

# Separation Risk over Union Duration: An Immediate Itch?

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## Abstract

This study examines the risk of separation over union duration. Previous research reports a rising-falling pattern of divorce risk over marriage duration consistent with psychological notions of ‘honeymoon’ and ‘seven-year itch’. Little is known about the variation of the separation risk over cohabitation duration or over marriage duration when the length of partnership is measured from the beginning of coresidence. We include data on non-marital and marital unions and propose a novel way of treating cohabitation and marriage as episodes of the same union. We use Finnish large-scale register data and control for individuals’ observed and unobserved characteristics. Our results show that in cohabitations, the separation rate is highest at the beginning of union. Entry into marriage is followed by a significant drop in separation levels and a modest rising-falling pattern, which is independent of the length of pre-marital cohabitation. Marriage entails permanence, with a short ‘honeymoon’ effect and a long-term ‘effect’, much of which probably reflects self-selection of committed and satisfied cohabiters to marriage.

## Introduction

In industrialized societies, the break-up of a coresidential partnership has become a common life event, and it has consequences for both adults and children (Amato, 2000; Härkönen, Bernardi and Boertien, 2017). Over the past few decades, a myriad of studies have been published on demographic, social, and economic factors that contribute to divorce or separation (Amato, 2010; Lyngstad and Jalovaara, 2010). The time dimensions of union dissolution, namely, the spouses’ ages, the union duration, the period, and the union cohort, received much attention in the 1980s but were little studied in the following decades. Inspired by some recent studies (Kulu, 2014; Schnor, 2015), we return to the classic topic of how individual time influences union stability with new methodological

solutions, better data, and—above all—contemporary realities of union formation and dissolution.

Previous research on marriage dissolution has consistently reported that the risk of divorce is low during the first years of marriage; it then increases, reaches its peak between third and sixth years of marriage, and then declines (Thornton and Rodgers, 1987; Andersson, 1995; Diekmann and Engelhardt, 1999; Kulu and Boyle, 2010; Lyngstad, 2011; Jalovaara, 2013). The psychological literature considers this pattern consistent with the notions such as ‘honeymoon-is-over’ and ‘seven-year itch’. Most married couples experience a decline in marital quality after the first years of marriage, with tensions tending to culminate near the seventh year of marriage (Kurdek, 1999; Lavner and Bradbury, 2010). By contrast, demographers have argued that the rising-falling

pattern of divorce risk may result from omitting important covariates or *unobserved heterogeneity* from the models. Divorce-prone individuals leave the risk population at a higher rate, leaving mostly individuals with low separation proneness in the sample, and the risk of divorce therefore declines over the marriage duration (Vaupe and Yashin, 1985; Kulu and Boyle, 2010).

Although research has added to our knowledge of the relationship between marriage duration and divorce, it has a clear limitation: only few studies include information on cohabitation outside marriage. In most European countries, non-marital cohabitation has become a common and widely accepted form of partnership. The Nordic countries are forerunners in this trend (Sobotka and Toulemon, 2008). Focusing on first unions in Finland, Jalovaara (2013) shows that the separation pattern over union duration is different for cohabitations and marriages. For marriages the risk of separation follows a rising-falling pattern, whereas the separation risk in cohabitations is much higher and highest at early points, supporting the notion that cohabitations tend to be short-term partnerships. However, cohabitations are also the main route to marriage: a notable proportion (around 4 of 10) of cohabiters marry, and looked from the other side, more than 9 of 10 Finnish couples that marry cohabited first (Jalovaara, 2012). Cohabitations are common especially in the young age groups: in 2015, among persons below age 30, most unions were cohabitations, and at age 35, one-third (Statistics Finland, 2017).

This article investigates separation risk over union duration, incorporating data on the formation and dissolution of both marriages and cohabitations. We extend previous research by proposing of a novel way of treating cohabitation and marriage as parts of the same union. This approach corresponds to and informs us about contemporary union dynamics and allows to assess the usefulness of conventional explanations for separation patterns over marriage duration. Another strength is that we use large-scale register data from Finland that allow a detailed analysis of the variation of separation risk over long (non-marital and marital) union durations when controlling for individuals' observed and unobserved characteristics, with symmetrical data on both partners but no sample bias arising from selective non-response.

## Previous Research

### Divorce over Marriage Duration

The psychological literature suggests that a marital relationship goes through various stages, as its quality changes over time (Levinger, 1983; Finkel, Simpson and Eastwick, 2017). The 'honeymoon period' is followed

by 'everyday routine', during which the differences between spouses' attitudes, values and behaviour come to light and are subject to discussion and arguments. During that period, spouses encourage some of their partners' behaviours and discourage others, and attempt to adapt to those behaviours that cannot easily be changed. The partners gradually accumulate knowledge of each other's characteristics and develop a view about whether to stay together. If the mutual adaptation is successful, a period of stability follows in the marital relationship when the risk of separation is low (Diekmann and Mitter, 1984; Diekmann and Engelhardt, 1999). Studies on marital quality show similar changes over marriage duration. Marital quality is perceived as high at the beginning of the relationship; it declines rapidly in the early years of marriage and possibly stabilizes thereafter (Kurdek, 1999; VanLaningham, Johnson and Amato, 2001; Umberson *et al.*, 2005; Lavner and Bradbury, 2010; Ermisch, Iacovou and Skew, 2011; Schmiedeberg and Schroder, 2016). The theories focus on marriage and are not explicit about how its duration is defined, while empirical studies tend to focus on marriage, often because there are limited data on unmarried couples.

Although marital quality and satisfaction are key issues in marital stability, other factors that individuals presumably assess when considering whether to end a partnership include barriers to separation (Levinger, 1976). Barriers are factors that keep partners together in addition to or even in the absence of mutual attraction. Examples include joint property, feelings of obligation towards the spouse and children, and normative pressures from surrounding social environment (*ibid.*). Barriers may increase over marriage duration (Kurdek, 1999), helping explain the decline in separation risk.

Empirical demographic and sociological research has focused on how divorce risk varies over marriage duration, showing a rising-falling pattern. Most studies report that the risk of divorce increases rapidly during the first years of marriage, peaks between the third and sixth years of marriage, and declines thereafter (Andersson, 1995; Kulu and Boyle, 2010; Lyngstad, 2011; Jalovaara, 2013). Thus, divorce levels are not the highest in the seventh year of marriage; however, a clear rising-falling pattern is observed, supporting the ideas of diminishing marital satisfaction along with increasing barriers over marital duration.

