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Martin Karlsson, Les Mayhew & Ben Rickayzen.

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Cass Business School
106 Bunhill Row
London EC1Y 8TZ
T +44 (0)20 7040 8470
www.cass.city.ac.uk

In Sickness and in Health?

Dynamics of Health and Cohabitation in the United Kingdom

Martin Karlsson[a] Les Mayhew[b] Ben Rickayzen[b]

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Abstract

The purpose of this paper is to analyse the dynamics of cohabitation and functional impairments among older people. Our research has three main aims. Firstly, we want to analyse the effects of cohabitation on disability. Secondly, we want to study time trends in disability and cohabitation jointly to explore relationships between the two. Thirdly, we examine socioeconomic differences -- as captured by educational attainment -- in disability.

These issues are of great interest from several points of view. Firstly, they address an emerging theoretical debate concerning the effects of cohabitation on health and contribute to a sparse empirical literature on the topic. Secondly, our findings are highly policy relevant. Concerning long-term care for older people, for example, cohabitation is of double importance: firstly, since people who cohabit tend to be healthier, and secondly, since a partner is the typical provider of informal care. In a time where family structures among the old are likely to change (due to changes in life expectancy and divorce rates), our research will be useful for planning purposes. Finally, the model can be used to simulate populations of certain characteristics. Hence, it can be used to derive insurance premiums in order to reduce the problem of selection effects in the market for long-term care insurance.

Using the British Household Panel Survey dataset, we apply panel data and simulation techniques to exploit the longitudinal characteristic of the panel. We estimate the two dependent variables -- cohabitation status and disability -- jointly, and allow for time trends, age effects and unobserved heterogeneity.

We find that there are systematic differences between single and cohabiting people so that a cross sectional analysis would overestimate the causal relationship; nevertheless, cohabitation has a strong and positive effect on health. Furthermore, we find that bereavement of a partner has a significant negative impact on health.

Keywords: disability, cohabitation, maximum simulated likelihood

1 Introduction

Most developed countries' populations are ageing rapidly with consequent implications for public spending on long-term care (LTC), pensions and health care. The UK dependency ratio (the

[a] Institute of Ageing, University of Oxford. Corresponding author: martin.karlsson@ageing.ox.ac.uk.

[b] Cass Business School, City University

number of retired people per 100 people of working age) is projected to increase from 24 today to 38 in 2040. Although substantial, the increase is lower than in many other countries. In Japan, for instance, the ratio is projected to increase from 30 today to 65 in 2040 (United Nations, 2002).

Such demographic changes are expected to have a significant impact on the demand for LTC. Most consumers of LTC are over age 80; for example, in England, almost 80 per cent of care home inhabitants belong to this age group (Bajekal, 2002). Since increasing life expectancy causes this group to grow at a faster rate than the general retired population, there is concern that the demographic burden could make the current system of financing LTC unsustainable. Indeed, in the UK, there is already a trend towards concentrating resources only on individuals with severe disability (Karlsson et al, 2004).

Still, relatively little is known about long-term trends and the determinants of the disablement process. One important issue that has not yet been resolved is the long-term trends in healthy life expectancy and disabled life expectancy. Three competing hypotheses have been proposed. The most optimistic one, suggesting a compression of morbidity, was proposed by Fries (1980). According to this perspective, adult life expectancy is approaching its biological limit so that if disability spells can be postponed to higher ages the result will be an overall reduction in the time spent disabled. By contrast, Gruenberg (1977) suggested an expansion of morbidity based on the argument that the observed decline in mortality was mainly due to falling accident rates. The third hypothesis was proposed by Manton (1987) according to whom the development in mortality and morbidity is a combination of the two, which could lead to an expansion of the time spent in good health as well as the time spent in disability.

Official statistics, however, are surprisingly inconclusive as to which of the three hypotheses prevails in reality (Bone et al, 1995, Bebbington & Darton, 1996, Bebbington and Comas-Herrera, 2000). In general, results seem to be sensitive to the definition of disability (activities of daily living (ADLs) or limiting long-standing illness) as well as to the severity of disability taken into account. Despite this ambiguity in the statistics, the long-term trends have very strong implications for the future funding of long-term care. In a long-term projection model, Karlsson et al (2006) find that an optimistic scenario ('compression of morbidity') implies some 2 million disabled older people fewer than the most pessimistic scenario ('expansion of morbidity'). The implications for public finances are similar: in the pessimistic scenario, the element of the tax rate necessary to finance formal long-term care would have to increase by around 80 per cent of its present level, whereas virtually no increase would be necessary in the optimistic scenario. Similar differences arise in the supply of demand for informal care (i.e. unpaid care provided by spouses, children or other members of the local community): with an optimistic scenario, there is virtually no shortfall of informal carers in the next few decades, whereas the pessimistic scenario leads to a serious deficit of informal care that will eventually strain public finances.

This paper focuses on one particular aspect of the disablement process, namely the effects of cohabitation on disability.[1] Cohabitation is of particular importance for several reasons. Firstly, cohabitation is strongly correlated with health (a relationship which seems to be stronger for higher ages; Lillard and Panis, 1996) and it is of great interest to know whether this correlation reflects a causal effect -- so that changing cohabitation patterns would have implications for health -- or merely reflects self-selection into and out of cohabitation (i.e. people who cohabit are

[1] In this paper, we refer to 'cohabitation' as a wide concept including all types of cohabitation between partners - including married couples.

healthier at the outset). Separating causation and correlation leads to a host of methodological challenges that will be dealt with below.

