# Partnership Status, Health and Mortality: Selection or Protection?

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Acknowledgements: This research was supported by the ESRC Centre for Population Change (CPC), grant number ES/K007394/1; and the ESRC Centre for Population Change Connecting Generations research programme, grant number ES/W002116/1. We are grateful for the opportunity to use data from the British Household Panel Survey and the UK Household Longitudinal Study, hosted and managed by the UK Data Service.

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**Abstract**

Married individuals have better health and lower mortality than non-married people. Studies show that once we distinguish cohabitants from other non-married groups, health differences between partnered and non-partnered individuals become even more pronounced. Some studies argue that partnered individuals have better health and lower mortality because of the protective effects that a partnership offers (protection); others state that partnered people have better health and lower mortality because healthy persons are more likely to form a union and less likely to dissolve it (selection). This study contributes to this debate by investigating health and mortality by partnership status in England and Wales and analysing the causes of mortality differences. We use combined data from the British Household Panel Survey and the UK Household Longitudinal Study and apply a simultaneous equations hazard model to control for observed and unobserved selection into partnerships. We develop a novel approach to identify frailty based on information on self-rated health. Our analysis shows significant mortality differentials by partnership status; partnered individuals have lower mortality than non-partnered people. We observe some selection into and out of union on unobserved health characteristics; however, the mortality differences by partnership status persist. The study offers strong support for the marital protection hypothesis and extends it to non-marital partnerships.

Keywords: mortality, health, marital status, partnership status, survival analysis, simultaneous equations hazard model, selection, UK

## Introduction

Research on marital status, health and mortality shows that married persons have better health and lower mortality than single, divorced, and widowed individuals. Differences in mortality by marital status were first observed in Britain as early as the mid-19th century (Farr 1858); subsequent research in industrialised countries has shown that these results hold over time and across countries with the differences being larger for men than for women (e.g., Dupre et al. 2009; Johnson et al. 2000; Guner et al. 2014; Hu and Goldman 1990; Murray 2000; Ben-Shlomo et al. 1993; Blomgren et al. 2012; Murphy et al. 2007; Brockmann and Klein 2004; Kravdal et al. 2018; Requena and Reher 2021). Mortality and health variation by marital status persists even when controlling for the socio-demographic and economic characteristics of individuals (Ebrahim et al. 1995; Cheung 2000; Murphy et al. 2007; Drefahl 2012; Staehelin et al. 2012).

The reasons for health and mortality differences by partnership status are far from clear. The *protection hypothesis* states that married individuals have low mortality because the presence of a spouse results in greater emotional and social support, constrains risk-taking behaviour and leads to higher income (Goldman 2001). In contrast, the *selection hypothesis* argues that married people have better health and lower mortality because healthy persons are more likely to marry than unhealthy individuals. Further, they are less likely to become widowed or divorced and more likely to remarry if a previous marriage dissolved. While many studies discuss these two competing hypotheses to explain observed patterns, only a few studies have empirically addressed the role of these two competing explanations in mortality differences by marital status (e.g., Lillard and Panis 1996; Waldron et al. 1996). These studies focused on mortality and health of married and divorced individuals in the United States. Lillard and Panis (1996) found that unhealthy men tend to (re)marry early and remain married longer than healthy men. They also found evidence for selection into marriage based on unobserved factors. Waldron et al. (1996) studied health differences between married and unmarried women and found evidence for both protection and selection effects but only among unemployed women. No effects were found among women in full-time employment.

This study investigates health and mortality by partnership status in England and Wales. We develop previous research in two ways. First, we analyse mortality and health by partnership status rather than by marital status. Although pre- and post-marital cohabitation have become common in recent decades, many previous studies treat cohabitants as either single or divorced individuals. However, studies show that cohabitants' health is more similar to that of the married than to that of single and divorced individuals (Drefahl 2012). Therefore, once we distinguish cohabitants from other non-married groups, health and mortality differences between partnered and non-partnered individuals may become even more pronounced (Franke and Kulu 2018). Second, we examine the role of partnership selection and protection in explaining health and mortality differences by partnership status. We model selection into and out of partnerships and adjust our models for observed and unobserved selection effects using simultaneous-equations survival models. We propose an approach to identifying unobserved frailty – we exploit repeated information on individuals' self-reported health. We also conduct sensitivity analysis using observed data on long-term health and assumptions on the distribution of unobserved frailty. To the best of our knowledge, this is the first study to investigate the role of unobserved selection in mortality differences by partnership status when unmarried partnered individuals are treated as a separate category rather than together with single or divorced individuals.

**Mortality and health by partnership status**

The *protection hypothesis* states that partnership and marriage provide a number of advantages which help protect individuals against various unhealthy activities. Having a partner discourages risk-taking behaviour and encourages healthy lifestyles. Partnered individuals have more regular habits and more healthy diets than non-partnered people (Mata et al. 2015; Oyebode et al. 2014). Partnership also results in greater emotional and social support and acts as a buffering mechanism in the presence of stress. Further, having a partner facilitates access to medical information and health services. This is especially important for men; women are considered to be 'closer' to medical services than men because of motherhood and children (Mesle 2006). Research has also shown that partnership leads to higher (and more stable) income and partners benefit from the economies of the scale (Goldman 2001; Drefahl 2012). For example, couples can prepare a meal for two rather than separate meals for each person, or heat one home rather than two homes (Franke and Kulu 2018). Previous studies show that men benefit more from the reduction in risk-taking behaviour, whereas women benefit from the financial support a partnership can offer (Lillard and Waite 1995). Additionally, married men tend to earn more than unmarried men (Wilcox et al. 2005), although it is not clear whether this premium is due to the greater support from home or the greater commitment of partnered men to paid work. Still, it is likely that the emotional support a partnership provides plays an important role as partners usually help each other cope with work-related stress and encourage career moves. Most of these protective effects vanish after separation (Rendall et al. 2011).

The *selection hypothesis* states that partnered and married people have better health and lower mortality because healthy persons are more likely to form a union. Individuals who are mentally and physically healthy have a higher likelihood of finding a partner and forming a union and are less likely to experience the dissolution of a union than those with health issues. Selection into marriage and partnership based on health also occurs indirectly. Studies show that mate selection is often based on factors related to the health status of an individual such as income or health-related habits. For example, heavy drinkers, smokers, or drug users are less likely to marry. Obesity and emotional instability are also considered factors limiting an individual's opportunities to find a partner (Goldman and Hu 1993; Goldman 2001), although attitudes may have changed over time as the average weight of populations has increased and poor mental health is less stigmatised now than it used to be in the past. Further, studies from several European countries have shown that married men have a higher BMI on average than non-married men, which may be related to a short-term weight gain among married individuals (Lipowicz et al. 2002; Mata et al. 2015).

Partnership patterns have changed significantly in industrialised countries in recent decades. Marriage rates have declined, cohabitation has become common, separation, divorce and repartnering have increased. While the median age of marriage for women born in the 1940s and 1950s was in the early twenties in the UK, the same figure for cohorts born in the 1970s was in the early thirties. The share of married individuals declined from nine-tenths among women born in the 1940s and 1950s to four-fifths among women born in the 1960s (Hannemann and Kulu, 2015). Only one-tenth of the individuals who were born in the 1940s ever cohabited by age 45, whereas more than half of the women who were born in the 1960s cohabited by the age of 30. For the younger cohorts, the share of individuals having ever cohabited by age 30 is about 70% (Ermisch and Francesconi, 2000; Murphy, 2000; Hannemann and Kulu, 2015). One-fifth of the marriages that were formed in the 1965–74 period ended in divorce before their 15th anniversary, whereas more than one-third of marriages have experienced separation in most marriage cohorts from 1995 onwards. Separation levels for non-marital unions have been even higher (Hannemann and Kulu, 2015). Repartnering has also increased, although re-marriage rates increased between the 1940s and 1950s cohorts but have declined thereafter. Instead, post-marital cohabitation has become common, increasingly also in older ages. Briefly, the UK provides an ideal context to study mortality of partnered and non-partnered individuals. Cohabitation spread in the UK slightly later than in the other countries of Northern Europe, but earlier than in Southern Europe.

Recent research has investigated mortality differences between partnered and non-partnered individuals by distinguishing cohabitants from married and divorced individuals. In a study on the US, Liu and Rezcek (2012) observed significant differences in mortality by partnership status: single and separated men had higher mortality than cohabiting men; interestingly, there were no such differences among women. However, subsequent studies have shown relatively similar patterns for both sexes. Using register data from Denmark on men and women ages 18 to 65, Drefahl (2012) showed that mortality was lowest among married individuals followed by cohabiting people. Cohabiting women, especially the low educated, had higher mortality than married women, but their mortality levels were significantly lower than those of single or non-partnered individuals. A study by Franke and Kulu (2018) on England and Wales reported similar results. The analysis showed that married individuals had lower mortality than unmarried persons. Men and women in premarital unions exhibited mortality levels similar to those of married men and women, whereas age-standardised mortality was slightly elevated for post-marital cohabitants, although mortality levels were still lower than those among single and separated individuals. The study also demonstrated that controlling for household size and the presence of children reduced mortality differences between partnered and non-partnered individuals, but significant differences persisted.

