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THE RELATIONSHIP BETWEEN PREMARITAL COHABITATION AND MARITAL STABILITY: Evidence for Swedish men born in 1936-1960

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Abstract

In recent years, the incidence of premarital cohabitation has increased dramatically in many countries of Western Europe and in the U.S.. In view of such developments, it is natural to raise issues related to the relationship between premarital cohabitation and the stability of subsequent marriages. The present paper examines such a relationship with reference to Swedish men born in 1936-1960. The data for the analyses come from the 1985 survey of Swedish males, which had about 3200 respondents. Multiplicative hazard models are used to estimate relative risks of marital dissolution and thereby test several hypotheses in connection with the effect of premarital cohabitation and other background factors, on the stability of subsequent marriages. Consistent with, but stronger than in earlier findings, the results show that previous cohabitors, compared to noncohabitors tend to be at much greater differential risk of dissolution at all durations of marriage. In addition, we found that age at marriage, marriage duration, and child(ren) are among the strongest determinants of the risk of marital dissolution.

Contents

* *In preparing this report, I have benefited much from comments by Professor Jan M. Hoem, for which I am most grateful.*

1. INTRODUCTION

During the last few decades, most countries, and in particular the industrialized western societies, have witnessed considerable social change which has affected every aspect of family life. Such change has been reflected in the various demographic indicators, and some of the largest effects have been in patterns of marriage, divorce and cohabitation. Among the most fundamental changes is that couples have been less influenced by traditional social norms and parental views on the choice of partner, on whether to live together before marriage, and so on. As a result, the nature and values of marriage have changed, and some of the features which used to distinguish legal from informal unions have become blurred.

In general, in most of the industrialized world, marriage rates have declined while the prevalence of cohabitation and divorce rates, have risen. Parallel to this trend, there has been a rise in social and economic independence of women; social norms recognizing the desire of women to have a greater degree of equality in matters of household responsibilities, social activities, and financial decisions. Consequently the roles of partners have become less differentiated in nature, or more flexibly allocated, so that the need - or desire- for one partner to provide for the other (as in the more traditional form of marriage), has diminished. More importantly, if the partnership does not live up to the expectations or hopes of either partner, it is socially and financially possible for couples to part and lead independent lives.

These factors, together with a decline in adherence to a religious view of marriage, may well have led many to a reexamination (reassesment) of making a promise of a life-long commitment. Some couples live together without wishing to marry, others intend a 'trial marriage', and still others, who are sure of their commitment, decide to live together before they marry.

Such changing attitudes to marriage have involved changing attitudes to divorce. If couples place less emphasis upon the aspect of the unconditional commitment of marriage, and are less willing to sacrifice some opportunities for individual growth and development, they may consider and accept divorce more readily. Cohabitation, too, may be an indirect recognition of the risk of a relationship failing, and of the emotional and financial penalties of a marriage ending in divorce, since the ending of a cohabiting union involves no formalities.

The above developments have been described in Haskey (1991), in tracing the trends in family dynamics in the different European countries. In view of such developments, it is natural to raise issues related to the relationship between premarital cohabitation and marriage as well as subsequent marital stability.

One way of viewing the subject of union dissolution is to consider two parallel systems of union formation and dissolution, one restricted to legal unions and concerned with marriage and divorce, and the other to cohabiting couples and concerned with the start and end of informal unions. In practice, this dichotomy is not particularly satisfactory for analytic purposes, since couples and individuals do not remain within one system; for instance, couples often progress from living together to marrying, and partners who have recently divorced often start cohabiting with new partners.

An appropriate strategy would therefore be to include the experience of premarital cohabitation as one among the many explanatory variables that are bound to affect the risk of marital disruption and confine the analysis solely to the dissolution of marital unions.

This strategy has been adopted by many scholars to examine the relationship between premarital cohabiattion and marital stability with reference to different countries; Sweden (Hoem and Hoem, 1992; Trussell et al., 1992; Bennet et al., 1988; Blanc, 1985; Trost, 1979), Norway (Blanc, 1985), Finland (Lutz, 1991), The Netherlands (Klijzing, 1992; Manting, 1992), Australia (Bracher et al., 1992), France (Leridon, 1990), Germany (Schneider, 1990), Canada (Balakrishnan et al., 1987; White, 1987; Trussell et al., 1989), The United Kingdom (Haskey, 1983, 1991) and the United States (Catlin et al., 1978; Newcomb and Bentler, 1980; Kitson et al. 1985; Booth and Johnson, 1988; Thornton, 1988; Teachman and Polonko, 1990; Teachman et al., 1991; Thomson and Colella, 1991; Tucker and O'Grady, 1991; Axinn and Thornton, 1992; DeMaris and Rao, 1992).

