

I Can't Smile Without You: Spousal Correlation in Life Satisfaction

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Abstract

This paper tests whether one partner's happiness significantly influences the happiness of the other partner. Using ten waves of the British Household Panel Survey, it utilizes a panel-based GMM methodology to estimate a dynamic model of life satisfaction. The use of the GMM-system estimator corrects for correlated effects of partner's life satisfaction and solves the problem of measurement error bias. The results show that, for both genders, there is a positive and statistically significant spillover effect of life satisfaction that runs from one partner to the other partner in a couple. The positive bias on the estimated spillover effect coming from assortative mating and shared social environment at cross-section is almost offset by the negative bias coming from systematic measurement errors in the way people report their life satisfaction. Moreover, consistent with the spillover effect model, couple dissolution at $t+1$ is negatively correlated with partners' life satisfaction at t .

Key words: life satisfaction; assortative mating; spillover; marriage; happiness; GMM

JEL: D1, D62, D64, I0

“You see, I feel sad when you’re sad, I feel glad when you’re glad. If you only knew what I’m going through, I just can’t smile without you.” - *Barry Manilow*

1. Introduction

The idea that married people care a great deal about the well-being of their partner is not new to economists (Becker, 1973, 1974; Friedman, 1986). The past three decades have seen a significant increase in the number of studies showing that people in marriage tend to behave altruistically towards their partner (see, for example, Foster and Rosenzweig, 2001; Ermisch, 2003). However, while it may be possible to make some inferences about the degree of caring between partners from their behaviour, the idea that there may be a direct spillover of well-being from one partner to the other has rarely been tested empirically.

This paper aims to do just that. Using a long-run panel of nationally representative randomly sampled married and cohabiting individuals, it examines the extent of spousal correlation in subjective well-being data, particularly self-rated life satisfaction (LS). It proposes that a positive correlation between partners’ LS may reflect three distinct processes. First, individuals who are born happy or are born with innate predispositions that make them happy may tend to marry those who are similar to them. In addition to this, people of the same family background or life styles – in other words, same unobserved social factors – may also tend to marry each other. This matching of fixed personal characteristics on the marriage market is analogous to the concept of assortative mating (Becker, 1974). Manski (1995) refers to such phenomena as correlated effects of social interactions.

Second, given that marriage allows individuals to share with their partner the kind of physical and emotional resources that may not be available for each person to obtain outside marriage (Waite and Gallagher, 2000), correlated effects may also arise from the shared social

environment (which can either be time-invariant or time-variant) that is simultaneously affecting LS for both spouses.

Lastly, the observed correlation may be the result of a direct spillover of LS within the couple. The idea is that, if a husband cares about his wife, then her LS becomes one of the main determinants of his own LS, and vice versa. This generates a possibility that a husband will be *ceteris paribus* happier when his wife is happier for whatever reasons that make her happy but not him directly. Hence, we would expect an increase in one partner's LS to be positively correlated with the other partner's LS even after allowing for all the factors that can affect both partners' LS at the same time. This phenomenon is likened to the endogenous effects in Manski's terminology, whereby the individual outcome is a function of group achievement.

In addition to the above confounding influences which make it difficult for the true relationship between partners' well-being to be identified, the estimates of spousal correlation in LS may also suffer from the negative measurement error bias. There may be, for example, a tendency for individuals to misreport their true LS in surveys. The low signal-to-noise ratio caused by misreporting can result in an estimated coefficient on partner LS that is biased towards zero in a large sample. In short, because there are both positive (correlated effects) and negative (measurement error) biases involved, the direction of bias is unclear on *a priori* ground.

This paper uses ten waves of the British Household Panel Survey (BHPS) data to examine the extent of spousal correlation in LS. In particular, it uses the "system GMM estimator" proposed by Arellano and Bover (1995) and Blundell and Bond (1998) to estimate the causal spillover effect that runs from one partner's life satisfaction to the other partner's life satisfaction. The use of the GMM-system estimator, which is a unique approach in the study of happiness, control for the correlated effects and solve the problem of measurement

error bias in self-rated life satisfaction through instrumentations and first-differencing. The results show that there is strong evidence of a spillover effect of LS, which suggests that well-being is transferable from one partner to the other. Consistent with the spillover effect model, partners' LS today are also associated with lower probabilities of partners separating or divorcing one period into the future.

There are similarities in terms of research questions and analytic strategy between this paper and previous studies that examined similarities in a husband's and wife's behaviour such as smoking (Clark and Etile, 2006), their political preferences (Kan and Heath, 2006), and their sporting activities (Farrell and Shields, 2002).

This article is organised as follows: Section 2 reviews relevant past research on marriage and well-being. Section 3 addresses theoretical issues revolving around the various interpretations of the correlation between partners' LS. Section 4 describes how to implement a test and the data set. Section 5 discusses the results, and Section 6 concludes.

2. Marriage, subjective well-being, and spillovers

Previous research on marital status and emotional well-being is clear on one point: married persons are significantly happier and more satisfied with life than those who are divorced, separated, widowed, or single, across various countries and time periods (Gove et al., 1983; Mastekaasa, 1994; Marks and Lambert, 1998). The large psychological benefits of marriage persist even when the selection of happy people into marriage is controlled for in the analysis (Frey and Stutzer, 2006; Mastekaasa, 1992), and such advantages are sometimes shown to be stronger for men than for women (see, for example, Gove et al., 1983). Both cross-sectional and longitudinal studies confirm the overall psychological benefits of marriage (for a review, see Oswald and Wilson, 2005).

There are several explanations for the protective effects of marriage. First, on the grounds that two can live almost cheaply as one, marriage may work simply because it provides higher real income per partner (Korenman and Neumark, 1991; Loh, 1996; Smock et al., 1999). Second, marriage provides the couple with a source of constant emotional and instrumental support, which may act as an important buffer against stress and depression for the person who experiences negative shocks in life events (Berkman, 1988; Marks and Lambert, 1998). In other words, the negative impacts of shocks in life events appear to be significantly lower for married individuals than for those of other marital groups. Third, marriage provides the couple with a sense of belonging and social reality, in which they are the only two people living and operating in their own world. This shared sense of meaning can be an important foundation for emotional well-being (Berger and Kellner, 1964; House et al., 1982). Marriage also encourages people to engage less in risky activities and more in healthy ones – perhaps for the sake of their partner. For example, married people smoke and drink less, and such healthy behaviour may provide an important source of both physical and emotional well-being for the couple. The results hold even when one allows for selection effects into marriage (see Power et al., 1999).

