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Gender and family stability: Dissolution of the first parental union in Sweden and Hungary

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Abstract

The increasing trend of partnership disruption among families with children in recent decades has been accompanied by substantial changes in traditional gender roles in industrialized countries. Yet, relatively little is known about the effects of changing gender relations on family stability in the European context. In this paper, we study such gender influences at the familial and societal level in Sweden and Hungary between the mid-1960s and the early 1990s. We focus on the disruption of the first parental union (i.e. the union in which a couple's first child was born). Our analysis is based on data extracted from the Swedish and Hungarian Fertility and Family Surveys of 1992/93. We use the method of hazard regression. The results suggest (i) that the establishment of the dual-earner family model influences family stability only if it is accompanied by some changes in traditional gender relations within the family, and (ii) that women's and men's labor-market behavior have different effects in spite of the relatively long history of women's (also mothers') labor-force participation in both Sweden and Hungary.

1. Introduction

In recent decades, family stability has decreased substantially in most industrialized countries. This was seen both in increasing divorce rates and in the growing prevalence of non-marital cohabiting relationships, which are usually more fragile than marriages (Da Vanzo and Rahman 1993, Kaa 1994). Partnership dissolution among families with children has also become more and more common.

As the demographic changes have been accompanied by greatly increasing labor-force participation among women, even among mothers with young children, theoretical explanations of new family patterns have focused mainly on women's increasing economic independence (Becker 1991) but also on ideational changes such as growing individualism and other value shifts (Lesthaeghe 1983, Kaa 1987). Although these influential theories point to the importance of gender relations for family stability at least indirectly, it has not been studied much in demographic research except for the US.

In spite of important changes in the gender context all over the developed world, we thus know relatively little about the mechanisms of the potential interplay between family stability and changing gender relations linked to the interrelationship between women's rapidly increasing employment rates and changes in (the demands on) men's involvement in childrearing and domestic responsibilities. Also, our knowledge about whether men's characteristics have different effects than women's characteristics on family dissolution is rather limited. Such gender differences are likely to exist based on the gendered nature of family life. Hence, the purpose of this paper is two-fold. First, we are interested in the impact of changing gender relations on family stability. Second, we intend to shed more light on gender differences that may affect family disruption.

2. Theoretical framework

2.1 General framework

In our attempt to study the potential influence of changing gender relations on family dissolution we use a theoretical framework based on previous research on the link between union break-up and (i) women's labor market work, (ii) the household division of labor in the light of theories about distributive justice, and (iii) gender-role attitudes, as in what follows.

In the extensive (and mainly American) literature women's employment has been associated with marital instability for a long time, based on three types of arguments. Women's labor-force participation destabilizes the marriage (1) by overthrowing traditional marriage norms, (2) by facilitating divorce in case of conflicts in the relationship, or (3) by taking women away from their traditional responsibilities at home, which in turn generates conflicts between the spouses. (For an overview see Greenstein 1990). A number of studies have found that the risk of family dissolution increases with the number of working hours per week for women. (For an overview see Greenstein 1995). The relationship between these factors is likely to be more complex, however, as also work schedules, gender ideology, and women's perceptions of the fairness of the division of domestic responsibilities within the family seem to have an important (direct or indirect) influence (Blair 1993, Greenstein 1995, Presser 2000, Sanchez and Gager 2000). Such perceptions are shaped by ”comparison referents” as women evaluate their own situation by comparing themselves to other women or by comparing their partner to other men in terms of sharing domestic work within the family (i.e. within-gender comparisons) (Thompson 1991).

These findings illuminate the effects of gender relations on union stability at the family level, at least in the US context. It is very likely that there is an interchange between societal-level gender arrangements and family-level gender relations. As women and men share the task of economic provision for the family, women's traditionally sole responsibility for domestic work becomes perceived not only as unfair but is also problematic given constraints on time and energy. This generates a demand among women that their male partner should contribute (more) to the family work. While female employment is typically accepted by both women and men in the industrialized world (Scott et al 1996, Panayotova and Brayfield 1997), the division of unpaid domestic work is a much more controversial issue in most countries, however (Shelton and John 1996, Braun et al. 1998). Its limited societal acceptance is likely to counteract women's claim on men's engagement in household duties. The lack of support for the idea that men should share family responsibilities with women is reflected in public policies. For example, working parents can take parental leave in many countries but the eligibility of fathers for such rights is often considerably more restricted than for mothers. In fact, men were not eligible for parental leave before the 1980s or 1990s, except in Scandinavia.

The actual policies and the policy discourse in a country are likely to influence the development of gender relations at both the family and the societal level either by promoting or by slowing down changes in traditional gender roles. In the process of altering gender relations such policy influence on women's perceptions of the fair share of family work is probably an important item beside its role for female labor-force participation. In addition, the mechanisms around women's employment and family instability are likely to be somewhat different in societies where the male-breadwinner norm is the ideal family model than in countries where the dual-earner family model has a fairly long history (Lewis 1992, Sainsbury 1996). While we have some understanding of these mechanisms in the male-breadwinner setting based on the rich American literature, our knowledge about societies that pursue the dual-earner family model is relatively limited. As the gap between male and female employment rates are rapidly diminishing all over the developed world, we need to increase our efforts to learn more about the gender dimension of family dissolution in the dual-earner context. This study makes a modest contribution in that direction.

We have two main hypotheses that we want to test in this study, namely (1) that changes in gender relations around paid and unpaid work influence family stability, and (2) that there are gender differences in the effects of the partners' characteristics, especially regarding labor-market behavior, that influence family dissolution even in the dual-earner-family setting.

2.2 Our analytical strategy

We focus on families with children, because the disruption behavior of such families is of much greater societal concern than the dissolution of childless relationships. Also, the bargaining process around paid and unpaid work within the family is more strongly gendered for families with children than it is for childless couples. (See Presser 1994, Ahrne and Roman 1997, McFarlane et al. 2000.) The dissolution of families with children is a rather complex event since at least three parties are involved: the woman, the man, and their child(ren). To reduce the complexity of the issues involved, we focus in this study on the first parental union, defined as the union in which the first child was born to a couple where neither partners have children from a previous relationship. Arguably, entry into parenthood constitutes the most important role transition in adult individuals' lives.

The combination of employment and family responsibilities is a key aspect of changing gender relations, and is a particularly important issue in countries where the dual-earner family model has become relatively well established. As this is a longish process, we need to study a calendar period in which traditional gender relations have been exposed to new requirements inducing a course of changes both at the family level and at the societal level. For such reasons we focus on Sweden and Hungary, and the period from the mid-1960s to the early 1990s. We apply a comparative research design because of the complex nature of the issues we intend to study. It helps us to avoid generalization based on results specific for a single country alone, and it can help us detect patterns and mechanisms which are part of a more general trend.

3. The comparative setting

We have selected Sweden and Hungary for our study for a number of reasons beyond the fact that we know both societies well and have access to the best data available for them. In some respects these two countries are similar, at least largely though not always in the details; in other respects they are really different.

Here is the list of arguments we base this comparison on.

First, in the period we study, the gap between the proportions of women and men with higher education has greatly diminished **both** in Sweden and Hungary. In parallel, women's labor-force participation has reached high levels as compared to other industrialized countries. Yet, while **Hungarian** women have long worked full time just like men have (Szalai 1991), the proportion of part-time workers has been rather high among employed women in **Sweden** (40-50 per cent; see Sundström 1987).

Second, the high level of female employment is commonly connected with the concept of gender equality. It has a relatively long history as an influential principle of public policy-making **both** in Sweden and Hungary. However, what gender equality consists in has been interpreted somewhat differently in these countries. In **Hungary**, the understanding of the issues involved was limited to equal labor force participation of women and men (Makkai 1994). In **Sweden**, the policy discourse has aimed more broadly at a general transformation of traditional gender roles into a system with equal participation in paid work but also in family responsibilities for women and men (Sainsbury 1996).

Third, families with children have received substantial state-support in **both** countries, and a wide range of social services have facilitated the combination of employment and parenthood for women (Table A). Some of these policy measures have induced and/or reinforced changes in gender relations. The parental-leave program is a clear example of this. In **Sweden**, fathers have been able to take parental leave on equal terms with mothers (with 90 per cent income-replacement) ever since 1974. In **Hungary**, mothers have been eligible for a long child-care leave with a flat-rate benefit since the late 1960s. Fathers became eligible first in 1982 and only for children above age 1. An income-related parental benefit at a 75 per cent replacement level was introduced in the mid-1980s. The flexibility of the Swedish program (single days and half- or quarter days can be used) in combination with the high replacement-level of the benefit have encouraged fathers to engage in active parenting and to use some parental leave. In contrast, the Hungarian program has been formed so as to show that childcare is women's task; fathers are considered as secondary care-givers at best (Adamik 1991). Not surprisingly, only a few per cent of Hungarian fathers have personal experience of parental leave. In Sweden one-quarter of all parental leave users were men by the mid-

1980s, though they had barely a ten per cent share of the total number of benefit days (Sundström and Duvander 1999).

Fourth, on some other counts these countries have been very dissimilar. Beside their divergent histories, geographical positions, economic structures and political regimes (Baxter and Kane 1995, Panayotova and Brayfield 1997), their family formation patterns have also been quite different. In **Hungary**, childbearing was essentially restricted to marriage before the 1990s (Klinger 1991), while in **Sweden**, non-marital cohabitation has been very common at least since the 1970s, and births have increasingly occurred in such relationships. In fact, since the mid-1980s the majority of Swedish couples have been unmarried when they have become parents for the first time (Hoem 1996). As a number of studies have demonstrated that previous cohabitation and premarital children usually increase the risk of partnership disruption (for an overview see White 1990, Lillard et al. 1995), the difference in our countries' family formation patterns may have implications for their trends in family stability.

Fifth, **both** Sweden and Hungary have a long history of liberal divorce legislation (Table B) and of rather high divorce rates also among families with children (Goode 1993). As for the economic consequences of divorce the rules have been based on the ”clean-break”-theory regarding spousal alimony (none), and property division (even). While such rules often lead to a severe deterioration of women's economic situation after union dissolution (Voydanoff 1990, Holden and Smock 1991, Gähler 1998a) in other countries, this is not necessarily the case in Sweden and Hungary given the high level of female employment and extensive social support for families with children, including for solo mothers. In fact Hungarian fathers also experience economic hardship after family dissolution as mothers retain the right to live in the couple's dwelling, while men, beside a rather high child-support obligation, have to find another place to live, which has been a hard task considering the housing shortage (Utasi 1999).

Sixth, legal rules allow divorced or separated parents to continue to have joint custody for their children in **both** countries (Table B). In **Sweden**, joint custody means that parents have to cooperate in important questions regarding their children, such as to agree on their residence, education, sports and other activities, even though they usually live in separate households after family break-up (SOU 1995). The joint-custody rule has led to increased involvement in the children by both parents even after the relationship ends (Bernhardt 1996), notwithstanding that relatively few parents choose to share also the physical custody of the children when the family dissolves. In **Hungary**, the law requires parents' cooperation in important decisions regarding their children even if one parent has sole custody after the dissolution of the union. Yet, many fathers practically disappear from the children's life after the family break-up (Kamarás 1986) as mothers have sole custody for the children in the vast majority of cases (Hoóz 1995). Although fathers have visitation rights, they either do not use it or the children refuse to meet their fathers, sometimes under the influence of their mothers. So while in Sweden divorce or separation does not necessarily mean the loss of the parental status for either parent, Hungarian men often risk losing contact with their children permanently after a family break-up.

Based on their differences and similarities these countries provide us with an opportunity for a systematic comparative analysis of gender relations and family stability.

4. Data and method

The empirical analysis in this study is based on data extracted from the Swedish Survey of Family and Working Life, conducted by Statistics Sweden in 1992/93, and from the Hungarian Fertility and Family Survey, conducted by Statistics Hungary in the same years. Both surveys are part of the European Family and Fertility Surveys program, concerted by the Population Activities Unit of the United Nations Economic Commission for Europe. It aimed at providing data suitable for cross-country comparisons in demographic analyses. Both women and men, independently selected, were respondents in the Swedish as well as the Hungarian surveys. This allows us to study the influence of characteristics of both female partners and male partners on family dissolution. At the same time we recognize the limitations of the data, given that the true unit of analysis for family dissolution is couples, not individuals. We have little information on the partners of the respondents, as is the case for most event history data. Nevertheless, parallel histories from women and men provide a simulation of sorts of what we might expect had we information on both male and female partners' characteristics for each union reported by a woman or by a man.

The empirical material provides us with data on 4984 individuals for Sweden and 5487 individuals for Hungary. We know their childhood family characteristics, their full retrospective histories of union formation (cohabitation, marriage), childbearing, and family dissolution, as well as of educational and occupational activities, covering the period from the 1960s to the early 1990s. Respondents in the Swedish survey were selected by simple random sampling from each of five national strata of women born in 1949, 1954, 1959, 1964, and 1969, and three strata of men born in 1949, 1959, and 1964. The National Population Register was used as a sampling frame. Interviews were successfully obtained with 78 per cent of the women and 77 per cent of the men of the original target sample in Sweden. (For further details see (Granström 1997).) In Hungary the figures were 87 and 77 per cent, respectively. The Hungarian sample was nationally representative for the female population at ages 18-41 years and the male population at ages 20-44 years. The sample was selected partly through one-stage proportional sampling (for the capital and the largest cities) and partly through two-stage stratified sampling (for the rest of the country). In the first stage for the latter, sampling areas were stratified according to the number of inhabitants and municipalities were randomly selected from each stratum. The final sample was drawn at random, given sex and age constraints, using the National Population Register as the sampling frame for each stratum (Kamarás 1999).

For the purpose of the present study, we have selected respondents who have reported one or more marital or non-marital unions and have born at least one child in a union. Individuals who do not have a recorded union are excluded, as are those who were childless at interview or whose first child was born outside of a recorded partnership. From the sample for Sweden we have also excluded respondents who grew up outside of Scandinavia in order to avoid problems of cultural differences which could influence family dissolution risks. From the Hungarian sample we have excluded respondents who were below age 20 at the interview; only few of them had children anyway. Also, respondents with incomplete records on partnership or childbearing history are excluded. As the effects of changing gender relations on parental union disruption may be difficult to detect, we try to make the samples as homogeneous as possible. Therefore, we have also excluded respondents who have an adopted child in their first parental union or whose partner had a child from a previous relationship, as well as those whose first child in their first parental union died (Note 1) or whose union ended in the same month as they had their first child. Censoring occurs at sixteen years after first birth, when a union ends because the respondent’s partner dies, or at interview, whichever event comes first. A total of 2730 individuals (1869 women and 861 men) are included in the Swedish working sample and 3500 individuals (2430 women and 1070 men) in the Hungarian one. The proportion of respondents who experienced the disruption of their union is about 20 per cent for Sweden, and 13 per cent for Hungary. (For further details see Table 1 in Appendix A.)

