Max-Planck-Institut für demografische Forschung
Max Planck Institute for Demographic Research
Doberaner Strasse 114 · D-18057 Rostock · GERMANY
Tel +49 (0) 3 81 20 81 - 0; Fax +49 (0) 3 81 20 81 - 202;
http://www.demogr.mpg.de

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Marital disruption in the Czech Republic: The role of personal characteristics, individuality, and premarital cohabitation

Kryštof Zeman (zeman2@natur.cuni.cz)

This working paper has been approved for release by: Jan M. Hoem (hoem@demogr.mpg.de)
Head of the Laboratory of Contemporary European Fertility and Family Dynamics.

MARITAL DISRUPTION IN THE CZECH REPUBLIC: THE ROLE OF PERSONAL CHARACTERISTICS, INDIVIDUALITY, AND PREMARITAL COHABITATION

Kryštof Zeman

Department of Demography and Geodemography
Faculty of Science
Charles University in Prague
Albertov 6, 128 43 Praha 2
Czech Republic

zeman2@natur.cuni.cz

ABSTRACT

In this paper, we apply event history analysis to examine the possible determinants of marital disruption in the Czech Republic. We use the method of hazard regression with the baseline captured by multiple duration clocks; the event under observation is the first marital union disruption. We use the Fertility and Family Survey data from 1997, which covers the period between the 1970s and the 1990s. We focus on personal characteristics, the attributes of individuality and on conditions of partnership formation. We are particularly interested in characteristics covering the development of respondent's individuality in early life stages, like being an only child, experiencing the parents' divorce, living alone after leaving parental home and cohabiting before marriage. We control among others for the effect of educational enrolment and attainment and for the effect of children on marital stability. Through introducing unobserved heterogeneity into model, we also control for unobserved personal characteristics and examine the role of selection in the marital dissolution process.

Some of our results are similar to the results found among Western societies: Parental divorce and premarital cohabitation, as well as young age at marriage and childlessness are shifting the probability of marital breakdown towards upper levels. Moreover, we show that having no siblings and living independently in early adulthood contribute to higher marital disruption proneness of individuals.

Marital Disruption in the Czech Republic: The Role of Personal Characteristics, Individuality, and Premarital Cohabitation

Kryštof Zeman

BACKGROUND

The Czech Republic ranks among European countries with the highest incidence of divorce. Since 1950, when new family legislation was adopted in the former Czechoslovakia, divorce has become the only way to dissolve marriage, as the new law did not allow for separations (Rychtaříková, 1994). Throughout the 1960s and 1970s, Czechoslovakia enacted simpler and less restrictive divorce legislation, allowing divorce to be a quite frequent means of 'solving' marital discord. Such utilization of divorce is being generally accepted in Czech society, according to public opinion surveys. In the 1970s, more than one quarter of marriages ended in divorce, and this fraction continued to rise during the 1980s and 1990s (see Table 1). The last important change in family law as of September 1998 is unfortunately not covered by our data.

Under Socialist rule, marriage was supported and requested at a very young age but on the other hand, was taken as a product of Roman Catholic legislation and a bourgeois society and therefore legislatively 'sabotaged'. During the socialist era, the main causes of a high divorce rate were unnaturally high marriage rates at a young age and a chronic housing crisis. Both often lead the couples having to live with parents. The relative economic independence of women in relation to the full employment policy also played a role (KDGD, 2002).

Premarital pregnancy, often caused by the lack of reliable and modern contraception, frequently led to early marriage. Fertility rates among young people in the 1970s and 1980s were quite high and the second child soon followed the first. In the prevailing two-child model of fertility, the family fulfilled its main objective while the partners remained young and the potential crisis often led in divorce (KDGD, 2002).

After the fall of communism in 1989, the central-planned economy was transformed into a market economy. It included a less family-friendly labour market and national income redistribution. Some researchers argue that the stress of the labour market restructuring, economic uncertainty and deteriorating social and economic conditions will lead to higher divorce rates (Lorenz et al., 2001). However, Czech living standards were not affected to as large an extent as in some other Central-Eastern European countries and the rates of inflation and unemployment were quite moderate in the 1990s. Even the official vital statistics data do not support the divorce growth notion.