Studies reporting the rising-falling pattern of divorce over marriage duration have controlled for sets of spouses' demographic and socio-economic characteristics, such as number and ages of children, and spouses' education and employment status or income, to avoid

selection bias (e.g. Kulu and Boyle, 2010; Lyngstad, 2011). However, it is likely that some important characteristics have not been included, particularly factors such as spouses' personality traits and social values. For example, the sample may contain individuals who are prone to divorce because they are adventurous and novelty-seeking (Boertien, von Scheve and Park, 2017), as well as individuals who are unlikely to divorce because their views are conventional. If this were the case, the estimates of the risk of divorce at longer durations would be downward biased. The high-risk group leaves the risk population at a higher rate, and therefore, as time passes, the share of the low-risk group increases and the hazard of divorce for the population approaches that group's (low) risk levels (Vaupel and Yashin, 1985; Kulu, 2014).

Vaupel and Yashin (1985) showed that the rising-falling pattern of divorce risk over marriage duration could be an outcome of two different risk patterns; while the risk increases for the high-risk group, it decreases for the low-risk group (or is constant at low levels). In reality, the rising-falling pattern of divorce risk can be a product of several underlying patterns that are all consistent with the psychological argument that couples face an increased risk of divorce after the 'honeymoon period'. For example, the risk may increase for both groups rapidly in the first years of marriage and slowly thereafter, but the two subpopulations may have different levels, which will lead to the declining divorce levels after initial increase. With high-quality data and advanced methods, it is possible to consider the influence of both observed and unobserved heterogeneity in models of union dissolution. A recent study (Kulu, 2014) reported that the rising-falling pattern of divorce risk persisted when both observed and unmeasured (time-invariant) characteristics of individuals were controlled for, suggesting that the selection bias is not a focal explanation for that pattern but may be somewhat informative.

### Separation in Cohabitations

Overall, cohabitation is considered a 'looser bond' (Schoen and Weinick, 1993). Compared to marriage, it is characterized by weaker legal support, less social recognition, and less clear normative structures (Nock, 1995). A usual argument is that with increasing prevalence of cohabitation, countries progress through stages where cohabitation develops from being marginal behaviour, then a prelude or alternative to marriage, to finally being indistinguishable from marriage (Heuveline and Timberlake, 2004). However, recent research suggests that even in forerunner countries, cohabitation and

marriage continue to have distinct meanings (Perelli-Harris *et al.*, 2014). For most, cohabitation represents a lower level of commitment, greater freedom and a way to test the relationship, and marriage represents an ideal for ultimate commitment (*ibid.*). In line with this, research reports other differences between cohabitations and marriages. A main observation is that cohabiting couples separate at a much higher rate than married couples (e.g. Liefbroer and Dourleijn, 2006; Wu and Musick, 2008; Schnor, 2014; Perelli-Harris and Lyons-Amos, 2015). Also in the Nordic countries, surveys show that cohabiters have lower commitment to and satisfaction with their relationships than do married persons (Wiik, Keizer and Lappegård, 2012); satisfaction and commitment are positively related to planning to marry (Wiik, Bernhardt and Noack, 2010) and actually marrying (Moors and Bernhardt, 2009).

The differences between cohabitation and marriage may reflect the causal effects of marrying or being married, such as more social support or pressure to stay together; however, it is very likely that they partly reflect self-selection of more committed and satisfied partners into marriage (Schoen and Weinick, 1993; Kulu and Boyle, 2010). The selectivity of cohabitations presumably weakens as cohabitation becomes more common in the society, but when almost all couples cohabit first, then marriage is, in turn, selective (Liefbroer and Dourleijn, 2006); this may now be the case in countries such as Finland, where the majority of cohabitations eventually lead to either separation or marriage.

### The Present Study

Although there is some knowledge of the variation in separation risk in cohabitations, previous research has examined the risk of separation in marital and non-marital unions separately. Given that cohabitation has become the majority route to marriage in many countries, we show that research on union dissolution benefits from viewing cohabitation and marriage as parts of the same union (see, e.g. Teachman, Thomas and Paasch, 1991). Our goal is to provide an adequate and informative description of how separation risk varies at non-marital and marital stages of unions. This allows us to assemble existing pieces of knowledge on the stability of cohabitations and marriages and to assess the validity of the usual explanations for the patterns.

The first question is: Should we expect to find a rising-falling pattern of separation risk when all coresidential unions are observed? Although many psychological theories on marriage duration and divorce are silent on the differences between marriage and life

together, they seem to concern the latter and therefore would also apply to time in non-marital cohabitation. Briefly, when a new couple moves in together, a honeymoon-like period follows during which satisfaction is high, while incompatibilities and problems take time to surface. We posit the first hypothesis:

*Hypothesis 1A ('All-embracing honeymoon'): A rising-falling pattern of separation risk over union duration is observed when coresidential unions regardless of marital status are included.*

However, competing ideas are plausible. With high rates of cohabitation and separation, the threshold of forming and dissolving cohabitations is low: selection into coresidence in terms of satisfaction and commitment is weak, as are the consequences of moving in (such as normative pressures to continue the union once it has begun). Research suggests that the household formation process is different for cohabitation than for marriage: many cohabiters move in with partners soon in the relationship, often 'sliding' into cohabitation for convenience or being pushed by other events, such as changes in housing or employment, and often without plans on a long-term commitment (Manning and Smock, 2005; Sessler, 2010). That many cohabitations end soon seems practically inevitable. Accordingly, previous empirical evidence suggests that there is no initial rise in the risk of cohabitation dissolution; the risk is highest at the beginning of the union (Jalovaara, 2013). We propose a competing hypothesis:

*Hypothesis 1B ('Immediate itch'): During cohabitation, the separation risk is highest at early points; the previously found rising-falling pattern only characterizes time after marrying.*

The question then remains what the rising-falling pattern of separation after the entry into marriage reflects: trouble-free first years of living together (as the psychological theories suggest), or factors specific to marrying that depress separation risk for some time. This can best be answered by examining whether (and how) separation patterns by marriage duration depend on the length of union duration, measured from the beginning of coresidence. We posit two competing hypotheses:

*Hypothesis 2A ('Post-honeymoon marriages'): The rising-falling pattern is observed for marriages with no or with a short period of premarital cohabitation but not for couples who marry after living together for several years.*

This result would imply that the rising-falling pattern reflects the effects of married couples' living together.

