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FAMILY INITIATION AMONG SWEDISH MALES BORN 1936-1964: THE CHOICE BETWEEN MARRIAGE AND COHABITATION

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Abstract

This paper examines family initiation behaviour among Swedish males born in the period 1936-64 with data from the 1985 survey of Swedish males, which had about 3200 respondents. The study provides a systematic description of the national pattern of conjugal-union formation, within the context of theories about the relationship between various demographic and socioeconomic variables on the one hand and family initiation on the other.

Analysis using separate multiplicative hazard models shows that cohabiting men are more likely to marry than single men. Results from a competing-risks analysis show a recent reversal of the great preference of unmarried cohabitation over marriage that had continued for the last few decades.

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1. INTRODUCTION

Entry into a first conjugal union is perhaps one of the most important life-cycle transitions made by men and women in their movement from adolescence to adulthood. Understanding the timing of such entry and changes in such timing over time, therefore, is an important step towards developing a theory of the transition from adolescence to adulthood.

Family-initiation behaviour is particularly important to our understanding of the linkages between social and economic change and demographic change. Traditionally marriage has signaled the beginning of the family formation process and, consequently, has played a central institutional role in determining overall levels of fertility. Continuous modification in family structure has, however, given rise to *cohabitation*, a new form of living arrangement, and with a different level of commitment to the partnership.

Significant changes in social, cultural and economic structures, which have been central features of Western Europe and the US during the era following World War II, create pressure for change in the family formation patterns of a society. While the mechanisms underlying such change are complex and often poorly understood, the historical experience of Europe and the USA, as well as more recent trends in the developing countries, have made it abundantly clear that both nuptiality and fertility are responsive to the structural changes that occur during the course of "modernization".

During the last two decades, unmarried cohabitation, both as a prelude to marriage and as a more or less permanent alternative to it, has strongly increased in popularity in many countries in Western Europe such as Sweden (J. Hoem and Rennermalm, 1985), Norway (Blanc, 1985; Blom, 1992), Denmark (J. Hoem & Selmer, 1984), France (Leridon, 1990), The Netherlands (Latten, 1984; Liefbroer, 1991), the United Kingdom (Brown and Kiernan, 1981), as well as in the United States (Thronton, 1988, 1989; Willis & Michael, 1988), Canada (Rao, 1990) and Australia (Santow & Bracher, 1993).

The practice of starting a union without marrying is not a new phenomenon in Sweden. (For references, see J. Hoem, 1986.) It has roots that go back more than a century. We do not really know how common cohabitational unions were before, however, since nonmarital cohabitation received little public attention until the later part of the 1960s, when marriage rates suddenly started to fall dramatically (B. Hoem, 1992a). The Swedish fertility survey of 1981 and the 1985 Survey of Swedish males provide the first opportunities to investigate in depth cohabitational behaviour among Swedish females and males respectively. The term cohabitation itself has no standard definition. Consistent with others (Blanc, 1985; J. Hoem and Rennermalm, 1985; Liefbror, 1991; Willis & Michael, 1988) we have adopted the following definition here: cohabitation is living with a partner (of opposite sex) in a marriage-like relationship for a period of time of one month or more. Such definition, of course, is subject to some interpretation, but we believe that it captures the essence of the phenomenon of interest and is sufficiently unambiguous.

Elsewhere, cohabitation has been conceived as a short-duration marriage and the commitment to the partnership has been closely related to its expected duration (Willis & Michael, 1988). Accordingly, attempts have been made to investigate why a couple would prefer a short duration for partnership. Some reasons involve uncertainty; uncertainty about whether one wants to have any partner, uncertainty about the particular person chosen as a partner, or uncertainty about some other high priority-event (such as another and preferred partner), that dictates avoiding a long term commitment to a partner. A short test period with a partner may appear to be a reasonable strategy for dealing with such uncertainties.

Another reason for a short-duration partnership, unrelated to uncertainty, is postponement. There may be a known, preferred opportunity (such as educational training) that will become available at a known future date, and the current partnership may be formed with certain knowledge that it is to be temporary.

Others (Bachrach & Horn, 1987; Liefbroer, 1991) suggest the dichotomy of a cohabitation as either a prelude to marriage or a substitute to it. Another dichotomy regarding cohabitation distinguishes cohabitation as a trial marriage from cohabitors as a "select group of people for whom relationships in general are characterized by a lack of commitment and stability" (Bennet, Blanc, & Blom, 1988; Blanc, 1985). This last distinction has implications about predicted subsequent behaviour of cohabitors.

In Sweden however, the rise in cohabitation after the mid-1960s have been coupled with a lengthening of the time spent in the cohabitational union and as noted in B. Hoem (1992a) the joint effect has resulted in a sharp drop in marriage rates among Swedish females during the mid-1980s. (See also Hoem and Hoem, 1988). This suggests that so far we have relatively weaker evidence to view cohabiation among Swedish adults, as a short-duration marriage.

The present study utilises the 1985 Mail Survey data in order to examine Swedish males' behaviour in union initiation after the mid-twentieth century. Issues addressed in this paper include: *What category of Swedish men prefer cohabitation to marriage as a first union? Why would men choose cohabitation instead of marriage? What is the most likely subsequent behaviour of cohabitors? What does the descriptive evidence tell us about recent changes in the institution of marriage and/or cohabitation*?

The analysis is carried out in two stages. At the initial stage, separate hazard regression models are used to estimate intensities of transition from the single status to marriage, cohabitation, or either of these forms of union formation, across some sociodemographic variables. This preliminary stage of analysis was also used to screen out insignificant variables from further analysis. In the second and more refined stage of analysis, the intensity of entry into a marital union and that of entry into nonmarital cohabitation are modelled simultaneously. (For previous applications, see Gomez de Leon & Potter, 1989; Liefbroer, 1991.)

Results from our preliminary analysis provide *no* support for the notion that cohabiting men purposefully avoid or reject marriage. On the contrary, cohabiting males are more likely to enter legal marriage than their single counterparts. This finding, which is consistent with findings for the United States (Willis and Michael, 1988), suggests that cohabitation in Sweden has served more as a trial union than a permanent replacement for a formal marriage.

Results from the competing risk analysis indicate that overall Swedish males initiate a union by cohabitation at a rate of nearly three times the rate of marriage, but that the relative intensities of entry into marriage and cohabitation vary across background variables. Further, it is shown that the greater popularity of unmarried cohabitation than marriage which persisted for the last few decades, has recently been reversed.

In our next section, the source of data is discussed together with the variables and the underlying substantive theory behind the choice of such variables. The statistical model and the methods of analysis are described too. Section 3 is devoted to a presentation and discussion of the empirical findings.

2. DATA AND METHODS

2.1 The Data Set

The data set providing the basis for the following analysis is 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. Sample sizes in each age group are shown in Table 1.

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.

Table 1. Sample sizes in the 1985 mail-survey, by birth cohort.

Birth cohort	Age at survey time	Sample size
1936-1940	44-48	587
1940-1945	39-43	738
1946-1950	34-38	799
1951-1955	29-33	631
1956-1960	24-28	653
1961-1964	20-23	587
Total (1936-64)	20-48	3995

Source: Lyberg (1988, p. 4).

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).

2.2 Variable Specification and Hypotheses

Our focus is on the choice a man makes between a formal marriage and a cohabitation at the beginning of his partnership. An analysis of such a choice is likely to benefit from consideration of the social context and the individuals' developmental level and concurrent role demands. The major explanations for the growing popularity of unmarried cohabitation since the 1960s focus on *opportunities provided by the social structure* and/or on changing *preferences* among young adults (Liefbroer, 1991). This has lead us to distinguish three sets of determining factors; intercohort factors, intracohort factors, and factors related to the individuals' current experiences.

The intercohort variable considered in the study is the *birth cohort* of the respondent. Birth cohort has been one of the most important demographic factors determining an individual's age at entry into a first union (Carter and Glick, 1970). Our society continually experiments with modifications in family structure. Therefore, period factors differ across cohorts because different birth cohorts live through different historical periods or experience the same historical periods but at different ages in their lives.

To account for variations in the timing of union formation due to differences in background factors, four intracohort variables are considered, namely, *social class position of the occupation of the respondent's father, social class position of the occupation of the respondent's mother, size of community of respondent's origin,* and *disruption history of his family of origin*. Generally, parents' social class is likely to affect individuals' access to resources. Parents use their resources to promote behaviour (such as the extent of schooling) that provide alternatives to premature marriage among their children. For details see Hogan (1978), Bernhardt & B. Hoem (1985), and Michael & Tuma (1985). Further, the socioeconomic status of the respondent's family is expected to influence his aspirations, values, and lifetime plans.

The size of the community in which the respondent grew up is expected to affect both expectations and opportunities in union formation.

