bobée, 2005). Furthermore, social differences in the risk of mortality persist and even have been increasing in the past decades for all causes of death in France (Jougla *et al*., 2000). Several authors suggested that between the 1950s and the 1990s there has been a more rapid reduction in the risk of mortality for white-color workers compared with blue-color workers (Desplanques, 1985, 1993; Monteil and Robert-Bobée, 2005). The situation is particularly alarming given that social differences in premature mortality are the highest in Europe (Kunst *et al*., 2000). This finding can be surprising considering that using five performance indicators including equity, the WHO (2000) ranked France as the best health care system in the world. Therefore it is important to better understand the mechanisms which contribute to social health inequalities in France.

This paper proposes an analysis of the relationship between mortality and income in France, which is one of the major questions in the field of research on health inequalities (Kawachi *et al*., 2002).

Understanding the relationship between income and mortality is critical for two reasons.

First, studying the shape of the health-income relationship would allow to better understand whether social differences in health status are caused by poverty (the poverty hypothesis), or whether there is a socio-economic gradient in health status, i.e health status rises with each level of socioeconomic status (the absolute income hypothesis).

Second, the analysis of the profile of the relationship between income and health would contribute to the debate on the impact of income inequality on health (e.g. Wilkinson, 1992; Deaton, 2003; Wagstaff and van Doorslaer, 2000; Mackenbach, 2002; Subramanian, 2003). If individual health is positively associated with income, but with diminishing returns, the well-documented correlation existing between income inequality and average level of health at the aggregate level could only be a statistical artefact (Gravelle, 1996) induced by a concavity of this relation, without any contextual effect of income inequality[1]. This assumption is supported by several American studies (e.g. Mc Donough *et al*., 1997; Mellor and Milyo, 1998; Smith and Kington, 1997). However, is also likely that national context and particularly social and health policies could influence this relationship.

In this context it is important to disentangle the impact of income from that of other factors determining socio-economic status such as occupation and education. The correlation between mortality and income is well-documented (e.g Kitagawa and Hauser, 1973; Duleep, 1986; Adams *et al*. 2003; Deaton, 2003) and several studies have showed that income is the most powerful socioeconomic predictor of

1. If the relationship between income and health is concave, any increase in income inequality, for a given average income, will result in a reduction in the average health status of the population, because the improvement in health of a wealthy person whose income increases, will be lower than the reduction in health status resulting from the reduction in income of a poorer person.

mortality (Geyer and Peter, 2000; Duncan *et al.*, 2002). Even where income and occupational status are correlated, the within-group variance is twice that of the inter-group variance (Jusot, 2003). In particular, the distribution of income among the most highly educated is more dispersed than for the population as a whole (Nauze-Fichet, 2002). If occupational status is an imperfect measure of income, the relationship between occupational status and mortality would produce an imperfect measure of the link between income and mortality (Wilkinson, 1986; Menchik, 1993). Basically, these two indicators (income and occupational status) do not reflect the same dimensions. French occupational classes are a synthetic indicator of education, working conditions, prestige of occupation, living conditions and lifestyles (Desrosières and Thevenoy, 2000). Indeed, all these dimensions are well-documented to be correlated with health status (Wilkinson, 1986; Marmot and Wilkinson, 1999; Smith, 1999). According to health capital models (Grossman, 1972; Erlich and Chuma, 1990), health status increases with disposable income, because income determines the resources available to an individual for investing in his health capital. Hence income and occupational status should have distinct effects on mortality. But most of these studies do not separate the impact of income controlling for educational and occupational status. By using separate indicators of socioeconomic status, this study separates the effects of both occupation and education, approximated by occupational status, from the effect of disposable income.

The analysis is based on mortality data as we consider that mortality rates are more appropriate for studying social inequalities in health. Self-assessed health and morbidity indicators are sensitive to the reporting differences between social groups (Idler and Benyamini, 1997; Lindeboom and van Doorslaer, 2004; Mackenbach, 1996)[2]. Furthermore, differences in life expectancy may be regarded as a synthetic indicator of the social differences which affect health during the life course (Aïach, 2000). Unfortunately, data on mortality differentials in France are scanty (Feinstein, 1993). Until recently, available data did not allow to describe the relation between income and mortality for the general population in France. Existing studies only document differences by occupational status (Mesrine, 1999; Monteil and Robert-Bobée, 2005) or employment status (Mesrine, 2000).

Given the lack of income data in the databases usually used in mortality studies, this analysis uses two fiscal data bases: the 1988 Wealth at Death Survey and the 1990 Household Taxable Income Survey. Hence the analysis is based on a comparison of the characteristics of persons who died in 1988 (the Wealth at Death Survey) and those of the persons surveyed in 1990 (the 1990 Household Taxable Income Survey), and therefore alive in 1988.

The results of our analysis first confirm that, there is a high correlation between income and mortality, even controlling for occupational status. Second, the observed relationship does not support the hypothesis of a concave relationship. Thus, we confirm that there is excess mortality associated with poverty, as well as a less expected effect, a protective impact of highest income on health.

This paper is organised as follows. In the next section, we provide a review of the theoretical background for this study with some empirical results on the relationship between income and health. The design of the case-control study and the

2. Reporting an illness presupposes awareness of that illness and a search for appropriate care for that illness. In contrast self-assessed health status is highly dependent on the distinction made by the respondent between a normal state of health and an impaired state of health.

methodology are described in section 3. The results are presented in section 4 and we close with a discussion in section 5.

# 2 Theoretical Background and Empirical evidence

The studies providing evidence of a positive correlation between socioeconomic status and health status are abundant in developed countries. In France, the existence of a strong social gradient in health status is also well documented (Leclerc at al., 2000). This correlation appears to be independent of the choice of health indicators, such as mortality and morbidity, and to the choice of socioeconomic status indicators, income, occupational status or employment status. While the relationship between socio-economic status and health is well established, the mechanisms behind and its implications for health policy are less obvious.

In literature, it is well accepted that the relationship between health and socioeconomic status is dual, socioeconomic status impacts health but health status impacts socioeconomic status as well. However, the shape of the relationship between health and socioeconomic status is not well defined. It is not clear if the income-health relationship concerns only the very poor (poverty effect), or the relationship between health and income occurs across the whole range of income distribution.

## 2.1 A Dual Relationship Between Health and Income

Health capital models[3] (Grossman, 1972, Erlich and Chuma, 1990) propose a theoretical approach to the analysis of the link between mortality and income.

These models suppose that a person has an initial stock of health capital (his genetic inheritance) which depreciates in successive periods (the biological effects of ageing). Death occurs when the level of health capital drops below some minimum level. In any given period a person chooses between the consumption of goods from which he derives utility, and investing in his health by consuming health care, medical goods or other activities which are beneficial to health (sport, a balanced diet etc.).

Hence, theoretically health status increases with disposable income, because income determines the resources available to an individual for acquiring goods, including medical goods. The existence of social inequalities in the consumption of health care is well known. Differences in the structure of consumption have been largely demonstrated in France, as in other countries (Bocognano *et al*., 1999; Newhouse, 1993; van Doorslaer *et al*., 2004). The poorest consult fewer specialists, dentists and opticians, and use less hospital and nursing care. In France, the under-

3. The endogenous determination of life expectancy in Grossman's original model has been widely disputed. Erlich and Chuma's demand model of longevity (1990) proposed a specific trade-off between longevity and quality of life. (For a more detailed discussion see Grossman, 2000).

consumption of ambulatory care corresponds to less frequent consultations rather than lower expenditure for a given episode of care (Breuil-Genier *et al.*, 1999). These differences are explained, at least in part, by the effect of income on the decision to invest in health capital. For example, 14% of persons interviewed in the 1998 Social Protection and Health Survey gave financial reasons for not seeking care, this percentage increasing to 24% for those with income of less than 3 000 FF per household (Bocognano *et al.*, 1999). It has also been shown that health expenditure increases to a certain extent with income (Grignon and Polton, 2000).

Apart from the direct effect of disposable income on access to care, the literature suggest also a reverse effect from health to socioeconomic status. In the health capital models (Grossman, 1972, Erlich and Chuma, 1990) health status affects the time during which a person can work, because his health capital determines the period of good health available for work, consumption of goods and investment in health capital. From the perspective of a work-leisure trade-off, employment prospects are partly determined by health status, worse health making work more difficult, or impossible (Couffinhal, 2002a, b). Moreover, efficiency wage models (Leibenstein, 1957) suppose an effect on productivity of health, and hence on wage.

