Demographic Research a free, expedited, online journal
of peer-reviewed research and commentary
in the population sciences published by the
Max Planck Institute for Demographic Research
Konrad-Zuse Str. 1, D-18057 Rostock · GERMANY
www.demographic-research.org

DEMOGRAPHIC RESEARCH

VOLUME 23, ARTICLE 31, PAGES 879-904
PUBLISHED 05 NOVEMBER 2010
http://www.demographic-research.org/Volumes/Vol23/31/
DOI: 10.4054/DemRes.2010.23.31

Research Article

Premarital cohabitation and divorce: Support for the "Trial Marriage" Theory?

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Demographic Research: Volume 23, Article 31
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Premarital cohabitation and divorce: Support for the "Trial Marriage" Theory?

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Paul J. Boyle[2]

Abstract

A number of studies show that premarital cohabitation is associated with an *increased risk* of subsequent marital dissolution. Some argue that this is a consequence of selection effects and that once these are controlled for premarital cohabitation has no effect on dissolution. We examine the effect of premarital cohabitation on subsequent marital dissolution by using rich retrospective life-history data from Austria. We model union formation and dissolution jointly to control for unobserved selectivity of cohabiters and non-cohabiters. Our results show that those who cohabit prior to marriage have a higher risk of marital dissolution. However, once observed and unobserved characteristics are controlled for, the risks of marital dissolution for those who cohabit prior to marriage are significantly lower than for those who marry directly. The finding that premarital cohabitation *decreases* the risk of marital separation provides support for the "trial marriage" theory.

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1. Introduction

The number of marriages ending in divorce rose rapidly in most developed nations during the 1970s and 1980s. Although divorce rates began to stabilise in the 1990s in some countries, they continued to rise in others. It has been shown that many factors are related to marriage dissolution, including women's increasing financial independence as their role in the labour market grows (Becker 1981; Hoem and Hoem 1992) and gender inequalities in wages gradually diminish (Davis and Joshi 1998); changes in gender roles (Kalmijn, de Graaf, and Poortman 2004; Lye and Biblarz 1993); factors related to the parental home, including parental separation (Amato 1996; Kiernan 1986); personal characteristics, such as educational qualifications (Hoem 1997a; Morgan and Rindfuss 1985); religious attitudes (Balakrishnan et al. 1987; Kalmijn, de Graaf, and Janssen 2005); the presence of one's own children (Erlangsen and Andersson 2001; Hoem 1997b; Morgan and Rindfuss 1985; Waite and Lillard 1991); the duration of the union (Chan and Halpin 2003); the partners' age at union formation (Tzeng and Mare 1995); the age gap between partners (Chan and Halpin 2003); the number of previous unions (O'Connor et al. 1999); and the place of residence and migration histories (Boyle et al. 2008; Muszynska and Kulu 2007). Recent research has also compared the relative influence of some of these factors across countries (Härkönen and Dronkers 2006; Liefborer and Dourleijn 2006). A detailed review of the factors associated with separation and divorce is provided by Amato (2010) and Lyngstad and Jalovaara (2010).

An additional factor that has generated considerable debate in the literature, and is becoming an increasingly common phenomenon, is the role of premarital cohabitation (Bumpass, Sweet, and Cherlin 1991; Ermisch and Francesconi 2000; Gabrielli and Hoem 2010; Hoem et al. 2009; Thornton and Philipov 2009). While some might imagine that premarital cohabitation would stabilise subsequent married unions, most of the literature suggests that it is in fact related to higher risks of marital dissolution. Various reasons for this empirical finding have been suggested, but one important factor that has not usually been controlled for adequately is the role of *unobserved selection*. Those who cohabit prior to marriage may have different unmeasured characteristics compared to those who do not and, if true, selection effects may mask the positive role that premarital cohabitation plays in subsequent marital stability.

The aim of this paper is to examine the effect of premarital cohabitation on divorce. We first study the relationship between premarital cohabitation and divorce with and without controlling for a set of demographic and socioeconomic characteristics of women and their partners. We then control for unmeasured characteristics of those who cohabit prior to marriage and those who marry directly to determine whether selection effects mask the "true" relationship between premarital cohabitation and divorce.

2. Premarital cohabitation and divorce

According to some theoretical arguments, we might expect premarital cohabitation to help stabilise subsequent married relationships, because those who cohabit will gain more information about their spouse than those who do not live together. Cohabiting partners who find they are well suited might consider marriage, while those who find they are incompatible will end the cohabitation (Teachman, Thomas, and Paasch 1991). Such "trial marriages" (Bennett, Blanc, and Bloom 1988) involve relatively low investment and are therefore easier to terminate; unsuccessful partnerships are effectively "weeded out" (Cherlin 1981; Klijzing 1992). Indeed, most young adults appear to believe that cohabitation improves the chances of a subsequent marriage (Kline et al. 2004), suggesting that lay people's views about premarital cohabitation concur with this theoretical perspective.