Whether the family of origin was intact or disrupted during the respondent's childhood also relates to the timing and type of union formation. Recent studies (Cherlin et al., 1991) suggest that the observed differences between children from families in which the parents have separated or divorced and children from two-parent families may be traced to three distinct sources. The first source is the effect of growing up in a dysfunctional family - a home where serious problems of the parents or the children make normal development difficult. Parents with psychological impairments are reportedly more prone to divorce and their children are more likely to experience developmental difficulties (Kiernan, 1986). A second source, often accompanying the first, is severe and protracted marital conflict, which is known to harm children's development and often leads to divorce (Emery, 1982). The third source is the difficult transition that occurs only after couples separate - the emotional upset, fall in income, diminished parenting, continued conflict, and so forth. Although some studies acknowledge the potentially adverse contribution of each source (Heterington et al. 1985), nearly all empirical studies, including the present, have focused exclusively on the third - the period after separation - and have collected information only after the separation occurred. Among the latter, Kiernan (1992) suggests that, in general, coming from a broken home is associated with a greater likelihood of early union formation, which suggests to some that union formation provides an escape from an unhappy, complex environment. Given the relative ease and lack of strong commitment in nonmarital cohabitation, one would expect the propensity to be higher for the formation of cohabitational unions than for marriage.

The concurrent and recent life-course experiences that are included in the present study are the *respondent's own social class at the time of the survey,* the *respondent's level of education at the time of the survey,* an indicator of whether the respondent lives in a *consensual (nonmarital) union,* and an indicator of whether the respondent has a *pregnant (preunion*) sex partner. The social class of the respondent was measured by the social class position of his occupation at the time of the survey, while the indicator of pregnancy was computed by subtracting seven months from the date of birth of a recorded live-born child fathered by the respondent.

We have several findings suggesting that the poorer the economic prospects and circumstances of the male at the time of the formation of the partnership the more likely he is to choose cohabitation over marriage. This is consistent with less commitment to the partnership among couples where the economic gains from marriage are less. See for instance Landale and Tolnay (1991), and Cooney and Hogan (1991).

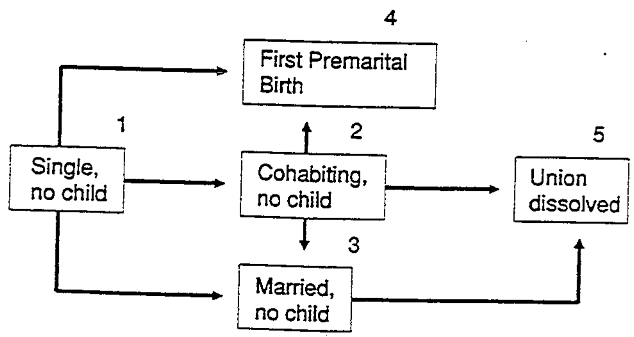
Goldscheider and Waite (1986) on the other hand argue that recent declines in the marriage rate have not resulted from increased barriers to marriage but from declines in relative preferences for marriage. This suggests the emergence of a more attractive alternative of union formation, namely unmarried cohabitation; an alternative that provides easier entry and demands less commitment. The investigation of whether this new phenomenon is an institution used to strategically search intensively for a marital partner or is an institution that is in some more fundamental way a replacement for marriage, is among the aims of our present study.

As we mentioned earlier, the traditional sequence of relatively early entry into legal marriage followed by child-bearing is no longer the course followed by the majority of men in modern societies. Instead, entry into nonmarital cohabitation has been rising. There are two competing hypotheses in connection with the effect of premarital cohabitation on entry into marriage. (See Santow and Bracher, 1993). On the one hand, cohabitation can be viewed as a rejection of the constraints of traditional, legal marriage, and cohabitants are expected to be less likely to enter a legal marriage than noncohabitants at all ages. An alternative perspective is that cohabiting males have already found a partner from the market of potential partners and therefore are expected to be more likely to enter legal marriage than their noncohabiting counterparts.

We deal with these two competing perspectives by including the *'cohabitating'* status as a time-varying covariate (see Appendix B) in part of the analysis, and by comparing the intensities of transition into marriage from the *'single'* and *'cohabiting'* statuses.

The event of pregnancy outside union is expected to accelerate union initiation (B. Hoem, 1987).

Fig. 1: Statuses and Transitions



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

In the present paper interest is focused on the following intensities of transition:

1) intensity of transition from 'Single' to 'Married' ($\mu_{1->3}$).
2) intensity of transition from 'Single' to 'Cohabiting' ($\mu_{1->2}$).
3) intensity of transition from 'Single' to 'In First Union' in general. (i.e. to either 'Married' or 'Cohabiting'.)

The initial stage of the analysis will involve fitting separate hazard models to each of the above three transition intensities and estimating the effects of the various sociodemographic variables on each of the intensities. As mentioned earlier, this initial stage of the analysis will mainly be for the purpose of screening out factors with no significant effects on one or more of the intensities. In addition we expect to get an overall picture of the pattern of union formation. In the analysis of the transition from '*single*' status to '*married*', we include '*cohabitational status*' as a time-varying variable. This allows us to analyse the two intensities $\mu_{1->3}$ and $\mu_{2->3}$ at the same time. In this way, we investigate whether cohabiting men are more/less likely to marry than single men.

After the analyses of separate intensities, the first two intensities of transition ($\mu_{1->3}$ and $\mu_{1->2}$) will be modelled simultaneously in the framework of competing-risks models with a view of demonstrating any methodological differences. At this latter stage '*cohabitational status*' is not included as a factor among the observed variables, for we are interested in comparing the patterns in transition from the '*single*' status into '*Married*' and into '*Cohabiting*' statuses.

We expect that the findings of the present paper will also serve as a stimulus for future investigations that focus on other transition intensities, say from 'Single' to 'First Premarital Birth', both directly and through the status 'Cohabiting' ($\mu_{1->4}$ and $\mu_{2->4}$ respectively); from the statuses 'Cohabiting' and 'Married' to the status 'Union Dissolved' ($\mu_{2->5}$ and $\mu_{3->5}$ respectively) after utilizing additional information relevant to the event of union dissolution for nulliparous men; and other transitions of interest. Analysis of transition to State 4 would give indications on levels, trends and determinants of nonmarital first births. Here it represents an unanalysed competing risk.

2.3 Statistical Methods

As we just explained, we use an approach where the intensities (hazards) of entering into a marital union and of entering into a nonmarital union will be modelled first separately and then simultaneously. In each case the approach involves a multiplicative model for each 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 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.

We have fitted separate models for each of the three transition intensities for men in the single status; namely i) the intensity of transition into marriage, ii) the intensity of transition into cohabitation, and iii) the intensity of transition into any type of first union, i.e., either into cohabitation or into marriage. For details, see Appendix A.

As the next step in our analysis we have simultaneously modelled the intensity of exit from the status 'single' into marriage and into nonmarital cohabitation. By formally including type of union as just another covariate in the analysis, it is possible to easily test whether the effects of the other characteristics on the process of union formation vary according to the type of union that is entered. This means that we have fooled the LOGLIN program into believing that the cause of transition is another factor affecting the intensity of union formation together with the real observed factors. This is achieved by specifying the same exposures for all levels of that new factor (union type). Details on how we achieve this, is explained in Appendix B.

The Maximum Likelihood method has been used to estimate the parameters in each model. We have used the LOGLIN computer program, Version 1.64 (Olivier & Neff, 1976). The program uses occurrences and exposure matrices as input data and an algorithm called Iterative Proportional Fitting, in estimating the parameters. See Appendices A and B.

3. EMPIRICAL FINDINGS

3.1 Results from separate analyses of intensities of entry into marriage and into cohabitation.

We have started our analysis by fitting separate hazard models for each type of union entered. As we mentioned earlier, the main goal at this stage in the analysis is to screen out non-significant factors. Social groups of respondent's parents and his level of education at survey time, which in the absence of his own social class, had moderate effects on entry into a first union, were found to be redundant when the respondent's own social group is included in the model. Accordingly, they were dropped out from subsequent analyses. Further, the size of community of his origin was dropped from further analysis on the basis of lack of any meaningful pattern in the intensities across its levels.

In parallel studies of Swedish women, social groups of respondent's parents have had significant effects on union initiation behaviour among Swedish women (Bernhardt and B. Hoem, 1985). Such studies have not included respondent's own social class, however, and we cannot make a strict comparison between males and females regarding the effect of the respondent's own social class on union formation behaviour.

Among the significant factors, we found that cohabiting men are more than four times more likely to enter into marriage than single men (details not shown). This provides supporting evidence to the hypothesis that cohabitation is a prelude to marriage rather than a permanent alternative to it. The estimated underlying age structure of the union-formation 'risks' is displayed in Fig. 2, and the cohort and social-group patterns are displayed in Figs. 3 and 4 respectively. Fig. 2 shows that, overall, Swedish males initiate their family unions in their early to late twenties and that they do so more frequently through cohabitation rather than through formal marriage. The difference is most pronounced for men in their early- and late- twenties. (The sum of the figures in the two lower curves of Fig. 2 is close to but not necessarily equal to the corresponding figure in the upper curve. This is so because the three curves have been estimated independently.)

The cohort pattern in Fig. 3 shows an initial decline in relative intensities of entry into marriage, and a corresponding rise in the relative intensities of entering into cohabitation, for the three oldest cohorts, and and a reversal of such a trend thereafter. Overall, union formation has declined during the 1980s.

The overall picture depicted in Fig. 4 is that the relative intensities of union formation rises with a rise in the social-class position over the positions that have a natural hierarchy (unskilled workers to higher white-collar employees). Unskilled and skilled workers enter cohabitation more readily (and marriage less readily) than white-collar employess do.

