Disruption of the first 'parental union' in Sweden and Hungary.

Focusing on policy and gender effects[i].

by Livia Sz. Oláh

Ph.D student
Demography Unit,
Stockholm University
S-106 91 Stockholm, Sweden.
Phone: +46 (8) 16 20 04,
FAX: +46 (8) 15 68 38.
e-mail: f63os5f8@students.su.se

Abstract

In this paper, we study the impact of public policies and of changing gender relations on union dissolution among families with children in Sweden and Hungary from the mid-1960s to the early 1990s. The results suggest that (i) changes in divorce legislation in either a liberal or a restrictive direction had little effect on parents' union disruption risks while the introduction of joint custody for children greatly influenced family dissolution; that (ii) gender relations in the union affect family stability; and that (iii) there are gender differences in the patterns of family dissolution risks.

1. Background

In the last decades of the 20th century, union dissolution became more and more common even among families with children in most industrialised countries. At the same time, women's employment rates have greatly increased there. This has challenged traditional gender relations which are based on the ”male breadwinner-female homemaker” model. In parallel, important policy changes have taken place, such as the introduction of no-fault divorce laws and rules of joint custody for children even after family break-up. Yet we know relatively little about the impact of such policies and that of changing gender relations on family stability. Another important question related to this is whether there is a gender-specific pattern in family dissolution behaviour. Some factors may be influential for women only or for men only, and various factors may affect relationship stability differentially for women and for men.

The purpose of this paper is to shed more light on these issues. We study the impact of public policies and of changing gender relations on union disruption among parents in Sweden and Hungary from the mid-1960s to the early 1990s. We focus on the first parental union, defined as the union in which the first child was born to a couple. Our choice of countries for the analysis has many reasons:

First, women's participation in higher education and in the labour force have reached high levels in both Sweden and Hungary as compared to other industrialised countries. Yet a large proportion of Swedish women have been part-time employed, while Hungarian women have worked full time just like men.

Second, family formation patterns have been rather different in these countries. In Hungary, childbearing was essentially constricted to marriage before the 1990s, while in Sweden the proportion of births occurring in non-marital cohabiting relationships has increased rapidly from the 1970s onwards. By the mid-1980s, the majority of Swedish couples who had their first child lived in consensual unions, and the trend has not changed since then.

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, mainly for women. Gender equality has been on the active policy agenda, but it was interpreted as limited to equal labour market participation of women and men in Hungary, while in Sweden it has aimed a more general transformation of gender roles with equal participation in paid work and in family responsibilities for women and men.

Fourth, both Sweden and Hungary have a long history of liberal divorce legislation and of rather high divorce rates also among families with children (Goode 1993). Sweden acknowledged some no-fault grounds as early as 1915. The divorce law reform of 1974 eliminated all fault grounds, and simplified the divorce procedure. A waiting period as short as six months has been required only for couples with minor children or if the couple disagreed on divorce. In Hungary, no-fault divorce was introduced in 1952. The 1974 reform permitted divorce by mutual consent. In mid-1987 divorce procedures became more complicated when a compulsory pre-divorce court hearing aiming the reconciliation of the couples was introduced. Spousal alimony is almost non-existent in both countries, and property is evenly divided between the former spouses independently of its source. A non-resident parent is obliged to pay child support if the couple has minor children.

Fifth, legal rules allow divorced or separated parents to continue to have joint custody for their children in both countries. In Sweden, the rule of joint (legal) custody was introduced in mid-1983 and has led to increased involvement in the children by both parents even after family break-up, 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 union disruption. Yet many fathers practically disappear from their children's life after the partnership ends as mothers have sole custody after the divorce in the vast majority of cases. Thus in Sweden family disruption does not necessarily mean the loss of the parent status for either women or men, while Hungarian fathers often risk losing contact with their children when the union dissolves.

Our main hypotheses in this study are the following: (i) Public policies and gender relations in the union influence family stability. Their impact should be significant even when we control for factors which have proved to affect union dissolution in previous studies. (ii) Disruption risks have a gender-specific pattern. We expect to find gender differences especially for factors that reflect the labour market attachments of the parents in the analysis.