However, the majority of empirical studies have found that premarital cohabitation is associated with *higher* risks of subsequent marital dissolution compared to couples who married without prior cohabitation (Wagner and Weiss 2004). Bennett, Blanc, and Bloom (1988), Hoem and Hoem (1992), and Trussell, Rodríguez, and Vaughan (1992) noted this effect in Sweden. Axinn and Thornton (1992), Bumpass, Sweet, and Cherlin (1991), DeMaris and Rao (1992), Schoen (1992), Teachman and Polonko (1990), Teachman, Thomas, and Paasch (1991), and Thomson and Colella (1992) observed the disruptive effect of premarital cohabitation in the US. Balakrishnan et al. (1987), Hall and Zhao (1995), and Trussell, Rao, and White (1989) found the same in Canada, Bracher et al. (1993) in Australia, Manting (1992) and Klijzing (1992) in the Netherlands, Berrington and Diamond (1999) and Haskey (1992) in Britain, and Kiernan (2002a) and Liefbroer and Dourleijn (2006) in a number of European countries. Indeed, premarital cohabitation is also associated with lower marital satisfaction (Brown and Booth 1996), higher rates of wife infidelity (Forste and Tanfer 1996), and lower commitment to the partnership (Stanley, Whitton, and Markman 2004). While there is some limited evidence that the effect of premarital cohabitation on the risk of marital dissolution may have reduced for more recent birth cohorts (Schoen 1992; Brown et al. 2006; Reinhold 2010), other research suggests that this is not the case (Dush, Cohan, and Amato 2003). Overall, then, the consistency of these results in a number of countries makes this finding particularly persuasive and raises the question of why the empirical evidence does not support the "trial marriage" theory.

Some have suggested that the duration of the union has an effect on this relationship. Bennett, Blanc, and Bloom (1988) and Thomson and Colella (1992) found that marriages were more susceptible to divorce for those who cohabited for longer periods of time. Teachman and Polonko (1990) found that while prior cohabitation raised the risk of dissolution of subsequent marriage, once the duration of the entire union was accounted for the effect disappeared. Similarly, Hall (1996) found that those who cohabited for at least one year prior to marriage did not have a higher risk of marriage dissolution. However, duration of prior cohabitation was not found to influence subsequent marital instability by Lillard, Brien, and Waite (1995) – short cohabitations appeared to offer no advantage compared to longer cohabitations. More recently, Kline et al. (2004) showed that those who are engaged at the point when the couple starts cohabiting are at much less risk of subsequent marital break-up. Hence, commitment to the relationship appears to be an important aspect influencing later partnership success (Stanley and Markman 1992).

Others have argued that premarital cohabitation raises the risk of marriage dissolution because of *selection* effects. Cohabiters may have unobserved characteristics that make them more prone to separation, such as less conventional attitudes about marriage and, perhaps, higher expectations about the quality of unions, or poorer relationship skills (Bennett, Blanc, and Bloom 1988; Hall 1996; Smock 2000; Thomson and Colella 1992). For example, we know that those who cohabit tend to be more liberal, less religious, and more supportive of egalitarian gender roles (Clarkberg, Stolzenberg, and Waite 1995; Lye and Waldron 1997). For cohabiters, relationships in general, be they marital or non-marital, may be characterised by a lack of commitment and stability and they may be more willing to contemplate divorce if a subsequent marriage proves unsatisfactory (Bennett, Blanc, and Bloom 1988). Early studies seemed to support this selection hypothesis: Carlson (1986) reported that cohabiters were much more likely to view marriage as a response to social pressure than married couples, while Axinn and Thornton (1992) showed that cohabitation was selective of those who were less committed to marriage and more approving of divorce.

A second hypothesis is that the *experience* of cohabiting may also change people's views about marriage, making them less strongly committed to the institution (Axinn and Barber 1997; Axinn and Thornton 1992; DeMaris and Leslie 1984; Hall and Zhao 1995; Magdol et al. 1998; Thomson and Colella 1992). Through cohabitation, people may come to accept the temporary nature of relationships and to recognise that there are alternative arrangements to the more formal marriage. Of course, it is also possible that both selection and causation effects are operating concomitantly.

In an influential paper, Lillard, Brien, and Waite (1995) examined selection effects explicitly, recognising that some people choose to cohabit because they fear that marriage may not be successful. They applied a two-equation model that simultaneously considered time-invariant unmeasured characteristics of women that influenced the choice to cohabit and the characteristics that influenced marital dissolution. The strategy recognised that the decision to cohabit prior to the marriage may be endogenous to subsequent divorce and this was accounted for by allowing the unobserved heterogeneity to be correlated across the two decisions. The objective was to test if women with above-average risks of disruption had above-average propensities to cohabit prior to marriage instead of marrying directly. Their results provided convincing evidence of strong selection effects and showed that observed differences in union dissolution between married couples who had and had not cohabited previously disappeared once these were accounted for.

Lillard, Brien, and Waite's (1995) approach has since been applied, with varying results. Woods and Emery (2002) have found that controlling for selection effects eliminates the significant relationship between premarital cohabitation and marital instability, as did Steele, Kallis, and Joshi (2006). And, controlling for selection effects in a German study, Brüderl, Diekmann, and Engelhardt (1997) found that premarital cohabitation actually decreases the risk of divorce, suggesting that gathering information about the spouse during this period does increase subsequent marital stability. A more recent study by Svarer (2004) in Denmark supported these results. At the very least, these findings suggest that accounting for potential selection effects is critical when examining the influence of premarital cohabitation on subsequent marital dissolution.