The effects of these and the other (significant) variables are closely similar to those obtained in the simultaneous analysis (see Section 3.2), and we defer a detailed discussion until then.

3.2 Simultaneous analysis of intensities of entry into marital and nonmarital unions.

After this demonstration of differential patterns in the two competing forms of union formation, it is time to investigate in greater depth the effect of the factors we have kept in our analysis, namely birth cohort, respondent's own social class, family of origin, and age, on young men's choice between marriage and cohabitation as a first union. This is the subject of the present section.

The estimated parameter values for a model that includes no interaction term among any of these factors are presented in Table 2.

As is evident from the table, the overall rate of entry into nonmarital cohabitation is close to three (actually 2.77) times the rate of entry into marriage. The cohort pattern in Table 2 again indicates that unmarried cohabitation generally rose in popularity from older to younger cohorts of men, but that there has been a recent reversal. (See Fig. 3). This result is consistent with that for Swedish women (J. Hoem and Rennermalm, 1985; B. Hoem, 1992b). There is a corresponding estimated countervailing reversal of the decline in the "risk" of entry into a marital first union. As we shall see in Section 3.3, this is essentially a compositional effect, however. Marriage patterns after the two oldest cohorts were not uniform across age groups, and the marriage revival is only apparent.

With regard to social class, the results again indicate that the rate of entry into marriage generally is higher for men who are higher up on the social-group ladder. (See Fig. 4). Higher white-collar employees, farmers and the self-employed men enter marriage at relatively higher rates than other categories of men. There is a corresponding systematic pattern in the rate of entry into cohabitation across social class, though it shows up better in Fig. 4 than in Table 2. The table and Fig. 4 both clearly show that the self-employed and men in the category 'others' enter cohabitation at a much lower rate than men in the other categories. In particular, men classified in the 'others' category seem to lose out on the union-formation market. This last category probably contains a 'residual' group of socially 'rejected' men who most likely are unable or unwilling to report their occupation. While farmers and the self-employed enter marriage at almost the same rate, the self-employed enter nonmarital unions at an extremely low rate. (The relative rate of 0.53 is by far less than 1.)

As we mentioned earlier, the social-group variable was measured at the time of interview and it is used as a fixed variable throughout the analysis despite the fact that most of the analysis is aimed at explaining behaviour which took place before the survey. Social mobility (changes in social group) are likely to occur between the time of union formation and the survey date. An interesting question at this stage, therefore, is to what extent the changes in union-formation patterns across social classes can be attributed to changes in the distribution of young adults across the social positions, and to what extent they reflect real differences in behaviour of men in different social-class positions.

We may safely assume that some of the respondents have improved their social-class position between the time of union formation and the date of the survey. Therefore, at each level of the recorded social-group factor except the lowest one there are some men who, at the time of union formation, belonged to a lower level. Men with a lower social group have lower rates of marriage and in general moderately higher rates of cohabitation. It is plausible, therefore, that the relative intensities of marriage corresponding to the recorded social-class factor in Table 2 are underestimates of the true rates, while those of consensual-union formation are overestimates. The degree of over/under estimation depends on the extent of social mobility and the actual values of the relative risks, both of which are unknown to us. This should not bias our findings concerning the signs of the slopes of the curves (or estimated sequences) of relative risks, however.

Men from disrupted families enter cohabitation somewhat more often (and marriage slightly less often) than men from intact families. This is in accordance with Kiernan's (1992) findings for men and women in England, and we interpret it as an after-effect of the experiences in the respondents' home of origin. In other words, men with a nonconducive atmosphere in their home of origin are more likely to view union formation as an escape option and when they do so they prefer unmarried cohabitation to marriage.

The effect of having a pregnant non-coresidential partner is another result worth elaborating. First, Table 2 indicates that men with a pregnant woman are much more likely to enter a union than men with no such woman, as is quite natural. The effect on entry into a marital union is about three times higher than the effect on entry into nonmarital cohabitation. In our initial analysis of the intensities of transition into marriage, in which the factor '*cohabitational status*' is included among the factors affecting the intensity (see Section 2.2), we found that single men with a pregnant sex partner are seventeen times more likely to initiate a union (of either form) than single men with no such woman (details not displayed here). Cohabiting men with a pregnant sex partner, on the other hand, marry at a rate of 4.7 times that of cohabiting men whose cohabitant is not pregnant. The corresponding figures for Swedish women are 10.1 and 4.2, respectively (B. Hoem, 1987). While the figures for cohabitants are consistent, the effect of a pregnancy on union formation appears stronger for single men than for single women. This probably does not reflect real differentials in union-formation behaviour, however. It is more likely that it indicates higher underreporting of pregnancies among men than among women. It is easier for a man to fail to report a child he has fathered and not really taken responsibility for than it is for a woman to leave out a child she has born.

3.3 Changes in age profiles of union formation across cohorts.

To further advance our understanding of how the pattern of age at first union formation has changed across our birth cohorts, we have added an interaction term between cohort and age to the factors in our analysis. The results are shown in Figures 5 to 9.

Shifts across cohorts in age at union formation again differ according to the type of union entered. The age pattern of marriage formation depicted in Fig. 5 shows an initial <u>increase</u> in marriage rates at younger ages between the two oldest cohorts. The age-pattern of marriage intensities in these two cohorts did not continue in later cohorts, however. Teenagers in the three youngest cohorts (born after 1950) enter marriage at higher rates than the corresponding men in the older cohorts[1]. This initial increased 'risk' is, however, outweighed by the much lower overall rates at ages beyond twenty. Furthermore, none of these cohorts show any systematic change in relative marriage rates as age increases. The overall decline in relative marriage rates between the oldest and youngest pairs of cohorts (see Table 2) has been largely due to a decline in

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1 To avoid cluttering, the curve for the cohort born in 1961-64 has been deleted from Fig. 5. Men born during this period have not much information to contribute beyond the early ages.

marriage rates at ages beyond 20 in the youngest four cohorts. (See Figs. 6 and 7.) In other words, the initial decline in the relative rate of marriage across cohorts as shown in Table 2 has has probably been a result of a decline in the rate at relatively older ages, while the upward shift in the marriage rate for the younger cohorts in the table is due to the contribution of men at the very young ages. The discrepancy between the results in Table 2 and these figures may probably be attributed to compositional effects. Figures 8 and 9, on the other hand, display a continuous rise in early entry into cohabitation for the first (oldest) four cohorts (Fig. 8) and a subsequent radical reversal for the last (youngest) two cohorts (Fig. 9), as we saw in Table 2 already.

As the right picture is already shown in Table 2, separate figures similar to Figures 6 and 7 are not included here to show age patterns of cohabitation. It is worth noting, however, that unmarried cohabitation started as a phenomenon among men in their mid- to late-twenties and only later became common practice among younger men. (Compare the peaks in Figures 8 and 9.)

Appendix A: The Analysis of Separate Decrements

The analysis in this paper is carried out by means of multiplicative hazard models. We use techniques which are now common in the analysis of life-history data.

In the analysis of separate decrements, the approach involves a multiplicative model for each transition studied. 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 marriage intensity. Let us further assume for illustration, that the intensity depends only on the three factors; Age (A), Birth Cohort (B), and Social Class (C). The model assumes that for a particular age group a, birth cohort b, and social class c, we can write the intensity function $\lambda(a,b,c)$ in the form

$$\ln \lambda(a,b,c) = \Delta + A(a) + B(b) + C(c) + AB(a,b) \tag{1}$$

where Δ is an overall average effect (grand mean), $A(a)$ is a parameter which represents the main effect specific to the age group, $B(b)$ is the corresponding main effect specific to the birth cohort, $C(c)$ is the effect specific to social class and $AB(a,b)$ is a parameter which represents the effect of an interaction between age and birth cohort.

For the purpose of expressing the relative intensity of a particular group on one of the factors considered (except the time variable), one level is selected as a baseline for that factor. Let b_0 be the cohort born in 1946-50, let c_0 be the class " Middle-level white collar employees ", and let us use b_0 and c_0 as baseline levels for factors B and C, respectively.

We make all parameters in (1) identifiable by defining $B(b_0) = C(c_0) = 0$ and $AB(a,b_0) = 0$ for all a. Thus,

$$\lambda(a,b_0,c_0) = \mathrm{Exp}(A(a)), \quad [\text{where } \mathrm{Exp}(A(a)) = e^{A(a)}] \tag{2}$$

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

$$\lambda(a,b,c)/\lambda(a,b,c_0) = \mathrm{Exp}(C(c)) = e^{C(c)} \tag{3}$$

for all a and b, is the risk in social class c, relative to the risk in class c_0. at all ages and in all cohorts. A similar formula holds for the risk in cohort b, relative to the risk in class b_0 at all ages and for all social classes.

Factors B and C in (1) above are fixed (constant) over age. The model can be extended easily to allow for explanatory variables that change in value over time. We have, for instance included the *'cohabitational status'* factor as a time-varying variable (with values of 0 for the 'single' status and 1 for the 'cohabiting' status) in part of our analysis. In the presence of one such additional variable, say S, (1) above can be extended to

$$\ln \lambda(a,b,c,s_a) = \Delta + A(a) + B(b) + C(c) + S(s_a) + AB(a,b). \tag{4}$$

The model in (4) says that in addition to factors A, B, and C, the intensity (of marriage in our case) at time a depends on the status of the individual at the same time a (i.e. on whether he is single or cohabiting at that time). 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).