Secondly, knowing the relationship between cohabitation and disability is important for analysing the implications of ageing for long-term care. Informal care comprises a substantial part of total long-term care resources and around 75 per cent of all LTC recipients receive informal care according to Karlsson et al (2006). It is a common concern that there may be a shortage of informal carers if certain discernible trends carry on in the future. These trends are, inter alia, the increase in single person households, the rising number of childless older people and the increase in the proportion of females in paid employment. It should be noted, however, that there are some trends that could be expected to countervail these threats to informal care provision. These could include, for instance, a decreasing age at which people retire, together with an improvement in health among younger retirees. This scenario implies that there will be a larger pool of able retirees available in the future to provide informal care; however, the opposite could also apply with an increasing retirement age. Another factor could be changing social values leading to increased male participation in this traditionally female activity.

One good source of information on informal care is the General Household Survey, which offers comparisons over time by studying different cohorts. Previous research (Pickard, 2002) shows that, as expected, the composition of the informal care provision has changed markedly over the last 15 years. There has been a marked drop in the provision of informal care coming from outside the household, whereas the proportion of people providing care within their own household has remained more or less constant. Overall, there has been a significant decrease in the number of people providing care to parents or parents-in-law, whereas the provision of care to spouses has increased significantly. Hence, research should be focused on the role of spouses in the provision of informal long-term care.

In this paper, we make use of all available waves of the British Household Panel Survey (BHPS) data in order to study the determinants of disability and cohabitation of males[2] over time. Our research has three main aims. Firstly, we want to analyse the effects of cohabitation on disability. Secondly, we want to study time trends in disability and cohabitation jointly to explore relationships between the two. Thirdly, we examine socioeconomic differences -- as captured by educational attainment -- in disability. Educational attainment is particularly important in this context. Firstly, it is very convenient as a socioeconomic indicator as it normally remains constant over most of the life course. Secondly, it is a well established result in health economics and epidemiology that education is an important factor in explaining socioeconomic differences in health (cf Fuchs, 2004). Thirdly, empirical studies of marital matching indicate that education is an important aspect of a person's 'marriageability' (Wong, 2003). Hence, excluding education might lead to an overestimation of the importance of cohabitation status for health.

There are two main methodological challenges. The first one is that we seek to estimate a dynamic model, where previous disability and cohabitation status influence current disability and cohabitation status. The second one is that we seek to distinguish causation from correlation in the relationship between cohabitation and disability. This requires using simulation techniques that allow for systematic differences between cohabiting and non-cohabiting individuals which are not discernible in the data. For instance, it might be that healthier people are considered more attractive, and our method is one way to correct for this type of reverse causality.

[2] We focus on men since, according to Wilson and Oswald (2005), health effects of cohabitation are more pronounced for men.

There is little previous empirical research in the field. Brown (2000) performs a simple empirical analysis of the National Survey of Families and Households (waves 87-88), estimating the effects of cohabitation and relationship characteristics, allowing for self-selection into cohabitation, on psychological well being. Brown found no evidence of self-selection, but observed that simple cohabitation is less beneficial to psychological health than marriage. The main explanation seems to be poorer relationship quality in cohabitation relationships.

Cheung (2000) looked at cohabitation and mortality amongst British women. Analysing the Health and Lifestyle Survey, Cox regressions were used in order to allow for self-selection into cohabitation. This is one of few studies allowing for reverse causality from health to marital status (i.e. people being healthy having a higher propensity to be married). Having adjusted for age and marital selection factors, being single was significantly associated with higher mortality, but being divorced or widowed was not. Another study that tries to compensate for reverse causation is Goldman et al (1995). They analyse marital status, health and mortality amongst older people, controlling for baseline health (i.e. before a change in marital status), socioeconomic status and social networks. The main finding is that marriage affects mortality only for men, and that the effect is modest. Widowed men were more likely to be disabled, whereas single women are actually healthier than married counterparts.

Finally, Lillard and Panis (1996) use a simultaneous equations approach to estimate the relationship between health, marital status and mortality, with instrumental variables to account for the reverse causality problem. One of their hypotheses is that the selection effect has a 'demand side' (i.e. healthy people are more attractive) and a supply side (i.e. unhealthy people have more to gain from marriage), and they find indications of both: the explained part of health status tends to be negatively correlated with marriage, whereas the unexplained part is positively correlated. Hence, if the good health status is attributable to personal characteristics, it tends to reduce the propensity to get married, whereas the propensity goes up for a person whose good health is not attributable to personal characteristics. For example, this result would imply that adverse health effects from unemployment (an observable characteristic) are connected with a reduced chance of being married, whereas the opposite holds for individual variations in health that cannot be explained by such personal characteristics. The paper by Lillard and Panis represents the most rigorous attempt to take the reverse causality issue into account; however, the models estimated do not allow for random changes over time in the dependent variable, or autocorrelation (i.e. that these random changes are persistent once they occur).

A good overview of the empirical research to date is provided by Wilson & Oswald (2005). After reviewing a great number of articles on the relationship between cohabitation and health - psychological, physical and mortality - they identify the following general conclusions:

- Marriage reduces the risk of psychological illness
- Marriage tends to increase life expectancy
- Marriage makes people healthier & happier
- Men tend to gain more from the advantageous effects of marriage.
- There is not only a guardian effect (i.e. changes in risk behaviours) - marriage seems to have other positive effects on health as well.
- The quality of the relationship is important

In conclusion, there is still a paucity of research into the issue of cohabitation and health, and this goes for the theoretical as well as the empirical side. It is the objective of this paper to shed some light on the empirical relationship, using econometric techniques that have previously not been applied to the issue.

Our main findings can be summarised as follows. We find that a substantial part of the variation in the data is due to unobserved heterogeneity in the population; i.e. cohabiting and non-cohabiting individuals are different, in terms of their health, from the outset. The correlation is strongly positive, so that people in good health are more likely to be cohabiting. An implication of this finding is that studies that disregard the hidden heterogeneity will overestimate the beneficial effects of cohabitation on health. A second finding is that bereavement of a partner is very detrimental to health, and this effect can, to a great extent, offset the beneficial effect enjoyed from previously cohabiting. Educational attainment, on the other hand, seems to be of little explanatory value once individual differences have been accounted for. In other words, educated people are not healthier because they are educated, but they are healthier because they are different from the outset.