The effect of health on inactivity and unemployment has been clearly shown in France (Saurel-Cubizolles *et al.*, 2002, Jusot *et al.*, 2006), as in others countries (Currie and Madrian, 1999). However the effect of health on wage is less well documented. It seems that the effect of health on productivity would appear to be limited to the least well-off in France. According to the RMI Beneficiary Survey, the rate of increase of monthly income, corrected for the number of hours worked, decreases with ill-health (Rioux, 2001). In contrast, the Living Conditions Survey shows that handicap or serious ill-health do not affect investment in human capital. The temporary reduction in income is explained entirely by the loss of experience entailed by absence from the labour market (Lechene and Magnac, 1998). Similarly, according to the Health and Retirement Survey, the effect of unanticipated health events is explained by the reduction in number of hours worked, and not by a reduction in wage rates (Smith, 1999).

Finally, some other factors such as unfavorable health-related habits or poor working conditions, could have a negative impact on health and then explain the correlation between socioeconomic status and health (Smith, 1999; Adams *et al.*, 2003; Marmot and Wilkinson, 1999).

This dual relation between socioeconomic status and health is one of the most important issue raised by the recent studies on social health inequalities (Adams *et al.*, 2003, Adda *et al.*, 2003, Meer *et al.*, 2003, Michaud and van Soest, 2004), in particular for the design of adequate public policies seeking to reduce social health inequalities. If health inequalities are due to a causal effect of socioeconomic status on health, reducing health inequalities would focus on income redistribution, equity in education, equity in access to health care or working conditions policies. Conversely, if health inequalities are principally the result of the impact of poor health on socioeconomic status, public policies must prevent health conditions that reduce ability to work and favor integration of disabled people into labor market. However the results do not provide unambiguous conclusions on the causal mechanisms.

Adams *et al.* (2003) propose to test the absence of causal links from socioeconomic status to health and from health to wealth using Granger causality test. The application of this method to the three waves of the AHEAD panel leads the authors to conclude that, among the elderly American population, the evidence supports the hypothesis of no direct causal link from socioeconomic status (income and wealth) to mortality and to incidence of new health conditions, when initial health conditions are controlled, but rejects the hypothesis of no direct causal link from health to wealth. However the hypothesis of no direct causal link from wealth to health is rejected for most of the chronic or mental diseases, as well as for self-assessed health. Using the same methodology (without testing for invariance), Adda *et al.* (2003) shows that, among the sample of British civil servants (Whitehall II Study) the hypothesis of no direct causal link from socioeconomic status to health cannot be rejected for some measures of new health conditions, but is rejected for some others, like self-rated heath. Among the Swedish population (ULF panel), the results are also comparable to the US. In particular the incidence of cancer does not appear to be linked to socioeconomic status, contrary to change in self-rated health. The replication of this methodology in the six waves of the Health and Retirement Study leads to reject the hypothesis of non causality from wealth to health, but the tests in a dynamic panel data model incorporating unobserved heterogeneity do not provide evidence of a causality from wealth to health. Conversely, both methodologies lead to the strong evidence of causal effect from health to wealth (Michaud and van Soest, 2004). Using quite similar models as Adams *et al.* (2003), Hurd and Kapteyn (2003) find that changes in health are more related to income in the U.S. (HRS/AHEAD) than in the Netherlands (CSS and SEP), the effect of health on change in income is greater in the Netherlands than in the U.S. Using inheritance as a suitable instrument for the change in wealth, Meer *et al.* (2003) show that wealth changes have a very small effect on health change, based on the PSID data. Finally, using the natural experiment in which the level of pensions was modified in 1977 in the US, affecting only the generations born after 1917, Snyder and Evans (2002) show a negative impact of exogenous income on mortality among elderly.

## 2.2 Absolute Poverty Versus Absolute Income

The shape of the relationship between income and health is also a major issue addressed in the empirical literature. There are two main hypotheses.

Under the first hypothesis, *the absolute poverty hypothesis*, mortality differences are limited to excess mortality amongst the poorest, resulting from very poor working conditions, poor housing and limited access to health care. Under this hypothesis health state increases with income, but only up to a specified poverty threshold (Wagstaff *et al.*, 2000). Above this threshold, health state does not improve with increasing resources.

The absolute poverty hypothesis has been seriously questioned, particularly in Britain, notably in the Black report (Townsend, 1982). This report showed that, despite the existence of a health system accessible to all (the National Health Service), and the virtual disappearance of absolute poverty, differences in the risk of mortality in men aged between 15 and 64 had widened between occupational categories between 1931 and 1971. However the validity of the absolute poverty hypothesis has been partly re-established, by demonstrating the importance of living conditions in childhood for adult health and the importance of social reproduction. If the poorer socioeconomic groups lived in absolute poverty during childhood, their poor state of adult health could be due to their childhood lifestyle. This hypothesis, known as the "early life hypothesis" depends on the significance of very long-term effects on health (Wadsworth, 1999; Case *et al.*, 2002). For example, Barker (1997) has shown that the risks of cardiovascular illnesses are partly determined in utero, and Ravelli (1998) has shown, that diabetes in adulthood could be explained by exposure to famine during pregnancy with a natural experiment in the form of the Dutch experience between 1944 and 1945.

Furthermore, the lack of health care consumption, induced by financial hardships, seems to partly explain the bad health status of the poorest part of the population. The Rand experiment in the US showed that free access to care did not produce any significant change in health status other than for the poorest populations and at risk groups, such as children and hypertensives (Newhouse, 1993). In France, before the establishment of the universal health insurance coverage (Couverture Maladie Universelle), problems of access to care for the poorest were due to lack of medical insurance, even given compulsory cover for care (Dourgnon *et al*., 2001). Social security covers only 75% of the cost of care, and the absence of supplementary insurance was largely correlated with poverty. In 1998 for example, only 52% of persons with a monthly income below 2 000 FF per consumption unit had supplementary insurance cover (Bocagnano *et al*., 1999). While it is difficult to establish empirically the impact of health care consumption on health, the absence of supplementary insurance could explain the bad health status of the poorest: failure to seek care or delays in seeking treatment (Couffinhal *et al*., 2002a, b).

Under the second hypothesis, known as the *absolute income hypothesis*, social inequalities in health cannot be reduced to a simple dichotomy between the poor and non-poor, manual and non-manual workers. Rather, disposable income has a continuous effect on health, and this underlies an economic gradient in health. In France a social gradient in mortality has been shown, with white collar workers living on average longer than blue collar workers, and executives living longer than intermediate professions (Mesrine, 1999, Monteil and Robert-Bobée, 2005). According to Deaton (2003) and Wagstaff *et al*. (2000), the improvement in health status resulting from an increase in resources declines with increase in the level of income. This supports the hypothesis of decreasing returns in the health capital production function in Erlich and Chuma's model (1990). Several studies, all using American data, have shown the existence of this continuous and concave relation between income and health (Mellor and Milyo, 1999; Smith and Kington, 1997; Wolfson, 1999). However, these studies have not controlled for occupation and have paid scant attention to the highest income categories. In these estimations income is usually introduced in a form which is likely to result in a concave relation between income and health: logarithmic (Wolfson, 1999) or quadratic (Mellor and Milyo, 1999) (which is additionally very restrictive). When income is specified in blocks, the highest block often does not enable analysis of the effect of the highest incomes. For example Smith and Kington (1997) study the concavity of the relation between income and functional limitation using three classes of income.

# 3 Data and Methodology

The dearth of studies examining the income-mortality relation in France reflects the lack of income data in the demographic data bases used to study mortality, and the fact that it is not possible to study mortality using French economic surveys. Most surveys which include questions about income are cross-sectional surveys, which by design only include living persons, (for example the Household Budget and Taxable Income Surveys). Longitudinal surveys (such as the DADS panel or the Health and Social Protection Survey) should enable the identification of deceased persons during the course of the survey. However they cannot be used for the study of mortality. Because the death rate in the general population is low, large sample sizes are required, which must then be followed for a long period, or include a high proportion of elderly persons, in order to obtain a statistically significant number of deaths. Furthermore, deaths are not always clearly identified among other possible causes of attrition. Demographic studies therefore rely on specific sources (such as the Permanent Demographic Sample or the longitudinal mortality studies of the INSEE[4]) which collect only death certificate and census data.

Here we use data from the only representative survey which includes information on the income of deceased persons, the 1988 Wealth at Death Survey. However because this survey collects information on deceased persons only, the analysis is based on a case-control study using data from this survey and the 1990 Taxable Income Survey.

## 3.1 The Wealth at Death Survey: a Selected Sample

The 1988 Wealth at Death Survey was carried out jointly by the National Institute for Statistics and Economic Studies (INSEE) and the Directorate General of Taxes (DGI) (Laferrère and Monteil, 1995). The sample consisted of 4 570 persons, aged 20 or more, living in metropolitan France outside Corsica, who died in 1988.