What has received much less attention is whether one partner's well-being is a function of the other partner's well-being (Becker, 1974). Previous studies on emotional spillover have often focused on daily transmissions of only negative effects, such as measures of stress and strain. One of the common findings in the literature is that stress experienced by one partner in the workplace has the tendency to heighten the level of stress being experienced by the other partner at home (see Bolger et al., 1989; Jones and Fletcher, 1993; Westman and Vinokur, 1998). Yet this is not a persuasive reason to believe that well-being is transferable within couples. One reason for this is that there is evidence in the psychology literature that one measure is not just a mirror of the other. For instance, whilst several studies have found a

moderate correlation between ill-being and well-being (Chamberlain, 1988; Michalos, 1991), others have shown that these components appear to behave differently over time and to have differing relationships with other variables (Liang, 1985; Stock et al., 1986; Huppert and Whittington, 2003). Research into the validity of the two constructs has also shown that there is a clear distinction in terms of determinants between measures of well-being and ill-being (see Bradburn, 1969; Diener et al., 1999, Headey et al., 1993; Pavot and Diener, 1993). Because measures of cognitive well-being such as LS frequently form a separate factor and correlate with predictor variables in a unique way, it seems worthwhile to separately assess this construct in the research.

Of the very few studies on the topic, Rose's (1955) was one of the first to report cross-sectional correlation between LS within marriage. The author showed that spouses may feel something is wrong with their marriage when one or both feel unhappy with life. In that case, even if there was nothing wrong with their marriage, low levels of happiness detected between couples may have shattered their confidence and lead to separation and divorce. Argyle (1999) made a conjecture in his study on the effect of marriage on subjective well-being that one spouse's happiness may encourage the happiness of the other in a marriage. More recently, Anderson et al. (2003) found that people in early dating relationships, i.e. the first six months of dating, tend to report similar levels of positive emotional experiences over time, such as happiness, amusement, and pride. Plug and Van Praag (1998) found similarities in the reported income satisfaction by members of the same household. In a similar study, Powdthavee and Vignoles (2008) also found significant spillover effects in reported well-being between parents and children.

To the best of my knowledge, the only paper that has conducted a longitudinal analysis on whether there is a substantial long-term interdependent relationship between spouses' LS within marriage is the innovative work by Schimmack and Lucas (2006). Their methods and

dataset differ from those set out in this paper, and the respective projects were begun independently. Using the German panel data and time-lagged cross-spouse correlation method, they found that spousal correlation in LS is due mainly to the shared stable component of LS within the couple, i.e. partners sharing similar traits and social environments. Little was discussed in their paper, however, on the possibility that there may be a spillover effect of LS from one partner to the other. This is the main difference between this paper's analysis and that adopted by Schimmack and Lucas (2006).

3. Theory

In this section, I will briefly discuss the three underlying mechanisms that may account for the raw correlation between a husband's and wife's LS levels: assortative mating, shared social environment, and spillover effect.

3.1 Assortative mating

The first explanation is that the observed spousal correlation in LS may have been the outcome of a matching of unobserved personality traits and/or social factors on the marriage market (Becker, 1974). Individuals may prefer partners who are phenotypically similar to them. Hence, people who are born with innate predispositions that make them happy or come from the same social backgrounds may tend to marry each other. One reason for this is that the decision to marry somebody who is like us could make living with them easier, as the latter may enjoy the same kind of lifestyle, such as leisure and sporting activities, whilst someone else with a completely different set of personalities or have a different social background may not. Such positive assortative mating or homogenous matching by personality traits and/or

social factors is supported by the evidence that a number of lifestyles are highly correlated within a couple (Contoyannis and Jones, 2004). It is also consistent with the evidence of positive assortative mating by education (Mare, 1991), professional backgrounds (Qian, 1998), productivity traits and desires for public goods (Lam, 1998). An assortative mating market may thus induce correlated effects in LS via correlation between respondent individual fixed effects and partner's LS.

Alternatively, correlated effects may result from partners sharing the same social environment, as discussed below.

3.2 Shared social environment

The second explanation also views the observed correlation as a correlated effect; what appears to be a direct spillover of LS from one partner to the other may be no more than the result of partners sharing the same social environment that simultaneously influences the well-being of them both. For example, a positive shock in one partner's income (i.e. one partner receiving a pay rise at work) can result in an increase in both spouses' LS through an increase in the family income. Moreover, under assortative mating, a couple with common lifestyles characteristics may also experience common life cycle events. Two people in the same occupation may be attracted to each other and as a result, they may experience the same cyclic shocks to income. Likewise, couples may share the same health habits and given their similarity in age and habits, they may also experience health shocks that are close in timing. As a result, the observed spousal similarity in LS could thus be a spurious relationship stemming from the fact that some life events either occur to both spouses simultaneously or occur to one partner but due to the nature of partnership affect both spouses simultaneously.

3.3 Spillover

The final interpretation of the raw spousal correlation in LS views the association as a result of a spillover effect of LS from one partner to the other (Becker, 1974; Friedman, 1986). Transmission of LS is assumed to occur between closely related partners who identify with and care for each other and share a great part of their lives together. In the spillover effect model, the LS of one partner acts as an externality for the other partner, which in turn increases the current level and influences future growth in LS for the latter (Larson and Almeida, 1999; Westman and Vinokur, 1998). Note, however, that there may be other kinds of interactions between partners' LS if there is no caring between partners. For instance, if something affects one spouse's LS positively and it enhances the desirability for the person to re-enter the marriage market (e.g. he or she wins at the lottery), this may affect the other negatively. In those cases, an external positive shock to one partner's LS may have a negative impact on the other's LS.

The idea of a spillover effect of LS within marriage is consistent with many studies that have found a positive relationship between self-rated well-being and altruistic or caring behaviour. For instance, Benson et al. (1980) found a positive correlation between LS and time spent in a variety of helping activities. Using panel data, Thoits and Hewitt (2001) found volunteer work leads to greater happiness, LS, self-esteem, and even physical health for the individual. Konow and Earley (2008) showed through various laboratory experiments that giving and helping others and many other selfless acts can raise and sustain happiness at a higher than average level compared with other goals such as the pursuit of material wealth. Frey et al. (2004) show how LS may often depend more on the processes (i.e. from helping others) than on the returns or outcomes of the actions. Following the collapse of the infrastructure of volunteering work in the German Democratic Republic in the late 1980s,

Meier and Stutzer (2008) studied the causal impact of loss of volunteer work on happiness. They found that a drop from volunteering monthly to less than monthly reduced LS by more than 0.2 point in an 11-point-scale (one-half of the effect of separation from partner). More closely related to this paper, in the German Panel data Schwarze and Winkelmann (2003) and Bruhin and Winkelmann (2007) found some evidence of altruism. As well as showing that predicted altruists are more likely to make transfer payments, they were able to demonstrate that an exogenous increase in children's LS can lead to an increase in LS for the parents.