2. Data and method

The empirical analysis in this study is based on data extracted from the Swedish Family and Working Life Survey of 1992/93, conducted by Statistics Sweden, and from the Hungarian Fertility and Family Survey of 1992/93, conducted by Statistics Hungary. (For details see Granström 1997; Kamarás 1999.) Both surveys are part of the European FFS-project and thus the data provided are suitable for cross-country comparisons.

For the purpose of this study, we have selected respondents who have one or more recorded marital or non-marital unions and at least one biological child who was born in a union. In order to avoid problems of cultural differences which are likely to affect family dissolution risks, we have excluded individuals of non-Nordic origin from the Swedish sample. From the Hungarian sample we have excluded respondents who were not yet twenty years old at interview and those whose records of partnership or childbearing histories were incomplete. As policy effects on parental union disruption may be difficult to detect, we try to make the samples as homogenous as possible. Therefore we have also excluded those 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 union ended in the same month when they had their first child, or whose first child in their first parental union died. Censoring occurs at 16 years after first birth, at the end of the union caused by the death of the respondent’s partner, or at interview, whichever event comes first. Our working samples include 1869 women and 861 men for Sweden, and 2430 women and 1070 men for Hungary.

We use the method of intensity regression to estimate the impact of various factors on the risk of family disruption in the first parental union. The analysis is based on a piecewise-constant proportional-hazards model. Exposure is measured in months (but presented in years), starting from the birth of the first child of our respondents and continued until the child turns 16 years old or until censoring for other reasons. The Windows-based software ”RocaNova”, developed at Statistics Sweden, is used for the model fitting. The results, produced as maximum-likelihood estimates of the effect parameters of the model, are presented in the form of relative risks.

3. Variables

Our main variables of interest are *current policy period* and gender relations in the first parental union. Our policy period variable represents a partitioning of calendar time. It is defined in consideration of major policy changes in the field of family dissolution. For Sweden, we distinguish between the following periods: (i) 1964-1973, when divorce was possible both on fault grounds and no-fault grounds; (ii) 1974- mid-1983, when the divorce-law reform eliminated all fault grounds and divorce procedure was shortened and simplified; and (iii) mid-1983 -1993, when joint custody for children after the parents' separation became the general rule. We distinguish between three policy periods for Hungary too, although the changes were less radical than in Sweden. No-fault divorce was admitted already in the first period, which goes from 1964 to 1973. An even more liberal divorce law denoted the second period (from 1974 to mid-1987), while in the third period (from mid-1987 to 1993) a more restrictive law made divorce procedures more complicated.

For a measure of *gender relations in the union*, we use the information in the Swedish data on whether the father took parental leave with the first child. For Hungary, we use an index of gender-role attitude based on questions regarding relationship and career.

Human capital variables are the second group of interest in this study as they can reveal gender differences in the patterns of union dissolution risk given the differences between women's and men's parenting and labour-force activities. Our variable *current educational attainment* refers to the level of schooling the respondent had up to any month, while the factor *current employment status* shows the respondent's labour-market attachments in any month after the first birth.

Further, we control for factors that have been found to greatly affect family disruption in previous studies. We divide these variables into the following groups:

1) <u>Individual characteristics</u>: 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 (not discussed in this paper) and religious activity level.

2) Maturity at family formation: The factor age at first birth grouped according to educational level at first birth shows the respondents' maturity at the time they become parents. Our variable first-birth interval reflects the maturity of respondents and their partner as a couple at the time of first birth. It is based on the interval between the start of the union and the first birth.
3) Union-specific characteristics: This group provides us with important additional information on the partnership, such as the order of the union in which the respondents' first child was born, 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, such as current national unemployment rate for Sweden, and changes in consumer price index (for food products) for Hungary.