Table 1: Divorce development in the Czech Republic, vital statistics, 1970-1995

Year:	1970	1975	1980	1985	1990	1995
Number of Divorces	21516	26154	27218	30489	32055	31135
Crude Divorce Rate	2.19	2.60	2.64	2.95	3.09	3.01
Total Divorce Rate (%)	26.2	30.1	30.8	35.9	38.0	38.4

Data sources: KDGD (2002), CR POPIN

DATA AND METHODS

We use Fertility and Family Survey (FFS) data of the Czech Republic, collected in November-December 1997. The analysis is restricted to women[1]. To avoid problems with the distinction between premarital cohabitation, cohabitation as an alternative to marriage, and marriage itself[2] we are interested only in marital unions. Cohabitation is spreading among Czech society, but probably still only as premarital cohabitation. Cohabitation is being transformed into marriage at least at the time of pregnancy of the female partner.

Of the total number of 1734 respondents at ages 15-44 (cohorts 1952-1980), 1278 have experienced at least one marriage, from which 272 later divorced or separated. There was a quite low number of second marital partnerships (116) and almost no higher order marriages (3) in the sample. Our research is restricted to women's first marriages. We have left out 364 respondents who have not experienced any partnership up to the date of the interview and 92 who have only participated in cohabitation. Twenty-two respondents experienced cohabitation with another partner before the union under observation, two of them had two different cohabitations. We do not control for such rare events.

The data covers marriages contracted in the period between 1969 and 1997, and disrupted between 1974 and 1997. The summary of the data is given in Table 2.

We use the method of hazard regression with the baseline captured by multiple duration clocks. The event under observation is the first marital union disruption, censored in the case of a partner's death (10 causes), by forced living apart together (1 cause) or by the survey (994 causes). The main baseline clock is the **duration since marriage** (marriage date set to zero). For a discussion of the difference between "duration since the initiation of the union" and "duration since the initiation of the marriage" see Bennett et al. (1988, p. 131). We have chosen the latter possibility because we are interested in the stability of the marital union. We also introduce the length of premarital cohabitation as an explanatory variable. As an additional duration spline we use the calendar period. We choose between two different age/period/cohort interpretations of time[3]:

The **first** sensible approach was to introduce the **woman's current age** represented by an additional spline and to use the **generation** (cohort, year of birth) of the respondent as a

[1] Czech sample of men was selected only among partners of interviewed women.
[2] For an extensive discussion about the meaning of cohabitation see Rindfuss and Van den Heuvel, 1990.
[3] The problem of the representation of time is discussed in several studies (Bracher et al., 1993; Lutz et al., 1991; Thornton and Rodgers, 1987).

set of dummy variables to capture the effect of the cohort group. In this model, the period could not be introduced into the model, or else the model would be overdetermined.
However, during preliminary computations we found the **second** approach more useful. We used the **period** as a secondary duration spline, capturing the influence of the different socio-economic atmospheres of the 1970s, 1980s, and then the introduction of a market economy and the restoration of democracy in the 1990s. The **age** is captured by the time constant variable 'age at marriage' (divided into four subcategories of dummy variables, 15-18, 19-22, 23-26 and 27+ years). Combined with the duration of marriage, the model includes the representation of current age (the first approach of both current age and duration of marriage virtually uses the union duration twice). In the second approach, we omit the influence of generation, so we get the effect of period instead of the effect of cohort[4].
Then the model has a form:

$$\ln h_i(t) = y(t) + p(u_i+t) + \Sigma_j \alpha_j x_{ij} + \Sigma_k \beta_k w_{ik}(t)$$

where h_i is the intensity of marital disruption for individual **i**, **t** is the basic duration variable (duration of marriage), **y** is a spline that picks up the effect of marriage duration, **p** is an additional spline that picks up the effect of the period, starting at individual **i** specific **u=year at marriage-1969** (1969 is the first observed date of marriage in our sample). The last two sums represent the sets of fixed covariates **x** indexed by **j** and time varying covariates **w** indexed by **k**, with corresponding vectors of parameters α and β, respectively.
In addition, we have included unobserved heterogeneity term into the model to control for the influence of individual unobserved characteristics, as explained below. This model has a form:

$$\ln h_i(t) = y(t) + p(u_i+t) + \Sigma_j \alpha_j x_{ij} + \Sigma_k \beta_k w_{ik}(t) + V_i$$

where the item V_i picks up the unobserved heterogeneity of individual **i**, assumed to be normally distributed across individuals.
All modelling and computations were made using aML software; data preparation was made by Stata statistical software.