There could be 'honeymoon' and 'post-honeymoon' stages, but couples who have lived together long have already passed the honeymoon and marrying therefore matters less. However, we also propose a competing hypothesis:

*Hypothesis 2B ('Significant marriage'): The rising-falling pattern is observed regardless of the duration of pre-marital cohabitation.*

In this case, support is found for explanations that emphasize marriage-specific processes, including the protective effect of marriage (e.g. greater social support and pressure to stay together once married) and self-selection of more satisfied and committed couples into marriage.

The inclusion of observed and unobserved heterogeneity in the models allows to assess the extent to which they influence the variation of separation risk over union duration. Previous research on marriages (Kulu, 2014) led us expect that the basic patterns are quite robust to the inclusion of observed and unobserved heterogeneity but that they explain some of the lowering of separation rates at longer durations. This would suggest that individuals who are prone to separate 'because of' their observed and unobserved characteristics are under-represented at longer durations, but that there may also be real processes such as psychological relationship dynamics built into union duration that cannot be attributed to other, measurable factors or individual-level separation proneness.

## Data and Methods

### Data

We used data prepared by Statistics Finland by linking data from a longitudinal population register and registers of employment, educational qualifications, vital events, and other register sources. The extract used in this study was an 11 per cent random sample of persons born between 1940 and 1995 who were counted in Finland's population between 1970 and 2009. The data included full histories of childbearing and coresidential partnerships for the sample persons, along with educational histories and annual measurements of economic activities, incomes, and other data for the sample members and all their partners until the year 2009. The sample included data on the timing of vital events, including union formation and dissolution, with the precision of 1 month.

From 1987 onwards, cohabitations and marriages are identifiable: Finnish registers are exceptional in that they contain information on the place of residence down

to the specific dwelling, enabling the linkage of individuals to coresidential couples even when they are childless and unmarried. (For details, see ‘Inference of cohabitations’ in [Supplementary Material](#)).

The analyses focused on cohabitations and marriages of women formed between January 1988 and September 2009. All the unions of each woman formed during that period were included.<sup>1</sup> The women were born between 1940 and 1992. Data on the unions of foreign-born women were eliminated due to the lack of information on the life histories of persons born abroad covering the time preceding immigration.

Exposure time (i.e. couple-months at risk) was calculated separately for three types of unions: all unions regardless of marital status, cohabitations, and marriages. Exposure time for all unions regardless of marital status was calculated as follows: the unions were followed from the time (i.e. the month) the partners moved in together or married, whichever came first. The unions were right censored at the death of either partner, emigration of the woman, or September 2009. The exposure for cohabitations was calculated in the same way except that entry into marriage was introduced as an additional right censor. Marriages were followed from entry into marriage and right censored just as with all unions. The outcome event in all the analyses was (permanent) separation. In the case of cohabitations, separation was defined as moving apart; for marriages, it was defined as moving apart or judicial divorce, whichever came first.

The analyses covered approximately 140,000 unions, of which 121,000 were included in the analysis as cohabitations and 57,000 as marriages. Of cohabitations, approximately 51,000 and of marriages 15,000 ended in separation. The number of separations per 100 years at risk was 7.5 for all unions, 11.8 for cohabitations, and 3.3 for marriages.

## Methods and Analytic Strategy

We use a continuous-time multilevel event history model to study the risk of separation over union duration (Kulu, 2014). The basic model is specified as follows:

$$\text{In}h_{ij}(t) = \text{In}h_0(t) + \beta X_{ij}(t) + \sum_k \gamma_k Z_{ijk}(t) + \varepsilon_i, \quad (1)$$

where  $h_{ij}(t)$  denotes the hazard of separation of  $j$ th union for woman  $i$ .  $\text{In}h_0(t)$  represents the baseline log-hazard, the duration of the union, which we specify as the piecewise constant. The piecewise constant specification provides a flexible way of measuring the shape of the baseline hazard. The time (or union duration) is, in most analyses, divided into 1-year intervals. Although the hazard is assumed constant within each 1-year category

of duration, it could vary between them.  $x_{ij}(t)$  represents the values of a variable for the union type (marital or non-marital), and  $\beta_k$  measures its effect on union dissolution. The model also includes time-constant and time-varying covariates denoted by  $z_{ijk}(t)$ , with parameters  $\gamma_k$  measuring their effect. We also include a woman-level residual (or random effect) to control for the time-invariant unmeasured characteristics of a woman that influence the hazard of separation for any of her unions.

Identification of the model was attained through within-person replication. Some women had experienced more than one partnership episode. Of the women, 30 per cent had more than one union observed in the data (70 per cent had one, 22 per cent had two, 6 per cent had three, and 2 per cent had four or more unions); therefore, it was both possible and necessary to include woman-level random effects (‘shared frailty’) (Hoem, 1990; Gutierrez, 2002). We experimented with gamma and inverse Gaussian-distributed shared frailty. The results were similar, and we present the results for gamma-distributed shared frailty, which is widely used in the literature because it has a flexible shape and is analytically tractable (Gutierrez, 2002). The basic model (equation (1)) described above includes a dummy variable to distinguish between episodes in which individuals are cohabiting from those in which they are married. Once an individual moves from one union status to another (i.e. marries), her separation risk can change. However, individuals are assumed to follow the same separation pattern over union duration whether or not they are married. A conventional approach to relax this assumption is to fit a model with separate baselines for cohabitations and marriages by examining the two union types separately, for instance. This is also what we do as the first step of the analysis. However, this solution is not satisfactory for the current study because the conventional approach treats marital and non-marital spells of the same union as two different unions (both start with duration 0), which does not correspond to real-life experience where many cohabiters marry and most marriages are preceded by cohabitation with the partner. We propose to extend the basic model by including in the analysis time since marriage for marital episodes:

$$\text{In}h_{ij}(t) = \text{In}h_0(t) + \beta m_{ij}(t) + \sum_k \gamma_k Z_{ijk}(t) + \varepsilon_i, \quad (2)$$

where  $m_{ij}(t)$  denotes the time since marriage formation of the  $j$ th union for woman  $i$ . In this study, we divide marriage duration into 1-year intervals and use a set of dummies to measure its effect. Note that the baseline now represents the shape of hazard of separation for non-marital unions. The (log) risk of marital separation

at any time point of union duration is a sum of the effects of cohabitation (or union) and marriage duration. We will later illustrate the computation of marital separation risks. The proposed model draws upon the notion of ‘multiple clocks’ proposed by Lillard (1993).

