EUI WORKING PAPERS

SPS No. 2006/04

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EUI Working Paper **SPS** No. 2006/04

ISSN 1725-6755

Printed in Italy
European University Institute
Badia Fiesolana
I – 50016 San Domenico di Fiesole (FI)
Italy

http://www.iue.it/
http://cadmus.iue.it/dspace/index.jsp

Abstract

In this paper we explore the changes in the relationship between female educational attainment and the risk of union disruption in seventeen countries: Austria, Estonia, Finland, Flanders (Belgium), France, West-Germany, Greece, Hungary, Italy, Latvia, Lithuania, Norway, Poland, Spain, Sweden, Switzerland, and the United States. We start from the hypothesis presented by William J. Goode (1962; 1979; 1993), stating that in the Western countries, the initially positive relationship between social class and divorce would gradually change during the modernization process and waning of barriers to divorce, so that eventually there would either be no relationship between the two, or that the lower classes would divorce more. We expand the examination to all unions – not just marriages – due to the increasing importance of non-marital cohabitation in many of our countries. We run separate models for all unions. We first analyse the data within each of the seventeen countries with discrete-time event-history analyses. We find important variation across countries in the relationship between education and union disruption, and find that the relationship has become more negative in five countries. Second, we use multi-level models for event-histories in discrete time to examine the macro-level correlates of this variation. We report that a higher level of employment in service sector and higher percentages of economically active women are associated with a more negative relationship between education and union disruption. Overall, we find support – although not unanimous – for Goode's hypothesis, and conclude that the waning of social, and economic barriers to union disruption increases the risk of union disruption relatively more among the less educated.

Previous versions of the paper have been presented at the 99th Annual Conference of the American Sociological Association in San Francisco, 14th-17th August 2004, Quantitative Methods in the Social Sciences workshop on Longitudinal Data Analysis in Southampton, 18th-27th August 2004, the 3rd Conference of the European Research Network on Divorce, December 2nd - 4th 2004, in Cologne, 37th World Congress of the International Institute of Sociology in Stockholm, 5th-9th July 2005, and the Notestein Seminar at Princeton University, 27th September 2005. We thank the participants of these seminars and conferences, and especially Gunnar Andersson, Hans-Peter Blossfeld, Josef Brüderl, Tak Wing Chan, Paul de Graaf, Matthijs Kalmijn, Torkild Hovde Lyngstad, Anne-Rigt Poortman and Michael Wagner for valuable comments. All remaining errors are ours. A companion paper focusing solely on divorce has been submitted for publication in the *European Sociological Review*, Special Issue on the Causes and Consequences of Divorce in Europe.
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Introduction

Increasing marital instability has been among the most visible features of family change in the Western countries. Even though the general trend has been similar across the industrialized world, this development has, however, taken different timings, levels, and paths in different societies, thus raising questions of the underlying societal factors responsible for these differences (*e.g.*, Cherlin, 1981; White, 1990; Castles and Flood, 1993; Lesthaeghe, 1995; Ono, 1999). In addition to societal factors, students of divorce have analysed the impacts of various individual and family related factors on the risk of divorce (*e.g.*, Goode, 1962; Bumpass and Sweet, 1972; Becker, 1981; Blossfeld *et al*., 1995; Dronkers, 2002). Results from these studies suggest important differences in divorce risks across social groups. These differentials reflect social inequalities not only in the opportunities for the dissolution of unhappy relationships, but in the possibilities for stable and satisfactory relationships as well. They also reflect social inequalities in the (mainly negative) consequences of divorce. The overrepresentation of divorce – and other "unconventional" family behaviour – in the lower social groups has raised concerns of the accumulation of disadvantages over different life spheres (McLanahan, 2004).

Results pointing to social differentiation in family forms and family behaviour do not mean that these differences are necessarily stable across time and space. For example, the current American experience of the low class character of single parenthood and early births does not characterize all earlier periods or other countries (Ellwood and Jencks, 2004; McLanahan, 2004). The same might apply to union dissolution. Studies on divorce trends have not for the most part considered the possibility of different trends across social groups (see, however, Martin and Bumpass, 1989; Hoem, 1997; Teachman, 2002; Chan and Halpin, 2005). Despite the widely applied implicit assumption of stability in divorce risk factors, there are theoretical reasons to expect different developments for social groups (Teachman, 2002: 332).

William J. Goode was probably the first to argue for a link between societal factors and the social composition of divorce. In a series of papers (Goode, 1962; 1979; 1993), he suggested that the relationship between the social composition of divorce and the level of modernization is inverse. He expected that the once positive relationship between social status and divorce – characteristic of early stages of modernization with high legal, social, and economic barriers to divorce – will gradually fade away by the lifting of these barriers. In the "mature" stages of modernization with low legal, social, and economic costs to divorce, divorce risks may finally be higher in the lower classes, which generally have higher marital strain. Following this line of thought, we can also expect cross-country differences in the social structure of divorce, which can be linked to particular variables reflecting the social environment of marriage and divorce.

In this paper, we examine the effects of female education on union disruption over time in seventeen countries (Austria, Estonia, Finland, Flanders (Belgium), France, West-Germany, Greece, Hungary, Italy, Latvia, Lithuania, Norway, Poland, Spain, Sweden, Switzerland, USA). We use education as the indicator of social status, due its central role in modern stratification systems. Because of the increasingly important role of unmarried cohabitation as a "trial marriage" and even an alternative to marriage, we examine both marriages and all unions together, regardless of their marital status.

Starting from Goode's hypothesis, we ask 1) whether there are cross-country differences in the educational gradient of union dissolution, 2) whether the effect of education on union dissolution has become more negative across time, and whether this is a consistent pattern across countries, 3) whether we arrive in significantly different results when we examine all first unions instead of first marriages only, and 4) whether the differences between countries and across time can be linked to macro-level variables reflecting the legal and socio-economic environment of family life? We focus our attention on first unions and first marriages, because of the well-known differences in the marital processes of higher order marriages/unions (Martin and Bumpass, 1989). With data from the Fertility and Family Surveys (FFS), we perform discrete-time event history analyses in each country, and then continue with multi-level discrete-time event history analyses to test for explanations for the patterns found.

Trends in divorce and union dissolution

The general story of divorce in post-war industrialized countries has been an often dramatic increase. Figure 1 presents trends in total divorce rates (share of marriages predicted to end up in a divorce) in selected European countries. The figure shows a generally upward trend across the countries, even though cross-national differences remain remarkable. In 1995, half of all Swedish marriages were expected to dissolve, whereas less than ten per cent of Italian marriages had a similar prediction. But even in Italy, the main trend has been an increase in divorce risks since the mid-1970s. Similar trends are found across Europe and the United States, with a main exception of Latvia, where divorce risks have notably decreased after the peak in the early 1980s (OECD, 2002; Council of Europe, 2003). A considerable amount of research has been devoted to explaining these trends and the country differences in them. Explanations have focused on cultural change, changing gender roles, and “modernization” as a more all-encompassing development (Goode, 1970; Becker, Landes and Michael, 1977; White, 1990; Lesthaeghe, 1995).

FIGURES 1 & 2

The increases in non-marital cohabitation are another notable change in post-war family behaviour. In all of the countries included in this study, non-marital cohabitation increased in the younger cohorts, and in the most liberal countries, first unions starting as marriages now present a considerable minority (Figure 2; Andersson and Philipov, 2002). Figure 2 presents the trend in the percentage of first unions starting as consensual unions in selected countries, Table 1 shows the development all seventeen countries. From these results we can first of all see that cohabitation has become increasingly popular in each country. We can also detect three separate country clusters, with Estonia and Sweden being the fore-runners, most of the Central-European countries, Finland, Norway and the US catching up with some delay, and the more traditional catholic/orthodox countries starting later and having relatively low rates of pre-marital cohabitation still in the 1990s.

TABLE 1

The nature of cohabitation has also changed, as cohabitation spells have increased and non-marital cohabitation has in some countries, such as Sweden, even begun to challenge marriage as a form of family life and longer-term commitment. Consensual unions have also become more widespread across the social structure (e.g., Andersson and Philipov, 2002; Villeneuve-Gokalp, 1991; Finnäs, 1995; Bumpass and Lu, 2000; Murphy, 2000). However, despite the changing nature of cohabitation, non-marital unions have a higher risk of dissolution than marriages, and a focus on marriages only does not tell the whole story of the formation and dissolution of intimate cohabitational relationships in modern times (e.g., Finnäs, 1996; Raley and Bumpass, 2003).

Female education and union disruption: theoretical approaches

The best-known hypothesis of the effects of female education on divorce comes from Gary Becker's economic theory of the family (e.g., Becker, 1981). According to Becker, women's educational attainment is positively related to their labour market opportunities – and thus chances of supporting themselves (and their children) regardless of the provision of the husband – while it is negatively related to (traditional) role specialization and mutual interdependence within the family. Since the benefits of marriage and cohabitation mainly stem from specialization (according to traditional roles) and interdependence, higher female educational attainment thus reduces the gains from marriage, and increases the risk of divorce. Other accounts have predicted a positive effect of female education on divorce by pointing the more liberal values these women are likely to hold (Levinger, 1976), and their better resources in handling the social, legal and economic aspects of the divorce process (Blossfeld *et al*., 1995).

Some theories have, however, led to the opposite predictions of the effect of education on union dissolution. It has been argued, for instance, that education improves resources – such as social, cultural, economic and cognitive skills – that increase the stability of relationships, either by successful partner matching, or by enhancing communication skills and other factors that make a relationship work (Amato, 1996; Ono, 1998; Hoem, 1997; Dronkers, 2002). Others have emphasised the economic returns of higher education and their positive impact on marital life. Thus, in line with the original Goode hypothesis, we can assume that those in lower social strata have more marital strain due to greater socio-economic hardship, and therefore a higher likelihood of marital disruption (also Hoem, 1997; Oppenheimer, 1997; Jalovaara, 2003). Women with low education may also feel less tied by "middle-class" family norms (Amato, 1996).

The theories discussed above predict different effects of female education on the risk of union disruption. However, they all have in common an assumption that unions are maintained as long as the well-being of the partner(s) exceeds that of dissolving the union (Teachman, 2002: 331-2). Their main difference is in the mechanisms (economic, social, cultural, cognitive) emphasised, and less in the direction in the effect of these mechanisms. Thus, there seems to be no *a priori* reason to rule out the possibility that these mechanisms operate simultaneously, even though in the different direction. Thus the resulting effect of female education on union dissolution depends on the net effect of these different mechanisms.