In most of the studies that give attention to the link between cohabitation and marital disruption, the hypotheses to be tested are formulated out of the belief that cohabitation is a temporary premarital phase. As a consequence, the incidence of cohabitation (and sometimes, the duration) is treated as a determinant in the process of dissolution of married couples

(Balakrishnan et al., 1987; Hoem and Hoem, 1992; Bennet et al., 1988). Few other studies follow a different formulation. Axinn and Thornton (1992) brought forward the idea that the experience of the period of cohabitation might actually change the individuals involved in their acceptance of divorce, while Manting (1992) among others, suggests that cohabitation can also be viewed as an alternative to singlehood or to marriage.

In almost all studies, the intuitive belief that premarital cohabitation improves the quality of a subsequent marriage has lacked empirical support. Of the few exceptions, White (1987) used the 1984 Canadian National Survey and found a positive effect of cohabitation on staying married, which remains when length of marriage and age at marriage are controlled, while Teachman and Polonko (1990) suggest that empirical differences in the risk of marriage dissolution between premarital cohabitors and non cohabitors in the US, can be explained by the greater time duration cohabitors have spent in the union.

Such conclusions were, however, only short-lived. Trussell et al (1989) argued that the results obtained (and conclusion drawn) in White (1987) were merely due to a methodological error and using alternative options, they found that nonmarital cohabitation had no effect on marital stability when marital cohort and age at marriage are controlled. Moreover, reinvestigations of U.S. data by DeMaris and Rao (1992), and findings for other areas such as Sweden (Hoem and Hoem 1992; Bennet et al. 1988) have proved that premarital cohabitation is associated with a greater hazard of dissolution even after counting the time spent in unmarried cohabitation as part of marital duration, thereby questioning the findings of Teachman and Polonko (1990).

In the present paper we use data from the 1985 Survey of Swedish males to investigate the relationship between the experience of premarital cohabitation and the stability of subsequent marriages. The basic issue to be addressed here is whether the risk of marital disruption is the same for all modes of entry into the married-life or is mode-specific. Apart from this we also examine whether the risk of marital dissolution varies across selected sociodemographic characteristics of men. We have adopted the type of analystic startegy discussed above within the framework of event history analysis, commonly known as intensity (hazard) regression.

The results build-on the bulk of previous findings; in which the experience of premarital cohabitation contributes to a higher risk of dissolution of subsequent marital unions. Further, it is demonstrated that the risk of marital disruption varies across background sociodemographic variables such as higher risks for marriages at younger ages and for those with no children.

In the next section two competing hypotheses with regard to the effect of premarital cohabitation on the risk of marital disruption are discussed at length. Description of the data set, variables, theoretical expectations, and the statistical method are given in Section 3, while Section 4 presents and discusses the empirical findings. The last section summarizes the contents of the paper.

2. COHABITATION & MARITAL STABILITY: THEORETICAL CONSIDERATIONS

2.1 The 'Weeding' Hypothesis

Intuitively, one could argue that men who marry after having lived with a partner in a nonmarital union for a while, ought to have a lower dissolution risk than comparable men who enter marriage directly (without preceding cohabitation) because the latter have taken less time to get to know their partner and her behaviour in situations important for daily family life (Hoem and Hoem, 1992). The less durable unions would be 'weeded-out' before marriage when the couple invests in a premarital trial period, while otherwise this weeding process is delayed into the early stages of the marriage. This is what is known as the 'weeding' hypothesis.

The 'weeding' hypothesis views cohabitation as a living arrangement playing the role of sorting-ground, a meeting place for potential partners (Hoem and Hoem, 1992; Klijzing, 1992). If it is successful, the partners may stick together and perhaps get married and/or have children. If not, then perhaps it will be with the next partner that life continues to be enjoyable. In view of this, the 'weeding' hypothesis predicts that cohabitation may be a healthy start for adult union life, and the longer one tests one's partner, the better the resulting relationship may be, at least if the partner stands up to the test. This is closely related to search behaviour, getting to know one's partner, exploring alternatives, and hence preparing oneself in the best possible way. According to this hypothesis, those who marry without a trial period of joint household are expected to have higher risk of divorce than those who cohabited before marriage.