In the models, we control for basic demographic and socio-economic characteristics of the unions and partners. Most of these control variables are time-varying covariates updated monthly (e.g. data on children and educational attainment) or yearly. To control for other time dimensions, we control for *period*, *age at union formation*<sup>2</sup> for the female partner (collapsed into seven categories), and *age difference* of the partners. *Union order* (first or subsequent) is based on civil status in 1987, and from 1988 onwards, on the data on all coresidential unions. (‘Union order’ is excluded from multilevel models.<sup>3</sup>) We control for the (woman’s) *number of children*, *age of the youngest child*, and a dummy indicating whether the *youngest child was a common child* of the partners. To control for socio-economic status and resources, we include *educational attainment* and annual *income* of each partner, and *home ownership*. Moreover, models for marriage only include a dummy indicating whether the couple cohabited before marriage. [Supplementary Table S3](#) shows the covariate categories and provides the distributions of total exposure and the number of separations by the control variables separately for cohabitations and marriages.

Stata software (Stata Corp 2017) was used to analyse the data. In Stata, fitting a piecewise constant hazard rate model is a multistep process: one first splits the time axis into intervals, defines a set of baseline dummy variables each representing one interval, and then estimates an exponential model using the `streg` command (Blossfeld, Golsch and Rohwer, 2007). The STPIECE module performs the steps more automatically (Sorensen, 1999). To obtain a model with separate baselines by a covariate—in our case the length of premarital cohabitation—one needs to define a set of dummy variables that represent each combination of time intervals and categories of the covariate (see Blossfeld *et al.*, 2007, pp. 125–127).

To illustrate the link between cohabitation and marriage in the study population, we present rates of separation and marriage among cohabiters from two simple hazard models. For another illustration, cumulative incidences (Coviello and Boggess, 2004) are used to calculate the cumulative probabilities of separation and marriage among cohabiters.

## Results

The empirical analysis proceeds as follows. We start by analysing separation patterns in cohabitations and

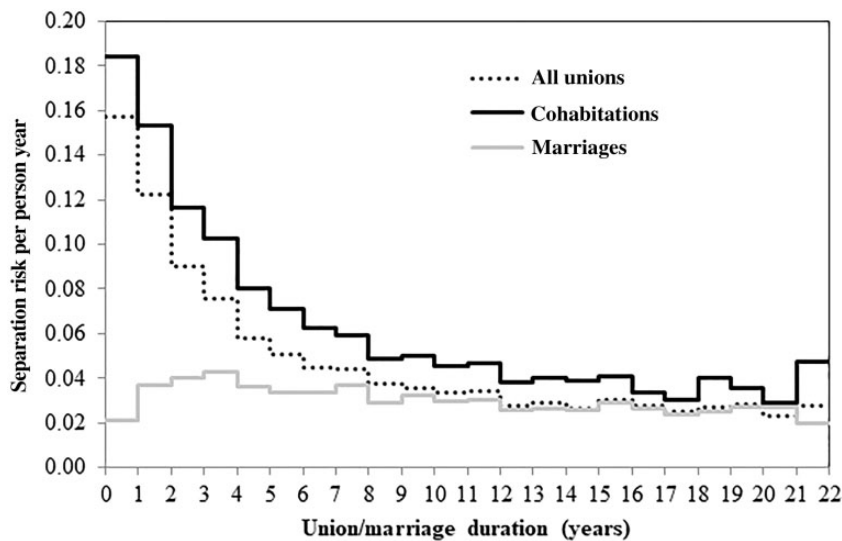
marriages separately. The introductory analysis first describes separation patterns for cohabitations, marriages, and all unions (i.e. cohabitations and marriages) by union duration. We continue by analysing cohabitations and marriages separately using different models that include controls for observed and unobserved characteristics of partners to identify their influence on the patterns. In the main analysis, we include all unions and consider the non-marital and marital episodes as parts of the same union and explore alternative ways of incorporating marital status into the models of union dissolution to distinguish between non-marital and marital episodes of unions. Models for all unions include two alternative measures of union type. We first distinguish between non-marital and marital episodes of the same union using a dummy variable, thus assuming the same risk patterns over union duration for marital and non-marital episodes but at different levels. We then use a measure that distinguishes between not only cohabitations and marriages but also the time in years elapsed since the entry into marriage. Although the latter is first included as a main effect only, it is ultimately included as a stratifying variable that enables us to examine whether the separation pattern of marriages is influenced by the length of premarital cohabitation.

## Introductory Models

[Figure 1](#) shows the yearly separation risks in different types of unions. They are based on introductory models that were fitted separately for each union type: cohabitations, marriages, and all unions. They do not include any control variables or shared frailty—just the baseline dummy variables. As expected, the separation pattern over union duration is very different for cohabitations and marriages. The separation rate for cohabitations is very high at early points (although unions lasting less than 3 months were not included), and the longer the cohabitation has lasted, the lower the separation rate is.<sup>4</sup> The curve for all unions is similar because cohabitations dominate the numbers, especially at early points. The marital separation rate is much lower, and there is a rising-falling pattern: the marital separation rate first increases, remains somewhat higher for a few years, and decreases thereafter. Notably, the initial increase in the rise-and-fall occurs as rapidly as after the first year in marriage. The magnitude of the initial increase appears modest in this graph. The rate nevertheless doubles between the first and second years.

## Inclusion of Observed and Unobserved Heterogeneity

The next step was to determine how controlling for observed and unobserved heterogeneity affects



**Figure 1.** Yearly separation risks by union/marriage duration for cohabitations, marriages, and all unions. Introductory models that were fitted separately for each union type and that do not include any control variables or shared frailty; only the baseline

separation patterns over union duration for cohabitations on the one hand and marriages on the other hand. Separate models were fitted for cohabitations and marriages. Figure 2 shows the relative separation risks by union duration for cohabitations and Figure 3 for marriages. In both figures, Model 1 includes only union/marriage duration; in Model 2, the control variables are added; and Model 3 includes the control variables along with a woman-level random effect (shared frailty). Here, and in all subsequent analyses, the baseline risks are presented as relative risks (rather than yearly or monthly separation risks provided by the model), since this facilitates comparisons of baseline shapes. When observed and unobserved heterogeneity is included, the shape of the baseline remains essentially the same for both union types. It nevertheless seems that observed and unobserved heterogeneity explain some of the lowering of separation rates at longer durations. This is expected since individuals who are less likely to separate ‘because of’ their observed and unobserved characteristics are overrepresented at longer union durations.