In this paper, we follow the approach adopted by Lillard, Brien, and Waite (1995) and examine the relationship between premarital cohabitation and marital stability, controlling for potential selectivity by modelling the processes of union formation and dissolution simultaneously. Few contemporary studies have adopted this approach and only two have supported the notion of "trial marriage". We use data from Austria for two reasons. First, the rich retrospective data set (see below) allows us to investigate the relationship between premarital cohabitation and divorce when controlling for many demographic and socioeconomic characteristics of women and their partners. Second, Austria has around average rates of both union dissolution and cohabitation compared to other countries in Europe (Neyer 2003). Using data from 15 of the European countries (and the US) that participated in the Fertility and Family Surveys (FFS) conducted between 1989 and 1997, Andersson (2003) showed that in the US 42% of all marriages ended in dissolution within 15 years, compared to 37% in Latvia, which experienced the highest divorce rates among the European countries, and less than 10% in Italy, Spain, and Slovenia, which had the lowest rates. In Austria the corresponding figure was 25%, which was slightly above average for Europe and similar to the levels in Germany.

3. Data

The retrospective event-history data for this analysis were drawn from the Austrian Family and Fertility Survey (FFS) conducted in 1995–96. This was one of a sweep of surveys conducted in a number of European nations, Canada, New Zealand, and the US. The surveys included mainly consistent questions, but there were some variations, and an advantage of the Austrian survey is that it was one of the few to include detailed retrospective partnership and residential histories. This allowed us to identify where individuals were living at the time of a particular union and also to control for the effect of residential context when examining the relationship between premarital cohabitation and divorce (Boyle et al. 2008).

The response rate for the Austrian survey was 72% (Hoem, Prskawetz, and Neyer 2001), resulting in 4,581 female and 1,539 male respondents born between 1941 and 1976. As in many other studies (e.g., Berrington and Diamond 1999), the nature of the data meant that unions were defined based on the co-residence of two intimate (heterosexual) partners. We excluded those born outside Austria, those living abroad at age 15, and those for whom significant parts of the data were incomplete, leaving 3,804 women of whom 3,118 had been in a union (at least once) during their life and were therefore included in our analysis.

We constructed a multi-episode data set for union dissolutions where individuals are at risk from union formation and are followed until union dissolution, interview, or death (if not separated). Of the 3,118 women who had at least one partnership, 397 had a second, 62 had a third and 10 had a fourth union. The number of union dissolutions was 669 for first unions, 103 for second, 22 for third and 1 for fourth unions. Separations outside Austria and after return to Austria were excluded (a total of 22 events). We were particularly concerned with the dissolution rates for cohabiters, for those who married directly, and for those who married following an episode of cohabitation.

We included a range of time-varying and time-constant demographic and socioeconomic explanatory variables in the models presented below. These were identified from the literature reviewed above and included variables relating to family structure, women's independence, qualifications, and place of residence and migration (Table 1 provides a list of the categorical variables and their subcategories).

Table 1: Person-years and union dissolutions by categorical variables

	Person-years		Union dissolutions	
	Number	**Percent**	**Number**	**Percent**
Demographic and socioeconomic variables				
Partnership status				
Cohabiting	6022.24	14	313	39
Married, after cohabitation	10583.32	25	193	24
Married, directly	26063.81	61	289	36
Unions				
One union	39513.28	93	669	84
Two or more unions	3156.08	7	126	16
Children				
No children	9053.72	21	329	41
One child	12921.31	30	268	34
Two or more children	20694.33	48	198	25
Stepchildren				
No stepchildren	35882.47	84	638	80
One or more stepchildren	6786.89	16	157	20
Educational level				
Basic	38564.19	90	690	87
Secondary	2389.43	6	81	10
Higher	1715.74	4	24	3
Educational enrolment				
Not enrolled	41337.07	97	735	92
Enrolled	1332.29	3	60	8
Religiousness				
No	11774.78	28	350	44
Yes	30894.58	72	445	56
Parental home				
Parental divorce				
No	39869.58	93	691	87
Yes	2799.78	7	104	13

Table 1: (Continued)

	Person-years		Union dissolutions	
	Number	**Percent**	**Number**	**Percent**
Woman's independence				
Comparative education				
Man better educated	6942.64	16	127	16
No difference	34398.08	81	612	77
Woman better educated	1328.64	3	56	7
Employment status				
Not employed	26357.90	62	422	53
Employed	16311.46	38	373	47
Employment status (at start of union)				
Man employed, woman employed	24691.21	58	432	54
Man employed, woman not employed	14959.07	35	266	33
Man not employed, woman employed	1630.39	4	38	5
Man not employed, woman not employed	1388.69	3	59	7
Relative ages of partners				
Man younger	3425.24	8	87	11
No difference	7252.75	17	140	18
Man older	31991.38	75	568	71
Place of residence, migration, and mobility				
Place of residence				
Rural areas	30010.87	70	424	53
Towns and cities	5261.46	12	130	16
Vienna	7397.03	17	241	30
Migrations (inter-county moves)				
No migrations	38241.44	90	704	89
One migration	3791.06	9	68	9
Two or more migrations	636.86	1	23	3
Residential moves (intra-county moves)				
No moves	28407.94	67	565	71
One move	11517.74	27	162	20
Two or more moves	2743.68	6	68	9
Total	42669.36	100	795	100

Source: Authors' calculations based on the Austrian FFS data.

