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Children and union dissolution across four decades: Evidence from Norway

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Abstract:

This study describes the association between having children and the risk of union disruption, and whether this association has changed over time. We expand upon previous research by including data on cohabiting as well as married couples, and by studying change over four decades. We use data from the Norwegian Gender and Generation Study (2007) (N = 14 892). Combining self-reported union histories with register data on fertility histories, we construct a data set of person months for all individuals at risk of union dissolution in the period 1970-2007. Results from the event history analysis confirm that couples with children have lower union dissolution risk. Union dissolution risk is lowest when children are young, and also varies by number of children. There is little change over historical time in the correlation between having children and the risk of union dissolution. However, the monthly risk of dissolving unions increases substantially over time among childless as well as among parental couples.

Keywords: Union dissolution, Fertility, Demographic trends

JEL classification: J11, J12, J13

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Address: Elina Vinberg, Statistics Norway, Research Department. E-mail: elina.vinberg@gmail.com

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Sammendrag

Denne studien undersøker betydningen av det å ha felles barn for samlivbrudd blant samboende og gifte par i Norge i perioden 1970-2007. Vi fokuserer på to spørsmål: Hva sammenhengen mellom det å ha barn og risiko for samlivsbrudd er, og hvordan denne sammenhengen eventuelt har endret seg over tid. Informasjon om samlivshistorier er hentet fra spørreundersøkelsen LOGG (N=14 892), og data om barnefødsler, utdanningsnivå og studentstatus er hentet fra administrative registre.

Sammenhengen mellom barn og bruddrisiko analyseres ved diskret tids hasardregresjon. Vi kontrollerer for alle tilgjengelige observerbare kjennetegn som kan konfundere sammenhengen: Samlivstype, samlivets varighet, antall tidligere samliv, alder ved samlivsinngåelse, utdanningsnivå og hvorvidt en er under utdanning. Resultater for perioden samlet viser at par med felles barn har signifikant lavere bruddrisiko enn barnløse par, men det er stor variasjon etter barnas alder og antall barn. Par med veldig små barn har lavest bruddrater, noe som peker mot at barn stabiliserer samliv på kort, men ikke på lang sikt.

Vi undersøker også hvorvidt sammenhengen mellom felles barn og samlivsbrudd har endret seg over tid. Selv om bruddratene økte markant fra 1970-2007 både blant par med og uten barn, var forskjellen gruppene imellom ganske stabil over tid. Bruddraten økte mer blant par med eldre barn og mindre blant par med yngre barn.

1. Introduction

Common children are usually seen as a prime source of cohesion in a co-residential relationship, and thus expected to contribute to the prevention of divorce (Waite and Lillard, 1991). A contrasting view has been taken by those who emphasize that becoming a parent and raising children can lower relationship satisfaction (Feeny et al., 2003; Keizer and Schenck, 2012; Twenge et al., 2003), potentially leading couples to divorce. In the literature on antecedents of union dissolution, the majority of studies reports lower dissolution risk among parental couples (Lyngstad and Jalovaara, 2010), while a small group of studies have found that the presence of common children actually increases dissolution risk (Chan and Halpin, 2002a, 2002b; Svarer and Verner, 2008).

Over the past decades, massive structural and cultural changes in the family and the economy have taken place throughout the Western world. These changes have great potential for profound changes in how childbearing is related to union stability. Union dissolution rates have surged (Lesthaeghe, 2010), and cohabitation has arrived as an accepted and increasingly regulated family form (Perelli-Harris and Gassen, 2012). As women increase their labour force participation, economic dependency will to a lesser extent keep couples together (Becker, 1991). The availability of economic support for single mothers, which has increased strongly in some countries, further reduces mother's dependency on their spouses (Tjøtta and Vaage, 2008). However, as union dissolution rates increase also among childless couples, the *difference* in dissolution rates between parental and childless couples may have remained unchanged – or even narrowed. For example, while sanctions against individuals breaking up their relationships have weakened markedly, norms prevail against dissolving a union when there is children involved (Chan and Halpin, 2002b; Liefbroer and Billari, 2010).

In spite of the increasing dissolution rates, there is still relatively little research on whether the determinants of dissolution have varied over historical time (Teachman, 2002). Previous research on the potentially changing role of children in union dissolutions has focused on the compatibility of cohabitation and childbearing by comparing the experiences of specific birth cohorts (Steele et al., 2006). Although very useful and admirably executed, such endeavours do not provide a birds-eye view of dissolution rates of the parous and childless from a period perspective. This gap in the literature may be important for the understanding of the rise in union dissolution rates, as other studies have indicated that at least changes in divorce rates (and potentially also dissolution rates of cohabitation) are chiefly a period phenomenon (Lutz et al., 1991; Lyngstad and Jalovaara, 2010). That is, period changes in structural or cultural factors lead to subsequent changes in dissolution risk for all cohorts. This study aims at filling this gap in the literature by investigating whether the relationship between children and union dissolution has changed over the period 1970-2007.

Norway has in several aspects been among the forerunners in family change (Kitterød and Rønsen, 2014). The country has gone from a being male-breadwinner to a dual-earner society across the study period (Kitterød and Rønsen, 2013a). We combine self-reported union histories from the Norwegian GGS (N study sample = 12 249 persons/3 028 987 person-months) with fertility histories drawn from administrative registers. Our data is particularly well suited to study change over time, partly because of its larger-than-usual sample size. As union dissolution was rather infrequent in earlier decades, a higher sample size is necessary for precise estimation of dissolution rates.

Several earlier studies have used very large samples drawn from registers (e.g. Andersson, 1995, 1997; Hoem, 1991), but these studies did not cover informal cohabitation. Some decades ago it was still very meaningful to study only divorce, and not include dissolution of cohabitations. Marriages made up the lion's share of unions, and most children were born to married parents. Cohabiting parental couples often married following a birth. Even if a study of divorce covered a subset of the population in coresidential unions, it was possible to argue that this subset of families at least represented the prime relationship context for childbearing.