4. Implementing a test

4.1 The utility model of couples

Consider first Gary Becker's (1974) simple utility function of an individual i in a marriage to individual j at any given time, which can be written as

$$U_i = U(X, U_j(X)), \tag{1}$$

where X is a vector of the consumption of commodities within the household, and U_j is the individual i 's partner's utility. The individual's utility is assumed to be increasing with X , which is divisible and can be shared between the couple. An increase in X therefore raises the individual's utility both through a direct effect upon $U_i(X)$ and an indirect effect, acting through a rise in the partner's utility, $U_j(X)$. The utility function is highly stylized and abstracts from a number of issues that can be expected to be important, i.e. I assume no other relations-specific investments in the household.

For a partnership between individual i and j , the empirical counterpart to equation (1), with application to life satisfaction, is of the following form:

$$LS_{it} = \alpha_{0t} + \alpha_1 LS_{jt} + \alpha_2 X_t + \lambda_i + \varepsilon_{it}, \quad (2)$$

$$LS_{jt} = \beta_{0j} + \beta_1 LS_{it} + \beta_2 X_t + \lambda_j + \varepsilon_{jt}, \quad (3)$$

where $t = 1, 2, \dots, T$. The parameters λ_i and λ_j represent individual i and j 's unobserved individual fixed effects. The error term, ε , is thought to subsume the inability of human beings to communicate accurately true satisfaction levels. Equation (2) and (3) imply that U_{it} and U_{jt} are jointly determined, i.e. endogenous. The parameters α_1 and β_1 can be defined such that the higher the value of α_1 (β_1) the more satisfaction individual i (j) can derive from the same increase in his or her partner's life satisfaction. The reduced form of (2) and (3) are given by:

$$LS_{it} = \Pi_1 + \Pi_2 X_t + v_{it}, \quad (4)$$

$$LS_{jt} = \Lambda_1 + \Lambda_2 X_t + v_{jt}, \quad (5)$$

where the parameters are defined as:

$$\Pi_1 = \frac{\alpha_0 + \alpha_1 \beta_0}{1 - \alpha_1 \beta_1}, \Pi_2 = \frac{\alpha_1 \beta_2 + \alpha_2}{1 - \alpha_1 \beta_1}, \Lambda_1 = \frac{\beta_0 + \beta_1 \alpha_0}{1 - \alpha_1 \beta_1}, \Lambda_2 = \frac{\beta_1 \alpha_2 + \beta_2}{1 - \alpha_1 \beta_1} \quad (6)$$

with the error terms: $v_{it} = \frac{\alpha_1(\lambda_j + \varepsilon_{jt}) + \lambda_i + \varepsilon_{it}}{1 - \alpha_1 \beta_1}$ and $v_{jt} = \frac{\beta_1(\lambda_i + \varepsilon_{it}) + \lambda_j + \varepsilon_{jt}}{1 - \alpha_1 \beta_1}$.

Therefore, unless it can be shown that $Cov(LS_{jt}, \lambda_j)$ and $Cov(LS_{jt}, \varepsilon_{it})$ are equal to zero, OLS on (2) will be inconsistent. The same applies to equation (3). Owing to the correlated effects

mentioned in the previous section, it is likely however that the partner's LS will be correlated with both the respondent's unobserved fixed effects and the time-varying error term.

4.2 A sketch of identification

The correlated effects due to assortative mating, i.e. $Cov(LS_{jt}, \lambda_j) \neq 0$, can be easily solved by applying the first-differencing approach on the LS data. By contrast, the correlated effects due to shared social environmental can be solved if there is a valid instrument that affects the respondent's partner's LS directly but is not correlated with the respondent's LS beyond its impact on the endogenous regressor.

One such instrument for partner LS may be its own lagged variables. To illustrate, we can incorporate the lagged variables of the respondent's own LS in each equation in the system. The inclusion of lagged LS is theoretically plausible as, according Graham and Oswald (2006), past happiness can be thought of as a "hedonic capital" or investment for future happiness. Assuming, for now, that the individual's past LS only correlates with the individual's current LS but is not correlated with his partner's LS, we can rewrite equations (2) and (3) as:

$$LS_{it} = \alpha_{0t} + \alpha_1 LS_{jt} + \alpha_2 X_t + \alpha_3 LS_{it-1} + \lambda_i + \varepsilon_{it}, \quad (2')$$

$$LS_{jt} = \beta_{0j} + \beta_1 LS_{it} + \beta_2 X_t + \beta_3 LS_{jt-1} + \lambda_j + \varepsilon_{jt}, \quad (3')$$

where the individual's past LS is assumed to be correlated only with the individual's current realisation of LS and not his partner's LS. It then follows that the reduced form of (2') and (3'), after first-differencing to eliminate the time-invariant parameter λ , are given by:

$$\Delta LS_{it} = \Pi_1 + \Pi_2 X_t + \Pi_3 \Delta LS_{it-1} + \Pi_4 \Delta LS_{jt-1} + v_{it}, \quad (4')$$

$$\Delta LS_{jt} = \Lambda_1 + \Lambda_2 X_t + \Lambda_3 \Delta LS_{it-1} + \Lambda_4 \Delta LS_{jt-1} + v_{jt}, \quad (5')$$

where $\Pi_3 = \frac{\alpha_1 \beta_3}{1 - \alpha_1 \beta_1}$, $\Pi_4 = \frac{\alpha_3}{1 - \alpha_1 \beta_1}$, $\Lambda_3 = \frac{\beta_3}{1 - \alpha_1 \beta_1}$, $\Lambda_4 = \frac{\beta_1 \alpha_3}{1 - \alpha_1 \beta_1}$, with the other

parameters have remained the same as defined in equation (6) and the error terms are now

defined as: $v_{it} = \frac{\alpha_1 \varepsilon_{jt} + \varepsilon_{it}}{1 - \alpha_1 \beta_1}$ and $v_{jt} = \frac{\beta_1 \varepsilon_{it} + \varepsilon_{jt}}{1 - \alpha_1 \beta_1}$. We can now solve for α_1 and β_1 from the

structural equation that will give the true effects of partner LS on the respondent's LS

$$\frac{\Pi_3}{\Lambda_3} = \frac{\alpha_1 \beta_3}{1 - \alpha_1 \beta_1} \bullet \frac{1 - \alpha_1 \beta_1}{\beta_3} = \alpha_1, \quad (7)$$

$$\frac{\Pi_4}{\Lambda_4} = \frac{\beta_1 \alpha_3}{1 - \alpha_1 \beta_1} \bullet \frac{1 - \alpha_1 \beta_1}{\alpha_3} = \beta_1. \quad (8)$$

In addition to this, provided that the lagged dependent variable is a valid instrument, i.e., it does not correlate with the measurement error and the equation error (or, for individual i , $Cov(LS_{jt-1}, \varepsilon_{it}) = 0$ and $Cov(LS_{jt-1}, LS_{jt}) \neq 0$), the IV method will also correct for the negative bias associated with the misreporting of LS by removing the variance-in-noise and only use variance-in-signal to estimate the true well-being effects from one partner to the other.