A comparative European study by Perelli-Harris et al. (2018) supported that partnered individuals had better self-reported health in mid-life than non-partnered individuals. However, once they adjusted the models for the presence of children and socio-economic status, differences in self-reported health were reduced and in some countries were even eliminated. Kravdal et al. (2023) investigated the effect of cohabitation and marriage on mental health using register data on GP consultations because of mental health conditions. The analysis of longitudinal data showed that individuals' mental health improves over several years before cohabitation. For those who marry, there is a small reduction in the number of GP consultations before the marriage, but no change thereafter. The results suggest that the mental health benefits of cohabitation and marriage are similar. Interestingly, the patterns are similar for women and men suggesting that in egalitarian societies both men and women equally enjoy the benefits of a partnership. These findings are in contrast to some earlier cross-sectional studies, which showed that while there are no significant differences in depressive symptoms between married and cohabiting women, married men enjoy significantly lower depression scores than cohabiting men suggesting that cohabitation does not provide men with the same level of mental health benefits as marriage (Brown et al. 2005).

In summary, past research reports significant mortality and health differences by marital status. Recent studies show that once we distinguish cohabitants from other non-married groups, mortality and health differences between partnered and non-partnered individuals become even more pronounced. With the spread of pre- and post-marital cohabitation and divorce, the distinction between non-married individuals who cohabit and those who do not has become critical to improve our understanding of the causes of good health and low mortality among partnered and married individuals. However, none of these studies have controlled for selection into and out of partnerships when examining mortality differences by partnership status. This study contributes to the existing literature on mortality differences by partnership status in two ways. First, we distinguish between partnered and non-partnered individuals. Second, we use simultaneous-equations survival models to examine the role of partnership selection and protection in explaining mortality differences by partnership status.

## Methodology and modelling strategy

We use survival analysis to study mortality differences by partnership status. The basic model is formalised as follows:

$$ln\mu_i(t) = ln\,\mu_0\,(t) + \textstyle\sum_k \beta_k x_{ik}(t), \qquad (1)$$

where $\mu_i(t)$ denotes the risk of dying for individual $i$ at age $t$. $\ln\mu_0(t)$ denotes the baseline log-risk, with an individual's age as the duration variable. We specify the baseline log-risk via linear splines with the nodes at the following ages: 35, 40, …, 80 and 85. The use of a parametric Gompertz model leads to very similar results. The model includes a set of time-constant and time-varying covariates denoted by $x_{ik}(t)$, with parameters $\beta_k$ measuring their effect. Partnership status is one of these covariates.

There may be unobserved factors (e.g., 'frailty', long-term health conditions, or health habits) that influence both an individual's risk of dying and their likelihood of being in a specific partnership status (e.g., non-partnered or partnered). If this is the case, then the estimated effect of partnership status on mortality (in equation 1) would be biased because the variable is endogeneous to mortality, i.e., it correlates with the error term of the mortality equation (not specified). For example, if healthy individuals were more likely to form partnerships and unhealthy individuals were more likely to remain non-partnered, then we would over-estimate the (protective) effect of partnership on an individual's health. To detect and control for unobserved selection effects, we could consider the following simultaneous-equations model:

$$
\begin{aligned}
&ln\,\mu_i^M(t) = ln\,\mu_0^M(t) + \sum_k \beta_k^M x_{ik}(t) + u_i^M \\
&ln\,\mu_{ij}^U(t) = ln\,\mu_0^U(t) + \sum_k \beta_k^U x_{ijk}(t) + u_i^U \quad (2) \\
&ln\,\mu_{ij}^D(t) = ln\,\mu_0^D(t) + \sum_k \beta_k^D x_{ijk}(t) + u_i^D
\end{aligned}
$$

where $\mu_i^M(t)$ denotes the risk of dying for individual $i$, $\mu_{ij}^U(t)$ and $\mu_{ij}^D(t)$ denote the hazard of $j$th union formation or union dissolution for individual $i$, respectively. $u_i^M$, $u_i^U$, $u_i^D$ are individual-level time-invariant residuals (or random effects) for the mortality, union formation, and union dissolution equations, respectively. However, while the individual-level error terms in the second and third equations are identifiable if repeated episodes are available for some individuals, the identification of the individual-level error term of the first equation (i.e., the mortality equation) is not possible (without strong assumptions) because 'we only live once'.

To solve the issue, we include an individual's health history as a separate process in the model. We use information on an individual's health status measured several times during their lives. The simultaneous-equations model is then specified as follows:

$$
\begin{aligned}
&y*_{it}^H = \alpha_t^M + \sum_k \beta_k^H x_{ikt} + u_i^H + \varepsilon_{it} \\
&ln\,\mu_i^M(t) = ln\,\mu_0^M(t) + \sum_k \beta_k^M x_{ik}(t) + \lambda u_i^H \quad (3)
\end{aligned}
$$

$$ln\, \mu_{ij}^{U}(t) = ln\, \mu_{0}^{U}(t) + \sum_{k} \beta_{k}^{U} x_{ijk}(t) + u_{i}^{U}$$

$$ln\, \mu_{ij}^{D}(t) = ln\, \mu_{0}^{D}(t) + \sum_{k} \beta_{k}^{D} x_{ijk}(t) + u_{i}^{D}$$

where $y*_{it}^{H}$ is a variable measuring the health status of individual *i* at age *t*, specified in the form of a linear model; $u_i^H$ is an individual-specific time-invariant residual (or random effect) for both the health and mortality equations. The identification of the random effect is based on repeated measures of individual health. We assume that $u_i^H$ captures an individual's long-term health conditions that are unobserved; $\lambda$ is a loading factor allowing different effects of the random effect on mortality and health[1].

The residuals of the model are assumed to follow a multivariate normal distribution. We estimate the variances of the person-specific residuals and the covariances between the residuals. A non-zero covariance between $u_i^H$ and $u_i^U$ would suggest that individuals whose unobserved characteristics place them at an above-average level of health are more (or less) likely to form a partnership. By making this correlation a part of our model, we control for unobserved selection effects and calculate unbiased estimates of the effect of partnership status on mortality and health. We use the statistical software aML to estimate simultaneous-equations survival models (Lillard and Panis 2003).

Our model shares several features with that proposed by Lillard and Panis (1996). Their strategy was to allow heterogeneity components from marriage formation, marriage dissolution, and health to affect mortality via the correlation of the three residual components. In contrast, we propose to first extend the heterogeneity component from health to mortality (i.e., borrow information from the health equation to identify frailty) and then include the correlations between the residual terms in the models. We also study all partnerships rather than marriages only. With increased complexity of individuals' partnership histories, the proposed strategy allows us to include in a simultaneous equations model further equations with heterogeneity terms (e.g., separately for cohabitation, for marriage after cohabitation and for direct marriage) to determine and control for various selection mechanisms and effects when investigating mortality by partnership status (cf. Kulu and Boyle 2010).

The identification of the model is based on repeated partnership episodes available for part of the sample: a) 4,201 individuals were at risk of forming a second union (1,186 second union formation events); b) 2,180 individuals were at risk of experiencing a second separation (739 separation events). The identification of the individual-level error term of the mortality equation is not possible as we only have a single episode for each individual. However, we can "borrow" information on individuals' frailty from data

[1] The residual of the health equation ($u_i^H$) measures unobserved health determinants, which may include frailty, but also other factors such as lifestyle preferences.

on their self-reported health. Hence, we also include in the simultaneous-equations model a health equation, which investigates self-reported health. Because we have annually repeated health measures (on average, there are 12.3 measurements per individual), we can identify the individual-level error term of the health equation. This error term is assumed to measure unobserved long-term health conditions or (even) health determinants; we extend this term (or random effect) also to the mortality equation (see equation 3).We conduct further analysis to examine the sensitivity of the results to an alternative specification of frailty. We measure individuals' underlying long-term health conditions (or frailty) directly and include this measure both in our basic mortality model (as specified in equation 1) and the simultaneous-equations model (in equation 3) to test the robustness of the results. To do so, we use self-reported health and calculate a cumulative health measure, which is an average value for health, updated annually. However, some caution is needed when interpreting these results. Although the cumulative health variable measures individuals' longer-term health conditions, its values are still shaped by partnership changes and experiences (e.g., marriage may lead to better self-reported health and separation to poorer health). Hence, their inclusion in the model will likely over-estimate selection effects. We thus consider the results of joint modelling without cumulative health more conservative and thus superior to this approach.