This sample is drawn from the Permanent Demographic Sample (EDP). The EDP is a random sample of individuals selected on the basis of their birth date, and for whom information has been collected from the civil register and censuses since 1968. Thus for a representative sample of 1% of the population, we have information on their civil status: date of birth, marital status, number of children, date and place of death, and profession. To supplement this information, fiscal information relating to income was collected by the tax authorities from the tax declaration for the year preceding death for persons in the EDP who died in 1988. Some information collected in the census was not recorded in the PAD survey. Specifically, educational attainment and economic activity at the time of the census (inactive, retired, unemployed, employed) is not known.

Although information on age at death and the income declared by the tax unit of the deceased in the year preceding death is available for a sufficiently large and representative sample, this does not enable us to properly estimate the relation between income and mortality.

Given the hypothesis of a stationary population, the determinants of risk of death may be analysed for a representative sample of deaths in one year, using survival models (Cox, 1984) as the distribution of their characteristics is exactly the same as the distribution of the characteristics for a representative sample of an extinct cohort. In the case of a non-stationary population, indirect methods of mortality estimation, qualified by the variable *r procedure* (Preston, 1999), enable the cal-

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culation of survival functions for the population corrected for changes in the size of cohorts and in mortality rates, on the basis of the distribution of age at death in one year, and annual rates of change of the population in each age group. However, these methods can only be used to demonstrate differences in mortality according to income under in the one hand the assumption that income is not affected by generation effects and in the other hand the assumption that differences in mortality between the income groups are stable over time (Jusot, 2003). Without theses assumptions, the determinants of the risk of mortality can not be analysed on the basis of the Wealth at Death Survey, because this sample is strictly selected on the basis of the variable we wish to study – mortality.

## 3.2 **The 1990 Taxable Income Survey: a Relevant Control Group**

To resolve the problem of response-based sampling (Manski, 1995) we decided to design a case-control study. Case-control studies are often used in epidemiology to measure the association between a risk factor and a disease (Schlesselman, 1982, Bouyer *et al*., 1995), particularly for those diseases which are too rare to be studied in the general population. A case group consisting of persons with the disease, and a control group of persons without it, are selected. By comparing risk factors in both groups, risk factors for the outcome in question can be determined (disease in epidemiological studies and death in this case).

In this context the sample in the Wealth at Death Survey constitutes the case group, consisting of persons died in 1988. To select the control group *i.e.* a group of survivors, we used data from the 1990 Household Taxable Income Survey because this source is most directly comparable with the Wealth at Death Survey. Carried out jointly by INSEE and the Directorate General of Taxes, the 1990 Household Taxable Income Survey is part of a series which began in 1956 (Campagne *et al*., 1996). For a random sample of ordinary households this survey collected data on income declared by their tax unit using tax declarations, and, by aggregation, of the income declared by the household.

We assume that the two samples are derived from the same population, alive at the beginning of 1988, and that the determinants of mortality in 1988 may be estimated by comparing the characteristics of a randomly selected sample of persons deceased at the end of 1988, the PAD Survey sample, with the characteristics of a representative sample of persons alive at the end of 1988, the sample of the Taxable Income Survey. However the latter sample is not totally representative of 1988 survivors, because it contains only those persons alive in 1990, and therefore still alive at the end of 1989. However any bias caused by the absence of those dying during 1989 would be very small, because the low death rate in the general population would ensure that the distribution of characteristics in the population of survivors at the end of 1989 would be similar to that of 1988 survivors.

## 3.3 **The Study Sample**

Because this study is based on an analysis of two populations, from two different surveys, the robustness of the results depends on the representativeness of both surveys, and on the comparability of their content. Thus, this has limited the information we have been able to use. The representativeness of the surveys is partly ensured by the random sampling method, and partly by taking into account the weighting of each survey in the econometric analysis. The homogeneity of the sources is essentially due to the method of collection of income data, namely retrospective collection of data on income declared to the tax authorities. For both surveys we know, for each individual, the income declared[5] by his tax unit and the number of persons in that tax unit.

Hence information on taxable income used in this study differs from that normally used. Firstly we do not have data on individual incomes from the PAD Survey. Secondly "taxable household" is a legal term which does not correspond exactly to other concepts of household. To determine available income, we have used the equivalence scale estimated by INSEE for the Household Budget Survey of 1989 (INSEE, 1995). This scale gave a weight of 1 to the first person in the tax unit and 0.35 to other persons[6]. This scale does not take age into account, because the age of other persons in the taxable household is not recorded in the PAD Survey. Finally, taxable income includes only declared income and does not include transferred income (allocations, the minimum retirement pension and income support from 1990 onwards). In order to avoid any bias caused by specificities of tax declaration for agricultural, commercial and industrial income we have excluded farmers and self-employed persons (including the liberal professions).

Our sample included only heads of household from the Taxable Income Survey. This restriction was necessary given the differences in coding of occupational categories between the two surveys. The classification used for all individuals in the Household Taxable Income Survey differs from that used in the census, and hence in the Wealth at Death Survey. In the Household Taxable Income Survey, occupations are coded according to the census occupation coding system which enables the reclassification of retired persons according to their former occupation, for household heads only. The change in nomenclature results in discrepancies in a third of cases (Campagne *et al.*, 1996). To avoid bias induced by the selection of heads of household in the Taxable Income Survey we restricted the sample to the male population, given that men are more often heads of household. We also excluded economically inactive persons (except for reclassified retired persons) and retired persons whose former occupation was not known.

Thus our study sample consists of 13 399 men in total, 1 438 of whom comes from the Wealth at Death Survey (the deceased sample) and 11 961 were from the 1990 Taxable Income Survey (the survivor sample) (Appendix 2, Table 1). The results obtained will be specific for the selected sample of employed men, still active or retired, and cannot be extrapolated to the whole population.

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5. The incomes of persons deceased in 1988 are adjusted for inflation at a rate estimated by the INSEE for all urban households.
6. The results are not sensitive to changes in the equivalence scale (for a scale of 1 for the first person in the household, 0.7 for the second and 0.5 for the third; and for a scale of 1, 0.5 and 0.3 respectively.)

## 3.4 Methodology

Given that the validity of case control studies depends on the properties of the odds ratio, the determinants of the probability of death may be estimated by a LOGIT model.

Our model assumes that the unobservable individual health capital Hi is explained by the observable characteristics Xi and a residual u, distributed according to a logistical distribution:

(1) $$H_i = \alpha + X_i\beta + u_i$$

In the health capital model (Grossman, 1972; Erlich and Chuma, 1990), death occurs when the health capital drops below a critical minimum level Hmin. Assuming that Hmin is equal to zero, the probability of death conditionally to the characteristics Xi, P(D|Xi) is therefore equal to:

(2) $$\mathrm{P}(\mathrm{D}|\mathrm{X_i}) = P(H_i < H\min) = 1 - F(\alpha + X_i\beta) = \frac{1}{1+\exp(\alpha + X_i\beta)}$$

It is straightforward to show that in the case of a LOGIT model, the estimated value of the β parameters is related only to the differences in distribution of the X variables in the population of deceased and survivors; it is independent of the size of each sub-population. Given the low probability of death (1.4% for men aged over 20 in 1988) the estimated odds ratio may be interpreted as the relative risk of death with regard to the reference category (see Appendix 1).

Our dependent variable is vital status, and the latent variable can be interpreted as health status. The explanatory variables used are: age, occupational status, marital status, geographical environment, and taxable income per C.U.

Age is introduced by a piecewise-linear function. Four occupational groups are distinguished: worker, employee, intermediate profession and executive. Indeed these are social class codes (PCS at one digit) constructed by the INSEE taking into account a number of social characteristics of the individuals in particular the occupation and working environment. Marital status variable allows to distinguish married people, single people, divorced and widowed. The geographical environment is taken into account by introducing dummies for each Territorial Development and Planning Zone (ZEAT), the sample size being too small to introduce a fixed effect by regions. Taxable income per C.U. is introduced in several specifications.

Eight models are estimated. The results presented in the appendix 2 show estimations made for the whole sample, separate analyses of the under 65 and over 65 population giving very similar results.

In order to test the representativeness of our database, we first estimates in the model 0 the conditional effect on the probability of death of the 4 socio-demographic dimensions: age, occupational status, marital status and geographical environment.

In the models 1 to 7, we analyse the relationship between income and mortality.

In the model 1, we test the association between the logarithm of equivalent income and the risk of death (Model 1), controlling for age, marital status, and area of residence. In Model 2 (and subsequent models) we introduced occupational status. In fact, the analysis of the distribution of income within each occupational group shows that occupational status and income are different concepts. These variables are imperfectly correlated (Appendix 2, Table 2). By using two indicators of socioeconomic status we are able to separate the effects of working conditions and of education, approximated by occupational status, from the effect of disposable income.