The analysis is based on a piecewise-constant proportional-hazards model. Exposure is measured in months, starting from the birth of the first child of the respondents included and continued until the child turns 16 years old or until censoring for other reasons. We have divided this period into the following intervals: infancy (age below 1), toddler years (up to age 3), pre-school years (up to age 6), early and middle school-age years (up to age 12) (Note 2), and teenage years (up to age 16). We behave as if the disruption intensity is constant over each of these pre-selected time intervals, but let it vary between intervals. Information for those who did not experience the disruption of their first parental union during the period of observation is also taken into account.

Our computations are based on exposures in half-month units. We pretend that the interview and all recorded demographic and other events happened at the middle of a calendar month, while changes in period variables occur at the beginning of a calendar month. The Windows-based software ”RocaNova”, developed at Statistics Sweden, is used to fit the model. The results, produced as maximum-likelihood estimates of the effect parameters of the model, are presented in the form of relative risks. We analyze women and men separately for each country in order to detect possible gender differences in the effects of partners’ characteristics on family disruption.

5. Variables

5.1 Gender relations

As we have discussed earlier (Section 2.1) changes in gender relations which are potentially influential on family stability have occurred both at the societal level and at the family level. We use therefore separate measures in our attempt to estimate their effects on union dissolution risks.

We start with our measure of changes in policies that affect gender relations at the societal level but also have implications at the family level. We use a period approach to define this variable, since policy changes that denote stages of development towards gender equality in society take place in clearly identifiable calendar years. Alternatively, we could have used two dichotomous factors for measuring the effects of changes in legislation around family dissolution (one for divorce laws and the other for custody rules) in combination with a variable which picks up changes in the parental leave program. However, even this solution would be period-based in some sense. Also, given the several policy measures involved it would not have simplified our task. Thus we have chosen to use a single factor which we call *current policy period*. It represents a partitioning of calendar time based on periods of major policy changes linked to gender equality and to family dissolution.

For Sweden, we distinguish between three periods. The first period goes from the mid-1960s to 1973. This was a period of intensive public debate on gender equality, a time of consensus building concerning the meaning of the concept, with outcomes that strongly influenced subsequent policies and later changes in traditional gender relations (Baude 1992). In the second period (from 1974 to the middle of 1983), a range of policy reforms reflected an increasing influence of the idea of gender equality on policy-making. Men’s role as parents was recognized in the same way as women’s through the parental-leave program, but the leave provided was relatively short (6-9 months). The divorce-law reform aimed at a clean break between the spouses by allowing a quick and easy divorce procedure but no spousal alimony after divorce. In these years the emphasis was mostly on strengthening women’s economic provider role as one aspect of gender equality. Between mid-1983 and 1993 (our third period), the attention turned towards men. Policies aimed at increasing men's involvement in parenting as the second main component of gender equality. For example, a reform in mid-1983 introduced joint custody as a general rule for children after the parents' separation. In 1989, the parental leave was extended to 12 months. The longer leave provides better opportunities for men to strengthen their relations with their children by taking out parental leave after the end of the breastfeeding period. Thus the Swedish policy environment has not only been ”women-friendly” (Hernes 1987) but also has facilitated the transformation of the traditional male gender-role even though the changes in men's behavior have been rather limited (Ruggie 1988, Lewis 1992).

For Hungary, our policy variable is based also on three periods but they mark much less radical policy changes than in Sweden, at least for gender relations. In the first period (from the mid-1960s to 1981), a long child-care leave (2 years) with a job-guarantee was introduced for mothers in Hungary (in 1967) as a recognition of women's double role as mothers and workers in policy-making. Women's economic provider role was also emphasized in the two relatively liberal divorce laws from that time (the first from 1952 and the second from 1974 (Note 3) that greatly limited the possibility of spousal alimony after divorce since both women and men were expected to be gainfully employed. In the second period (from 1982 to mid-1987) fathers' right to use parental leave on a par with the mothers was introduced (in 1982), but only for children above age one. This rule was far from promoting Swedish-type equal parenting as it was not backed up with a policy discourse which would have encouraged men's active participation in parenting. The third period (from mid-1987 to 1993) started with major changes in family law. Divorce procedures became more complicated and took longer because a compulsory pre-divorce court hearing was introduced. Unmarried fathers could get joint custody if they lived with their children's mother. The latter is, again, a rule which strengthened men's role as fathers. However, the lack of supportive policy discourse suggests that policy makers had no sincere intention to change traditional gender relations in Hungary (or in other Central and East European countries), notwithstanding that women's labor-force participation on conditions similar to men's was seen as the precondition of gender equality during the forty years of state socialism (Watson 1993).

Our next central covariate is a measure of *gender relations in the union.* We constructed this factor somewhat differently in the two countries because of differences in the data. The Swedish survey provides information on whether the father took any parental leave with the first child (Note 4). This gives us a clue both to the relations between the partners and to the father's involvement in parenting. In unions based on traditional gender relations, men do not take parental leave regardless of the societal expectations around the modern paternal role. In other relationships where the idea of gender equality is accepted by the partners, men tend to take parental leave both because they wish to spend more time with their children and also to take their share of parenting responsibilities and thus help the mothers. We do not distinguish according to the length of the leave taken by the father, as previous studies have shown that mothers are quite unwilling to let their partner use more than a small part of the parental leave (Haas 1992, Sundström and Duvander 1999). Thus father's use of parental leave functions as a signal of whether he is prone to share family responsibilities with the mother or not, which can influence union stability.

For Hungary, we apply a gender-role-attitude variable. It is an index based on several questions regarding relationship and career. (For details, see Table H.5 in Appendix A.) Also, we use the information on the respondent's attitude to parenthood provided in the data. These indirect measures should be more useful than information on paternal leave could have been, since Hungarian men hardly ever take such leave even when they have the opportunity (S. Molnár 1992). Such attitude variables may, however, be somewhat problematic as respondents might have adjusted their attitudes to their previous behavior (in a kind of ”post-hoc rationalization”). Since there is no additional information on gender relations in the data for Hungary, we apply these attitude measures, but we shall be cautious in our interpretation of the findings.

As women's and men's equal responsibility for both paid work and family work has been part of the everyday discourse (of both policy discussions and public debate) in Sweden since the 1960s (Sainsbury 1996), whereas there was no real effort to challenge traditional gender relations in the family in Hungary (Szalai 1991), we expect to find a stronger effect of changing gender relations on family stability in Sweden than in Hungary (Note 5). This should be the case not only for our societal-level variable but also for our family-level factor as traditional gender relations have changed very little within the family in Hungary while Swedish women expect their partners to take more responsibility for domestic duties. (For evidence of the latter see (SCB 1982, Roman 1999).)

5.2 Labor-force attachment

In order to test our second hypothesis on gender differences in the effects of the partners' characteristics regarding labor-force attachment on family disruption, we use two factors, one to estimate the effect of schooling and the other for labor-market strategies.

We apply *current educational attainment* to measure differences in individuals' human capital which are likely to influence parenting and employment strategies but also other skills that may be important in a partnership. Our variable refers to the level of schooling the respondent had attained up to any month after the first birth. This is a multifaceted factor in union-disruption risks. Since improved education is likely to lead to a higher income, we expect men with higher education to be considered as more attractive partners and therefore to have stabler families. Also, highly educated men turn out to share childrearing responsibilities (Näsman 1992, McFarlane et al. 2000) and household tasks (Shelton and John 1996) with their partners more often than less educated men. This should make them even more desirable as partners in a union.

For women, the reverse relationship may obtain. On the one hand, the higher salary of highly educated women increases their attraction value on the partnership market too (Oppenheimer 1994). On the other hand, it also provides them with greater economic independence from their partners. Thus they may be more likely to leave a less satisfactory relationship than other women (see Hobson 1990), especially as they probably also are more certain about their abilities to live as single parents than less educated mothers are. Furthermore, highly educated women more often pursue gender equality in the division of domestic tasks (Presser 1994, Ahrne and Roman 1997) and probably are less likely to accept it if their partners fail to share the responsibilities of home life. On the other hand, one could also argue that given their higher salaries highly educated women can purchase domestic services more easily in the market (in the form of hiring domestic help) and thus need much less help from their partners than less educated women. In Sweden and Hungary such a solution was, however, rather atypical even among women with higher income in the period we study. Another aspect of higher education is that it can improve individual skills to find a good partner and to solve conflicts. This in turn would strengthen union stability.

Thus, while the dissolution risk is likely to decrease when a man's educational level improves, we are uncertain about whether disruption risks increase or decrease with improving educational attainment for women when we take all these various items together. Only an empirical analysis can provide us with answers as to how these issues balance against each other in Sweden and Hungary.

To represent labor-market strategies, we use a measure called *current employment status*, which shows the respondent's labor-market attachment in any months after the first birth. Alternatively, we could have measured previous work experience. However, such a measure would be less informative than current employment status since women, even mothers, of working ages normally are in gainful employment in Sweden and Hungary. In such a context our measure seems to be a sounder choice as it also informs us about how much time our female respondents may have for their partner, for children, and for domestic tasks, given the constraints of working hours. Unfortunately, being on parental leave was not recorded as a separate category for Hungary but as a continuation of the respondents' previous employment status.

As mothers are still the main carer for children, parenting probably has a stronger influence on their work strategies than that of fathers. Furthermore, doing part-time rather than full-time work is likely to have different implications for women than for men because of their differing family roles and particularly their differing labor market roles. For men, the rule is full-time employment in industrialized countries. By contrast, because women are expected to give priority to family responsibilities, they are allowed to follow a wider range of work strategies, often strongly related to the age of their children. Giving priority to labor market roles over family obligations is considered ”normal” for men, but much less so for women. This is very much the case in Sweden (see e.g. Lundén Jacoby and Näsman 1989, Ahrne and Roman 1997). Thus we expect to find differences between women’s and men’s risk patterns for employment status for family dissolution there. In Hungary, by contrast, opportunities for part-time work were extremely limited during our period of observation. (Only about 3 per cent of all jobs were part-time jobs, Frey 1993.) Consequently, full-time employment has been the rule for both women and men. Thus gender differences are less likely to show up in the pattern of dissolution risks by employment status for Hungary than for Sweden.

5.3 Analytic model

The analytic model we use to test our two main hypotheses includes then four explanatory variables, i.e. current policy period, gender relations in the union, current educational attainment, and current employment status. These are time-varying covariates with the exception of the ‘gender relations in the union’ measure, which is a fixed factor. The model also includes factors that are not of main interest to us but that have been found to greatly influence family disruption in previous studies, or that may otherwise reasonably be assumed to affect union dissolution (see Appendix A). We divide these control variables into the following groups:

1. Individual characteristics: Some of these refers to the respondents’ childhood family experiences (composition of family of origin, number of siblings), others relate to features such as the respondents’ own birth cohort and own religiosity.

2. Maturity at family formation: This group of factors is linked to the respondents’ maturity as an individual (age at first birth grouped according to educational level at first birth) and as a member of a couple at the time when the relationship became a parental union. (The latter is measured by taking into account the interval in months between (i) the start of the first parental union and (ii) the first birth. We call this factor ”first-birth interval”.)

3. Characteristics specific to the parental-union: These factors contain important further information on the relationship, namely a representation of the order of the union

in which the first birth occurs (called ”first-birth union order”), the marital status, the number of children in the household, and the current age of the youngest child.

4. Business-cycle variations: These are measured with country-specific factors (current unemployment rate for Sweden, changes in consumer price index for Hungary).

As our time variable we use the *age of the first child* (i.e. union duration since first birth). Our choice is based on the assumption that the birth of the first child radically changes the nature of a relationship. Individuals who are parents make calculations in their decisions on whether or not to dissolve their unions that are different from those of childless couples. Parenting tasks also change as the child grows older, as do parents' perceptions on how harmful family disruption is to children. These in turn influence parents' propensity to dissolve their families. We have taken these changes into account in the definition of categories of the time variable. We measure the age of the first child in months, but our age-categories are defined in years (see Section 4 for details) as this is easier to understand.

We use a stepwise approach for the model fitting. In a first step we include only individual characteristics. Next we add the group of covariates that represent maturity at family formation. Then we include the rest of the control variables. This stepwise introduction of factors into the model corresponds to the sequence in which they appear in the respondents' life. This in turn determines their causal proximity to the current life situation of the respondents (see Figure 1). This procedure will also help us to see whether the inclusion of a factor at a later stage in this process influences the estimated effect of an earlier factor. When it does, the indirect effect of the early factor which works via the later factor disappears when we include the latter, and we can see more clearly the direct effect of the earlier factor. This allows us to exclude control variables that do not have a significant direct effect on first parental union dissolution. In our last step we keep the control variables that have proved to be important for the current analysis, and include our explanatory variables in the model.

In the present part of our account we discuss only the main results of our final model. (See Appendix B for details on the stepwise model fitting in our multivariate analysis.)

6. Findings and discussion

6.1 Sweden

6.1.1 Gender relations

As we see for our societal-level variable (Table S; for the full model see Table S.II in Appendix B), public policies, measured as *current policy period* in our analysis, appear to influence family disruption. The effect is stronger in the male sample, in spite of its much smaller size as compared to the female sample. In both samples, however, the patterns of dissolution risks across policy periods are much alike.