The chapter is organised as follows. In the next section, our methodological approach is outlined and the dataset presented. After that, in Section 3, we present our results. The last section concludes.

2 Methodological Approach

Our econometric model includes two estimating equations; one for cohabitation and one for health. We follow closely the estimation technique taken by Börsch-Supan et al (1993) and adapt it to our problem. Firstly, we allow for unobserved person-specific attributes in both cohabitation and health. In other words, we do not assume that all differences in health trajectories are attributable to observable characteristics (age, gender, education) but we exploit the longitudinal character of the dataset to allow for systematic differences between individuals which emerge from the analysis.

Disability status varies over time, but it also has an important, time-invariant component reflecting the fact that some people are "structurally" healthier than others (due to genetic predisposition or preferences towards risk factors, for example). The same goes for cohabitation status, where it can be assumed that some people are likely to be "structurally" more successful than others (for reasons such as, for example, physical appearance or personal charisma). Furthermore, if there is selection into marriage based on health (or variables correlated with health) we would expect the person-specific attributes of the two estimating equations to be correlated as well.

However, not all unobserved differences can be captured by components which do not vary over time. Hence, we also allow for time-varying disturbances, which are potentially correlated across the equations and potentially exhibiting autoregression, meaning that there is persistence in unobservable characteristics.

2.1 Estimating Equations

We now define the two estimating equations and then investigate the error structure more closely. The health variable is discrete and takes on four different values: healthy, moderately disabled, severely disabled, and dead. Hence, we choose to estimate an ordered probit model. This involves a latent, unbounded and continuous health variable H_t^*:

$$(1)\quad H_t^* = \delta_1 E_t + \delta_2 A_t + \delta_3 A^2{}_t + \delta_4 A^3{}_t + \delta_5 A^4{}_t + \delta_6 \hat{C}_t + \delta_7 \hat{C}_{t-1} + \delta_8 \hat{H}_{t-1} + \delta_9 t + \varepsilon_t^H$$

where E_t is education, A_t is the age[3] in year t, δ_9 captures a linear time trend, C_t and C_{t-1} represent the cohabitation status in the current and the previous period, respectively, and (with a slight abuse of notation[4]) H_{t-1} refers to whether the individual was moderately or severely disabled in the previous year. Hence, we allow for state dependence in both dependent variables in that previous disability and cohabitation status also influence current disability. One problem with state dependence is that unless the first year of observations is treated differently, there is a potential bias in the estimate since it does not account for the fact that the system might not be in equilibrium in the first period.[5]

Remedies to this problem have been proposed by Heckman (1983) and Wooldridge (2000). The problem, however, is that both approaches are unsuitable for our purposes, since the Heckman

[3] We need to go as far as the 4th power since other specifications tend to have health improving at extreme ages.

[4] The previous health status is actually captured by two dummies, $Moderate_{t-1}$ and $Severe_{t-1}$ which for simplicity have been summarised as one (i.e. H_{t-1}) in equation (1).

[5] In other words, the initial observation is not drawn from an unconditional distribution, which is implicitly assumed.

approach requires estimating a large number of extra parameters, and the Wooldridge approach involves estimating fixed effects using an average of the independent variables. This is not a very attractive option, since some of the independent variables are dummies; taking on values zero or one only. Hence, we are not able to correct for this problem, but will use a very long sequence of observations in order to mitigate it.

Then, just as in the standard model, the actual state of disability is defined according to a switching function:

$$(2)\quad \hat{H}_t = \begin{cases} 1 & if \quad H_t^* \le \alpha_1 \\ 2 & if \quad \alpha_1 < H_t^* \le \alpha_2 \\ 3 & if \quad \alpha_2 < H_t^* \le \alpha_3 \\ 4 & if \quad \alpha_3 < H_t^* \end{cases}$$

where the values of H_t, ranging from 1 to 4, correspond to healthy, moderately disabled, severely disabled and dead, respectively and the alphas are cut-off points defining the limit between the various health states. Needless to say, death is treated as an absorbing state, meaning that no recovery from death is possible.

For cohabitation status, we estimate a binomial probit. Hence, the latent function is

$$(3)\quad C_t^* = c + \beta_1 E_t + \beta_2 A_t + \beta_3 A^2{}_t + \beta_4 A^3{}_t + \beta_5 A^4{}_t + \beta_7 \hat{C}_{t-1} + \beta_8 \hat{H}_{t-1} + \beta_9 t + \varepsilon_t^C$$

where c is a constant, and the rest of the independent variables are the same as in equation (1) above. The switching function is then

$$(4)\quad \hat{C}_t = \begin{cases} 1 & if \quad C_t^* \ge 0 \\ 0 & otherwise \end{cases}$$

Finally, we look at the error structures. In order to ensure that the estimated causal effects of age, education and cohabitation are not confounded with systematic differences between individuals, we allow for a very general structure in the error terms. The various parameters estimated in this part can be summarised as:

- *Fixed individual attributes:* some individuals are more likely than others to be disabled, and there is some variation in the propensity to be cohabiting as well. These effects, or rather their variances, are captured by the parameters ω^{hh} (disability) and ω^{cc} (cohabitation) below. A high value of ω^{hh} implies that a great share of the variation in disability is attributable to unobserved structural differences which are unrelated to the independent variables.

- *Correlation in unobservables:* this effect, represented by the parameter ω^{hc} below, measures the degree to which people who are structurally predisposed to be unhealthy are also more likely to be cohabiting. This is an important parameter since it captures systematic differences between people which would otherwise (wrongly) be identified as a causal effect of cohabitation on health. A high value of ω^{hc} implies that a great deal of the observed correlation between disability and cohabitation is due to people being different at the outset and not due to causation.

