



The impact of the first child on family stability

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Abstract

Statistical investigations invariably show that couples run a lower risk of seeing their (marital or consensual) union break up if they have a child together than if they do not. Common explanations posit that only couples with a durable relationship have children, that responsible parents stay together for the sake of the child, and that the child itself produces an additional bond between the parents. In the inimitable terminology of economists, a child is a marriage-specific investment whose proceeds the parents do not want to lose by dissolving their union.

The insights of family counselors suggest that things may be a bit more complicated. In their experience, the arrival of the first child may serve as a destabilizing factor in the relationship between the parents, sometimes causing their union to break up. Tensions and conflicts that the partners can control as long as they only have themselves to care for, may become unmanageable when the child arrives, and the great needs of the newborn may cause new problems to surface. Unfortunately, much of the evidence for such an observation seems to be experiential and anecdotal, so the question arises whether it is caused by the counselors' special vantage point. Could the daily exposure to families that have difficulties with their internal relationships give the impression that such problems are universal when they are not? Could we be faced with yet another instance of biased observation centered on high-risk outcomes and lacking the counterbalance of less problematic cases? Will the apparent risk of union disruption in families with a newborn first child disappear in a statistical investigation based on a more balanced observational design?

The message of the present paper is that it will not, but that the risk of union disruption depends strongly on the family situation. When we control for a number of personal and structural variables (such as the mother's age, how long the union has existed, whether the union is consensual or marital, what calendar period we are in, and so on), it turns out that the risk is reduced in the first few months after the birth of the first child, but that it rises subsequently and reaches roughly the same level as before the woman became pregnant as soon as the (first) child gets to be one-and-a-half to two years old. A telltale feature is the strongly reduced risk of disruption during pregnancy, i.e., at a time when many partners with latent problems of cohesion may still have illusions about their relationship.

We speculate that these issues apparently have not surfaced in previous statistical work partly because to our knowledge no one seems to have given them sufficient attention and partly because much previous statistical investigations have concentrated on formal divorce, not actual union break-up. The delay normally caused by divorce proceedings can easily hide subtler effects of the kind that show up in the present investigation.

1. Introduction

Statistical investigations typically show that married couples have a lower risk of divorce if they have one or more children than if they are childless,² and this pattern extends to the risks of union disruption³ for couples who live in (marital or nonmarital) unions.⁴ The risk reduction is often explained partly as an effect of selection into parenthood, partly as a consequence of adaptation to new circumstances. The selection element arises because the arrival of a child signals a more durable union. People do not usually become parents willingly if they perceive their relationship as unstable in the first place; furthermore, the lack of a child may destabilize an existing union because many couples regard having children as an important part of what makes their relationship meaningful. New parents are also thought to adapt to the new and positive dimension that the child represents in their own lives, and they make greater efforts to overcome any difficulties because they feel responsible for the child. In economic theory (Becker, 1981), children are seen as similar to investment goods and parents as prone to protect their marriage-specific investments by avoiding union disruption and divorce.

Family counselors tend to tell a different story. In their view, the arrival of the first child serves in many cases to destabilize the relationship between the parents, both because pre-existing tensions and problems may come to a head and also because the new parents may discover that they have incompatibilities they never recognized before. The newborn child needs a lot of attention and creates new adult roles for its parents, so the parents may find that they have much less time for each other than they used to⁵ and their willingness to share household chores and give emotional support to each other may be put to a new and sometimes severe test.⁶ Growing dissatisfaction with the marital relationship may mature into a decision to bring the union to an end.⁷

This process seems to be well known among family counselors, and it has been described to us as one of the most painful that they experience in their professional lives. Nevertheless, we have been hard put to find any statistical report displaying a reduction in union cohesiveness caused by the arrival of the first child (beyond any deleterious effect on marital happiness). A constant exposure to problem cases and an apparent lack of systematic counter-evidence may lead to a danger of selectivity in the family counselor's experience. A statistician may suspect that an imbalance in the field-worker's case-oriented observational design may lead to a focus on high-risk outcomes that may be adequate for a counselor's practice but not for an assessment of disruption risks in the general population. In fact, Glenn and McLanahan (1982, p. 69) have argued that the presence of one or more children may deter many unhappily married couples from divorcing, at least for a time.