The risk of disrupting a first parental union hardly changed over our first two periods (1965-1973, and 1974-middle of 1983), notwithstanding the large increase in female labor force participation which took place at that time in Sweden (Sundström 1987) and notwithstanding the introduction of one of the most liberal divorce laws of the world (in effect since 1974). Couples with children were apparently much less affected by these changes than childless couples (Note 6) were, at least when it comes to long-term effects. In the third period (July 1983-1993) however, the risk of family dissolution rose significantly among families with children, and we notice it particularly in the male sample. In this period, policy emphasis lay on increasing men's involvement in parenting and joint custody for children after family break-up was introduced as a main rule. Perhaps as this rule gave both parents a better chance to remain active parents even if the children do not live in their households permanently after family dissolution, parents felt less obliged to stay in a union which they found unsatisfactory. In this sense, the joint custody reform may have reduced family stability. Yet, the fear of (oneself or the other parent) losing contact with the children is not a good basis for keeping a relationship intact as it may create an atmosphere in the family which is more harmful than a break-up for the children and also for the parents. Moreover, previous research has shown that a ”good divorce” may actually be better for the children than living in an intact family with a lot of tensions between the parents (see e.g. Amato et al. 1995, Gähler 1998b, Jekielek 1998, Morrison and Coiro 1999). An alternative explanation for the increase of disruption risks in the third policy period as compared to the previous periods points to women's rising expectations of their partner sharing family responsibilities (see Nordenstam 1985, Roman 1999) and the gap between such expectations and men's actual involvement in domestic tasks (Nermo 1994).

Looking at *gender relations in the union* (measured through the fathers' involvement in childrearing) we notice that they greatly affect family stability in Sweden, and significantly so in the female sample. The patterns of disruption risks are very similar in both samples. If the father took some parental leave with the first child, the risk of union dissolution is lower than otherwise. This suggests that both partners' engagement in economic and care responsibilities can strengthen their relationship, possibly because the sources of potential conflicts are reduced in such unions. Alternatively, this risk pattern may show a selection effect as men who take parental leave are likely to be more family-oriented than other fathers and thus may also have stabler unions.

We have also tested for possible interaction effects between current policy period and gender relations in the union. The results are not significant in the male sample (and are not presented here), again probably because of its smaller sample size and correspondingly greater random variation. In the female sample, the results are significant at the 6 per cent level and the pattern is quite interesting (Figure 2S). Before the early 1980s, the risk of family disruption in unions where the father took parental leave with the first child, was about half that of other unions. During the 1980s and early 1990s, however, the risk of family dissolution increased for all unions, and more so in partnerships where the father used parental leave. Their disruption risk was still well below that of unions where men did not engage in such active parenting. How can we interpret this pattern?

The very low dissolution risk for active fathers before the early 1980s may reveal a selection effect. Men who took parental leave at that time may have been more highly family-oriented than others and more so then than later on when it became more common for fathers to take some parental leave. However, this has not been accompanied by a comprehensive change of attitude among men in terms of them sharing domestic tasks more equally with their female partner (see Nermo 1994), notwithstanding that men in other countries are even less engaged in family work than Swedish men are (Flood and Klevmarken 1990). Swedish women's expectations on the more equal distribution of unpaid domestic work have increased in the 1980s and 1990s (Nordenstam 1985, Roman 1999) given their experiences of gainful employment and the Swedish policy discourse on gender equality. Men's failure to meet these expectations then resulted in reduced family stability for all unions. Men's engagement in active parenting still had some stabilizing effect. As we do not have information in the data sets on the division of household work (Note 7), our explanation cannot be tested fully but remains somewhat speculative. In any case, these findings (i.e. both the results of the main effect model and the interaction term) suggest that changing gender relations have influenced family stability in Sweden.

6.1.2 Combining parenthood and employment

As dual-earner families became prevalent in Sweden, the combination of parenting and paid work became an important aspect of parents' family stability. Our analysis shows that factors related to mothers' labor-market attachment influence family disruption risks significantly but this is not the case for fathers (Table S). This is in line with our expectations, given that women still have a larger share of parenting responsibilities than men do.

Parents' *current educational attainment* proved to be rather important for family stability. Although mother's education has a stronger effect than does father's education, the patterns of dissolution risks are very similar. Those who have the lowest educational level have the highest risk of family disruption. There are no significant differences in the disruption risks between the various levels of those with more than compulsory education. This suggests a selection effect for those with the least education, but can also refer to their weaker labor market situation and lower incomes. These in turn can create serious conflicts between the partners in the relationship, who may also have less skills to solve them. In any case, we do not find important differences in effects of mother's versus father's education on the risk of family disruption in Sweden.

Current employment status also has a strong impact on family disruption risk. Compared to mothers employed full time, those who are students have a higher risk of disruption, while those who are employed part-time or on parental leave or not employed have the lowest risks. Men's employment has nearly opposite effects, with higher risks of dissolution for fathers who are employed part time or unemployed (but not for those on parental leave which is included in the ”other non-employed” category). Male students do not have a higher disruption risk than do full-time employed fathers. These results taken together indicate that couples with a more traditional division of labor are less likely than couples in which the man's and woman's employment is similar or the couple has reversed traditional roles to end their first parental union. This is linked to societal expectations for men to be ”good providers” and for women to take the lion's share in childrearing and household tasks in Sweden too (Björnberg 1992). Hence, the male gender role seems to change very slowly, while great changes in the female gender role in terms of labor-force participation are accepted as long as gender relations in the family concerning the division of responsibilities for domestic tasks are hardly affected.

We have also tested for interaction effects between our two labor-force-attachment factors. The effect was not significant in the male sample, and even in the female sample it was significant only at 12 per cent, but the pattern is interesting nevertheless. As we see in Figure 3S, the least stable families are those of mothers with compulsory education who work or study full time, and of students with higher educational attainment. For all other work categories (except ”other non-employed” women), the least educated mothers have the higher risk of union disruption. Hence, being a student has a negative impact on family stability independently of educational level (Note 8), but full-time work does not automatically lead to reduced union stability for mothers who have more than compulsory education. This pattern is in line with findings of Swedish time-budget studies showing that the partners of more highly educated women are more likely to share domestic duties than is the case for women with compulsory education (Nermo 1994). Based on such findings and on our own results we may speculate that in families where parents share both economic and domestic responsibilities, changing gender relations would not have as detrimental effect on union stability as in families where only the woman follows the new gender behavioral pattern while her partner fails to change.

In addition, we tested for interactions with our variables measuring gender relations at the societal level and at the family level respectively. None of these interaction terms were found to be significant in the male sample. In the female sample, we found that women's education interacted with policy period and paternal leave-taking in its effects on the disruption risk (significant at the 5 per cent level).

Figure 4S shows that mothers with the least education had the highest risk of family disruption in all three policy periods. However, the variation of dissolution risks over time is rather different for them than the pattern we see for mothers with more than compulsory education. Family instability decreased somewhat for less educated women between our first and second period, but it increased strongly thereafter. In contrast to this pattern, the risk of union disruption nearly doubled for mothers with more than compulsory education from the first to the second period, but it hardly changed in the decade from 1983 to 1993.

Given that family instability did not increase for less educated mothers in the second policy period as compared to the earlier period, but it did for mothers with some education, the explanation of the increased disruption risk for the latter group is not so likely to be found in the introduction of no-fault divorce law, for there is no reason why the outcomes would be so different for these two educational groups. Instead we offer an interpretation based on the notion that gender ideas spread like an innovation process. We speculate that the increased risk for the more educated mothers may have depended on their changed perceptions of a fair division of domestic tasks. They may have wanted their partners to take more responsibility in family work than the partners would have according to traditional gender relations. They may also have been less tolerant towards the partners' failure to meet their higher expectations than women with the least education were initially. We suppose that in the third period similar ideas had spread to less educated mothers too. The joint-custody rule established in the third period may have contributed additionally to the increase of disruption risk specifically for women with the least education. For more educated women, their partners may have been more likely to remain part of their children's life after family break-up (as has been shown for the US context (Stephens 1996) even before the introduction of the joint-custody rule. For less educated mothers this may not have been the case before the mid-1980s, thus the parents in those families may have been more likely to stay together ”for the sake of the children”. Such sacrifice became unnecessary for them when parents retained joint custody for their children after family dissolution as the general rule.

The interesting pattern seen for Figure 5S reveals a further aspect of the link between family-level gender relations and union stability. We notice that the risk of union disruption is lower for women at all educational levels when the father took parental leave with the first child than otherwise. Women with the lowest and the highest education whose partner did not take parental leave have the least stable families among all mothers. The risk of union disruption is twice as high for highly educated mothers whose partner proved to be a less engaged parent as for those whose partner took parental leave. The difference in disruption risks between those with a more caring partner and those with a less caring partner is much lower for women with less than higher education. This indicates that women with post-gymnasium education are less likely than other mothers to accept the failure of their partner to adjust to modern gender roles and not to engage in active parenting.

In sum, our findings for Sweden point to the importance of gender relations in family stability as well as to changes in gender perceptions over time. In addition, the disruption-risk patterns by employment status revealed substantial gender differences in the combination of parenting and employment and in its consequences on union stability.

6.2 Hungary

6.2.1 Gender relations

In contrast to the Swedish findings, the overall impact of public policies, measured through *current policy period*, on family disruption risks in Hungary is not significant (Table H; for the full model see Table H.II in Appendix B). This was what we expected, since the Hungarian policy discourse around gender equality and the changes in relevant policies were less aimed at changing gender relations in the family than it was the case in Sweden. The disruption-risk patterns are alike in the female and male samples, with much higher risk in the early and mid-1980s than in the previous period. The difference is not significant in the male sample, again probably because of its much smaller sample size. In the third period, the risk of family break-up decreased somewhat but the change was not significant in either sample.

The higher disruption risk for the second period than in the first period suggests that men might have felt having a better bargaining position to keep contact with their children even after a divorce due to the new rules to strengthen their role as fathers. This may have made them more likely to accept an end of their union if they found it unsatisfactory. Alternatively, the higher disruption risk was connected to increasing individualism in the Hungarian population as Hungary became more open to ”Western influences” than other Central and East European countries, especially in the 1980s. As for the somewhat lower although not significantly different divorce risk in the third period, it seems that the more restrictive divorce law from 1987 could not strengthen family stability in the long run, possibly because of the dramatic changes of the political and economic system after 1989. The collapse of state socialism and the rapidly increasing unemployment level generated tensions which a large number of individuals and families had difficulties to cope with. This could counteract the more restrictive divorce law.

Not surprisingly, the impact of gender relations at the family level on union stability in Hungary is not significant either, quite unlike our findings for Sweden. For both *gender-role attitude* and *attitude to parenthood*, the disruption risk patterns are rather similar in both samples, and the risk profiles are in line with our expectations. For the former factor we find that those with a traditional gender-role attitude have lower risk of union dissolution than parents with egalitarian or intermediate gender-role attitudes. For the latter factor we see that the risk of family disruption is higher for self-centered parents than for others.

The differences between the dissolution risks for the different categories are not significant for either of these factors. For the attitude-to-parenthood variable this might be explained by union dissolution being a widely accepted behavior in the Hungarian society, and parents whose main principle is their children’s best interest (i.e. our ”child-centered” category) are no exception. Perhaps some of those among them who dissolved their union might even argue that they have done so because of their children. Thus a given attitude to parenthood is less likely to function as an impediment of divorce or separation. As for the other factor, the lack of significant differences in dissolution risks among parents with different gender-role attitudes can be related to the fact that most Hungarian families follow the traditional gender division of labor in the home in spite of women’s full-time employment. Thus the ”comparison referents” (see Section 2.1) could hardly have changed women’s perceptions on the fair share of domestic tasks in Hungary. Also, the Swedish-type intensive public (and policy) debate on the new male gender role as linked to the idea of gender equality was missing in Hungary as we have noted before. Hence, the differences between our three types of gender-role attitudes remained relatively limited.

The lack of significant influence of these factors (i.e. attitude to parenthood and gender-role attitude) on family disruption in Hungary can also be connected to the fact that our gender-relation variables are attitude measures, while for Sweden we had a measure based on actual behavior (father's participation in active parenting). As attitude and behavior do not always coincide, these measures may have been less adequate for the dimension we try to tap, though they were the best we could obtain from the Hungarian data. However, this measurement problem is unlikely to be of major importance as the effect coefficients are indeed meaningful in the Hungarian context.

We tested also for interaction effects between our gender-relations variables (i.e. between gender-role attitude and attitude to parenthood, and these separately with current policy period). None of these interaction terms were significant or interesting in either sample.

6.2.2 Combining parenthood and employment

While the gender-relations factors had no significant impact on parents' family stability in Hungary, labor-market behavior appears to have stronger effect (Table H). *Current educational attainment* seems to be somewhat less important for union stability in Hungary than it is for Swedish parents, but effects of women's and men's education are similar. As is true in Sweden, those with low level of education have higher risk of family disruption than individuals with more schooling. However, the threshold for effects of men's and women's education differs in Hungary. Hungarian fathers with more than compulsory education have much lower risk of family disruption than the least educated men, while mothers have a reduced dissolution risk only at a gymnasium and higher level of schooling. The gender-specific thresholds may be connected to some gender-segregation in the labor market, and the disruption risk patterns relate to the relatively poorer economic situation of those with low education. It is thus a gender effect that for women more schooling was required for a job with decent earnings than for men for whom any education above the compulsory level could lead to a relatively good income. As individuals with the least education had fewer resources, their housing conditions were relatively poorer. This meant that they more often had to share a home with the parents of their spouse or with their own parents, which is likely to create tensions. Alternatively, the higher disruption risk for those with less schooling can be explained by their lower skills in solving conflicts that arise in the relationship.

The relatively strong influence of *current employment status* on family dissolution in Hungary is rather surprising, as the vast majority of women and men alike were employed full time during state socialism. However, we only find significantly different disruption risks when women are outside of the labor market (except for housewives) and when men work less than full time. The lack of significant differences in disruption risks between housewives and employed women who work full time or part time can be related to the limited influence of women's work strategies on gender relations within the family, especially regarding the persistent traditional division of domestic tasks in Hungary, unlike in Sweden. In addition, Hungarian women's student status does not increase the risk of disruption, but instead reduces it, contrary to the Swedish case. On the other hand, Hungarian men who study have a high risk of union disruption but the effect is not significant. In any case, students represent a very small fraction of parents in Hungary as childbearing was usually postponed until one finished education and had a job. Similarly to the findings for Sweden, father's part-time employment (and unemployment) greatly increases the risk of family break-up. As gender arrangements in the family remained traditional in Hungary, men's role as the main economic provider of the family is even more emphasized there than in Sweden.

We also tested for interaction effects between the labor-market attachment factors, and for these with each of the factors of gender relations. None of these interaction terms proved to be significant or interesting in either the female or the male sample.