4.3 A dynamic GMM model of life satisfaction

The solutions obtained from equations (7) and (8) would be valid only if each partner's lagged LS is a valid instrument, i.e. $Cov(LS_{jt-1}, \varepsilon_{it}) = 0$, and vice versa. However, this may still be a

strong assumption considering that partner's lagged LS may have an inertial effect on the respondent's behaviour today, which may affect his or her current LS. In addition to this, the presence of the lagged dependent variable in the respondent's LS equation may also give rise to a new endogeneity problem; that is, the autocorrelation issue or, as expressed in the first-differenced model, the correlation between the new error term, i.e. $\varepsilon_{it} - \varepsilon_{it-1}$, and the differenced lagged dependent variable, i.e. $LS_{it-1} - LS_{it-2}$.

Furthermore, the observed correlation may also be biased owing to the presence of measurement error associated with how each partner's LS is collected. Some people may have a systematic tendency to misreport their true LS scores in surveys. People with personalities that make them happy may always overrate their true LS when asked, for example. In addition to this, because each partner is asked to report their LS to the surveyor in confidence, the signal-to-noise ratio of his or her LS when used as an explanatory variable in the other partner's LS equation is also likely to be small. The presence of measurement error bias can therefore cause the estimated effect of partner LS on the respondent's LS to be biased towards zero in a large sample (see Angrist and Krueger, 2001).

In order to control for the correlated effects described in the previous section and solve the problem of measurement error bias described above, this paper's econometric methodology follows the system generalized methods of moments (GMM) estimator outlined by Arellano and Bover (1995), Blundell and Bond (1998), and more recently in an applied paper by Fajnzylber et al. (2002). Consider the following dynamic model with unobserved individual fixed effects:

$$LS_t^* = \alpha_0 + \alpha_1 LS_t^{p*} + \alpha_2 X_t + \alpha_3 LS_{t-1}^* + \lambda + \varepsilon_t, \quad (9)$$

where the terms, LS^* and LS^{p*} , denote self-rated life satisfaction of the respondent and the respondent's partner, respectively. Assuming that there are systematic errors to how each partner LS is measured and that this measurement error is being driven by (a) a fixed effect and (b) a time-varying component:

$$LS_t = LS_t^* + \xi + \omega_t, \quad (10)$$

$$LS_t^p = LS_t^{p*} + \eta + \phi_t. \quad (11)$$

Substituting (10) and (11) into (9)

$$LS_t = \alpha_0 + \alpha_1 LS_t^p + \alpha_2 X_t + \alpha_3 LS_{t-1} + [\lambda - \alpha_1 \eta + (1 - \alpha_3) \xi] + [\varepsilon_t - \alpha_1 \phi_t + \omega_t - \alpha_3 \omega_{t-1}],$$

$$LS_t = \alpha_0 + \alpha_1 LS_t^p + \alpha_2 X_t + \alpha_3 LS_{t-1} + \mu + v_t, \quad (12)$$

Therefore the measurement error of each partner's LS is subsumed in the unobserved individual fixed effects and the error term in our model. Because there are two components to the bias, i.e. assortative mating (+ve) and measurement error (-ve), the overall direction of bias coming from unobserved individual fixed effects is unclear on *a priori* ground.

To estimate the dynamic model of life satisfaction with unobserved individual fixed effects, I use the system GMM estimator mentioned above. The special feature of this dynamic GMM estimator is that it joins in a single system the regression equation in differences and in levels, each with its specific set of instrumental variables, thereby allowing both types of endogeneity to be controlled for.

In the difference equation, the unobserved individual fixed effects, μ , are eliminated from the model through first-differencing:

$$\begin{aligned}
LS_t - LS_{t-1} &= (\alpha_{0t} - \alpha_{0t-1}) + \alpha_1(LS_t^p - LS_{t-1}^p) + \alpha_2(X_t - X_{t-1}) \\
&+ \alpha_3(LS_{t-1} - LS_{t-2}) + (\varepsilon_t - \varepsilon_{t-1}).
\end{aligned} \tag{13}$$

The use of instruments is required to deal with two issues. First, the endogeneity of the explanatory variables – and in particular, the partner’s LS – which is reflected in the correlation between these variables and the error term. The second is the correlation, by construction, between the new error term, $(\varepsilon_t - \varepsilon_{t-1})$, and the differenced lagged dependent variable, $(LS_{t-1} - LS_{t-2})$. Assuming, for now, that (a) the error term is not serially correlated, and (b) the explanatory variables are weakly exogenous, i.e. they may be affected by past and current realisations of the dependent variable (the respondent’s LS) but not by its future innovations, the following moment conditions apply:

$$E[LS_{t-s}(\varepsilon_t - \varepsilon_{t-1})] = 0 \quad \text{for } s \geq 2, \quad t = 3, \dots, T, \tag{14}$$

$$E[LS_{t-s}^p(\varepsilon_{it} - \varepsilon_{it-1})] = 0 \quad \text{for } s \geq 2, \quad t = 3, \dots, T, \tag{15}$$

$$E[X_{t-s}(\varepsilon_t - \varepsilon_{t-1})] = 0 \quad \text{for } s \geq 2, \quad t = 3, \dots, T, \tag{16}$$

However, given that the partner’s LS is an endogenous variable, the natural candidate instrument in the above equation would instead be the first-difference of the second lag of partner LS rather than the first-difference of the first lag. This is simply because the second lag is not correlated with the error term, while the first lag is. The same applies to other explanatory variables which are endogenous to current realisations of the respondent’s LS. In addition to this, given our lagged dependent variable model, the presence of time-varying measurement error would make that the regression error follow a moving average of order 1. In other words, as will be explained later in the results section, the error term will be serially

correlated and, thus, we could not use the most recent lags of the dependent variable as instruments. Because the BHPS has $T \geq 3$, the use of higher-order lagged dependent variable as instruments in the GMM-system estimation is considered feasible. It will also, following Arellano's (2003, p.51) advice, help to solve the problem of time-varying measurement error bias.