This may be true for formal divorce, but modern Swedish parents show little such restraint in splitting up the co-resident family (union disruption). The findings presented in this paper give support to the interpretation given by family counselors. Based on a sample survey from the general population in Sweden, collected with no particular overrepresentation of problem cases,⁸ we are able to display

² The divorce risk is particularly low when the children are very young, in the preschool ages, say. For some contributions, see Cherlin (1977), Waite and Lillard (1991), Toulemon (1994), and Andersson (1997ab). Many of the patterns we describe in this paper are very common and when they are, our citations are confined mostly to recent contributions.

³ Note that we just considered formal divorce and now consider household dissolution.

⁴ See Hoem and Hoem (1989) and many others.

⁵ See Hill (1988); Bäckström (1996).

⁶ Entwistle and Doering (1981), Cowan and Cowan (1992).

⁷ For an explication of such features, see Entwistle and Doering (1981, pp. 252-266), Clulow (1982, pp. 21-34) and White, Booth and Edwards (1986) and their references. Cowan and Cowan (1995) discuss possible interventions aimed at easing the transition to parenthood.

⁸ If anything, one may suspect that problem cases are under-represented in our data, for they should have a higher nonresponse probability than run-of-the-mill cases do.

patterns in union-disruption risks that are largely compatible with the description that the counselors give. We find that disruption risks are the lowest during a woman's pregnancy, that they are low also during the first few months of a child's life, but that they subsequently rise as the child grows older, largely to re-attain the levels of a childless woman already when the child has become one-and-a-half to two years old. If any couples felt that a child would save a shaky relationship, they must soon have found out that they were mistaken. The low risk during pregnancy can be understood in the light of expectations held by prospective parents before any incompatibilities have become manifest after the birth. The subsequent rise fits the notion that it takes a little time before the decision to end the union becomes mature.⁹

This pattern is obscured in data where a formal divorce is used as the event of marriage dissolution, as it is in data from population registers. Even in Sweden, where divorce is available on demand and where the waiting time for a divorce is only six months (and that is when the divorce is contested or there are children present), the divorce risk is much reduced by the arrival of the first child and it remains low during the child's first three years of life (Andersson, 1997a, Figure 1).¹⁰ The process of obtaining a formal divorce introduces a delay that is sufficient to conceal the finer features of the time pattern of real break-up risks. This makes the otherwise excellent Swedish register data less adequate for our present purposes.

In all of our investigation, we control for the impacts of other demographic factors, such as the woman's age, whether she lives in her first or second union, how early in the union she had her baby, whether she is married or a cohabitant, and whether she became pregnant or even had her baby before marriage. Since divorce risks have risen over time (Anderson, 1997a), we have also controlled for calendar period. We display the effect of each factor and have studied its stability over time. Our concern is to establish the basic disruption pattern in the disruption risk and not to study its educational or social gradient, so we have not included factors that represent the woman's educational level or occupational or social status.¹¹

2. Data, method, and first results

Our analysis is based on the data from the second Swedish Family Survey, conducted by Statistics Sweden in 1992/93. The survey collected data from a random sample of 3318 women born in the years 1949, 1954, 1959, 1964, or 1969 and a smaller random sample of men. We are not now concerned with a comparison of men and women, so we have concentrated on the sample of women. The data set contains histories of the formation and dissolution of marital and nonmarital unions, of childbearing, education, and labor-force and other activities up to the time of the interview, as well as a host of other information.

We have eliminated women who did not report any marital or nonmarital union, as well as women who reported a first child before entry into a first union. Our sample consists primarily of 2592 women who ever entered a first union without prior childbearing. We have followed each such woman month by month through her first union and have noted any demographic events recorded in it up to

⁹ Its timing is also remarkably well in line with the return of many new mothers to their jobs in the labor market in Sweden, a stage at which one may suspect that problems connected with chore-sharing at home may be particularly acute. Such an interpretation is not supported by our data, however, for the disruption-risk profile is essentially the same for respondents with and without a job. The profiles are also essentially the same for respondents who are on maternity leave or are housekeepers on the one hand, and for those who are not on the other.

¹⁰ The divorce risk subsequently rises almost to the level pertaining to the childless well before age of entry into school, unless a second child is born in the meantime. The pattern is very similar in France (Toulemon, 1994, Figure 9).