These two models assess the association between income and health but can not provide an estimate of the causal effect of income on health. By introducing income directly into the econometric analysis a problem of endogeneity arises because of the dual relation between health and income. Thus the direct estimation produces an overestimate of the effect of income on health and the simultaneity bias is an increasing function of the effect of health on economic status. The rigorous separation of the direct effect of income on health from the reverse effect, called the "healthy worker effect", is only possible if an instrumental variable of income is available for the health equation. Such an instrument is very difficult to find, as Adams *et al.* (2003) have noted[7]. One solution is to study sub-populations for whom total resources are, by construction, independent of health state. Populations of children are potentially interesting here if one supposes that the health state of children does not affect parents' decisions to work (Case *et al.*, 2002). A direct effect of income on health can be shown using income of the partner, again under the assumption that the partner's income does not affect a person's decision to work. Another approach is to study the effect of income on health status, controlling for preceding health status (Adams *et al.*, 2003; Duleep, 1986; McDonough, 1997; Smith and Kington, 1997).

From the survey used here, we do not have information on health status before death, nor on the cause of death. Nevertheless it is possible to use information on the type of income received by individuals, rather than the level of income, to obtain information on their health status (Smith and Kington, 1997). For example, receipt of an invalidity pension indicates a poor state of health. In the surveys analysed in this study, invalidity pensions are not distinguishable from other pensions and allowances, particularly retirement pensions, and we are unable to identify retired persons in the Wealth at Death survey. However, according to the ESTEV survey, physical problems increase the probability of early retirement (Saurel-Cubizolles *et al.*, 2001). Hence we assume that income coming from pensions and allowances before age 60 indicates a poor initial state of health, thereby helping to avoid the bias of simultaneity (model 3 and subsequent models). Above the age of 60, we cannot control for the "healthy worker effect" when analysing mortality, which means that the results must be interpreted cautiously.

To explore the shape of the relationship between income and mortality and to separate the absolute poverty and absolute income hypotheses, we use various specifications for equivalent income. To test the concavity hypothesis, we use the logarithmic form (models 1,2 and 3) and the quadratic form (model 4). Then we use a cubic specification (model 5) in order to assess more specifically the association between the highest incomes and health. To study the income health relation across the whole distribution, income is introduced in quintiles, without

7. Note however the random attribution of different rates of sickness insurance cover in the Rand Experiment (Newhouse, 1993) and the natural experiment in which the level of pensions was modified in 1977 in the US, affecting only those generations born after 1917 (Snyder and Evans, 2002).

controlling for occupation (model 6), and then introducing occupational status in model 7. Quintiles of income are calculated taking into account the weights for each survey, and re-weighting the whole sample to make it representative of the population alive on January 1, 1988, such that the population of deceased and survivors correspond to a mortality rate for the male population aged 20 in 1988, namely 1.415%.

# 4 Results

## 4.1 Descriptive Analysis

Our study sample consists of 13 399 men in total, 1 438 of whom comes from the Wealth at Death Survey (the deceased sample) and 11 961 were from the 1990 Taxable Income Survey (the survivors sample) (Appendix 1, Table 1).

The average age is significantly higher in the deceased sample than in the survivors sample. 38% of the survivors are aged under 40 while the proportions is 4% in the deceased sample. Conversely, 46% of the deceased persons are aged over 75 and this proportion is only 5% in the survivors sample. The distribution of occupational status is also different among the deceased sample and the survivors sample. The proportion of manual workers is higher in the deceased sample than in the survivors one (52% against 36%), and the proportion of man in executive position is lower (10% against 24%). Marital status seems also to be correlated with mortality. The proportion of married man is higher in the sample of survivors (81% against 68%) while the proportion of widowers is higher in the deceased sample (19% against 2%). The geographical distribution of the two sample over the regions is not statistically different.

The descriptive analysis appears to confirm a link between mortality and income, because the average taxable income per consumption unit is significantly lower in the deceased sample than the survivors (84 000 FF compared to 97 000 FF). Moreover, the income distribution of survivors is above that of deceased persons at any point of the income distribution curve (Appendix 2, Table 3). However these results cannot be entirely explained by the carrier effect on income, because separate analyses for under and over 65 year olds give the same results. If the quintiles are defined for the whole population alive on January 1, 1988 (Appendix 2,Table 4), the deceased are more concentrated at the lower end of the distribution than survivors. In fact the proportion of deceased in each quintile decreases across the whole distribution of income, whatever the population analysed (total population, under 65's, over 65's). Although these results should be interpreted cautiously because income is strongly correlated with age, they suggest that death is not randomly distributed in the population – it affects those at the lower end of the income distribution more.

## 4.2 Traditional Determinants of Mortality

First of all, our multivariate analysis (Model 0, Appendix 2, Table 5) confirms the results of previous demographic studies in France which look into the impact of these variables on mortality just by controlling age and sex (e.g. Vallin, Meslé and Valkonen, 2001).

Risk of death increases with age. The social gradient of mortality corresponds to the occupational hierarchy, although the probability of death of manual workers is not significantly different from that of white collar workers[8]. Matrimonial status is also correlated with risk of mortality. Married persons are at lower risk of death than single persons, divorcees or widows/widowers. Finally, our multivariate analysis has confirmed the existence of geographic differences in mortality (Salem, Rican and Jougla, 2000) - our results show excess mortality in the Nord-Pas-de-Calais.

The consistency of our results confirms the robustness of our case-control methodology.

## 4.3 A Strong Correlation Between Mortality and Income

The positive correlation between income and health is confirmed by our multivariate econometric analysis. An increase in equivalent taxable income is significantly associated to a reduced risk of mortality (Model 1, Appendix 2, Table 5). Income appears to have an effect on risk of mortality independently of profession or education because both of the two indicators for socioeconomic status are significant (Model 2, Appendix 2, Table 6). The introduction of logarithmic equivalent income reduces only slightly the explanatory power of occupational status. Moreover, separate analyses show an association between income and mortality among each occupation groups (Jusot, 2003). These results confirm the fact that occupational status and income reflect different dimensions correlated with health.

The proportion of income derived from pensions and allowances seems to be a proxy of previous poor health status before age 60 because it is positively correlated with risk of mortality. When it is introduced in the model it reduces the coefficient of income in line with the "healthy worker effect" hypothesis (Model 3, Appendix 2, Table 6).

Marital status is also significantly linked to risk of death. This is at odds with the frequently cited hypothesis that the correlation between marital status and longevity is partly a function of the correlation between marriage and economic status (Vallin, Meslé and Valkonen, 2001). The excess mortality of single persons may be a consequence of a selection effect on the marriage market, or of the harmful effect of single status, or of a protective effect of marriage (Wilson and Oswald, 2005). Excess mortality among widowers, which cannot be explained by the selection effect, may however, reflect the protective effect of marriage, or the harmful consequences of widowhood on health.

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8. The fact that the risk of death is the same for both categories is consistent with the results of previous demographic analyses. Firstly it reflects the fact that we could not distinguish between skilled and unskilled manual workers. Secondly, according to Mesrine (1999), workers who provide direct services to private individuals, who are included in the occupational category of employees, have a higher risk of mortality than manual workers.

The excess mortality observed in the Nord-Pas-de-Calais region is significant after controlling for income and occupational status. This contradicts the hypothesis of Salem, Rican and Jougla (2000) who suggest that the high mortality rates in this region can simply be explained by the high proportion of manual workers and the economic problems of the region. Furthermore, the South-West zone is associated to a lower mortality risk, after controlling for income. This effect is then consistent with previous demographic literature.

## 4.4 Income Elasticity of Mortality Risk

The estimations of the impact of logged income on the probability of death allow us to calculate the elasticity of the probability of income-related death. This indicator is easier to interpret than the value of the coefficients estimated by logistic regression. Moreover a recent study which used the administrative files of the Inter-Régime Sample, proposes for a few cohorts of retired persons an estimation of the elasticity of the probability of survival for four years as a function of entitlements at retirement in France between 1997 and 2001 (Bommier *et al.*, 2003). Although these two samples are not directly comparable, a comparison of those results with our study enables us to validate the case-control approach used here.

The calculation of the elasticities requires several steps, given the case-control methodology. The elasticity of the probability of death corresponds to the relative variation of the probability of death p resulting from the relative variation of income $y$.

For a representative individual $i$, recipient of income $y_i$, equal to the average income of the sample, the elasticity corresponds to the marginal effect of income and the relation between average income and the associated probability of death $p_i$:

$$\eta = \frac{\partial p}{\partial y} \times \frac{y_i}{p_i}$$

In the specific case of the LOGIT model, the marginal effect of logarithmic income on death can be deduced simply from the probability of death of the representative individual defined above and the estimated coefficient β corresponding to logarithmic income:

$$\frac{\partial p}{\partial \ln y} = p_i \left(1 - p_i\right) \beta$$

The income elasticity of the probability of death therefore corresponds to:

(3) $$\eta = \left(1 - p_i\right) \beta$$

The calculated elasticity value is therefore a function of the probability of death specified above for the representative individual. However, in our study the relative proportions of deceased and survivors does not permit the calculation of a realistic probability of death. As there are approximately ten times more deceased men in our case-control study, this would result in an underestimate of about 10% of the calculated elasticity. There are two possible ways to correct for this.