Are there reasons to expect that the educational gradient of union dissolution is different if we look at marriages only compared to all unions? Despite the increase in and the changing forms of cohabitation, the behaviour of married partners differs from that of non-married ones in many respects, even in countries like Sweden (e.g., Henz and Sundström, 2001). Marriages also have a different legal status than consensual unions, and the higher rates of union dissolution in non-marital unions further suggest that marriages provide more stability than consensual unions. Therefore, in line with the arguments presented below, one could expect that the effects of education on divorce are different – possibly more positive – if we focus on marriages instead of all unions. On the other hand, several commentators have pointed to possible selection effects. For example, Hoem (1997: 26) suggested that not only will entry into marriage be more selective on commitment to the marital institution, but also that as educational attainment levels rose, those with low education may be negatively selected also on their chances of entering marriages and maintaining them. This might suggest that, at least in some countries, those less educated women who do enter marriage are more committed to marriage than their higher educated sisters, and thus, the educational gradient of divorce would be more positive than the educational gradient in union disruption regardless of marital status. American results by Raley and Bumpass (2003) seem to support this hypothesis.

Cross-country and period differences in the effects of female education?

We can discuss the expected cross-country and period differences in the effects of education from a cost-benefits point of view outlined above (Teachman, 2002). Teachman argued that the effects of the disruption risk factors might change because of the changes in the social environment is not perceived by all couples in the same way, and all couples do not thus react similarly to these changes. In general, when divorce is costly (in social, economic, and legal terms), any traits positively related to disruption risks will be suppressed, whereas when these costs are lower, such traits can "flourish". He also argued that these changes might not affect all unions in the first place.

From such a viewpoint, we can speculate the effects of three different environmental factors, which affect not only the costs of union disruption, but also the benefits of staying in a union. First, we can point to the legal environment of union disruption. The effect of the liberalization in divorce laws on divorce rates has been a topic of great interest, especially in the United States. Recent results suggest that divorce legislation does have a positive effect on divorce rates, at least in the short run (Friedberg, 1998; Wolfers, 2003). In line with the discussion above, we can expect that the strictness of divorce laws does not have equal effects on all unions. Goode (1970 [1963]: 85-86) made the hypothesis that strict divorce laws mainly suppress the divorce chances of the lower classes, while the upper classes are more resourceful to find their ways around. Goode continued to argue that with more liberal laws, "the normal difficulties of lower-class family life were permitted an expression in divorce" (ibid.).

We can expect that social norms and conventions surrounding family life can produce similar differences. First of all, strict social norms against union disruption may require extra resources (such as high education) to overcome them, whereas loosened norms of union disruption reduce the importance of such resources. Second, if divorce and union disruption is relatively rare in a society, such behaviour is more innovative, thus, again, requiring more resources. Later, through social learning, such behaviour can diffuse to the wider population. Similar to other forms of demographic behaviour, union disruption patterns can also "trickle down" from the higher social groups to the lower ones (Chan and Halpin, 2004).

The economic environment, including the welfare state, has an apparent possibility of affecting the relative costs and benefits gained from union disruption. First, good female labour market possibilities decrease women's dependency on their husbands. When female participation in the labour market is acceptable and widespread, even the women with lower human capital have better chances for economic independence. The same can be said of the welfare state: when welfare state benefits and services are extensive and generous, women, especially with small children, can provide themselves independently, either by working or through benefits (cf. Orloff, 1993). Here again, more extensive and generous welfare states might be assumed to improve the disruption chances of the lower groups more. On the other hand, welfare states might also reduce the economic strain of the lower classes (cf. Hoem, 1997; Oppenheimer, 1997; Jalovaara, 2003), thus reversing the relationship.

In line with the tentative argumentation above, we can expect that the macro-level factors affect marriages differently compared to all unions.

Previous results and hypotheses

What is the empirical evidence regarding the effects of education on union disruption, in different countries and at different times? In general, the results are mixed. American, Nordic, and British studies generally find a negative effect (e.g., Bumpass and Sweet, 1972; Berrington and Diamond, 1999; South, 2000; Hoem, 1997; Jalovaara, 2003; Lyngstad, 2004). Similar results were already found in the early American and British studies on the class gradient of divorce (Goode, 1951; Gibson, 1974). Other studies have, however, found different results. German and Dutch research has generally reported a positive effect (e.g., Diekmann and Klein, 1991; Kalmijn *et al*., 2004). Support for cross-national differences were further given by Blossfeld and colleagues (1995), who found the positive effects to be the strongest in Italy, weaker in Germany, and the weakest in Sweden. Comparative differences were also found in the cross-national case studies in Chester (1977).

Few studies have explicitly tested whether there has been a change in the effect of female education on union disruption. Again, the results from the few studies, which have considered this possibility, have given conflicting results. Martin and Bumpass (1989) found that the effect had generally become more negative in the younger American marriage cohorts, whereas Teachman (2002) did not find such instabilities in the effect of education. However, again, the findings by Raley and Bumpass (2003) do seem to suggest that union disruption has increased in the lower groups, but stayed stable in the upper ones. They also reported that the change has been less pronounced among marriages than all unions. For the UK, Chan and Halpin (2005) found that the relationship between female education and divorce has changed from a positive to a negative one, and similar changes were also found in Sweden by Hoem (1997). In the Netherlands Dronkers (2002) reports a change in the relation between intelligence and union disruption during the second half of the 20th century.

These results send a conflicting message. On the one hand, the cross-country (and some cross-cohort) differences in the effects of education suggest that societal level factors have an effect, as suggested by the theoretical discussion above. On the other, one might expect that the within-country trends would be more consistent, knowing the profound changes in all post-war Western societies, and in the family institution in particular.

Based on the theoretical discussion and previous research, we formulate five hypotheses.

Hypothesis 1: The effect of education on the risk of union disruption varies across countries.

Hypothesis 2: The effect of education on union disruption is more positive if we focus only on marriages.

Hypothesis 3: The effect of education on the risk of union disruption becomes more negative across time, and this is a consistent pattern across countries.

Hypothesis 4: The change to a more negative educational gradient is less pronounced in marriages than in all unions.

Hypothesis 5: More liberal divorce legislation and normative environment towards family issues, more generous social welfare systems, more prosperity and "modernity", and more female labour market opportunities all change the effect of education on union disruption more negative.

Data and methods

Micro-data

In the subsequent analyses we use data for our seventeen countries from the Fertility and Family Surveys (FFS), collected by the Population Activities Unit of the United Nations Economic Commission for Europe (see Andersson and Philipov, 2002). The FFS is a retrospective survey, which includes information on the fertility, family, education, and occupational histories of the interviewed. The data were collected between 1989 and 1999, in different years in different countries. For the analyses we selected the first partnerships (whether cohabiting or married) of women who had entered such relationships[1]. To be suitable for discrete-time event-history analysis, we re-organized the data into person-year form[2] (Yamaguchi, 1991). After considerable data cleaning, we ended up with a sample of 52 150 women, 604 178 person-years, and 12 880 union disruptions in 17 countries (see table 2). Our dependent variable is union disruption, which was coded 1 if the union dissolved during a particular year, and 0 otherwise[3].

TABLES 2 & 3

The independent variable of most interest to us is educational attainment. We could not find complete educational histories for many of the countries, and thus had to resort to using the educational attainment at the time of the interview. Although the measure is subject to reverse causation and the distributions in some countries look unfamiliar, we chose this measure as the most straightforward of the ones available. We ruled out the use of years of education due to serious problems in this measure in some countries, and as mentioned, many countries lacked educational histories, in particular reliable ones. The education variable was coded into three categories, according to the ISCED scheme: low (0-2), middle (3), and high (4-6). Table 3 gives the percentages of the latter two categories, both total and per country, based on year-person units.

Our other explanatory variables are duration, duration squared, year of the start of the first union, age at start of the first union, a dummy indicating parental divorce, a dummy indicating a birth before the union and a dummy indicating whether the woman was married to her partner at a specific point in time (see Table 3). Of these variables, parental divorce is the only proper control variable, whereas the last three are better seen as intervening variables. The Norwegian data missed the parental divorce variable and for that reason it will only be included in the within-country analyses and without a control for parental divorce.

In the analyses, we use union cohort (the year the couple started living together) to measure the changing social environment of union disruption. However, because we include duration (linear and quadratic terms) in the models, and since cohort plus duration equals period, the effects of union cohorts can be interpreted either as cohort or as period effects (see Allison, 1995: 142-3; Teachman, 2002). Following Teachman (2002), we interpret the union cohort variables as capturing the period effect of a change in the social context of family life[4]. This

[1] The processes affecting the survival and dissolution of higher order unions are notoriously different from those of first unions (e.g. Martin and Bumpass, 1989). In an earlier version of the paper we experimented with different union types without major differences in the conclusions.

[2] The FFS would have allowed us to build a person-month file as well. Since handling and analysing the data was computationally burdensome already as it is, we did not want to change to a person-month file, especially as preliminary analyses suggested no major differences in the results.

[3] Because in some countries (Italy, Poland, Greece, Spain) there were only a few union dissolutions in the early periods, we did not censor durations of ten, fifteen, or twenty years. However, the models with a censoring on the 15th year of the union (if still intact) gave very similar results (not shown). We coded death and "forced living apart together" also as censored. A competing-risks analysis with these categories did not change our results.

[4] Following Thornton and Rodgers (1987) and Teachman (2002), we also tested whether risk of union disruption at different durations varied across the partnership cohorts. In a model with the main and squared effect of duration, and an interaction between duration and partnership cohort (not shown), we did not find stability of dissolution risk at different durations across cohorts. We thus conclude that, if we disregard the assumption of a

interpretation is also supported by the large literature pointing to the importance of period effects over cohort effects, whether of birth cohorts or marriage cohorts (Thornton and Rodgers, 1987; Heaton, 1991; Lutz *et al.*, 1991; however, Ono 1999)[5].