¹¹ Duvander (1996) and Strandberg (1997), who have analyzed the same data with focuses different from ours, did include some such factors. They did not find much effect of educational level or social origin in disruption risks in either first or second unions. Strandberg (1997) found that students had a particularly high disruption risk in second unions, and women whose occupational behavior suggests a family orientation had lower disruption risks than other women.

the time of the interview. We have used the information up to a given censoring time, defined as the time of the interview or of the occurrence of the first of the following events if it occurred before the interview:

First, suppose a respondent had a first birth in her first union. We have then censored her record at the time of that birth if it resulted in twins. We have also censored her record at the time of adoption if the first child was adopted.¹² If not, we have censored her record when she became pregnant with any second child in the union.¹³ Otherwise, we have treated the termination of the union as a censoring event if it ended because her husband or partner died.

Secondly, each record is also censored when the first child reaches age four, because most women who have a second child, have at least become pregnant with it by that time, and the group of those who do not becomes too small for our analysis.

If a respondent has ever entered a second union while still childless (487 cases), we have followed her through her second union as well and have treated it in the same manner as the first union. We have not pursued respondents beyond any second union. There are too few cases.

Our analysis thus essentially concerns disruption risks for childless women and mothers of one child below age four, for women who live in first and second unions that they entered at parity 0.¹⁴ Union formation and union disruption are self-reported by the respondents.

Our method of analysis of the subset of data selected in this manner is known as intensity regression or hazard regression, or alternatively as (multivariate) indirect standardization. In our application, it is a way of accounting for the simultaneous influence of several factors on disruption behavior. To describe this procedure, let us start with a relatively simple situation that has many though not all of the features that we need. Consider a woman of parity 1 who lives in her first union, which she started at age i . Suppose that she married without cohabiting first, suppose that she had her (first and so far only) child in the union's j -th year, and suppose that we consider her disruption risk at some time in calendar year k , at which time her child is t months old. Let us specify her disruption risk as, say,

$$\lambda_{ijk}(t) = b_i c_j d_k f_t.$$

The coefficient b_i measures the specific effect of starting the union at age i , c_j measures the effect of waiting until the union's j -th year to have the child, d_k measures the period effect of calendar year k , and f_t measures the effect of the child's age t . For our analysis, we have grouped the women's starting ages, the calendar years, and so on, as indicated in Table 1, which also contains maximum-likelihood estimates of the coefficients b_i , c_j , d_k , and f_t . The table can be read as follows:

First disregard women who started their first union without marrying directly.¹⁵ Among those who started their first union as a marital and not as a marital union, we have selected a baseline group consisting of women who entered their first union at age 20-23 ($i=2$) and had their first child in the union's second year of existence ($j=2$). These factor levels appear with a "relative risk" of 1 (without decimals) in the table. In any year no later than 1973, women in the baseline group have a disruption risk of f_t when the child is in age group t , where estimates of f_t for the various age groups t are given in

¹² We suspect that the post-partum behavior of the few mothers who bore twins or who adopted their first child is different from comparable mothers who had single births of biological children.

¹³ Actually, we have censored the record seven months before the second birth because that is when she must have known that she was pregnant.

¹⁴ We say that a woman has parity n if she has given birth to n live children.

¹⁵ The subscript 1 in $\lambda_{ijk}(t)$ indicates that the woman married without cohabiting first. Below, we will use a subscript of 2 to indicate that she entered a consensual union at initial union formation.

the final panel of the table.¹⁶ The risk in subsequent calendar years appears by multiplication of f_t by the corresponding estimated value of d_k in the next-to-the-last panel. We see, for instance, that for women in the baseline group, the disruption risk was estimated as $1.44 f_t$ in the second calendar period. This means that at any age of the first child, the risk rose by 44 percent between the first and second periods, i.e., from before to after a divorce law reform which took effect at the beginning of 1974. We also see that the estimated risk rose further in subsequent years.

The other coefficient estimates in the table have similar interpretations. For instance, the value $b_1 \equiv 2.87$ for the youngest group of starting ages shows that starting the first union as a teenager gives a woman a strongly elevated disruption risk, and the sequence of values for c_j shows that the disruption risk is reduced if she waits beyond the union's first year of existence before she has her first child.¹⁷