**Data and variables**

We use combined data from the British Household Panel Survey (BHPS) and the UK Household Longitudinal Study (UKHLS). The BHPS is a nationally representative sample of 5,000 households and approximately 10,000 individuals (Institute for Social and Economic Research 2010, 2014; Taylor et al. 2010). Between 1991 and 2008, the same sample of adults were interviewed each year. If the composition of a household changes, the survey follows original household members and interviews new household members. Children were interviewed once they reached the age of 16. In our analyses, we use information on original sample members and two additional subsamples (the European Community Household Panel and the Wales Extension Sample). We exclude Scotland and Northern Ireland because the sample design and some control variables (e.g., area of residence) differ from the England and Wales sample.

In 2009, the UK Household Longitudinal Study (UKHLS) was launched. The structure and design of the UKHLS is very similar to that of the BHPS; the same sample of adults is interviewed each year (Institute for Social and Economic Research and NatCen Social Research 2015). From wave 2 onwards, UKHLS includes information on BHPS respondents who completed an individual interview in the last BHPS wave and agreed to participate in the UKHLS (Lynn et al. 2012). We have linked the BHPS and UKHLS data and thus follow our sample members until 2015. Using this combined dataset allows us to follow individuals for much longer (more than 20 years) than if we only analysed UKHLS respondents; this is important for the identification of our selection models which require information on repeated partnership events and health measurements. Individuals are observed from age 16 or from the date of entry into the study (if later) until death, or the end of observation window (2015), whichever happens first. Individuals are also censored when they drop out from the survey. We have a sample of 7,059 men and 7,788 women. We have full monthly partnership histories of the respondents and annual information on their self-reported health. Individuals who die between two consecutive waves are identified using information on the 'activity status' of the respondents. It is assumed that the respondent had died 6 months after the end of the last interview. This rich dataset allows for a detailed analysis of the effect of partnership status on mortality and health controlling for observed and unobserved selection effects.

The outcome variable is age at death measured in months. Our main explanatory variable of interest is partnership status, which is a time-varying variable. We distinguish between never partnered single, cohabiting, married, separated, and widowed individuals. It is important to clarify that all partnered individuals are either cohabiting or married, whereas non-partnered can be single, separated, or widowed. For example, if a person separates and then forms another union, they will be classified as cohabiting (if the union is non-marital) following separation. We thus explicitly distinguish between partnered and non-partnered individuals. Our main control variables are an individual's age, sex, and period (1991–1994, 1995–1999, 2000–2004, 2005–2009, 2010–2015). We also control for educational level (low, medium, high), the number of children (no child, one child, two children, three or more children), housing tenure type (homeowner, social rent, private rent), and the area type of residence. We used the size of the local authority district and its population density to classify areas as: (1) the capital city of London; (2) other large cities with a population of more than 400,000 (large city); (3) cities with 200,000–400,000 inhabitants (medium city); (4) local authority areas with less than 200,000 inhabitants, but with a population density of 1,000 or more individuals per km2 (town); (5) local authority areas with less than 200,000 inhabitants and with a population density of 250–1,000 individuals per km$^2$ (small town); and (6) areas with less than 200,000 inhabitants and with less than 250 individuals per km$^2$ (rural area) (for further details, please see Kulu and Washbrook 2014). Table 1 provides the distribution of the number of deaths and risk time by the variable categories.

We also model union formation, union dissolution, and health in our simultaneous analysis. In the union formation equation, additional variables are partnership status (never partnered single, separated, widowed), time since separation (no separation, 0-1, 1-3, 3-5, 5+ years), annual health status (excellent, good, fair, poor, very poor, missing), and employment status (self-employed, employee, unemployed, other, missing). For the separation equation, we also included information on union duration (0–1, 1–3, 3–5, 5+ years), partnership status (cohabitation, marriage), age at union formation (16–19, 20–24, 25–29, 30–34, 35+ years), order of separation (first, second or higher order); housing tenure type (homeowner, social rent, private rent, missing), employment status, and annual health status. Models on health use a continuous variable (1 – excellent, … 5 – very poor, 6 – missing); our sensitivity analysis with a binary variable (1 – excellent and very good, 0 – fair, poor, very poor, missing) led to very similar results (also with and without the missing category). The models include the following covariates: age, sex, period, partnership status, place of residence, tenure, employment status, and educational level. Tables A1 to A3 in the Appendix provide information on the distribution of risk time and events by categories of variables.

## Results: Mortality by partnership status

*Main results*

Table 2 shows relative mortality rates in England and Wales by partnership status (Table 3 shows the corresponding log mortality rates and the results of the full models). The rates have been calculated relative to those of married individuals. We estimate three models stepwise. In Model 1, we only control for age, sex, and period. We see that non-partnered individuals have higher mortality than partnered people. Single (i.e., never-partnered) individuals exhibit 47% (90% CI: 1.26, 1.72), separated (or divorced) 58% (90% CI: 1.34, 1.86), and widowed individuals 23% (90% CI: 1.09, 1.38) higher mortality than married people. There are no significant differences between married and cohabiting individuals. Although the standard errors of the estimates are large and confidence intervals are wide, the differences between the groups are significant. Most importantly, the results are very similar to those based on the analysis of the ONS Longitudinal Study, a one-percent sample of the population of England and Wales (Franke and Kulu 2017). This is reassuring and provides further evidence of the high quality of the combined BHPS and UKHLS data. In Model 2, we adjust for further demographic and socio-economic characteristics of individuals: the number of children they have, their educational level, housing tenure type, and place of residence. We see that mortality differences between the groups are reduced, but they persist – non-partnered individuals still have significantly higher mortality rates than partnered people. Further analysis showed that housing tenure accounts for some initial mortality differences observed between partnered and non-partnered individuals: married and cohabiting individuals are more likely to be homeowners than single and separated individuals among whom many are social renters.

Next, we account for possible selection effects while investigating mortality by partnership status. There may be unobserved factors (e.g., frailty or long-term underlying health conditions) that influence both individuals' likelihood of being in or out of a union and their risk of dying. To detect and control for unobserved selection effects, we model mortality and partnership change jointly allowing for the correlation between individual-level residuals across the equations. The analysis shows that all correlations between the random effects are significantly different from zero but vary in their magnitude. Individuals who are frail (i.e., with long-term underlying health issues) are more likely to form and particularly to dissolve unions (Table 4, Model 3). The correlations between the random effects are 0.15 (health and union formation) and 0.51 (health and separation), accordingly. The latter result corresponds to expectations, the former is somewhat surprising – if the selection theory was true, one would expect a negative correlation between health and union formation, although a previous study has shown similar results (for men) (Lillard and Panis 1996). Our further analysis showed that the cross-equation correlation is significantly different from zero for cohabitation, but not for marriage suggesting that there might be some selection of individuals with poor health into post-marital cohabitation. However, the relationship between the residuals is not very strong. What are the implications of the unobserved selection effects for our main results of interest, especially the finding that individuals with underlying health problems are more likely to separate than healthy people? Interestingly, mortality differences by partnership status do not change very much across the models – the results of Models 2 and 3 are very similar (Table 2), except that the differences between single and married individuals slightly increase. Nevertheless, the results suggest that unobserved selection effects, if present, are not very strong.

*Sensitivity analysis*

Next, we measure individuals' underlying long-term health conditions directly and include this measure in our mortality model to test the robustness of the results. We first include a cumulative health measure in a basic mortality model (Table 2, Model 2a) and our simultaneous equations model (Table 2, Model 3a). In both cases, the differences between the groups marginally decrease but they are very close to those observed in Models 2 and 3 (Table 2). However, most importantly, the group differences persist: single, separated, and widowed individuals have significantly higher mortality than partnered individuals. Some caution is needed when interpreting these results. Although the cumulative health variable directly measures individuals' longer-term health conditions, its values are still shaped by partnership changes and experiences. Hence, their inclusion in the model will likely over-estimate selection effects. Interestingly, when comparing models with (Model 3a) and without (Model 3) cumulative health the loading factor for mortality declines from 0.80 to 0.51. This suggests that the cumulative health measure explains away some variation in frailty in the mortality equation, but some is still left. This is largely expected.

The effects of the control variables are consistent with the findings of previous research (Table 3). Women are less likely to die than men. Mortality has declined over the years, although the analysis shows that mortality rates were higher in the late 1990s than in the early 1990s, which may be related to the study design. Mortality is lower among individuals with children, high education, and among homeowners compared to those without children, with low education, and those living in socially rented accommodation.