We use a stepwise approach for the model fitting. First we include only the individual characteristics. Next we add to the model our second group of control variables which reflects maturity at family formation. Then we include the rest of the control variables and our human capital variables. The stepwise introduction of factors into the model corresponds to the sequences in which they appear in the respondents' life. This in turn determines their causal proximity to the current life situation of the respondents. This procedure also allows us to exclude those control variables that do not have a significant direct impact on the disruption of the first parental union. Thus in the last step only the important control variables are kept in the model, when we add our explanatory variables.

4. Findings

4.1 Sweden[ii]

In our discussion of findings we follow the steps of the model fitting described above. *Religious activity level* has a stronger impact on union disruption for mothers than for fathers, but the patterns are very similar. As usual, those who are religiously active have a much lower risk of family dissolution than other individuals. With the inclusion of marital status in the model, the effect of religiosity disappears. This means that religiosity has no direct impact on family stability in the secularised Swedish society, only an indirect effect which works through marital status.

Childhood family is a very important determinant of family stability in Sweden. The disruption-risk patterns are alike for women and men. Individuals whose parents divorced before the respondent's 16th birthday, have nearly twice as high risk of family dissolution as those who came from intact families. Thus we have found evidence of intergenerational transmission of divorce also for Sweden (see McLanahan and Bumpass 1988; Amato 1996 for the US; Kiernan and Cherlin 1999 for the UK; Diefenbach 1997 for Germany). The risk of disruption is also very high for individuals from other non-intact families.

Number of siblings[iii] has no influence on family disruption behaviour for either women or men 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.

Age at first birth (conditional on educational level at first birth) is important for family stability. For both mothers and fathers we find 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 regarding early start of union formation (see Morgan and Rindfuss 1985; Castro Martin and Bumpass 1989 for the US;

Berrington and Diamond 1999 for the UK; Hoem and Hoem 1992 for Sweden; Finnäs 1996 for Finland).

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

First-birth union order has a significant impact on family stability for both women and men. 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.

Marital status, which is a time-varying covariate along with the rest of control variables we present below, is another factor of great importance for family stability in Sweden. Living in a non-marital consensual union strongly increases the risk of family dissolution for both women and men. Direct marriages are the most stable relationships, while marriages transformed from cohabiting relationships have an intermediary position. This is in line with findings of previous studies (see Lillard et al. 1995 for the US; Berrington and Diamond 1999 for the UK; Bennett et al. 1988 for Sweden; Finnäs 1996 for Finland).

Our variable *current age of the youngest child* controls also for the effect of having one child only, which in previous analysis was found to be linked to much higher dissolution risk than having two or more children. For both women and men we find that the protective effect of having another child in the family works only while this child is very young. With children above age three the risk of family dissolution is nearly the same as for one-child families.

Current unemployment rate had hardly any influence on family break-up, probably because unemployment levels were rather low in Sweden during the period covered with our data. The macro-economic situation had very little effect on mothers' family dissolution intensity. For fathers, however, we see that the risk of union disruption is lower in times of slightly higher unemployment rates than otherwise.

The disruption-risk patterns are rather similar for women and men for *current educational attainment*. Those who have only compulsory education have the least stable families. There are no significant differences in family break-up risk among the other educational levels. What we see here is thus probably a selection effect for those with the least education. Alternatively, it refers to their weaker labour market situation and lower income which can create serious conflicts in the relationship, and in the long run can lead to family break-up.

Our other human-capital variable *current employment status* is also important for family stability, especially for women. As expected, we find clear gender differences in the disruption-risk patterns. Women who work or study full time have the highest risk of union dissolution among mothers in Sweden, while fathers who pursue similar labour-market strategies have the most stable families. Short-part-time employed and unemployed fathers have a very high risk for family break-up, but the disruption intensities of women in such positions are rather low. Housewives, women on parental leave and other non-employed mothers have the lowest disruption risks. Although these findings seem to support Becker's (1991) argument, the causality might work in a direction opposite to what he suggests. Perhaps women who wish to leave their partner shortly increase their labour market activities in order to reduce their economic dependence. Men who do not behave in line with the traditional male gender role as main economic providers have apparently great difficulties to maintain their unions. Alternatively, this might be a selection effect, and these men are less capable for both work and family relationships for some reason.