We extend this specification to women who began a consensual rather than a marital union when they first started, by introducing a further multiplicative coefficient a and specifying their disruption risks similarly as $\lambda_{2ijk}(t) = a b_i c_j d_k f_t$. The second row of Table 1 gives a maximum-likelihood estimate of 4.72 for a , which indicates that the estimated disruption risk for women who start union formation by entering a consensual is almost five times as high as for those who start by forming a marital union. We may write the two separate formulas for $\lambda_{1ijk}(t)$ and $\lambda_{2ijk}(t)$ above as a single formula by introducing the notation $a_1 = 1$ and $a_2 = a$, to get $\lambda_{hijk}(t) = a_h b_i c_j d_k f_t$, for $h = 1, 2$. Coefficients like a_2 and b_i are called relative risks because $a_2 \equiv \lambda_{2ijk}(t) / \lambda_{1ijk}(t)$ for any combination of i, j, k , and t , and $b_i \equiv \lambda_{hijk}(t) / \lambda_{h2jk}(t)$ for any combination of h, i, j, k , and t . (For the latter example, recall that $i=2$ is the baseline level of our second factor, i.e., the impact on the disruption risk of any starting-age group i is found by comparing the group to the one that has starting age 20-23.)¹⁸

Baseline risks and relative risks like those of Table 1 are easily converted to estimated probabilities of union disruption. We can compute, for instance, that an estimated 1.3 percent will experience union disruption before their first child's fourth birthday if we consider women who (i) were married at the time of the first birth and of age 20-23 at marriage formation, (ii) had the birth in their first union's second or third year, (iii) had no second birth in the meantime, and (iv) for whom the disruption rates of the periods before 1974 pertain. This does not sound like such a high risk, but it rises to 6 percent if the disruption rates of cohabiting women pertain throughout those four first years of their child's life, and to almost 14 percent if in addition the rates for 1988-92 are used. In a worst-possible scenario, we estimate that almost half the women¹⁹ will experience a disruption of their first union before their child's fourth birthday if we apply the disruption rates for 1988-92 for cohabitants who started their union as teenagers, had their first birth in the first year of their union, and had no second birth before that fourth birthday. A disruption risk of this size order highlights the seriousness of the situation for the real high-risk groups.

¹⁶ Note that the estimated values of f_t are given as the number of disruptions per 10 000 person months of exposure to the risk of disruption.

¹⁷ The decline in the disruption risk may be caused by the increasing maturity of the partners in the union, or perhaps also by a selection process that weeds out the less cohesive partnerships before the arrival of the first child.

¹⁸ The final column of Table 1 shows how the exposure time is distributed over the levels of each of the factors included. Note in particular how most of the time was spent in unions that did not start as marriages. The total exposure time in this part of our data set was of 42 062 person months altogether, corresponding to 1532 cases, of which only 30 started their first unions after age 28. None of the latter had a recorded union disruption (noted as "no cases" in Table 1).

¹⁹ We compute this estimate as follows: Multiply each baseline risk in the final panel of Table 1 by the length of the corresponding interval and add up to get the figure 0.0130776. Pick out the largest relative risks 4.72, 2.87, 1.55 and 2.36 from the first four panels of the table. The four-year disruption risk for the corresponding factor combination is then estimated to equal $1 - \exp(-0.0130776 \times 4.72 \times 2.87 \times 1.55 \times 2.36) \equiv 0.476$, or 47.6 percent. This kind of illustrative synthetic computation is commonly used to illuminate the consequences of risk tables like ours.

We have introduced the model described above mainly to explain our mode of analysis. The model picks up some major features of the pattern of disruption risks in the group of women we have applied it to, but more complex models are needed to really address the issues we have set out to investigate. To study the impact of the arrival of a woman's first child on the cohesiveness of her (marital or nonmarital) union, we need above all to also include life segments when the woman is childless. We also need to take into account the fact that a union which is consensual at the outset may be turned into a marriage later, for we can expect this to have consequences for the disruption risk during the time that follows marriage formation. It is a standard finding that women who cohabit, or become pregnant before marriage, or even have a baby that early, have elevated disruption risks in marriage. We have already indicated that we want to include second unions in addition to first unions in our analysis. Finally, since people mature as they grow older, it may also be sensible to replace a woman's age at union formation by her current age in any month of observation. In what follows, we will include these features by extending our initial model specification.