To sum up our findings for Hungary, traditional gender relations within the family have hardly changed there, unlike in Sweden, notwithstanding the societal acceptance of women's full-time employment (also for mothers). Nevertheless, we found gender-specific disruption risk patterns for parents' labor-market behavior in Hungary similarly to the results for Sweden.

7. Concluding remarks

As the male-breadwinner model has increasingly been challenged (or even replaced) by a dual-earner family model in industrialized countries in the last decades of the 20th century, male-female relationships within the family operate in a new terrain. While women's economic dependence on their spouses was undoubtedly an important reason of family stability in previous times, the new conditions under which families function nowadays have changed the nature of the relationship between women and men (Oppenheimer 1994). Yet increasingly overlapping gender roles in the labor market do not automatically result in similar changes in the home. The gap between gender roles in the public and private spheres can then become a source of family instability, which can be overcome by a more equal share of domestic responsibilities between women and men. In this study, we have found evidence for such a development in Sweden. While the policy discourse and the actual policies around equal participation in work and family life have encouraged this change, there is still a gap between ideology and practice.

Hungary remained a ”traditional society” in many ways and preserved a traditional gender hierarchy during state socialism in spite of the dominant position of a (full-time) dual-earner family model for over 40 years. In this country gender relations seem to be of much less importance for family stability than other factors. Perhaps economic constrains have been more severe for shaping family decisions in Hungary than in Sweden. In addition, there may be a threshold in individual awareness of the need to change traditional gender-behavior patterns also at the family level, and this threshold may not yet have been reached in Hungary. Such awareness has started to grow in the countries of Central and Eastern Europe in the transition period which followed the collapse of state socialism, while the conditions under which women combine paid work and family life have become harder (Watson 1993). A new wave of family surveys in the early 21st century may then provide us with data that help us further clarify the relationship between changing gender relations and family stability.

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Notes

1. Both the death of a child and having the partner's child from a previous relationship in the household can create tensions in the family which may lead to union dissolution. Such cases are not suitable for a study of the effect of changing gender relations on family stability.

2. In a preliminary analysis we distinguished also between the categories 'years of school start' (up to age 8) and 'mid-school-age years' (up to age 12). We decided to combine them into a single category (i.e. 'early and middle school-age years') based on their very similar disruption intensities.

3. We have seen in a previous analysis that the further liberalization of the Hungarian divorce legislation (in effect from 1974) had no significant long-term effect on family stability among families with children (Oláh 2000).

4. Swedish fathers' eligibility for parental leave was limited before 1974, and under certain circumstances even later on. Thus we also include a category called 'other' for this variable.

5. One might think that Hungary being a more traditional society the effect of non-traditional gender attitudes should be stronger than in Sweden based on the greater selectivity of that group. However, the lack of change in the home sphere in Hungary in terms of labor division suggests that non-traditional attitudes were not accompanied by less traditional everyday practices. Thus the effect of attitudes were probably relatively limited. Therefore we expect to find stronger impact of changing gender relations in Sweden than in Hungary.

6. Andersson's (1997) annual index of divorce-risk level for Sweden rose dramatically in the 1970s and 1980s. His study included only married couples, while we also include non-marital unions as we would have expected a bandwagon or contagion effect for cohabiting couples in times of radically increasing divorce indices. Yet as we found also in a previous analysis (Oláh forthcoming), the risks of family dissolution among families with children changed very little in the period before 1983 both for non-marital and for marital unions (and independently of whether a marriage started as a cohabiting relationship or not).

7. The Swedish questionnaire did not address these issues. For Hungary, there is some information about the share of domestic work, but only for respondents who were living in a relationship at the time of the interview, and this is insufficient for our purposes. We do not have data on the gender division of labor in previous unions for Hungary either.

8. Sweden has a highly developed system of adult-education comprising all levels of education from primary to tertiary levels.

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Table A: *Public support provided to families with children in Sweden and Hungary*

	Sweden	**Hungary**
Cash benefits:		
1, Child/family allowance	introduced in 1947	introduced in 1938
	universal benefit	until the 1970s employment in state sector required; 1990-1995: universal benefit; means-tested from 1996
	paid to the mother	paid to the father (or to the mother if ineligible father)
	one-child families eligible from the start	originally for large families; from 1983 also one-child families eligible
	higher allowance for the third child (and additional children)	higher allowance for single parents, for those with disabled children, and for large families
2, Housing-related benefit	introduced in the mid-1930s, for large families	introduced in the 1970s
	means-tested housing allowance	"baby bonus" - part of the bank loan for young couples building or buying a dwelling transformed into a benefit if they had one child within three years or two children within six years
	one-child families eligible since 1958	only couples below age 35 eligible
3, Maternity/ parental benefit	1962: six months maternity leave with 65% income replacement	1967: child-care leave until the child becomes 2.5 years old; extended to the third birthday of the child in 1969. Flat-rate child-care allowance "GYES" (about 40% of the average earnings of a young female unskilled worker)
	gradually extended parental leave since 1974 with income-related benefit (90% replacement level until the mid-1990s)	only mothers eligible at first; since 1982 also fathers after the first birthday of the child
	1974: 6 months parental leave	1985: income-related child-care pay "GYED" (75% replacement level); only mothers eligible with children below age one
	1975: 7 months parental leave	1986: "GYED" extended by half year; fathers eligible after the child's first birthday
	1978: 8 months parental leave, plus one month with flat-rate benefit	1987: "GYED" extended until the second birthday of the child
	1980: 9 months parental leave, plus three months with flat-rate benefit; also 10 "daddy days" with 90% income replacement	"GYES" paid by the number of children below age three cared for at home; "GYED" is independent of the number of children
	1989: 12 months parental leave, plus three months with flat-rate benefit	1996: "GYED" abolished, "GYES" means-tested
	1995: one month leave reserved for the father and another for the mother with 90% income replacement, but only 80% for the other 10 months leave	
	1996: 75% replacement level	

(continued on the following page)

Table A (continuation):

	Sweden	**Hungary**
3, Maternity/ parental benefit	1998: 80% replacement level	
	leave on a full-time or part-time basis, or spread out over the years up to the child's eighth birthday	leave on full-time basis; since 1982 "GYES" on half-time basis in combination with part-time work
	taxable benefit with job-guarantee and pension-entitlement	both benefits with job-guarantee and pension entitlement; only "GYED" taxable from 1988 when individual income tax was introduced
Social services:		
1, public childcare	since the 1960s, rapid expansion after the mid-1970s	since the early state-socialist period
	run by municipalities	run by municipalities and by big companies
	financed mainly by government subsidies, but also by parents' fees based on the family income	financed mainly by state subsidies, but also by parents' fees based on the family income
	employed parents and students with pre-school children (age 1.5-7years) eligible	employed parents and students with children of age 0.6-6 years eligible
2, school lunch	free or low fee for school-lunch for pupils in primary school	school lunch for pupils in primary school and secondary education; fees based on the family income

Table B: *Divorce legislation and child custody rules in Sweden and Hungary*

	Sweden	**Hungary**
Divorce legislation:	1915: divorce on fault grounds; also some no-fault grounds acknowledged	1952: no-fault divorce introduced; a marriage could be dissolved if it had broken down irretrievably
	1974: all fault grounds eliminated; no-fault divorce only	1974: more liberal law; divorce by mutual consent
	shortened and simplified divorce procedure; a waiting period of six months required only for couples with minor children or if the couple disagreed on divorce	1987: more complicated divorce procedure; compulsory pre-divorce court-hearing aiming the reconciliation introduced
	spousal alimony almost non-existent	spousal alimony legally possible, in practice nearly non-existent
	property divided evenly between the former spouses independently of its source; only the family home (if acquired for the couple's joint living) and the household goods divided evenly at separation from non-marital cohabiting relationships	property divided evenly between the former spouses independently of its source
	non-resident parent obliged to pay child-support (relatively modest amount); if non-payment state-provided maintenance allowance (requested from the debtor later on)	non-resident parent obliged to pay child-support based on income (20% for one child, 40% for two children, 50% for three or more children)
Child custody:	unmarried mother: sole custody; married parents: joint custody	unmarried mother: sole custody; married parents: joint custody
	joint custody also on request of parents in non-marital consensual union through joint application to the court (since 1977) or to the local tax authority if both parents and the child Swedish citizens (since 1983)	joint custody also on request of the father in non-marital consensual union through application to the court (since 1987), mother's consent required
	mid-1983: divorced/separated parents retain joint custody unless one of them files for its annulment	
	October 1998: court can decide on continued joint custody according to the child's best interest even if one parent disagrees	

Figure 1: *Causality sequence diagram*

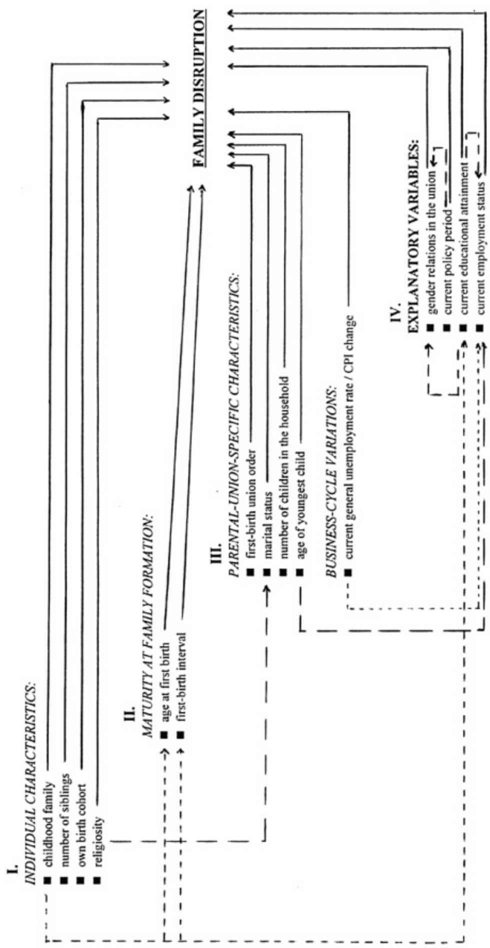


Table S: *Relative risks of dissolution of the first parental union for Swedish women and men. Main results from the final model.*

Standardized for childhood family, age at first birth (conditional on education at first birth), first-birth interval, first-birth union order, marital status, current age of the youngest child, and age of the first child[#].

	women	**men**
current policy period:	*(p = 0.091)*	*(p = 0.014)*
Jan. 1964 - Dec. 1973	0.92	0.91
Jan. 1974 - June 1983	1	1
July 1983 - June 1993	1.30**	1.74***
father took leave after first birth:	*(p = 0.006)*	*(p = 0.529)*
yes	0.70***	0.81
no	1	1
other	1.07	0.94
current educational attainment:	*(p = 0.003)*	*(p = 0.284)*
compulsory education	1	1
lower level vocational school	0.70***	0.75
gymnasium	0.56***	0.78
post-gymnasium	0.63***	0.63*
current employment status[¤]:	*(p = 0.000)*	*(p = 0.116)*
full-time employed	1	1
long part-time employed	0.76*	1.81
short part-time employed	0.41***	3.26*
on parental leave	0.45***	
own household work	0.60***	
unemployed	0.87	3.10**
student	1.51*	1.18
other non-employed	0.47***	0.86
log likelihood	**-2870.3**	**-1189.7**
no. of independent parameters	**33**	**31**
null model log likelihood	**-3016.6**	**-1253.8**

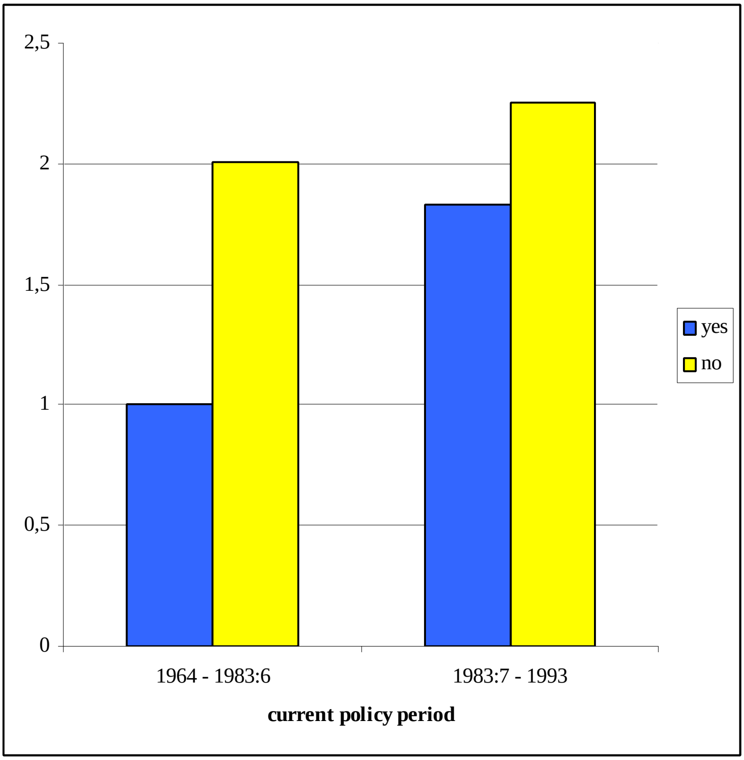
*** significant at the 1%-level; ** at 5%; * at 10%

Note: For each variable, risks and their significance are given relative to the reference level, indicated by 1 (no decimals). The p-value of the entire factor is given in the row containing the variable name

[#] For the full model, see Table S.II in Appendix B.

[¤] The categories "on parental leave" and "own household work" for men are included in the "other non-employed" category.