Macro-variables
In order to analyze the effects of macro-level factors, we collected data on divorce legislation, social policies, values, family practices, and the "general level of modernization" (see Appendix). The time-dependent nature of the variables varied. For some measures (for instance divorce legislation, extra-marital births, female economic activity, and urbanization levels) we were able to find data for rather long time periods. For others (such as value measures), we had to restrict ourselves to a few time-points, or even only one. In general, with the exception of divorce legislation, a decade was the basic unit of time used. For some societies, in particular the former Eastern Bloc, the accuracy of some data might, of course, be questioned. Their averages and standard deviations are given in table 3, based on the year-person categories, both total and per country. If the standard deviation of a macro-variable of a country is zero that variable is not time dependent, but varies between the countries only. If the standard deviation of a macro-variable of a country is not zero that variable is time dependent and varies both between the countries and in time. In the case of the divorce-laws characteristics the lack of variation in time reflects the stability of the divorce law of a country and not a lack of data. In the case of the opinion on 'whether divorce is justifiable' the lack of variation in time reflects a lack of more than one wave to measure this value.

Divorce legislation. The strictness of divorce legislation is measured with a single time-dependent variable. During the period of study, divorce legislation varied considerably from prohibition of divorce to unilateral no-fault divorce. We use a three-fold categorisation as our divorce law measure (cf. Glendon, 1987; Castles and Flood, 1993):

(1) Divorce not permitted (1a), or permitted on the grounds of fault or other major disruption of marital life (1b). Institutionalisation of marriage remains the leading principle, and the divorce process (if permitted) hard and lengthy.
(2) Divorce permitted, possibly alongside (1b), on mutual consent of the spouses, prolonged separation, other measure of factual breakdown of marriage, or other less restrictive legislation. Shows more understanding for the will of the spouses.
(3) No (or very minor) judicial ground to deny divorce: unilateral non-fault divorce granted on the basis of the will of either spouse with very short waiting or "reconsideration" times.

We joined (1a) and (1b) into a common category, because the small number of events in (1a) did not permit sustainable analysis. We use this category also as the reference category.

Data on the divorce laws of each country were collected from Boschan (1972), Chester (1977), Chloros (1978), Lobodzinska (1982), Moskoff (1983), Glendon (1987; 1989), Goode (1993), Nakonezny et al. (1995), Friedberg (1998), and Hamilton and Perry (2002). In some cases, classifying a country into only of the categories was not very straightforward. The trickiest case was the United States, where individual states have their own divorce laws (Glendon,

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strong interaction effect between partnership cohort and duration, our data support the hypothesis of period effects instead of cohort effects. We included the linear and quadratic effects of duration instead of the more widely used strategy of comparing risks at two different durations because in some countries the numbers of dissolutions at particular intervals were not big enough to allow sustained analyses. Our conclusion for the importance of period versus cohort is also supported by the similarity of our results, when the models were ran with a censoring of the longest intact unions (>15 years).

[5] Using a direct period measure would have been inconvenient, since in Italy, Spain, Greece, and Poland the number of union disruptions in the early periods was too small to permit sustainable analyses.

1987, 1989; Nakonezny et al., 1995; Friedberg, 1998). Since we could not distinguish between different states, we treated the US as a single case. We used the year 1970 (when California as the first state enacted a no-fault legislation) as a breakpoint between categories 1 and 2, and the year 1985 (when South Dakota was the last state to enact no-fault divorce) as a breakpoint in a move from category 2 to category 3. This solution admittedly provides only an approximation, but since "divorce tourism" between states was possible (Castles and Flood, 1993), and since the law in many cases was a dead letter with pressures from neighbouring states, we regard this as a second-best option.

Social policies. Social policy generosity is measured by two variables, social expenditure per GDP and family cash benefits per GDP. The former was used to capture general welfare state generosity and social protection (and the extent to which one can gain a living independently of the labour market or the family), while the latter was chosen to reflect more targeted social expenditure. For most countries, we were able to construct good time-series of the developments of these social policy measures with data from the ILO (1967; 1988) and the OECD (1997). The exception was Poland, for which we found data for only one point in time.

Values. Values were measured with two variables, the percentage of denounced atheists to measure (non)religiosity (Barrett *et al*., 2001) and the national mean of a ten-point scale of the question of whether one finds divorce justifiable or not, from the World Values Study (1981; 1990; 1995) and the European Values Study (for Greece) (1999). For the former variable, the first data were found from 1970 onwards (projected back for earlier periods), for the latter, mainly from 1981 only.

Family practices. We measure family practices with a single time-dependent variable of "unconventional family types", which is a sum measure of the percentage of extra-marital births, the share of divorces per 100 marriages, and the percentage of 25-year olds who have ever lived in a consensual union (OECD 2002; Council of Europe 2000; FFS standard country tables http://www.unece.org/ead/pau/ffs/f_h_151b.htm). These variables were strongly correlated (0.7-0.8), and therefore they were combined to proxy the social costs of divorce and the "conventionality" of the family institution.

Modernity and the labour market. We use three variables as indicators of labour market conditions and "modernity": the degree of urbanization, the percentage of employment in the service sector, and the percentage of economically active women of all working aged women (World Bank World Development Indicators; ILO Labour statistics http://laborsta.ilo.org/; Scharpf and Schmidt, 2000: 349; United Nations Common Database http://unstats.un.org/unsd).

Models

Within-country discrete-time event history models

In our first analyses, we model the effects of female education on the risk of union disruption separately in each country with discrete-time event history analysis techniques (Yamaguchi, 1991). Event-history analysis regresses the conditional probability of experiencing an event at time *t* (union disruption), conditional on that it has not happened before, on selected covariates (discussed above).

Multi-level models

To explain the patterns found from the within-country event history analyses, we continue our analyses by using multi-level discrete-time event history models to test for the effects of the macro-variables. Replacing countries with variables has been generally regarded as a valuable strategy in comparative research (e.g., Przeworski and Teune, 1970). Here, we have an additional dimension, historical time, and therefore we replace country at a specific period (measured by union cohort in the within-country models) with direct measures of divorce legislation, social policies, values, demographical practices, and the level of the economy. To analyze these data, the data file is restructured into a two-level data structure: countries and duration of the first union (organized in person-years). This representation allows us to use models for binary response variables in a multilevel context (Hox, 2002).

The model can be written as:

$$\operatorname{logit}\left(h_{ij}\left(t\right)\right)=\alpha\left(t\right)+\beta x_{ij}\left(t\right)+\beta z_{j}+u_{0j}\,. \tag{1}$$

The hazard function $h(t)$ is the probability of the event occurring in interval t conditional upon no earlier occurrence. In our case, the time variable t is the length of the union at time t. In this equation $\alpha(t)$ is the baseline hazard at union-year t, x_{ij} represents the micro level predictors and z_j represents the country predictors. The u_{0j} are the country level residual errors; since this is a logit model for binary outcomes there is no women level error term (cf. Hox, 2002). The regression coefficient α for the effect of union-duration may or may not vary across countries; in our case there was between country variation which did not disappear when all available individual predictors were included in the model. The regression coefficients β for the women level predictors may or may not vary across countries. The model was estimated using MLwiN (Rasbash et al., 2000).

Results from within-country event-history analyses

Gross and net effects of education

Table 4 presents results from the discrete-time event history analyses by country. Models A in the first columns give the "gross" effect of educational attainment, when year of start of union, parental divorce, and the linear and quadratic terms of duration are the only variables adjusted for. Models B in the second columns show the effect of education, net of the effect of the three important confounding variables (age at start of union, child before union, and the time-dependent term for marriage). The estimates for the other explanatory variables are mostly as expected. Women who have divorced parents, who start their first union at a young age and who have a child before the union have higher risks of union disruption than other women. The duration terms show a familiar shape, and in most countries there is a trend towards more union disruptions in the later cohorts.

TABLE 4

The estimates for the education dummies show cross-country variation in both of the models. The gross effect of education is consistently negative Austria, Lithuania and the USA. In Latvia high education decreases the risk of union dissolution, but the effect is only weakly significant. Women with high or middle-level educated women have higher dissolution risks in France, Germany, Italy and Spain. In Greece and Switzerland only the high educated have more union disruptions and in Poland the positive effect is limited to those with middle education. In all other cases the differences are not statistically significant.

Inclusion of the confounding variables in Models B does not change the result of the cross-national differences, but does make the estimates of education more positive in practically all cases. Germany, Greece, Spain and Switzerland are exceptions in this regard. In Latvia and

Lithuania the originally negative effect of high education, and in the US the negative effect of middle education, become non-significant in Model B. In Germany and Switzerland the originally positive estimates become non-significant, and in the Nordic countries the non-significant effects of high education in Model A become positive and significant. With the exceptions of Germany, Greece, Spain and Switzerland, the change in the effects of education between the two models means that on the aggregate level, women with lower education are more likely to engage in behaviours that increase the risk of union disruption. On the basis of these models, we can conclude that there are cross-national differences in the effects of education on the risk of union disruption, which cannot be explained by cross-national differences in dissolution risk increasing behaviours by educational level. The results in Table 4 thus support out first hypothesis.

Does a focus on marriages make a difference?
In Table 5 we analysed all unions. What if we restrict the analyses to marriages only? Table 5 compares estimates between Model A for first unions and a comparable Model A for first marriages. The later estimates for first marriages are from our companion paper (Härkönen and Dronkers, forthcoming). The first column repeats the estimates for education from the previous table, and the second column shows the estimates for the effects of education for first marriages. The third column gives the difference between the estimates and the fourth the standard error for this difference.

TABLE 5

The clear result is that in most cases focusing on first marriages instead of first unions does not make the conclusions about the effects of education different. The estimates show statistically significant differences for both educational levels in Germany, Lithuania and Switzerland. We can detect a (more or less) significant difference for high education in Greece, Latvia, Poland and the United States, and a weakly significant difference for middle education in Hungary. The difference in the estimates does not seem to correlate with the predominance of unmarried cohabitation. Furthermore, we do not find differences in countries like Sweden or Estonia, where cohabitation is the most common (cf. Hoem, 1997). In the cases where the difference is statistically significant, the sign of the coefficient is in most cases negative, although in others it is positive. Only in Switzerland does the sign of the coefficient change. All in all, we do not find support for our second hypothesis, and conclude that, in general, the choice between selecting all first unions versus selecting first marriages does not change the conclusions considerably.

Changes in the educational gradient of union dissolution?
Next we move on to test our third and fourth hypotheses, that is, we examine whether there has been a consistent change towards more a negative educational gradient in union dissolution, and whether this change has been more pronounced for marriages compared to all unions.