3. Inclusion of life segments before first birth, and inclusion of second unions

Recall that we follow each respondent month by month through her first (and possibly second) family union. To extend the simple model of Table 1, we replace its fixed age variable (namely, age at union formation) by the respondent's current age, updated continuously to any month of analysis. We add the factor "union order" to separate first from second unions, and we include union duration to pick up any separate effect of the current length of the union in any month. To include life segments before first birth,²⁰ we introduce a new factor that we call "motherhood status" and give it the following levels: Before the respondent becomes pregnant with any first child, we assign the motherhood status "childless" to her. During her pregnancy, if any, her motherhood status is "pregnant".²¹ In the month when her first child is born, her motherhood status is "mother with child at age 0 months". In subsequent months, her motherhood status is "mother with child at age t months" for $t = 1, 2$, and so on. We group these months and the months of union duration as indicated in Table 2, which contains maximum-likelihood estimates of the relative risks of each factor. We have also included a factor to represent the woman's current civil status and her civil-status history (see the fifth panel in Table 2), but we will explain and discuss it later.

Table 2 shows that women in their second unions have higher disruption risks than comparable women in first unions;²² in fact their overall super-risk is estimated to be 67 percent. There is a strongly negative gradient in the effect of the woman's own age, which reflects a decline in disruption risks as respondents mature. As in Table 1, disruption risks are seen to increase over the years for which we have observations.²³

Our main issue is addressed by means of the variable we have called "motherhood status". Table 2 shows that the disruption risk drops to a quarter of its previous level when a woman becomes pregnant with her first child ($0.65/2.51 \approx 0.26$). When the child is born, the risk rises from that low level by a factor of 1.5 ($=1/0.66$), and it increases further as the child becomes older, to such an extent that the risk gets back up to the level of childless (and non-pregnant) women already when the child is

²⁰ Recall that we have eliminated women with one or more births before union formation, so in our data segment, all births occur in the union in question.

²¹ In practice, we have assigned that status to her during the seven months just preceding her first birth.

²² The word "comparable" indicates that this comparison is made for women who have the same level on all factors other than the factor "union order". All of our comparisons are made under a similar assumption of *ceteris paribus*.

²³ The final column in most panels of Table 2 and the figures in parentheses in its fifth panel show how the exposure time is distributed over the levels of each factor. Including life segments before the first birth almost quadruples the number of person months of exposure, to 156 572 months for 2592 first unions and 487 second unions. There were 855 disruptions in all.

about eighteen months old. We conclude that the disruption risk is particularly low only during the months when the woman knows she is pregnant with her first child, and that the subsequent period of risk reduction is restricted to the first months of the child's life. By the time many that the child is about two years old, overall disruption risks are largely unaffected by the child's arrival. This pattern holds in any given calendar period, and it repeats itself across union orders, across age groups, and across various specifications of a woman's civil status.²⁴ It is also independent of a woman's labor-force status.

Given the small size of our data set, there is no need to pay much heed to the difference between the observed relative risk of disruption for childless women on the one hand and the risk for mothers of a child aged above some eighteen months on the other. It is actually remarkable that the risk for the first of these groups is not substantially higher than for the latter, for there must be considerable selection of people with a low personal disruption risk out of the group of childless couples and into the group of parents. Suppose that we could determine an unobserved-heterogeneity factor to represent an underlying level of disruption-proneness, specific for each couple. Suppose further that the couple's risk in any life situation is this heterogeneity factor multiplied by a parameter that represents the specific life situation, the latter parameter being the same for all comparable couples. Then surely the distribution of the heterogeneity factor would be loaded more toward high values among the childless couples than among the parents; in fact this is what the selectivity of entry into parenthood would mean. The empirical risks that we use to compute the relative values in Table 2 are averages over the heterogeneity distribution. If each couple had the same risk when they were parents of a two-year-old (say) as when they were childless, then this empirical average would be lower for the parents than for the childless couples. If they had a lower risk as parents than during the childless stage, then the empirical average should be lower still. Since the two empirical averages are of the same size order instead, the life-situation parameter of such parents must have a higher value than the corresponding parameter for the childless. In other words, if we could control for the underlying heterogeneity of our couples, it would be revealed that couples must have a higher disruption risk when they are parents of one child of an age around two years than when they are childless.²⁵ Quite contrary to being a stabilizing factor in the union, the child destabilizes it²⁶ after a rather brief initial period of reduced disruption risks.

²⁴ Technically, we investigate such a feature by including an interaction between motherhood status and another factor, inspecting a plot of the interaction, and testing whether the interaction significantly improves the model fit to the data. In no case did such an interaction improve the model significantly. These computations are not displayed here.