2.2 The '(mode of entry) self selection' hypothesis

The growing acceptance of concensual unions as a possibly long-lasting predecessor of marriage, or in some cases even a replacement for it, has been explained as a reflection of ideational developments; changes in attitudes and norms or in tastes and performances, and possibly even in deeper life values in the general population (Lesthaeghe and Meekers, 1986). Such ideational changes have extended to family cohesion and have produced a higher and increasing incidence of dissolution for those who marry only after they have lived in a premarital cohabitation (Hoem and Hoem, 1992).

An alternative argument is therefore, that as part of the change in the mode of union formation, direct marriage must progressively have become a manifestation of a particular religious or other convictions (or of particular regional or other cultural influences). As a consequence, we observe lower and decreasing dissolution risks for this group (Hoem and Hoem, 1992; Klijing, 1992). All in all, the explanation is one of a progressive self-selection in the choice of mode of union formation and of consequential changes in norms concerning union-dissolution. This is what we call the '(mode of entry) self-selection' hypothesis.

The '(mode of entry) self selection' hypothesis views nonmarital cohabiting relationships as being selective of those who are least committed to marriage and most accepting of divorce (Hoem and Hoem, 1992; Axinn and Thornton, 1992). According to this argument, cohabitors are drawn disproportionately from the ranks of divorce-prone, and their original susceptibility to divorce contributes to a higher rates of marital instability. In addition some scholars have argued that the experience of cohabitation may change the way individuals view marriage and may teach them something about relationships that alters their view of divorce (Booth and Johnson, 1988; Thomson and Colella, 1991; Axinn and Thornton, 1992). According to this perspective, it possible for cohabitation to have a direct negative influence on marital stability by producing relationships, attitudes or values that increase susceptibility to divorce.

Axinn and Thornton (1992), for instance, use multiwave panel data of young Americans and their mothers to investigate the relationship between cohabitation and susceptibility to divorce. Their analysis shows that nonmarital cohabiting relationships indeed are selective of those who are least committed to marriage and most accepting of divorce. The evidence also supports the existence of causal factors linking nonmarital cohabiting experiences with subsequent approval of divorce. This result is consistent with the idea that cohabitation may change the way individuals view marriage and may teach them something about relationships that alters their view of divorce.

3. DATA AND ANALYTIC METHOD.

3.1 The data set

The data set providing the basis for the following analysis come from the 1985 Mail Survey of Swedish Men, which was conducted by Statistics Sweden (the Swedish National Central Bureau of Statistics). A simple random sample of men was drawn from each of the five-year cohorts born in 1936-40, 1941-45, 1946-50, 1951-55, 1956-60 as well as from the four-year cohort born in 1961-64.

From each male who responded, the survey obtained data on the community in which he grew up, his current occupation, education, leisure time and financial situation at the time of the survey, his previous marital and cohabitational history, present family situation, and on attitudes and future plans on fatherhood and children. Those who did not respond to the questionnaire were followed up by telephone. A total of 3171 males responded. Of these, 3115 records were usable for our particular purposes.

The overall response rate of 79% was very good for a mail survey, but lower than the corresponding figure in the 1981 survey of women, which was 87% (Arvidsson et al. 1982). About half (56%) of the nonrespondents were refusers while one-third could not be reached. As with the survey for the females, nonresponse rates varied by subgroup, ranging from 13% for the married men to 31% for divorced men. The corresponding figure for the never-married was 22%. Detailed tabulations of results from the survey can be found in Johansson (1991) and Lyberg

(1988). The youngest cohort (born in 1961-64) would only have reached its early twenties at the time of the survey and would have very little information to contribute beyond the initial stages of family formation. We have, therefore, dropped it from further consideration in the present analysis.

3.2 Sociodemographic correlates of marital disruption

In addition to the event (and duration) of premarital cohabitation, we have included some sociodemographic background variables in our analysis. The choice of most of these covariates is based on their statistical and/or substantive significance in the analysis of family initiation carried out on the same set of data (See Ghilagaber, 1993).

Birth cohort In general the cohort variable is expected to (vaguely) measure general secular developments such as changes in attitudes, values and norms, that are not picked up explicitly by other covariates. Societies continually experience modifications in family structure. Therefore, period factors differ across birth cohorts because different birth cohorts live through different historical periods or experience the same historical periods but at different ages in their lives. In view of the trends in attitudes and norms discussed in the introduction, we would expect higher risks of marital dissolution in younger than in older birth cohorts.