To improve the model's efficiency, a second *level* equation is augmented to the first *difference* equation. One reason for this is because sometimes the lagged dependent variables are poor instruments for the first-differenced regressors, and by adding the second equation additional instruments can be obtained. For the regression in levels, the unobserved individual fixed effects are not eliminated but must be controlled for by the use of instrumental variables. Assuming that there is no correlation between the *differences* of the explanatory variables and the unobserved individual fixed effects, their values can be instrumented with their own first differences. This assumption results from the following stationarity property:

$$E[LS_{t+l}\lambda_i] = E[LS_{it+m}\lambda_i], \quad (17)$$

$$E[LS_{t+l}^p\lambda_i] = E[LS_{it+m}^p\lambda_i], \quad (18)$$

$$E[X_{t+l}\lambda_i] = E[X_{it+m}\lambda_i], \quad (19)$$

for all l and m . Thus, the additional moment conditions for the second part of the system are given by:

$$E[(LS_{t-s} - LS_{t-s-1})(\mu + \varepsilon_t)] = 0 \quad \text{for } s = 1, \quad (20)$$

$$E[(LS_{t-s}^p - LS_{t-s-1}^p)(\mu + \varepsilon_t)] = 0 \quad \text{for } s = 1, \quad (21)$$

$$E[(X_{t-s} - X_{t-s-1})(\mu + \varepsilon_t)] = 0 \quad \text{for } s = 1. \quad (22)$$

Using the moment conditions (14-16) for the difference equation and (20-22) for the level equation, I employ the system GMM estimator to generate consistent estimates for partner's LS on the respondent's LS and their asymptotic variance-covariance (Arellano and Bond , 1991; Arellano and Bover, 1995; Blundell and Bond, 1998). Because the impact of partner LS on own LS may be different for men and for women, I conduct all statistical analysis by gender.

A crucial assumption for the validity of GMM is that the instruments are exogenous. There are two specification tests that we can use to check whether the lagged values of the explanatory variables are valid instruments in the LS regression equation. The first is the Hansen test of overidentification (Hansen, 1982), which tests the null hypothesis of overall validity of the instruments by analysing the sample analogue of the moment conditions used in the estimation process. Failure to reject the null hypothesis gives support to the model. Note that the Hansen test is robust compared to Sargan test of overidentification but can be weakened by having too many instruments, i.e. more instruments than the number of individuals used in the estimation (Roodman, 2007). The second test examines the hypothesis that the error term ε_{it} is not serially correlated. This is done by testing whether the differenced error term (that is, the residual of the regression in differences) is first- or second-order serially correlated. First-order serial correlation of the differenced error term is expected even if the original error term (in levels) is uncorrelated, unless the latter follows random walk. Second-order serial correlation of the differenced error term specifies that the original error term is serially correlated and that the instruments are misspecified. By contrast, if we can accept the null hypothesis of no second-order serial correlation then we can conclude that the original error term is serially uncorrelated and the moment conditions are well-specified.

4.4 Data

The present investigation uses data from the BHPS. This is a nationally representative sample of persons aged 16 and over in 1991, who have been interviewed every subsequent year. The study interviewed separately all adult members of the household with respect to their income, employment status, marital status, health, and attitudes. There is both entry into and exit from the panel, leading to unbalanced data with an increasing number of individual interviews over time. This is due to the inclusion of children from the original household sample who turn 16, of refresher samples, and of new members of households formed by original panel members.

As well as questions on socio-economic status, individuals were also asked from Wave 6 onwards to indicate how satisfied they are with their life, from 1 (*very dissatisfied with life*) to 7 (*very satisfied with life*). The LS question is located in a self-completed section of the survey, which is strategically placed at the end of the questionnaire after individuals had been asked about their household and individual characteristics.

I consider all married and cohabiting individuals (including same-sex partnership) observed consecutively over two periods with information on own and partner lagged LS for the years 1996–2007 (Waves 6-16). Couples who remained with the same partner are treated the same way in the analysis as those who changed partners during the observed panel. Note that Wave 11 is omitted from the analysis because of the omission of LS questions in that survey year. The unbalanced panel with non-missing information on both on and partner LS (not including lags) includes 38,161 female observations (7,468 individuals) and 38,311 male observations (7,464 individuals). Of these observations, 6,493 of female observations, and 6,644 of male observations, are currently cohabiting. The average age for women is 46 and 48 for men. Around 51% of women and 59% of men are in full-time employment. Approximately 70% of households have at least one child under the age of 16 in the household.

4.5 Accounting for selection bias

Although I am interested only in live-in couples, there is likely some selection bias involved in moving from other marital statuses (e.g. single, divorced, widowed, or separated) to being in a relationship. One could imagine, for instance, that married or cohabiting individuals present specific characteristics that influence the way they are affected by their partner's well-being compared to other non-married persons.

To correct for any selection bias in moving from non-married status to being in a couple, I compute an inverse Mills ratio using a selection variable that equals 1 if the person is either married or cohabiting and 0 otherwise. This couple equation is estimated on the whole BHPS sample, as shown in Table A1 in the appendix, as a function of gender, age, age-squared, number of young (aged 11 and under) and old children ($12 \leq \text{age} < 18$) in the household, subjective health (5 dummies), log of real household income per capita, education (2 dummies), employment status (3 dummies), wave dummies and the divorce rate of other households within the region. This last variable is used to satisfy the exclusion restrictions, which is possible as the divorce rate of other households within the region is correlated with whether or not the respondent is married or cohabiting, but should not be correlated with how each partner is affected by the other partner's LS.

5. Results

5.1 Life satisfaction spillovers

Table 2 reports results from the GMM-system estimator described in the previous section. The dependent variable is the respondent's self-rated life satisfaction measured cardinally (on a

scale of 1 to 7)¹. For comparative purposes, I also present results from the OLS estimator, as well as the GMM-levels estimator, which does not control for individual fixed effects. In contrast to the GMM-system, first-order serial correlation is a sign of misspecification in the case of levels estimator. To prevent the case of having too many instruments that would lead to an overfitting of the endogenous variables, I follow Roodman's (2007) advice and "collapse" instruments in all GMM regressions. Finally, it should also be noted that, in cases where the lagged dependent variable is included as a right-hand side variable in the regression, each estimated coefficient represents the short-run effect of the respective variable.

The first two columns of Table 2 report the OLS results for men and women, respectively. Here, only partner LS, inverse Mills ratio, and the exogenous variables – namely age, age-squared, and wave dummies – are included as the right-hand side variables in the LS equation. As anticipated, partner LS enters the respondent LS equations in a positive and statistically significant manner, i.e. there seems to be strong evidence that partners' LS are positively correlated at cross-section. The relationship appears to be moderate in size; an increase of one life satisfaction score in partner LS is associated with approximately 25 percentage-points increase in respondent LS for men and 27 percentage-points for women. However, as described in the previous section, this observed correlation between partners' well-being is likely to be biased owing to either or both correlated effects and measurement errors in the LS variables.