Figure 2S: *Relative risks of dissolution of the first parental union for Swedish women, by current policy period and according to whether the father took parental leave with the first child*.*

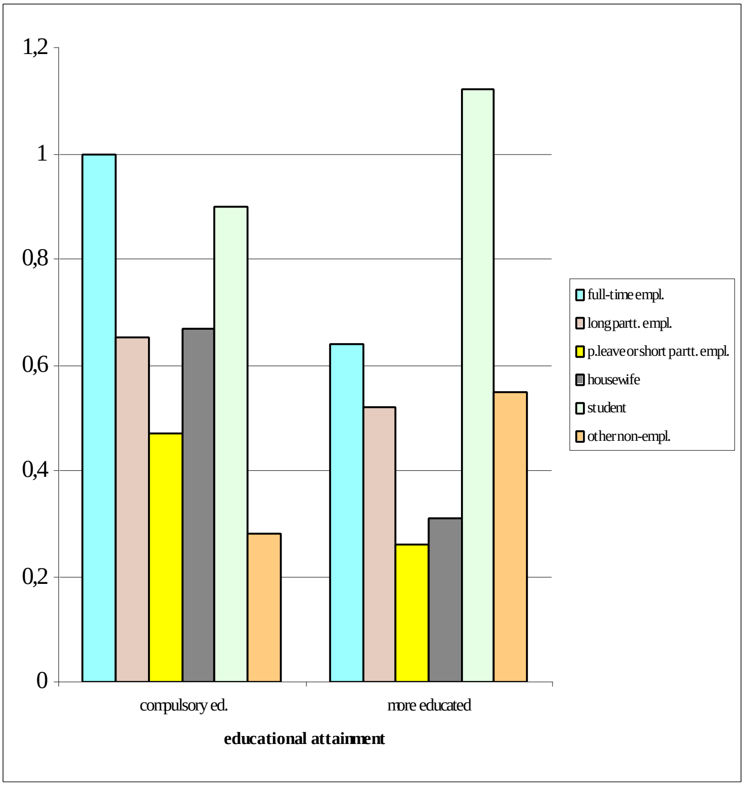


* Standardized for childhood family, age at first birth (conditional on education at first birth), first-birth interval, first-birth union order, marital status, current age of the youngest child, current educational attainment, current employment status, and age of the first child.

Note: (p = 0.056)

Since men were not eligible for parental leave in the first policy period, we collapsed the first two periods into a single category. Also, the categories ”no” and ”other” for father’s leave were collapsed into a single category (called ”no”) given their very similar disruption intensities.

Figure 3S: *Relative risks of dissolution of the first parental union for Swedish women, by current educational attainment and current employment status*.*

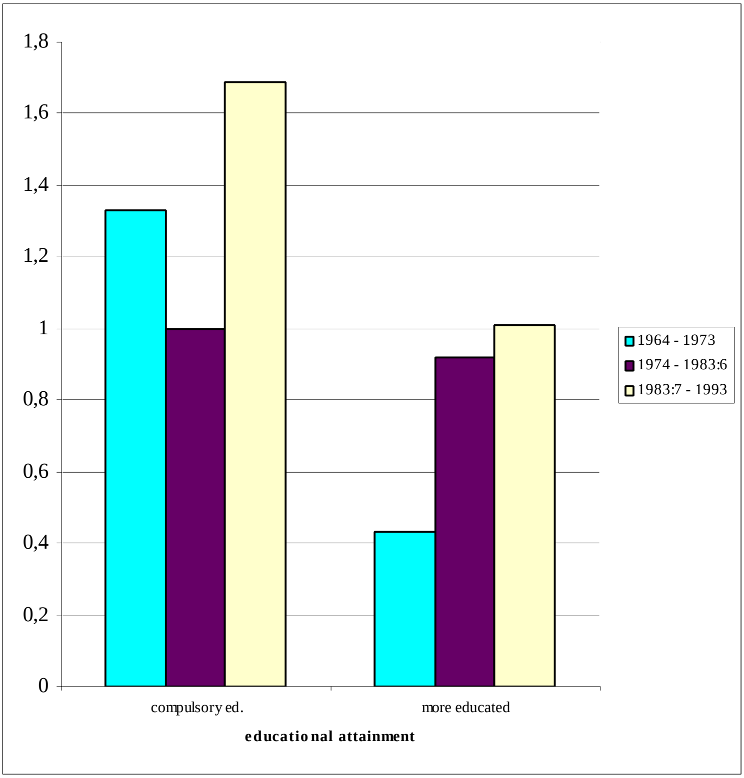


* Standardized for childhood family, age at first birth (conditional on education at first birth), first-birth interval, first-birth union order, marital status, current age of the youngest child, father's use of parental leave, current policy period, and age of the first child.

Note: (p = 0.123)

As the disruption intensities for the categories above compulsory education were not significantly different from each other, we have collapsed them into a single category. For current employment status the category of "unemployed" was included in the "other non-employed" category.

Figure 4S: *Relative risks of dissolution of the first parental union for Swedish women, by current policy period and current educational attainment *.*

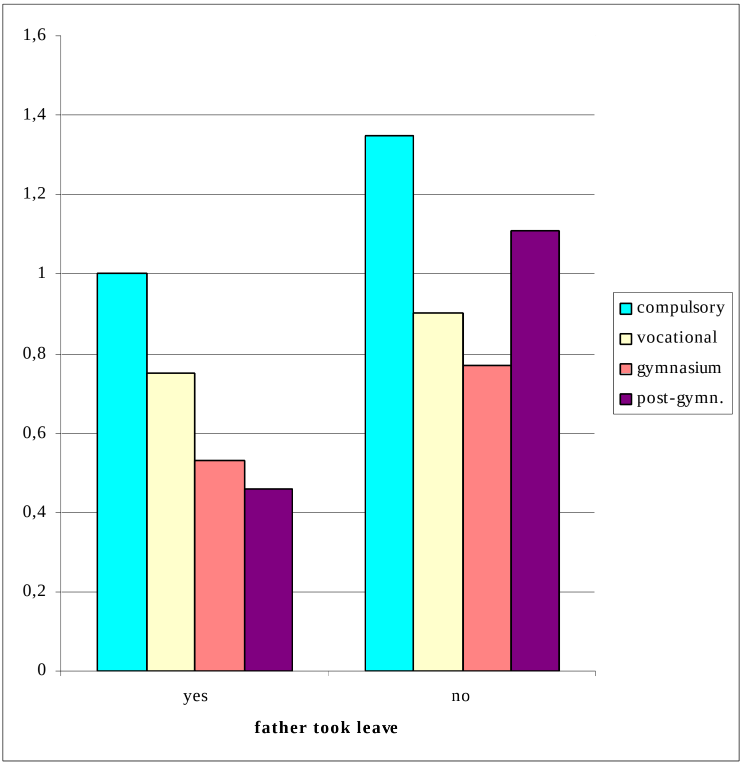


* Standardized for childhood family, age at first birth (conditional on education at first birth), first-birth interval, first-birth union order, marital status, current age of the youngest child, father took leave after first birth, current employment status, and age of the first child.

Note: (p = 0.032)

Those with an educational level above compulsory education are included in the ”more educated” category.

Figure 5S: *Relative risks of dissolution of the first parental union for Swedish women, by current educational attainment and according to whether the father took parental leave with the first child *.*



* Standardized for childhood family, age at first birth (conditional on education at first birth), first-birth interval, first-birth union order, marital status, current age of the youngest child, current policy period, current employment status, and age of the first child.

Note: (p = 0.036)

The category ”no” for father’s leave includes also fathers who were not eligible for parental leave.

Table H: *Relative risks of dissolution of the first parental union for Hungarian women and men. Main results from the final model.*

Standardized for religiosity, childhood family, age at first birth (conditional on education at first birth), first-birth union order, marital status, current age of the youngest child, and age of the first child[#].

	women	**men**
current policy period:	*(p = 0.168)*	*(p = 0.490)*
Jan. 1964 - Dec. 1981	1	1
Jan. 1982 - June 1987	1.34*	1.41
July 1987 - Dec. 1993	1.22	1.33
gender-role attitude:	*(p = 0.686)*	*(p = 0.235)*
egalitarian	1	1
intermediate	1.01	1.29
traditional	0.83	0.61
attitude to parenthood:	*(p = 0.278)*	*(p = 0.639)*
child-centered	1	1
self-centered	1.25	1.33
other	0.89	1.04
current educational attainment:	*(p = 0.148)*	*(p = 0.317)*
compulsory education	1	1
lower level vocational school	1.13	0.65*
gymnasium	0.83	0.60*
post-gymnasium	0.82	0.64
current employment status[¤]:	*(p = 0.049)*	*(p = 0.223)*
full-time employed	1	1
part-time employed	0.75	4.17**
own household work	1.50	
unemployed		1.81
student	0.60**	3.62
other non-employed	0.36*	1.53
log likelihood	**-2716.9**	**-933.9**
no. of independent parameters	**28**	**28**
null model log likelihood	**-2787.4**	**-964.1**

*** significant at the 1%-level, ** at 5%, * at 10%

Note:

For each variable, risks and their significance are given relative to the reference level, indicated by 1 (no decimals).

The p-value of the entire factor is given in the row containing the variable name.

[#] For the full model, see Table H.II in Appendix B.

[¤] For women, the category ”unemployed” is included in the ”other non-employed” category.

Appendix A. Control variables

A.1 Individual characteristics

Our first group of control variables consists of a diversity of individual characteristics. A factor called c*hildhood family* reflects the composition of the family in which the respondent was brought up. For Sweden, we distinguish between intact families and various types of non-intact families. For Hungary, we do not have information about the death of the respondents' parents, so we distinguish only between intact, divorced, and other non-intact families. We include this factor into our analysis because there is convincing evidence of intergenerational transmission of divorce in a large literature from several countries. These earlier findings show that parental divorce greatly increases offspring's family dissolution risks (see McLanahan and Bumpass 1988; Amato 1996 or Feng et al. 1999 for the US; Kiernan and Cherlin 1999 for the UK; Diekman and Engelhardt 1995; Diefenbach 1997 for Germany).

A variable called *number of siblings* counts full and half siblings of the respondents. Those who are single children or who have one sibling only are probably often more strongly individualistic than those who come from larger families and may have different social skills. Everything else equal, the latter group should therefore be less likely to dissolve their relationships.

A variable called *own birth cohort* groups individuals born in the same calendar years. For Sweden we use the data's single year cohorts: 1949, 1954, 1959, 1964, and 1969 for women and 1949, 1959, and 1964 for men. For Hungary we work with the following categories: 1951-1952, 1953-1957, 1958-1962, 1963-1967 and 1968-1972 for women and 1949-1953, 1954-1958, 1959-1963, 1964-1968 and 1969-1973 for men, based on the categorization of the FFS Standard Country Report for Hungary (Kamarás 1999). As family disruption has become more frequent in recent decades, the attitudes toward family stability have also changed. Divorce has no great stigma in the perception of younger cohorts, but it may be more difficult to accept as a part of a 'normal' family career for older cohorts.

A respondent's *religious activity level* is linked to their frequency of church attendance. Those who participate in religious meetings at least once a month are defined as religiously active. The information we use for this variable relates to the time of the interview. We recognize that this may be problematic. We assume, however, that religiosity is relatively constant over a person's life-time, even if dramatic events can alter a person's religious commitment. Given that the countries of our analysis are rather secularized and that the age-range we look at is limited to young adulthood and early-middle age, changes in one's level of religiosity is probably relatively rare among our respondents. As for the effect of religiosity, it has been shown that religiously active individuals have more stable families than others (Hoem and Hoem 1992; Finnäs 1996).

A.2 Maturity at family formation

The second group of control variables allows us to estimate the effect on family stability of different aspects of maturity at family formation. We study the impact of age at first birth and of first-birth interval as the former represents individual maturity, while the latter shows the maturity of the couple at the time of first birth.

Our age variable is defined in relative terms given that the grouping takes into account the respondents' educational level at first birth but also the fact that men have their first child at higher ages than women. The construction of the factor *age at first birth* is based thus on the reasoning that "… the effect of a given covariate on behavior … must depend on its social meaning among a person's peers and not on its average meaning in the population as a whole" (Hoem 1996:334). For each educational level we grouped together ages in approximately 20 per cent intervals (Table S.4; Table H.4). We have included this factor in our analysis on the basis of findings reported in the literature from a number of countries. It is a common finding that an early start of family formation greatly increases the risk of union dissolution (see Morgan and Rindfuss 1985; South and Spitze 1986; Castro Martin and Bumpass 1989 for the US; Berrington and Diamond 1999 for the UK; Hoem and Hoem 1992; Trussell et al. 1992 for Sweden; Finnäs 1996 for Finland).

First-birth interval represents the time between the formation of the first parental union and the first birth. It should pick up how much the partners have matured together as a couple before they become parents. Previous research has shown that the risk of family disruption is higher if a couple has the first child in the first year of their union already, and that the risk decreases greatly if the first birth takes place in the third or a later year of the union (Hoem 1997).

A.3 Parental-union-specific characteristics

Our third group of control variables includes additional characteristics of the respondents' first parental union. *First-birth union order* informs us about whether the first child was born in the respondents' first co-residential relationship or in a higher-order union. A lot of previous research has shown that disruption risk is higher in second or later unions. (See e.g. Hoem 1997; Berrington and Diamond 1999.)

While the previously presented factors were fixed covariates, the rest of our control variables are time-varying covariates. To represent *marital status*, we distinguish between (i) non-marital consensual unions, (ii) “direct” marriages, where the couple married at the start of their first parental union, and (iii) ”transformed” marriages, in which the union started as non-marital cohabiting relationship but was transformed into a marriage by the time the couple’s first child was born or later. As there is no distinction in either Sweden or Hungary between children born to married couples and children of unmarried mothers, we treat in the analysis marriages that were preceded by cohabitation with the same partner as a singly union type after marriage formation independently of when the marriage occurred (i.e. before or after the first birth). The importance of marital status for family stability has been shown for a number of countries. One has found that the risk of separation is much higher in non-marital cohabiting relationships than in marriages, and that a marriage preceded by cohabitation with the same partner has a higher dissolution risk than a direct marriage (see Axinn and Thornton 1992; Thomson and Colella 1992; Lillard et al. 1995 for the US; Berrington and Diamond 1999 for the UK; Bennett et al. 1988, Hoem and Hoem 1992 for Sweden; Finnäs 1996 for Finland).

The *number of children in the household* gives the actual family size of the respondents in any month after the first birth. It has not been a decisive factor of union stability in the US (Waite and Lillard 1991) or Canada (Wu 1995), but previous studies for Sweden have shown that families with two or three children are significantly more stable than other families (Hoem and Hoem 1992; Andersson 1995; Andersson 1997).

A variable called *current age of the youngest child* shows how old the youngest child is in any month if there is more than one child in the family. Otherwise it informs us that there is only one child in the household. In that case the age of the child is of course measured by the time variable and we do not need a second covariate to pick it up. Previous research has shown that the presence of young children (below age 5 or 6) strongly reduces the risk of family disruption (see Bracher et al. 1993 for Australia; Trussell et al. 1992, Andersson 1997 for Sweden).