Parental Disruption: There is little concensus concerning the reasons for the relationship between parental divorce and the divorce of their children. The *intergenerational transmission* hypotheses has not been supported by many. Small, but positive, associations between parental divorce and children's risk of being divorced have been reported (Bumpass and Sweet, 1972).

There are theoretical reasons for expecting that a parental marital disruption will lead to more personal apprehensions about marital success and more negative attitudes toward marriage as an institution (Blechman, 1982). The hypothesis of more negative attitudes toward marriage among children of divorce has been supported by empirical evidence (Thornton and Freeman, 1982). One would expect that these more cautious attitudes toward marriage would slow the pace of entry into marriage. Such negative attitudes toward marriage could also lead to increased cohabitation. (For empirical evidence see, for instance, Ghilagaber, 1993). In view of the '(mode of entry) self selection' hypothesis then, the experiences accumulated while in a broken family will lead to a lesser commitment to marriages, and thus to higher risk of dissolution relative to those from an intact family.

Age at Marriage: In virtually every study of marital breakdown it has been found that marriage at young ages is the most powerful descriminant between marriages that survive and those that do not. Teenage marriages are more likely to end-up in divorce than marriages contracted at later ages (Ghilagaber, 1992; Haskey, 1987; Hoem, 1992; Moore and Waite, 1981).

At first, it might seem that age differentials in marital disruption are due to the longer exposure to the risk of breakdown for those married at younger ages, but the methods of analysis used here and in previous works usually control for the duration of exposure. Hence the explanation lies elsewhere. Perhaps, the impact of young marriages on the struggle for independence from the family of origin and the effect of subsequent changes in young adult role perceptions may explain the high incidence of divorce for young marriages.

Social class: In the present study, the social class of the respondent is measured by the social class position of his occupation at the time of the survey. For the most part, empirical studies have demonstrated that individuals with low-status occupations and lower income have a higher risk of family dissolution than those with high-status occupations. With an increase in the number of women in the labour force, it is important to assess the occupational status of one's wife in examining marital instability. For instance, Mott and Moore (1979) report an 'independence' effect as the cause of marital dissolution; women in the labour force develop resources and economic security apart from those of their husbands and hence couples in which the woman is of this type are expected to be more liable to dissolution that corresponding couples with no such woman.

Education: More attention needs to be directed toward understanding why and how education plays a part in marriage dissolution. The timing and the level of the husband's and wife's education and how each spouse views his (her) education in reference to the marriage need to be addressed. Here we shall focus on the following two views:

One could argue that there is an inverse relationship between marital disruption and level of education. The reason is that with increase in the level of education, net of other factors, the 'maturity' to handle difficult situations within the marriage develops.

On the other hand, it is common knowledge that one's educational level is an indicator of his social and economic independence, in particular his ability to cope up with the consequences of family disruption (Blossfield et al., 1993). This is expected to work in the direction of increasing the educational gradient in the disruption risk of highly educated men.

Children: The timing, planning, number and age of children are among variables associated with family dissolution. The union from which the chid(ren) have come has also been associated with dissolution. In view of economic theory, children are investments common to the partners in the union, they are union specific capital (Becker, 1981; 1991). The arrival of a child(ren) should, therefore be a signal of a well functioning relationship with a low risk of dissolution.

A different view is held by other investigators. Chester (1972) demonstrates that the common belief that childlessness is positively associated with instability of marriage, is untenable. It is concluded that the alleged relationship between childlessness and instability of marriage is probably either non-existent or the reverse of that normally assumed and that in any case measurement of the net overall effect of childlessness does not provide a helpful datum. According to Cherlin (1977), children are detterent to divorce and separation only when they are in the pre-school ages, when the time and effort required for child care are at their peak. In view of this perspective, children prevent marital dissolution not because they build new bonds between parents but rather because early child care may be too expensive and time consuming for one spouse to manage alone. For lack of information, the present study does not include age of children.