A nonzero interaction term in (1) means that while $\mathrm{Exp}(A(a))$ is the age structure of the intensity for cohort b_0, each other cohort has its own age structure for this intensity. If we plot $\lambda(a,b,c_0) = \mathrm{Exp}[A(a)+B(b)+AB(a,b)]$ as a function of a for fixed b we get different curves for different values of b. Many of our diagrams and tables contain result from plots and expressions like those in (2) and (3) above.

In the absence of a nonzero interaction term, the model assumes that there is an underlying age-structure (age pattern) $\mathrm{Exp}(A(a))$ characterizing the intensity curves, that is the same for all levels of the other factors included in the model.

For each subgroup, that is, for each combination of factors except the time variable (factor), transition intensities are computed by multiplying the underlying intensity curve by the relevant parameters, one specific for each factor included in the model; and the corresponding relative intensities are found using (3) above.

The relative intensities are indicators of how often a transition occurs to individuals at a particular level of a factor, relative to the other levels of the same factor.

We have fitted separate models for each of the three transition intensities for men in the single status; namely i) the intensity of transition into marriage, ii) the intensity of transition into cohabitation, and iii) the intensity of transition into any type of first union, i.e., the sum of the two former intensities.

The Maximum Likelihood method has been used to estimate the parameters in each model. We have used the LOGLIN computer program, Version 1.64 (Olivier & Neff, 1976). The program uses occurrences and exposure matrices as input data and an algorithm called Iterative Proportional Fitting in estimating the parameters. Several support programs developed at the Demography Unit, Stockholm University, have also been used to make necessary modifications both before and after the use of the LOGLIN program.

In search of a parsimonious model for each transition, we have started with a basic model that includes only the time variable A, and subsequently extended it by adding various combinations of main effects and/or interactions of other factors. At each step the fit of the model is examined and differences between various models are tested for significance using the fact that the difference (or likelihood ratio) is approximately distributed as Chi-square with appropriate degrees of freedom.

Appendix B: Simultaneous Analysis of Competing Risks

The basic form of a competing-risks model can be written as

$$\lambda_{(i)jkml} = a_{(i)j} b_{(i)k} c_{(i)m} \tag{5}$$

for i=1,..,I;j=1,...,J;k=1,...,K, m=1,...,M and where $\lambda_{(i)jkm}$ is the hazard rate of the i^{th} risk, $a_{(i)j}$ are I arbitrary underlying hazard functions for each i and $b_{(i)k}$ and $c_{(i)m}$ are effects specific to factors B and C, respectively, on the intensity of risk i, i=1,...,I. Note that for simplicity we are assuming that there are only three factors, A, B, and C, in order to be consistent with the presentation of most of Appendix A. In the present connection, we do not include any interactions between A, B, and C.

Further, we assume that the same individuals are at risk of two 'causes of transition', namely marriage and cohabitation. We also assume that the factors other than the time variable are measured categorically, and that the time-duration dimension of each risk function is divided into J given intervals of time.

The data can be configured into matrices of occurrences (cases) and exposures, with the exposure contributed by each individual studied classified in a J x K x L matrix, and the occurrences in a 2 x J x K x L matrix.

Let the recorded exposures in group (j,k,m) be denoted by R_{jkm} and let the occurrences corresponding to each risk be denoted by $D_{(1)jkm}$ and $D_{(2)jkm}$. Then the respective likelihoods of the two competing risks are given by,

$$\Lambda_1 = \Pi_j \Pi_k \Pi_m [\mathrm{Exp}(-\lambda_{(1)jkm} R_{jkm}).(\lambda_{(1)jkm})^{D_{(1)jkm}}] \quad (6)$$

and

$$\Lambda_2 = \Pi_j \Pi_k \Pi_m [\mathrm{Exp}(-\lambda_{(2)jkm} R_{jkm}).(\lambda_{(2)jkm})^{D_{(2)jkm}}]. \quad (7)$$

The total likelihood is

$$\Lambda = \Lambda_1 \Lambda_2 = \Pi_i \Pi_j \Pi_k \Pi_m [\mathrm{Exp}(-\lambda_{(i)jkm} R_{jkm}).(\lambda_{(i)jkm})^{D_{(i)jkm}}] \quad (8)$$

Within the log-linear model, the hazard in any specified category can be represented as resulting from contributions by each of the different covariates. For example, if we had only one transition, namely first union formation and if we let the rate of first union formation be dependent on four covariates, in this case, the type of union formed (with i=1 for marriage and i=2 for cohabitation), the age of the respondent a_j (with j=1,...,J) his birth cohort b_k (with k=1,...,K), and his social class c_m (with m=1,...,M), then the likelihood function would look like

$$\Pi_i \Pi_j \Pi_k \Pi_m [\mathrm{Exp}(-\lambda_{ijkm} R_{ijkm}).(\lambda_{ijkm})^{D_{ijkm}}] \quad (9)$$

The only essential difference between (8) and (9) is that (9) contains two different exposure matrices (R_{1jkm} and R_{2jkm}) while (8) contains one and the same set of exposures R_{jkm}, though it appears twice, namely both in the product $\lambda_{(1)jkm}R_{jkm}$ and $\lambda_{(2)jkm}R_{jkm}$ (see also (6) and (7).) In other words, the exposure matrix of men initiating a union through marriage is identical to the exposure matrix of men initiating a union through cohabitation, while the occurrences matrices are different.

Otherwise, there are only notational differences between (8) and (9). If we write $\lambda_{(i)jkm}$ as λ_{ijkm} and $D_{(i)jkm}$ as D_{ijkm}, (8) reduces to

$$\Lambda = \Pi_i \Pi_j \Pi_k \Pi_m [\mathrm{Exp}\,(-\lambda_{ijkm}R_{jkm}) . (\lambda_{ijkm})^{D_{ijkm}}] \tag{10}$$

and if we define R_{1jkm} and R_{2jkm} both as being equal to R_{jkm} in (10), then (10) reduces to (9).

Therefore, we may analyze both competing risks simultaneously by fooling the LOGLIN program into believing that the cause of transition is another factor affecting the intensity of union formation together with the real observed factors. This is achieved by specifying the same exposures for all levels of that new factor (union type). This is essentially what we did above in reducing (10) to (9).

To use LOGLIN for such purposes we first append each of the two pairs of occurrence and exposure matrices corresponding to the two causes of transition (we recall that the exposure matrices are identical for the two causes) into a single occurrence (exposure) matrix. Thus, if the occurrence matrices corresponding to the two causes of transition were in files OCC1 and OCC2 and the corresponding exposures were in files EXP1 and EXP2 (where EXP1 and EXP2 are one and the same), then the new matrices, say OCC and EXP, obtained by appending the corresponding matrices, contain the occurrences and exposures respectively, organized in the right sequence, with the cause of transition (union type) as the first factor.

Then we analyze the new pair of matrices (OCC and EXP) as if the cause of transition were the first of the factors involved in the model. In our particular case, the preferred model whose estimated parameters are presented in Table 2 was of the following format:

$$\lambda_{(i)jkmnr} = a_{(i)j}b_{(i)k}c_{(i)m}d_{(i)n}e_{(i)r} \qquad (11)$$

where i refers to union type, j to respondent's age, k to birth cohort, m to respondent's social class, n to disruption status of respondent's family of origin, and r refers to pregnancy status of respondent's preunion partner.

In the illustrative three factor-model above there were no interaction terms in the model. However, LOGLIN's output looks as if there were interactions between the first factor say T (type of union entered, indexed by i) and each of the other factors (A, B, and C).

In other words, the results from the program look as if the model were not (5), but

$$\lambda_{ijkm} = (TA)_{ij}(TB)_{ik}(TC)_{im} \qquad (12).$$

It is only because we know how the outcome should be interpreted that we realize that the model is as in (5) and not as in (12).

In the usual notation of log-linear models, (5) can be rewritten as

$$\ln \lambda_{(i)jkm} = \Delta + T_{(i)} + A_j + B_k + C_m + (TA)_{(i)j} + (TB)_{(i)k} + (TC)_{(i)m} \tag{13}$$

or equivalently as

$$\lambda_{(i)jkm} = \mathrm{Exp}\ (D) \tag{14}$$

where D represents the right hand expression in (7). In view of this, the relative risks for factor B say, are of the form

$$\lambda_{(i)jkm}/\lambda_{(i)jk^0m} = \mathrm{Exp}\,[(b_k - b_{k^0}) + (TB)_{(i)k} - (TB)_{(i)k^0}] \tag{15}$$

When used properly, the INTERACT program, originally developed for use with analysis of separate causes of transition, produces four standard table panels in which each panel contains risks relative to appropriately selected reference points (baselines). In the case of simultaneous analysis of competing risks of the type considered in Table 2 of this paper, not all four panels are of use. However, one panel gives precisely what is needed in (15) above. This means that in a competing-risks model of the type used here, one of the panels used by INTERACT gives the relative risks sought for a selected factor, separately for each cause of decrement.