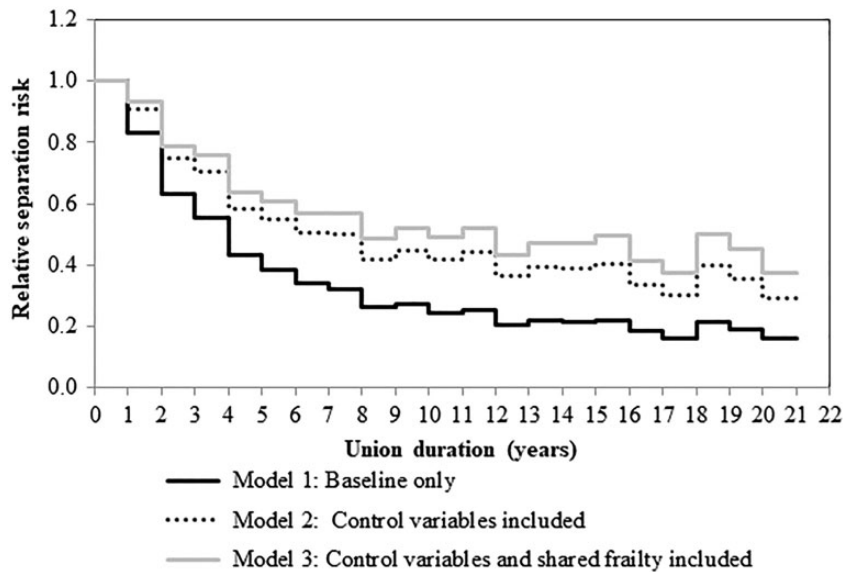
### Marriage among Cohabitors

Before proceeding to analyses in which cohabitations and marriages are viewed as stages of the same union, we show how the two are linked in the study population with respect to cohabiting couples converting their unions into marriage. As Supplementary Figure S7 shows, entry into marriage is most common during the first 4 years. As the cumulative incidences in

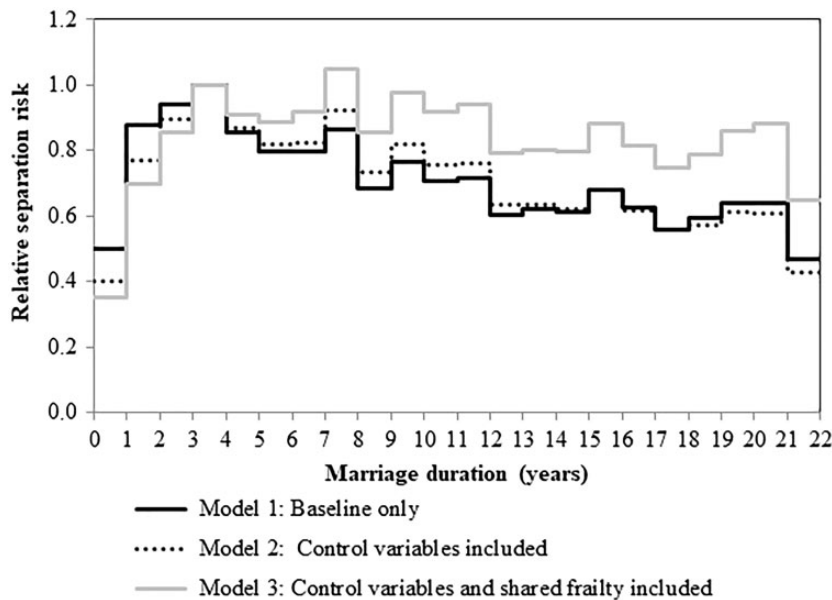
Supplementary Figure S8 show, 40 per cent of cohabiting couples eventually marry. What the cumulative incidences also show is that long cohabitations are uncommon: during the first 8 and 15 years, 80 and 90 per cent of the couples, respectively, have either separated or married, with the median being 2.7 years.

### Separation in All Unions and the Effect of Marrying

We now proceed to analyses of all unions in which we view cohabitations and marriages as stages of the same union. To illustrate the situation, we first fit a model with a dummy variable indicating whether the union was a marriage or a cohabitation (for the model specification, see equation (1)). The model includes the control variables (except union order) and shared frailty. (For the hazard ratios all covariates in the model, see Model 1 in Supplementary Table S4). The hazard ratio for marriages is 0.44. If this model is used to estimate the separation baselines for marriages, the baseline hazards (representing cohabitations) are multiplied with that ratio. The results of this simple calculation, that is, separate baselines for cohabitations and marriages, are provided in Table 1. Thus, according to this model, the separation risk for marriages is, at each duration, 56 per cent lower than for cohabitations. Figure 4 provides an illustration: let us assume that in the third year of their coresidential union, a couple marries. We also assume that the cohabitation dissolution



**Figure 2.** Relative separation risks by union duration for cohabitations from different models: Model 1 includes only union duration; in Model 2, the control variables are added, and Model 3 includes the control variables and shared frailty. Reference (Relative risk = 1) is the first year of cohabitation



**Figure 3.** Relative separation risks by marriage duration for marriages from different models: Model 1 includes only marriage duration; in Model 2, the control variables are added, and Model 3 includes the control variables and shared frailty. Reference (Relative risk = 1) is the fourth year of marriage

baseline applies to couples until entry into marriage. At entry into marriage, the separation risk would drop 56 per cent and remain at that lower level thereafter. This type of marital status dummy is what is

typically used in models of union dissolution if cohabitations are included. However, we already know that the shape of the baseline hazard is different for marriages than for cohabitations, and



therefore, it oversimplifies the patterns, at least for the purposes of this article.