As for our main variables of interest, we find that gender equality in the union (here: *father took leave after first birth*) greatly affects family stability in Sweden, at least for women, but the patterns are similar for women and men. The risk of union disruption is lower if the father took parental leave with the first child than otherwise. Thus a more equal share of domestic responsibilities beside that of economic provision seems to strengthen the relationship.

Public policies (measured as *current policy period*) also influence disruption behaviour, especially for men. The risks of family dissolution are very similar in the first two policy periods. This suggests that the introduction of one of the most liberal divorce laws of the world had relatively little effect on the stability of families with children. In the third period however, we see a much higher disruption risk for both women and men. It seems that the introduction of joint custody for children after family dissolution as a main rule has facilitated union disruption for parents as it gave them better chance to remain an active parent even if the children do not live with them permanently after the family break-up.

4.2 Hungary[iv]

As we followed the same procedure for the model fitting for Hungary as for Sweden, we discuss first the risk profile for the individual characteristics. *Religious activity level* has a strong impact on family disruption for Hungarian women, but less so for men. The patterns are nevertheless similar. Those who are religiously active have much lower dissolution risk than others. In contrast to the results for Sweden, the effect of religiosity remains for Hungary even after we added to the model all our control variables. Thus the religiously active individuals seem to be a distinct group regarding family disruption behaviour as compared to other families with children in Hungary.

Childhood family seems to influence women more strongly than men, but the patterns are similar. Individuals whose parents divorced before their 16th birthday have much higher disruption risk than those coming from intact families, like in Sweden. Women who were brought up in other non-intact families have the least stable families of all as adults. However, as we included marital status in the model, the influence of childhood family becomes weaker (and not significant) and appears only for women coming from other non-intact families. This suggests 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.

Number of siblings has no impact on women's family stability, but it affects men. Hungarian fathers who have siblings have about half of the risk to dissolve their union as men without sibling. Men in the latter category 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 findings 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 in Hungary. 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 in order to preserve harmony in the family.

Age at first birth (conditional on educational level at first birth) has a strong effect on women's risks of family disruption but less influence on men. Yet the patterns are quite similar and resemble the findings for Sweden. Those who start family formation at very young ages have much higher risk of union dissolution than later starters.

First-birth interval has little impact on family stability in Hungary for both women and men, unlike 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). As parental consent may be a condition of young couples' co-residence, the partners also had more time to learn to know each other before moving in together.

First-birth union order has little influence on Hungarian men's family disruption behaviour, but it affects women. The patterns are similar for women and men, nevertheless. Those who have their first child in their first co-residential relationship have more stable unions than others, like in Sweden.

As usual, *marital status* is very influential for family stability for both women and men. The patterns resemble the findings for Sweden. Parents who live in consensual unions have a very high disruption risk. The most stable relationships are direct marriages, and marriages which started as consensual unions have an intermediary position.

Current age of the youngest child is also an important indicator of family stability in both countries. When there are at least two children in the family, the risk of family disruption increases with the age of the youngest child. Unlike in Sweden however, the dissolution risk for such families remains well below that of one-child parents in Hungary even when the children grow older.

CPI change had no influence on family dissolution 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 *current educational attainment* we find that those who have less than gymnasium education have higher disruption risk than those with more schooling for both women and men, but the effect is stronger for women. This suggests either a selection effect for those with lower education or relates to their poorer economic situation, like in Sweden.

Current employment status is, again, more important for women's family disruption risks than for men. Like in Sweden, the dissolution-risk pattern is strongly gendered. Students and other non-employed mothers have the lowest risk to disrupt their unions, while there are no significant differences among the other categories. For men, only part-time workers have significantly different dissolution risk than full-time workers, but their disruption risk is three times as high as for those who are employed full time. The findings for women might indicate that those with less resources have more stable families because of their greater economic dependence on their partners in line with Becker's (1991) argument. However, the fact that the dissolution risk for housewives is not significantly different from that of full-time workers suggests that women's increased economic independence is not really the reason of growing family instability in Hungary. The high dissolution risk of short-part-time working men is probably a selection effect.