A.4 Business-cycle variations

Our last group of control variables includes country-specific factors referring to business-cycle variations. For Sweden, we use the annual national *current unemployment rate*. For Hungary we use the consumer price index (CPI), since unemployment did not officially exist during state-socialism and therefore it was not registered before the late 1980s. Our measure for Hungary is based on the level of changes in CPI (for food products) from one year to the next as this greatly affected people’s living standard. We define the categories of these variables lagged by half a year. Previous research has shown that the divorce rate tends to fall following periods of economic prosperity and tends to rise after periods of business-cycle downturns (South 1985).

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Table 1: *Number of respondents excluded from and included in the analysis of the disruption of the first parental union in Sweden and Hungary.*

	Sweden	Hungary
Total number of respondents:	**4984**	**5487**[1]
Cases excluded because of:		
no recorded partnership	492	1108
no recorded biological child	1155	558
grew up in a non-Nordic country (for Sweden only)	269	-
below age 20 at interview (for Hungary only)	-	20
incomplete union and/or childbearing history (for Hungary only)	-	21
single (neither married nor cohabiting) at first birth	223	211
had adopted[2] child and/or partner's child in the first parental union	94	11
first parental union ended in the month when first birth occurred	5	8
first child died in the first parental union	16	50
Total number of excluded cases:	*2254*	*1987*
Total number of cases used:	**2730**	**3500**
Number of women included in the analysis:	**1869**	**2430**
Number of cases where the first parental union was disrupted:	383	333
Number of cases censored 16 years after the first birth:	360	406
Number of cases censored at the death of the partner in this union:	9	42
Number of cases censored at interview:	1117	1649
Number of men included in the analysis:	**861**	**1070**
Number of cases where the first parental union was disrupted:	156	111
Number of cases censored 16 years after the first birth:	210	202
Number of cases censored at the death of the partner in this union:	2	7
Number of cases censored at interview:	493	750

1 The number of respondents for Hungary is larger than it is in the Standard Recode File that can be requested from the Population Activities Unit in Geneva because only fully completed interviews are included there. In the present study we work with the original Hungarian FFS files provided by Statistics Hungary.

2 Since there were only one such respondent left in our Swedish working sample plus two respondents in the Hungarian sample, we have chosen to exclude them instead of to censor them at the adoption date.

Table S.2: *Distributions of Swedish respondents at the various levels of fixed covariates. Parents in their first-birth union.*

	Women		men	
	# of respondents	per cent	# of respondents	per cent
own birth cohort:				
1949	464	24.8	425	49.3
1954	471	25.2		
1959	469	25.1	209	24.3
1964	327	17.5	227	26.4
1969	138	7.4		
childhood family:				
intact family	1506	80.6	712	82.7
parents divorced	234	12.5	97	11.3
parent died	58	3.1	22	2.5
other non-intact family	71	3.8	30	3.5
number of siblings:				
no sibling	141	7.5	67	7.8
one sibling	587	31.4	261	30.3
two siblings	526	28.2	256	29.7
three or more siblings	615	32.9	277	32.2
religious activity level:				
active	174	9.3	58	6.7
not active	1695	90.7	803	93.3
first-birth union order:				
1	1565	83.7	728	84.6
2	261	14.0	112	13.0
3+	43	2.3	21	2.4
age at first birth[#] (conditional on educational level at first birth):				
very young	311	16.6	171	19.9
rather young	478	25.6	196	22.8
medium	421	22.5	213	24.7
rather old	310	16.6	137	15.9
oldest	349	18.7	144	16.7
first-birth interval (in months):				
< 8	180	9.6	88	10.2
8 - 17	362	19.4	169	19.6
18 - 35	546	29.2	243	28.2
36 - 59	405	21.7	214	24.9
60 +	376	20.1	147	17.1
father took leave after first birth:				
yes, 3+ months	212	11.3	91	10.6
yes, 1 - 2 months	189	10.1	83	9.6
yes, < 1 month	627	33.6	335	38.9
no	745	39.9	296	34.4
other	96	5.1	56	6.5
total:	**1869**	**100.0**	**861**	**100.0**

[#] See Table S.4 for details.

Table S.3: *Exposure time in person-half-months at the various levels of time-varying covariates and of the time variable for Swedish respondents.*

	women exposure time	per cent	men exposure time	per cent
marital status:				
non-marital consensual union	109254	29.4	49666	28.0
transformed marriage (cohab. at start)	223256	60.2	111910	63.0
direct marriage	38600	10.4	15972	9.0
no. of children in the household:				
one	149052	40.2	71534	40.3
two	168712	45.4	80970	45.6
three or more	53346	14.4	25044	14.1
current age of the youngest child:				
< 1 year	40818	11.0	18462	10.4
1 - 2 years	66640	18.0	30190	17.0
3 - 5 years	60254	16.2	28746	16.2
6 + years	54346	14.6	28616	16.1
only one child in the household	149052	40.2	71534	40.3
current educational attainment:				
compulsory education	104116	28.1	53832	30.3
lower level vocational school	150458	40.5	65180	36.7
gymnasium	35374	9.5	23710	13.4
post-gymnasium	81162	21.9	34826	19.6
current employment status:				
full-time employed	83358	22.5	164734	92.8
long part-time employed	75388	20.3	2118	1.2
short part-time employed	62378	16.8	1418	0.8
on parental leave	75454	20.3	778	0.5
own household work	48692	13.1	342	0.2
unemployed	3076	0.8	1260	0.7
student	10720	2.9	4348	2.4
other non-employed	12044	3.3	2550	1.4
current policy period:				
January 1964 - December 1973	18578	5.0	7211	4.1
January 1974 - June 1983	127571	34.4	65404	36.8
July 1983 - June 1993	224961	60.6	104933	59.1
current (national) unemployment rate:				
< 2.0 %	113670	30.6	54726	30.8
2.0 - 2.9 %	159544	43.0	76054	42.8
> = 3.0 %	97896	26.4	46768	26.4
age of first child (time variable):				
< 1 year	43162	11.7	19940	11.2
1 - 2 years	75114	20.2	35002	19.7
3 - 5 years	89162	24.0	40736	23.0
6 - 7 years	46834	12.6	21554	12.1
8 - 11 years	71338	19.2	34958	19.7
12 - 15 years	45500	12.3	25358	14.3
total:	**371110**	**100.0**	**177548**	**100.0**

Table S.4a: *Age at first birth, conditional on educational level at first birth. Swedish women.*

Age	age-groups according to education at first birth							
	compulsory education		low vocational school		gymnasium		post-gymnasium	
	ages in years	per cent	ages in years	per cent	ages in years	per cent	ages in years	per cent
very young	15 - 18	19.0	17 - 20	14.9	18 - 21	12.0	21 - 24	19.3
rather young	19 - 20	23.8	21 - 22	28.4	22 - 23	26.1	25 - 26	22.5
medium	21 - 22	23.8	23 - 24	21.6	24 - 26	26.6	27 - 28	20.6
rather old	23 - 24	15.5	25 - 26	17.4	27 - 28	15.0	29 - 30	17.4
oldest	25 +	17.9	27 +	17.7	29 +	20.3	31 +	20.2
Total	499		749		207		414	
per cent of total N (1869)		26.7		40.1		11.1		22.1

Table S.4b: *Age at first birth, conditional on educational level at first birth. Swedish men.*

Age	age-groups according to education at first birth							
	compulsory education		low vocational school		gymnasium		post-gymnasium	
	years	per cent	years	per cent	years	per cent	years	per cent
very young	18 - 21	19.3	18 - 22	20.1	20 - 23	20.1	21 - 25	20.0
rather young	22 - 23	21.7	23 - 24	23.0	24 - 25	19.3	26 - 27	22.5
medium	24 - 25	22.2	25 - 26	22.0	26 - 27	29.9	28 - 29	22.6
rather old	26 - 27	16.6	27 - 28	16.3	28 - 29	16.7	30 - 31	13.6
oldest	28 +	20.2	29 +	14.6	30 +	14.0	32 +	17.3
Total	253		339		114		155	
per cent of total N (861)		29.4		39.4		13.2		18.0

Table H.2: *Distributions of Hungarian respondents at the various levels of fixed covariates. Parents in their first-birth union.*

	women		men	
	# of respondents	**per cent**	**# of respondents**	**per cent**
cohort:				
women: 1951-52; men: 1949-53	174	7.2	299	27.9
women: 1953-57; men: 1954-58	828	34.0	319	29.8
women: 1958-62; men: 1959-63	639	26.3	258	24.1
women: 1963-67; men: 1964-68	500	20.6	151	14.2
women: 1968-72; men: 1969-73	289	11.9	43	4.0
childhood family:				
intact family	2023	83.2	945	88.3
parents divorced	257	10.6	83	7.8
other non-intact family	150	6.2	42	3.9
number of siblings:				
no sibling	336	13.8	156	14.6
one sibling	1041	42.8	477	44.6
two siblings	528	21.8	224	20.9
three or more siblings	525	21.6	213	19.9
religious activity level:				
active	336	13.8	106	9.9
not active	2094	86.2	964	90.1
first-birth union order:				
1	2343	96.4	1018	95.1
2+	87	3.6	52	4.9
age at first birth* (conditional on educational level at first birth):				
very young	356	14.7	217	20.3
rather young	667	27.4	251	23.5
medium	564	23.2	262	24.5
rather old	446	18.4	178	16.6
oldest	397	16.3	162	15.1
first-birth interval (in months):				
< 8	512	21.1	216	20.2
8 - 17	955	39.3	394	36.8
18 - 35	633	26.0	307	28.7
36 - 59	229	9.4	114	10.7
60 +	101	4.2	39	3.6
attitude to parenthood*:				
child-centered	1754	72.2	774	72.3
self-centered	227	9.3	108	10.1
other	449	18.5	188	17.6
gender-role attitude*:				
egalitarian	589	24.2	302	28.2
intermediate	1684	69.3	721	67.4
traditional	157	6.5	47	4.4
total:	**2430**	**100.0**	**1070**	**100.0**

* See Tables H.4 and H.5 for details.

Table H.3: *Exposure time in person-half-months at the various levels of time-varying covariates and of the time variable for Hungarian respondents.*

	women		men	
	exposure time	**per cent**	**exposure time**	**per cent**
marital status:				
non-marital consensual union	7108	1.3	1716	0.7
transformed marriage (cohab. at start)	46882	8.9	25444	10.5
direct marriage	474932	89.8	214520	88.8
no. of children in the household:				
one	235550	44.5	111600	46.2
two	247952	46.9	111696	46.2
three or more	45420	8.6	18384	7.6
current age of the youngest child:				
< 1 year	44974	8.5	19882	8.2
1 - 2 years	75740	14.3	33976	14.1
3 - 5 years	81174	15.4	36474	15.1
6 + years	91484	17.3	39748	16.4
only one child in the household	235550	44.5	111600	46.2
current educational attainment:				
compulsory education	157572	29.8	38238	15.8
lower level vocational school	120280	22.8	108298	44.8
gymnasium	187438	35.4	66040	27.3
post-gymnasium	63632	12.0	29104	12.1
current employment status:				
full-time employed	436828	82.6	208944	86.5
long part-time employed	16240	3.1	1182	0.5
short part-time employed	7678	1.5	594	0.2
irregular time employed	17458	3.3	23514	9.7
unemployed	1168	0.2	1230	0.5
own household work	13364	2.5	0	0.0
student	5412	1.0	780	0.3
other non-employed	30774	5.8	5436	2.3
current policy period:				
January 1964 - December 1981	127853	24.2	51039	21.1
January 1982 - June 1987	178681	33.8	75948	31.4
July 1987 - December 1993	222388	42.0	114693	47.5
CPI change (for food products only):				
< 5.0 %	134925	25.5	56072	23.2
5.0 - 9.9 %	152756	28.9	64110	26.5
> = 10.0 %	241241	45.6	121498	50.3
age of first child (time variable):				
< 1 year	56926	10.8	25104	10.4
1 - 2 years	103148	19.5	46504	19.3
3 - 5 years	130402	24.7	59608	24.7
6 - 7 years	70840	13.4	32430	13.4
8 - 11 years	106538	20.1	49408	20.4
12 - 15 years	61068	11.5	28626	11.8
total:	**528922**	**100.0**	**241680**	**100.0**

Table H.4a: *Age at first birth, conditional on educational level at first birth. Hungarian women.*

Age	age-groups according to education at first birth							
	compulsory education		low vocational school		gymnasium		post-gymnasium	
	ages in years	per cent	ages in years	per cent	ages in years	per cent	ages in years	per cent
very young	14 - 17	16.8	16 - 18	15.7	17 - 19	12.1	20 - 22	13.6
rather young	18 - 19	31.2	19	17.4	20 - 21	30.3	23 - 24	30.3
medium	20	16.4	20 - 21	32.5	22 - 23	24.3	25	17.7
rather old	21 - 22	18.7	22 - 23	18.7	24 - 25	17.8	26 - 27	18.1
oldest	23 - 34	16.9	24 - 34	15.7	26 - 37	15.5	28 - 39	20.3
Total	763		587		859		221	
per cent of total N (2430)		31.4		24.1		35.4		9.1

Table H.4b: *Age at first birth, conditional on educational level at first birth. Hungarian men.*

Age	age-groups according to education at first birth							
	compulsory education		low vocational school		gymnasium		post-gymnasium	
	ages in years	per cent	ages in years	per cent	ages in years	per cent	ages in years	per cent
very young	17 - 20	22.4	17 - 21	20.7	18 - 22	18.1	21 - 24	20.2
rather young	21 - 22	20.3	22 - 23	24.1	23 - 24	25.3	25	21.2
medium	23 - 24	24.5	24 - 25	22.6	25 - 26	25.6	26 - 27	30.3
rather old	25 - 26	18.8	26 - 27	17.2	27 - 28	16.2	28 - 29	11.1
oldest	27 - 36	14.0	28 - 39	15.4	29 - 39	14.8	30 - 35	17.2
Total	192		482		297		99	
per cent of total N (1070)		17.9		45.0		27.8		9.3

Table H.5: *Composition of the two attitude variables for Hungarian respondents.*

I. ”Attitude to parenthood”:

Respondents had to choose the statement with which they agreed most strongly:

A, ”It is the parents’ duty to do their best for their children even at the expense of their own well-being”
B, ”Parents have lives of their own and should not be asked to sacrifice their own well-being for sake of their children”
C, Neither statement
D, Don’t know

Our categories for this variable:

”Child-centered”	- those who chose statement A
”Self-centered”	- those who chose statement B
”Other”	- all other respondents

II. ”Gender-role attitude”:

Agreement or disagreement with the following statements:

1, ”If a single woman wants to have children without living together with a man, she should be allowed to do what she likes.”
2, ”I do not mind any sacrifices in order to have a good relationship with my spouse/partner, even if it jeopardizes my other goals.”
3, ”My career is very important for me.”