[4] In preliminary results the first approach showed that the youngest generations (1972-1980) have a 35 % higher divorce rate, compared to older categories. But this result had no statistical significance, probably because of the small number of young respondents and cases among them.

Table 2: Main characteristics of used data sample, women experienced at least one marital union, Czech Republic, FFS 1997

Characteristics:	TOTAL	Censored	Disrupted
1st marital unions	**1278**	**1006**	**272**
Direct marriages	944	760	184
Marriage after premarital cohabitation	334	246	88
Number of children in 1st marriage:			
0	130	93	37
1	382	257	125
2	593	502	91
3+	173	154	19
Level of education at the time of marriage:			
Education not yet finished	257	192	65
Low level	567	428	139
Middle level	416	349	67
High level	38	37	1
Partnership begun during pregnancy	555	437	118
Divorced partner	43	33	10
Woman older than partner	76	67	9
Lived alone after leaving parental home	165	121	44
Parents' union disrupted before own age of 18	160	110	50
Respondent is an only child	97	68	29
Religious	168	138	30
Year of birth:			
1952-1965	755	567	188
1966-1971	340	278	62
1972-1980	183	161	22
Age at marriage:			
15-18	289	200	89
19-22	743	587	156
23-26	193	172	21
27+	53	47	6
Cohabitation:			
Moved in together after marriage	118	89	29
Direct marriage	826	670	156
Premarital cohabitation 1-5 months	90	61	29
Premarital cohabitation 6-23 months	178	139	39
Premarital cohabitation 2 years or more	66	47	19

Note: Categories are time-constant, hence sometimes do not correspond entirely with categories in Table 3.

We focus on how personal and partnership characteristics affect marital stability. We are particularly interested in the role of premarital cohabitation and conception and in following characteristics capturing personality and individuality.

As an indicator of individuality and independence we use the **living alone** variable, constructed by comparing the date of a partnership's beginning with the date of leaving parental home (LPH). We want to test the hypothesis that more individual and self-contained respondents, recognised by leaving their parental home and living alone for some time before starting cohabitation or directly marrying, later display less stable marital behaviour.

We use the number of respondent's siblings to capture possible exceptionality of respondents with **no siblings**. According to Blake (1981), lone children are usually found to be intellectually advantaged, more mature, but somewhat less sociable than children with siblings. They come from more educated and advantaged families, but they also are more likely to come from broken families (an only child is frequently a result of marital disruption). When married, they tend to have fewer children. Our hypothesis is that in the Czech Republic lone children have a higher propensity for disruption after controlling for other factors (especially parental divorce).

We are also interested in the intactness of respondent's parental family since intergenerational transmission of marital instability is generally detected in research on marital instability (i.e. Feng et al., 1999; Wolfinger, 1999). The transmission can be caused by an unfavourable socio-economic and demographic status among broken families, like lower income and educational attainment (Feng et al., 1999), which are more likely to cause marital instability. Children of divorced partners also probably have fewer barriers to divorce and more alternatives to marriage. They display a higher propensity to cohabit before marriage than children from non-divorced families do. We use an indicator of parental union disruption to capture the effect of the union instability transmission. The **parents' divorce** is taken into account only if it occurred before the respondent's age of 18.

Finally, we include an indicator of **religiousness** (responds "Yes" to the question "Are you religious?") because higher marital stability among religious people[5] is generally recognised. According to some authors (Bennett, 1988), religious people are less likely to cohabit and more likely to marry directly. We are observing the effect in the atmosphere of high secularisation in Czech society.