We do not observe significant differences between residential contexts once individuals' demographic and socio-economic characteristics have been controlled for.

## Discussion and Conclusions

This study investigated health and mortality by partnership status in England and Wales using combined data from the British Household Panel Study and the UK Household Longitudinal Study. This study proposed a novel way of estimating unobserved frailty to identify and control for potential selection into and out of partnerships on unobserved long-term health conditions. We proposed to "borrow" information on individuals' frailty from data on their health (or repeated health measurement). We also conducted sensitivity analysis using observed data on long-term health as a measure of individual frailty. Our analysis suggests significant health differences by partnership status; partnered individuals have lower mortality than non-partnered people. Our simultaneous analysis of mortality, health, and partnership changes showed some selection into and out of union on unobserved health characteristics; interestingly, however, the mortality differences by partnership status persisted even after controlling for unobserved selection. Our further analysis showed that the results are robust to different model specifications (including annually measured cumulative health) suggesting that partnership protection may play an important role in the health advantage of partnered individuals compared to non-partnered people. This is the first new finding of this study.

As we discussed in the literature review, a partnership provides advantages, which protect people against unhealthy behaviour and activities. Partnered people have more regular habits, more healthy diets and they exhibit more healthy lifestyles than non-partnered individuals (Mata et al. 2015; Oyebode et al. 2014). Partnership is also protective in various other ways: it reduces risk-taking behaviour; provides emotional and social support and acts as a buffering mechanism in the presence of stress. Furthermore, partnerships lead to higher (and more stable) income and partners benefit from the economies of the scale (Drefahl 2012). This is all indirectly supported by previous studies showing that partnered individuals have lower mortality than non-partnered people from circulatory, respiratory, digestive, nervous system, as well as alcohol- and accident-related (including self-harm) causes (Franke and Kulu 2018). Although differences in cancer mortality by partnership status are smaller than for other causes, separated individuals still have higher cancer mortality than married and cohabiting people (Franke and Kulu 2018).

The results mostly corroborate the protection hypothesis; however, we also observed some selection into and out of unions. First, the results for union formation were inconclusive. Our simultaneous analysis showed a surprising but weak selection of individuals with poor underlying health into partnerships. Whether this is related to pre-marital partnerships or post-marital cohabitations in later ages as observed by Lillard and Panis (1996) quarter of a century ago remains unclear[2]. Most importantly, these effects were not strong and, as our further analysis showed, were sensitive to different model specifications. Although our additional analysis using annual self-reported health showed that individuals with good health were slightly more likely to form unions than those with poor health (see Table A4 in Appendix), we could not use this as conclusive evidence either as annually measured health status may be shaped by the anticipation of a positive life event (e.g., marriage). Clearly, possible selection into first and subsequent cohabitations and marriages based on health status is a topic which requires further investigation.

Second, our analysis showed that individuals with underlying long-term health issues were more likely to experience union dissolution than those with better long-term health. Although this selection was not very strong, our models using both unobserved and observed health conditions supported its presence. It is likely that selection into partnership formation has declined over the years, i.e., partners' health characteristics are less important than they used to be (unless they have a strong impact on other life domains and activities, which is very unlikely at younger ages). However, underlying health issues may directly or indirectly (e.g., via employment, income etc.) increase partnership instability and lead to separation. This is the second new finding of this study.

Should we now accept that there is little selection by health status into and out of partnership and our analysis of mortality by partnership status supports the marriage protection theory? It is possible that self-reported health is not a good measure to identify underlying long-term health problems or frailty. One option is to use different (objective) health measures if they are available (this is not the case in this study). Alternatively, we can assume some heterogeneity in frailty across individuals, make this part of our model, and see what the effects are. This idea is not new (see the seminal paper by Vaupel et al. 1979). We can fix the value of residual variance of the mortality equation in a simultaneous equations model assuming that some individuals have poorer and some better underlying long-term health conditions (which will not change during their lives). Our further analysis showed that to observe a significant change in mortality by partnership status the variance of unobserved heterogeneity (or frailty) should be more than four times larger than observed in this study based on self-reported health. Such a huge variation in frailty and unequal distribution by partnership status is theoretically possible, but very unlikely in reality, especially given that our cumulative health measure explained away some variation in estimated frailty (although largely based on self-reported health). We thus have little evidence to support the presence of large unobserved heterogeneity in mortality net of observed covariates and strong selection effects, which would explain observed mortality differences by partnership status.

[2] Potential mechanisms of the selection of individuals with poor underlying health conditions into partnerships in older ages could include inheritance laws incentivizing unhealthy individuals to marry or health insurance and care benefits tied to partnership status.

Mortality by partnership status may vary by gender. Previous studies show that the differences are slightly larger for men than for women, but the patterns are very similar (Drefahl 2012; Franke and Kulu 2018). However, it is possible that there are gender differences in health selection, e.g., separated men with health issues may be more likely to search for a partner. This is an issue, which requires a larger sample than was available for this study. Our (further) analysis showed relatively similar patterns for men and women, but the sample was too small to conduct a detailed investigation on health selection. A larger sample might be also needed to replicate the study in a different context. The BHPS and UKHLS data have information on deaths, health, and partnership histories including information on non-marital unions, which are not available in population registers. However, the study may still be not long enough to follow the same individuals from the early part of their life course until the end of their lives, which may be ideally needed to investigate in detail selection based on frailty. Large cohort studies could be used if they have a long observation window and include information on mortality, health, as well as non-marital unions. In this study, the risk time for cohabitants is not negligible, but the number of death events among cohabitants is small. This is because spells of pre-marital cohabitation dominate in the data; long-term cohabitations and post-separation cohabitation are still not common and thus contribute little to the mortality risk time. Again, mortality among the latter group could be investigated only using large cohort studies. Previous research shows that mortality among post-marital cohabitants is higher than among married individuals, although their mortality levels are still lower than those of single and separated individuals (Drefahl 2021; Franke and Kulu 2018).

To conclude, using rich combined data from the British Household Panel Study and the UK Household Longitudinal Study, our study showed significant mortality differences by partnership status: partnered individuals had lower mortality than non-partnered people. Although the study detected some unobserved selection effects, i.e., individuals with underlying health issues were more likely to separate than those without health problems, the mortality differences between the groups persisted supporting the marital protection hypothesis and (with some caveats) its extension to non-marital partnerships.

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Table 1. Person-years and deaths by categories of variables.

| | Person-years | | Deaths | |
|---|---|---|---|---|
| | **Number** | **Percent** | **Number** | **Percent** |
| *Partnership status* | | | | |
| Single | 32640.45 | 19.2 | 153 | 11.0 |
| Cohabiting | 13686.60 | 8.0 | 19 | 1.4 |
| Married | 93846.30 | 55.2 | 629 | 45.3 |
| Separated | 17932.95 | 10.5 | 123 | 8.9 |
| Widowed | 11951.18 | 7.0 | 464 | 33.4 |
| *Age* | | | | |
| 16-39 years | 64111.35 | 37.7 | 24 | 1.7 |
| 40-44 years | 15311.81 | 9.0 | 21 | 1.5 |
| 45-49 years | 15618.68 | 9.2 | 33 | 2.4 |
| 50-54 years | 14602.21 | 8.6 | 41 | 3.0 |
| 55-59 years | 13043.69 | 7.7 | 55 | 4.0 |
| 60-64 years | 11845.25 | 7.0 | 100 | 7.2 |
| 65-69 years | 10521.31 | 6.2 | 121 | 8.7 |
| 70-74 years | 9279.07 | 5.5 | 188 | 13.5 |
| 75-79 years | 7469.50 | 4.4 | 225 | 16.2 |
| 80-84 years | 5004.04 | 2.9 | 237 | 17.1 |
| 85+ years | 3250.58 | 1.9 | 343 | 24.7 |
| *Sex* | | | | |
| Male | 78520.94 | 46.2 | 700 | 50.4 |
| Female | 91536.54 | 53.8 | 688 | 49.6 |
| *Period* | | | | |
| 1991-94 | 25889.35 | 15.2 | 173 | 12.5 |
| 1995-99 | 38541.68 | 22.7 | 392 | 28.2 |
| 2000-04 | 45800.00 | 26.9 | 415 | 29.9 |
| 2005-09 | 38816.04 | 22.8 | 292 | 21.0 |
| 2010-15 | 21010.42 | 12.4 | 116 | 8.4 |
| *Number of children* | | | | |
| No children | 54996.44 | 32.3 | 328 | 23.6 |
| One child | 25785.69 | 15.2 | 285 | 20.5 |
| Two children | 51006.21 | 30.0 | 389 | 28.0 |
| Three or more children | 38269.15 | 22.5 | 386 | 27.8 |
| *Educational level* | | | | |
| High | 30050.78 | 17.7 | 107 | 7.7 |
| Medium | 28854.29 | 17.0 | 82 | 5.9 |
| Low | 111152.42 | 65.4 | 1199 | 86.4 |
| *Tenure* | | | | |
| Homeowner | 123002.89 | 72.3 | 866 | 62.4 |
| Social rent | 29107.95 | 17.1 | 429 | 30.9 |
| Private rent | 17210.37 | 10.1 | 78 | 5.6 |
| Tenure missing | 736.28 | 0.4 | 15 | 1.1 |