Our categories for this variable:

”Egalitarian”	- those who agreed with all three statements
”Traditional”	- women who disagreed with statement 1 and 3
	- men who disagreed with statement 1 and 2
”Intermediate”	- all other respondents

APPENDIX B. Stepwise model presentation

B1. Sweden

Our main findings (for our explanatory variables) are discussed in detail in the main text of our paper. Here we present the disruption-risk profile for the control variables and for the time factor.

We started the model fitting with the control variables (Table S.I) referring to individual characteristics, then we included factors on maturity at family formation. In the process of adding new groups of control variables to the model, we kept only those factors from the previous groups that proved to be important for the first-parental-union break-up in either the female or the male samples (i.e. significantly influenced the disruption intensity).

Own birth cohort (Model 1), as expected, shows that the risks of family dissolution increased over cohorts. The patterns are similar in the female and male samples, but the disruption risks more than doubled for the 1959 and 1964 male cohort compared with that for the 1949 male cohort, while the increases for female cohorts of the same years are more moderate. We see, however, a three-fold increase of dissolution risk for the youngest female cohort as compared with the oldest one. Although cohort influenced family stability significantly, we did not include it in the next models. Adding age at first birth in combination of duration since first birth (i.e. age of the first child) as our time variable would otherwise result in an overdetermined model.

Mother's *religious activity level* has a stronger impact on union disruption than that of father's, but the patterns are very similar. Those who are religiously active have a much lower risk of family dissolution than other individuals, as usual.

Childhood family proved to be very important for family stability also in Sweden. Individuals whose parents divorced before the respondent's 16th birthday, have about twice as high risk of family dissolution as those who came from intact families. The risk of disruption is also very high for individuals from other non-intact families. There is no significant difference in the probability of union dissolution between those who experienced the death of a parent in their childhood and those who were brought up by both of their parents. As we see in Models 2 - 4 in (Table S.I), childhood family has a strong direct impact on family stability as its effect hardly changed when we included factors of maturity at family formation and the rest of the control variables in the model.

A parent's *number of siblings* (Model 1) has apparently no influence on family disruption in Sweden. This suggests that those who came from small families are not more individualistic and are not less capable of compromises in family life than individuals who were brought up in larger families.

Maturity at family formation, both in its individual and couple aspects (Model 2), is important for family stability. *Age at first birth (conditional on educational level at first birth)* shows that those who start family formation at younger ages have much higher risk of union disruption than later starters. This is in line with previous findings in the literature.

For *first-birth interval*, we find that the partners should mature together as a couple before they become parents, for about three years. Those who wait with childbearing have much more stable families than couples who became parents within relatively short time after they had moved in together.

Next, we added to the model the rest of the control variables (Models 3 and 4). With the inclusion of marital status in the model, the effect of religiosity disappeared. This means that religiosity has no direct impact on family stability in the secularized Swedish society, only an indirect effect which works through marital status.

First-birth union order has a significant impact on parent’s family stability. Individuals who had their first child in their second or higher-order union, have more than 1.5 times as high a risk of family dissolution as those who became parents in their first co-residential relationship. This, again, is in line with previous findings in the literature.

Marital status, which is a time-varying covariate as are the control variables we discuss in the followings, is another factor of great importance for family stability. Living in a non-marital consensual union strongly increases the risk of family dissolution. Direct marriages are the most stable relationships, while marriages transformed from cohabiting relationships have an intermediary position, as expected.

For *number of children in the household* (Model 3) we find that those who have one child only have the highest risk of family disruption among the parents we studied. There is no significant difference among two-child parents and those with three or more children in the risks for family break-up. Since our current age of the youngest child variable controls also for the effect of having one child only (and the effect of the child’s age is measured then by our time variable), we used this factor instead of number of children in the household in the further analysis.

Current age of the youngest child (Model 4) also proved to be important for union stability. The protective effect of having another child in the family works only while this child is very young, that is below age 1. With children of age 3 and older, the risk of family dissolution is nearly the same as for one-child families.

Current unemployment rate (Models 3 and 4), which controls for business-cycle variations, had hardly any influence on family break-up, probably because unemployment levels were rather low in Sweden during the period covered with our data. In the male sample however, we see that the relative risk of disruption of the first parental union is lower in times of slightly higher unemployment rates than otherwise.

Finally, we included our explanatory variables in the model (Table S.II) which are discussed in details in the main text of our paper (Section 6.1). Thus it only remains to describe the impact of our time variable, *age of the first child.* While we hardly see any changes in union-disruption risks with children below 6 years of age in the female sample, the male sample produces steeply increasing dissolution intensities from the first child's infancy to pre-school years, which thereafter decrease. When the first child enters her/his teenage years, the risk of family disruption for parents is at the level of that of just becoming a parent or even lower, ceteris paribus, as seen in both samples. This may be a selection effect, namely that relationships that did not break up by the time the first child becomes a teenager are increasingly selected to be more stable than other unions.

B2. Hungary

As we followed the same procedure for the model fitting for Hungary (Table H.I) as for Sweden, we discuss first the risks profile for the individual characteristics. The patterns are quite similar in the female and male samples for *own birth cohort* (Model 1). Disruption risks increased substantially (more than three times) for the youngest cohorts as compared to the oldest cohorts, like in Sweden. However, we do not find significant changes in dissolution risks for cohorts born in the 1950s and early 1960s for Hungary, in contrast to Sweden.

Mothers' and fathers' *religious activity level* shows similar patterns with much lower disruption risks for those who are religiously active than for others, as usual. The impact of women's religiosity is much stronger than is that of men's, similarly to the results for Sweden. In contrast to Sweden however, the effect of religiosity on family stability remains for Hungary even after we added to the model all our control variables (Models 2 - 4). This suggests that the religiously active individuals are a distinct group regarding family disruption behavior as compared to other families with children in Hungary. This also means that the Hungarian society is less secularized than the Swedish society.

Men's *childhood family* does not have a significant impact on family disruption in Hungary, but the childhood families of women do influence the risk (Model 1). Yet we see in both samples that those who experienced their parents' divorce before their 16th birthday have much higher disruption risk than those coming from intact families, as in Sweden. Women who were brought up in other non-intact families have the least stable families of all as adults. Interestingly, as we included marital status in the model, the influence of childhood family becomes weaker and appears only for women coming from other non-intact families (Models 3 and 4). This finding indicates that the daughters of divorced parents probably more often live in consensual unions or transformed marriages than other women, and thus when we controlled for the type of the relationship, the influence of parental divorce disappears.

Women's *number of siblings* is not important for family stability, but men's sibship size is negatively related to the risk of disruption. Hungarian fathers who do not have any sibling have twice as high risk of family disruption than men with siblings and the difference is significant (Model 1). These men seem to have fewer skills of solving problems that arise in a union, or they may be less sensitive to the needs of other family members. We had no similar finding for Sweden, interestingly though. This might be explained by the Hungarian society being more traditional than the Swedish especially regarding gender roles within the family. Only sons are treated differently than only daughters by their parents. For example, the daughters usually are required to actively participate in domestic work but not the sons. When there are at least two children in the family, even the male children learn to pay attention to others and to cooperate. However, as we added variables that accounted for union-specific characteristics, the influence of the sibling factor on family stability was reduced and lost significance even in the male sample (Models 3 and 4). Thus it was not included in our final model (Table H.II).

In the next step we added the group of variables reflecting maturity at family formation to the model (Model 2). We find that women's *age at first birth (conditional on educational level at first birth)* has a strong effect on risks of family disruption but men's age does not. The results for women's age at first birth resemble the Swedish patterns. Those who start family formation at very young ages have much higher risk of union dissolution than later starters in both countries. For men's age in Hungary we find significantly lower disruption risk only for those who became fathers at a medium age, but not for older ages at first birth.

First-birth interval has little impact on family stability in Hungary in both samples, unlike in Sweden. This may be explained by a longer dating period before marriage or non-marital cohabitation in Hungary as compared to Sweden. Because of the housing shortage, Hungarian couples often have had to search for a long time to find a dwelling for themselves. Even among married couples, the majority started their lives together in the home of the parents of one of the spouses (Kamarás 1999). Thus parental consent may be a condition of young couples' co-residence, and most parents want to get to know their children's partner better before giving their consent. Since the young couples also have more time to learn to know each other before moving in together, many who have weak bonds to each other will break up before they can start living together. Therefore, this variable is a less efficient a measure of the maturity of the partners as a couple for Hungary than it is for Sweden.

Next we included in the model the rest of the control variables (Models 3 and 4). Men's *first-birth union order* has little influence on family disruption, but women's union order does. Nevertheless, the patterns are similar; those who have their first child in their first co-residential relationship have stabler unions than others, like in Sweden.

As usual, *marital status* is very influential for family stability. The patterns resemble the findings for Sweden, even though consensual unions are a tiny fraction of all families with children in Hungary, and direct marriages represent the typical family form in contrast to Sweden. Parents who live in consensual unions have a very high disruption risk in both countries. The most stable relationships are direct marriages, and marriages which started as consensual unions (i.e. transformed marriages) have an intermediary position.

Number of children in the household (Model 3) influences family disruption very strongly for parents in Hungary. Similarly to what is the case in Sweden, one-child parents have the highest risk of family disruption, and there is no significant difference between two-child parents and those with more children.

Current age of the youngest child (Model 4) is also an important indicator of family stability in both countries. When there are at least two children in the family, the risk of union dissolution increases with the age of the youngest child in both Sweden and Hungary. Yet, only the disruption intensities of one-child parents are significantly higher than that of larger families with at least one infant in Hungary.

CPI change (Models 3 and 4), which measures the effect of business-cycle variations, had no influence on family stability in Hungary. As our data covers mostly the period of state socialism, we can conclude that the relatively stable macro-economic situation of that time made family-disruption decisions less sensitive to macro-economic changes. For Sweden, business-cycle variations were somewhat more important for family stability (as we saw in the male sample), probably due to the differences in the economic structures of the two countries.

Our final model (Table H.II) shows a somewhat different disruption-risk profile for the time variable (*age of the first child)* than the one we saw for Sweden. As the first child grows older the risk of family dissolution increases for Hungarian parents and declines somewhat only during the teenage-years of the first child. In Sweden the disruption intensities decrease already from the early- and middle school-age years of the first child.

We can only speculate about the reason of the continuously high dissolution risk for parents with school-age children in Hungary. The educational system in Hungary strongly encourages competition among children, who constantly face a rather high level of stress, and they may possibly transfer this to the parents. Thus the tension in the family would not decrease as the children grow older. If the parents have problems in their own relationship, this can make the situation worse and in the long run eventually cause the break-up of the family. The Swedish school-system is much less competitive, at least in the first seven grades. Our other explanation relates to the economic difficulties a family with school-aged children might experience. As children (or their parents) were expected to buy their textbooks and other school materials in Hungary, which were more and more expensive as the child progressed in the higher grades, this was quite a heavy burden on the family budget. As we know financial problems often create tensions. If the relationship was not strong enough, such tensions could lead to the dissolution of the union, especially in combination with other types of difficulties. In Sweden, school materials and books are usually provided free of charge by the school, at least below secondary-school level. Thus the family’s expenses for education are relatively limited in Sweden.

Table S.I: *Relative risks of dissolution of the first parental union for Swedish women and men. Models with the control variables.*

	(1) women	(1) men	(2) women	(2) men	(3) women	(3) men	(4) women	(4) men
own birth cohort:	*(p = 0.000)*	*(p = 0.000)*						
1949	1	1						
1954	1.19							
1959	1.36**	2.29***						
1964	1.66***	2.68***						
1969	3.16***							
religiosity (activity level):	*(p = 0.024)*	*(p = 0.167)*	*(p = 0.015)*	*(p = 0.127)*	*(p = 0.339)*	*(p = 0.757)*	*(p = 0.366)*	*(p = 0.720)*
active	0.64**	0.59	0.62**	0.56	0.82	0.87	0.83	0.85
not active	1	1	1	1	1	1	1	1
childhood family:	*(p = 0.000)*	*(p = 0.004)*	*(p = 0.000)*	*(p = 0.001)*	*(p = 0.000)*	*(p = 0.004)*	*(p = 0.000)*	*(p = 0.003)*
intact family	1	1	1	1	1	1	1	1
parents divorced	1.85***	2.03***	1.92***	2.11***	1.83***	1.90***	1.82***	1.90***
parent died	1.15	0.66	1.06	0.48	1.03	0.48	1.03	0.46
other non-intact family	1.88***	2.36**	1.98***	2.20**	1.71**	2.10**	1.70**	2.13**
number of siblings:	*(p = 0.374)*	*(p = 0.752)*						
none	1.23	1.06						
one	1	1						
two	1.24	1.13						
three or more	1.21	0.90						
age at first birth (conditional on education at first birth):			*(p = 0.079)*	*(p = 0.025)*	*(p = 0.001)*	*(p = 0.008)*	*(p = 0.001)*	*(p = 0.006)*
very young			1	1	1	1	1	1
rather young			0.90	0.81	0.85	0.83	0.85	0.82
medium			0.84	0.70	0.76*	0.66*	0.76*	0.66*
rather old			0.68**	0.43***	0.58***	0.40***	0.58***	0.40***
oldest			0.60***	0.45***	0.44***	0.36***	0.44***	0.35***
first-birth interval:			*(p = 0.000)*	*(p = 0.196)*	*(p = 0.001)*	*(p = 0.464)*	*(p = 0.001)*	*(p = 0.510)*
< 8 months			1	1	1	1	1	1
8 - 17 months			0.90	1.18	0.92	1.11	0.92	1.11
18 - 35 months			0.78	1.02	0.82	1.08	0.82	1.07
36 - 59 months			0.43***	0.80	0.49***	0.99	0.49***	0.98
60 + months			0.52***	0.53	0.63**	0.57	0.63**	0.58
first-birth union order:					*(p = 0.000)*	*(p = 0.007)*	*(p = 0.000)*	*(p = 0.006)*
1					1	1	1	1
2 +					1.82***	1.92***	1.82***	1.93***
marital status:					*(p = 0.000)*	*(p = 0.000)*	*(p = 0.000)*	*(p = 0.000)*
consensual union					1.80***	2.09***	1.81***	2.18***
transformed marriage					1	1	1	1
direct marriage					0.55***	0.82	0.54***	0.81
no. of children in the household:					*(p = 0.000)*	*(p = 0.034)*		
one					1.83***	1.45*		
two					1	1		
three or more					0.97	0.61		