Among partnership characteristics, we are particularly interested in the effects of **premarital cohabitation** and its length. Theories of marital search and marital stability usually focus on the role of incomplete information in later disruption (Becker, 1991; Oppenheimer, 1988). This could indicate the advantage of premarital cohabitation for marital stability as a source of precious and relevant information on mates. However, most empirical studies have found that marital unions preceded by cohabitation, also called transformed marriages, are generally less stable than direct marriages (Bennett et al., 1988; Hoem and Hoem, 1992; Thomson and Colella, 1992; Hoem, 1997). Some studies argue that the substantially lesser stability of transformed marriages results from selectivity (Bennett et al., 1988; Lillard et al., 1995) and vice versa: "…direct entry into marriage has become progressively selective in favour of those with high marriage cohesion" (Hoem and Hoem, 1992, p. 76). After removing the impact of self-selection using advanced statistical methods, premarital cohabitation is seen to cause a decreased risk of divorce (Lillard et al., 1995; Brüderl et al., 1999). In our research, we include the occurrence of premarital cohabitation as an important determinant of the union disruption process. We want to examine the difference between the hazard of disruption of premaritally cohabiting couples in comparison with those couples who did not live with each other before marriage. We

[5] This is true for catholic in particular, but because of lack of indicators on confession in the Czech data, we use only data on religiousness. The self-determination of religiousness may be time varying but we could not control for that in the model.

expect the results to be comparable to those from Western European societies and the USA, even when the incidence of premarital cohabitation is evidently less in the Czech Republic. Moreover, we add an indicator for the length of premarital cohabitation. Our hypothesis is that (apart from the above described difference) the marital stability will be more fragile for couples with long cohabitation, and more stable for couples cohabiting about one year before marriage. According to Thomson and Colella (1992), couples who cohabited for long periods differ more strongly from the directly married; according to Bennett et al. (1988) these couples develop individualistic modes of behaviour which are incompatible with roles in marriage. Our model recognises three categories of cohabitation duration: Less than half a year, from half a year until two years and more than two years. We have introduced the short duration category because of a possible difference from regular premarital cohabitation: "Those who cohabit for a short time may be either formally or informally engaged and do so merely for logistical reasons, having at the outset already committed themselves to marrying" (Bennett, 1988, p. 132).

Apart from direct and transformed marriages, we have also observed in the Czech Republic a category of couples moving in together only after marriage (117 respondents in our sample). This is probably due to a housing shortage or other problems of such character. We include these unions as a separate category into a model, but generally we treat them as direct marriages.

We are also interested, if the **union started during pregnancy** of respondent. We distinguish whether a first child was conceived during premarital cohabitation or before the beginning of the first union (cohabitation or marriage). A first child born up to 9 months after the start of cohabitation indicates the pre-union conception, while child born up to 9 months after marriage indicates the premarital conception. Especially the latter event was very frequent in the Czech Republic during the period under observation. The share of unwanted pregnancies is probably higher outside marriage than in marriage, hence we expect that some share of conceptions out of marriage could be unwanted but were then realised and 'legalised' by a forced ('shotgun') marriage. Such marriages should tend to have a higher propensity to end in separation than marriages concluded under normal circumstances: "It is a standard finding that women who … become pregnant before marriage … have an elevated disruption risk in marriage" (Hoem, 1997, p. 7). We test the hypothesis whether the unions that begun during the pregnancy of the respondent are less stable than other unions. The situation when cohabitation is transformed into marriage under the influence of pregnancy is quite common in modern societies, with the underlying mechanism being far different from the situation of an unwanted pregnancy. Our model uses only a variable indicating whether the partnership (direct marriage or premarital cohabitation) began during pregnancy; we are not interested in premarital conceptions that took place during cohabitation.

Moreover, we control for the woman's age at marriage, for the age difference between partners (whether female respondent was older than her partner), and for previous divorces of male partner.
We also control for the impact of education on marital stability captured by a time-varying variable combining the level of education with educational enrolment. It is recommended for studies on family formation to make a distinction between an achieved level of education and the effect of being a student (Hoem, 1986). We use such an approach in this paper on union disruptions: We distinguish between the period prior to finishing education and the period after the education is finished. The latter one is divided further into three subcategories according to the highest attained education level: Low level of education covers basic level (primary school) or no education. Middle level of education covers secondary or high school education. High level means university (college) education was completed[6]. Our expectation is that the propensity to disrupt goes down the higher the education of the respondent. The effect of educational enrolment after marriage is not clear due to its possible correlation to the age at marriage as well as to other factors.
Finally we control for the effect of the number of children, their age and pregnancy. Childless unions are generally found to be less stable than couples with children. The lowest relative risk of disruption is usually observed during the first pregnancy (pregnancy with the first child) (Hoem, 1997). The risk also remains low after the birth of the child, especially when the children are very young (Andersson, 1997). Nevertheless, as Hoem argues (1997) the risk caused by the first birth remains low only for some time (one to two years), after which (unless there is another pregnancy) the risk returns to childless values. The next pregnancy again lowers the risk. However, earlier research includes only one-child couples, which could cause the overrepresentation of those who tried to have a second child but did not succeed. As the first child grows older, the group of women remaining under observation is being selected progressively towards more disruption-prone ones (Hoem, 1997, p. 9). As Gunnar Andersson (1997, p. 117) found of Sweden: "The Swedish two-child norm and its typical pattern of family formation may explain part of the very high divorce risk observed for one-child mothers with a child at age 3–5 years or more. At, or rather before, this age of the first child, a Swedish woman is normally expected to deliver her second child. If she instead remains at parity 1 and does not proceed with her child-bearing (to the category pregnant, parity 2) this may be an indication of some kind of marital problem or of a lower commitment to family life." We expect similar results for the Czech Republic, which also has a very pronounced two-child norm.