This state of affairs has decidedly come to an end. It is, at least in the Nordic countries, not at all recommended to consider only married couples when studying the dissolution of coresidential unions. The increase in cohabitation and births outside of marriage is an important part of the change in family patterns over the past decades, and a range of studies find that there are differences in relationship satisfaction and dissolution rates between married and cohabiting unions (Lyngstad and Jalovaara, 2010; Wiik et al., 2009), although several studies also conclude that some of these differences are due to selection mechanisms (Lillard et al., 1995; Svarer 2004). Even though studies of dissolution risk that make use of register data also on cohabiting couples have appeared the last few years, these remain a minority, and are able to cover shorter spans of historical time (Jalovaara 2013; Maenpää and Jalovaara 2014). We will include both married and cohabiting relationships in our analysis, and consistently use the term partnership to cover both marriage and cohabitation. This eases comparison over historical time, as the changing selectivity into marriage will not confound the estimated relationship between dissolution risk and joint childbearing history.

2. Theoretical perspectives and previous research

Children are often considered to be the prime example of "marital-specific capital" (Becker et al., 1977) and both partners are expected to have more to lose after a union dissolution if they have common children: The parent who retains custody (usually the mother) typically experiences a decline in economic well-being upon union dissolution, as childrearing limits the time for market work and running a one-income household is costly. Single mothers are also less likely to re-partner than childless women (Becker, 1991: 330). The risk of lost or restricted access to children and loss of control over the children's well-being might prevent fathers (usually the non-custodial parent) from dissolving unions (Chan and Halpin, 2002b; Kalmijn, 1999). Parents may also stay together because they expect union dissolution to have negative consequences for their children (deGraaf and Kalmijn, 2006) and other people's sanctions following dissolution are expected to be stronger for parental unions - preventing even dissatisfied couples with children from breaking up (Thornton, 1977). In short, the expected costs of union dissolution are expected to be higher when couples have children.

Entering parenthood is among the most challenging and stressful life transition a couple faces (Feeney et al., 2001): It implies dealing with new social roles, reduces individual freedom and increases duties and routine work (Twenge et al., 2003). Childrearing limits "couple time" (Waite and Lillard, 1991) and may also interfere with the parents' sex lives. For these reasons, childbearing may lower relationship satisfaction (Keizer and Schenck, 2012), increasing the risk of dissolution.

Most previous studies find that couples with children have significant lower union dissolution rates than childless couples (Andersson, 1997; Erlagsen and Andersson, 2001; Kravdal, 1988; Vuri, 2003; Waite and Lillard, 1991). However, as more stable and committed couples are more likely to have children, these findings may also be driven by selection (Lillard and Waite, 1993). Couples with young children may be a particularly selected group of satisfied and happy partners, as they recently were sufficiently satisfied with their relationship to decide to have a(nother) child together. Additionally, being unable to have a child may destabilize a union because many couples regard having children as an important part of a relationship (Hoem 1997). There are, however, also some studies pointing in the opposite direction, finding a higher risk of union dissolution for couples with children (Chan and Halpin, 2002a, 2002b; Svarer and Verner, 2008). The variations in findings might reflect different operationalization of the child-related indicators (the age and number of children), or variation in other factors such as definition of partnership type, cohort composition of the sample or variation between countries (Steele et al., 2006).

Change over time in Norway

As the association between having children and union stability likely depends on the social context couples are situated within, mechanisms linking children to union dissolution may strengthen or weaken over historical time. Because changes in union stability of *childless* couples also affects how the differences in dissolution rates between couples with and without children develops, we include theoretical perspectives on possible changes in both groups.

Women's increased economic independence is the most striking relevant change in our period of study. In the 1960s, Norway was among the countries in Europe with the lowest proportion of women in paid work, but has currently one of the highest proportions (Kitterød and Rønsen, 2012). The use of child care for children of preschool ages has increased from 3% in 1972 to 97% in 2010 (Kitterød and Rønsen, 2012). Together with increased economic support for single parents, this strengthens the economic prospects of mothers after union dissolution (Kavli et al., 2010, Rijken and Liefbroer, 2012). If women have stayed together with the father of their children due to economic constraints, union dissolution rates of parental couples is expected to increase over time. However, while the share of highly specialized childless couples is minuscule in the later part of the period (Cools, 2012), parents (at least of young children) tend to have partial specialization throughout the period (Kitterød and Rønsen, 2013b). Hence, while the (potentially stabilizing) specialization has weakened somewhat among parental couples, it has decreased to a much larger extent among childless couples.

Studies from several Western countries have shown a growth in fathers' physical and emotional care of children over the past decades (Hook, 2006; Hook and Wolfe, 2012; Kalmijn, 1999), particularly in Scandinavian countries (Brandth and Kvande, 2013; Kitterød and Rønsen, 2013a). More involved fathering strengthens ties between father and child, and can affect the stability of parental couples' in two opposing ways: By spending time with their children, fathers may prevent loss of contact if the union is dissolved (Aarseth, 2011: 73), lowering their cost of a union dissolution. On the other hand, as mothers remain more likely to get sole custody of the children (Kitterød and Lyngstad, 2012) fathers who develop strong emotional ties to their children may be more reluctant to leave their partner (Kalmijn, 1999).

Social norms against union dissolution have weakened over the period of study (Thornton and Young-DeMarco, 2001). A parental couple who stayed together early in the period of study, partly or wholly motivated by a belief of potential welfare benefits for their children, may later have chosen to dissolve their union. However, norms against union dissolution remain substantially stronger for parental couples than for couples without children. Using Dutch data, Liefbroer and Billari (2010) found that 9% disapproved of divorce if no children were present, whereas 44% disapproved when the divorcees had young children. Similar findings are reported from the UK (Chan and Halpin, 2002b). If social norms and sanctions prevent people from dissolving unions, dissolution rates may have increased less for parental couples than for childless couples, giving a relatively more stabilizing effect of children over the period.

Social theorists have argued that so-called "late modern relationships" are expected to be "pure", based on mutual self-disclosure and appreciation of each other's unique qualities (e.g. Giddens, 1992: 49-64). Particularly for parental couples, ideals based of intimacy, equality and mutuality may be difficult to attain in the presence of structural gender inequalities (Jamieson, 1999: 486). The increasing instability of heterosexual couples has been suggested to be a consequence of such a tension (Jamieson 1999): When the division of labour is perceived as unfair, the risk of union dissolution increases (Greenstein, 2009; Cooke, 2006). As parental couples increasingly struggle to live up to the ideals of a "pure" relationship, the increase in dissolution rates may be more rapid for couples with children than for childless couples.