As a remedy to this problem, we propose a model on separation risk in all unions in which the civil-status dummy is replaced with a variable that not only distinguishes between cohabitations and marriages but also includes marriage duration in years. Again, the model includes the control variables (except union duration) and shared frailty. (For the hazard ratios of the covariates, see Model 2 in [Supplementary Table S4](#)). The relative separation hazards for the new, more refined civil-status variable are also shown in [Table 2](#), Panel A. The hazard ratios indicate that the longer the marriage has lasted, the higher the separation risk is, eventually

almost reaching the level of the reference category, that is, cohabitations. However, these hazard ratios alone have little substantive meaning, because they are isolated from the baseline values; to obtain the meaningful values of separation risk over marriage duration, the separation risk over cohabitation duration and marriage duration should be analysed together. Therefore, we now use these hazard ratios to calculate the duration-specific separation risks for marriages as follows. Again, we assume that the cohabitation baseline applies to the couple until they marry. (The model baseline is shown in Panel B of [Table 2](#), presented as relative risks). Thereafter, the couple's separation risk moves to the level of marriage, which is calculated by multiplying, from that duration year onwards, the baseline risk (representing cohabitations) with the corresponding hazard ratio for marriage duration (obtained from the more refined civil-status variable). Panel C in [Table 2](#) shows the resulting baselines. For example, the baseline for a couple who marries during the third year of their coresidential union is calculated as follows (numbers bolded in the Table):  $0.82 \times 0.195 = 0.16$  (first year of marriage, third in union),  $0.80 \times 0.390 = 0.31$  (second year of marriage, fourth in union), and  $0.67 \times 0.504 = 0.34$  (third year of marriage, fifth in union); these rates are relative to the separation levels for the first year of cohabitation. The result is illustrated in [Figure 5](#), again assuming a couple marries during the third year of their union. We observe a significant drop in separation risks after the event of marriage followed by an increase and perhaps a slight decline thereafter. We thus observe a modest rising-falling pattern of marriage separations; however, the risk levels for marriages remain lower than for cohabitations, including at long durations (although the difference diminishes). The baseline shape characteristic of marriages, including an initial rise in the separation risk, is now integrated into the picture.

The previous calculations are based on the assumption that the separation risk in marriages is the same regardless of whether and how long the couple has cohabited before marriage. To determine whether this is a reasonable assumption, we fitted a model with separate baselines by the length of premarital cohabitation and, to obtain a comparison point, another model for cohabitation dissolution. [Supplementary Table S5](#) shows how the model was estimated with Stata. First, to obtain separation risks for marriages by length of premarital cohabitation, dummy variables were created that represent each

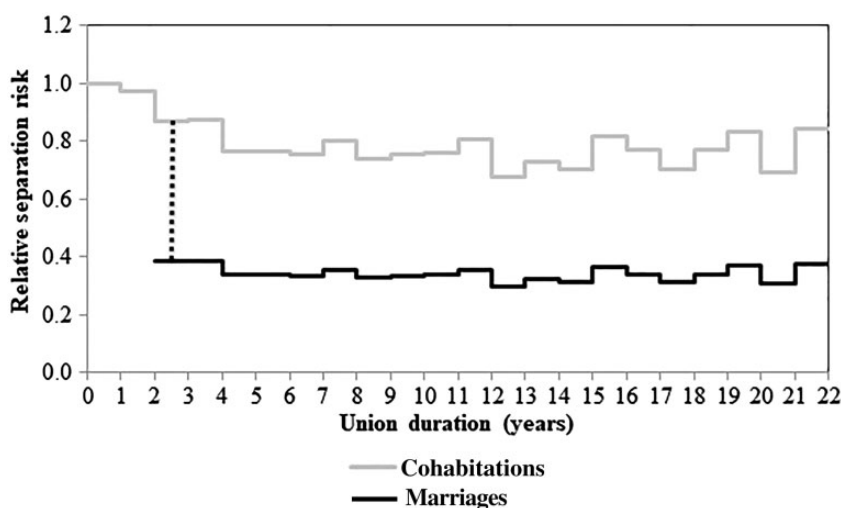
**Table 1.** Relative separation risks by union duration for cohabitations and marriages from the model that includes the civil-status dummy. The model also includes the control variables (except union order) and shared frailty. Reference (Relative risk = 1) is the first year of cohabitation

(A) Hazard ratios for the civil-status dummy (ref: Cohabitation) <sup>a</sup>		
Married	0.44	
Union duration	Relative separation risk	
(years)	(B) Cohabitations <sup>b</sup>	(C) Marriages <sup>c</sup>
1	1	0.44
2	0.97	0.43
3	0.87	0.38
4	0.87	0.39
5	0.77	0.34
6	0.76	0.34
7	0.75	0.33
8	0.80	0.35
9	0.74	0.33
10	0.76	0.33
11	0.76	0.34
12	0.81	0.36
13	0.68	0.30
14	0.73	0.32
15	0.70	0.31
16	0.82	0.36
17	0.77	0.34
18	0.70	0.31
19	0.77	0.34
20	0.83	0.37
21	0.69	0.31
22	0.84	0.37

<sup>a</sup>Hazard ratio for a covariate in the model.

<sup>b</sup>Relative risks calculated from the baseline risks of the model.

<sup>c</sup>Relative separation risks calculated by multiplying the hazard ratio (A) and (relative) baseline risks (B).



**Figure 4.** Relative separation risks by union duration for couples marrying during the third year of their coresidential union, according to the model with a marital-status dummy. The model includes the control variables (except union order) and shared frailty. Reference (Relative risk = 1) is the first year of cohabitation. The figure illustrates results shown in [Table 1](#)

combination of time intervals and categories of the covariate. To decrease random variation, wider time intervals were used: 1-year intervals for the first 3 years and 2-year intervals from that onwards. The model included the resulting  $12 \times 6 = 72$  dummy variables and all control variables, including union order. Shared frailty was not included owing to convergence problems. The yearly separation risks obtained from the model (by multiplying monthly risks by 12) are shown in [Supplementary Table S5](#), Panel A. Another model was fit to obtain separation risks for cohabitations using the wider time intervals. Again, all control variables were included. The resulting yearly separation risks are shown in [Supplementary Table S5](#), Panel B. Finally, the separation risks from these two models were used to calculate relative separation risks by union duration and length of premarital cohabitation by dividing each yearly risk with the yearly risks for the reference category (first year in cohabitation). The results are given in [Supplementary Table S5](#), Panel C. The same relative separation risks by marriage duration are shown in [Figure 6](#).<sup>5</sup> Upon visual inspection, it appears that the shape of the marital separation baseline does not depend on the length of premarital cohabitation. In other words, the influence of marriage formation on the risk of separation is the same regardless of how long the couple has resided together in that it is followed by a similar rise and fall. Thus, it seems that the previous model, illustrated in

[Table 2](#) and [Figure 5](#), provides an adequate representation of the data.

In [Figure 6](#), the group ‘married after <1 years’ cohabitation’ includes those who seemed to marry directly without cohabiting first. In supplementary analyses, they were distinguished. The shape of the separation baseline is very similar for the direct marriers, but, consistent with research comparing direct marriages and previous cohabiters ([Kulu and Boyle, 2010](#)), the level is somewhat lower.