In the presence of an additional common factor, say F, whose effect on first-union formation does not vary across the type of union entered, the formulae corresponding to (5), (12) and (15) would be, respectively,

$$\lambda_{(i)jkmn} = a_{(i)j} b_{(i)k} c_{(i)m} f_n \tag{16}$$

$$\lambda_{ijkmn} = (TA)_{ij} (TB)_{ik} (TC)_{im} F_n \tag{17}$$

$$\lambda_{(i)jklm} / \lambda_{(i)jklm^0} = \mathrm{Exp}[f_m - f_{m^0}] \tag{18}$$

In such a case, the INTERACT program cannot be used to compute relative risks corresponding to factor F. Instead relative risks of factors F as given in (18) can easily be computed by hand.

Without F, there is one baseline hazard for each cause of transition and is computed as follows:

$$a_{ij} = \mathrm{Exp}[\Delta + T_i + A_j + B_{k^0} + C_{m^0} + (TA)_{ij} + (TB)_{ik^0} + (TC)_{im^0}]. \tag{19}$$

From this we can compute, for instance, quantities like

$$a_{2j^0}/a_{1j^0} = \mathrm{Exp}[T_2 - T_1 + (TA)_{2j^0} - (TA)_{1j^0} + (TB)_{2k^0} - (TB)_{1k^0} + (TC)_{2m^0} - (TC)_{1m^0}]. \tag{20}$$

In our case, an estimate of this quantity (from a model including more than the factors listed above) is found to be 2.77 and is shown at the top of table 2.

When a factor that is not common to both causes of transition is involved, its effects on the intensity of the cause it is related to is better determined through the analysis of separate decrements.

At last, the case where two or more factors (other than the union type) interact with each other in a simultaneous analysis of competing risks, is handled by combining the relevant factors into a single new factor and proceeding with the simultaneous analysis as discussed above. Consider the case her where we have five factors A, B, C, D and E, apart from the first factor T (union-type). The occurrence and exposure matrices are organized in the order B C D E A.

At one point in our analysis, we have considered an interaction between B and C (see section 3.4). This was achieved by first letting LOGLIN read in the matrices with B and C locked together as one combined factor, say N = (B,C). In this way we let LOGLIN believe that it reads data for four factors (in addition to the union type) in which the factor order is N D E A with factor N having the combined number of levels on B and C. We then fitted the simultaneous model in which, as before, the results look as if there were interaction between union type T and the new factor N. We used the INTERACT program to compute the four panels of relative risks and used the one appropriate panel to display patterns of the seemingly three-way interaction between T, B and C. A similar procedure was used in Section 3.3 whose results are displayed in some of our figures.

Thus one can use LOGLIN to develop an intensity model with interactions at any level of complexity. INTERACT will produce the usual four standard tables of relative intensities. The initial output from INTERACT may not look like what we need right at the outset. However, some manipulation of the intensity formula in the LOGLIN format may show that some of the output can of use.

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Table 2: Simultaneously estimated parameters of competing risks of entry into a first union, by type of union. (Swedish males born 1936-1964)*

	Union-specific effects	
	Marriage	Cohabitation
Overall rate per 10000	65.80	182.00
Overall rate (relative to marriage)	1+	2.77
BIRTH COHORT		
1936-40	2.32	0.48
1941-45	2.08	0.71
1946-50	1	1
1951-55	0.98	1.35
1956-60	1.56	1.17
1961-64	1.40	0.67
RESPON. OWN SOCIAL GROUP		
Unskilled worker	1	1
Skilled worker	1.28	1.26
Lower white collar	1.49	1.20
Middle white collar	1.75	1.18
Higher white collar	2.00	1.20
Farmer	1.96	1.34
Self-employed	1.98	0.53
Other	1.35	0.32
FAMILY OF ORIGIN		
Intact	1	1
Disrupted	0.94	1.22
PREGNANCY		
Not pregnant	1	1
Pregnant	32.20	10.07
AGE (Years)		
16-19	0.16	0.18
20-21	0.33	0.78
22-23	0.63	1.13
24-25	0.92	1.09
26-27	1	1
28-29	0.80	0.87
30-32	0.40	0.45
33-35	0.13	0.32
36+	0.08	0.22

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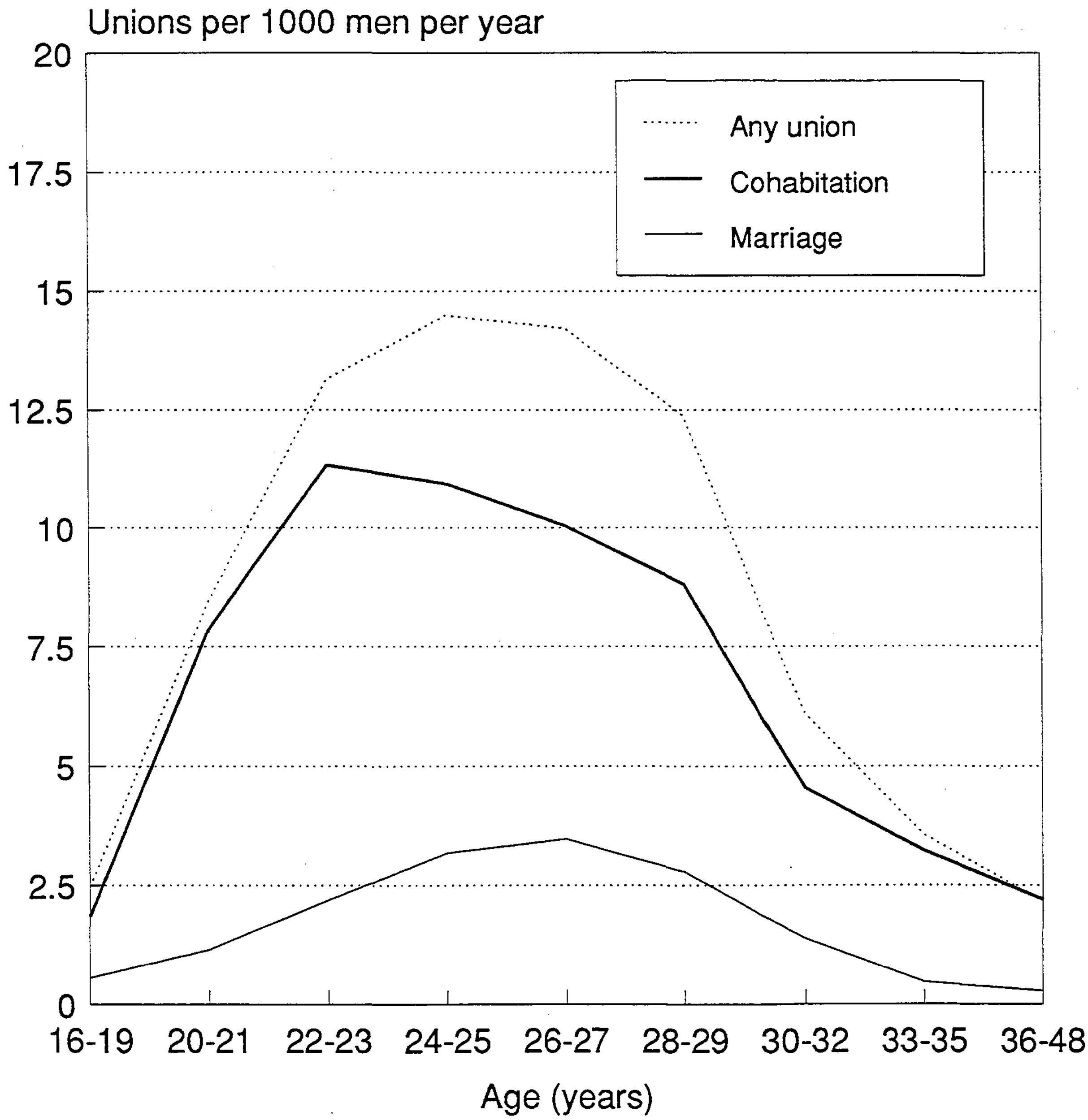
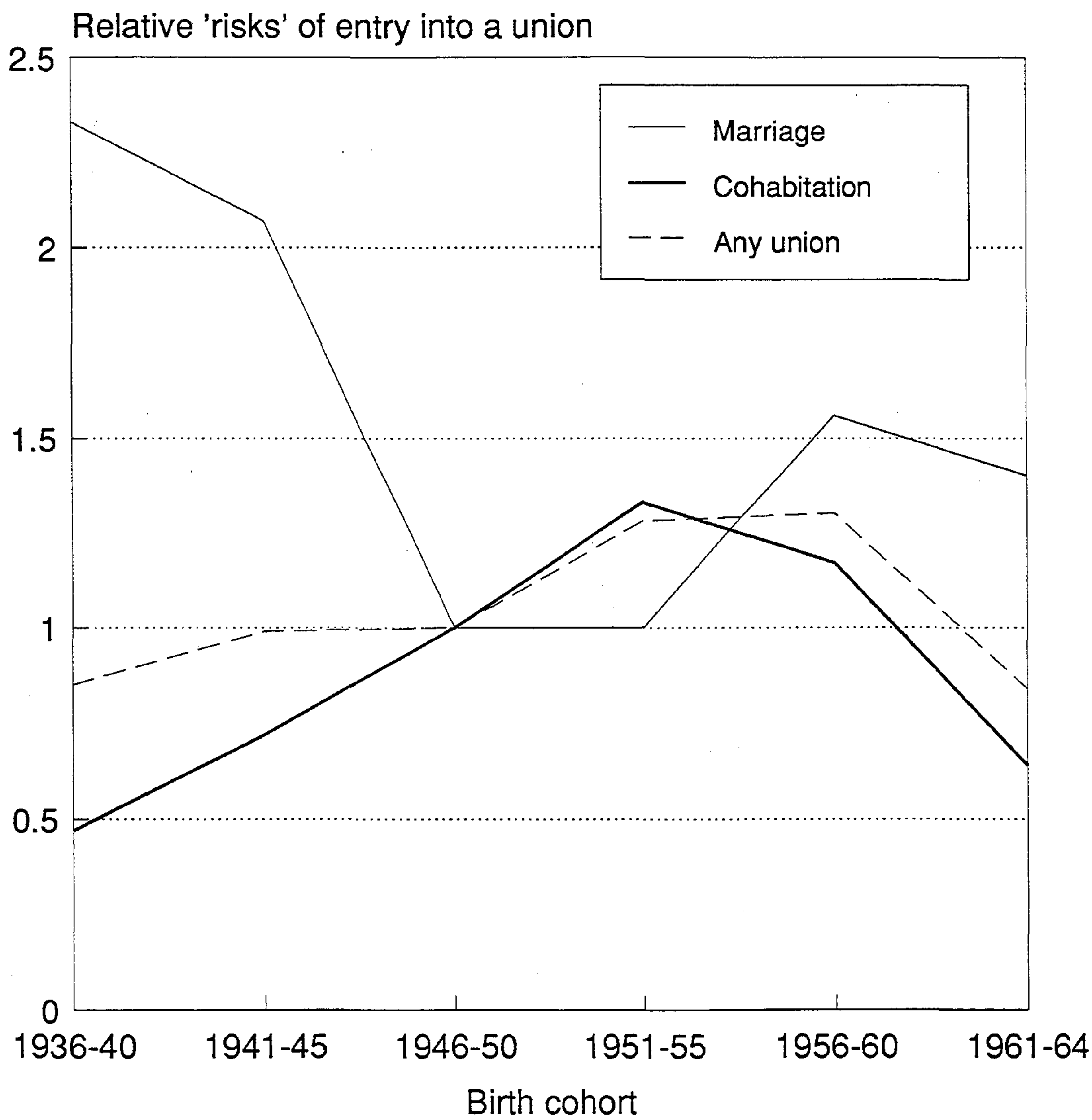
* *Only factors that are significant (from both statistical and substantive points of view) are included in this table. Social groups of the respondent's parents, size of community of his origin, and his level of education at the time of the survey, all turned out to have insignificant effects in a preliminary analysis whose results are not presented in this paper.*
+ For each factor, baseline levels are indicated by a parameter value of 1 (without decimals) in all panels of the table.