In contrast to the Swedish findings, our main variables of interest seem to have relatively little influence on family disruption behaviour of both women and men in Hungary. As most Hungarian families follow the traditional gender division of labour in the home, we find no significant differences in dissolution risks among parents with different *gender-role attitude*s.

Changes in divorce legislation either in a more liberal or in a more restrictive direction (measured as *current policy period*) had also very little impact on family dissolution risks of parents in Hungary.

5. Discussion

In this paper we have examined the relationship between increasing union instability among families with children and changing gender relations as seen in policy changes and at the family level in Sweden and Hungary. In the context of high female employment rates and of no spousal alimony after divorce, we have found very little impact of changes in divorce legislation on parents' family dissolution behaviour in either countries. Although changes in divorce laws have relevance only for married couples, we have included also those who live in consensual unions in the analysis. In times of radically increasing divorce indices like in the 1970s in Sweden (Andersson 1997), we would have expected a bandwagon or contagion effect for cohabiting couples which is a relatively large group in Sweden, but not in Hungary. We have seen, however, that parents in any type of unions were hardly effected by divorce law reforms, unlike couples without children. Yet child custody rules seem to be important for parents' union dissolution, as the Swedish results suggest.

As for gender relations in the union, we have found for Sweden that the disruption risk is much lower for relationships where the tasks of economic provision and of active parenting are more equally shared between the partners, than for other unions. The lack of effect for this variable for Hungarian parents might be explained through the ambivalence of equally shared economic responsibilities but lack of share in domestic tasks as the general pattern which in turn results in women's double burden.

Furthermore, we have found gender differences in the patterns of family dissolution risks for both countries, as expected. Gender differences have appeared for the factors current unemployment rate, and current employment status for Sweden, and for number of siblings, and current employment status for Hungary. Apparently, labour market strategies are still gendered to some extent even in countries where the dual-earner family model is well established.

Based on our findings in this study, we conclude that (i) changes in gender relations affect family stability; and that (ii) we should study both women and men to deepen our understanding on family dissolution behaviour because the patterns of union disruption risks are gender-specific to some extent.

References

Amato, P. R. (1996), Explaining the intergenerational transmission of divorce, *Journal of Marriage and the Family*, 58, 628-640.

Andersson, G. (1997), The impact of children on divorce risks of Swedish women, *European Journal of Population*, 13, 109-145.

Becker, G. S. (1991), *A Treatise on the Family,* Harvard University Press, Cambridge, Massachusetts.

Bennett, N. G., Blanc, A. K., and Bloom, D. E. (1988), Commitment and the modern union: assessing the link between premarital cohabitation and subsequent marital stability, *American Sociological Review*, 53, 127-138.

Berrington, A. and Diamond, I. (1999), Marital dissolution among the 1958 British birth cohort: The role of cohabitation, *Population Studies*, 53, 19-38.

Castro Martín, T. and Bumpass, L. L. (1989), Recent trends in marital disruption, *Demography*, 26, 37-51.

Diefenbach, H. (1997), Intergenerationale Scheidungstransmission in Deutschland: Relevanz und Erklärungsansätze, *Zeitschrift für Rechtssoziologie*, 18, 88-105.

Finnäs, F. (1996), Separations among Finnnish women born between 1938-1967, *Yearbook of Population Research in Finland,* 33: 21-33.

Goode, W. J. (1993), *World Changes in Divorce Patterns*, Yale Univesity Press, New Haven, London.

Granström, F. (1997), *Fertility and Family Surveys in Countries of the ECE Region. Standard Country Report. Sweden*. United Nations, New York and Geneva.

Hoem, B. and Hoem, J. M. (1992), The disruption of marital and non-marital unions in contemporary Sweden, in: Trussell, J. et al. (Eds.) *Demographic Applications of Event History Analysis*, Clarendon Press, Oxford, 61-93.

Kamarás, F. (1999), *Fertility and Family Surveys in Countries of the ECE Region. Standard Country Report. Hungary*. United Nations, New York and Geneva.