| | | | | |
|---|---|---|---|---|
| *Place of residence* | | | | |
| London | 21745.26 | 12.8 | 153 | 11.0 |
| Large city | 19678.68 | 11.6 | 170 | 12.2 |
| Medium city | 32002.72 | 18.8 | 280 | 20.2 |
| Town | 20735.36 | 12.2 | 174 | 12.5 |
| Small town | 42108.78 | 24.8 | 347 | 25.0 |
| Rural area | 32882.13 | 19.3 | 261 | 18.8 |
| Missing | 904.56 | 0.5 | 3 | 0.2 |
| | | | | |
| Total | 170057.49 | 100.0 | 1388 | 100.0 |

Source: Calculations based on combined BHPS and UKHLS data; N=14847.

Table 2. Relative mortality rates by partnership status.

| Partnership status | Model 1 | | Model 2 | | Model 2a | | Model 3 | | Model 3a | |
|---|---|---|---|---|---|---|---|---|---|---|
| Single | 1.47 | ** | 1.23 | † | 1.24 | † | 1.35 | ** | 1.32 | * |
| Cohabiting | 0.99 | | 0.94 | | 1.02 | | 1.04 | | 1.05 | |
| Married | 1 | | 1 | | 1 | | 1 | | 1 | |
| Separated | 1.58 | ** | 1.44 | ** | 1.43 | ** | 1.46 | ** | 1.45 | ** |
| Widowed | 1.23 | ** | 1.16 | * | 1.15 | † | 1.25 | ** | 1.20 | * |

Source: Calculations based on combined BHPS and UKHLS data.
Significance: $^{\dagger}p < .10$; $^{*}p < .05$; $^{**}p < .01$.
Model 1: controlled for individuals' age, sex, and calendar time.
Model 2: additionally controlled for the number of children, educational level, housing tenure type and place of residence. (Model 2 vs Model 1: LR = 65.6, df = 13, $p < .001$.)
Model 2a: Model 2 and additionally controlled for cumulative self-reported health.
Model 3: Model 2 and additionally controlled for unobserved selection by health status. (Model 3 vs Model 2: LR = 18,969.2, df = 4, $p < .001$.)
Model 3a: Model 3 and additionally controlled for cumulative self-reported health.
Note: Full model results as log mortality rates are reported in Table 3.

Table 3. Log mortality rates by partnership status and other variables (parameter estimates and standard errors).

| Variables | Model 1 | | Model 2 | | Model 2a | | Model 3 | | Model 3a | |
|---|---|---|---|---|---|---|---|---|---|---|
| *Partnership status* | | | | | | | | | | |
| Single | 0.388 | ** | 0.210 | † | 0.213 | † | 0.298 | ** | 0.280 | * |
| | (0.095) | | (0.110) | | (0.110) | | (0.111) | | (0.110) | |
| Cohabiting | -0.012 | | -0.061 | | 0.015 | | 0.040 | | 0.047 | |
| | (0.236) | | (0.236) | | (0.237) | | (0.237) | | (0.237) | |
| Married | 0 | | 0 | | 0 | | 0 | | 0 | |
| Separated | 0.456 | ** | 0.367 | ** | 0.360 | ** | 0.379 | ** | 0.371 | ** |
| | (0.100) | | (0.101) | | (0.101) | | (0.102) | | (0.101) | |
| Widowed | 0.205 | ** | 0.149 | * | 0.138 | † | 0.225 | ** | 0.180 | * |
| | (0.072) | | (0.073) | | (0.073) | | (0.074) | | (0.074) | |
| *Age (baseline)* | | | | | | | | | | |
| Constant | -8.168 | ** | -8.163 | ** | -8.781 | ** | -8.225 | ** | -8.512 | ** |
| | (0.492) | | (0.505) | | (0.507) | | (0.506) | | (0.510) | |
| 16-34 years (slope) | 0.009 | | 0.012 | | 0.012 | | 0.005 | | 0.008 | |
| | (0.040) | | (0.040) | | (0.040) | | (0.040) | | (0.040) | |
| 35-39 years (slope) | 0.222 | † | 0.223 | † | 0.214 | † | 0.217 | † | 0.216 | † |
| | (0.118) | | (0.118) | | (0.118) | | (0.118) | | (0.118) | |
| 40-44 years (slope) | 0.152 | | 0.153 | | 0.151 | | 0.154 | | 0.152 | |
| | (0.095) | | (0.095) | | (0.095) | | (0.095) | | (0.095) | |
| 45-49 years (slope) | 0.024 | | 0.022 | | 0.012 | | 0.022 | | 0.016 | |
| | (0.075) | | (0.075) | | (0.075) | | (0.075) | | (0.075) | |
| 50-54 years (slope) | 0.083 | | 0.080 | | 0.071 | | 0.083 | | 0.076 | |
| | (0.066) | | (0.066) | | (0.066) | | (0.066) | | (0.066) | |
| 55-59 years (slope) | 0.128 | * | 0.125 | * | 0.119 | * | 0.125 | * | 0.123 | * |
| | (0.055) | | (0.055) | | (0.055) | | (0.055) | | (0.055) | |
| 60-64 years (slope) | 0.107 | * | 0.103 | * | 0.106 | * | 0.117 | ** | 0.113 | * |
| | (0.044) | | (0.044) | | (0.044) | | (0.044) | | (0.044) | |
| 65-69 years (slope) | 0.043 | | 0.039 | | 0.039 | | 0.045 | | 0.043 | |
| | (0.038) | | (0.038) | | (0.038) | | (0.038) | | (0.038) | |
| 70-74 years (slope) | 0.145 | ** | 0.142 | ** | 0.140 | ** | 0.153 | ** | 0.148 | ** |
| | (0.032) | | (0.032) | | (0.032) | | (0.032) | | (0.032) | |
| 75-79 years (slope) | 0.057 | * | 0.055 | * | 0.057 | * | 0.069 | * | 0.065 | * |
| | (0.028) | | (0.028) | | (0.028) | | (0.028) | | (0.028) | |
| 80-84 years (slope) | 0.130 | ** | 0.129 | ** | 0.127 | ** | 0.133 | ** | 0.131 | ** |
| | (0.025) | | (0.025) | | (0.025) | | (0.025) | | (0.025) | |
| 85+ years (slope) | 0.098 | ** | 0.100 | ** | 0.110 | ** | 0.123 | ** | 0.118 | ** |
| | (0.014) | | (0.014) | | (0.015) | | (0.015) | | (0.015) | |

| | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| *Sex* | | | | | | | | | | |
| Male | 0 | | 0 | | 0 | | 0 | | 0 | |
| Female | -0.446 | ** | -0.458 | ** | -0.540 | ** | -0.517 | ** | -0.534 | ** |
| | (0.056) | | (0.057) | | (0.057) | | (0.058) | | (0.058) | |
| *Period* | | | | | | | | | | |
| 1991-94 | 0 | | 0 | | 0 | | 0 | | 0 | |
| 1995-99 | 0.294 | ** | 0.297 | ** | 0.472 | ** | 0.334 | ** | 0.405 | ** |
| | (0.091) | | (0.091) | | (0.092) | | (0.092) | | (0.092) | |
| 2000-04 | 0.081 | | 0.111 | | 0.397 | ** | 0.154 | † | 0.312 | ** |
| | (0.091) | | (0.091) | | (0.095) | | (0.092) | | (0.095) | |
| 2005-09 | -0.225 | * | -0.164 | † | 0.131 | | -0.128 | | 0.034 | |
| | (0.097) | | (0.098) | | (0.101) | | (0.098) | | (0.101) | |
| 2010-15 | -0.641 | ** | -0.556 | ** | -0.285 | * | -0.544 | ** | -0.380 | ** |
| | (0.121) | | (0.122) | | (0.125) | | (0.123) | | (0.125) | |
| *Number of children* | | | | | | | | | | |
| No children | | | 0 | | 0 | | 0 | | 0 | |
| One child | | | -0.082 | | -0.095 | | -0.018 | | -0.065 | |
| | | | (0.089) | | (0.089) | | (0.090) | | (0.089) | |
| Two children | | | -0.184 | * | -0.139 | | -0.069 | | -0.105 | |
| | | | (0.086) | | (0.086) | | (0.087) | | (0.086) | |
| Three or more children | | | -0.186 | * | -0.200 | * | -0.151 | † | -0.186 | * |
| | | | (0.086) | | (0.086) | | (0.087) | | (0.086) | |
| *Place of residence* | | | | | | | | | | |
| London | | | 0.017 | | 0.073 | | 0.048 | | 0.064 | |
| | | | (0.112) | | (0.112) | | (0.112) | | (0.112) | |
| Large city | | | 0 | | 0 | | 0 | | 0 | |
| City | | | 0.128 | | 0.102 | | -0.009 | | 0.065 | |
| | | | (0.097) | | (0.098) | | (0.098) | | (0.099) | |
| Town | | | 0.033 | | 0.040 | | 0.024 | | 0.036 | |
| | | | (0.108) | | (0.108) | | (0.108) | | (0.108) | |
| Small town | | | 0.113 | | 0.183 | † | 0.111 | | 0.182 | † |
| | | | (0.094) | | (0.095) | | (0.094) | | (0.095) | |
| Rural | | | -0.062 | | 0.050 | | -0.029 | | 0.064 | |
| | | | (0.099) | | (0.101) | | (0.100) | | (0.101) | |
| *Tenure* | | | | | | | | | | |
| Homeowner | | | 0 | | 0 | | 0 | | 0 | |
| Social rent | | | 0.306 | ** | 0.177 | ** | 0.216 | ** | 0.188 | ** |
| | | | (0.063) | | (0.063) | | (0.063) | | (0.063) | |

| | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Private rent | -0.127 | | -0.143 | | -0.103 | | -0.128 | |
| | (0.121) | | (0.121) | | (0.121) | | (0.121) | |
| Tenure missing | 0.744 | ** | 0.371 | | 0.566 | * | 0.422 | |
| | (0.269) | | (0.272) | | (0.273) | | (0.273) | |
| High | -0.278 | † | -0.207 | | -0.154 | | -0.165 | |
| | (0.148) | | (0.148) | | (0.149) | | (0.148) | |
| Medium | 0 | | 0 | | 0 | | 0 | |
| | | | | | | | | |
| Low | 0.070 | | -0.008 | | 0.008 | | 0.002 | |
| | (0.118) | | (0.118) | | (0.119) | | (0.119) | |
| *Cumulative self-reported health* | | | | | | | | |
| Excellent | | | 0 | | | | 0 | |
| | | | | | | | | |
| Good | | | 0.588 | ** | | | 0.277 | ** |
| | | | (0.087) | | | | (0.098) | |
| Fair | | | 1.288 | ** | | | 0.605 | ** |
| | | | (0.095) | | | | (0.137) | |
| Poor and very poor | | | 1.962 | ** | | | 0.963 | ** |
| | | | (0.122) | | | | (0.187) | |
| Missing | | | 0.390 | ** | | | 0.022 | |
| | | | (0.108) | | | | (0.123) | |

Source: Calculations based on combined BHPS and UKHLS data.
Notes: For age we present slope estimates which show how the log-hazard increases or decreases over a certain duration.
Significance: : † p <.10; * p <.05; ** p <.01.

Table 4. Standard deviations and correlations between person-specific residuals.

| | **Model 1** | | **Model 2** | | **Model 2a** | | **Model 3** | | **Model 3a** | |
|---|---|---|---|---|---|---|---|---|---|---|
| *Standard deviations* | | | | | | | | | | |
| Union formation | 3.22 | ** | 3.22 | ** | 3.22 | ** | 3.21 | ** | 3.21 | ** |
| Union dissolution | 1.50 | ** | 1.50 | ** | 1.50 | ** | 0.35 | ** | 0.35 | ** |
| Health | 0.52 | ** | 0.52 | ** | 0.52 | ** | 0.61 | ** | 0.61 | ** |
| Loading factor for mortality | | | | | | | 0.80 | ** | 0.51 | ** |
| *Correlations* | | | | | | | | | | |
| Union formation and dissolution | | | | | | | 0.15 | † | 0.15 | † |
| Union formation and health | | | | | | | 0.15 | ** | 0.15 | ** |
| Union dissolution and health | | | | | | | 0.51 | ** | 0.51 | ** |

Source: Calculations based on combined BHPS and UKHLS data.
Significance: † p <.10; * p <.05; ** p <.01.
Model 1: controlled for the individual's age, sex and calendar time.
Model 2: additionally controlled for the number of children, educational level, tenure type and place of residence. (Model 2 vs Model 1: LR = 65.6, df = 13, p < .001.)
Model 2a: Model 2 and additionally controlled for cumulative self-reported health.
Model 3: Model 2 and additionally controlled for unobserved selection by health status. (Model 3 vs Model 2: LR = 18,969.2, df = 4, p < .001.)
Model 3a: Model 3 and additionally controlled for cumulative self-reported health.

Table A1. Person-years and union formations by categories of variables.

| | Person-years | | Union formation | |
|---|---|---|---|---|
| | Number | Percent | Number | Percent |
| *Partnership status* | | | | |
| Single | 30954.60 | 51.1 | 3875 | 68.0 |
| Separated | 17874.83 | 29.5 | 1746 | 30.6 |
| Widowed | 11725.18 | 19.4 | 79 | 1.4 |
| *Time since separation* | | | | |
| No separation | 30955.89 | 51.1 | 3875 | 68.0 |
| 0-1 year | 2990.29 | 4.9 | 475 | 8.3 |
| 1-3 years | 4805.47 | 7.9 | 568 | 10.0 |
| 3-5 years | 3750.18 | 6.2 | 290 | 5.1 |
| 5+ years | 18052.78 | 29.8 | 492 | 8.6 |
| *Age* | | | | |
| 16-19 years | 8993.55 | 14.9 | 330 | 5.8 |
| 20-24 years | 9885.86 | 16.3 | 1058 | 18.6 |
| 25-29 years | 5401.73 | 8.9 | 1004 | 17.6 |
| 30-34 years | 3644.09 | 6.0 | 729 | 12.8 |
| 35-39 years | 3232.28 | 5.3 | 555 | 9.7 |
| 40-44 years | 3201.97 | 5.3 | 431 | 7.6 |
| 45-49 years | 3213.18 | 5.3 | 371 | 6.5 |
| 50-54 years | 2907.18 | 4.8 | 330 | 5.8 |
| 55-59 years | 2648.89 | 4.4 | 252 | 4.4 |
| 60-64 years | 2607.68 | 4.3 | 221 | 3.9 |
| 65-69 years | 2887.42 | 4.8 | 171 | 3.0 |
| 70-74 years | 3262.29 | 5.4 | 135 | 2.4 |
| 75+ years | 8668.49 | 14.3 | 113 | 2.0 |
| *Sex* | | | | |
| Male | 25779.09 | 42.6 | 2657 | 46.6 |
| Female | 34775.52 | 57.4 | 3043 | 53.4 |
| *Period* | | | | |
| 1991-94 | 8830.98 | 14.6 | 877 | 15.4 |
| 1995-99 | 13631.83 | 22.5 | 1716 | 30.1 |
| 2000-04 | 16234.38 | 26.8 | 1785 | 31.3 |
| 2005-09 | 14618.65 | 24.1 | 895 | 15.7 |
| 2010-15 | 7238.76 | 12.0 | 427 | 7.5 |
| *Number of children* | | | | |
| No children | 35107.73 | 58.0 | 2980 | 52.3 |
| One child | 7267.38 | 12.0 | 804 | 14.1 |
| Two children | 9310.80 | 15.4 | 1122 | 19.7 |
| Three or more children | 8868.70 | 14.6 | 794 | 13.9 |
| *Educational level* | | | | |
| High | 9113.61 | 15.1 | 1139 | 20.0 |
| Medium | 11669.69 | 19.3 | 1168 | 20.5 |
| Low | 39771.31 | 65.7 | 3393 | 59.5 |
| *Place of residence* | | | | |
| London | 8688.50 | 14.3 | 586 | 10.3 |
| Large city | 7232.43 | 11.9 | 468 | 8.2 |

| | | | | |
|---|---|---|---|---|
| Medium city | 11737.86 | 19.4 | 1248 | 21.9 |
| Town | 7047.26 | 11.6 | 616 | 10.8 |
| Small town | 15141.59 | 25.0 | 1468 | 25.8 |
| Rural area | 10503.64 | 17.3 | 1269 | 22.3 |
| Missing | 203.33 | 0.3 | 45 | 0.8 |
| *Employment status* | | | | |
| Self-employed | 2600.41 | 4.3 | 427 | 7.5 |
| Employee | 25493.65 | 42.1 | 3383 | 59.4 |
| Unemployed | 3913.07 | 6.5 | 343 | 6.0 |
| Other | 27952.30 | 46.2 | 1415 | 24.8 |
| Missing | 595.17 | 1.0 | 132 | 2.3 |
| *Self-rated health* | | | | |
| Excellent | 13664.86 | 22.6 | 1411 | 24.8 |
| Very good | 21235.85 | 35.1 | 2374 | 41.6 |
| Fair | 15174.20 | 25.1 | 1263 | 22.2 |
| Poor | 6760.94 | 11.2 | 472 | 8.3 |
| Very poor | 2630.01 | 4.3 | 130 | 2.3 |
| Missing | 1088.74 | 1.8 | 50 | 0.9 |
| Total | 60554.61 | 100.0 | 5700 | 100.0 |

Source: Calculations based on combined BHPS and UKHLS data; N=10393.