[6] Not stated and not classifiable were included into the low level. According to FFS coding, low level means the responses 0-2 and 7 of the questions v801 and v805, middle level means 3-4 and high level 5-6.

We control for all described events adding one combined time-varying covariate on childlessness, pregnancy, the number of children (1, 2, 3+) and a simple age grouping for the youngest child (younger than one year, older).

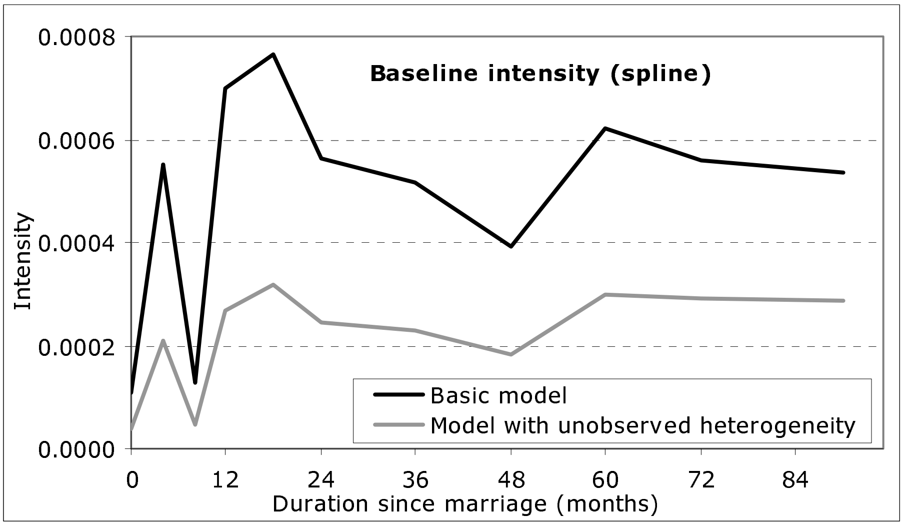
In examining the effect of children we raise the question of **unobserved heterogeneity**. According to Hoem (1997), there must be a selection of women with a low personal propensity for disruption out from the childless group into the group of parents. Then the distribution of heterogeneity factor is loaded toward higher values among childless women than among parents. Our hypothesis is, that after controlling for unobserved characteristics, the difference in marital instability between childless unions and those with children should diminish. We also expect some reduction in marital duration baseline risk because couples with higher divorce-prone unobserved characteristics tend to quit the union earlier, causing the selectivity in favour of more stable partnership in later stages of marital duration. Through the introduction of a normally distributed unobserved heterogeneity $\mathbf{V_i}$ into our model, we want to control for heterogeneity and improve our understanding of the process of union disruption.

The model does not include any variable capturing the role of employment. During the era of real socialism, full employment was guaranteed (and required) by the state, and maternal leave was one way of evading the obligation to work. We do not use the variable of employment because of its ambiguous meaning (compared to Western European societies) and because of a possible distortion and misinterpretation of the results. Although some changes in divorce risk during the introduction of a market economy after 1989 could be caused by labour market transformations, in this paper we are not interested in such issues.
Also we did not include any variables on values and beliefs (apart from religiousness), because such variables tend to change during the life course of respondents. Thus they reflect the situation at the time of surveying rather than at the time of the event.