- *Correlation in shocks:* This parameter, denoted σ^{hc} below, captures how random shocks to health and cohabitation are correlated. Hence, this parameter is similar to the 'correlation in unobservables' parameter above, with the difference that the correlation refers to temporary effects and not structural characteristics of the individual. If this variable is significant, there are factors not captured by the independent variables that influence both disability and cohabitation. Again, disregarding this effect would lead to an overestimation of the causal effect of cohabitation on health.

- *Autocorrelation:* This effect, represented by parameters ϱ_1 and ϱ_2 below, shows to what extent shocks to health and cohabitation are persistent over time. These do not have an obvious interpretation but are necessary once we allow for persistence (i.e. state dependence) in the dependent variables since otherwise the coefficients would be biased. Bertrand et al (2004) have shown that as soon as there is persistence in the dependent variables, a causal effect will be picked up even when there is none unless autocorrelation is allowed for.

2.2 Estimating Procedure: Maximum Simulated Likelihood

We estimate the model outlined in equations (1) and (3) using maximum likelihood. However, given that the two dependent variables are both limited - taking on discrete values only - estimating a dynamic model with the type of error structures we have outlined above poses some challenges. The main problem is that the likelihood function attains so many dimensions that it becomes intractable.

However, maximum simulated likelihood offers a solution to this problem. The idea of this estimator is to draw several series of error terms which are consistent with the data actually observed. We employ an algorithm proposed by Geweke (1989). In short, it means that we draw a series of numbers from a uniform distribution and then transform them (in a straightforward application of the integral transform theorem) into a truncated normal variable that fits the observed data. In terms of the cut-off point in equation (2), the simulated error terms must be such that each individual ends up in the disability category actually observed in the data. The Geweke algorithm produces unbiased estimates of the parameters, and once it has been implemented, standard maximum likelihood techniques can be used to estimate the model.

2.2.1 Discussion

In general, the simulation estimator produces consistent estimates of the parameters of the econometric model. Furthermore, Börsch-Supan and Hajivassiliou (1993) find that for 20 simulations per observation, the simulation bias is negligible. Hence, the estimator seems to be appropriate for our purposes.

There is, however, one practical problem related to the assumption of persistence in the two dependent variables. This problem has to do with the treatment of initial observations, as a simple estimation along the lines outlined above would be based on the erroneous assumption that the system is in equilibrium in the first period. This will lead to inconsistent estimates. Two different approaches have been suggested to remedy this problem. The first one, proposed by Heckman (1981), is to estimate the initial conditions separately and allow for any type of correlation pattern between the initial conditions and any subsequent condition. An alternative to this is provided by Wooldridge (2005) who proposes modelling the distribution of the heterogeneity conditional on the initial condition and any time varying regressors that may be present. Doing this does not require internal consistency with the underlying statistical model nor does it require computations that are as involved as the Heckman (1981) method, but it does require additional distributional assumptions.

Unfortunately, none of these approaches will be useful for our purposes. Heckman's approach increases the number of parameters to be estimated substantially, and given the size of the dataset this becomes a hopeless task. With Wooldridge's method, we are left with the problem that most variables used for estimation, apart from age and the dependent variables themselves, tend to be time-dependent. Hence, using that approach would prevent us from estimating the parameters of interest.

The initial condition problem decreases with the number of waves in the panel, however. Since we have many waves at our disposal, the problem is likely to be relatively small in our case. Furthermore, since we are focusing on older people in the population, it could be argued that the model should be close to equilibrium once their conditions are being recorded in the BHPS.

2.3 The Dataset

For the estimation, we use the twelve first waves of the British Household Panel Survey. In this subsection, we define the variables used, report the treatment of missing values and provide some summary statistics.

2.3.1 Variables

The variables used for estimation are presented in *Table 1*. The definitions are mostly obvious, but the health variable H requires some further explanation. For this variable, we make use of information as to whether the individual is alive in a certain year or not (for dead individuals, the variable takes on the value $H = 4$). For survivors, we use the questions concerning whether the health status of the respondent limits daily activities. The categories allowed for in these questions are roughly equivalent to Activities of Daily Living.[6] Furthermore, as is common in long-term care insurance underwriting, we assume that 'moderate disability' corresponds to having failed two activities, and 'severe disability' corresponds to having failed three or more activities. Respondents who report that their health does not limit their daily activities are coded as healthy. Furthermore, we have suppressed the education category *E5*, which does not have any of the qualifications mentioned in categories *E1-E4*. Hence, this is the group with the lowest level of educational attainment.

[6] Activities of daily living are activities related to personal care and include bathing or showering, dressing, getting in or out of bed or a chair, using the toilet, and eating.

Table 1. *Definition of Variables.*

Variable	*Definition*
A	Age (Calendar year minus birth year)
E1	University Degree
E2	Teaching/Nursing Qualifications
E3	A Levels
E4	O Levels or equivalent
C	Individual married or cohabiting
H	Health limits daily activities

2.3.2 Treatment of Missing Values

Missing variable values are a particularly large problem in this work, since excluding individuals with missing observations is not an option as it would bias the mortality rates. In general, some 2-3 per cent of observations were missing. Some of these were quite easy to impute from earlier or later observations: for instance; somebody who has a university degree in a certain year will have a university degree in any subsequent year.

In a second step, we assumed that if an individual has the same cohabitation status or health status in the two years either side of a missing observation, we assume that the missing observation had the same value as the two surrounding ones. This seems reasonable given how slowly health and cohabitation status may change, but it might be problematic since it would bias the estimates if there is a substantial probability of two transitions occurring over that time period. Given, however, the low number of missing observations of this kind, the impact on the parameter estimates must be relatively low.

For all observations that were still missing after this exercise, we simply assumed that they belong to the most common categories (i.e. cohabitation, healthy and no education). This is certainly not unproblematic, but it would be equally problematic to make imputations based on variables used in the estimations. Besides, it can again be argued that the small number of missing cases will mean that this practice has a limited impact on the results.