In sum, as (most) stabilizing mechanisms are weakened and most destabilizing mechanisms are strengthened, we expect that the probability of union dissolution among parental couples will increase over historical time. On the other hand, with more non-obligating unions spreading, few factors should prevent dissatisfied childless couples from breaking up. This would imply that the *difference* in dissolution risk between couples with and without children will increase (or remained unchanged) over the observation period, making for an increased (or constant) stabilizing effect of children across time.

3. Data & Methods

Data

To study union dissolution patterns over the period 1970-2007, we use data from the nationally representative Norwegian Gender and Generations survey (Lappegård and Veenstra, 2010). The survey had a 60% response rate, and was supplemented with individual-level information from administrative registers. The linking to registry data generated no sample attrition. Most importantly for our purposes, both educational level and fertility histories are taken from registers. The Norwegian GGS is well suited for analysing the relationship between children and union dissolution for several reasons. A comparatively large sample size (N=14,892) and respondents from a wide range of birth cohorts (1927–1988) facilitates comparisons of union dissolution rates across historical time. In the early part of our study period, the 1970s, divorce rates were increasing but still rather low, requiring large samples for precise estimation.

Compared to Norwegian register data, the GGS has an advantage of identifying childless cohabiting couples in an unproblematic way.

The analysis sample consists of respondents who have ever entered a union (N=12,249). Partnership histories were collected retrospectively by asking all respondents to recall the start and (if relevant) end dates of all co-residential relationships (cohabitation and marriage). Data from official registers of any change in civil status were used as a supplement to the survey dataset. Information about the respondent's own children was taken from official registers and then supplemented with self-reporting, making for very high reliability. In addition, respondents were also asked if the partner had any children from a previous relationship. Nearly all union histories contained monthly updated information about when the couple moved in together, when (if so) they got married and when (if so) their children were born.[1]

Statistical approach

Our analysis consisted of a set of discrete-time hazard regression models (Allison, 1995). This standard method for union dissolution research implies that each union record in the data set is split up into a number of observations of union-months, started at the month of union entry (restricted to observations over 15 years) and ended at whatever did come first of union dissolution, or to censoring due to the death of one of the partners or end of the observation period (i.e. December 2007). The dependent variable in the statistical models is the conditional probability (measured in log-odds) of a couple's union dissolution in month t, given that the couple was at risk at the start of month t.

[1] If information of union entry n+1 were missing, it was set to two months after union n was dissolved. If information on the entry of first union was missing, this union was deleted. Missing months of entry were set to June. If union n+1 was entered before union n was dissolved, union n was recordeded as dissolved to two months before union n+1 was entered.

The explanatory variable, (common) children, was constructed by combining the number of children and the age of the youngest child, which is a standard approach in the literature (Lyngstad 2004). If the date of birth of a child lied within the interval of entry and dissolution/censoring of a union, the child was considered a common to that couple. The age of the youngest child was categorized into three different groups: 0–2 years, 3–6 years, and 7 years or more. Parity was categorized as childless, one child, two children, and three or more children. This led to a set of ten categories (childless, one child, youngest aged 0–2; one child, youngest aged 3–6; one child, youngest aged 7+; two children, youngest age 0-2; and so forth). Childless couples serves as the reference category in the models, while the other categories are represented by nine dummy variables.

Table 1. Descriptive statistics of analysis variables, person months.

	1970-07		1970-85		1986-96		1997-07	
	Freq.	%	Freq.	%	Freq.	%	Freq.	%
Union dissolution								
No	3023692	100	784962	99	929844	99	1308886	99
Yes	5295	0	923	0	1790	0	2582	0
No. and age of ch.								
No children	777316	26	187610	23	246376	26	343330	26
1 Ch, <2 y	188846	6	69322	8	56747	6	62777	4
1 Ch, 2-6 y	181257	6	66848	8	53377	5	61032	4
1 Ch, >6y	195187	6	35881	4	63996	6	95310	7
2 Ch, youngest <2 y	147441	5	52768	6	43543	4	51130	3
2 Ch, youngest 2-6y	260722	9	96709	12	72523	7	91490	6
2 Ch, youngest >6y	568728	19	99256	12	182845	19	286627	21
> 2 Ch, youngest <2y	79368	3	29004	3	23586	2	26778	2
> 2 Ch, youngest 2-6y	167357	6	62670	7	45359	4	59328	4
> 2 Ch, youngest >6y	462765	15	85817	10	143282	15	233666	17
Step children								
No step child	2473734	82	692804	88	769600	82	1011330	77
Stepchild(ren)	555253	18	93081	11	162034	17	300138	22
Union type								
Cohabiting	613905	20	82028	10	202082	21	329795	25
Married	2415082	80	703857	89	729552	78	981673	74
Union number								
1	2580498	85	745337	94	810881	87	1024280	78
2	400643	13	38406	4	110434	11	251803	19
>2	47846	2	2142	0	10319	1	35385	2
Age at union entry								
< 20	435498	14	138364	17	143192	15	153942	11
20-24	1415798	47	408599	51	444163	47	563036	42
25-34	943642	31	217628	27	281549	30	444465	33
> 34	234049	8	21294	2	62730	6	150025	11
Educ. attainment								
Compulsory	684307	23	215628	27	225724	24	242955	18
Upper Secondary	1485240	49	405139	51	460168	49	619933	47
University, low	648849	21	121430	15	189460	20	337959	25
University,high	182783	6	32438	4	49265	5	101080	7
Information missing	27808	1	11250	1	7017	0	9541	0
Educ. enrollment								
Not enrolled	2742949	91	631560	80	880070	94	1231319	93
Enrolled	157929	5	26216	3	51564	5	80149	6
Information missing	128109	4	128109	16	.		.	
Sex								
Man	1473813	49	374129	47	447810	48	651874	49
Woman	1555174	51	411756	52	483824	51	659594	50
Birth cohort								
1927-34	536039	18	234066	29	156588	16	145385	11
1940-49	853864	28	335588	42	255484	27	262792	20
1950-59	787890	26	195723	24	284684	30	307483	23
1960-69	563412	19	20508	2	205118	22	337786	25
1970-79	257412	9	.		29760	3	227652	17
1980-88	30370	1	.		.		30370	2
	Mean	**(SD)**	**Mean**	**(SD)**	**Mean**	**(SD)**	**Mean**	**(SD)**
Duration (years)	15,1	(11,6)	10,2	(7,5)	14,5	(10,4)	18,4	(13,3)
N	3028987		785885		931634		1311468	