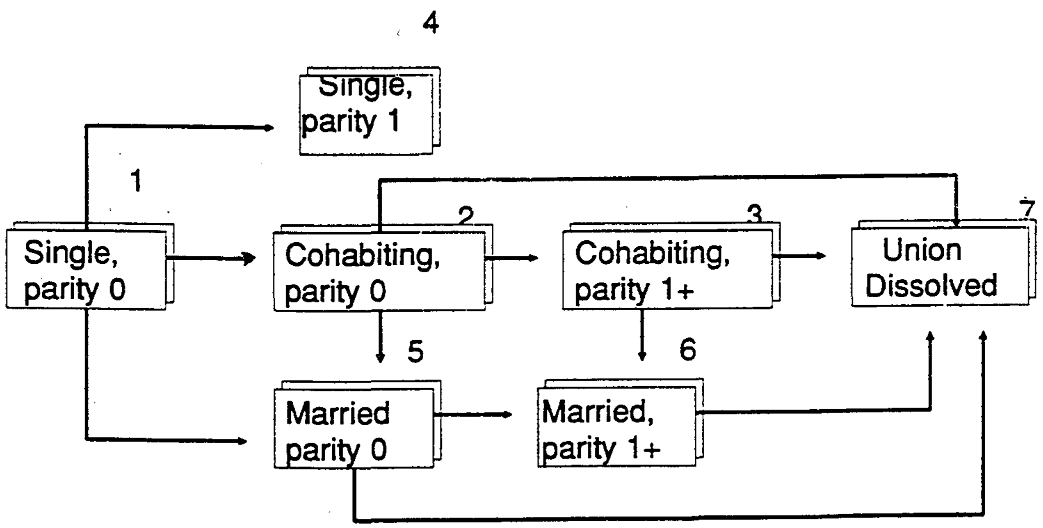
Marriage duration: Marriage duration is also important determinant of marital dissolution both, separately and jointly with other factors such as cohort, age at marriage, and social class.

Early dissolving couples show certain characterstics, for example, a tendency for the husband and the wife to have had a low social class at marriage, or to have married as teenagers, or to have had their contracts in civil ceremonies (Haskey, 1987). Further, it is a common knowledge that during the early days, and under normal situations, divorce petition could only be filed after a number of years of marriage had been elapsed. In modern days this time bar has been shortened in most societies which had had it. It is thus reasonable to suppose that some of the marriages which, in the past, could be dissolved only say, between their third and fourth anniversaries might nowadays dissolve between say, thier first and second anniversaries. This, in our case will be taken care, by the joint effect of birth cohort and marriage duration.

3.3 Method of Analysis

The individual-level population processes that are worth investigating in relation to the 1985 Mail Survey can be represented by a seven-statuses-and-transitions diagram as indicated by the boxes and arrows shown in Fig. 1.

Fig. 1: Statuses and transitions



Issues related to family initiation behaviour (intensities of transition from State 1 to State 2; as well as from States 1 and 2 to State 5, have been addressed in Ghilagaber (1993). In the study, correlates of the choice between marriage and cohabitation as a first union are examined. Further, differences in the intensities of marriage formation of cohabiting men and single men, are examined.

In the present paper, the later phases of the family in the same cohort of men are investigated in order to assess correlates of marital dissolution. The basic issue to be addressed here is whether the risk of union disruption is the same for all modes of entry into the married life or is mode-specific. In relation to Fig. 1, the focus of the present study is on the intensity of transition from 'Married' to 'Union dissolved' ($\mu_{5->7}$ and $\mu_{6->7}$). In so doing, transitions from states 1 and 2 to state 5 will represent the two levels of the factor '*Mode of entry into marriage*' while the transition from state 5 to 6 will represent a level of the *Parity* (*Children*) factor. In other words, the *parity* factor is included as a time-varying covariate while *mode of entry* and all other factors are fixed overtime.

The analysis consists of fitting a series of hazard models to the transition intensity and estimating the effects of the *mode of entry* and the other sociodemographic variables on the intensity of marriage dissolution.

From the preceding discussions, it is clear that the factors involved in cohabitation, marriage, childbearing, & dissolution of unions, are inter-linked in complex ways. For our case, we have approached the problem by including the experience of premarital cohabitation as one among the many fixed explanatory variables and analysing the risk of dissolution of marital unions. The approach involves fitting a multiplicative hazard model for the transition studied. For the purpose of expressing the relative intensity of a particular group on one of the factors considered, one level is selected as a baseline for that factor. The relative intensities are then, indicators of how often a transition occurs to individuals at a particular level of a factor, relative to the baseline level of the same factor.

The Maximum Likelihood method has been used to estimate the parameters in each model. We have used the ROCANOVA computer program, Version 1.0 (Martinelle, 1993). The program uses occurrences and exposure matrices as input data and an algorithm called Iterative Proportional Fitting, in estimating the parameters.