In Table 6, we test whether there has been a change in the educational gradient of union disruption. We focus here on all unions, and compare Models A with Models A with the interaction between education and union dissolution. We test the fit of these models with ordinary chi-square tests. The first column shows the difference in chi-squares and degrees of freedom between the models in each country, the second column gives the statistical significance of the difference, and the last four columns report the coefficients of the main terms and interaction terms of the equation.

TABLE 6

Table 6 shows that in six countries (France, Greece, Lithuania, Poland, Sweden and the USA), the model with the interaction term provides a better fit than the one without at the 5 % level of significance, and in one additional country (Hungary) the fit is better at the 10 % level of significance. Therefore, in seven out of the seventeen countries there has been a change in the educational gradient of union disruption. With the exception of Greece, the coefficients show that the educational gradient has become increasingly negative, that is, the risk of union dissolution has increased among the less educated relative to those with more education. In two countries (France and Poland), both interaction coefficients have a negative sign, and in Greece, both have a positive sign. In the other countries, there has been a decrease of the relative risks of dissolution only for women with tertiary education. It should also be noted that in Finland, Italy, and Latvia, the interaction term between tertiary education and year of marriage is negative and significant, although the model with both interaction terms does not fit better than the one without. In conclusion, our third hypothesis (the educational gradient becomes more negative) receives full support for two countries (France and Poland), and partial support (the relative risk has decreased for women with tertiary education) for seven countries. Greece is the obvious outlier, there the change in the educational gradient strongly contradicts our prediction. Finally, figures 2a to 2q show these developments in the educational gradient of union disruption in each of the countries in a more visual form. The graphs show the risk of union disruption for women with different educational levels at different periods (measured by union cohort), relative to the disruption risk of the low educated women in the oldest cohort (reference group), after controlling for all independent variables of model B in table 4.

FIGURES 2A TO 2Q

Would our conclusions be different if we focused on first marriages only? We examine this in Table 7. The first two rows report whether the difference in the χ^2 -statistic between the main effects model and the interaction term model was significant or not, for the models for first unions and first marriages, respectively. The next four columns show the coefficients for the interaction terms for comparison.

TABLE 7

The results in Table 7 lead to a similar conclusion as those in Table 5: the choice between first unions and first marriages does not, in most countries, change our conclusions. In Flanders, Finland and Italy we can find a significant change in the educational gradient of disruptions in the case of first marriages but not in the case of all unions (this result supporting our fourth hypothesis), while in Greece the situation is the opposite. In all the other countries, the choice of the specific type of union does not change the main conclusions.

Results from multilevel event-history analyses[6]

Table 8 begins the multilevel analyses (model A) with union disruption as the dependent variable and only the duration and duration squared as independent variables, but with the

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[6] We restrict ourselves her to the multilevel analysis of the union disruption of all unions. The results for only marriages are basically similar (Härkönen & Dronkers, forthcoming).

merged file of all countries[7]. As one can see in the bottom row, significant between-country-variance remains. Model B adds the all the micro variables and the interactions 'middle education*union year' and 'high education*union year'. The two parameters of 'high education' and 'middle education' are both significant and positive, but the interaction 'high education*union year' is also significant and negative. This means that in the oldest union cohort the educational gradient on union disruption is positive but that it is smaller in the younger union cohorts. We tested (results here not shown) whether the parameter estimates of 'high education' and 'middle education' vary across the countries. Because the random-variance of these two parameters was significantly large, these parameters are indeed different between countries.

TABLE 8

Models C to K add separately each macro-variable and its interaction with the two educational level dummies to model B in order to explain the change in the effect of education on union disruption. Most macro-variables have a significant effect on union disruption. More liberal divorce laws, higher social expenditure, higher percentage atheists, higher acceptance of divorce as justifiable, higher percentages unconventional family types, a higher urbanisation degree, more employment in services and higher percentage economic active women increase the odds of union disruption. But more family cash benefits decreases the odds of union disruption. The parameter estimates of the micro-variables remain considerably stable. The interactions between the macro-variables and educational level are in many cases significant. The results show the risk of union disruption is relatively lower for women with middle or high education than for those with low education when unconventional family types are common, the urbanisation level is higher, the service sector is larger, and when women are more economically active at the labour market. This suggests support for our fourth hypothesis. However, with increasing government spending on social policies in general and family cash benefits in particular, the relationship between education and disruption risk seems to become more positive, contrary to what we expected. Next we examine how these results change when we include all the significant macro-level variables into the same model.

TABLE 9

In table 9 we combine the multilevel analyses of table 8. Model L is equal to model B of table 7, but without the insignificant interaction 'middle education*union year'. Model M includes the main effects of those macro variables that had a significant influence on the odds of union disruption and declined the amount of between-country-variance: unconventional family types, employment in service sector and percentages of economically active women. The main effects of these three macro variables are as expected. Model N further includes all interactions between these three macro-variables and high educational level. The parameter estimates of the interaction terms mainly have the same sign as those in Table 8, as suggested by our fifth hypothesis. However, the interaction 'unconventional family type*union year' is not longer significant. Therefore, Model M supports our fourth hypothesis only partially, with divorce legislation being the main, and unexpected, exception. Our final model O has only two significant interactions 'employment in service sector*union year' and 'percentages of economically active women*year of union'. Moreover by the inclusion of these two interaction the parameter of the interaction 'high education*union year' has decreased substantially. But our fifth hypothesis is our partial confirmed by these results More liberal

[7] Norway excluded, since it misses information on parental divorce.

divorce legislation and normative environment towards family issues, more generous social welfare systems and "modernity" does not change the effect of education on union disruption. Only female labour market opportunities and more prosperity makes the effect of education on union disruption more negative.

Discussion

In this paper, we have examined the relationship between female educational attainment and the risk of union disruption across union cohorts and sixteen European countries and the USA with data from the Fertility and Family Surveys. On the basis of the theoretical discussions and previous research, we expected that this effect varies across countries, and that the effect of education on union disruption would become more negative across time. We also aimed to link this variation to cross-national differences in divorce legislation, social policies, values, family practices, and social and labour market conditions across time and between countries.

Although we report important cross-national variation in the effects of education on union disruption (positive or zero in some, negative in others), we did not find a constant pattern towards a more negative relationship through time. Instead, we found cross-national differences with regard to this development: in many countries, the effect remained stable, while in some it did become more negative over time, and in Greece, it seemed to have become more positive, contrary to our expectations. Therefore, William Goode's (1962) hypothesis, which predicts that the increase in divorce risks will be more rapid in the lower than the upper strata, received partial support. But the support is both for the change of the educational gradient for disruption of first unions as for the divorce of first marriages. The different selectivity of marriages and unions is not a relevant distinction for the explanation of the change of the educational gradient for union or marriage disruption, contrary to Hoem' s assumption (1997).

We sought to explain this variation across countries and across time with several societal level variables reflecting the social, legal, and economic costs of union disruption. Here we again follow Goode's original hypothesis, which claims that these costs have a more important impact on the disruption risks of the lower classes (or education, as in our case). We find support for some, but not all, of our society-level measures. When we entered the macro-level variables one by one, we found results in the expected direction (lower costs of disruption associated with a more negative relationship between education and dissolution risk) in most cases, social policies being the exception. The latter, unexpected, result may suggest that social policies can reduce economic strain of the less-off, thus reducing their risks of union dissolution. In the next step, we added the four macro-variables which had significant effects and reduced between-country variance (commonness of unconventional family types, size of the service sector, and female labour market activity rates) into a single multi-level event history model. The interaction term estimates remain rather similar, except in the case of commonness of unconventional family types.

If we accept this summary of our results, Goode seems to have been right. When the (social) costs of union dissolution are high, one needs extra resources to dissolve the union, while when they are low, one needs more resources to maintain a relationship. Therefore, it seems that strict divorce regimes bias the composition of union dissolution towards the more well-off, and make union disruption extremely hard in the lower ranks of society. Lax regimes, on the other hand, have the consequence of increasing dissolution risks especially for those with less education.

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Tables and Figures

Table 1: Percentage of first unions, first started as consensual unions, by decade when union began

	1950s	1960s	1970s	1980s	1990s	Total
Austria		2,5	11,4	41,2	74,1	34,4
Flanders			1,5	15,2	48,7	12,8
Estonia		65,5	66,1	76,6	90,8	78,1
Finland	1,1	2,5	18,8	72,8		32,4
France		2,7	14,1	60,7	94,6	47,5
West-Germany			16	41,7	84	42,2
Greece			1,7	9,9	27,1	11,5
Hungary			2	8,6	34,7	9
Italy		0	0,5	4,1	14,8	4,6
Latvia		2,8	4,7	10,6	44,4	14,4
Lithuania		0	0,7	2,8	14,4	4,7
Norway		2	16,4	78,2		43,2
Poland		0,5	0,9	3,4	12,3	2,3
Spain		0,8	0,9	5	22,5	6,9
Sweden		33,3	71,5	88,3	93,9	82,1
Switzerland		5,9	18,3	46,3	67,6	37,7
USA		2,7	16,7	31,9	57,6	31,8

Source: UNECE-PAU: Fertility and Family Surveys

Table 2: Year of data collection, number of cases, person-years units and events (dissolution) by country

	Year collected	Cases	Person-years	Dissolutions
Austria	1995-96	3 860	54 447	928
Flanders	1991-92	2 648	25 451	324
Estonia	1994	1 442	16 512	416
Finland	1989-90	3 689	48 075	884
France	1994	2 493	28 112	1 012
West Germany	1992	1 697	15 155	426
Greece	1999	1 930	28 792	207
Hungary	1992-93	2 911	30 786	544
Italy	1995-96	3 245	47 125	211
Latvia	1995	2 204	25 918	693
Lithuania	1994-95	2 307	27 525	366
Norway	1988-89	3 185	32 207	756
Poland	1991	3 267	45 792	226
Spain	1994-95	2 812	36 747	216
Sweden	1992-93	2 994	30 364	1 104
Switzerland	1994-95	3 468	41 317	852
USA	1995	7 998	69 853	3 715
Total		52 150	604 178	12 280

Table 3: Descriptive statistics (mean of year-person units and standard deviation) of the individual characteristics and the macro-variables