The models control for several potential confounders previously found in the literature of union dissolution and children (Lillard and Waite, 1991; Lyngstad and Jalovaara, 2010). First, we included a dummy variable for children born prior to union entry, capturing any children the respondent (or his/her partner) may have had in an earlier relationship.[2] This variable was coded 1 if either the respondent or the respondent's partner had children from a previous relationship. We also controlled for relationship duration, measured from date of union formation and updated monthly. Due to evidence of a nonlinear association between relationship duration and union dissolution (Kulu 2014), a quadratic polynomial was included. To facilitate interpretation, the duration variable was divided by 12, counting duration in years instead of months. Next, we included monthly updated information as to whether the couple was currently married (coded 1) or cohabiting (coded 0).

A control for union order was included to control for potential confounders for higher order unions. First unions constituted the reference group, and we included dummy variables for one previous union, and two or more previous unions. The variable measuring respondent's age at union formation was grouped into four broad age categories (< 20, 20–24, 25–34, 35+). Register data on the respondent's highest completed educational attainment was included as a time-varying covariate, converted into a set of dummy variables that corresponds to basic education (primary and lower secondary) (0–9/10 years of schooling)[3], high school (upper secondary) (10/11–13 years), university low (equals to a bachelor degree, 3–4 years), and university high (master's degree or PhD, 5–8 years). Information on education is included as it is found to be a key determinant of union dissolution (Hoem, 1991; Lyngstad, 2004). Moreover, the distribution of education has changed over cohorts, and its role in dissolution processes may have changed over time (Härkönen and Dronkers, 2006). We also included a dummy variable for educational

[2] We do not distinguish between the two partners' children because we do not have any residential or custody information of their children born prior to union entry.

[3] From 1997, the mandatory «grunnskole» was expanded from 9 to 10 years.

enrolment, based on yearly updated information. Since data on student status was only available from 1974 onward, all exposures before 1974 were coded as missing.

Ideally, one would want to measure both spouses' characteristics in studies of union dissolution (Lyngstad and Jalovaara, 2010). Using retrospective survey data, our analysis is limited to information reported by the survey respondent. For previous partners, we know only whether they brought (presumably own) children into the household. A conventional approach is then to estimate models separately by respondent sex, in order to avoid conflating different (and potentially opposite) coefficients of covariates related to the dissolution rate. As we are chiefly concerned with historical change and want to preserve statistical power, we take a different approach. We pool data on male and female respondents, including a covariate for the respondents' sex, and test for sex-specific effects of covariates wherever theoretically plausible or previously documented in the scientific literature.

Previous research suggests that the period effect dominates the cohort-effect when estimating change (over historical time) in union dissolution and divorce rates (Andersson, 1997; Kulu, 2014; Lutz et al., 1991). Similarly, the development in women's labour-force participation has been found to be period-driven rather than cohort-driven (Koren, 2012: 29). We therefore take a period approach to estimating change over historical time. Balancing the need for statistical power (requiring larger subsamples) and detailed information on period change (requiring a larger number of subsamples), we measure children's role in disruption risk in three different time periods (1970-1985, 1986-2006, 1997-2007). The cut points were chosen to aquire subsamples of equal size, giving comparable test strength across subsamples.

4. Results

Our analysis proceeded in three main steps. First, we estimated a discrete-time event history regression for the whole study period. Second, we estimated the same model separately for the periods 1970-1985, 1986-1996, and 1997-2007, and these models provides the bulk of our empirical results. Finally, we report the outcomes of several sensitivity analyses done to investigate, and hopefully rule out several potential problems of our analysis.

Table 2. Model 1. Results from discrete-time logistic model of union dissolution risk. 1970–2007

Variable	OR	(95% CI)		
Intercept	**0,005**	(0,004	–	0,005)
Joint children (ref= none)				
1, 0–2 yrs	**0,530**	(0,471	–	0,596)
1, 3–6 yrs	**0,830**	(0,745	–	0,924)
1, 7+ yrs	0,921	(0,793	–	1,071)
2, 0–2 yrs	**0,343**	(0,287	–	0,409)
2, 3–6 yrs	**0,545**	(0,480	–	0,618)
2, 7+ yrs	**0,682**	(0,596	–	0,780)
3+, 0–2 yrs	**0,215**	(0,156	–	0,296)
3+, 3–6 yrs	**0,336**	(0,275	–	0,410)
3+, 7+ yrs	**0,566**	(0,479	–	0,667)
Step children	**0,618**	(0,564	–	0,676)
Union type (ref = cohabiting)				
Married	**0,332**	(0,309	–	0,357)
Union number (ref=1st)				
2nd	**1,149**	(1,052	–	1,255)
3rd +	**1,711**	(1,442	–	2,029)
Duration	**1,056**	(1,042	–	1,069)
Dur. squared	**0,997**	(0,997	–	0,998)
Age at union formation (ref = 20-24)				
<20	**1,465**	(1,360	–	1,578)
25–34	**0,846**	(0,788	–	0,909)
>34	**0,644**	(0,565	–	0,733)
Education (ref= Upper secondary)				
Compulsory	**1,108**	(1,034	–	1,186)
University, low	1,024	(0,954	–	1,100)
Univerity, high	0,981	(0,865	–	1,112)
Information missing	**1,521**	(1,221	–	1,893)
Educational enrolment (ref=no)				
Yes	**1,507**	(1,389	–	1,636)
Information missing	**0,431**	(0,345	–	0,538)
Sex (ref = man)				
Woman	1,046	(0,988	–	1,108)
-2 LL With covariates (intercept only)	72935,8	(77819,0)		
ChiSq (DF)	**4883,2**	(25)		
N	3028987			
Dissolutions	5295			

Note: Estimates in bold are significant at the 95% level.