2.3.3 Descriptive Statistics

In what follows, we will provide some simple cross-tabulations of the raw data which we use in the estimates. We include all 6,690 permanent members of the panel. We start out by showing the main variable of interest - disability - and how it evolves with age in *Table 2*.[7] We have defined disability quite widely here as failing one or more ADLs.

[7] In principle, the data could also be partitioned by gender. Due to the few observations in some age brackets, however, it is better to present the pooled dataset. Males and females exhibit the same patterns of gradually deteriorating health, but female survivors tend to have slightly worse health than males.

Table 2. *Health Status (ADLs) by age, 1991. (Number of individuals, percentage in italics).*

Age	*Healthy*	*Disabled*	*Total*
<60	841	188	1,029
	81.73	*18.27*	*100.00*
60-69	788	234	1,022
	77.10	*22.90*	*100.00*
70-79	529	247	776
	68.17	*31.83*	*100.00*
80-89	171	132	303
	56.44	*43.56*	*100.00*
90+	13	16	29
	44.83	*55.17*	*100.00*
Total	**2,342**	**817**	**3,159**
	74.14	*25.86*	*100.00*

Table 2 shows the well documented relationship between health and age. For instance, among people in their fifties, fewer than 20 per cent have any physical impairment, whereas at the highest ages, the majority of people have at least one impairment.

Next, we look at the role of cohabitation. In *Table 3*, we cross-tabulate the initial wave by health status and cohabitation status; again, the disabled status is assumed to be when one or more ADLs are failed. The two seem to be correlated; a person not cohabiting is fifty per cent more likely to be disabled as a person who is cohabiting.

Table 3. *Health Status (ADLs) by cohabitation status, 1991.*

Cohabit	*Healthy*	*Disabled*	*Total*
No	762	372	1,134
	67.20	*32.80*	*100.00*
Yes	1,580	445	2,025
	78.02	*21.98*	*100.00*
Total	2,342	817	3,159
	74.14	*25.86*	*100.00*

Furthermore, the cohabitation status in the initial year seems to be quite a good predictor of the health status (including death) in subsequent years. In *Table 4* we cross-tabulate the cohabitation status in 1991 with the health status in 1996. Clearly, people who were cohabiting in 1991 had a higher chance of being alive and healthy in 1996. The mortality rate, in particular, seems to be high for non-cohabiting people when compared to cohabiting people.

Table 4. *Health Status 1996 (ADLs) by cohabitation status 1991.*

Cohabit	*Healthy*	*Moderate*	*Severe*	*Dead*	*Total*
No	739	60	194	141	1,134
	65.2	*5.3*	*17.1*	*12.4*	*100.0*
Yes	1,539	79	256	151	2,025
	76.0	*3.9*	*12.6*	*7.5*	*100.0*
Total	2,278	139	450	292	3,159
	72.1	*4.4*	*14.2*	*9.2*	*100.0*

Finally, we look at the relationship between health and education. Figures are presented in *Table 5*. The education variable reflects the self reported educational attainment, where for simplicity we have merged the educated categories into one single group corresponding to an educational level of at least a GCSE (or its predecessors: GCE, Higher School Certificate etc). The health variable is disability in 1996. As expected, a higher educational attainment is correlated with better health. The effect of education seems to be particularly strong for moderate disability, where the prevalence amongst non-educated people is more than twice as high as the corresponding figure for educated people.

Table 5. *Health Status 1996 (ADLs) by educational attainment (GCSE+ equivalent) 1991.*

Education	*Healthy*	*Moderate*	*Severe*	*Dead*	*Total*
No	1,443	113	338	227	2,121
	68.0	*5.3*	*15.9*	*10.7*	*100.0*
Yes	835	26	112	65	1,038
	80.4	*2.5*	*10.8*	*6.3*	*100.0*
Total	2,278	139	450	292	3,159
	72.1	*4.4*	*14.2*	*9.2*	*100.0*

3 Results

3.1 Parameter Estimates

Estimation results for cohabitation are presented in *Table 6*. In the table, the first group of variables are the exogenous variables - constant, age and education. The label C_{t-1} refers to the cohabitation status in the previous year, and '$Moderate_{t-1}$' and '$Severe_{t-1}$' refer to whether the individual was moderately or severely disabled in the previous year. Parameters σ^{bc}, ω^{bc}, ω^{cc} and ϱ_1 refer to the structure of the error terms.

Some aspects of *Table 6* are surprising. Firstly, all the exogenous explanatory variables are insignificant - i.e. education seems to have very little explanatory power for the cohabitation variable (as can be seen by the high p values for the variables *Edu2-4*). Concerning age, the estimates (*Age*, Age^2 etc) are individually insignificant but taken as a group they are significant, which reflects the fact that the probability of cohabitation is declining with age, but at a non-linear rate. We also notice that there is a negative time trend - the probability of cohabitation decreases over time - which is significant at the 10 per cent level.

Even more interestingly, we find evidence of assortative mating into cohabitation, as demonstrated by the negative coefficient estimated for the ω^{bc} parameter. The implication is that people predisposed to be healthy are also more likely to be cohabiting. On the other hand, the results also suggests that there is an adverse selection effect in cohabitation (i.e. people with disabilities have stronger incentives to remain with their partners), since the occurrence of disability increases the probability of staying in the cohabitation state. This result is not statistically significant, however.

Furthermore, we estimate a positive coefficient for the σ^{bc} parameter, implying that the changes in unobserved differences between people actually exhibit a negative correlation between cohabitation and health. This has a similar interpretation to the ω^{bc} parameter since it also measures the degree of non-causal correlation between cohabitation and health. In this case, however, the parameter measures how the correlation evolves due to shocks that occur over time. In other words, this finding suggests that people who experience adverse shocks to their health are more likely to stay in their relationship. This is to be contrasted with the previous finding of a positive correlation between cohabitation and health in *fixed* unobservables (as captured by ω^{bc}).