	Total	Austria	Flanders	Estonia	Finland	France	Germany	Greece	Hungary	Italy	Latvia	Lithuania	Norway	Poland	Spain	Sweden	Switzer land	USA
Duration in years	7.90 6.42	9.70 7.76	5.90 4.63	7.77 6.38	8.66 6.87	7.76 6.40	5.97 4.87	8.87 6.64	6.63 5.17	9.00 6.78	7.88 6.36	7.97 6.43	7.02 5.68	8.71 6.62	8.10 6.25	7.29 6.01	7.92 6.34	6.65 5.73
Parental divorce	.10 .31	.09 .29	.07 .26	.21 .41	.06 .23	.12 .33	.10 .31	.03 .17	.15 .36	.03 .16	.20 .40	.15 .36	-	.03 .16	.04 .19	.11 .31	.10 .30	.22 .42
Starting age union	21.60 3.28	21.26 3.38	21.39 2.43	21.68 2.99	21.87 3.35	21.05 2.99	21.07 3.03	21.51 3.80	20.26 2.69	22.33 3.50	21.71 3.05	22.05 2.92	21.91 2.91	21.67 3.05	22.52 3.32	20.70 3.01	22.79 3.28	20.98 3.51
Child before union	.06 .23	.13 .34	.01 .08	.04 .18	.05 .22	.04 .20	.06 .24	.02 .13	.03 .17	.03 .17	.04 .19	.05 .22	.08 .27	.04 .20	.02 .14	.05 .21	.03 .18	.11 .31
Partnership cohort	74.39 7.25	72.77 8.37	76.34 4.93	74.29 6.96	68.54 7.48	73.94 6.83	77.07 5.35	78.46 6.40	76.27 5.33	74.77 6.90	75.56 6.80	75.82 7.16	71.29 6.31	71.44 6.86	76.26 6.63	73.82 6.43	75.14 6.83	76.99 6.46
Middle education level	.43 .50	.48 .50	.37 .48	.38 .49	.56 .50	.38 .49	.37 .48	.43 .50	.39 .49	.31 .46	.70 .50	.36 .48	.42 .49	.34 .47	.14 .35	.57 .50	.75 .43	.37 .48
High education level	.24 .43	.43 .50	.27 .44	.22 .42	.13 .34	.17 .37	.10 .30	.19 .39	.10 .30	.07 .26	.23 .42	.60 .49	.38 .49	.13 .34	.11 .31	.29 .45	.12 .32	.43 .50
Strict, Institutionalized divorce laws	.09 .28	.00 .00	.86 .35	.00 .03	.00 .00	.14 .35	.00 .00	.18 .38	.00 .00	.46 .50	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.01 .07
Less strict, more individual divorce laws	.84 .37	1.00 .00	.14 .35	1.00 .03	.83 .38	.86 .35	1.00 .00	.82 .38	1.00 .00	.52 .50	1.00 .00	1.00 .00	1.00 .00	1.00 .00	.81 .40	.09 .28	1.00 .00	.99 .07
Pure unilateral divorce laws	.06 .24	.00 .00	.00 .00	.00 .00	.17 .38	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.00 .00	.91 .28	.00 .00	.00 .00
% social expenditure of GDP	19.01 5.25	23.45 2.42	24.95 1.40	13.50 .61	18.49 4.63	21.90 4.63	21.63 1.96	19.57 4.29	18.24 1.47	22.56 2.44	13.54 .55	13.54 .52	21.01 3.72	16.20 .00	17.94 2.09	29.95 4.11	16.99 3.61	12.93 .96
% of family cash benefits of GDP	12.28 8.46	22.25 2.58	25.34 3.63	2.89 .32	12.59 4.50	21.88 1.51	12.71 2.88	7.87 3.24	27.10 3.27	8.40 2.14	2.92 .28	2.93 .26	12.33 2.47	18.00 .00	3.16 1.50	18.86 4.02	9.19 2.76	2.57 1.18
% atheists or non-religious	11.49 9.62	7.10 1.06	7.21 .04	39.61 2.40	5.28 .60	19.05 1.18	13.99 6.42	1.85 .05	12.55 .26	15.77 .57	35.38 1.66	14.17 1.78	2.14 .25	3.36 1.24	5.41 .29	29.31 .64	7.23 .78	8.82 .34
'divorce justifiable'	52.61 7.94	49.00 .00	46.95 3.78	54.00 .00	64.92 7.00	55.72 1.89	55.35 3.24	63.00 .00	48.86 2.10	53.30 1.27	54.00 .00	40.00 .00	50.70 2.49	39.00 .00	54.25 3.57	62.33 .95	59.95 6.96	48.55 .84
unconventional family types	82.41 46.14	105.85 24.54	55.01 16.89	135.57 9.173	105.55 41.41	93.99 30.90	86.34 7.73	40.96 5.68	67.71 13.53	18.45 5.29	109.81 9.04	74.24 3.63	99.09 38.10	25.29 3.70	25.44 10.05	170.46 14.67	84.23 16.11	122.91 19.32
degree urbanization	68.62 9.63	67.12 .16	95.98 .57	70.10 1.58	58.40 4.19	73.48 .77	84.14 1.40	58.49 .71	59.69 2.81	66.57 .47	69.16 1.85	64.81 4.42	70.32 1.95	58.52 2.51	74.30 1.84	82.97 .50	58.44 1.56	74.60 .74
Employment in services	53.02 10.87	51.71 3.92	65.98 4.04	42.79 1.73	53.98 5.46	51.30 5.11	53.58 5.12	47.38 4.16	43.00 4.57	55.55 6.04	43.09 1.81	38.14 3.55	62.93 5.04	33.72 3.07	51.49 5.65	61.94 4.24	57.13 3.86	68.84 2.94
% women economically active	58.47 12.22	54.31 1.83	44.69 3.39	77.22 1.44	67.99 3.95	55.61 2.31	55.19 2.18	39.75 4.66	60.44 1.44	42.49 3.35	76.16 1.16	71.94 2.06	59.91 9.35	66.55 1.26	37.69 5.18	73.39 8.16	56.69 4.83	62.90 4.77

Table 4: Relationship between female educational attainment and the risk of union disruption, discrete-time event history models (standard errors in parentheses)

	Austria		Flanders		Estonia		Finland		France		Germany	
Model	A	B	A	B	A	B	A	B	A	B	A	B
Middle education	-0.223**	-0.248**	0.035	0.115	-0.075	-0.036	-0.070	0.074	0.191**	0.235**	0.253*	0.102
	(0.078)	(0.079)	(0.131)	(0.134)	(0.160)	(0.163)	(0.085)	(0.086)	(0.073)	(0.074)	(0.108)	(0.111)
High education	-0.231*	-0.270*	-0.214	0.104	-0.285	-0.078	-0.155	0.221+	0.228*	0.367**	0.310+	0.285
	(0.105)	(0.105)	(0.158)	(0.173)	(0.190)	(0.196)	(0.113)	(0.117)	(0.091)	(0.095)	(0.167)	(0.174)
Year of union	0.045**	0.032**	0.019	0.018	-0.008	-0.008	0.044**	0.015*	0.062**	0.059**	0.053**	0.057**
	(0.005)	(0.005)	(0.014)	(0.015)	(0.008)	(0.008)	(0.006)	(0.006)	(0.006)	(0.006)	(0.011)	(0.012)
Duration	0.013	0.112**	0.075+	0.135**	0.050	0.060+	-0.052**	0.043*	0.090**	0.107**	-0.075*	0.058
	(0.017)	(0.019)	(0.045)	(0.046)	(0.033)	(0.034)	(0.018)	(0.019)	(0.019)	(0.019)	(0.036)	(0.038)
Duration squared	-0.002*	-0.005**	-0.004	-0.007*	-0.006**	-0.007**	0.001+	-0.002*	-0.003**	-0.003**	0.004*	-0.001
	(0.001)	(0.001)	(0.003)	(0.003)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
Parental divorce	0.777**	0.583**	1.023**	0.740**	0.376**	0.307**	0.713**	0.581**	0.368**	0.274**	0.789**	0.519**
	(0.089)	(0.091)	(0.149)	(0.155)	(0.112)	(0.113)	(0.111)	(0.112)	(0.086)	(0.088)	(0.125)	(0.129)
Out-of-wedlock		0.013		0.592		0.468+		0.512**		0.175		0.127
		(0.113)		(0.472)		(0.242)		(0.142)		(0.156)		(0.218)
Age at start union		-0.106**		-0.160**		-0.098**		-0.085**		-0.074**		-0.095**
		(0.013)		(0.030)		(0.021)		(0.012)		(0.013)		(0.021)
Marriage (t)		-1.309**		-1.362**		-0.109		-1.449**		-0.309**		-1.389**
		(0.093)		(0.155)		(0.105)		(0.096)		(0.084)		(0.118)
Constant	-7.425**	-3.619**	-6.197**	-1.894	-2.843**	-0.871	-6.869**	-2.375**	-8.746**	-6.835**	-7.827**	-5.581**
	(0.415)	(0.489)	(1.165)	(1.251)	(0.625)	(0.760)	(0.426)	(0.520)	(0.485)	(0.567)	(0.953)	(0.978)
Person-years	52697	52697	25338	25338	16353	16353	47473	47473	27540	27540	14758	14758
Log-likelihood	-4239.41	-4087.74	-1688.28	-1639.45	-1871.27	-1858.44	-4215.21	-4060.22	-4177.18	-4150.83	-1851.59	-1769.99
Chi-square	297.32	600.66	48.58	146.23	82.31	107.95	196.74	506.73	181.07	233.77	93.35	256.55

Table 4.ctd.