Kiernan, K. E. and Cherlin, A. J. (1999), Parental divorce and partnership dissolution in adulthood: Evidence from a British cohort study, *Population Studies*, 53, 39-48.

McLanahan, S. and Bumpass, L. (1988), Intergenerational consequences of family disruption, *American Journal of Sociology*, 94, 130-152.

Lillard, L. A., Brien, M. J. and Waite, L. J. (1995), Premarital cohabitation and subsequent marital dissolution: A matter of self-selection?, *Demography*, 32, 437-457.

Morgan, S. P. and Rindfuss, R. R. (1985), Marital disruption: Structural and temporal dimensions, *American Journal of Sociology*, 90, 1055-1077.

---

[i] **Acknowledgements**
I am indebted to Eva M. Bernhardt and Jan M. Hoem for their valuable suggestions during the work with this paper. Help from András Klinger and Ferenc Kamarás regarding detailed information on the Hungarian Family Survey is also highly appreciated. I thank Statistics Sweden and Statistics Hungary for providing access to their Family Surveys, and the Demography Unit at Stockholm University for permission to work with the Swedish data set. The Swedish Council for Social Research provided economic support to the Demography Unit for data management and processing (Grant 93-0204:3A). Financial support from the Demography Unit, Stockholm University is gratefully acknowledged.
[ii] See Table S in Appendix.
[iii] The variable number of siblings is not presented in Table S.
[iv] See Table H in Appendix.

Appendix

Table S. Relative risks of dissolution of the first parental union for Swedish women and men. Models with individual characteristics, maturity at family formation, parental-union-specific characteristics, and business-cycle variations.

	(1) women	(1) men	(2) women	(2) men	(3) women	(3) men
religiosity (activity level):	*(p = 0.015)*	*(p = 0.127)*	*(p = 0.517)*	*(p = 0.718)*		
active	0.62**	0.56	0.87	0.85		
not active	1	1	1	1		
childhood family:	*(p = 0.000)*	*(p = 0.001)*	*(p = 0.000)*	*(p = 0.010)*	*(p = 0.001)*	*(p = 0.016)*
intact family	1	1	1	1	1	1
parents divorced	1.92***	2.11***	1.67***	1.83***	1.64***	1.75**
parent died	1.06	0.48	0.93	0.48	0.93	0.47
other non-intact family	1.98***	2.20**	1.76**	1.89*	1.69**	1.79
age at first birth (conditional on educational level at first birth):	*(p = 0.079)*	*(p = 0.025)*	*(p = 0.000)*	*(p = 0.001)*	*(p = 0.000)*	*(p = 0.000)*
very early / early	1	1	1	1	1	1
early	0.90	0.81				
medium	0.84	0.70	0.80*	0.73	0.79*	0.70*
late	0.68**	0.43***	0.54***	0.41***	0.55***	0.37***
very late / late	0.60***	0.45***				
first-birth interval:	*(p = 0.000)*	*(p = 0.196)*	*(p = 0.004)*	*(p = 0.198)*	*(p = 0.003)*	*(p = 0.122)*
< 8 months / < 36 months	1	1	1	1	1	1
8 - 17 months / < 36 months	0.90	1.18				
18 - 35 months / < 36 months	0.78	1.02				
36 - 59 months	0.43***	0.80	0.61***	0.92	0.60***	0.85
60 + months	0.52***	0.53	0.81	0.53*	0.79	0.50**
first-birth union order:			*(p = 0.000)*	*(p = 0.011)*	*(p = 0.001)*	*(p = 0.026)*
1			1	1	1	1
2 +			1.82***	1.84***	1.72***	1.70**
marital status:			*(p = 0.000)*	*(p = 0.000)*	*(p = 0.000)*	*(p = 0.000)*
non-marital consensual union			1.82***	2.16***	1.84***	2.04***
transformed marriages			1	1	1	1
direct marriages			0.59**	0.77	0.56***	0.72
current age of the youngest child:			*(p = 0.066)*	*(p = 0.007)*	*(p = 0.064)*	*(p = 0.004)*
< 1 year			1	1	1	1
1 - 2 years			1.59*	1.94*	1.61*	1.94*
3 - 5 years			1.83**	3.18***	1.82**	3.26***
6 + years			1.51	3.33***	1.50	3.18***
only one child in the household			2.03***	3.03***	2.04***	3.18***