Table 2 show results from Model 1 where the outcome is the risk of union dissolution in a given month over the whole period of study (i.e. 1970-2007). The overall pattern is a negative association between having common children and a couple's monthly risk of splitting up, net of the other covariates. There are, however, quite large variations in dissolution risk by the number of children and the age of the youngest child. In line with the most common finding in the literature, young children are correlated with a lower rate of dissolution (Andersson, 1997; Cherlin, 1977; Erlangsen and Andersson, 2001; Lillard and Waite, 1993; Waite and Lillard, 1991). There is also a separate and negative parity effect, where dissolution risk decreases significantly with parity, echoing the consensus from the earlier Scandinavian literature (Andersson, 1997; Jalovaara, 2001; Lyngstad, 2004). The direction of the parameter estimates for the control variables are largely as expected from previous research (Lyngstad and Jalovaara, 2010).

To assess potential changes over historical time, we divided the data into three segments covering three different periods and reestimated Model 1 separately for these periods.[4] Model 2a covers the period 1970–1985, model 2b the period 1986–1996, and model 2c the period 1997–2007. As splitting the sample reduces statistical power markedly, we also mark up and comment upon results significant at the 10% level in these models.

[4] When comparing estimates across different logistic regression models, one should be aware that large reductions in unexplained variance in itself will lead to stronger estimates (logits further away from 0) (Allison, 1999; Mood, 2010). Our models display the same pattern when expressed as marginal effects, which are unaffected by changes in explained variance. Changes in variance across models are thus unlikely to drive our results.

Table 3. Model 2a-c. Results from discrete-time logistic models of union dissolution risk. Three sub-periods

Sub-period	**1970-1985**			**1986-1996**			**1997-2007**		
Variable	**OR**	**(95% CI)**		**OR**	**(95% CI)**		**OR**	**(95% CI)**	
Intercept	**0,003**	(0,002	-0,003)	**0,004**	(0,004	-0,005)	**0,007**	(0,007	-0,008)
Joint children (ref=none)									
1, 0–2 yrs	**0,670**	(0,512	-0,878)	**0,518**	(0,422	-0,636)	**0,479**	(0,402	-0,569)
1, 3–6 yrs	1,056	(0,838	-1,332)	**0,801**	(0,662	-0,970)	**0,719**	(0,610	-0,847)
1, 7+ yrs	**0,751**	(0,523	-1,078)	0,889	(0,689	-1,146)	1,028	(0,825	-1,281)
2, 0–2 yrs	**0,369**	(0,251	-0,542)	**0,411**	(0,309	-0,546)	**0,265**	(0,200	-0,351)
2, 3–6 yrs	**0,400**	(0,298	-0,537)	**0,543**	(0,433	-0,682)	**0,601**	(0,503	-0,719)
2, 7+ yrs	**0,533**	(0,392	-0,725)	**0,619**	(0,491	-0,781)	**0,804**	(0,659	-0,979)
3+, 0–2 yrs	**0,231**	(0,127	-0,418)	**0,215**	(0,125	-0,370)	**0,179**	(0,105	-0,305)
3+, 3–6 yrs	**0,240**	(0,157	-0,367)	**0,281**	(0,192	-0,411)	**0,425**	(0,320	-0,564)
3+, 7+ yrs	**0,488**	(0,337	-0,706)	**0,590**	(0,444	-0,783)	**0,612**	(0,475	-0,788)
Step children	0,836	(0,664	-1,040)	**0,591**	(0,503	-0,696)	**0,570**	(0,503	-0,646)
Union type (ref=coh.)									
Married	**0,332**	(0,273	-0,403)	**0,367**	(0,320	-0,421)	**0,374**	(0,337	-0,416)
Union number (ref=1^{st})									
2nd	**1,426**	(1,065	-1,909)	1,166	(0,990	-1,374)	**1,149**	(1,052	-1,255)
3rd +	**1,490**	(0,546	-4,069)	**1,521**	(1,038	-2,228)	**1,711**	(1,442	-2,029)
Duration	**1,149**	(1,104	-1,195)	**1,076**	(1,051	-1,101)	**1,056**	(1,042	-1,069)
Duration squared	**0,995**	(0,993	-0,996)	**0,997**	(0,996	-0,997)	**0,997**	(0,997	-0,998)
Age at union formation (ref=20-24)									
	<20	**1,548**	(1,313	-1,824)	**1,578**	(1,397	-1,783)	**1,465**	(1,360
25–34	**0,704**	(0,589	-0,841)	**0,842**	(0,742	-0,955)	**0,846**	(0,788	-0,909)
>34	**0,425**	(0,261	-0,693)	**0,549**	(0,427	-0,705)	**0,644**	(0,565	-0,733)
Education (ref= Upper secondary)									
Compulsory	1,099	(0,940	-1,285)	1,102	(0,983	-1,237)	**1,108**	(1,034	-1,186)
University, low	**1,339**	(1,113	-1,611)	1,000	(0,881	-1,137)	1,024	(0,954	-1,100)
Univerity, high	1,351	(0,959	-1,902)	1,112	(0,886	-1,395)	0,981	(0,865	-1,112)
Information missing	1,500	(0,964	-2,333)	**1,977**	(1,360	-2,873)	**1,521**	(1,221	-1,893)
Educational enrollment (ref=no)									
Yes	**1,828**	(1,458	-2,292)	**1,536**	(1,331	-1,772)	**1,507**	(1,389	-1,636)
Information missing	**0,611**	(0,484	-0,772)				**0,431**	(0,345	-0,538)
Sex (ref= man)									
Woman	1,104	(0,963	-1,27)	0,967	(0,876	-1,068)	1,046	0,988	-1,108)
-2 LL Interc. (w/cov.)	14300	(13686		25968	(24538		37332	(34414	
ChiSq (DF)	**613,4**	(25)		**1430,4**	(24)		**2918,2**	(24)	
N (dissolutions)	78588	(923)		931634	(1790)		1311468	(2582)	

Note: Estimates in bold are significant at the 95% level.