RESULTS

Using the aML software, we have obtained the results of the final hazard regression model, first the basic one without heterogeneity and second the model that incorporated unobserved heterogeneity. The main duration baseline hazard was reached by splitting marriage duration into intervals with nodes at 4, 8, 12 and 18 months and 2, 3, 4, 5 and 6 years. We chose the intervals based on preliminary results so as to express the nature of the process in the best possible way. The baseline is shown in Figure 1. The intensity of disruption reaches its maximum 18 months after marriage, followed by a decrease up to the fourth year of marriage and subsequent stabilisation five years after marriage and later. In the beginning of the marital union, we witness a strong increase in the intensity of disruption, an exception being duration of 4-7 months. This discontinuity is probably caused by a high occurrence of premarital conceptions, and therefore quite common pregnancy and first childbirth (both known as reducing disruption-proneness) in this period.[7]

Figure 1: Baseline of first marital union disruption according to the duration since marriage



After incorporating unobserved heterogeneity into the model, the baseline risk capturing the effect of marital duration is reduced to approximately one half. This was also the main difference in the results of the models (the basic one and the one with incorporated unobserved heterogeneity). We can interpret this phenomenon in the following way: Individuals with personal, unobserved characteristics that make them more prone to divorce, disrupt their unions earlier and the share of 'normal' individuals among the population at risk thus rises.

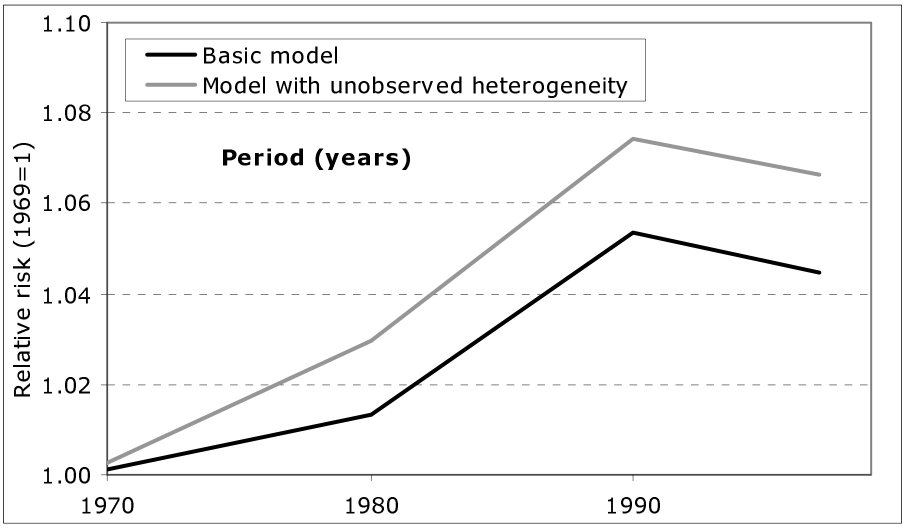
Figure 2 shows the relative risk of disruption according to the period. Nodes were placed at the years 1980 and 1990 so the trend of the 1970s and 1980s can be compared to the development after the fall of communism in 1989. As already indicated by vital statistic data, no enormous changes in divorce behaviour followed the socio-economic change in the 1990s. The last part of the baseline even indicates some reduction in disruption rates in the 1990s after controlling for other factors.

We understand the gap between the risk gained from the basic model and the risk gained from the heterogeneity model as follows: During the observed period, personal unobserved characteristics leading to higher disruption proneness of women increased (were strengthened across generations). This can be linked to increasing individualism, the spread of previously uncommon types of partnerships, the weakening of family-oriented values, a less family-friendly socio-economic background and generally social and demographic

[7] When we run the model after dropping all respondents with premarital conceptions, the peak in the 4th month fades away and does not occur.

behaviour connected to the second demographic transition (e.g. Lesthaeghe, 1995).
However, including unobserved heterogeneity into model we gain shifts in levels of both duration splines (which counterbalance each other) but not in the patterns of risks. Model estimates are quite robust against the inclusion of heterogeneity factor.