²⁵ If we strip down an underlying mathematical argument to its bare essentials, the disruption-proneness would be represented by some random variable Θ , a value of which is assigned to each couple, and the couple's disruption risk in life situation i would be $\varphi_i \Theta$ for some (nonrandom) parameter φ_i . If $i=0$ represents childlessness, $i=1$ indicates that the woman is pregnant with her first child, and $i=2, \dots, 6$ represent couples with a child at the various ages listed in the fourth panel of Table 2, then empirical disruption risks for childless couples would be estimates of $\varphi_0 \theta$, where $\theta = E(\Theta | \text{childless})$, and empirical disruption risks for parents with a two-year-old child (say) would be $\varphi_5 \theta'$, where $\theta' = E(\Theta | \text{parent})$. Our empirical risks in Table 2 suggest that $\varphi_0 \theta = \varphi_5 \theta'$ to a fair approximation, and our theory about the selectivity of entry into parenthood would mean that $\theta' < \theta$. Therefore, $\varphi_5 / \varphi_0 = \theta / \theta' > 1$, or $\varphi_5 > \varphi_0$. In other words, when we correct for the presence of heterogeneity, parents of a two-year-old child would have higher disruption risks than corresponding childless couples.

As Tore Schweder has pointed out in an oral discussion, this argument could break down if there were massive underreporting of childless unions that have ended in disruption. This could lead to the underestimation of the parameter φ_0 , and in reality we could have $\varphi_0 \theta > \varphi_5 \theta$ instead of $\varphi_0 \theta = \varphi_5 \theta'$. Even though $\theta' < \theta$ would still hold, we would not then know whether φ_5 were smaller or larger than φ_0 . Given the eagerness with which our respondents reported their unions, however, this objection does not seem very salient.

²⁶ As the first child grows older, the group of parents who remain at parity 1 instead of having their second child, will be selected progressively toward those who have a high underlying disruption-proneness. This would tend to raise the empirical relative disruption risk that we observe. Given normal intervals between the first two births in our population, however, this feature cannot have been prominent at the salient ages of the first child in Table 2.

4. The effects of civil status and premarital cohabitation, pregnancy and childbearing

In the model of Table 2, we have also included a factor that picks up the effect of a woman's civil status (i.e., whether she lives in a marital or a nonmarital union), and we update it every month during the period of observation. For married women, we combine this factor with the demographic circumstances at marriage formation, in that we indicate (i) whether she converted her union from a nonmarital to a marital one or whether she married without premarital cohabitation, (ii) whether she was pregnant at marriage, and (iii) whether she already had a child before marriage. This produces the combined factor in the last-but-one panel of Table 2.

Since we have eliminated respondents who had a first child before the current union, married women in our data segment who were mothers at marriage, had their first child during a premarital phase of that union.²⁷ This makes it impossible for a married woman in our data segment who did not cohabit before marriage to have parity 1 at marriage formation, and the corresponding cell in the lower-rightmost ("south-eastern") corner of the civil-status panel is blank. Similarly, other cells of the panel are blank when they represent a meaningless or impossible combination of the sub-factors involved. (We have written the civil-status panel in Table 2 as a five-by-three table to see the risk pattern more easily. In reality, the model simply includes a single factor with the six levels represented by the non-blank cells in the panel.)

The civil-status panel in Table 2 shows that consensual unions are much less cohesive than marital unions, and that premarital cohabitation seems to increase the disruption risk after marriage. Furthermore, pregnancy at marriage seems to increase the marital disruption risk somewhat and premarital childbearing seems to increase it further. These patterns are what we have become used to seeing in such investigations.²⁸

5. Concluding remarks

An appreciable segment of the population has a considerable risk of union disruption even when they have a very young child. The segment consists mainly of cohabitants, women in their second unions, and women who have married in their first union after premarital cohabitation and premarital childbearing. Disruption risks increase as the child grows older and soon reaches a level similar to that of comparable childless women.

Social workers we have talked to during our work with this study have the impression that the arrival of the first child may be a particular strain for those who have lived together very briefly before the first birth (because the partners really never got to know each other first) or very long before the birth (because they got used to living together without the interference of a very small child). We do not find elevated union-disruption risks at very brief durations of co-residence, nor do we see any tendency for them to increase at long durations. We have not been able to locate any pattern in risks by union duration that is particular to the high-risk groups either.