4. Methods

We modelled the time from union formation to dissolution using a series of hazard regression models (Blossfeld and Rohwer 1995; Hoem 1987; 1993). We were particularly interested in whether marital dissolution was influenced by premarital cohabitation, controlling for a range of other factors, as represented by equation 1:

$$\ln \mu_{ij}(t) = y(t) + \sum_k z_k(u_{ijk} + t) + \sum_l \alpha_l x_{ijl} + \sum_m \beta_m w_{ijm}(t), \quad (1)$$

where $\mu_{ij}(t)$ denotes the hazard of dissolution for the *j*th union for individual *i* and $y(t)$ denotes a piecewise linear spline that captures the impact of baseline (i.e., union) duration on the log-hazard. We used a piecewise linear spline specification, instead of the widely used piecewise constant approach, to pick up the baseline log-hazard. Parameter estimates are thus slopes for linear splines over user-defined time periods. With sufficient nodes (bend points), a piecewise linear specification can capture any log-hazard pattern in the data.[3] Here $z_k\left(u_{ijk} + t\right)$ denotes the spline representation of the effect of a time-varying variable that is a continuous function of *t* with origin u_{ijk} (e.g., a woman's age), x_{ijl} represents the values of a time-constant variable (e.g., parental divorce), and $w_{ijm}(t)$ represents a time-varying variable whose values can change only at discrete times (e.g., place of residence or partnership status).

A crucial part of the modelling was to investigate the possible role of unobserved selectivity bias (see Figure 1). In particular, we were interested in whether the apparent relationship between marital dissolution and premarital cohabitation would persist once the time-invariant unmeasured characteristics of those who enter cohabitation or marry without prior cohabitation were accounted for. The aim was to test if women with above-average risks of disruption have above-average propensities to cohabit (prior to marriage) and whether women with below-average risks of disruption have above-average propensities to marry without prior cohabitation. Thus we built a simultaneous-equations model to estimate jointly an equation for union dissolution and three equations for partnership formation (distinguishing between direct marriage and cohabitation and marriage for cohabiters). Person-specific heterogeneity terms were included in all four equations, allowing us to test for a correlation between these

[3] The value of the linear spline function between the points (t_n, y_n) and (t_{n+1}, y_{n+1}) is computed as follows: $y(t) = y_n + s_{n+1}(t - t_n)$ for $n = 0, 1, 2$ ..., where s_{n+1} is the slope of the linear spline over the interval $[t_n, t_{n+1}]$. To compute the linear spline function we thus need to define nodes and estimate from the data constant y_0 and slope parameters $s_1, s_2, \ldots$.

residual terms (cf. Lillard, Brien, and Waite 1995). The simultaneous equations model can be formalised as:

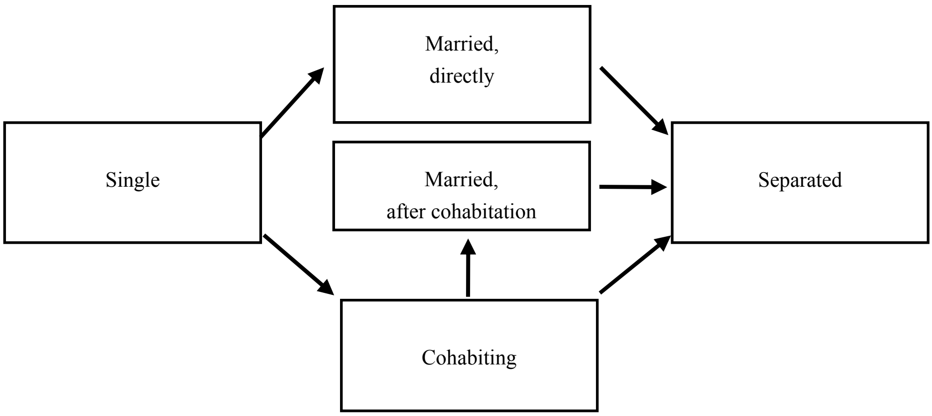
$$\begin{aligned}
\ln \mu_{ij}^{D}(t) &= y^{D}(t) + \sum\nolimits_{k} z_{k}^{D}(u_{ijk} + t) + \sum\nolimits_{l} \alpha_{l}^{D} x_{ijl} + \sum\nolimits_{m} \beta_{m}^{D} w_{ijm}(t) + \varepsilon_{i}^{D} \\
\ln \mu_{ij}^{C}(t) &= y^{C}(t) + \sum\nolimits_{k} z_{k}^{C}(u_{ijk} + t) + \sum\nolimits_{l} \alpha_{l}^{C} x_{ijl} + \sum\nolimits_{m} \beta_{m}^{C} w_{ijm}(t) + \varepsilon_{i}^{C} \\
\ln \mu_{ij}^{M}(t) &= y^{M}(t) + \sum\nolimits_{k} z_{k}^{M}(u_{ijk} + t) + \sum\nolimits_{l} \alpha_{l}^{M} x_{ijl} + \sum\nolimits_{m} \beta_{m}^{M} w_{ijm}(t) + \varepsilon_{i}^{M} \\
\ln \mu_{ij}^{CM}(t) &= y^{CM}(t) + \sum\nolimits_{k} z_{k}^{CM}(u_{ijk} + t) + \sum\nolimits_{l} \alpha_{l}^{CM} x_{ijl} + \sum\nolimits_{m} \beta_{m}^{CM} w_{ijm}(t) + \varepsilon_{i}^{CM}
\end{aligned} \tag{2}$$