Figure 2: Relative risk of first marital union disruption according to the period



In measuring personal characteristics, we have found a strong influence of living alone (out of cohabitation or marriage) before union formation. Such individuals have a 65 % higher propensity to divorce than persons who left their parental home in order to begin a partnership. The effect is comparable to that of being an only child (having no siblings), reflecting the influence of the respondents' personality and individual development in the early stages of their life spans on subsequent marital stability. The effect of intergenerational transmission of marital instability is also quite strong in the Czech Republic: Women whose parents divorced display a 41 % higher risk of marital union disruption than respondents from intact families. The religiousness is not as important in secularised Czech Republic: The 18 % less disruption-proneness of religious people was not significant even on the 10 % significance level.

We have found a significant importance of premarital cohabitation on marital stability. Following previous research (Bennett et al., 1988; Hoem and Hoem, 1992; Thomson and Colella, 1992; Hoem, 1997 and others) we have shown that marriages preceded by premarital cohabitation display a substantially lower stability than direct marriages (and than unions which began to live together after the act of marriage, which is in our case not significant due to their low number). Moreover, we have verified the notion of Bennett et al. (1988), that couples who cohabited for long periods differ more strongly from direct marriages (124 % superfluence of those who cohabited more than 2 years; 48 % of those cohabiting one half to two years). Short cohabitations of less than six months duration surprisingly display more than twice the risk of later dissolution compared to direct marriages. The relative risks of premarital cohabitants are even higher after the introduction of unobserved heterogeneity. Our explanation is that cohabiting respondents have unobserved characteristics that make them even more prone to dissolve their union.

Table 3: Results of hazard regression, final model, marital disruption in the Czech Republic, FFS 1997

MODEL	BASIC MODEL			HETEROGENEITY MODEL		
Splines:	**Intensity**		**Sig.**	**Intensity**		**Sig.**
Duration of marriage (in months)	-9.1167		***	-10.1117		***
0-3	0.4035			0.4112		
4-7	-0.3676			-0.3694		
8-11	0.4261		*	0.4314		*
12-17	0.0155			0.0281		
18-23	-0.0508			-0.0455		
24-35	-0.0073			-0.0044		
36-47	-0.0229			-0.0196		
48-59	0.0381			0.0413		
60-71	-0.0087			-0.0019		
72+	-0.0024			-0.0008		
Period 1969-79	0.00119			0.00264		
1980-89	0.00391		*	0.00425		*
1990-97	-0.00120			-0.00103		
Variables:	**Intensity**	**EXP**	**Sig.**	**Intensity**	**EXP**	**Sig.**
Personal characteristics:						
Lived alone after leaving parental home	0.498	1.65	***	0.592	1.81	**
Parents divorced before own age of 18	0.341	1.41	**	0.396	1.49	*
The only child (no siblings)	0.493	1.64	**	0.524	1.69	*
Religious person	-0.200	0.82		-0.295	0.74	
Education						
Not yet finished	-0.223	0.80		-0.220	0.80	
Low level educated	0.236	1.27	*	0.300	1.35	*
Middle level educated	0	1		0	1	
High level educated	-0.463	0.63		-0.506	0.60	
Partnership characteristics:						
Partnership begun during pregnancy	0.084	1.09		0.099	1.10	
Partner divorced before marriage	0.224	1.25		0.229	1.26	
Woman older than partner	-0.322	0.72		-0.310	0.73	
Cohabitation						
Moved together after marriage	0.274	1.32		0.330	1.39	
Direct marriage	0	1		0	1	
Premarital cohabitation 1-5 months	0.775	2.17	***	0.933	2.54	***
Premarital cohabitation 6-23 months	0.394	1.48	**	0.499	1.65	*
Premarital cohabitation 2 years or more	0.804	2.24	***	0.910	2.48	**
Age at marriage						
15-18	0.406	1.50	***	0.483	1.62	**
19-22	0	1		0	1	
23-26	-0.632	0.53	**	-0.694	0.50	**
27+	-0.707	0.49		-0.714	0.49	
Children from current partnership						
No children	0.938	2.55	***	1.105	3.02	***
Pregnant with 1st ch. (conc. in marr.)	-1.052	0.35		-0.892	0.41	
One children 0-11 months old	0.086	1.09		0.203	1.22	
One child 12+ months old	0.624	1.87	***	0.728	2.07	***
Two children, 2nd 0-11 months old	-0.521	0.59		-0.423	0.66	
Two children, 2nd 12+ months old	0	1		0	1	
Three or more children	-0.255	0.77		-0.309	0.73	
Standard deviation of heterogeneity	*not used*			**0.9864**	-	
Log-likelihood	**-1757.2**			**-1756.8**		