Treating partner LS as endogenous – i.e. using the second- rather than the first-lag of partner LS as instruments, Columns 3 and 4 of Table 2 re-estimate the first two columns' specification using the GMM-levels estimator. GMM-levels yields a positive coefficient on partner LS for both men and women, although the relationship is statistically significant only

¹ One objection is that measures of subjective well-being, such as happiness and life satisfaction, are ordinal rather than cardinal variables. However, many research papers have shown that it virtually makes no difference in the estimation whether one assumes ordinality or cardinality in the reported well-being data (see, e.g., Ferrer-i-Carbonell and Frijters, 2004).

in the male sub-sample regression; the coefficients on partner LS are 0.561 (*S.E.* = 0.205) for men and 0.228 (*S.E.* = 0.188) for women. However, the model performs poorly; this specification is rejected by the test of first-order serial correlation, which indicates that the model is misspecified in the case of the levels estimator.

Columns 5 and 6 of Table 2 introduce a lag one of the dependent variable into the model. Using the most recent lags of the dependent variable as instruments, the lagged respondent LS variable is positive and statistically significant at the 1% level for both men and women, which is consistent with the hedonic capital theory, i.e. our happiness today is determined by how happy we were in the past (Graham and Oswald, 2006). Yet the inclusion of the lagged dependent variable has now made the partner LS variable insignificant in both male and female sub-sample regressions. Furthermore, this specification is also rejected by the tests of serial correlation for both men and women, as well as the Hansen test of overidentification in the case of male sub-sample regression. In other words, we may still have (a) omitted variables that are correlated with both respondent and partner LS, which can either be time-varying or relatively stable over-time, and/or (b) serial correlations in the error term caused by the presence of time-varying measurement error bias.

To solve the problem of unobserved heterogeneity, the next two columns use the GMM-system estimator to control for unobserved individual fixed effects that are potentially correlated with the explanatory variables. As described in the previous section, the presence of unobserved individual fixed effects can bias the estimated coefficients on partner LS through assortative mating and shared social environment with high over-time persistence (+ve) and measurement errors (-ve), although the extent of each bias on the estimated correlation between partners' LS is unknown on *a priori* ground.

The GMM-system estimator generates coefficients on partner LS that are notably larger than the ones obtained by the GMM-levels estimator; a one-satisfaction-point increase

in partner LS is now associated with approximately 0.58 increase in the respondent's LS for men and 0.75 for women. What this implies is that, in regressions where unobserved individual effects are ignored, the negative bias stemming from the relatively stable component of the measurement error is larger compared to the positive bias coming from assortative mating and shared social environment with high over-time persistence. Nevertheless, the estimates are still poorly defined, considering that the specification is rejected by both the Hansen test of overidentification and the tests of serial correlation. As mentioned earlier, a rejection of the second-order serial correlation in the case of the GMM-system estimator implies that the original error term is serial correlated. This serial correlation is expected, however, as we have yet to correct for the measurement error term that varies over time.

To solve the problem of time-varying measurement error bias, Columns 9 and 10 drop the most recent lags of the dependent variable from the instrument list and use only the second-lag and beyond as instruments in the GMM-system estimation. The specification is now supported by the tests of serial correlation, i.e. we can now accept the null hypothesis of no second-order serial correlation with the use of higher-order lags of the dependent variable as instruments. This is the case for both men and women. However, this specification is still rejected in the female sub-sample regression by the Hansen test of overidentification. In addition to this, the process also increases the size of the estimated coefficients on the lagged dependent variable; the estimated coefficients on respondent lagged LS are 0.13 and 0.09 for men and women, respectively. This supports the way we modelled the measurement error term as having a time-varying component as well as a fixed effect. However, this specification also produces large standard errors for the estimated lagged dependent variable. By contrast, the

The last two columns (Columns 11 and 12) control for a number of household and individual characteristics known in the literature to be important LS predictors (see, e.g.,

Oswald and Powdthavee, 2008), adding log of real household income per capita, employment status, a dummy of whether the respondent is cohabiting, number of years stayed together as a couple in the BHPS (maximum of 16 years), number of children, education, and subjective health status, all of which are treated as endogenous variables in the GMM-system estimation. The estimated coefficients on these control variables are presented in Table A2 in the appendix.

In this specification, we can see that for both men and women the p -values for the Hansen test are now, according to Roodman (2007), sufficiently high and within the recommended range ($0.1 \leq p < 0.25$). This confirms that important time-varying components has been omitted in the previous specifications and by controlling for them in the estimation, the models are now well-specified. With a full set of controls, the point estimates on the partner LS variable are reduced to 0.278 ($S.E. = 0.047$) for men and 0.329 ($S.E. = 0.057$) for women, which are not that much different to the point estimates obtained in the OLS estimation (see Columns 1 and 2). What this implies is that there is almost a full offsetting effect between omitted variables (i.e. unobserved individual fixed effects, time-varying variables, and dynamic effects coming from the lagged dependent variable) and the measurement error bias on the estimated correlation between partners' LS at cross-section.

With respect to Table 2's other results, there is a U-shaped relationship between age and LS for both partners, minimizing at around early 40s. The negative and statistically significant inverse Mills ratio indicates that the selection into partnership and LS is correlated. Interestingly, household income is negatively correlated with LS in the GMM-system estimation. Better health is found to be highly correlated with higher LS levels for both genders. Nonetheless, while almost all of the other estimated coefficients have the correct signs, they are also estimated with large standard errors.

5.2 Termination of partnership

Up to this point, this paper has concentrated on the estimation of causal effect of partner LS on respondent LS. Such an approach seems to be of some worth in its own right. However, in order to make further justification on the importance of the previous spillover effect model presented in the last section, I now estimate a marital dissolution equation.

Of the married and cohabiting individuals, there were approximately 621 couples (roughly 1% of the sample) who moved from being married or cohabiting with a partner at period t to separation or divorce at $t+1$. The key hypothesis to be tested here is that there is a short-run association between partners' LS and their decision to stay together. More specifically, couples with higher LS levels at t are less likely to be separated or divorced at $t+1$.

Table 3 presents marginal effects (reported in percentages) obtained from a probit model on whether the couple terminates their relationship at period $t+1$. The model controls for a number of household and personal characteristics, including age, age-squared, cohabiting dummy, income, employment status, education, health, number of years stayed as a couple in the panel, selection into a couple (i.e. inverse Mills ratio), number of children, regional and wave dummies. It also includes as the variables of interest (a) LS measured at t and (b) average LS over time for male and female partners in a couple. Following one of the referees' advice, I also introduce variables for the interaction between partners' LS at t and the same interaction for the average LS levels. The probit is estimated on the male sub-sample, although the female sub-sample regression produces qualitatively similar results.