## Conclusions

Using large-scale register data from Finland, this study investigated the variation of the separation risk over union duration, incorporating data on both cohabitations and marriages. We proposed a novel way of treating cohabitations and marriages as parts of the same union. This contributes to earlier research that has been confined to marriages (e.g. [Kulu, 2014](#)), with some evidence on a different pattern for cohabitations ([Jalovaara, 2013](#)).

Our results showed that separation levels are highest at the beginning of coresidential unions and decline over union duration. Entry into marriage is followed by a significant drop in separation risk, followed by a rise and a fall, albeit at modest levels. It seems that some of the stabilizing ‘effect’ of marrying is short term and is followed by a rise as soon as after the first year (creating the rising-falling shape of the marriage baseline). However,

**Table 2.** Relative separation risks by union duration for cohabitations and marriages from the model that includes the civil-status variable with marriage duration. The model also includes all control variables but no shared frailty. Reference (Relative risk = 1) is the first year of cohabitation

(A) Hazard ratios for civil status and marriage duration<sup>a</sup>

Ref: Cohabitation (HR = 1)

Married, first year	0.195
Married, second year	0.390
Married, third year	0.504
Married, fourth year	0.570
Married, fifth year	0.662
Married, sixth year	0.611
Married, seventh year	0.646
Married, eighth year	0.702
Married, ninth year	0.717
Married, 10th year	0.776
Married, 11th year	0.741
Married, 12th year	0.696
Married, 13th year	0.836
Married, 14th year	0.740
Married, 15th year	0.750
Married, 16th year	0.710
Married, 17th year	0.881
Married, 18th year	0.722

Relative separation risk

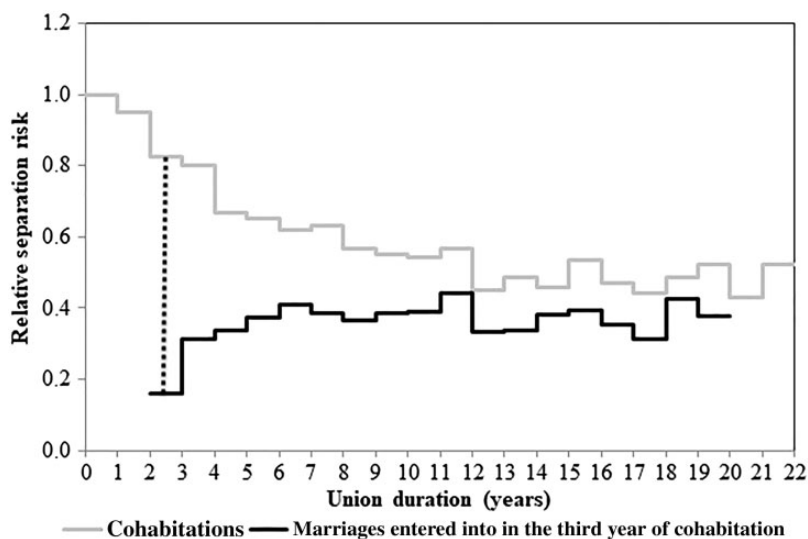
Union duration (years)	(B) Baseline risk: Cohabitations <sup>b</sup>	(C) Marriages <sup>c</sup>					
		Length of premarital cohabitation (years)					
		< 1	2	3	4	5	6+
1	1	0.19					
2	0.95	0.37	0.18				
3	<b>0.82</b>	0.42	0.32	<b>0.16</b>			
4	<b>0.80</b>	0.46	0.40	<b>0.31</b>	0.16		
5	<b>0.67</b>	0.44	0.38	<b>0.34</b>	0.26	0.13	
6	0.65	0.40	0.43	0.37	0.33	0.25	
7	0.62	0.40	0.38	0.41	0.35	0.31	0.12
8	0.63	0.44	0.41	0.39	0.42	0.36	0.25
9	0.57	0.40	0.40	0.37	0.35	0.37	0.29
10	0.55	0.43	0.39	0.39	0.36	0.34	0.31
11	0.54	0.40	0.42	0.39	0.38	0.35	0.36
12	0.57	0.40	0.42	0.44	0.41	0.40	0.35
13	0.45	0.38	0.31	0.33	0.35	0.32	0.29
14	0.48	0.36	0.41	0.34	0.36	0.38	0.34
15	0.46	0.34	0.34	0.38	0.32	0.34	0.33
16	0.53	0.38	0.40	0.40	0.45	0.37	0.41
17	0.47	0.41	0.33	0.35	0.35	0.39	0.35
18	0.44	0.32	0.39	0.31	0.33	0.33	0.31
19	0.48		0.35	0.43	0.34	0.36	0.40
20	0.52			0.38	0.46	0.37	0.39
21	0.43				0.31	0.38	0.32
22	0.52					0.38	0.37

Bold values: calculation explained in the text.

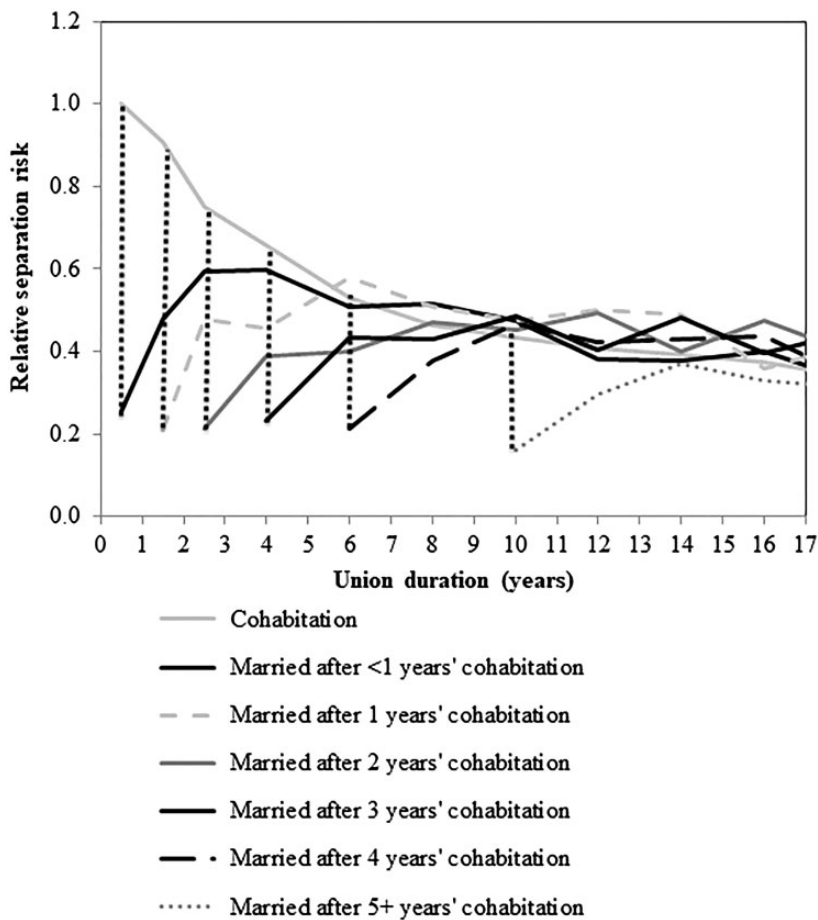
<sup>a</sup>Hazard ratios for a covariate in the model.