	Greece		Hungary		Italy		Latvia		Lithuania		Norway	
Model	A	B	A	B	A	B	A	B	A	B	A	B
Middle education	0.357	0.432+	0.023	0.142	0.429**	0.538**	-0.213	-0.043	-0.542**	-0.485*	-0.026	0.118
	(0.232)	(0.235)	(0.097)	(0.100)	(0.158)	(0.161)	(0.148)	(0.150)	(0.204)	(0.204)	(0.112)	(0.114)
High education	1.114**	0.842**	-0.038	0.172	1.012**	1.112**	-0.296+	-0.021	-0.352+	-0.245	0.008	0.213+
	(0.210)	(0.227)	(0.143)	(0.152)	(0.204)	(0.221)	(0.164)	(0.170)	(0.209)	(0.211)	(0.110)	(0.116)
Year of union	0.019	0.009	0.015	0.007	0.052**	0.056**	0.010	0.006	0.022*	0.021*	0.077**	0.026**
	(0.013)	(0.014)	(0.009)	(0.010)	(0.012)	(0.013)	(0.006)	(0.006)	(0.009)	(0.009)	(0.007)	(0.008)
Duration	-0.182**	-0.058	-0.071*	-0.011	0.029	0.079*	-0.020	0.021	0.060+	0.074*	-0.136**	0.011
	(0.045)	(0.049)	(0.031)	(0.033)	(0.038)	(0.039)	(0.023)	(0.024)	(0.032)	(0.032)	(0.023)	(0.025)
Duration squared	0.005*	0.001	0.002	-0.001	-0.000	-0.002	-0.001	-0.003**	-0.004*	-0.004**	0.006**	0.000
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Parental divorce	1.204**	0.907**	0.417**	0.302**	1.240**	0.869**	0.415**	0.305**	0.620**	0.527**	-	-
	(0.277)	(0.284)	(0.109)	(0.111)	(0.256)	(0.265)	(0.088)	(0.090)	(0.124)	(0.127)		
Out-of-wedlock		0.420		0.554**		0.733*		0.322+		0.194		0.277*
		(0.596)		(0.210)		(0.293)		(0.175)		(0.221)		(0.135)
Age at start union		-0.030		-0.072**		-0.104**		-0.076**		-0.057**		-0.109**
		(0.025)		(0.021)		(0.024)		(0.015)		(0.020)		(0.015)
Marriage (t)		-1.855**		-1.535**		-2.207**		-0.913**		-0.870**		-1.628**
		(0.215)		(0.136)		(0.202)		(0.120)		(0.223)		(0.105)
Constant	-6.176**	-3.602**	-4.918**	-1.737*	-10.03**	-6.241**	-3.966**	-1.528*	-5.816**	-3.799**	-9.043**	-2.413**
	(1.143)	(1.213)	(0.767)	(0.828)	(1.068)	(1.148)	(0.521)	(0.598)	(0.748)	(0.872)	(0.592)	(0.713)
Person-years	22803	22803	29587	29587	45856	45856	25675	25675	27318	27318	31409	31409
Log-likelihood	-857.47	-818.90	-2551.99	-2490.85	-1289.92	-1232.97	-3102.76	-3063.34	-1892.83	-1882.36	-3332.17	-3167.11
Chi-square	111.32	188.46	49.08	171.35	69.05	182.95	96.32	175.16	55.23	76.16	270.81	600.91

Table 4.ctd.

	Greece		Hungary		Italy		Latvia		Lithuania	
Model	A	B	A	B	A	B	A	B	A	B
Middle education	0.357	0.432+	0.023	0.142	0.429**	0.538**	-0.213	-0.043	-0.542**	-0.485*
	(0.232)	(0.235)	(0.097)	(0.100)	(0.158)	(0.161)	(0.148)	(0.150)	(0.204)	(0.204)
High education	1.114**	0.842**	-0.038	0.172	1.012**	1.112**	-0.296+	-0.021	-0.352+	-0.245
	(0.210)	(0.227)	(0.143)	(0.152)	(0.204)	(0.221)	(0.164)	(0.170)	(0.209)	(0.211)
Year of union	0.019	0.009	0.015	0.007	0.052**	0.056**	0.010	0.006	0.022*	0.021*
	(0.013)	(0.014)	(0.009)	(0.010)	(0.012)	(0.013)	(0.006)	(0.006)	(0.009)	(0.009)
Duration	-0.182**	-0.058	-0.071*	-0.011	0.029	0.079*	-0.020	0.021	0.060+	0.074*
	(0.045)	(0.049)	(0.031)	(0.033)	(0.038)	(0.039)	(0.023)	(0.024)	(0.032)	(0.032)
Duration squared	0.005*	0.001	0.002	-0.001	-0.000	-0.002	-0.001	-0.003**	-0.004*	-0.004**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
Parental divorce	1.204**	0.907**	0.417**	0.302**	1.240**	0.869**	0.415**	0.305**	0.620**	0.527**
	(0.277)	(0.284)	(0.109)	(0.111)	(0.256)	(0.265)	(0.088)	(0.090)	(0.124)	(0.127)
Out-of-wedlock		0.420		0.554**		0.733*		0.322+		0.194
		(0.596)		(0.210)		(0.293)		(0.175)		(0.221)
Age at start union		-0.030		-0.072**		-0.104**		-0.076**		-0.057**
		(0.025)		(0.021)		(0.024)		(0.015)		(0.020)
Marriage (t)		-1.855**		-1.535**		-2.207**		-0.913**		-0.870**
		(0.215)		(0.136)		(0.202)		(0.120)		(0.223)
Constant	-6.176**	-3.602**	-4.918**	-1.737*	-10.03**	-6.241**	-3.966**	-1.528*	-5.816**	-3.799**
	(1.143)	(1.213)	(0.767)	(0.828)	(1.068)	(1.148)	(0.521)	(0.598)	(0.748)	(0.872)
Person-years	22803	22803	29587	29587	45856	45856	25675	25675	27318	27318
Log-likelihood	-857.47	-818.90	-2551.99	-2490.85	-1289.92	-1232.97	-3102.76	-3063.34	-1892.83	-1882.36
Chi-square	111.32	188.46	49.08	171.35	69.05	182.95	96.32	175.16	55.23	76.16

Table 5: Significant differences between the effects of education on union disruption for first unions or divorce for first marriages

		First unions	First marriages	Difference (mar – unions)	Standard error (difference)
Austria	Middle	-0.223**	-0.205*	0.018	0.052
	High	-0.231*	-0.273*	-0.042	0.075
Flanders	Middle	0.035	-0.040	-0.075	0.064
	High	-0.214	-0.372*	-0.158	0.096
Estonia	Middle	-0.075	-0.092	-0.017	0.159
	High	-0.372	-0.285	0.252	0.173
Finland	Middle	-0.070	-0.035	0.035	0.053
	High	-0.155	-0.139	0.016	0.082
France	Middle	0.191**	0.195*	0.004	0.047
	High	0.228*	0.154	-0.074	0.069
West-Germany	Middle	0.253*	-0.090	-0.343**	0.094
	High	0.310+	-0.376	0.686**	0.191
Greece	Middle	0.357	0.338	-0.019	0.101
	High	1.114**	0.697**	0.417**	0.137
Hungary	Middle	0.023	-0.060	-0.083+	0.043
	High	-0.038	-0.074	-0.036	0.062
Italy	Middle	0.429**	0.437**	0.008	0.005
	High	1.012**	0.876**	-0.136	0.126
Latvia	Middle	-0.213	-0.126	0.087	0.077
	High	-0.296+	-0.160	0.136+	0.081
Lithuania	Middle	-0.542**	-0.608**	-0.066**	0.020
	High	-0.352+	-0.445*	-0.093**	0.029
Norway	Middle	-0.026	-0.009	0.015	0.096
	High	0.008	0.086	0.078	0.101
Poland	Middle	0.342*	0.350*	0.008	0.039
	High	0.299	0.363+	0.064+	0.035
Spain	Middle	0.743**	0.798**	0.054	0.115
	High	0.617**	0.390	-0.227	0.178
Sweden	Middle	0.024	-0.012	-0.036	0.115
	High	0.064	0.020	-0.044	0.120
Switzerland	Middle	0.030	-0.187	-0.217**	0.063
	High	0.314*	0.014	-0.300**	0.112
USA	Middle	-0.221**	-0.277**	-0.056	0.036
	High	-0.289**	-0.409**	-0.120**	0.038

Table 6: Significance and direction of the change of the educational gradient of union disruption for the first unions

	χ² / df difference	p	Middle education (ref: low)	High education (ref: low)	Middle * year of mar	High * year of mar
Austria	0.57 / 2	0.753	-0.236	-0.085	0.001	-0.008
Flanders	1.91 / 2	0.385	0.139	0.210	-0.015	-0.046
Estonia	0.32 / 2	0.853	-0.017	-0.117	-0.007	-0.016
Finland	4.01 / 2	0.134	0.111	0.432	-0.012	-0.033*
France	5.99 / 2	0.050	0.493**	0.603**	-0.022*	-0.027*
Germany	0.51 / 2	0.773	0.219	0.523	0.003	-0.021
Greece	10.46 / 2	0.005	-0.177	-0.032	0.064†	0.106**
Hungary	4.73 / 2	0.094	0.227	0.530†	-0.023	-0.058*
Italy	3.96 / 2	0.138	0.604†	1.904**	-0.014	-0.061*
Latvia	4.11 / 2	0.128	0.128	0.270	-0.027	-0.044*
Lithuania	8.45 / 2	0.015	-0.429	-0.935*	-0.004	0.042
Norway	1.54 / 2	0.462	0.207	0.305	-0.017	-0.021
Poland	15.84 / 2	0.000	1.359**	1.633**	-0.072**	-0.091**
Spain	1.90 / 2	0.369	0.869†	1.385*	-0.008	-0.044
Sweden	13.16 / 2	0.001	0.143	0.623**	-0.013	-0.046**
Switzerland	3.52 / 2	0.172	-0.025	0.658*	0.004	-0.024
USA	28.85 / 2	0.000	-0.297*	0.134	0.005	-0.026**

Table 7: Differences between the effect of the interaction education*year formation on union disruption for first unions and the effect of the interaction education*year marriage divorce for first marriages

	p of model χ² - difference First unions	First marriages	Year formation *middle education First unions	Year marriage *middle education First marriages	Year formation *high education First unions	Year marriage *high education First marriages
Austria	n.s.	n.s.	0.001	0.013	-0.008	-0.004
Flanders	n.s.	**	-0.015	-0.034	-0.046	-0.138**
Estonia	n.s.	n.s.	-0.007	-0.012	-0.016	-0.032
Finland	n.s.	*	-0.012	-0.034*	-0.033*	-0.054*
France	†	†	-0.022*	-0.026*	-0.027*	-0.033+
Germany	n.s.	n.s.	0.003	-0.007	-.0.021	-0.071
Greece	**	n.s.	0.064+	0.049	0.106**	0.010
Hungary	†	*	-0.023	-0.045*	-0.058*	-0.072*
Italy	n.s.	†	-0.014	-0.023	-0.061	-0.092*
Latvia	n.s.	n.s.	-0.027	-0.016	-0.044*	-0.032
Lithuania	*	*	-0.004	-0.027	0.042	0.017
Norway	n.s.	n.s.	-0.017	-0.039	-0.021	-0.018
Poland	**	**	-0.072**	-0.071**	-0.091**	-0.090**
Spain	n.s.	n.s.	-0.008	0.007	-0.044	-0.068
Sweden	**	**	-0.013	-0.057*	-0.046**	-0.096**
Switzerland	n.s.	n.s.	0.004	0.009	-0.024	-0.011
USA	**	**	0.005	0.000	-0.026**	-0.030**