The first is to re-weight observations to make the sample more representative of the mortality rate in 1988. This procedure modifies only slightly the value of the coefficients calculated for the parameters, but does change their significance, because this weighting reduces the amount of information included for the deceased. As we noted earlier, this weighting reduces the statistical power and therefore enlarges the confidence intervals of the coefficients. Elasticity is calculated separately for the whole sample, the sample of under 65s, and the sample of over 65s, for an individual with average income for the sub-sample (97 000 FF for the whole sample, 96 000 FF for the under 65s and 102 000 FF for the over 65s), employed, married, resident in the Paris Basin and with average age 65 for the whole sample, 55 for those under 65 and 75 for those over 65.

The second correction consists of approximating the value of the elasticity based on the value of the odds ratio, which, for rare events, is close to that for relative risk.

The estimation $\eta$ of the elasticity $\eta^e$ for a small increase $\Delta y$ of income around the value $y_i$ can be written as a function of relative risk associated with this increase in income:

$$\eta^e = \frac{P(D=1\,|y_i+\Delta y)-P(D=1\,|y_i)}{P(D=1\,|y_i)}\times\frac{\Delta y}{y_i}=(R.R_i-1)\times\frac{\Delta y}{y_i}$$

As the probability of death is very close to 0, the relative risk can be approximated by the value of the odds ratio $O.R._i$ associated with the increase in income $\Delta y$ around the value $y_i$. Hence:

$$\eta^e = (O.R_i-1)\times\frac{\Delta y}{y_i} \tag{4}$$

The elasticity of the probability of death is therefore calculated separately for the whole sample, the sample of under 65s and that of over 65s. This calculation is done without re-weighting the members of the sample for the probability of death in 1988. The value of the odds ratio $O.R_i$ considered corresponded to an annual increase in income by consumption unit of 12 000 FF for the representative individual:

$$O.R._i = \frac{\dfrac{P(D=1|y_i+\Delta y)}{1-P(D=1|y_i+\Delta y)}}{\dfrac{P(D=1|y_i)}{1-P(D=1|y_i)}}$$

These two methods for calculating the income elasticity of the probability of death give fairly similar results. Without controlling for profession (Model 1, Appendix 2, Table 7) the income elasticity of the probability of death is about –0.43 for the whole population, -0.5 for the under 65s, and –0.34 for the over 65s.

Thus the risk of death is strongly associated with income, for all ages, even if this correlation declines with age. The decrease in the effect of income with age can undoubtedly be explained by the selective effect of declining health in the elderly population at risk. This finding partly contradicts the results of McDonough *et al.* (1997) who found that the correlation was substantially lower for the over 65s. It may then be the case that the bigger differential between these two age groups in sickness insurance cover among the poorest population in the US explains the more marked decrease in the effect of income on risk of death. In fact general access to Medicare cover for over 65s has undoubtedly a greater impact on the reduction in health social inequalities in the US, where the poorest do not invariably have access to Medicaid before age 65. These elasticities can be quite high. However they are below the elasticities estimated by Bommier *et al.* (2003), which range from –0.62 for 67 year old men to –0.14 for 91 year old men.

The introduction of occupational status in Model 2 reduces the value of elasticities by about 0.1 (Model 2, Appendix 2, Table 7). The elasticity of the probability of death is about –0.34 for the whole population, -0.4 for the under 65s and finally –0.24 for the over 65s. On the other hand, if pensions and allowances are introduced (Model 3, Appendix 2, Table 7) the value of the elasticities changes very little, this variable reducing the absolute value of the calculated elasticities by about 0.01.

## 4.5 **Impact of Very High and Very Low Income**

A comparison of Models 3 to 7 enables a more specific analysis of the shape of the relationship between mortality and income.

Specifying income in quintiles (Models 6 and 7, Table 9) shows a specific risk related to poverty, independently of working conditions and culturally-specific behaviour, approximated by occupational status. The mortality risk for the first lowest quintile is significantly higher than that for all other income levels, for all age groups. Without controlling for profession, the mortality risk for the first quintile is 2.5 times greater than that of the fifth quintile (Model 6, Appendix 2, Table 9). After adjusting for occupational status status, the relative risk for the first quintile is till twice as high (Model 7, Appendix 2, Table 9).

However, our results invalidate the absolute poverty hypothesis, because the association between income and mortality is not confined to the lower end of the income distribution. The value of the odds ratios associated with income quintiles shows more clearly the extent of social health inequality in France. In line with the analysis of McDonough *et al.* (1997), the probability of death decreases across the whole income distribution.

Nor is the hypothesis of decreasing returns of the health capital production function proposed by Erlich and Chuma (1990), and which underlies Gravelle's statistical artefact (1996), confirmed by our analysis. Specification under a logarithmic form (Models 2 and 3, Appendix 2, Table 6) and under a quadratic form (Model 4, Appendix 2, Table 8) might suggest a concave relation between income and health because the coefficients of income and income squared are significant and of opposite signs. However the cubic (Model 5, Appendix 2, Table 8) and quintile specifications (Models 6 and 7, Appendix 2, Table 9) invalidate this hypothesis. The cubic form suggests that the health and income relationship does in fact have a logistical form. The results of Model 6 confirm that the marginal effect of income is only reduced in the middle of the distribution. The mortality risks are not significantly different for the second and third income quintiles ($2^{nd}$ and $3^{rd}$ lowest), but they are for the third and the fourth quintile. Furthermore we show that the-mortality rates are significantly lower for those with the highest incomes. Whatever the age groups studied, the risk of death for the fourth quintile is significantly higher than that of the fifth quintile (the highest income group). However, when we control for the occupational status, the risk of death of the fourth quintile is still significantly lower than the risk of death of the third quintile. Nevertheless, the risk of death for the fifth quintile is not significantly lower than the risk of death of the fourth quintile (Model 7, Appendix 2, Table 9). This could be explained by the fact that the last quintile is characterised by a high proportion of executives.

The correlation between poverty and mortality could undoubtedly be explained by differences in access to care or financials hardship. The reduction of the risk of mortality associated to increase of income beyond the third quintile is more surprising. The difference of average income is rather weak between the third and the fourth quintile and differences of material living conditions cannot obviously explain the decrease of the risk of death. So the question is no longer why poverty kills, but rather why highest incomes protect.

# 5 Discussion

Using a case control methodology, this study provides an analysis of the relationship between income and the risk of death for male population in France.

We found a strong correlation between the risk of death at any age and income. This correlation is robust even when we control for occupational status. Therefore, our results suggest that there is a distinct impact of income on health, independent from education and working conditions as proxied by occupational status.

Moreover the relation between income and health holds for all levels of income distribution. Our results suggest a strong effect of poverty on mortality risk as well as a protective effect of the highest incomes on health. In other words, our results rejected the hypothesis of diminishing returns of income on health. This unexpected association between health and high income levels has not so far been traced in the literature. However, this finding does not appear to be the result of a sampling bias between the two surveys, as the results by quintile are robust after excluding extreme values[9].

According to the health capital model, income affects health because more health care is consumed. Empirical analyses of heath care consumption in countries where access to health care is guaranteed almost for all, do not show a strong relation between the volume of care consumed and income. In general the effect of health services use on health status is not established, except for people with very low incomes in very poor states of health. Therefore, it seems difficult to explain the income-health relation demonstrated here entirely by the relative use of health

9. Estimations for the sample reduced to the first 98 centiles of the distribution of income by consumption unit in the Taxable Income Survey give equivalent results.

care services. On the one hand the most serious health problems are covered by social security and on the other the problem of access to supplementary insurance cover affect principally those at the bottom end of the distribution (Dourgnon *et al.*, 2001).

Even if problems of access to care can partly explain why poverty kills, in addition to financial hardship, it seems to us that the protective effect of highest incomes is less readily explained in terms of differences in material conditions. A first line of enquiry might be to look at differences in the quality of care consumed, and particularly rates of specialist consultation. Further investigation of income could also be interesting. The continuous effect of income on health might be in part due to the effect of relative social position on the risk of death, as income quintile reflects an individual's position in the society. It might also be that the availability of financial resources (being rich) has a distinct positive impact on health. This is consistent with the pathogenic effect of stress caused by a person's low position in a hierarchy and lack of control of his environment, (Marmot, 2000). Taken together these results call for more work on the impact of social interactions on health.