Significance: '*'=10%; ''=5%; '***'=1%**

Our interest focused also on the pre-union pregnancies. However, partnerships that began during the pregnancy of the female respondent showed only an insignificant 9 % increase in disruption risk. Our hypothesis about the particular importance of this covariate was not proved.[8]

Among partnership characteristics, we controlled for whether the male partner was divorced before the marriage under observation (25 % of disruption risk in addition) and whether the woman respondent was older than her partner (28 % lesser disruption risk).

The dummy variable 'age at marriage' shows that, as expected, marriages of young women between 15-18 years of age have one-half higher probability to be disrupted compared to the group of women who married at ages 19-22. Marriages contracted at a more mature age of 23+ displayed on the contrary one-half lesser risk to dissolve. Age at marriage, one of the strongest and most consistently documented determinants of union stability[9] (Martin and Bumpass, 1989; Cherlin, 1977; Bracher et al., 1993) was also found to be important in the Czech Republic.

Controlling for the effect of children in unions, we have found a strong proneness among childless couples to divorce, displaying 255 % of the risk of the control group. As the control group we chose the common situation of couples who have two children, the second one being older than one year. As found in other studies (Hoem, 1997; Andersson, 1997), the union is most stable during pregnancy. However, our result of a 65 % reduction is not statistically significant. We ascribe this fact to the short exposure time during pregnancy and we have no doubt about the mechanism of this phenomenon. The presence of young children or a higher number of children seem to be stabilising factors for marriage as well. The result is important in relation to one older child (87 % higher risk of dissolution): As already discussed, in the two-child climate, the presence of just one child of higher age can cause or display some problems among parents that may alternatively flow into a family breakdown. After the introduction of unobserved heterogeneity, this figure rises further. However, the result for childless couples also increases, so the interpretation of this phenomenon is not clear.

The results concerning education fulfilled our expectations: Less educated people have a 27 % higher risk of marriage disruption and university educated respondents a 37 % lower risk. Women still enrolled in education display 20 % reduction.

[8] No effect of premarital conception was found by Bracher et al. (1993) in Australia, where the effect disappeared after controlling for the age of marriage. The important factor thereby was not the bride's pregnancy status itself, but whether she married young or not.

[9] Some researchers use the age at the start of partnership rather than marriage age for studying the effect of premarital cohabitation on subsequent marital stability (Brüderl et al., 1999).

CONCLUSIONS

Following previous research on the topic in Western Europe and USA, this paper tries to find important explanations of marital instability in the Czech Republic using the Fertility and Family Survey data from 1997. We have used sophisticated methods of event history analysis to analyse the process of first marital union disruption. Our results are similar to the results found among Western societies in the following cases:

- strong intergenerational transmission of marital instability;
- significant and intense effect of premarital cohabitation on subsequent marital instability;
- higher divorce-proneness of marriages concluded in premature ages;
- high divorce risks of childless couples and least risk of divorce during pregnancy;
- positive influence of children on marital stability, with the exception of one older child.

In the Czech area we have found also as important:

- higher marital instability of lone children;
- high marital instability of persons living independently before union formation;
- weak importance of religiousness;
- no impact of pregnancy during the start of the union.

Our main conclusion is that events and characteristics that contribute to the development of individualistic mode of behaviour also contribute to higher instability of marital unions, especially parental divorce, no siblings, independent living and premarital cohabitation.

The main change in the model estimates after the implementation of unobserved heterogeneity was the reduction of baseline intensity and an even higher disruption risk of premaritally cohabiting persons. However, the patterns of risks remained unchanged. For better utilisation of statistical models that include unobserved heterogeneity we shall use repeated events data on second and higher order marital unions. We also want to focus on joint modelling of union stability with earlier life-course processes, represented by leaving parental home and premarital cohabitation, and on simultaneous modelling of marital childbearing and marital disruption. Only then we will be able to identify the role of selection and self-selection among more disruption-prone individuals during these processes, to pick up the potential endogeneity of distinct processes, and to find the consequences on marital stability.

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