where $\mu_{ij}{}^{D}(t)$ denotes the hazard of dissolution for the *j*th union of individual *i*, $\mu_{ij}{}^{C}(t)$ and $\mu_{ij}{}^{M}(t)$ represent the risk of transition to the *j*th cohabitation or direct marriage in the competing risk framework, and $\mu_{ij}{}^{CM}(t)$ denotes the hazard of marriage for the *j*th cohabitation. Here $\varepsilon_i{}^{D}$, $\varepsilon_i{}^{C}$, $\varepsilon_i{}^{M}$ and $\varepsilon_i{}^{CM}$ are person-specific heterogeneity terms for the dissolution, cohabitation, direct marriage, and cohabitation to marriage equations respectively. We assumed that the residuals would follow a joint multivariate normal distribution:

$$\begin{pmatrix} \varepsilon_i^{D} \\ \varepsilon_i^{C} \\ \varepsilon_i^{M} \\ \varepsilon_i^{CM} \end{pmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \\ 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2_{\varepsilon^D} & \rho_{\varepsilon^C \varepsilon^D} & \rho_{\varepsilon^M \varepsilon^D} & \rho_{\varepsilon^{CM} \varepsilon^D} \\ \rho_{\varepsilon^D \varepsilon^C} & \sigma^2_{\varepsilon^C} & \rho_{\varepsilon^M \varepsilon^C} & \rho_{\varepsilon^{CM} \varepsilon^C} \\ \rho_{\varepsilon^D \varepsilon^M} & \rho_{\varepsilon^C \varepsilon^M} & \sigma^2_{\varepsilon^M} & \rho_{\varepsilon^{CM} \varepsilon^M} \\ \rho_{\varepsilon^D \varepsilon^{CM}} & \rho_{\varepsilon^C \varepsilon^{CM}} & \rho_{\varepsilon^M \varepsilon^{CM}} & \sigma^2_{\varepsilon^{CM}} \end{pmatrix} \right) \tag{3}$$

where $\sigma^2_{\varepsilon^D}$, $\sigma^2_{\varepsilon^C}$, $\sigma^2_{\varepsilon^M}$, and $\sigma^2_{\varepsilon^{CM}}$ denote the variances of the person-specific residuals and $\rho_{\varepsilon^D \varepsilon^C}$, $\rho_{\varepsilon^D \varepsilon^M}$, $\rho_{\varepsilon^D \varepsilon^{CM}}$, $\rho_{\varepsilon^C \varepsilon^M}$, $\rho_{\varepsilon^C \varepsilon^{CM}}$, and $\rho_{\varepsilon^M \varepsilon^{CM}}$ are the covariances between the residuals.

Figure 1: Research model



Note: The arrows indicate the processes that were explicitly modelled in the simultaneous equations framework.

The model thus controlled for unobserved selection into cohabitation and direct marriages and selection of cohabiters into marriage. While Lillard, Brien, and Waite (1995) controlled for unmeasured selection into premarital cohabitation among those who were married, we modelled all partnership transitions of women and corresponding selection effects. Choosing a more complex strategy allowed us to explore and explicitly control for selection effects for various partnership transitions (e.g., selection into cohabitation, selection from cohabitation into marriage). We used hazard regression to study partnership formation in order to exploit all the information that was available; this strategy allowed us to include time-varying covariates in the analysis. The aim was to identify unobservables, having controlled properly for all the observed variables. The identification of our model was attained through within-person replication (see Lillard, Brien, and Waite 1995; Kulu 2005; 2006; Steele, Kallis, and Joshi 2006). There were 397 women who experienced multiple partnerships and 103 women who experienced multiple separations during the observation period. The model was implemented in aML (Lillard and Panis 2003) and the parameters were obtained using maximum-likelihood estimation.

5. Results

Table 2 provides the results for six models, which become increasingly complex. In Model 1 we considered the effect of partnership status on union dissolution, controlling only for the duration of the relationship and the age of the woman. Cohabiters were most likely to separate and those who cohabited prior to marriage were significantly more likely to separate than those who married directly. In Model 2 we also controlled for calendar time. The differences in the risk of union dissolution between various groups diminished, suggesting that cohabitants and those who cohabited prior to marriage are over-represented in more recent years, when the risk of union dissolution is higher than in earlier years (for various reasons). Still, the levels of divorce remained significantly higher for those who cohabited prior to marriage compared to those who married directly (the results of the likelihood ratio test to measure the improvement in the model fit are presented at the bottom of Table 2). Next we also included (other) demographic and socioeconomic characteristics in the model (Model 3). The differences in the dissolution levels decreased further between various groups of women. While the risk of union dissolution was still considerably higher for cohabiters than for the married, there were no significant differences between women who cohabited prior to marriage and women who married directly. The results were consistent with those obtained for Austria in other studies (see Kiernan 2002b; Liefbroer and Dourleijn 2006).