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# Appendix 1 : The Case Control Methodology

Suppose that the unobservable individual health capital Hi is explained by the observable characteristic x and a residual u, distributed according to a logistical distribution.

$$H = \alpha + \beta x + u$$

Suppose that death occurs when the health capital drops below a critical minimum level Hmin. Assuming that Hmin is equal to zero, the probability of death conditionally to the characteristic x, $\pi(x)$ is therefore equal to:

$$\pi(x) = \mathrm{P}(\mathrm{D}\,|\mathrm{x}) \;= P(H < H\min) = \frac{1}{1 + \exp(\alpha + \beta x)}$$

It is straightforward to show that in the case of a LOGIT model, the estimated value of the $\beta$ parameters is invariant to the overrepresentation of the deceased people in the total sample (Leblanc, Lollivier, Marpsat, Verger, 2000).

Suppose that the study sample is selected according to vital status. We note $\rho_0$ the sampling probability conditionally to the death and $\rho_1$ the sampling probability conditionally to survival:

$$\rho_0 = \mathrm{P}(\mathrm{S}\,|\mathrm{D}) \quad \text{and} \quad \rho_1 = \mathrm{P}(\mathrm{S}\,|\overline{\mathrm{D}})$$

According to Bayes rule, the probability of death conditionally to the characteristic x and to the sampling, $\mu(x)$, is as follows:

$$\mu(x) = \mathrm{P}(\mathrm{D}|\mathrm{S,x}) \;= \frac{\mathrm{P}(\mathrm{S}\,|\mathrm{D,x}) \times \mathrm{P}(\mathrm{D}|\mathrm{x})}{\mathrm{P}(\mathrm{S}|\mathrm{x})}$$

$$\mu(x) = \frac{\rho_0 \times \pi(x)}{\rho_0 \times \pi(x) + \rho_1 \times (1 - \pi(x))}$$

Hence:

$$\mu(x) = \frac{1}{1 + \dfrac{\rho_1}{\rho_0} \times \exp(\alpha + \beta x)} = \frac{1}{1 + \exp\left(\alpha + \log\left(\dfrac{\rho_1}{\rho_0}\right) + \beta x\right)}$$

The respective size of each sub-population does not bias the value of the parameter β and only influences the value of the constant α*, with

$$\alpha^* = \alpha + \log\left(\frac{\rho_1}{\rho_0}\right)$$

The determinants of the probability of death can also be assessed for the sample selected from the Wealth at Death Survey and Household Taxable Income Survey, without weighting the samples from each survey in order to obtain a realistic estimate of the probability of death as, for a global sample size (number of cases plus number of controls), the maximum power is achieved where the number of cases is equal to the number of controls (Schlesselman, 1982; Bouyer *et al.*, 1995).

Given the low probability of death (1.4% for men aged over 20 in 1988) the estimated odds ratio, equal to the exponential of the parameter β, may be interpreted as the relative risk of death with regard to the reference category. Indeed, it is straightforward to show that the value of the relative risk RR associated with the variable x tends towards the value of the odds ratio OR, when the probability of the event considered tends towards zero (Cornfield, 1951).

For a dichotomous variable x, the relative risk of death is equal to:

$$RR = \frac{P(D|x=1)}{P(D|x=0)} \frac{\dfrac{P(x=1|D)\times P(D)}{P(x=1|D)\times P(D) + P(x=1|\overline{D}) \times (1-P(D))}}{\dfrac{P(x=0|D)\times P(D)}{P(x=0|D)\times P(D) + P(x=0|\overline{D}) \times (1-P(D))}}$$

Hence RR tends towards the value of the odds ratio OR, *i.e.* the ratio of likelihoods between individuals with characteristic x = 1 and individuals with characteristic x = 0:

$$\lim_{P(D)\to 0} RR = \frac{\dfrac{P(x=1|D)}{P(x=1|\overline{D})}}{\dfrac{P(x=0|D)}{P(x=1|\overline{D})}} = \frac{\dfrac{P(D|x=1)}{P(\overline{D}|x=1)}}{\dfrac{P(D|x=0)}{P(\overline{D}|x=1)}} = OR = \exp\beta$$

# Appendix 2: Results

TABLE 1

***Description of the study sample (without tacking into account the weighting of each survey)***

| | **Survivor sample** | | **Deceased sample** | |
|---|---|---|---|---|
| **Characteristics** | number | frequency | number | frequency |
| Age < 40 | 4605 | 38.% | 57 | 4.0 % |
| 40-54 | 3866 | 32.3% | 140 | 9.7% |
| 55-64 | 1811 | 15.1% | 250 | 17.4% |
| 65-74 | 1069 | 8.9% | 317 | 22.0% |
| 75-84 | 526 | 4.4% | 446 | 31.0% |
| 85 and over | 84 | 0.7% | 228 | 15.9% |
| Quintile 1 | 2098 | 17.5% | 313 | 21.8% |
| Quintile 2 | 2086 | 17.4% | 373 | 25.9% |
| Quintile 3 | 2204 | 18.4% | 312 | 21.7% |
| Quintile 4 | 2511 | 21.0% | 236 | 16.4% |
| Quintile 5 | 3062 | 25.6% | 204 | 14.2% |
| Manual | 4328 | 36.2% | 752 | 52.3% |
| Employee | 1672 | 14.0% | 309 | 21.5% |
| Intermediate | 3092 | 25.9% | 238 | 16.6% |
| Executive | 2869 | 24.0% | 139 | 9.7% |
| Married | 9960 | 80.8% | 981 | 68.2% |
| Single | 1487 | 12.4% | 118 | 8.2% |
| Widower | 247 | 2.1% | 275 | 19.1% |
| Divorce | 567 | 4.7% | 64 | 4.5% |
| Ile-de-France | 2432 | 20.3% | 241 | 16.8% |
| Bassin Parisien | 2298 | 19.2% | 285 | 19.8% |
| Nord | 826 | 6.9% | 123 | 8.6% |
| Est | 1118 | 9.3% | 123 | 10.5% |
| Ouest | 1513 | 12.6% | 184 | 12.8% |
| Sud-Ouest | 1148 | 9.6% | 126 | 8.8% |
| Centre-Est | 1438 | 12.0% | 164 | 11.4% |
| Méditerranée | 1188 | 9.9% | 164 | 11.4% |
| All | 11961 | 100% | 1438 | 100% |

TABLE 2
***Income per C.U. distribution by occupational status***

| | Quintile 1 | Quintile 2 | Quintile 3 | Quintile 4 | Quintile 5 | All |
|---|---|---|---|---|---|---|
| Manual | 1671 (32.9%) | 1395 (27.5%) | 1117 (22.0%) | 683 (13.4%) | 214 (4.2%) | 5080 |
| Employee | 336 (17.0%) | 465 (23.5%) | 485 (24.5%) | 448 (22.6%) | 247 (12.5%) | 1981 |
| Intermediate | 304 (9.1%) | 484 (14.5%) | 675 (20.3%) | 959 (28.8%) | 908 (27.3%) | 3330 |
| Executive | 100 (3.3%) | 115 (3.8%) | 239 (7.9%) | 657 (21.8%) | 1897 (63.1%) | 3008 |
| All sample | 2411 | 2459 | 2516 | 2747 | 3266 | 13399 |

TABLE 3
***Income per C.U. distribution by vital status***

| | Total sample | | Under 65s | | Over 65s | |
|---|---|---|---|---|---|---|
| Income per U.C. | survivors | deceased | survivors | deceased | survivors | deceased |
| Mean | 97198 | 83842 | 96032 | 77013 | 103019 | 87017 |
| Standard deviation | 67737 | 56186 | 67888 | 52885 | 78079 | 57406 |
| Decile 1 (max value) | 37893 | 37727 | 36471 | 27321 | 46514 | 41484 |
| Decile 2 | 51471 | 48771 | 49934 | 40913 | 58129 | 52039 |
| Decile 3 | 62413 | 56627 | 61438 | 51167 | 67993 | 59039 |
| Decile 4 | 72721 | 64235 | 72061 | 57250 | 75775 | 67364 |
| Median | 83027 | 73062 | 82724 | 65216 | 84727 | 76187 |
| Decile 6 | 94416 | 81588 | 93890 | 74998 | 96437 | 83804 |
| Decile 7 | 109357 | 92116 | 109233 | 85500 | 109891 | 94440 |
| Decile 8 | 129231 | 109852 | 128962 | 105242 | 130367 | 111411 |
| Decile 9 | 166923 | 134864 | 166168 | 132104 | 170721 | 137163 |
| Decile 10 | 1727560 | 816855 | 1727560 | 612009 | 1176817 | 816855 |