Fig. 2: Age profiles of intensities of entry into first unions, by union type. (Swedish males born 1936-1964)



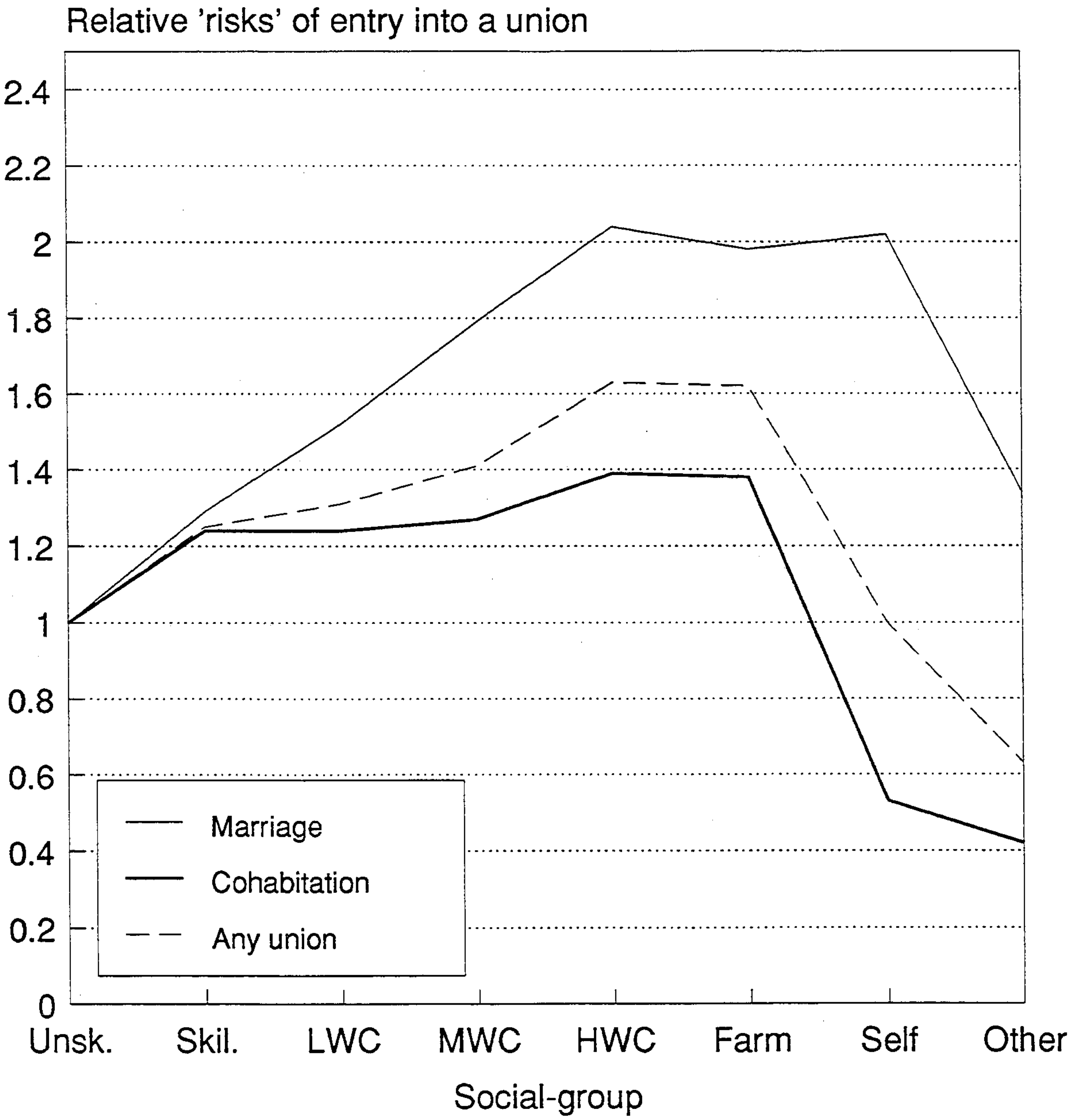
Baseline levels: Men born 1946-60, unskilled worker, from intact families, and with no pregnant partner.

Fig. 3: Cohort profiles of relative
'risks' of entry into first union,
by union type.



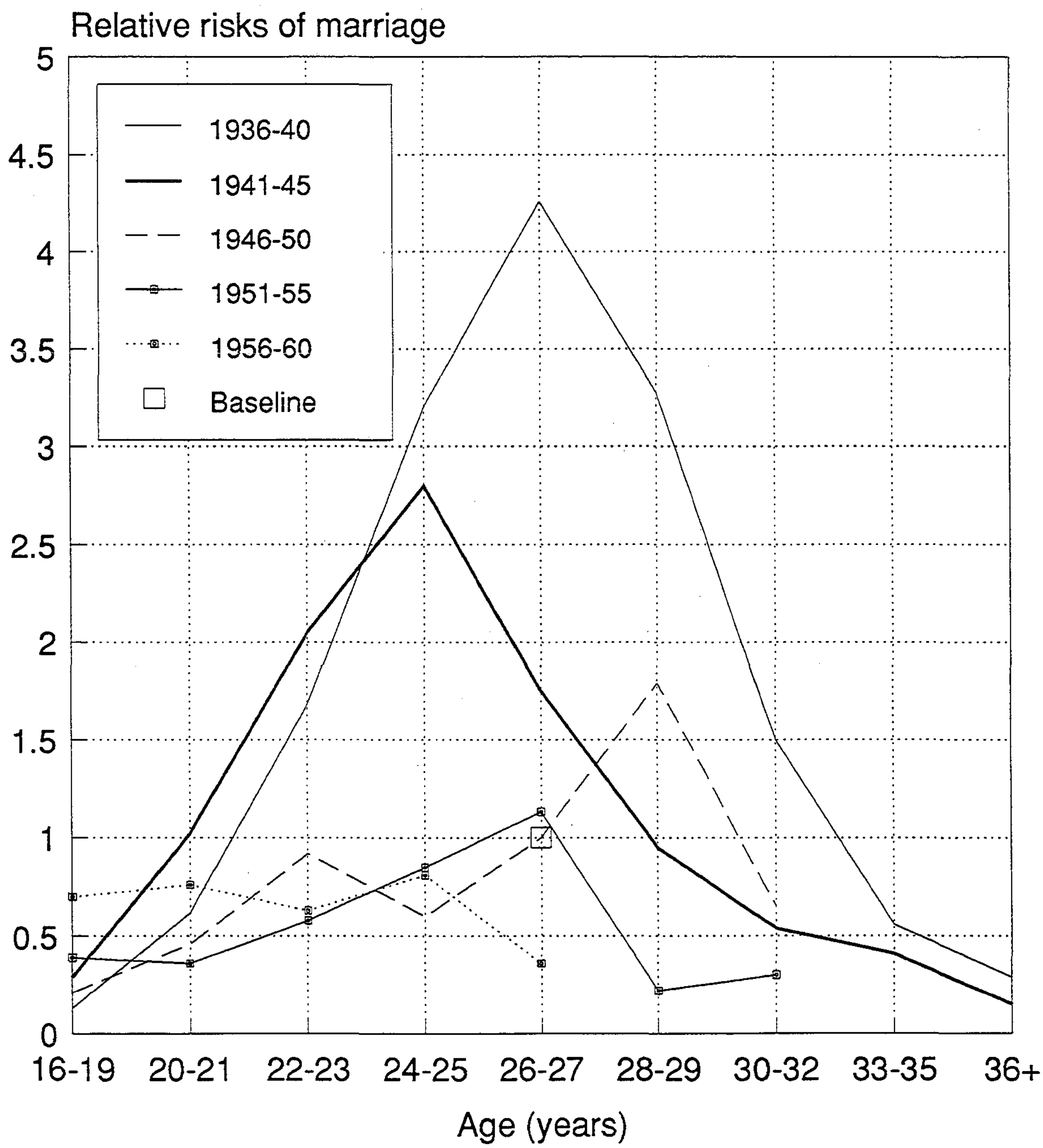
Baseline levels: Men born 1946-50,
unskilled worker, from intact families,
aged 26-27, and with no pregnant partner