Controlling for household and personal characteristics, as well as the average LS levels of both partners, we can see that both partners' LS at t are associated negatively and statistically significantly with separation and divorce at $t+1$. A unit increase in either partner

LS at t is associated with approximately 0.3 percentage-point decrease in the probability of separation or divorce at $t+1$. This is a large effect, considering that only 1% of the sample made a transition from being married/cohabiting at t to being separated or divorced at $t+1$. Interestingly, the interaction term between partners' LS at t is positive and statistically significant at the 10% level; the negative effect of each partner's LS on the couple's propensity to separate or divorce is smaller in the case when both partners' LS at t are high compared to when they are both low.

By contrast, the mean values of both partners' LS are positively correlated with the probability of separation or divorce at $t+1$. This is somewhat consistent with the notion that there may be important peak-end effects in the marriage market, whereby the decision to separate/divorce depends positively on past variation in LS, i.e. the difference between current LS and some reference level that would be proxied by the averaged LS levels (see, e.g., Kahneman, 1999; Clark and Georgellis, 2004). Furthermore, the peak-end theory is supported by the negative interaction effect between the average LS levels, which is statistically significant at the 5% level.

In short, the link between partners' decision to end partnership at $t+1$ and their LS at t is consistent with the spillover effect model, whereby changes in either partner's LS today influence the decision of whether or not to stay together as a couple in the next period.

Finally, Table 3's other results are also interesting in their own right. Income this period is associated negatively with the probability that the couple will split one period into the future, whilst people who are self-employed at t are more likely than those in full-time employment to end their relationship at $t+1$. Interestingly, being a housewife (or look after home) this year lowers the probability of couple dissolution in the subsequent year by approximately 3 percentage-points. This is perhaps due to the fact that at least one partner is

always home when the other partner is back from work, something which can be considered as good for the relationship.

6. Conclusion

This paper has used ten waves of BHPS data to study intra-spousal correlations in self-reported life satisfaction data. Its primary objective was to determine whether the observed correlation is due largely to partners' fixed traits are similar through assortative mating by personality traits on the marriage market, partners sharing the same social environment that simultaneously affects their well-being, or a spillover effect of life satisfaction from one partner to the other.

A simple OLS model reveals that there is indeed a positive and statistically significant correlation between partners' LS. This correlation persists even when the GMM-system estimator is used to estimate the model, i.e. partners' LS remain positively and statistically significantly correlated even after controlling for the presence of measurement error bias and unobserved heterogeneity and allowing partner LS to be instrumented by the first-differences of its lags. In other words, there seems to be strong evidence of a causal spillover effect of well-being that runs from one partner to the other in a couple. This conclusion is also supported by the evidence that partners' LS can be used to predict observable behaviours: there is a negative and statistically significant association between partners' life satisfaction this year and the likelihood of couple dissolution in the subsequent year. These results are consistent with models of spillover effects within couples. The findings thus provide strong statistical support in terms of validity for many economic models that were built around the assumption that utility is interdependently related between members of the same household

(Becker, 1974). It is also consistent with studies that found evidence of caring preferences between partners within marriage (Foster and Rosenzweig, 2001; Ermisch, 2003).

More generally, the empirical approach of this paper can be extended and applied to distinguish between various explanations of other types of similarity in couples' behaviours and characteristics that are not specific to a partner's subjective well-being.

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Table 1: Descriptive Statistics of the Variables used in the GMM estimation, BHPS 1996-2006

Variables	Men	Women
Life satisfaction	5.300 (1.196)	5.347 (1.250)
Partner Life satisfaction	5.347 (1.249)	5.299 (1.197)
Age	48.039 (15.422)	45.695 (14.826)
Age-squared/100	25.456 (15.855)	23.161 (14.826)
Inverse Mills ratio	0.377 (0.281)	0.483 (0.326)
Ln(real household income per capita)	9.570 (0.628)	9.569 (0.628)
Self-employed	0.127 (0.333)	0.042 (0.201)
Unemployed	0.031 (0.174)	0.017 (0.130)
Retired	0.195 (0.396)	0.170 (0.376)
Maternity leave	0.000 (0.011)	0.011 (0.107)
Family care	0.005 (0.074)	0.159 (0.366)
Full-time student	0.004 (0.067)	0.008 (0.093)
Disabled, long-term sickness	0.044 (0.206)	0.035 (0.185)
Government training scheme	0.000 (0.022)	0.000 (0.016)
Other type of employment	0.003 (0.056)	0.004 (0.065)
Number of years as a couple in the BHPS	11.381 (4.660)	11.430 (4.636)
Number of years as a couple-squared	151.269 (99.379)	152.156 (99.406)
Cohabit	0.173 (0.378)	0.170 (0.375)
Number of children (aged<16)	0.714 (1.036)	0.721 (1.040)
Completed first degree	0.105 (0.307)	0.102 (0.303)
Completed higher degree	0.033 (0.179)	0.022 (0.147)
Health: poor	0.073 (0.261)	0.086 (0.281)
Health: fair	0.212 (0.409)	0.224 (0.417)
Health: good	0.452 (0.497)	0.453 (0.497)
Health: excellent	0.241 (0.427)	0.211 (0.408)
N	38135	37985

Note: Standard deviations are in parentheses.

Table 2: OLS and Dynamic GMM estimates of life satisfaction, BHPS 1996-2007

	OLS		GMM-levels		GMM-levels, Instrument from first-lag dependent variable		GMM-system, Instrument from first-lag dependent variable		GMM-system, Instrument from second-lag dependent variable		GMM-system with full controls and instrument from second-lag dependent variable	
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Dependent variable: Own LS at t												
Partner LS at t	0.245	0.270	0.561	0.228	0.069	0.164	0.575	0.749	0.543	0.765	0.278	0.329
	[0.008]**	[0.009]**	[0.205]**	[0.188]	[0.136]	[0.113]	[0.094]**	[0.117]**	[0.116]**	[0.129]**	[0.047]**	[0.057]**
Own lagged LS (t-1)					0.668	0.448	0.064	0.064	0.130	0.093	0.018	-0.038
					[0.098]**	[0.115]**	[0.017]**	[0.017]**	[0.121]	[0.121]	[0.045]	[0.049]
Age	-0.067	-0.055	-0.044	-0.074	-0.022	-0.043	-0.026	-0.024	-0.025	-0.022	-0.039	-0.035
	[0.006]**	[0.006]**	[0.027]	[0.020]**	[0.017]	[0.015]**	[0.012]*	[0.012]+	[0.012]*	[0.013]+	[0.017]*	[0.019]+
Age-squared/100	0.073	0.062	0.048	0.083	0.024	0.047	0.030	0.026	0.029	0.024	0.047	0.044
	[0.006]**	[0.007]**	[0.029]+	[0.022]**	[0.018]	[0.016]**	[0.012]*	[0.013]*	[0.012]*	[0.013]+	[0.017]**	[0.019]*
Inverse Mills ratio	-0.504	-0.336	-0.340	-0.526	-0.216	-0.316	-0.202	-0.116	-0.211	-0.116	-0.456	-0.379
	[0.054]**	[0.045]**	[0.263]	[0.171]**	[0.165]	[0.130]*	[0.129]	[0.119]	[0.126]+	[0.120]	[0.176]**	[0.200]+
Constant	5.541	5.121	3.326	5.835	1.941	3.094	2.468	1.554	2.292	1.295	3.595	5.731
	[0.173]**	[0.171]**	[1.670]*	[1.355]**	[0.999]+	[0.950]**	[0.635]**	[0.718]*	[0.666]**	[0.805]	[1.023]**	[1.530]**
Observations	38135	37985	38135	37985	27520	27659	27520	27659	27520	27659	27518	27655
Number of individuals			7438	7442	6432	6494	6432	6494	6432	6494	6432	6494
Number of instruments			34	34	41	41	47	47	46	46	357	358
Wave dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
- Hansen test of overidentification			0.243	0.936	0.036	0.271	0.153	0.028	0.190	0.040	0.235	0.144
- First order serial correlation			0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
- Second order serial correlation			0.288	0.503	0.000	0.001	0.006	0.014	0.186	0.243	0.299	0.792