The models 2a-2c (Table 3) reveal that the relationships of parental couples are more stable than those of childless couples throughout the observation period. We find surprisingly little change in the general pattern across sub-periods. The patterns outlined above, that dissolution risks are lower for the parous and those with very young children, can be observed in for all three periods.

There are a few exceptions to this rather comprehensive constancy. For those with two or more common children in the two older age groups (3–6 and 7+), the negative coefficients are strongest in the first sub-period (1970-1985), indicating that the dissolution risk of couples with older children increases over time. The dissolution risk of couples with one child aged 7+ no longer differs significantly from that of childless couples in the latest sub-period (i.e. 1997-2007). Couples with very young children (age 0–2 years) display an opposite pattern of change over time. As compared to childless couples, having a child younger than two years old is correlated with the strongest reduction in dissolution risk in the last sub-period. This pattern of change over time is consistent across parities, but not statistically significant (Table 3).

Consistent with previous research on variations in union dissolution factors over historical time, the parameter estimates for the control variables presented in Table 3 display little change over time (Teachman, 2002; de Graaf and Kalmijn, 2006). First, being married is correlated with a substantially lower dissolution risk in all three periods, confirming that union type is a strong predictor of union disruption throughout the observation period. Dissolution risk is significantly higher among respondents in higher-order unions, and highest among those who have lived with more than one partner before. Having a step child is associated with lower dissolution risk. This counterintuitive result could be due to the inclusion of non-resident stepchildren, who presumably matter little for union stability, but who may proxy stabilizing factors such as (non-linearities in) duration. Being a student increases the likelihood of disruption in all sub-periods, in line with previous studies (Lyngstad 2004), although the heightened risk seems to decline over time in relative terms.[5]

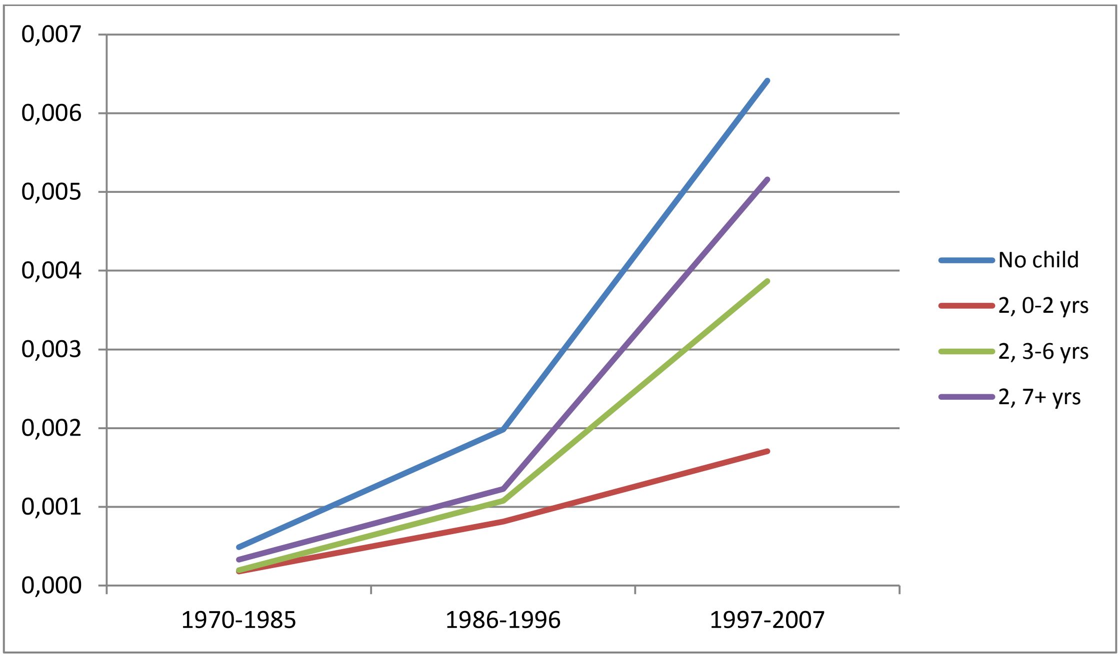
[5] As this variable takes missing values only up to 1974, results this variable from the earliest sub-period are not fully comparable with results for the two later periods.

Dissolution risks in parental unions across periods

The results presented in Table 3 reveal that the difference in dissolution risk between childless and parental couples has been relatively stable across historical time. To investigate whether parental couples have become increasingly likely to dissolve their union over time or not, we recalculate the estimates from Model 2 to predicted probabilities. Thus, we move from a description of relative dissolution risks (quantified by odds ratios), to a description of absolute dissolution risks (quantified by the absolute monthly risk of dissolving a union).

Figure 1 shows predicted probabilities of union dissolution setting duration to the median (11 years) for a union with the reference category values on the control variables; that is those cohabiting in their first union, secondary educated, and with union formation age of 20–24, without any stepchildren. Predicted probabilities are calculated for unions with two children, as the most common parity for the majority of cohorts in our study. For comparison, probabilities for childless couples in the three periods are also estimated.

Figure 1. Predicted probability of dissolution risk for childless couples and couples with two children in different age groups. 1970–2007



Note: Predicted probabilities are calculated based in the estimates shown in Table 3.

Figure 1 shows that the dissolution risks of parents with older children and childless couples increase in parallel over historical time. However, from the period 1985-1996 the dissolution rates have stayed quite stable for couples with children in the youngest age group. As the difference in dissolution risk between this group and childless couples increases, couples with very young children run a comparatively lower risk of dissolution over time.

Sensitivity analysis

We conducted several sensitivity tests on the set of period models to assess whether the results are robust to different model specifications and restrictions in the dataset. First, we considered potential bias stemming from longer average relationship durations in the first sub-period, which could indicate that couples in this period were selected on stability. If unions with above-average stability are pooled with more recent entrants with lower stability, change across historical time may be obscured. We reconfigured the sample, treating unions as censored once the union duration exceeded 20 years, and estimated the full model (i.e. Model 2a-c) on the reconfigured sample (Model 3a-c, Table 4). The coefficients of the reconfigured sample (Models 3a-c) are similar to those estimated from the full sample (Models 2a-c), indicating that variation in duration between the periods does not drive our findings.