The model counterpart of a rate of transition is the corresponding hazard (intensity) function. Assume that λ is such a function for a particular man and for a particular transition, say the intensity of marital dissolution. For purposes of illustration, let us further assume that the intensity depends only on four factors; age at marriage (A), *birth cohort* (C), *marriage duration* (D), and *Mode of entry* into marriage (M) (an indicator of whether or not the man cohabited before marriage.) The model assumes that for a particular *age group at marriage* a (a = 1,2,3), *birth cohort* c (c = 1,...,5), *marriage duration* d, (d = 1,...,6), and *mode of entry* m (m = 1,2), we can write the intensity function $\lambda(a,c,d,m)$ in the form

$$\lambda(a,c,d,m) = \mathrm{A}(a)\mathrm{C}(c)\mathrm{DM}(d,m) \qquad (1)$$

where A(a) and C(c) are factors (parameters) representing the effects specific to the *age group at marriage* and *cohort* covariates respectively, and DM(a,m) is a factor (parameter) representing an effect due to an interaction (if any) between the *duration* and *mode of entry* covariates.

Let a_0 be the age group 20-25, c_0 be the cohort born in 1946-1950, and let m_0 be the mode of entry " *direct marriage*", and let us use a_0, c_0, and m_0 as baseline levels for factors A, C, and M, respectively.

We make all parameters in (1) identifiable by defining $A(a_0) = C(c_0) = DM(d_0,m_0) = 1$. Thus,

$$\lambda(d, a_0, c_0, m_0) = D(d), \tag{2}$$

is a baseline duration structure in the intensity function λ, while

$$\lambda(a,c,d,m)/\lambda(a,c_0,d,m) = C(c) \tag{3}$$

for all a, d and m, is the risk of marital dissolution for men born in cohort c, relative to the risk for men born in cohort c_0 at all marriage durations, all modes of entry and all age groups at marriage. A similar formula holds for the risk in age group a, relative to the risk in age group a_0 at all durations, for all modes of entry, and for all birth cohorts; and/or for the risk in mode m, relative to the risk in mode m_0, at all durations, for all birth cohorts, and for all ages at marriage.

Factors A, C, and M in (1) above are fixed (constant) over duration of marriage. The model can be extended easily to allow for explanatory variables that change in value over time. We have, for instance included the *parity (number of children)* factor as a time-varying variable (with values of 0 for the nulliparous men and 1, 2, ... for men with positive parities.) In the presence of one such additional variable, say P (parity), (1) above can be extended to

$$\lambda(a,c,d,m,p_d) = A(a)C(c)P(p_d)DM(d,m). \tag{4}$$

The model in (4) says that in addition to factors A, C, D and M, the intensity (of marital dissolution in our case) at marriage duration *d* depends on the number of children the individual has at the same marital duration *d*. Models with time-varying variables can be estimated using the same method that is discussed below in connection with models with fixed variables. For previous applications, see for instance Allison (1984); Ghilagaber (1993).

A nonzero interaction term in (1) means that while D(*d*) is the duration structure of the intensity for marital disruption for *mode of entry* m_0, the other mode of entry has its own duration structure for this intensity. If we plot

$$\lambda(a_0,c_0,d,m) = DM(d,m)$$

as a function of *d* for fixed *m* we get different curves for different values of *m*. (See Fig. 2 below).

In the absence of a nonzero interaction term, the model assumes that there is an underlying duration-structure (duration pattern) D(*d*) characterizing the intensity curves, that is the same for all levels of the other factors included in the model.

In search of a parsimonious model and for the purpose of evaluating the relative importance of each factor in explaining the risk of marital dissolution, we have started with a basic model that includes only the cohort factor together with the time variable D (marriage duration), and subsequently extended it by adding, step by step, other factors in the order as may be experienced by the individual. Apart from the specific effects, we have also fitted models which include interaction terms between *the mode of entry* and selected other factors. At each step the fit of the model was examined and differences between various models were tested for significance using the fact that the difference (or likelihood ratio) is approximately distributed as Chi-square with appropriate degrees of freedom.

4. Results

4.1 Dissolution patterns across sociodemographic covariates

Table 1 below contains estimated relative risks of marital dissolution across covariates under various models. As mentioned earlier we have included the covariates in the order they may be experienced by men in their life-course. The following description refers to the first six columns of Table 1 (before the introduction of *premarital cohabitation* into the model).