Table A2. Person-years and union dissolutions by categories of variables.

| | Person-years | | Union dissolution | |
|---|---|---|---|---|
| | Number | Percent | Number | Percent |
| *Partnership status* | | | | |
| Cohabiting | 13599.76 | 12.7 | 1330 | 40.0 |
| Married | 93072.71 | 87.3 | 1994 | 60.0 |
| *Union duration* | | | | |
| 0-1 year | 5450.28 | 5.1 | 545 | 16.4 |
| 1-3 years | 9209.27 | 8.6 | 702 | 21.1 |
| 3-5 years | 7777.22 | 7.3 | 400 | 12.0 |
| 5+ years | 84235.69 | 79.0 | 1677 | 50.5 |
| *Age at union formation* | | | | |
| 16-19 | 9371.86 | 8.8 | 443 | 13.3 |
| 20-24 | 39604.03 | 37.1 | 1117 | 33.6 |
| 25-29 | 21789.92 | 20.4 | 605 | 18.2 |
| 30-34 | 10696.99 | 10.0 | 375 | 11.3 |
| 35+ | 25209.66 | 23.6 | 784 | 23.6 |
| *Order of separation* | | | | |
| First | 83665.00 | 78.4 | 2247 | 67.6 |
| Second+ | 23007.46 | 21.6 | 1077 | 32.4 |
| *Sex* | | | | |
| Male | 51215.88 | 48.0 | 1544 | 46.5 |
| Female | 55456.58 | 52.0 | 1780 | 53.5 |
| *Period* | | | | |
| 1991-94 | 16755.78 | 15.7 | 569 | 17.1 |
| 1995-99 | 24486.17 | 23.0 | 705 | 21.2 |
| 2000-04 | 29149.49 | 27.3 | 985 | 29.6 |
| 2005-09 | 23922.46 | 22.4 | 842 | 25.3 |
| 2010-15 | 12358.56 | 11.6 | 223 | 6.7 |
| *Number of children* | | | | |
| No children | 19832.79 | 18.6 | 1053 | 31.7 |
| One child | 17162.63 | 16.1 | 576 | 17.3 |
| Two children | 40807.04 | 38.3 | 978 | 29.4 |
| Three or more children | 28870.01 | 27.1 | 717 | 21.6 |
| *Educational level* | | | | |
| High | 20683.41 | 19.4 | 581 | 17.5 |
| Medium | 16827.91 | 15.8 | 667 | 20.1 |
| Low | 69161.15 | 64.8 | 2076 | 62.5 |
| *Place of residence* | | | | |
| London | 12782.02 | 12.0 | 387 | 11.6 |
| Large city | 12104.45 | 11.3 | 342 | 10.3 |
| Medium city | 19748.91 | 18.5 | 640 | 19.3 |
| Town | 13355.32 | 12.5 | 424 | 12.8 |
| Small town | 26364.26 | 24.7 | 888 | 26.7 |
| Rural area | 21938.56 | 20.6 | 622 | 18.7 |
| Missing | 378.96 | 0.4 | 21 | 0.6 |
| *Employment status* | | | | |
| Self-employed | 9541.86 | 8.9 | 335 | 10.1 |

| | | | | |
|---|---|---|---|---|
| Employee | 56649.72 | 53.1 | 2027 | 61.0 |
| Unemployed | 3620.63 | 3.4 | 191 | 5.7 |
| Other | 35704.89 | 33.5 | 738 | 22.2 |
| Missing | 1155.36 | 1.1 | 33 | 1.0 |
| *Self-rated health* | | | | |
| Excellent | 24726.32 | 23.2 | 793 | 23.9 |
| Very good | 36352.61 | 34.1 | 1059 | 31.9 |
| Fair | 28451.38 | 26.7 | 867 | 26.1 |
| Poor | 11962.07 | 11.2 | 390 | 11.7 |
| Very poor | 3754.08 | 3.5 | 134 | 4.0 |
| Missing | 1426.00 | 1.3 | 81 | 2.4 |
| Total | 106672.47 | 100.0 | 3324 | 100.0 |

Source: Calculations based on combined BHPS and UKHLS data; N=11449.

Table A3. Descriptive statistics for annual health by categories of variables.

| | Mean | SE |
|---|---|---|
| *Partnership status* | | |
| Single | 2.13 | 0.004 |
| Cohabiting | 2.10 | 0.008 |
| Married | 2.25 | 0.003 |
| Separated | 2.50 | 0.009 |
| Widowed | 2.64 | 0.008 |
| *Age* | | |
| 16-19 years | 1.99 | 0.007 |
| 20-24 years | 2.04 | 0.007 |
| 25-29 years | 2.04 | 0.007 |
| 30-34-54 years | 2.05 | 0.007 |
| 35-39 years | 2.10 | 0.007 |
| 40-44 years | 2.15 | 0.007 |
| 45-49 years | 2.22 | 0.007 |
| 50-54 years | 2.30 | 0.008 |
| 55-59 years | 2.39 | 0.008 |
| 60-64 years | 2.42 | 0.009 |
| 65-69 years | 2.49 | 0.009 |
| 70-74 years | 2.55 | 0.010 |
| 75+ years | 2.75 | 0.008 |
| *Sex* | | |
| Male | 2.20 | 0.003 |
| Female | 2.32 | 0.003 |
| *Period* | | |
| 1991-94 | 2.12 | 0.005 |
| 1995-99 | 2.19 | 0.004 |
| 2000-04 | 2.23 | 0.004 |
| 2005-09 | 2.20 | 0.005 |
| 2010-15 | 2.62 | 0.006 |
| *Number of children* | | |
| No children | 2.33 | 0.003 |
| One child | 2.12 | 0.005 |
| Two children | 2.08 | 0.006 |
| Three or more children | 2.17 | 0.009 |
| *Educational level* | | |
| High | 2.40 | 0.003 |
| Medium | 2.18 | 0.006 |
| Low | 2.16 | 0.003 |
| Missing | 2.19 | 0.015 |
| *Place of residence* | | |
| London | 2.52 | 0.020 |
| Large city | 2.67 | 0.021 |
| Medium city | 2.67 | 0.016 |
| Town | 2.61 | 0.019 |
| Small town | 2.67 | 0.013 |
| Rural area | 2.62 | 0.014 |

| | | |
|---|---|---|
| Missing | 2.20 | 0.002 |
| *Tenure* | | |
| Homeowner | 2.19 | 0.002 |
| Social rent | 2.57 | 0.006 |
| Private rent | 2.26 | 0.008 |
| Tenure missing | 2.32 | 0.021 |
| *Employment status* | | |
| Self-employed | 2.01 | 0.007 |
| Employee | 2.04 | 0.003 |
| Unemployed | 2.36 | 0.011 |
| Other | 2.57 | 0.004 |
| Missing | 2.16 | 0.031 |
| | | |
| Total | 2.26 | 0.002 |

Source: Calculations based on combined BHPS and UKHLS data; N=14847.
The mean of measurements is 12.3.