(continued on the following page)

Table S.I (continuation):

	(1) women	(1) men	(2) women	(2) men	(3) women	(3) men	(4) women	(4) men
current age of the youngest child:							*(p = 0.000)*	*(p = 0.008)*
< 1 year							1	1
1 - 2 years							2.01***	1.99*
3 - 5 years							2.39***	3.18***
6 + years							2.07**	3.28***
only 1 child in the household							3.17***	3.01***
current unemployment rate:					*(p = 0.521)*	*(p = 0.192)*	*(p = 0.573)*	*(p = 0.180)*
< 2.0%					1	1	1	1
2.0% - 2.9%					0.97	0.73*	0.97	0.72*
>= 3.0%					1.12	0.76	1.11	0.75
age of first child (time factor):	*(p = 0.003)*	*(p = 0.032)*	*(p = 0.000)*	*(p = 0.002)*	*(p = 0.054)*	*(p = 0.005)*	*(p = 0.015)*	*(p = 0.001)*
< 1 year	1	1	1	1	1	1	1	1
1 - 2 years	1.26	1.89**	1.19	1.74*	1.44**	2.04**	1.50***	2.19***
3 - 5 years	1.05	2.37***	0.92	1.88**	1.63***	3.05***	1.61***	3.22***
6 - 11 years	0.70*	1.81*	0.55***	1.15	1.25	2.56***	1.07	1.93*
12 - 15 years	0.81	1.25	0.55***	0.66	1.40	1.63	1.19	1.05
	[0.694]	**[0.327]**	**[1.946]**	**[0.879]**	**[0.667]**	**[0.391]**	**[0.384]**	**[0.185]**
log likelihood	**-2967.8**	**-1221.5**	**-2952.6**	**-1221.0**	**-2909.2**	**-1201.8**	**-2903.7**	**-1198.3**
no. of independent parameters	**16**	**14**	**17**	**17**	**24**	**24**	**26**	**26**

*** significant at the 1%-level, ** at 5%, * at 10%.

Note: For each variable, risks and their significance are given relative to the reference level, indicated by 1 (no decimals). The p-value of the entire factor is given in the row containing the variable name. Absolute risk (per 1000 person-half-months) for age < 1 year of the first child is given in the last row of the time factor in boldface letter.

Table S.II: *Relative risks of dissolution of the first parental union for Swedish women and men. Final model.*

	women	men
childhood family:	*(p = 0.001)*	*(p = 0.022)*
intact family	1	1
parents divorced	1.63***	1.73**
parent died	0.93	0.49
other non-intact family	1.69**	1.82
age at first birth (conditional on educational level at first birth):	*(p = 0.000)*	*(p = 0.000)*
very young	1	1
medium (incl. rather young)	0.77**	0.73
older (i.e. rather old and oldest)	0.48***	0.33***
first-birth interval:	*(p = 0.004)*	*(p = 0.130)*
< 3 years	1	1
3 - 4 years	0.61***	0.88
5 + years	0.79	0.50**
first-birth union order:	*(p = 0.000)*	*(p = 0.026)*
1	1	1
2 +	1.75***	1.71**
marital status:	*(p = 0.000)*	*(p = 0.000)*
non-marital consensual union	1.80***	1.99***
transformed marriage	1	1
direct marriage	0.55***	0.74
current age of the youngest child:	*(p = 0.066)*	*(p = 0.004)*
< 1 year	1	1
1 - 2 years	1.61*	1.96*
3 - 5 years	1.83**	3.29***
6 + years	1.53	3.22***
only one child in the household	2.04***	3.16***
father took leave after first birth:	*(p = 0.006)*	*(p = 0.529)*
yes	0.70***	0.81
no	1	1
other	1.07	0.94
current policy period:	*(p = 0.091)*	*(p = 0.014)*
Jan. 1964 - Dec. 1973	0.92	0.91
Jan. 1974 - June 1983	1	1
July 1983 - June 1993	1.30**	1.74***
current educational attainment:	*(p = 0.003)*	*(p = 0.284)*
compulsory education	1	1
lower level vocational school	0.70***	0.75
gymnasium	0.56***	0.78
post-gymnasium	0.63***	0.63*
current employment status[¤]**:**	*(p = 0.000)*	*(p = 0.116)*
full-time employed	1	1
long part-time employed	0.76*	1.81
short part-time employed	0.41***	3.26*
on parental leave	0.45***	
own household work	0.60***	
unemployed	0.87	3.10**
student	1.51*	1.18
other non-employed	0.47***	0.86

(continued on the following page)

Table S.II (continuation):

	women	**men**
age of first child (time variable):	*(p = 0.022)*	*(p = 0.000)*
< 1 year	1	1
1 - 2 years	1.07	2.21***
3 - 5 years	0.99	3.30***
6 - 11 years	0.61**	1.76
12 - 15 years	0.58*	0.80
	[1.445]	**[0.177]**
log likelihood	**-2870.3**	**-1189.7**
no. of independent parameters	**33**	**31**

*** significant at the 1%-level, ** at 5%, * at 10%.

Note: For each variable, risks and their significance are given relative to the reference level, indicated by 1 (no decimals). The p-value of the entire factor is given in the row containing the variable name. Absolute risk (per 1000 person-half-months) for age < 1 year of the first child is given in the last row of the time factor in boldface letter.

¤ The categories ”on parental leave” and ”own household work” for men are included in the ”other non-employed” category.

Table H.I: *Relative risks of dissolution of the first parental union for Hungarian women and men. Models with the control variables.*

	(1) women	(1) men	(2) women	(2) men	(3) women	(3) men	(4) women	(4) men
own birth cohort:	*(p = 0.003)*	*(p = 0.142)*						
w: 1951-52; m: 1949-53	1	1						
w: 1953-57; m: 1954-58	1.36	0.71						
w: 1958-62; m: 1959-63	1.39	0.97						
w: 1963-67; m: 1964-68	1.74**	1.08						
w: 1968-72; m: 1969-73	3.31***	4.07*						
religiosity (activity level):	*(p = 0.002)*	*(p = 0.274)*	*(p = 0.002)*	*(p = 0.290)*	*(p = 0.003)*	*(p = 0.477)*	*(p = 0.003)*	*(p = 0.501)*
active	0.58***	0.68	0.58***	0.69	0.59***	0.77	0.59***	0.79
not active	1	1	1	1	1	1	1	1
childhood family:	*(p = 0.024)*	*(p = 0.202)*	*(p = 0.028)*	*(p = 0.188)*	*(p = 0.071)*	*(p = 0.349)*	*(p = 0.069)*	*(p = 0.315)*
intact family	1	1	1	1	1	1	1	1
parents divorced	1.38*	1.62	1.35*	1.59	1.22	1.27	1.23	1.29
other non-intact family	1.55**	0.60	1.57**	0.56	1.54**	0.53	1.54**	0.52
number of siblings:	*(p = 0.672)*	*(p = 0.052)*	*(p = 0.846)*	*(p = 0.073)*	*(p = 0.962)*	*(p = 0.134)*	*(p = 0.964)*	*(p = 0.126)*
none	1.14	2.13***	1.09	2.08***	1.07	1.90**	1.07	1.92**
one	1	1	1	1	1	1	1	1
two	1.11	1.37	1.07	1.32	1.06	1.35	1.06	1.35
three or more	1.18	1.45	1.13	1.36	1.06	1.36	1.06	1.36
age at first birth (conditional on education at first birth):			*(p = 0.000)*	*(p = 0.264)*	*(p = 0.001)*	*(p = 0.293)*	*(p = 0.001)*	*(p = 0.305)*
very young			1	1	1	1	1	1
rather young			0.56***	0.74	0.63***	0.74	0.63***	0.76
medium			0.52***	0.61*	0.61***	0.61*	0.61***	0.61*
rather old			0.46***	0.98	0.52***	0.94	0.52***	0.95
oldest			0.52***	1.14	0.49***	1.11	0.49***	1.13
first-birth interval:			*(p = 0.316)*	*(p = 0.340)*				
< 8 months			1	1				
8 - 17 months			0.99	1.27				
18 - 35 months			1.24	0.91				
36 - 59 months			1.38	0.69				
60 + months			0.85	1.60				
first-birth union order:					*(p = 0.078)*	*(p = 0.737)*	*(p = 0.093)*	*(p = 0.745)*
1					1	1	1	1
2 +					1.68*	1.16	1.64*	1.16
marital status:					*(p = 0.000)*	*(p = 0.002)*	*(p = 0.000)*	*(p = 0.002)*
consensual union					2.27***	3.15*	2.25***	3.15*
transformed marriage					1	1	1	1
direct marriage					0.60***	0.52***	0.59***	0.51***
no. of children in the household:					*(p = 0.000)*	*(p = 0.000)*		
one					2.23***	2.59***		
two					1	1		
three or more					0.87	0.94		

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Table H.I (continuation)

	(1) women	(1) men	(2) women	(2) men	(3) women	(3) men	(4) women	(4) men
current age of the youngest child:							*(p = 0.000)*	*(p = 0.000)*
< 1 year							1	1
1 - 2 years							0.91	1.67
3 - 5 years							1.17	2.06
6 + years							1.61	2.65
only 1 child in the household							2.60***	4.87***
CPI change (food products only):					*(p = 0.644)*	*(p = 0.570)*	*(p = 0.636)*	*(p = 0.546)*
< 5.0%					1	1	1	1
5.0% - 9.9%					0.95	0.90	0.95	0.90
>= 10.0%					0.88	0.77	0.88	0.77
age of first child (time factor):	*(p = 0.483)*	*(p = 0.804)*	*(p = 0.583)*	*(p = 0.894)*	*(p = 0.001)*	*(p = 0.049)*	*(p = 0.003)*	*(p = 0.081)*
< 1 year	1	1	1	1	1	1	1	1
1 - 2 years	0.98	1.52	0.93	1.45	1.10	1.75	1.11	1.80
3 - 5 years	1.14	1.54	1.00	1.37	1.61**	2.44**	1.71***	2.59***
6 - 11 years	1.25	1.48	1.02	1.29	2.08***	2.95***	1.94***	2.76***
12 - 15 years	0.95	1.32	0.74	1.20	1.69*	3.07**	1.36	2.57*
	[0.362]	**[0.266]**	**[0.951]**	**[0.305]**	**[0.642]**	**[0.219]**	**[0.553]**	**[0.117]**
log likelihood	**-2767.5**	**-954.2**	**-2763.0**	**-952.6**	**-2728.7**	**-938.2**	**-2726.1**	**-936.9**
no. of independent parameters	**15**	**15**	**19**	**19**	**22**	**22**	**24**	**24**

*** significant at the 1%-level, ** at 5%, * at 10%.

Note: For each variable, risks and their significance are given relative to the reference level, indicated by 1 (no decimals). The p-value of the entire factor is given in the row containing the variable name. Absolute risk (per 1000 person-half-months) for age < 1 year of the first child is given in the last row of the time factor in boldface letter.

Table H.II: *Relative risks of dissolution of the first parental union for Hungarian women and men. Final model.*

	women	**men**
religiosity (activity level):	*(p = 0.004)*	*(p = 0.681)*
active	0.60***	0.86
not active	1	1
childhood family:	*(p = 0.119)*	*(p = 0.327)*
intact family	1	1
parents divorced	1.16	1.25
other non-intact family	1.49**	0.51
age at first birth (conditional on educational level at first birth):	*(p = 0.000)*	*(p = 0.082)*
very young	1	1
medium (incl. rather young)	0.61***	0.59**
older (i.e. rather old and oldest)	0.47***	0.80
first-birth union order:	*(p = 0.074)*	*(p = 0.560)*
1	1	1
2 +	1.69*	1.30
marital status:	*(p = 0.000)*	*(p = 0.013)*
non-marital consensual union	2.41***	2.43
transformed marriage	1	1
direct marriage	0.59***	0.56**
current age of the youngest child:	*(p = 0.000)*	*(p = 0.000)*
0 - 2 years	1	1
3 + years	1.38*	1.59
only one child in the household	2.73***	3.50***
gender-role attitude:	*(p = 0.686)*	*(p = 0.235)*
egalitarian	1	1
intermediate	1.01	1.29
traditional	0.83	0.61
attitude to parenthood:	*(p = 0.278)*	*(p = 0.639)*
child-centered	1	1
self-centered	1.25	1.33
other	0.89	1.04
current policy period:	*(p = 0.168)*	*(p = 0.490)*
Jan. 1964 - Dec. 1981	1	1
Jan. 1982 - June 1987	1.34*	1.41
July 1987 - Dec. 1993	1.22	1.33
current educational attainment:	*(p = 0.148)*	*(p = 0.317)*
compulsory education	1	1
lower level vocational school	1.13	0.65*
gymnasium	0.83	0.60*
post-gymnasium	0.82	0.64
current employment status:	*(p = 0.049)*	*(p = 0.223)*
full-time employed (incl. irregular time employed)	1	1
part-time employed (i.e. long and short part-time employed)	0.75	4.17**
own household work	1.50	
unemployed		1.81
student	0.60**	3.62
other non-employed (incl. unemployed for women only)	0.36*	1.53

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Table H.II (continuation):

	women	men
age of first child (time variable):	*(p = 0.026)*	*(p = 0.125)*
< 1 year	1	1
1 - 2 years	1.09	1.78
3 - 5 years	1.57**	2.55**
6 - 11 years	1.73***	2.54**
12 - 15 years	1.30	2.36*
	[0.523]	**[0.179]**
log likelihood	**-2716.9**	**-933.9**
no. of independent parameters	**28**	**28**

*** significant at the 1%-level, ** at 5%, * at 10%.

Note: For each variable, risks and their significance are given relative to the reference level, indicated by 1 (no decimals). The p-value of the entire factor is given in the row containing the variable name. Absolute risk (per 1000 person-half-months) for age < 1 year of the first child is given in the last row of the time factor in boldface letter.