Fig. 4: Social-group profiles of relative intensities of entry into first union, by union type



Baseline levels: Men born 1946-50, unskilled worker, from intact families, aged 26-27 and with no pregnant partner.

Fig. 5: Age profiles of relative risks
of entry into a first marriage,by cohort



Baseline levels: Men born 1946-50,
unskilled worker, from intact families,
aged 26-27, and with no pregnant partner

Fig. 6: Cohort profiles of relative risk of entry into a first marriage: younger age groups.

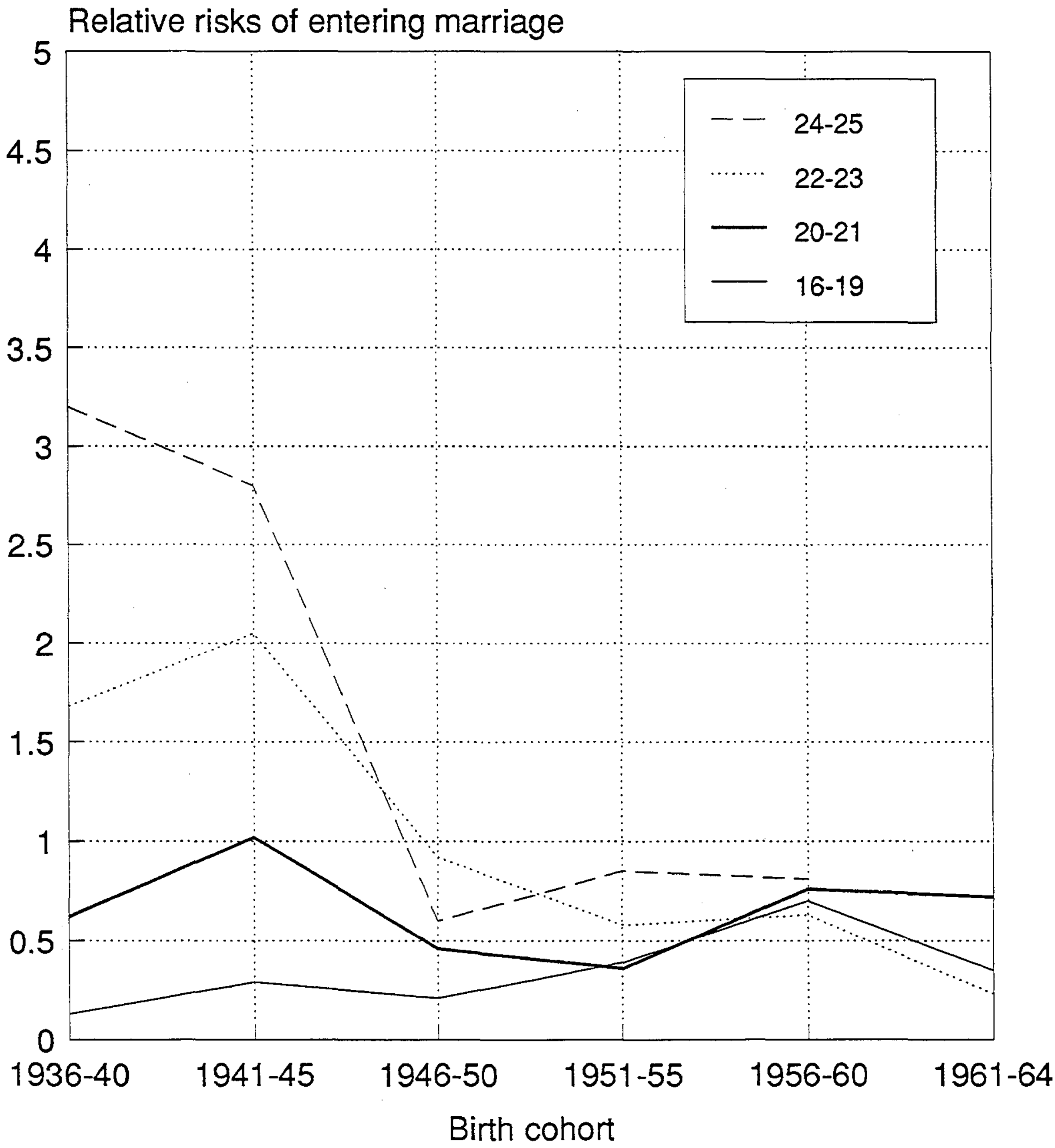


Fig. 7: Cohort profiles of relative risk of entry into a first marriage: Older age groups.