In general, our results suggest that the relationship between cohabitation and health is not straightforward: people who are generally healthy seem to have higher chances of finding a partner, but once people are cohabiting, it seems that the occurrence of disability increases the chances of staying together. Whereas the causal effect of disability is not significant (although the positive effect of moderate disability is almost significant at the 10 per cent level), the significant estimate of parameter σ^{bc} suggest that adverse health shocks coincide with increased probability of cohabitation. This finding suggests that we have observed the 'supply side' effect discussed by Lillard and Panis (1996).

Table 6. *Estimation Results, Cohabitation.*

Variable	*Coefficient*	*Std Error*	*T Stat*	*P Value*
Constant	-0.0140	0.3074	-0.045	0.964
Age	-0.2271	0.1614	-1.4076	0.160
Age^2	0.3668	0.3049	1.1855	0.236
Age^3	0.0059	0.0069	0.8505	0.395
Age^4	-0.2025	0.1686	-1.2011	0.230
Edu2	0.5868	0.4263	1.3767	0.169
Edu3	0.3262	0.6042	0.5398	0.589
Edu4	-0.3259	0.2301	-1.4162	0.157
C_{t-1}	4.4599	0.3261	13.6765	0.000
$Moderate_{t-1}$	0.6746	0.4157	1.6229	0.105
$Severe_{t-1}$	0.1244	0.3118	0.3989	0.690
Year	-0.0768	0.03898	-1.9285	0.054
σ^{hc}	0.4427	0.0880	5.0330	0.000
ω^{cc}	0.0444	0.0671	0.6625	0.508
ω^{hc}	-0.2371	0.1008	-2.3526	0.019
ϱ_1	-0.3125	0.1078	-2.9003	0.004
Loglik	-811.06			
$Loglik_0$	-1,974.50			
Pseudo R^2	0.5892			
N	7,986			

In *Table 7*, we present parameter estimates for the disability variable. Apart from the constant and the variable C_t (current cohabitation status), the variables are the same as in *Table 6* above.

Again, we find that the education level of the individual has low explanatory power (as can be seen by the high *p* values of the variables *Edu2-4*). The age parameters also tend to be insignificant, but they are jointly significant; hence, health is declining with age but at a non-linear rate.

The parameter estimates provide many interesting insights. Firstly, we notice that cohabitation has a very strong (and strongly significant) effect on health (C_t is significant even at the one per cent level). However, the effect of losing a partner is even stronger, and is comparable to the effect of severe disability in the previous year. In order to see this, notice that the net effect of cohabitation for a person who was cohabiting in both periods is equal to -0.44 (i.e. 1.87 – 2.31), whereas for a person who had a partner and lost him or her in the last year has a net effect of 1.87 (this is, again, the parameter estimate for C_{t-1}; remember that a higher value implies higher degree of disability). Thus, in comparison to a person who does not suffer bereavement, the adverse health effects are substantial and indeed seem to be even greater than the effect of previous severe disability (variable $Severe_{t-1}$ has parameter estimate 0.99). Hence, the cohabitation state clearly has a strong impact on health, even after unobserved heterogeneity has been accounted for.

Furthermore, we find that there is strong and statistically significant state dependence in the disability variable. Being moderately disabled reduces future health prospects substantially (which is captured by a point estimate of 0.82 in the table), and the effect of severe disability is slightly higher (point estimate 0.99). We also find that the time trend is towards worse health, as demonstrated by the positive parameter estimate of the year variable. This is at odds with the general increase in life expectancy that has been observed in the UK as in many other countries. This result has two possible interpretations. One possibility is that the observed increases in life expectancy are due to changes in the independent variables we are studying. This would imply that younger cohorts have higher levels of education and different family structures than older cohorts. Given that the explanatory value of education is limited according to our results, and given the development towards more unstable family structures, this explanation seems unlikely, however.

An alternative interpretation is that the model used here simplifies too much. It might not be appropriate to assume that the cut-off values between different disability states (and death) – i.e. a_1, a_2 and a_3 - are constant over time. Indeed, data for the development of healthy life expectancy (HLE) over the last few decades suggest that a substantial proportion of the gains in life expectancy are spent in moderate disability, whereas the time spent in severe disability has tended to decrease. This might in turn be due to improvements in medical technology. If this is the case, the three cut-off values might well be diverging over time and this is then the effect picked up by the time coefficient estimated here. This explanation would imply that an individual with given characteristics and age would have greater chances of being healthy the later he was born. Thus it remains an issue for future research to analyse whether the boundaries between the healthy state, disability, and death diverge over time.

Finally, we find that unobserved systematic differences between people are very significant in explaining the dynamics of health. Fixed individual characteristics (point estimate 0.67 in *Table 7*) account for a substantial share of the variation in the data. Again, we can conclude that causal effects from education seem to be dominated by variation in these types of unobserved characteristics. This also helps us to understand why the education variables come out insignificant: educated people are not healthier because they are educated, but because they are different from the outset. This is a finding which is very relevant for policy.

Table 7. *Estimation Results, Health.*

Variable	*Coefficient*	*Std Error*	*T Stat*	*P Value*
Age	-0.0450	0.0494	-0.9117	0.362
Age^2	-0.0003	0.0019	-0.1431	0.886
Age^3	0.0005	0.0030	0.1555	0.876
Age^4	0.0480	0.0238	2.0160	0.044
Edu2	-0.1806	0.2677	-0.6746	0.500
Edu3	-0.1456	0.5527	-0.2633	0.792
Edu4	0.1382	0.2697	0.5125	0.608
C_t	-2.3147	0.3298	-7.0172	0.000
C_{t-1}	1.8693	0.2882	6.4873	0.000
$Moderate_{t-1}$	0.8199	.01934	4.2386	0.000
$Severe_{t-1}$	0.9866	0.1925	5.1258	0.000
Year	0.0565	0.0211	2.6804	0.007
σ^{hc}	0.4427	0.0880	5.0330	0.000
ω^{hh}	0.6666	0.2189	3.0449	0.002
ω^{hc}	-0.2371	0.1008	-2.5326	0.019
ϱ_2	-0.1838	0.0949	-1.9374	0.053

3.2 Illustrations

Since we have built a dynamic model for disability and cohabitation, there are several different uses to which our research could be put: for example, public planning purposes, or pricing long-term care insurance policies. In this section, we present some examples of ways in which our results may be used to calculate survival curves and healthy life expectancy for chosen subcategories of the population. First, however, we give a brief introduction to the concept of survival curves.