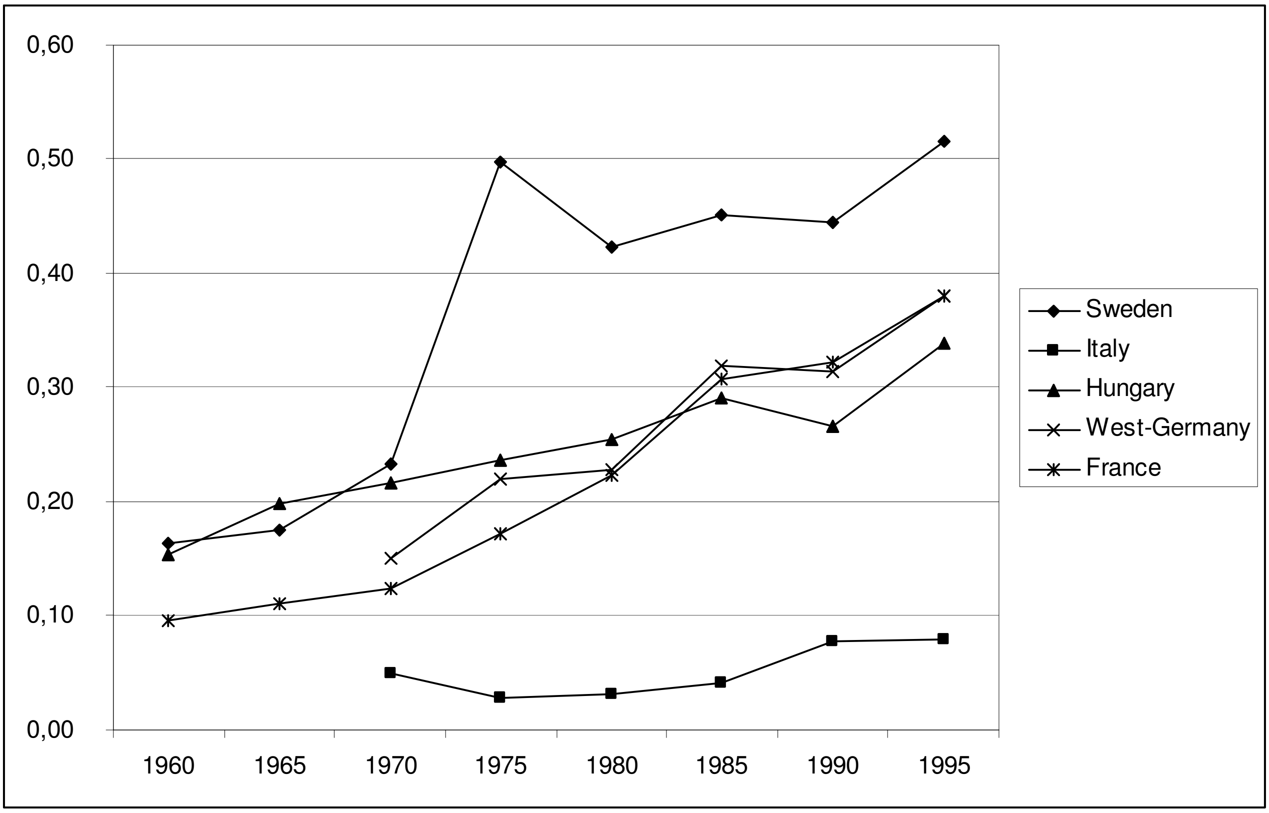
Table 8: Results from multi-level discrete-time event history models with each macro variable introduced separately. Standard errors between parentheses

	No macro-variables		Divorce laws		Social policies		Values		Family practices	Modernity and the labour market		
Micro & Macro variables	A: Empty event-history model	B: Equation with micro variables only	C: B & divorce laws (no & strict divorce as reference category)		D: B & social expenditure % GDP	E: B & Family cash benefits % GDP	F: B & % atheists or non-religious	G: B & % 'divorce justifiable'	H: B & unconventional family types	I: B & degree urbanization	J: B & Employment in services	K: B & economically active women (%)[1]
			Less strict, more individual	Pure unilateral divorce								
Middle education level		0.319 (0.232)	0.337 (0.239)		0.103 (0.236)	0.090 (0.239)	0.277 (0.239)	-0.012 (0.273)	0.301 (0.250)	0.394 (0.265)	0.576 (0.241)	0.926 (0.264)
High education level		1.386 (0.280)	1.350 (0.292)		1.121 (0.289)	1.207 (0.289)	1.403 (0.291)	1.081 (0.340)	1.398 (0.293)	1.873 (0.322)	1.639 (0.284)	2.466 (0.313)
Union year		0.032 (0.002)	0.027 (0.003)		0.026 (0.003)	0.031 (0.002)	0.032 (0.003)	0.028 (0.003)	0.016 (0.003)	0.030 (0.003)	0.013 (0.003)	0.030 (0.003)
Duration	-0.094 (0.004)	-0.073 (0.005)	-0.076 (0.005)		-0.079 (0.005)	-0.072 (0.005)	-0.073 (0.005)	-0.077 (0.005)	-0.086 (0.005)	-0.075 (0.005)	-0.089 (0.005)	-0.073 (0.005)
Duration^2	0.002 (0.000)	0.001 (0.000)	0.001 (0.000)		0.001 (0.000)	0.001 (0.000)	0.001 (0.000)	0.001 (0.000)	0.001 (0.000)	0.001 (0.000)	0.001 (0.000)	0.001 (0.000)
Parental divorce		0.504 (0.026)	0.506 (0.026)		0.506 (0.026)	0.506 (0.026)	0.504 (0.026)	0.505 (0.026)	0.513 (0.027)	0.503 (0.026)	0.506 (0.025)	0.503 (0.026)
Middle* Union year		-0.004 (0.003)	-0.002 (0.003)		-0.005 (0.003)	-0.003 (0.003)	-0.004 (0.003)	-0.004 (0.003)	0.000 (0.003)	-0.003 (0.003)	0.002 (0.003)	-0.003 (0.003)
High* Union year		-0.019 (0.004)	-0.017 (0.004)		-0.018 (0.004)	-0.018 (0.004)	-0.019 (0.004)	-0.017 (0.004)	-0.011 (0.004)	-0.017 (0.004)	-0.011 (0.004)	-0.016 (0.004)
Main effect macro variable			0.514 (0.062)	0.549 (0.137)	0.016 (0.005)	-0.009 (0.004)	0.012 (0.006)	0.013 (0.004)	0.008 (0.001)	0.014 (0.006)	0.040 (0.004)	0.014 (0.003)
Middle*macro variable			-0.225 (0.083)	-0.082 (0.136)	0.016 (0.005)	0.009 (0.003)	0.003 (0.002)	0.006 (0.003)	-0.003 (0.001)	-0.002 (0.002)	-0.013 (0.002)	-0.011 (0.002)
High*macro variable			-0.134 (0.108)	-0.239 (0.164)	0.011 (0.005)	0.007 (0.003)	0.000 (0.003)	0.004 (0.004)	-0.006 (0.001)	-0.009 (0.003)	-.016 (0.003)	-0.021 (0.002)
Variance at country level	0.362 (0.125)	0.362 (0.128)	0.328 (0.117)		0.404 (0.143)	0.358 (0.127)	0.352 (0.125)	0.359 (0.127)	0.219 (0.078)	0.325 (0.116)	0.252 (0.090)	0.345 (0.122)
-2 log likelihood	-680825	-670348	-684613		-657452	-667449	-678622	673785	-729520	-679913	-708957	-683424

Table 9: Union disruption Results from multi-level discrete-time event history models with all significant macro-variables and their interactions with high and middle education. Standard errors in parentheses

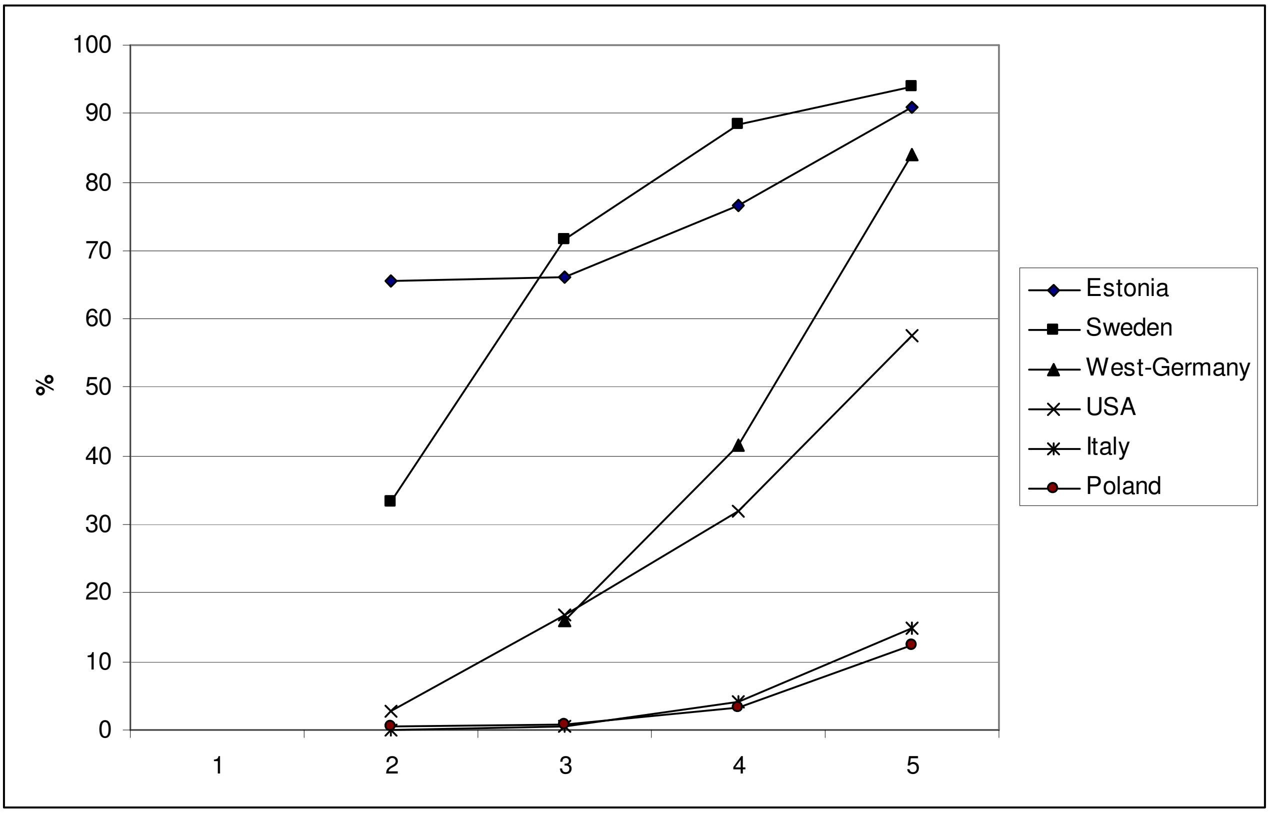
	L: B minus insignificant effects	M: L & macro-variables	N: M & interactions between the macro-variables and high education	O: N without insignificant interaction terms
High education level	1.178 (0.244)	1.115 (0.237)	1.698 (0.306)	1.927 (0.262)
Union year	0.029 (0.002)	0.009 (0.003)	0.008 (0.003)	0.008 (0.003)
Duration	-0.073 (0.005)	-0.096 (0.005)	-0.095 (0.005)	-0.095 (0.005)
Duration^2	0.001 (0.000)	0.002 (0.000)	0.001 (0.000)	0.001 (0.000)
Parental divorce	0.505 (0.026)	0.519 (0.025)	0.516 (0.025)	0.516 (0.025)
High education level * union year	-0.016 (0.003)	-0.015 (0.003)	-0.010 (0.003)	-0.010 (0.003)
Employment in services		0.031 (0.004)	0.031 (0.004)	0.032 (0.004)
% women economically active		-0.010 (0.003)	-0.006 (0.003)	-0.006 (0.003)
Unconventional family types		0.004 (0.001)	0.004 (0.001)	0.004 (0.001)
% women economically active* High education level			-0.009 (0.003)	-0.013 (0.002)
Unconventional family types * High education level			-0.001 (0.001)	
Employment in services * High education level			-0.005 (0.003)	-0.007 (0.002)
Variance at country level	0.361 (0.128)	0.225 (0.080)	0.214 (0.076)	0.214 (0.077)
-2 log likelihood	-670232	-716230	-722394	-721994