The theoretical expectation in relation to the trends across birth cohorts seems to be confirmed by our findings here. Relative risks of dissolution indeed rise as we move from older

Table 1: Estimated relative risks of union dissolution in different (nested) models. (Swedish men born 1936-60)

Covariates/levels	Relative risks of dissolution in different models							
Birth cohort	*1*	*2*	*3*	*4*	*5*	*6*	*7*	*8*
1936-40	.52	.52	.57	.56	.60	.65	.98	.97
1941-45	.93	.92	.91	.91	.95	1.01	1.32	1.33
1946-50	1+	1	1	1	1	1	1	1
1951-55	1.42	1.41	1.41	1.40	1.46	1.57	1.45	1.49
1956-60	2.32	2.31	1.44	2.34	2.36	2.45	2.09	2.16
Family of Origin								
Intact		1	1	1	1	1	1	1
Disrupted		1.38	1.37	1.35	1.37	1.30	1.23	1.21
*Age at marriage**								
15-19			1.60	1.58	1.66	1.66	1.93	1.92
20-25			1	1	1	1	1	1
26+			1.55	1.59	1.57	1.46	1.25	1.32
Social class								
Unsk./Skil.				1.14	1.39	1.47	1.37	1.34
WC				1	1	1	1	1
Farmer				1.09	1.23	1.22	1.18	1.18
Self				.27	.32	.34	.48	.49
Educational level								
Primary					1.01	1.01	1.08	1.10
Secondary					1	1	1	1
University					1.59	1.59	1.58	1.59
*Parity progression***								
0						3.73	3.14	3.12
1+						1	1	1
*Mode of entry into marriage***								
Direct							1	
After cohabiting							4.83	
*Duration of Cohabitation***								
0 (direct marriage)								1
1-5 months								3.98
6-12 months								5.86
13-35 months								5.22
36+ months								3.40

+ For each factor, baseline levels are indicated by a parameter value of 1 (without decimals) in all panels of the table. Note that the (duration of cohabitation) factor is only a further partitioning of the mode of entry factor.

** significant at 5% level*

*** significant at 1% level. In addition, marriage duration, whose relative risks are not displayed, has been significant at 1% level.*

to younger birth cohorts. In model 6 (the model that includes all covariates, except the *mode of entry*) we note that the risk of dissolution among men in the oldest cohort is only .65 times (while that of men in the youngest cohort is about 2.5 times) that of the reference point (men born in 1946-50). Men from disrupted families end up in dissoultion at a rate of 30% higher than corresponding men brought up in intact families. This, again is in accordance with our expectation.

We hypothesized that the risk of dissolution falls with rise in age at marriage. Indeed, our results show that marriages initiated at the very young ages are about 70% more likely to dissolve than marriages contracted at ages of early twenties. The aparent lack of consistency of this trend beyond age 25 is probably a compositional effect. Table 1 provides only an overall picture of patterns of marital dissolution across the covariates included in the study. But whether the pattern across each covariate is the same for all modes of entry into the married life, or not, has not yet been investigated.

To examine this, we have fitted a model that includes an interaction term between the *mode of entry* and *age at marriage*. The results (not displayed here) show that for those directly married, the risk of dissolution consistently decreases with increase in age at marriage. For those who marry only after cohabiting, early and late marriages are respectively 80% and 40% more likely to dissolve than marriages contracted at ages 20-25.

Men in lower occupations (unskilled and skilled workers) are about 50% more likely to end up in dissolution than white collar employees. In fact, this 'super-risk' is only a lower bound. We recall that this factor refers to the respondent's occupational class at the time of the survey. It is thus likely for the 'white collar' category to include men who, at the time of marriage, belonged to a lower occupational class. In view of the higher risk of dissolution in the latter category, therefore, we suspect that the relative risk could have been higher if the variable referred to the time of marriage. Among the nonhierarchical categories, farmers are about 20% more likely, while the self-employed are about 70% less likely to experience marital dissolution than white collar employess.

Overall, the educational gradient supports the arguement that one's level of education is an indicator of his potential to cope up with the consequences of marital breakdown. Men with university level education are about 60% more likely to dissolve their marriages than those at lower levels of education.