TABLE 4
***Distribution of deceased and survivors by equivalent income quintile (without tacking into account the weighting of each survey)***

| | | Total sample | | Under 65s | | Over 65s | |
|---|---|---|---|---|---|---|---|
| Income quintile | Max value | deceased | survivors | deceased | survivors | deceased | survivors |
| Quintile 1 | 51413 | 21.8% | 17.5% | 27.7% | 17.3% | 25.9% | 17.6% |
| Quintile 2 | 72793 | 25.9% | 17.4% | 27.7% | 17.6% | 21.7% | 17.9% |
| Quintile 3 | 94203 | 21.7% | 18.4% | 18.3% | 18.4% | 21.2% | 18.7% |
| Quintile 4 | 128948 | 16.4% | 21.0% | 13.0% | 21.0% | 16.6% | 21.1% |
| Quintile 5 | 1712560 | 14.2% | 25.6% | 13.2% | 25.8% | 14.5% | 24.7% |
| All | | 1438 | 11961 | 447 | 10282 | 991 | 1679 |

TABLE 5
***Probability of death in 1988 (LOGIT) (men over 20's): models 0 and 1***

| | Model 0 | | | Model 1 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef.[1] | O.R[2] | C.I. 95%[3] | Coef. | O.R | C.I. 95% |
| Log equivalent income | | | | -0.466*** | 0.628 | (0.569-0.692) |
| Manual | 1.035*** | 2.814 | (2.177 - 3.637) | | | |
| Employee | 1.069*** | 2.911 | (2.208 - 3.838) | | | |
| Intermediate | 0.776*** | 2.173 | (1.641 - 2.878) | | | |
| Executive | ref | 1 | | | | |
| Under 40 | 0.046** | 1.047 | (1.008 - 1.087) | 0.042** | 1.043 | (1.004 - 1.083) |
| 40-55 | 0.112*** | 1.119 | (1.091 - 1.147) | 0.118*** | 1.125 | (1.098 - 1.154) |
| 55-65 | 0.076*** | 1.079 | (1.050 - 1.109) | 0.081*** | 1.085 | (1.055 - 1.115) |
| 65-75 | 0.057*** | 1.059 | (1.031 - 1.087) | 0.059*** | 1.060 | (1.033 - 1.088) |
| 75-85 | 0.123*** | 1.130 | (1.094 - 1.168) | 0.114*** | 1.121 | (1.085 - 1.158) |
| Over 85 | 0.074** | 1.077 | (1.003 - 1.156) | 0.074** | 1.077 | (1.003 - 1.156) |

TABLE 5 (CONTINUED)

| | Model 0 | | | Model 1 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef.[1] | O.R[2] | C.I. 95%[3] | Coef. | O.R | C.I. 95% |
| Married | ref | 1 | | ref | 1 | |
| Single | 0.483*** | 1.620 | (1.288 - 2.038) | 0.438*** | 1.533 | (1.217 – 1.933) |
| Widower | 0.702*** | 2.018 | (1.636 - 2.491) | 0.731*** | 2.077 | (1.685 - 2.562) |
| Divorce | 0.402*** | 1.495 | (1.120 - 1.996) | 0.358*** | 1.430 | (1.071 - 1.911) |
| Ile-de-France | -0.056 | 0.945 | (0.766 - 1.167) | -0.036 | 0.965 | (0.781 - 1.191) |
| Bassin Parisien | ref | 1 | | ref | 1 | |
| Nord | 0.299** | 1.349 | (1.040 - 1.749) | 0.263** | 1.301 | (1.004 - 1.685) |
| Est | 0.148 | 1.160 | (0.914 - 1.471) | 0.133 | 1.142 | (0.900 - 1.449) |
| Ouest | 0.121 | 1.129 | (0.903 - 1.410) | 0.100 | 1.105 | (0.885 - 1.381) |
| Sud-Ouest | -0.203 | 0.816 | (0.635 - 1.050) | -0.248* | 0.780 | (0.606 - 1.004) |
| Centre-Est | -0.012 | 0.988 | (0.782 - 1.249) | -0.018 | 0.982 | (0.776 - 1.241) |
| Méditerranée | -0.137 | 0.872 | (0.689 - 1.104) | -0.189 | 0.827 | (0.653 - 1.047) |
| Constante | -5.009 | | | 1.045 | | |
| -2 Log L (constante) | 9134.858 | | | 9134.86 | | |
| -2 Log L | 6635.167 | | | 6635.21 | | |

1. Significance level: * 10%, ** 5%, *** 1%.
2. Odds ratio.
3. 95% confidence interval for conditional odds ratio.

TABLE 6
***Probability of death in 1988 (LOGIT) (men over 20's): models 2 and 3***

| | Model 2 | | | Model 3 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef. | O.R | C.I. 95% | Coef. | O.R | C.I. 95% |
| Log equivalent income | -0.366*** | 0.693 | (0.620 – 0.775) | -0.347*** | 0.707 | (0.632-0.792) |
| Manual | 0.719*** | 2.053 | (1.561 – 2.700) | 0.726*** | 2.067 | (1.570 – 2.720) |
| Employee | 0.858*** | 2.358 | (1.775 - 3.133) | 0.829*** | 2.290 | (1.723 - 3.045) |
| Intermediate | 0.631*** | 1.880 | (1.414 - 2.500) | 0.641*** | 1.899 | (1.427 - 2.527) |
| Executive | ref | 1 | | ref | 1 | |
| Pension under 60s | | | | 0.011*** | 1.011 | (1.007 - 1.015) |
| Under 40 | 0.045** | 1.046 | (1.007 - 1.087) | 0.048** | 1.049 | (1.010 - 1.089) |
| 40-55 | 0.115*** | 1.122 | (1.095 - 1.151) | 0.099*** | 1.104 | (1.076 - 1.133) |
| 55-65 | 0.079*** | 1.082 | (1.053 - 1.113) | 0.103*** | 1.108 | (1.076 - 1.141) |
| 65-75 | 0.057*** | 1.059 | (1.031 - 1.087) | 0.056*** | 1.058 | (1.030 - 1.086) |
| 75-85 | 0.118*** | 1.125 | (1.089 - 1.162) | 0.118*** | 1.125 | (1.089 - 1.162) |
| Over 85 | 0.076** | 1.079 | (1.004 - 1.159) | 0.076** | 1.079 | (1.005 - 1.159) |
| Married | ref | 1 | | ref | 1 | |
| Single | 0.417*** | 1.517 | (1.204 - 1.911) | 0.395*** | 1.485 | (1.178 – 1.871) |
| Widower | 0.709*** | 2.032 | (1.645 - 2.509) | 0.701*** | 2.017 | (1.633 - 2.490) |
| Divorce | 0.349** | 1.418 | (1.060 - 1.897) | 0.320** | 1.377 | (1.028 - 1.843) |

TABLE 6 (CONTINUED)

| | Model 2 | | | Model 3 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef. | O.R | C.I. 95% | Coef. | O.R | C.I. 95% |
| Ile-de-France | 0.008 | 1.009 | (0.815 - 1.247) | -0.021 | 1.021 | (0.825 - 1.264) |
| Bassin Parisien | Ref | 1 | | ref | 1 | |
| Nord | 0.267** | 1.306 | (1.008 - 1.693) | 0.242* | 1.274 | (0.982 - 1.653) |
| Est | 0.138 | 1.148 | (0.905 - 1.457) | 0.131 | 1.140 | (0.898 - 1.448) |
| Ouest | 0.105 | 1.110 | (0.888 - 1.388) | 0.091 | 1.095 | (0.875 - 1.369) |
| Sud-Ouest | -0.243* | 0.784 | (0.608 - 1.010) | -0.240* | 0.779 | (0.605 - 1.004) |
| Centre-Est | -0.009 | 0.991 | (0.783 - 1.253) | 0.009 | 1.009 | (0.797 - 1.277) |
| Méditerranée | -0.171 | 0.843 | (0.665 - 1.068) | -0.182 | 0.834 | (0.658 - 1.057) |
| Constante | -0.696 | | | -0.889 | | |
| -2 Log L (constante) | 9134.86 | | | 9134.86 | | |
| -2 Log L | 6596.25 | | | 6665.69 | | |

TABLE 7
***Income elasticity of the probability of death in 1988***

| | Total sample | | | Under 65s | | | Over 65s | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 | Model 2 | Model 3 | Model 1 | Model 2 | Model 3 | Model 1 | Model 2 |
| $\beta$ | -0.438 | -0.355 | -0.343 | -0.517 | 0.444 | -0.430 | -0.356 | -0.253 |
| $p_i$ | 0.0210 | 0.0204 | 0.0210 | 0.0091 | 0.0074 | 0.0062 | 0.0303 | 0.0325 |
| Marginal effect | -0.009 | -0.007 | -0.007 | -0.005 | -0.003 | -0.003 | -0.010 | -0.008 |
| $\eta$ | -0.429 | -0.347 | -0.335 | 0.512 | -0.441 | -0.428 | -0.345 | -0.244 |
| $O.R._i$ | 0.947 | 0.958 | 0.961 | 0.938 | 0.949 | 0.952 | 0.961 | 0.972 |
| $\eta^e$ | -0.428 | -0.338 | -0.317 | -0.498 | -0.409 | -0.386 | -0.334 | -0.236 |