Table 4. Model 3a-c. Results from logit models of union dissolution risk in three different time periods. Union durations of less than 20 years**.**

Sub-period	**1970-1985**			**1986-1996**			**1997-2007**		
Variable	**OR**	**95% CI**		**OR**	**95% CI**		**OR**	**95% CI**	
Intercept	**0,002**	(0,002	-0,002)	0,004	(0,003	-0,005)	**0,007**	(0,006	-0,008)
Joint children (ref=none)									
1, 0–2 yrs	**0,644**	(0,492	-0,843)	0,487	(0,396	-0,599)	**0,463**	(0,388	-0,551)
1, 3–6 yrs	0,966	(0,763	-1,223)	0,725	(0,596	-0,882)	**0,684**	(0,577	-0,810)
1, 7+ yrs	0,737	(0,502	-1,080)	1,007	(0,763	-1,328)	1,060	(0,817	-1,374)
2, 0–2 yrs	**0,338**	(0,229	-0,498)	0,364	(0,273	-0,487)	**0,252**	(0,190	-0,335)
2, 3–6 yrs	**0,368**	(0,274	-0,496)	0,507	(0,402	-0,641)	**0,595**	(0,494	-0,716)
2, 7+ yrs	**0,490**	(0,348	-0,690)	0,626	(0,468	-0,838)	**0,860**	(0,667	-1,110)
3+, 0–2 yrs	**0,218**	(0,120	-0,397)	0,179	(0,100	-0,321)	**0,182**	(0,107	-0,312)
3+, 3–6 yrs	**0,250**	(0,163	-0,383)	0,278	(0,185	-0,418)	**0,451**	(0,334	-0,611)
3+, 7+ yrs	**0,595**	(0,387	-0,913)	0,751	(0,499	-1,132)	0,670	(0,444	-1,011)
-2 LL									
Intercept	13365,52			22324,87			31136,25		
W/ cov.	12785,03			21233,05			29422,15		
ChiSq (DF)	**580,5**	(25)		**1091,8**	(24)		**1714,1**	(24)	
N	665483			699512			783678		
Dissolutions	869			1586			2276		

Note: Estimates in bold are significant at the 95% level.

In another set of models, birth cohort was included as a control variable in the different period models (see Appendix). The estimates were all significant at the 1% level with increasing divorce risk in younger cohorts, but without significantly changing the estimates for other covariates.

As mentioned in the data section, union dissolutions happen to couples, not individuals. Many of our analysis variables, notably educational attainment, age at union entry, and union number, are measured only for one of the two partners. In the regression models reported here, we ignore this complication. To make sure no important pattern is hidden, we checked if the estimates for age at union formation, union number, and educational attainment were different for men and women. We estimated the models with interaction terms between sex and these control variables. These models did not yield any significant interactions. This was the least surprising for age at union entry, given the very high correlation between partners' ages. For education this finding is also in line with previous results from the Scandinavian countries, showing that the net association between education and divorce is similar for husbands and wives (Jalovaara, 2003; Lyngstad, 2004). This indicates that the gender of the partner reporting the survey data does not affect our results. In sum the sensitivity analyses show that our results are robust to different model specifications and subgroup estimations.

5. Concluding discussion

Whether children stabilize or destabilize unions has been a recurrent question in sociodemographic research (Andersson, 1997; Lillard and Waite, 1993; Lyngstad and Jalovaara 2010). This study contributes to this literature by studying how the relationship between childbearing and union dissolution change across historical time in Norway. By taking a period perspective and analyzing data covering nearly four decades, we provide new insights on how social context shapes the role children play at different stages of their parent's relationship.

In the study period, increased female labour force participation and a stronger social security net made mothers less economically dependent on their partners, norms against union dissolution were relaxed (Thornton and Young-DeMarco, 2001), and relationship ideals changed (Giddens, 1992; Lesthaeghe, 2010). We expected these changes to increase the risk of union disruption among parents, and our results confirmed our expectation.

Our results demonstrate a rather stable negative association between having common children and dissolution risk when comparing parents to childless couples. This finding indicates that the various mechanisms that increase union instability across historical time affect both childless couples and parents in rather similar ways, which is in line with previous findings of change in dissolution determinants (Teachman, 2002; DeGraaf and Kalmijn, 2006). The universality of period changes in demographic behavior was succinctly summarized by Thornton and Rogers (1987: 20) who stated that “the degree of uniformity of the historical period effects across population subgroups has been remarkable”.

It is striking that couples with children aged two years or younger, as the only group in our study, display quite stable dissolution rates throughout the two last periods. Relative to childless couples, whose dissolution rates rose rapidly, young children increasingly stabilized unions across time. The mechanisms linking common children to union stability may remain important across historical time mainly for couples with relatively young children. For couples with children below two years, gender specialization is still very common (Kitterød and Rønsen, 2013b) and for parental couples who have split up, shared residence is not recommended before children are three years old (Stortinget 2009). In other words, fear of loss of contact as well as the fact that taking care of young children requires the joint efforts of two adults may make these couples reluctant to dissolve their unions. We cannot exclude that the selection of satisfied and stable couples with young children might have increased across the period due to better fertility control. However, as pregnancies among cohabitants are more often unplanned (Hayford and Guzzo, 2010; Kravdal, 1996) the increase in non-marital childbearing across the period could also drive the correlation in the opposite direction.

The surge in union dissolution throughout the Western world has been linked to several societal changes. Our findings indicate that these societal changes destabilize parental and childless unions alike. Though social norms against union dissolution arguably has been relaxed more for childless couples than for parental couples (Chan and Halpin, 2002b; Liefbroer and Billari, 2010), the change over time in the actual dissolution rates in these two groups has been strikingly similar. However, we can of course not exclude the possibility that different mechanisms drive a similar development in the two groups: Parental couples may struggle more to attain the prevalent relationship ideals, but simultaneously be held together by lack of time and money to a larger extent than childless couples.

The current study contributes to the literature on union dissolution in several ways. First, it provided new evidence on how the association between a coresiding couple's number and age of common children and their risk of union dissolution has developed over the course of four decades. During this period, massive social changes have taken place in the family, with the potential of upsetting the typical pattern of lower dissolution among the parous and parents of younger children. Our empirical analysis, based on high-quality survey and register data covering both married and cohabiting unions over a long time period, suggests that although the absolute probability of union dissolution among parental couples has increased substantially over time, the difference in union dissolution rates between childless and parents has remained relatively stable.