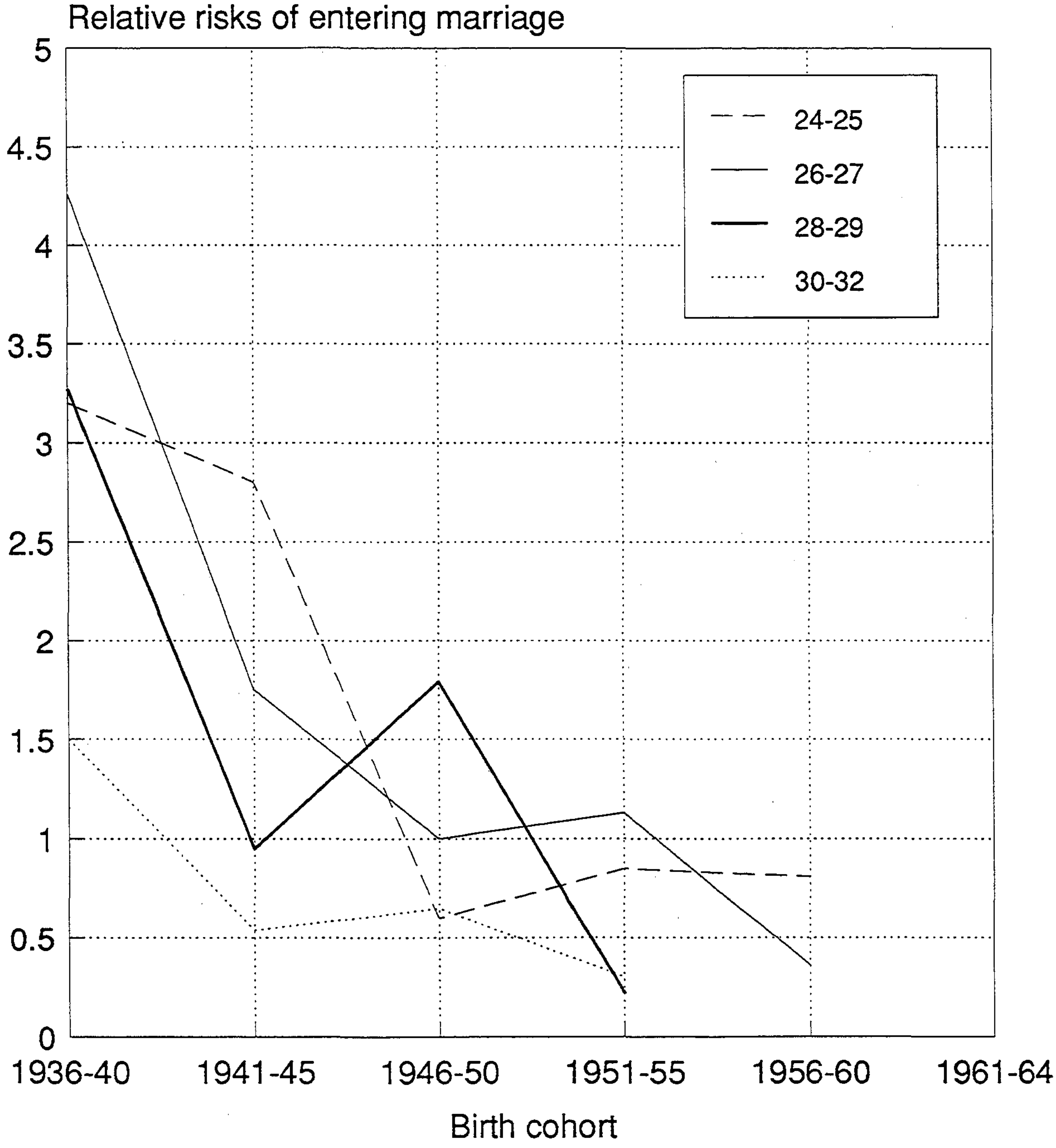
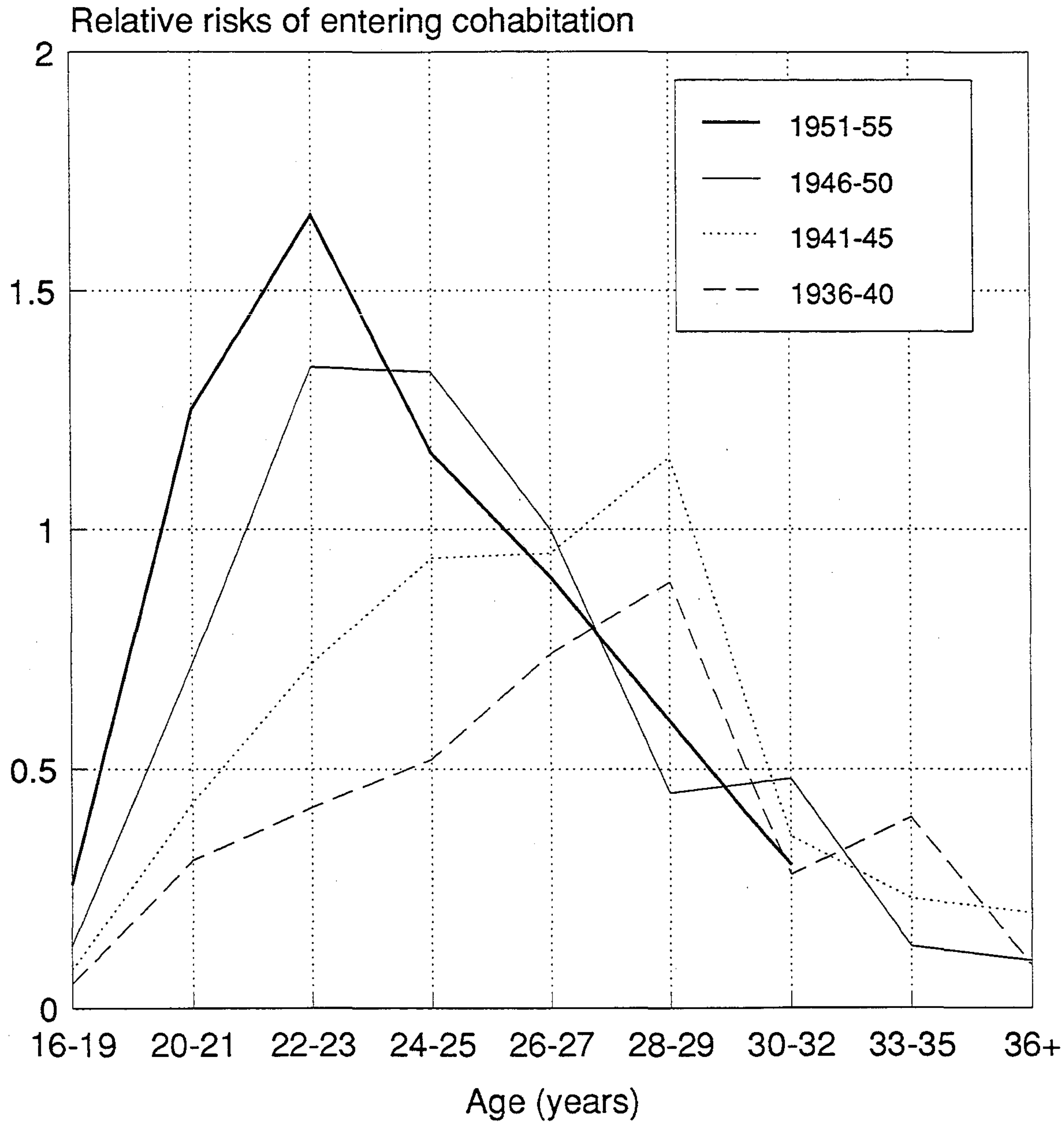
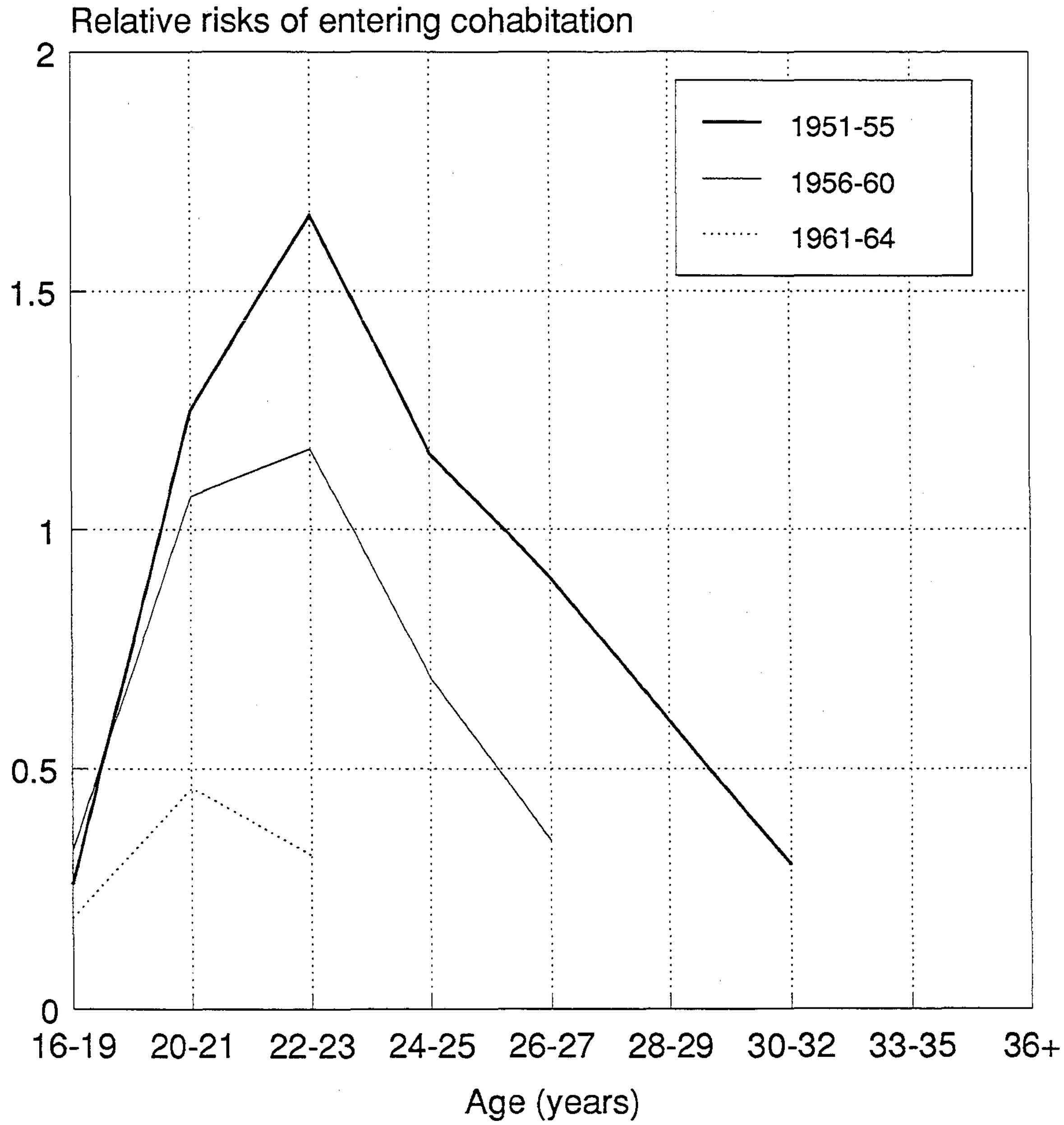


Fig. 8: Age profiles of relative risks
of entry into a nonmarital cohabitation:
the four oldest birth cohorts.



Baseline levels: Men born 1946-50,
unskilled worker, from intact families,
aged 26-27, and with no pregnant partner

Fig. 9: Age profiles of relative risks
of entry into a nonmarital cohabitation:
the three yongest birth cohorts.



Baseline levels: Men born 1946-50,
unskilled worker, from intact families,
aged 26-27, and with no pregnant partner