Figure 1 is a survival curve for males and females based on English Life Table 15 produced by the Office for National Statistics in conjunction with the Government Actuary. A life table does not represent the actual population but what the population would look like if age specific mortality were to apply to a synthetic population, usually 100,000 people, hence the values on the vertical scale. Let us assume on average that disability tends to be both progressive as well as permanent and is concentrated in the period leading up to death. The shaded area of *Figure 1* represents the proportion of the surviving population that is disabled. Diagrams like this are a useful tool for illustrating morbidity and mortality (Mayhew, 2003). It is fairly obvious that, in a stationery population, the horizontal width of the shaded area gives an indication of the expected duration of disability at a given age whereas the vertical height gives an estimate of the number of disabled of a given age.

The average 'stock' of disabled of a given age is given by measuring A-C and the duration by A-B. In fact it is striking how very nearly the duration tends to be constant in older age but is longer if disability begins at a younger age, say between 40 and 50 years. The overall average is 9.91 years. If we were to construct the same diagram but only represent on it the most severely disabled group our shaded strip would be much narrower. This group is the most severely disabled and contains those likely to be in need of intensive nursing or palliative care. It turns out that for this group the duration of severe disability averages 1.48 years.

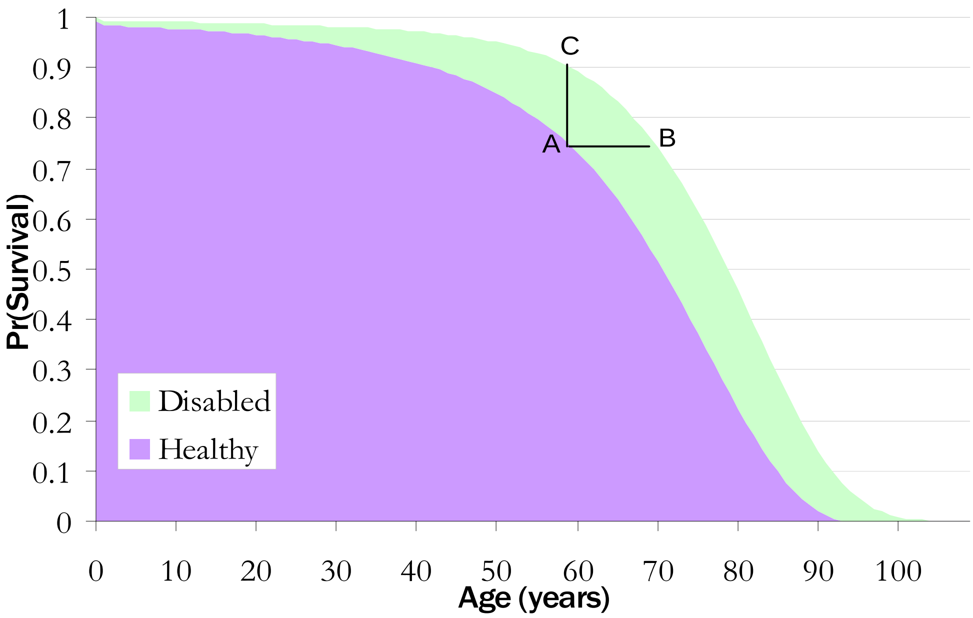

Figure 1. *Survival curve based on English Life Table 15.*

Using our estimation procedure, we are able to break down the population figures in *Figure 1* for every type of starting condition - defined in terms of age, education, cohabitation status and initial disability status. We now provide some examples. In *Figure 2*, we show a typical survival curve for a male aged 50 in 1991, who is healthy at the outset.

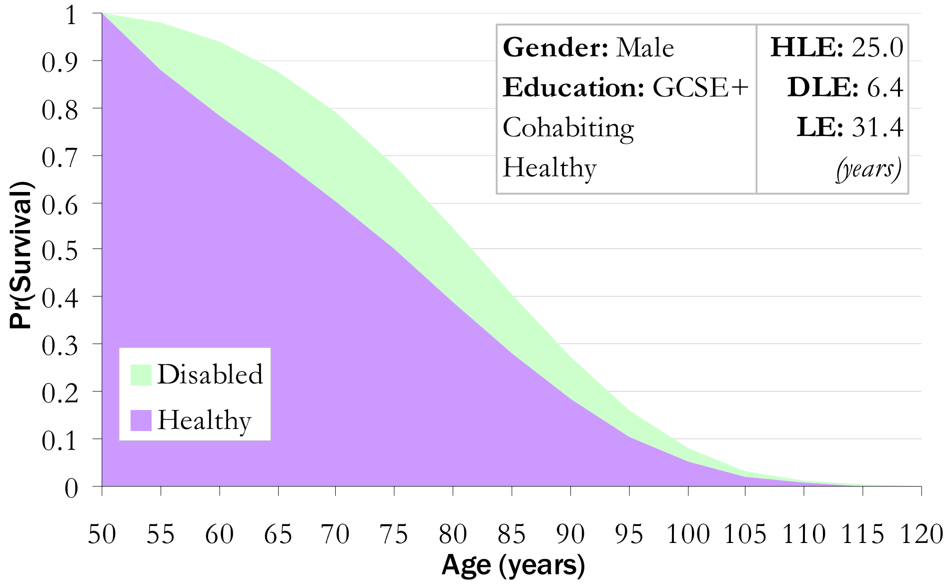

Figure 2. *Survival Curves for Cohabiting, well educated males.*

Hence, a male aged 50 in 1991 who was healthy, cohabiting and had at least GCSE level of education, would have a total life expectancy (LE) of 31.4 years, comprising 25 years spent healthy (healthy life expectancy, HLE) and 6.4 years spent in disability (disabled life expectancy, DLE).