	(1) women	(1) men	(2) women	(2) men	(3) women	(3) men
current unemployment rate:			*(p = 0.577)*	*(p = 0.161)*		
< 2.0 %			1	1		
2.0 % - 2.9 %			0.96	0.72*		
>= 3.0 %			1.10	0.72		
current educational attainment:			*(p = 0.002)*	*(p = 0.410)*	*(p = 0.000)*	*(p = 0.074)*
compulsory school			1	1	1	1
lower vocational/ more than compulsory			0.71***	0.79		
gymnasium/ more than compulsory			0.54***	0.81	0.67***	0.73*
post-gymnasium/ more than compulsory			0.62***	0.66		
current employment status[a]:			*(p = 0.000)*	*(p = 0.050)*	*(p = 0.000)*	*(p = 0.119)*
full-time employed			1	1	1	1
long part-time employed			0.75*	2.00	0.77*	1.85
short part-time employed			0.41***	3.30*	0.42***	3.05*
on parental leave			0.44***		0.45***	
own household work			0.60***		0.60***	
unemployed			0.86	3.90***	0.89	3.14**
student			1.53*	1.26	1.51*	1.15
other non-employed			0.49**	0.87	0.48***	0.83
father took leave after first birth:					*(p = 0.004)*	*(p = 0.512)*
yes					0.69***	0.81
no					1	1
other					1.08	0.92
current policy period:					*(p = 0.108)*	*(p = 0.013)*
Jan. 1964 - Dec. 1973					0.94	0.99
Jan. 1974 - June 1983					1	1
July 1983 - June 1993					1.30**	1.77***
age of first child (time variable):	*(p = 0.000)*	*(p = 0.002)*	*(p = 0.073)*	*(p = 0.001)*	*(p = 0.027)*	*(p = 0.000)*
< 1 year	1	1	1	1	1	1
1 - 2 years	1.19	1.74*	1.08	2.22***	1.08	2.22***
3 - 5 years	0.92	1.88**	1.03	3.36***	1.01	3.37***
6 - 11 years	0.55***	1.15	0.66*	2.06**	0.62*	1.82
12 - 15 years	0.55***	0.66	0.70	1.09	0.60	0.82
	[1.946]	**[0.879]**	**[1.213]**	**[0.204]**	**[1.264]**	**[0.157]**
log likelihood	**-2952.6**	**-1221.0**	**-2876.1**	**-1192.2**	**-2871.5**	**-1189.7**
no. of independent parameters	**17**	**17**	**32**	**30**	**31**	**29**

*** 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 beside the variable name. Absolute risk (per 1000 person-half-months) for age <1 year of first child is given in the last row for the time factor in boldface letter.

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

Table H. Relative risks of dissolution of the first parental union for Hungarian women and men. Models with individual characteristics, maturity at family formation, parental-union-specific characteristics, and business-cycle variations.