TABLE 8
***Probability of death in 1988 (LOGIT) (men over 20's): models 4 and 5***

| | Model 4 | | | Model 5 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef. | O.R | C.I. 95% | Coef. | O.R | C.I. 95% |
| Income | -0.005*** | 0.995 | (0.993 – 0.997) | -0.011*** | 0.989 | (0.985-0.993) |
| $Income^2$ | 4 E-6*** | 1.000 | (1.000 – 1.000) | 3 E-5*** | 1.000 | (1.000 – 1.000) |
| $Income^3$ | | | | -2 E-8** | 1.000 | (1.000 – 1.000) |
| Manual | 0.637*** | 1.891 | (1.417 – 2.522) | 0.660*** | 1.935 | (1.442 – 2.596) |
| Employee | 0.736*** | 2.087 | (1.555 – 2.802) | 0.793*** | 2.210 | (1.636 – 2.987) |
| Intermediate | 0.537*** | 1.763 | (1.317 - 2.360) | 0.636*** | 1.889 | (1.402 - 2.546) |
| Executive | ref | 1 | | ref | 1 | |
| Pension under 60s | 0.011*** | 1.011 | (1.011 - 1.015) | 0.011*** | 1.011 | (1.007 - 1.014) |
| Under 40s | 0.049** | 1.050 | (1.011 - 1.091) | 0.049** | 1.050 | (1.011 - 1.091) |
| 40-55 | 0.099*** | 1.104 | (1.076 - 1.133) | 0.100*** | 1.105 | (1.077 - 1.134) |
| 55-65 | 0.100*** | 1.105 | (1.073 - 1.138) | 0.101*** | 1.106 | (1.074 - 1.140) |
| 65-75 | 0.056*** | 1.058 | (1.030 - 1.086) | 0.056*** | 1.057 | (1.030 - 1.085) |
| 75-85 | 0.117*** | 1.124 | (1.088 - 1.162) | 0.117*** | 1.125 | (1.088 - 1.162) |
| Over 85s | 0.079** | 1.082 | (1.007 - 1.163) | 0.075** | 1.078 | (1.004 - 1.158) |
| Married | ref | 1 | | ref | 1 | |
| Single | 0.405*** | 1.500 | (1.191 - 1.889) | 0.388*** | 1.474 | (1.169 – 1.857) |
| Widower | 0.702*** | 2.018 | (1.634 - 2.492) | 0.709*** | 2.031 | (1.644 - 2.509) |
| Divorce | 0.325** | 1.384 | (1.034 - 1.853) | 0.314** | 1.370 | (1.023 - 1.834) |

TABLE 8 (CONTINUED)

| | Model 4 | | | Model 5 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef. | O.R | C.I. 95% | Coef. | O.R | C.I. 95% |
| Ile-de-France | 0.039 | 1.040 | (0.839 - 1.288) | 0.036 | 1.037 | (0.837 - 1.284) |
| Bassin Parisien | ref | 1 | | ref | 1 | |
| Nord | 0.241* | 1.273 | (0.981 - 1.651) | 0.237* | 1.267 | (0.977 - 1.644) |
| Est | 0.127 | 1.136 | (0.895 - 1.442) | 0.129 | 1.138 | (0.896 - 1.445) |
| Ouest | 0.090 | 1.094 | (0.875 - 1.368) | 0.086 | 1.090 | (0.872 - 1.363) |
| Sud-Ouest | -0.241* | 0.786 | (0.610 - 1.012) | -0.253* | 0.777 | (0.602 - 1.001) |
| Centre-Est | 0.009 | 1.009 | (0.797 - 1.278) | 0.009 | 1.009 | (0.796 - 1.277) |
| Méditerranée | -0.179 | 0.836 | (0.659 - 1.060) | -0.188 | 0.829 | (0.653 - 1.052) |
| Constante | -4.261 | | | -4.032 | | |
| -2 Log L (constante) | 9134.86 | | | 9134.86 | | |
| -2 Log L | 6567.51 | | | 6555.33 | | |

TABLE 9

***Probability of death in 1988 (LOGIT) (men over 20's): models 6 and 7***

| | Model 6 | | | Model 7 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef. | O.R | C.I. 95% | Coef. | O.R | C.I. 95% |
| Quintile 1 | 0.916*** | 2.499 | (1.995 - 3.130) | 0.688*** | 1.989 | (1.542 - 2.566) |
| Quintile 2 | 0.793*** | 2.211 | (1.778 - 2.749) | 0.557*** | 1.745 | (1.368 - 2.226) |
| Quintile 3 | 0.622*** | 1.862 | (1.495 - 2.319) | 0.394*** | 1.482 | (1.166 - 1.884) |
| Quintile 4 | 0.358*** | 1.431 | (1.139 - 1.798) | 0.175 | 1.191 | (0.938 - 1.513) |
| Quintile 5 | ref | 1 | | ref | 1 | |

TABLE 9 (CONTINUED)

| | Model 6 | | | Model 7 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef. | O.R | C.I. 95% | Coef. | O.R | C.I. 95% |
| Manual | | | | 0.670*** | 1.955 | (1.466 - 2.607) |
| Employee | | | | 0.797*** | 2.219 | (1.651 – 2.981) |
| Intermediate | | | | 0.642*** | 1.901 | (1.421 - 2.542) |
| Executive | | | | ref | 1 | |
| Pension under 60s | 0.011*** | 1.011 | (1.007 - 1.015) | 0.011*** | 1.011 | (1.007 - 1.014) |
| Under 40s | 0.049** | 1.050 | (1.011 - 1.091) | 0.050*** | 1.052 | (1.012 - 1.092) |
| 40-55 | 0.101*** | 1.106 | (1.078 - 1.134) | 0.099*** | 1.103 | (1.075 - 1.132) |
| 55-65 | 0.101*** | 1.106 | (1.074 - 1.139) | 0.101*** | 1.106 | (1.073 - 1.139) |
| 65-75 | 0.057*** | 1.058 | (1.031 - 1.087) | 0.055*** | 1.057 | (1.029 - 1.085) |
| 75-85 | 0.115*** | 1.122 | (1.086 - 1.160) | 0.118*** | 1.126 | (1.089 - 1.163) |
| Over 85s | 0.074** | 1.077 | (1.003 - 1.156) | 0.074** | 1.077 | (1.003 - 1.157) |
| Married | ref | 1 | | ref | 1 | |
| Single | 0.420*** | 1.522 | (1.209 - 1.918) | 0.405*** | 1.499 | (1.190 - 1.889) |
| Widower | 0.728*** | 2.071 | (1.678 - 2.556) | 0.710*** | 2.033 | (1.646 - 2.512) |
| Divorce | 0.335** | 1.398 | (1.046 - 1.869) | 0.326** | 1.386 | (1.035 - 1.855) |

TABLE 9 (CONTINUED)

| | Model 6 | | | Model 7 | | |
|---|---|---|---|---|---|---|
| Characteristics | Coef. | O.R | C.I. 95% | Coef. | O.R | C.I. 95% |
| Ile-de-France | 0.016 | 1.016 | (0.821 - 1.257) | 0.039 | 1.039 | (0.839 - 1.288) |
| Bassin Parisien | ref | 1 | | ref | 1 | |
| Nord | 0.229* | 1.258 | (0.969 - 1.631) | 0.237* | 1.267 | (0.977 - 1.645) |
| Est | 0.119 | 1.127 | (0.887 - 1.430) | 0.129 | 1.138 | (0.896 - 1.445) |
| Ouest | 0.080 | 1.083 | (0.867 - 1.355) | 0.084 | 1.088 | (0.870 - 1.361) |
| Sud-Ouest | -0.250* | 0.779 | (0.605 - 1.003) | -0.249* | 0.780 | (0.605 - 1.005) |
| Centre-Est | -0.004 | 0.996 | (0.787 - 1.262) | 0.006 | 1.006 | (0.794 - 1.274) |
| Méditerranée | -0.193 | 0.824 | (0.651 - 1.044) | -0.182 | 0.834 | (0.657 - 1.058) |
| Constante | -4.702 | | | -5.119 | | |
| -2 Log L (constante) | 9134.86 | | | 9134.86 | | |
| -2 Log L | 6592.14 | | | 6561.34 | | |