<sup>b</sup>Relative risks calculated from the baseline risks of the model.

<sup>c</sup>Relative separation risks calculated by multiplying hazard ratios (A) and (relative) baseline risks (B).



**Figure 5.** Relative separation risks by union duration for couples marrying during the third year of their coresidential union, according to the model that includes the marital-status variable with marriage duration. The model includes the control variables (except union order) and shared frailty. Reference (Relative risk = 1) is the first year of cohabitation. The figure illustrates results shown in [Table 2](#)



**Figure 6.** Relative separation risks by union duration in all unions; separate baselines by the length of premarital cohabitation. The models include the control variables but no shared frailty. Reference (Relative risk = 1) is the first year of cohabitation

some of it lasts longer: levels of marital dissolution remain low, although they approach those of cohabitations, which significantly decline with union duration.

With respect to the well-known rising-falling pattern (Kulu, 2014), it seems to be robust but specific to marriages, thus being only a part of a larger picture of union stability over union duration. If we focus on coresidence only, no initial period with lower separation risk is observed, supporting Hypothesis 1B ('Immediate itch'). Also, the rising-falling pattern is independent of the length of premarital cohabitation; supporting our Hypothesis 2B ('Significant marriage'), it is not confined to couples who have recently moved in together. These findings suggest that the stability-promoting 'honeymoon effect' is not about trouble-free couple interaction and high satisfaction in new partnerships but about processes specifically related to marrying. To what extent that effect is a protective effect of marriage (e.g. pressure to stay together created by a public declaration of love) and to what extent it is caused by the self-selection into marriage remains an open question. Given that 40 per cent of the cohabiters in our sample marry and half move apart during the first 15 years, it is almost a question of take-it-or-leave-it—or, in other words, choosing between marriage or splitting up. If marriage and separation are the outcomes of opposite forces, then there must be clear self-selection of satisfied and committed couples into marriage. This is in line with Brown (2003) stressing the marriage–separation polarity in quality factors: when most cohabiters expect to marry before long, couples failing to do so face low levels of relationship interaction and happiness.

As expected (Kulu, 2014), the patterns analysed are notably robust to the inclusion of observed and unobserved heterogeneity, although some of the decrease in separation risk in longer durations are explained by the characteristics of individuals and unions. Such robustness suggests that there are real processes, including relationship dynamics, built into the duration clocks of coresidence and marriage that are not easily explained by other correlated factors or self-selection processes.

The high separation rates for cohabitations suggest that there is a low threshold of forming and dissolving such unions. At least in the Nordic countries, cohabitation is in some respects increasingly a social substitute for marriage; however, judging from differences in permanence, these two relationship stages are far from equal. Thus, this study adds to recent research suggesting that despite high prevalence, cohabitation is generally not a replacement for or indistinguishable from

marriage (Perelli-Harris *et al.*, 2014; Perelli-Harris and Lyons-Amos, 2015). What also highlights the significance of marriage was that the union-stabilizing effect of marriage was independent of the length of premarital cohabitation. Apparently, the transition to marriage starts a new duration clock, and its effect is so pervasive that it abolishes the influence of the length of previous coresidence.

The role of childbearing should also be studied in more detail. In the Nordic context, the birth of a child often coincides with the transition from a non-marital to a marital union. Childbearing could thus be seen as one factor that potentially explains the drop in separation risk at marriage formation. Alternatively, the birth of a child can be seen as merely an indicator of other factors that influence both the relationship stability and the decision to have a child. Although our analysis controlled for both the number of children and their ages, a detailed analysis of the role of parenthood would lead to a better understanding of the factors determining the shape of separation risks for non-marital and marital unions (see Schnor, 2014). Furthermore, recent research (Schnor, 2014, 2015) suggests that the length of partnership prior to coresidence influences union stability and its duration patterns. Studying this question would require survey data.

Another issue to study is the role of self-selection into marriage. One option is to jointly model the processes of marriage and separation. Doing so would help determine whether unobserved characteristics of individuals that make them less prone to (marital) separation also increase their likelihood of marrying after a period of cohabitation (e.g. values, personality traits). This analysis would thus improve our knowledge of the causes of low separation risk after entry into marriage. However, such models have limitations because they cannot detect and control for unobserved factors that are union-specific and that influence the likelihood of cohabitants to marry (e.g. an excellent match between partners).

Research should also be conducted in other countries with similar partnership patterns. We believe that patterns in many countries are similar to those observed in this study, although there may also be significant differences caused by housing markets and policies and potentially by other institutional factors. In Nordic countries, with their flexible housing markets and welfare-state policies that support people during studies, for example, young couples have relatively easy access to rental housing; this suggests that they can easily form and dissolve coresidential unions, contributing to high separation levels.

## Notes

- 1 Since we have symmetrical data on men, women, and all their partners, it is practically irrelevant whether we include women's or men's unions.
- 2 Age at union formation is, for marriages, age at marriage, as this is consistent with the logic of our analyses. For a more general discussion, see Kuperberg (2014).
- 3 'Union order' variable and shared frailty are strongly correlated; even simple models with union order and shared frailty and only a few control variables lead to convergence problems. We also compared parameter estimates of pairs of models, i.e. one with union order and other with shared frailty. There were no differences between the models that would affect the conclusions (results are available upon request).
- 4 In supplementary analyses, the first year was divided into 3-month intervals and the second year to 6-month intervals. The result was the same: that the hazard is highest at early points—with the exception that it is very low during the first 3 months, which follows from the minimum duration of 90 days set for cohabitations in our data.
- 5 To improve readability, Figure 6 shows straight lines from hazard ratio to another rather than 'steps' as the previous graphs.

## Supplementary Data

Supplementary data are available at ESR online.

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