	(1) women	(1) men	(2) women	(2) men	(3) women	(3) men
religiosity (activity level):	*(p = 0.002)*	*(p = 0.290)*	*(p = 0.004)*	*(p = 0.630)*	*(p = 0.005)*	*(p = 0.525)*
active	0.58***	0.69	0.60***	0.84	0.60***	0.80
not active	1	1	1	1	1	1
childhood family:	*(p = 0.027)*	*(p = 0.188)*	*(p = 0.113)*	*(p = 0.267)*		
intact family	1	1	1	1		
parents divorced	1.34*	1.59	1.18	1.27		
other non-intact family	1.57**	0.56	1.49**	0.48		
number of siblings:	*(p = 0.713)*	*(p = 0.031)*	*(p = 0.635)*	*(p = 0.028)*	*(p = 0.473)*	*(p = 0.063)*
none	1	1	1	1	1	1
one / one +	0.91	0.48***	0.93	0.57**	0.89	0.63*
two + / one +	1.01	0.65*				
age at first birth (conditional on educational level at first birth):	*(p = 0.000)*	*(p = 0.258)*	*(p = 0.000)*	*(p = 0.127)*	*(p = 0.000)*	*(p = 0.137)*
very early	1	1	1	1	1	1
early / medium	0.56***	0.74				
medium	0.52***	0.61*	0.62***	0.65*	0.60***	0.64*
late	0.46***	0.98	0.50***	0.94	0.47***	0.88
very late / late	0.52***	1.14				
first-birth interval:	*(p = 0.321)*	*(p = 0.339)*				
< 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 = 0.568)*	*(p = 0.090)*	*(p = 0.761)*
1			1	1	1	1
2 +			1.68*	1.30	1.64*	1.15
marital status:			*(p = 0.000)*	*(p = 0.006)*	*(p = 0.000)*	*(p = 0.002)*
non-marital consensual union			2.43***	2.88*	2.48***	2.84*
transformed marriages			1	1	1	1
direct marriages			0.58***	0.54**	0.59***	0.51***
current age of the youngest child:			*(p = 0.000)*	*(p = 0.000)*	*(p = 0.000)*	*(p = 0.000)*
< 1 year / 0 - 2 years			1	1	1	1
1 - 2 years / 0 - 2 years			0.90	1.67		
3 - 5 years			1.14	2.09	1.22	1.41
6 + years			1.56	2.74	1.66**	1.76
only one child in the household			2.60***	4.84***	2.80***	3.32***

	(1) women	(1) men	(2) women	(2) men	(3) women	(3) men
CPI change (food products only):			*(p = 0.703)*	*(p = 0.524)*		
< 5.0 %			1	1		
5.0 % - 9.9 %			0.96	0.91		
>= 10.0 %			0.89	0.76		
current educational attainment:			*(p = 0.182)*	*(p = 0.294)*	*(p = 0.028)*	*(p = 0.336)*
compulsory school / lower education			1	1	1	1
lower vocational/ lower education			1.17	0.66		
gymnasium/ middle or higher education			0.86	0.60*	0.78**	0.82
post-gymnasium/ mid- or higher education			0.87	0.58		
current employment status:			*(p = 0.053)*	*(p = 0.180)*	*(p = 0.044)*	*(p = 0.246)*
full-time employed			1	1	1	1
part-time employed			0.75	4.03**	0.76	3.60*
own household work			1.48		1.49	
unemployed				2.15		1.95
student			0.35*	3.91	0.33*	3.73
other non-employed			0.61*	1.77	0.61*	1.66
gender-role attitude:					*(p = 0.711)*	*(p = 0.299)*
egalitarian					1	1
intermediate					1.01	1.26
traditional					0.84	0.64
current policy period:					*(p = 0.827)*	*(p = 0.908)*
Jan. 1964 - Dec. 1973					1	1
Jan. 1974 - June 1987					1.35	1.49
July 1987 - June 1993					1.36	1.53
age of first child (time variable):	*(p = 0.584)*	*(p = 0.895)*	*(p = 0.003)*	*(p = 0.072)*	*(p = 0.006)*	*(p = 0.110)*
< 1 year	1	1	1	1	1	1
1 - 2 years	0.93	1.45	1.11	1.85	1.10	1.77
3 - 5 years	1.00	1.37	1.69***	2.69***	1.64***	2.54***
6 - 11 years	1.02	1.29	1.93***	2.80***	1.85***	2.59***
12 - 15 years	0.74	1.20	1.36	2.57*	1.26	2.32*
	[1.039]	**[0.633]**	**[0.673]**	**[0.322]**	**[0.543]**	**[0.187]**
log likelihood	**-2763.1**	**-952.6**	**-2718.8**	**-933.1**	**-2721.4**	**-935.4**
no. of independent parameters	**18**	**18**	**27**	**27**	**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 beside the variable name. Absolute risk (per 1000 person-half-months) for age <1 year of first child is given in the last row for the time factor in boldface letter.