Table 2: The factors influencing union dissolution (parameter estimates)

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Constant (baseline)	-5.611***	-6.993***	-6.650***	-6.926***	-7.010***	-7.024***
Demographic and socioeconomic variables						
Union duration (baseline)[1]						
0–1 years (slope)	1.850***	1.845***	1.960***	2.028***	2.025***	2.012***
1–5 years (slope)	0.111***	0.105***	0.149***	0.192***	0.189***	0.179***
5–10 years (slope)	0.040	0.034	0.051	0.075**	0.070*	0.069*
10+ years (slope)	-0.002	-0.010	0.000	0.014	0.012	0.013
Age						
15–19 years (slope)	-0.135	-0.177*	-0.229**	-0.242**	-0.232**	-0.241**
20–24 years (slope)	-0.007	-0.025	-0.059*	-0.060	-0.051	-0.054
25–29 years (slope)	-0.102***	-0.114***	-0.161***	-0.167***	-0.167***	-0.169***
30–34 years (slope)	-0.048	-0.056	-0.104***	-0.118***	-0.122***	-0.124***
35+ years (slope)	-0.022	-0.031	-0.055**	-0.067**	-0.073**	-0.076***

Table 2: (Continued)

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Year						
–1969 (slope)		0.096	0.078	0.070	0.075	0.081
1970–79 (slope)		0.038*	0.047**	0.048**	0.054**	0.057**
1980–89 (slope)		0.029**	0.030**	0.036**	0.044***	0.046***
1990+ (slope)		0.051*	0.057**	0.060**	0.066**	0.068**
Partnership status						
Cohabiting	1.612***	1.332***	0.802***	0.865***	0.514***	0.354*
Married, after cohabitation	0.385***	0.202**	-0.050	-0.055	-0.419**	-0.373*
Married, directly	0	0	0	0	0	0
Unions						
One union			0	0	0	0
Two or more unions			0.525***	0.186	0.092	0.009
Time since first/last conception[1,2]						
0–0.75 years (slope)			-1.206***	-1.166***	-1.150***	-1.195***
0.75–2.75 years (slope)			0.609***	0.590***	0.583***	0.589***
2.75+ years (slope)			0.016	0.016	0.016	0.018
Children						
One child			0	0	0	0
Two or more children			-0.399***	-0.459***	-0.479***	-0.467***
Stepchildren						
No stepchildren			0	0	0	0
One or more stepchildren			0.162	0.173	0.160	0.184
Educational level						
Basic			0	0	0	0
Secondary			-0.258	-0.262	-0.260	-0.245
Higher			-0.411*	-0.447*	-0.440*	-0.456*
Educational enrolment						
Not enrolled			0	0	0	0
Enrolled			0.236	0.238	0.237	0.242
Religiousness[3]						
No			0	0	0	0
Yes			-0.333***	-0.392***	-0.444***	-0.474***
Parental home						
Parental divorce						
No			0	0	0	0
Yes			0.448***	0.489***	0.540***	0.553***

Table 2: (Continued)

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Woman's independence						
Comparative education						
Man better educated			0.001	-0.013	-0.014	-0.019
No difference			0	0	0	0
Woman better educated			0.904***	0.972***	0.983***	0.989***
Employment status						
Not employed			0	0	0	0
Employed			0.361***	0.393***	0.405***	0.399***
Employment status (at start of union)						
Man employed, woman employed			0	0	0	0
Man employed, woman not employed			0.024	0.041	0.035	0.040
Man not employed, woman employed			-0.025	-0.046	-0.038	-0.019
Man not employed, woman not employed			0.403**	0.454**	0.451**	0.477***
Relative ages of partners[4]						
Man younger			0.336**	0.381**	0.387**	0.388**
No difference			0	0	0	0
Man older			-0.103	-0.109	-0.101	-0.108
Place of residence, migration, and mobility						
Place of residence						
Rural areas			0	0	0	0
Towns and cities			0.342***	0.370***	0.402***	0.411***
Vienna			0.656***	0.727***	0.777***	0.794***
Migrations						
No migrations			0	0	0	0
One migration			0.039	0.040	0.067	0.049
Frequency of migrations						
One migration			0	0	0	0
Two or more migrations			0.933***	1.006***	1.013***	1.018***
Residential moves						
No moves			0	0	0	0
One move			-0.252***	-0.265***	-0.267***	-0.279***
Frequency of residential moves						
One move			0	0	0	0
Two or more moves			0.552***	0.577***	0.576***	0.559***

Table 2: (Continued)

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Standard deviation of residuals						
Union dissolution				0.714***	0.790***	0.856***
Cohabitation				1.366***	1.553***	1.551***
Direct marriage				1.654***	1.776***	1.782***
Cohabitation to marriage						0.696***
Correlation between the residuals						
Dissolution – cohabitation					0.225*	0.229*
Dissolution – marriage					-0.132	-0.125
Dissolution – cohabitation to marriage						-0.464**
Cohabitation – marriage					0.767***	0.765***
Cohabitation – cohabitation to marriage						-0.119
Marriage – cohabitation to marriage						-0.270*

Significance: * = 10%; ** = 5%; *** = 1%.