Table A4. Parameter estimates and standard errors for models on union formation, union dissolution, and health.

| **Union formation** | | | **Union dissolution** | | | **Health** | | |
|---|---|---|---|---|---|---|---|---|
| **Variables** | **Model 3** | | **Variables** | **Model 3** | | **Variables** | **Model 3** | |
| *Age* | | | *Union duration* | | | *Age* | | |
| Constant | -10.277 | ** | Constant | -2.857 | ** | Constant | 1.951 | ** |
| | (0.309) | | | (0.131) | | | (0.042) | |
| 16-19 years (slope) | 1.163 | ** | 0–1 years (slope) | 0.533 | ** | 16-19 years | 0 | |
| | (0.076) | | | (0.113) | | | | |
| 20-24 years (slope) | 0.617 | ** | 1-3 years (slope) | -0.271 | ** | 20-24 years | 0.099 | ** |
| | (0.030) | | | (0.044) | | | (0.010) | |
| 25-29 years (slope) | 0.427 | ** | 3-5 years (slope) | -0.126 | ** | 25-29 years | 0.106 | ** |
| | (0.024) | | | (0.037) | | | (0.012) | |
| 30-34 years (slope) | -0.080 | ** | 5+ years (slope) | -0.042 | ** | 30-34 years | 0.108 | ** |
| | (0.021) | | | (0.002) | | | (0.014) | |
| 35-39 years (slope) | -0.024 | | *Age at union formation* | | | 35-39 years | 0.166 | ** |
| | (0.024) | | 16-19 years | 0.375 | ** | | (0.014) | |
| 40-44 years (slope) | -0.084 | ** | | (0.061) | | 40-44 years | 0.203 | ** |
| | (0.026) | | 20-24 years | 0 | | | (0.015) | |
| 45-49 years (slope) | 0.035 | | | | | 45-49 years | 0.268 | ** |
| | (0.028) | | 25-29 years | -0.354 | ** | | (0.016) | |
| 50-54 years (slope) | -0.037 | | | (0.054) | | 50-54 years | 0.323 | ** |
| | (0.031) | | 30-34 years | -0.405 | ** | | (0.016) | |
| 55-59 years (slope) | 0.029 | | | (0.067) | | 55-59 years | 0.370 | ** |
| | (0.034) | | 35+ years | -0.709 | ** | | (0.017) | |
| 60-64 years (slope) | -0.067 | † | | (0.057) | | 60-64 years | 0.348 | ** |
| | (0.037) | | *Union order* | | | | (0.018) | |
| 65-69 years (slope) | -0.084 | * | First | 0 | | 65-69 years | 0.384 | ** |
| | (0.041) | | | | | | (0.018) | |
| 70-74 years (slope) | -0.069 | | Second and subsequent | 0.345 | ** | 70-74 years | 0.478 | ** |
| | (0.048) | | | (0.048) | | | (0.019) | |
| 75+ years (slope) | -0.367 | ** | *Union type* | | | 75+ years | 0.714 | ** |
| | (0.035) | | Marriage | 0 | | | (0.020) | |
| *Partnership status* | | | | | | *Partnership status* | | |
| Single | 0 | | Cohabiting | 0.679 | ** | Single | 0.021 | † |
| | | | | (0.049) | | | (0.011) | |
| Separated | -6.608 | ** | *Sex* | | | Cohabiting | 0.034 | ** |
| | (0.069) | | Male | 0 | | | (0.011) | |
| *Sex* | | | | | | Married | 0 | |
| Male | 0 | | Female | -0.033 | | | | |
| | | | | (0.038) | | Separated | 0.056 | ** |
| Female | 0.149 | * | *Period* | | | | (0.011) | |
| | (0.060) | | 1991-94 | 0 | | Widowed | 0.058 | ** |

| | | |
|---|---|---|
| *Period* | | |
| 1991-94 | 0 | |
| 1995-99 | 0.520 | ** |
| | (0.062) | |
| 2000-04 | 0.386 | ** |
| | (0.066) | |
| 2005-09 | 0.085 | |
| | (0.074) | |
| 2010-15 | 0.081 | |
| | (0.087) | |
| *Number of children* | | |
| No children, not pregnant | 0 | |
| Pregnant | 3.522 | ** |
| | (0.114) | |
| One child | 3.349 | ** |
| | (0.092) | |
| Two children | 3.414 | ** |
| | (0.090) | |
| Three or more children | 3.349 | ** |
| | (0.093) | |
| *Place of residence* | | |
| London | 0.063 | |
| | (0.127) | |
| Large city | 0 | |
| City | 0.502 | ** |
| | (0.113) | |
| Town | 0.630 | ** |
| | (0.122) | |
| Small town | 0.513 | ** |
| | (0.110) | |
| Rural | 0.749 | ** |
| | (0.108) | |
| *Employment status* | | |
| Employed | 0 | |
| Self-employed | 0.207 | * |
| | (0.084) | |
| In education | -1.234 | ** |
| | (0.111) | |
| Unemployed | -0.245 | ** |

| | | |
|---|---|---|
| 1995-99 | -0.154 | ** |
| | (0.057) | |
| 2000-04 | -0.019 | |
| | (0.056) | |
| 2005-09 | 0.125 | * |
| | (0.058) | |
| 2010-15 | -0.508 | ** |
| | (0.083) | |
| *Number of children* | | |
| No children, not pregnant | 0 | |
| Pregnant | -1.335 | ** |
| | (0.159) | |
| One child | -0.081 | |
| | (0.057) | |
| Two children | -0.084 | |
| | (0.054) | |
| Three or more children | 0.008 | |
| | (0.060) | |
| *Place of residence* | | |
| London | 0.057 | |
| | (0.075) | |
| Large city | 0 | |
| City | 0.014 | |
| | (0.068) | |
| Town | 0.048 | |
| | (0.074) | |
| Small town | 0.101 | |
| | (0.064) | |
| Rural | -0.004 | |
| | (0.069) | |
| *Employment status* | | |
| Employed | 0 | |
| Self-employed | 0.174 | ** |
| | (0.061) | |
| In education | 0.499 | ** |
| | (0.117) | |
| Unemployed | 0.125 | |
| | (0.079) | |
| Other | -0.203 | ** |

| | | |
|---|---|---|
| | (0.013) | |
| *Sex* | | |
| Male | 0 | |
| Female | 0.026 | * |
| | (0.011) | |
| *Period* | | |
| 1991-94 | 0 | |
| 1995-99 | 0.103 | ** |
| | (0.006) | |
| 2000-04 | 0.157 | ** |
| | (0.006) | |
| 2005-09 | 0.156 | ** |
| | (0.007) | |
| 2010-15 | 0.446 | ** |
| | (0.034) | |
| *Number of children* | | |
| No children | 0 | |
| One child | -0.026 | ** |
| | (0.007) | |
| Two children | -0.051 | ** |
| | (0.008) | |
| Three or more children | -0.048 | ** |
| | (0.012) | |
| *Place of residence* | | |
| London | -0.106 | ** |
| | (0.024) | |
| Large city | 0 | |
| City | -0.014 | |
| | (0.021) | |
| Town | -0.018 | |
| | (0.023) | |
| Small town | -0.016 | |
| | (0.020) | |
| Rural | -0.005 | |
| | (0.021) | |
| Missing | -0.214 | ** |
| | (0.038) | |
| *Tenure* | | |
| Homeowner | 0 | |

| | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | (0.085) | | | (0.050) | | | | |
| Other | -0.128 | * | Missing | -0.032 | | Social rent | 0.105 | ** |
| | (0.065) | | | (0.178) | | | (0.009) | |
| Missing | 0.682 | ** | *Educational level* | | | Private rent | 0.033 | ** |
| | (0.197) | | High | -0.109 | † | | (0.009) | |
| *Educational level* | | | | (0.059) | | Tenure missing | 0.226 | ** |
| High | 0.135 | | Medium | 0 | | | (0.016) | |
| | (0.088) | | | | | *Employment status* | | |
| Medium | 0 | | Low | 0.044 | | Employed | 0 | |
| | | | | (0.048) | | | | |
| Low | -0.024 | | *Self-reported health* | | | Self-employed | -0.047 | ** |
| | (0.067) | | Excellent | 0 | | | (0.010) | |
| *Self-reported health* | | | | | | Unemployed | 0.105 | ** |
| Excellent | 0 | | Very good | -0.081 | † | | (0.011) | |
| | | | | (0.049) | | Other | 0.215 | ** |
| Very good | 0.188 | ** | Fair | -0.039 | | | (0.006) | |
| | (0.054) | | | (0.053) | | Missing | 3.035 | ** |
| Fair | -0.067 | | Poor | 0.006 | | | (0.013) | |
| | (0.062) | | | (0.071) | | *Educational level* | | |
| Poor | -0.238 | ** | Very poor | 0.078 | | High | -0.106 | ** |
| | (0.083) | | | (0.109) | | | (0.008) | |
| Very poor | -0.554 | ** | Missing | 0.622 | ** | Medium | -0.033 | ** |
| | (0.141) | | | (0.122) | | | (0.009) | |
| Missing | -0.915 | ** | | | | Low | 0 | |
| | (0.201) | | | | | | | |
| | | | | | | Missing | 0.339 | ** |
| | | | | | | | (0.018) | |

Source: Calculations based on combined BHPS and UKHLS data; N=14847.