Future research in this field should try to replicate this study, to assess the generality of our findings and highlight any cross-national or cross-cultural variation. Furthermore, researchers should move beyond this design by overcoming one of its major limitations concerning the endogeneity of fertility and relationship dissolution. The selectivity of couples from childless to parous and beyond may have changed across historical time. If those who remain childless now are markedly different from those who remained childless in 1970, this may require a more involved explanation of our results.

Studies of actual behaviour and subjectively reported motivations for that behaviour are complimentary endeavours. While we have shown that the dissolution probability of parental and childless couples has changed in a similar fashion over time, it is very possible that the *motivation* for dissolving unions varies greatly between the two groups. To understand whether the motivation for dissolving unions varies between parental and childless couples, further research is called for.

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Appendix

Table 5. Model 4a-c. Results from discrete-time logistic models of union dissolution risk in three different time periods. Birth cohort included

Sub-period	**a. 1970-1985**			**b. 1986-1996**			**c. 1997-2007**		
Variable	**OR**		**(95% CI)**	**OR**		**(95% CI)**	**OR**		**(95% CI)**
Intercept	**0,002**	(0,002	-0,003)	**0,004**	(0,003	-0,005)	**0,007**	(0,005	-0,009)
Children (ref=none)									
1, 0–2 yrs	**0,663**	(0,507	-0,868)	**0,514**	(0,418	-0,631)	**0,483**	(0,406	-0,575)
1, 3–6 yrs	1,039	(0,824	-1,310)	**0,794**	(0,656	-0,962)	**0,719**	(0,609	-0,849)
1, 7+ yrs	0,730	(0,507	-1,050)	0,871	(0,675	-1,125)	0,963	(0,771	-1,203)
2, 0–2 yrs	**0,362**	(0,246	-0,533)	**0,405**	(0,305	-0,538)	**0,267**	(0,202	-0,354)
2, 3–6 yrs	**0,391**	(0,291	-0,526)	**0,532**	(0,423	-0,669)	**0,594**	(0,496	-0,713)
2, 7+ yrs	**0,519**	(0,380	-0,708)	**0,602**	(0,476	-0,762)	**0,735**	(0,601	-0,900)
3+, 0–2 yrs	**0,228**	(0,125	-0,413)	**0,210**	(0,122	-0,362)	**0,177**	(0,103	-0,302)
3+, 3–6 yrs	**0,241**	(0,157	-0,369)	**0,272**	(0,185	-0,399)	**0,407**	(0,306	-0,542)
3+, 7+ yrs	**0,490**	(0,338	-0,711)	**0,581**	(0,437	-0,773)	**0,554**	(0,428	-0,716)
Step children	0,831	(0,664	-1,040)	**0,600**	(0,509	-0,706)	**0,574**	(0,506	-0,651)
Union type (ref=coh)									
Married	**0,336**	(0,275	-0,409)	**0,372**	(0,324	-0,427)	**0,373**	(0,335	-0,415)
Union number (ref=1st)									
2nd	**1,412**	(1,054	-1,893)	1,162	(0,986	-1,369)	1,009	(0,902	-1,129)
3rd +	1,493	(0,545	-4,087)	**1,497**	(1,020	-2,198)	**1,452**	(1,189	-1,772)
Duration	**1,159**	(1,111	-1,208)	**1,084**	(1,056	-1,113)	1,015	(0,996	-1,034)
Duration squared	**0,995**	(0,993	-0,997)	**0,997**	(0,996	-0,998)	**0,998**	(0,998	-0,999)
Age at union formation (ref=20-24)									
<20		1,476	(1,240	-1,757)	1,509	(1,327	-1,714)	1,325	(1,175
25–34	0,772	(0,636	-0,938)	0,915	(0,796	-1,051)	0,900	(0,806	-1,004)
>34	0,584	(0,340	-1,003)	0,739	(0,542	-1,008)	0,836	(0,669	-1,044)
Education (ref=Upper secondary)									
Compulsory	1,106	(0,946	-1,294)	1,104	(0,984	-1,238)	1,199	(1,081	-1,330)
University, low	1,317	(1,095	-1,585)	0,995	(0,876	-1,131)	0,899	(0,815	-0,992)
Univerity, high	1,361	(0,966	-1,916)	1,112	(0,886	-1,395)	0,784	(0,662	-0,929)
Information missing	1,503	(0,966	-2,340)	1,961	(1,349	-2,851)	1,305	(0,925	-1,841)
Educ. enrolment (ref =no)									
Yes	1,842	(1,468	-2,309)	1,521	(1,317	-1,757)	1,264	(1,128	-1,418)
Information missing	0,692	(0,535	-0,896)	.			.		
Sex (ref=man)									
Woman	1,102	(0,960	-1,265)	0,962	(0,872	-1,062)	1,096	(1,010	-1,190)
Birth cohort (ref=1950-59)									
1927-34	0,638	(0,450	-0,905)	0,569	(0,404	-0,800)	0,406	(0,264	-0,625)
1940-49	0,919	(0,751	-1,124)	0,855	(0,711	-1,027)	0,728	(0,598	-0,886)
1960-69	1,050	(0,798	-1,383)	1,104	(0,942	-1,293)	0,972	(0,835	-1,130)
1970-79				1,235	(0,976	-1,563)	0,984	(0,809	-1,197)
1980-89							1,213	(0,943	-1,559)
-2 LL Interc. (w/cov.)	14299,76	(13679)	25968,47	(24526)	37332,38	(34384)			
ChiSq (DF)	**621,3**	(28)		**1442,9**	(28)		**2948,7**	(29)	
N	785885			931634			1311468		
Dissolutions	923			1790			2582		

Note: Estimates in bold are significant at the 95% level.

Statistics Norway

Postal address:
PO Box 8131 Dept
NO-0033 Oslo

Office address:
Akersveien 26, Oslo
Oterveien 23, Kongsvinger

E-mail: ssb@ssb.no
Internet: www.ssb.no
Telephone: + 47 62 88 50 00

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