[1] For linear splines we present slope estimates that show how the hazard increases or decreases over a certain time period. For example, during pregnancy (see time since conception) the log-risk of dissolution changes by -1.195 per year (Model 6), reaching a level of -0.896 (0.75×(-1.195)) by the time of birth. In relative terms, the risk is then 59% lower than prior to conception ((exp(-0.896)-1)×100%). The log-hazard of union dissolution increases 0.589 per year during the first two years of the child's life, reaching a level of 0.282 (-0.896+(0.589×(2.75-0.75))) when the child is two, which is a 33% higher risk than prior to conception ((exp(0.282)-1)×100%).

[2] The reference category for the first conception is parity zero.

[3] Women were asked whether they were religious or not. Those women who answered "certainly yes" or "rather yes" were defined as religious.

[4] The age difference was defined to be present when one of the partners was older/younger than the other by more than one year.

Notes: Likelihood ratio test statistic (LR)

Model 2 versus Model 1: LR = 40.5, df = 4, p < 0.001; Model 3 versus Model 2: LR = 343.1, df = 25, p < 0.001; Model 4 versus Model 3: LR = 271.7, df = 3, p < 0.001; Model 5 versus Model 4: LR = 112.5, df = 3, p < 0.001; Model 6 versus Model 5: LR = 16.5, df = 4, p < 0.01; the likelihood of simultaneous-equations models was compared to a sum of the likelihoods of models for union dissolution and those for union formation. As our research focus is on union dissolution, we have not reported the parameter estimates for the union-formation equations (except for the standard deviation of the women-level residuals and the correlation between them).

Source: As for Table 1.

In Model 4, we included a person-level heterogeneity term. The standard deviation of the person-level residuals was significantly different from zero (0.714), indicating that there were women-specific unmeasured characteristics that affected all unions. Controlling for woman-level unobserved heterogeneity removed initial differences in the dissolution risk between first and higher order unions, as expected, but had no significant effect on the coefficients for premarital cohabitation. We extended the analysis in Model 5 by fitting a simultaneous-equations model that included the equation for union dissolution, as fitted in Model 4, and equations of the risk of marriage and cohabitation for single and separated people (but excluded an equation for the hazard of marriage for cohabiters). This allowed us to test whether women with above-average risk of union dissolution were also more likely to cohabit before marriage and whether women who were less prone to union disruption tended to marry without prior cohabitation. Each equation included person-specific residuals and the correlations between them are reported at the end of Model 5. The significant positive correlation between the residuals from the cohabitation and dissolution equations (0.225) indicates that women who cohabited had unobserved characteristics that increased the risk of union dissolution, whereas the negative correlation (although not significant) between marriage and dissolution (-0.132) suggests that women who married directly had unobserved characteristics that decreased the risk of union dissolution. Once unobserved selection effects were accounted for women who cohabited prior to marriage had a significantly *lower risk* of union dissolution than women who married directly.

It is likely that average cohabiters were more prone to union disruption than cohabiters who married subsequently (the latter were more prone to disruption than those who married directly). Therefore we might overestimate the marriage-stabilising effect of premarital cohabitation in our previous analysis. To control for unobserved selection from cohabitation into marriage, we extended our simultaneous-equations model by including also an equation for the hazard of marriage formation for cohabiters. The results are presented in Table 2, Model 6. The significant negative correlation between the residuals from the marriage-formation equation (for cohabiters) and the residuals from the dissolution equation (-0.464) suggests that women who married after an episode of cohabitation had unmeasured characteristics that reduced their risk of union dissolution compared to average cohabiters. As expected, the differences that were previously observed (Model 5) in the divorce levels between women who cohabited prior to marriage and those who married directly decreased; however, these remained significant. Women who cohabited prior to marriage still had a 31% lower risk of union dissolution than women who married directly. The 90%

confidence interval around this estimate was between 5% and 50% lower risk of union dissolution.[4]

Lillard, Brien, and Waite (1995) showed how important it is to control for selection effects when considering premarital cohabitation effects on union dissolution. Our results confirm this, but are even more dramatic than those presented by Lillard, Brien, and Waite (1995). In their study a higher risk of premarital cohabitation ceased once selection effects were accounted for. In this study the higher risk of premarital cohabitation switched to a lower risk once observed and unobserved selection effects were accounted for – premarital cohabitation actually *decreased* the risk of marital separation.

Our analysis of selection effects revealed how unobserved selection into marriage operated through cohabitation. Women with above-average risk of separation selected themselves into cohabitation; among cohabiters, women with lower risks of separation were more likely to marry. However, the latter were still more prone to divorce than those women who married directly. Another interesting issue is the effect of cohabitation, which decreased significantly across the models, particularly after having controlled for unobserved selection effects. Although cohabiters had a higher risk of union dissolution than those who were married, observed characteristics and unobserved selection effects accounted for many of the initial differences in the risk of union disruption between women who cohabited and those who were married.