Our family-counselor contacts have told us of an increasing demand for their services over time. Such an increase may have several components. Couples who feel they are in trouble may seek advice more readily because of a greater awareness that it is available, because of an increasing belief that it may be helpful, and also because of the increasing complexity of post-disruption arrangements

²⁷ In fact, they also live in their first union.

²⁸ See, for instance, Andersson (1997a) and his precursors. Bennett et al. (1988), Hoem and Hoem (1992), and several subsequent authors have shown that premarital cohabitation signals values and experiences conducive to an increased disruption risk in marriage. This effect dominates anything that prevents shaky nonmarital relationships from being turned into marriages. Strictly speaking, however, none of the relative risks 0.86, 0.95, 1.14, and 1.34 in the civil-status panel in Table 2 are significantly different from the level 1 of the baseline group. The contributions from these four cells in our data set are too small for safe conclusions.

since the introduction of joint custody. The period effects in our Tables 1 and 2 show that the growing demand also reflects an appreciable increase in union-disruption risks. This increase seems to have hit all groups in a similar fashion. We have not been able to discern any particularly strong (or weak) increase for the high-risk groups among mothers of a small child.

The lack of special patterns in disruption risks, by union duration, by age of child, or over calendar time, does not necessarily mean that they do not exist. With the strong subdivision of our data set implied by our rather plentiful set of covariates and with the very skewed distribution of our data over the factor levels (indicated by the percentages in our tables), interaction terms of the kind needed to locate special risk patterns for special groups are so plagued by random variation that the noise may drown any signal contained in the data. Much larger data sets, and primarily different kinds of data, are needed if we want to make much further progress in assessing the importance of the substantive issues raised here. In addition, it would be very helpful if we could find out to what extent the quick rise in the disruption risk to pre-parenthood levels after the initial drop during pregnancy is present in other populations. Given all the other commonalities in disruption patterns across Western nations, it is hard to believe that this issue is particular to Sweden.

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Table 1. Risk of union disruption for mothers in a first union, by age at union formation, age of first child, etc.

Factor	Relative risk	Percent of <u>exposures</u>
Status of union at union formation		
	$\{a_h\}$	
marriage	1	8
cohabitation	4.72	92
Age of mother at union formation		
	$\{b_i\}$	
-19 years	2.87	44
20-23	1	43
24-28	0.92	12
29-35	(No cases)	1
36+ years	(No cases)	0 ²⁹
First birth occurs in the union's		
	$\{c_j\}$	
1 st year	1.55	17
2 nd year	1	20
3 rd year	1.00	17
4 th year	0.46	14
5 th year	0.23	9
6 th or later yr	0.24	23
Period		
	$\{d_k\}$	
-1973	1	14
1974-77	1.44	20
1978-82	1.81	24
1983-87	2.82	22
1988-92	2.36	20
Baseline risk of disruption		
	$\{f_t\}$ per 10 000 person months	
Age of first child		
0- 5 months	1.96	
6-11 months	2.66	
12-17 months	2.32	
18-23 months	3.94	
24-47 months	2.73	

²⁹ A zero means that the item is less than 0.5 percent.

Table 2. Risk of union disruption for women in first or second union, by current (own) age, age of first child, etc.

Factor	Relative risk	Percent of exposures	
Union order			
first	1	87	
second	1.67	13	
Period			
-1973	1	12	
1974-77	1.25	17	
1978-82	1.65	23	
1983-87	1.88	24	
1988-92	2.25	24	
Woman's current age			
<20 years	1.40	14	
20-23	1	41	
24-28	0.73	34	
29-35	0.51	10	
36+ years	0.51	1	
Motherhood status			
childless	2.51	65	
pregnant	0.65	7	
one child at age			
0- 5 months	1	6	
6-11 months	1.53	6	
12-17 months	1.49	5	
18-23 months	2.86	4	
24-47 months	2.53	7	
Current civil status			
Status at marriage	currently	married	married, no
	cohabiting	after	premarital
		cohab.	cohabitation
	2.99 (74) ³⁰		
parity 0			
- not pregnant		1 (15)	0.86 (5)
- pregnant		1.14 (3)	0.95 (0)
parity 1		1.34 (3)	
Baseline risk of disruption			
	per 10 000		
	person months		
Duration of union			
<9 months	4.18		
9-17	6.17		
18-29	6.54		
30-47	6.03		
48-71	6.89		
72-95 months	5.54		

³⁰ Figures in parentheses represent percentages of the total time of exposure to union disruption.