Note: + < 10%; * < 5%; ** < 1%. Standard errors are in parentheses. Life satisfaction is recorded on a 7-point-scale, with 1 = very dissatisfied and 7 = very satisfied. All explanatory variables whenever they are appeared in the model – except for age, age-squared, and year dummies – are treated as endogenous variables, i.e. used as GMM-type instruments. Age, age-squared, and year dummies are treated as exogenous variables orthogonal to the individual fixed effects, i.e. their levels are used as instruments the level equations. Control variables in the last two columns include log of real household income per capita, employment status, a dummy of whether the respondent is cohabiting, number of years stayed together as a couple in the panel, number of children, education, and subjective health status, all of which are treated as endogenous variables in the GMM-system estimation.

**Table 3: Marginal Effects Probit Model of Termination of
Partnership at Period t+1**

Dependent variable: separation or divorce at t+1	Male sub-sample
Own LS at t	-0.00281 [0.00085]**
Partner's LS at t	-0.00330 [0.00081]**
Own LS at t*Partner's LS at t	0.00031 [0.00016]+
Average own LS over time	0.00256 [0.00163]
Average partner's LS over time	0.00358 [0.00160]*
Average own LS over time*Average partner's LS over time	-0.00074 [0.00034]*
Ln(real household income per capita)	-0.00181 [0.00074]*
Self-employed	0.00292 [0.00122]*
Unemployed	0.00219 [0.00204]
Retired	-0.00054 [0.00178]
Family care	-0.00302 [0.00153]*
Disabled, long-term sickness	0.00042 [0.00188]
Other type of employment	0.00411 [0.00714]
Number of years as a couple in the BHPS	0.00014 [0.00059]
Number of years as a couple-squared	0.00000 [0.000037]
Age	-0.00006 [0.00030]
Age-squared/100	-0.00017 [0.00032]
Cohabit	0.00086 [0.00112]
Number of children	-0.00002 [0.00047]
Completed first degree	-0.00078 [0.00098]
Completed higher degree	-0.00043 [0.00189]
Health: poor	0.00096 [0.00235]
Health: fair	0.00115 [0.00237]
Health: good	0.00161 [0.00219]

Health: excellent	0.00183 [0.00258]
Inverse Mills ratio	-0.00810 [0.00265]**
<hr/>	
Regional dummies	Yes
Wave dummies	Yes
Observations	32356
Pseudo R-squared	0.1118
Log-likelihood	-1557.9151
<hr/>	

Note: + < 10%; * < 5%; ** < 1%. Standard errors are in parentheses. Reference groups include in full-time employment, married, education level less than first degree, health: very poor.

Table A1: Instrumental regression for selection bias

Dependent variable: In a couple = 1, 0 otherwise	β
Divorce rate of other households within the region (in %)	-0.055 [0.009]**
Male	0.218 [0.017]**
Age	0.166 [0.002]**
Age-squared/100	-0.152 [0.003]**
Number of younger children (aged under 12)	0.577 [0.020]**
Number of older children (12 <= age < 18)	-0.018 [0.017]
Health: poor	0.059 [0.030]+
Health: fair	0.066 [0.033]*
Health: good	0.071 [0.034]*
Health: excellent	0.041 [0.036]
Log of real household income per capita	0.389 [0.013]**
Completed first degree	-0.124 [0.030]**
Completed higher degree	-0.208 [0.063]**
In full-time employment	0.100 [0.017]**
Unemployed	-0.229 [0.027]**
Disabled	-0.377 [0.034]**
Constant	-7.033 [0.147]**
Wave dummies	Yes
Observations	204845
Log likelihood	-101618.8

Note: + < 10%; * < 5%; ** < 1%. Standard errors are in parentheses. Reference groups include women, not in full-time employment (e.g. self-employed, retired, family care, student, in government training, maternity leave, and other type of employment), education level less than first degree, and health: very poor.

Table A2: Estimated GMM-system coefficients on the control variables**(Columns 11 and 12, Table 2)**

Dependent variable: Own LS at t	GMM-system with full controls	
	Men	Women
Ln(real household income per capita)	-0.043 [0.046]	-0.215 [0.094]*
Self-employed	0.162 [0.116]	0.068 [0.194]
Unemployed	0.130 [0.248]	-0.350 [0.370]
Retired	0.226 [0.144]	0.102 [0.165]
Maternity leave	3.677 [5.896]	0.508 [0.340]
Family care	-0.272 [0.502]	0.097 [0.126]
Full-time student	0.253 [0.417]	-0.506 [0.336]
Disabled, long-term sickness	-0.173 [0.289]	-0.332 [0.291]
Government training scheme	-2.076 [2.595]	-0.026 [2.507]
Other type of employment	0.489 [0.576]	-0.264 [0.700]
Number of years as a couple in the panel	-0.061 [0.100]	0.009 [0.106]
Number of years as a couple-squared	0.002 [0.005]	-0.002 [0.005]
Cohabit	0.063 [0.064]	0.018 [0.092]
Number of children (aged<16)	-0.015 [0.042]	-0.051 [0.061]
Completed first degree	0.154 [0.159]	0.177 [0.141]
Completed higher degree	-0.213 [0.207]	-0.235 [0.250]
Health: poor	0.970 [0.395]*	0.213 [0.379]
Health: fair	1.363 [0.369]**	0.680 [0.355]+
Health: good	1.763 [0.366]**	1.257 [0.351]**
Health: excellent	2.125 [0.383]**	1.443 [0.368]**

Note: + < 10%; * < 5%; ** < 1%. Standard errors are in parentheses. Reference groups include in full-time employment, married, education level less than first degree, health: very poor.