6. Discussion

The aim of our research was to examine whether premarital cohabitation influences subsequent marital dissolution, controlling for measured characteristics of women and their partners and unmeasured selection effects. We used data from Austria. Our initial analysis showed that those who cohabited prior to marriage had a higher risk of marital dissolution. However, once observed characteristics of women and unobserved selection effects were properly controlled for, the risks of marital dissolution for those who cohabited prior to marriage were significantly lower than for those who married directly. Premarital cohabitation – net of self-selection – actually *decreased* the risk of separation. Thus it would appear that the “trial marriage” theory may indeed be relevant, with premarital cohabitation providing information that allows for a more

[4] In further analyses we excluded the variables on fertility and residential changes, which were potentially endogenous in the union-dissolution process. The results did not change (not shown). We also modelled migration and divorce jointly in one of our previous papers (see Boyle at al. 2008); the results showed that women with unmeasured “risky” characteristics were not over- or underrepresented among the migrants.

precise estimate of the match quality with the prospective spouse. As a consequence, marriages that involve prior cohabitation are more stable than direct marriages.

The study thus supported the notion of “trial marriage”. The results are similar to those of Brüderl, Diekmann, and Engelhardt (1997) in West Germany and Svarer (2004) in Denmark; in both studies unobserved selection effects were explicitly addressed. In Lillard, Brien, and Waite’s (1995) study in the US a higher risk of premarital cohabitation disappeared once selection effects were accounted for. Furthermore, the risks of marital dissolution for those who cohabited prior to marriage were lower than for those who married directly. However, the differences between the two groups were not significant. One reason for some differences between the studies might be that the effect of premarital cohabitation – net of selection effects – varies across countries. Other explanations might be related to modelling strategies. While Lillard, Brien, and Waite (1995) controlled for unmeasured selection into premarital cohabitation among those who were married, we explicitly modelled all partnership transitions of women and corresponding selection effects. We used hazard regression in our study of partnership-formation processes in order to exploit all the information that was available, particularly the time-varying covariates. The aim was to identify unobservables, having controlled properly for the observed variables. Our strategy might provide a better measurement of selection effects. Nevertheless, the results from these four studies demonstrate the importance of explicitly addressing unobserved selectivity into cohabitation and marriage when considering the effect of premarital cohabitation on marital dissolution.

Future research may benefit from a detailed investigation of three topics. First, the effect of premarital cohabitation could be analysed by cohort. Studies suggest that the relationship between premarital cohabitation and marital divorce may depend on how prevalent cohabitation is in a society (Liefbroer and Dourleijn 2006). Among cohorts where cohabitation is rare, those who do cohabit may have a high risk of subsequent marital dissolution because they are a select group. Alternatively, when premarital cohabitation is common, cohabiters may have a high risk of subsequent marital dissolution because those who marry directly are a select group with a low risk of divorce (Hoem and Hoem 1992). There may be little difference in the initial dissolution risks between the two groups when about half of the cohort cohabits prior to marriage and the other half marries directly. We agree that selection effects may be stronger for some cohorts than for others, but are reassured that they exist for all cohorts, and once selection effects are controlled properly premarital cohabitation *decreases* the risk of separation, regardless of how common cohabitation is. Our further analysis by cohort provided some support for this argument, but the data set was too small for a detailed analysis and for firm conclusions to be drawn. Most importantly, this study reveals that strong selection effects exist in a population where some cohabit prior to marriage and some marry directly.

Another interesting line of investigation is the effect of premarital cohabitation on subsequent marital separation by union order (cf. Kiernan 1999; Reinhold 2010). Our further analysis suggested that premarital cohabitation – net of self-selection – *decreased* the risk of separation for both first and higher order unions. Again, the data set was too small for a detailed analysis. Finally, the essence of selection effects needs further clarification. Being "disruption-prone" may simply mean being more liberal than others, but it may also reflect some unmeasured personality traits. In further analysis we included a variable measuring how liberal or conservative a woman is. To construct the variable, we used several questions on opinions and values, and once included in the model with other observed characteristics women who cohabited prior to marriage had a lower risk of union dissolution than those who married directly (although the differences were not significant). However, as this "values" variable was measured at the time of the survey, it is endogenous in the union-dissolution equation; the experience of divorce by a woman might have shaped or even determined her values and opinions at the time of the survey or there might have been other unmeasured characteristics that shaped both her values and her risk of union dissolution.

Finally, our study shows that there are significant geographical variations in union dissolution within Austria. Separation rates were highest in the capital city, Vienna, and lowest in the rural areas. Given that we controlled for a range of observed and unobserved individual characteristics in our models, these results provide convincing support that residential context also influences the risk of union dissolution. Significant differences in union dissolution by place of residence would need more attention in the field of divorce studies (cf. Kulu and Boyle 2009).

7. Acknowledgements

We wish to thank two referees for very useful comments on a previous version of this paper. They encouraged us to conduct the analyses, which we might have not conducted otherwise. We also thank the Max Planck Institute for Demographic Research (MPIDR) for funding, which supported Paul Boyle's research stay at the Laboratory of Contemporary European Fertility and Family Dynamics of the MPIDR. This research was mostly conducted when Hill Kulu was at the Laboratory of Contemporary European Fertility and Family Dynamics of the MPIDR.

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