The survival curve in *Figure 2* can be compared with the results in *Figure 3*. The change in educational status and cohabitation status at the outset is followed by a one year reduction in life expectancy (i.e. 30.3 years compared to 31.4). More importantly, however, the less advantaged individual is expected to spend three years more in disability (i.e. 9.5 years compared with 6.4).

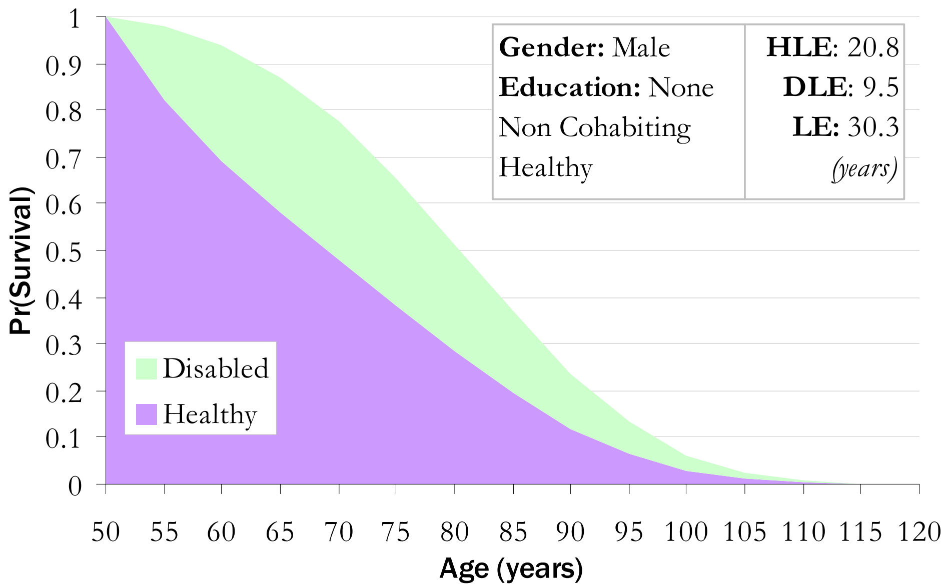

Figure 3. *Survival Curves for Non-cohabiting, uneducated males.*

4 Conclusion

The main future challenge to the UK long-term care system is not dealing with general trends in health, but ensuring that there will be enough informal carers to take care of those in need. Karlsson et al (2006) suggest that there might be a shortfall of up to 40 per cent of the informal care needed if the trends in the next 3-4 decades are unfavourable. Since the vast majority of disabled older people only receive informal care, this would put a considerable strain on public finances if the care instead became formal.

Hence, it is of utmost importance for planning purposes to study recent trends in disability and cohabitation jointly. One objective of this paper has been to develop an empirical model that can be used for these purposes. Previously, there has been a lack of empirical studies which account for the full complexity of the problem. By using simulation techniques developed for limited dependent variable models (i.e. where the dependent variable is categorical), we are able to derive a model which takes all relevant aspects of the problem into account.

We applied the model to a subsample of the British Household Panel Survey, including all men aged 65 and older at the beginning of the panel. In general, we find that the causal relationship from education on health is relatively weak, and that much variation in the data is due to correlation in unobservable characteristics instead. Hence, there is a strong correlation in fixed individual attributes relevant for health and cohabitation, and there is a strong persistence in the dependent variables. In other words, people with higher education are healthier not because they are educated, but because they are healthier at the outset. This finding obviously has strong policy implications.

We have also found evidence of assortative mating into cohabitation, i.e., people with better health prospects seem to be more likely to be cohabiting. However, the relationship is not as straightforward as we would imagine, since the effect of disability on the probability of cohabitation is positive: disabled people are more likely to remain with their cohabitation partner. In effect, this means that people with good health prospects are considered to be more attractive at the outset, but our results indicate that adverse changes in health actually increases the probability of remaining in a relationship. Whether this effect is due to social norms concerning the duties of a partner or behavioural changes of the individual becoming disabled (branded the 'supply side effect' by Lillard and Panis, 1996), remains an issue for future research. Likewise, it would be of great interest to analyse to what extent this effect remains even in times when the volatility of relationships increases and people's commitment to their partners has a tendency to diminish.

Another important finding is that cohabitation has a strong direct effect on health which remains even after unobserved differences between people have been allowed for. This might be one reason disabled people are more likely to remain with their partner. On the other hand, losing one's partner seems to have an even stronger - negative - effect on health. Quite remarkably, the effect on current health status of having lost a partner during the last year is comparable in magnitude to the effect of having become severely disabled during the last year.

Finally, we have demonstrated one of many possible applications of our results, which is to produce survival curves (and calculate healthy life expectancies) for people in different circumstances. It transpires that for males, the gap between cohabiting, educated individuals and non-cohabiting individuals without education, correspond to a one-year difference in healthy life expectancy and a three-year difference in disabled life expectancy.

There are several issues open for future research. First and foremost, the analysis should be extended to include females as well. It is likely that females exhibit different dynamics in their disability and cohabitation paths than males, and it is a plausible hypothesis that cohabitation and bereavement have different effects on women than on men. Another important issue for public policy is how the effects of bereavement can be mitigated so that it does not have such a strong detrimental effect on the individual's health. Knowing this would enable us to design policies aimed at this group in particular.

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