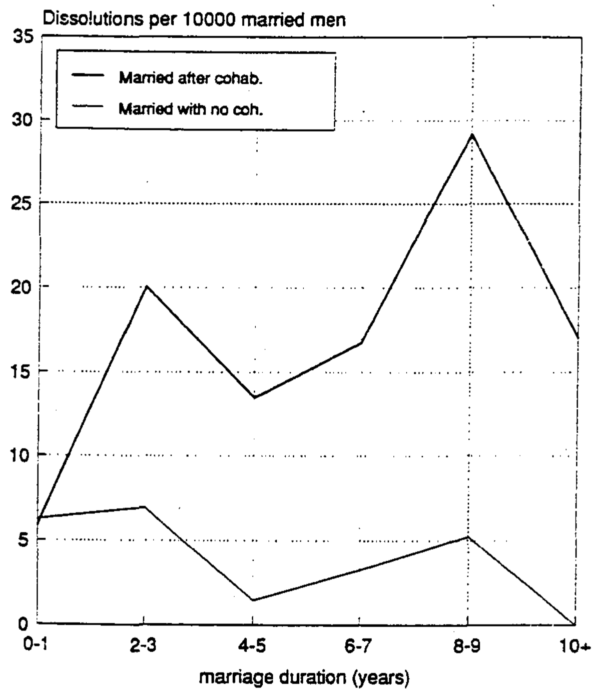
Whether or not a man has a child(ren), turns to be a strong descriminator of the survival of his marriage. Those with no children within the marriage are nearly 4 times as likely to end up in dissolution that those with at least one child within the marriage. The parity (here representing the number of children within marriage) has been measured by counting the number of live-born children born at any time beginning the first month of marriage. This includes children which were conceived before marriage (at most seven months before marriage). An explicit representation of this latter event, by means of another factor (*pre-marital pregnancy*) could have been a better way of capturing the effect. As we have not done this, we pass by reminding the reader that the effect of parity, as shown in Table 1, would have been deflated if *premarital pregnancy* was explicitely taken care of. In winding up this subsection, we note that cohort, family of origin, educational level and social class have not been statisticaly significant.

4.2 The role of premarital cohabitation

The seventh column in Table 1 (model 7), shows that previously cohabiting men are nearly 5 times more likely to end their present marriage in dissolution than men who never cohabited before marriage. The direction of the effect is in accordance with findings for Swedish women and that of men and/or women elsewhere. The magnitude of the effect is, however, by far greater here (for Swedish men) than in any of the previous findings we are aware of. For instance, the corresponding relative risk for Swedish women was only about 1.5 (Hoem and Hoem, 1992), that for Dutch (men and women) was about 1.3 (Klijzing, 1992), and that of US men and women were respectively, 1.80 and 1.53 (DeMaris and Rao, 1992).

The duration profiles of baseline intensities of marital disruption, by mode of entry, are displayed in Fig. 2 below. At all marriage durations, men who initiate marriage only after having lived in cohabitation, experience the event of marital dissolution, at higher rates than those who enter the married life directly. The difference is highly pronounced after marital

Fig. 2: Baseline intensities of family-
disruption (Swedish men born 1936-64)



durations of over seven years. Else, the trends are the same in both curves, an initial increase in the number of dissolutions, followed by a fall between 2 and 5 years of marriage, another rise between 6 and 9 years, followed by a drop after 10 years of marriage.

To account for differences in the length of cohabitational period, we have fitted a model in which the second level of the '*mode of entry*' factor is partitioned into smaller periods. As shown in the bottom of the last column of Table 1 (model 8) the 'super risk' of cohabitors persists for all premarital durations, though cohabitation of between 6 months and 3 years seem to be relatively more sensitive relative to the other durations.

5. Summary and conclusions

This research report has examined the patterns, determinants, and differentials in the risk of marital dissolution among Swedish men. The central issue has been the relationship between premarital cohabitation and the risk of marital dissolution. To this end, data from the 1985 Survey of Swedish men has been analysed using proportional hazard models.

Consistent with earlier findings, the results show that first marriages preceded by cohabitation are less stable than first marriages that were not. The influence of premarital cohabitation on the risk of marital dissolution is however, stronger in our present study than in earlier findings. Previous cohabitors, compared to noncohabitors tend to be at much greater differential risk of dissolution at all durations of marriage. This is in contrast with the view that emphasizes the importance of the transition to marriage, and argues that cohabitors adjust their behaviours and expectations after marriage. Apart from premarital cohabitation we found that age at marriage, marital duration, and children are strong determinants of risk of marital dissolution.

The intended contribution of the study was to contribute to our understanding of the complex process by which the family, one of the most fundamental social institutions, operates in a modern welfare society. Consistent with previous explanations offered for the cohabitation effects on marital stability, we suggest that cohabitors end up in a lower quality marriages, have lower committment to the institution of marriage, and that men who cohabited had more individualistic views of marriage than those who had not. These suggestions are only tentative, however, and further investigations that include information on quality of marriage and individuals' perceptions of marriage are required in order to draw firmer conclusions. Given this, it is likely that cohabitation will continue to be a phenomenon of interest in future research.

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