Figure 1. Divorce trends (total divorce rate) in France, Hungary, Italy, Sweden and West-Germany



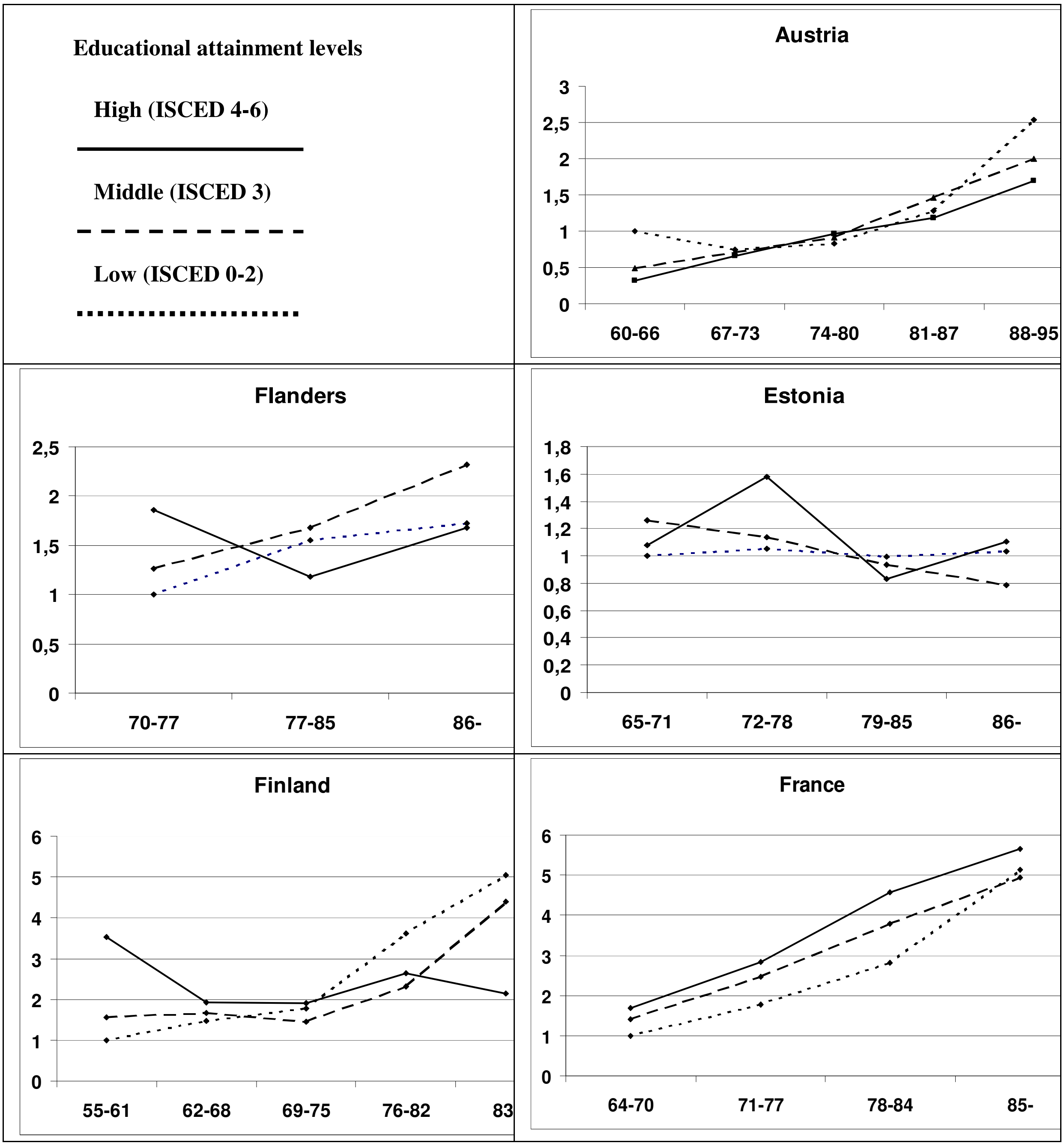
Source: Council of Europe (2003) Recent Demographic Developments in Europe

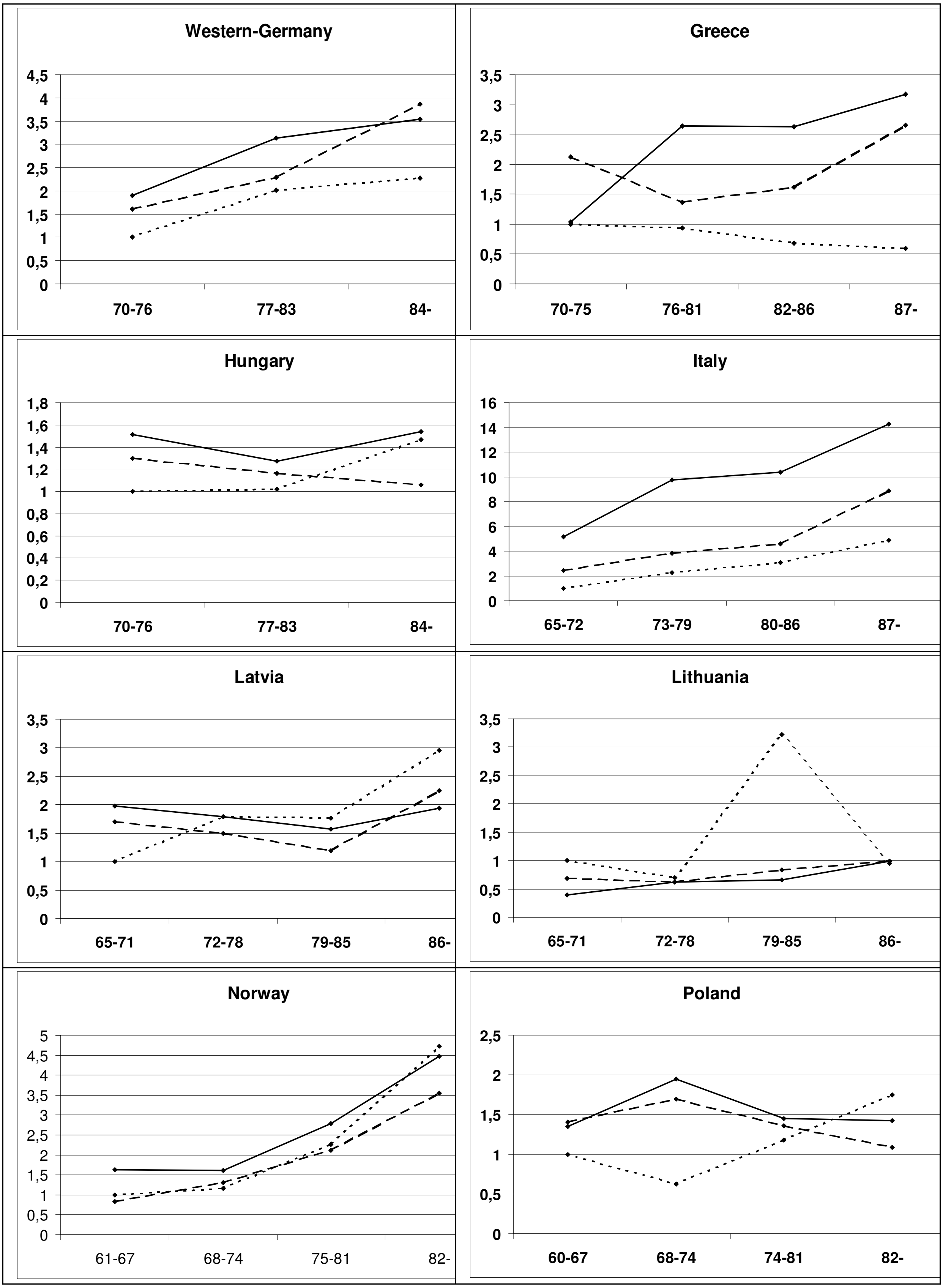
Figure 2. Percentage of first unions, first started as consensual unions, by decade when union began

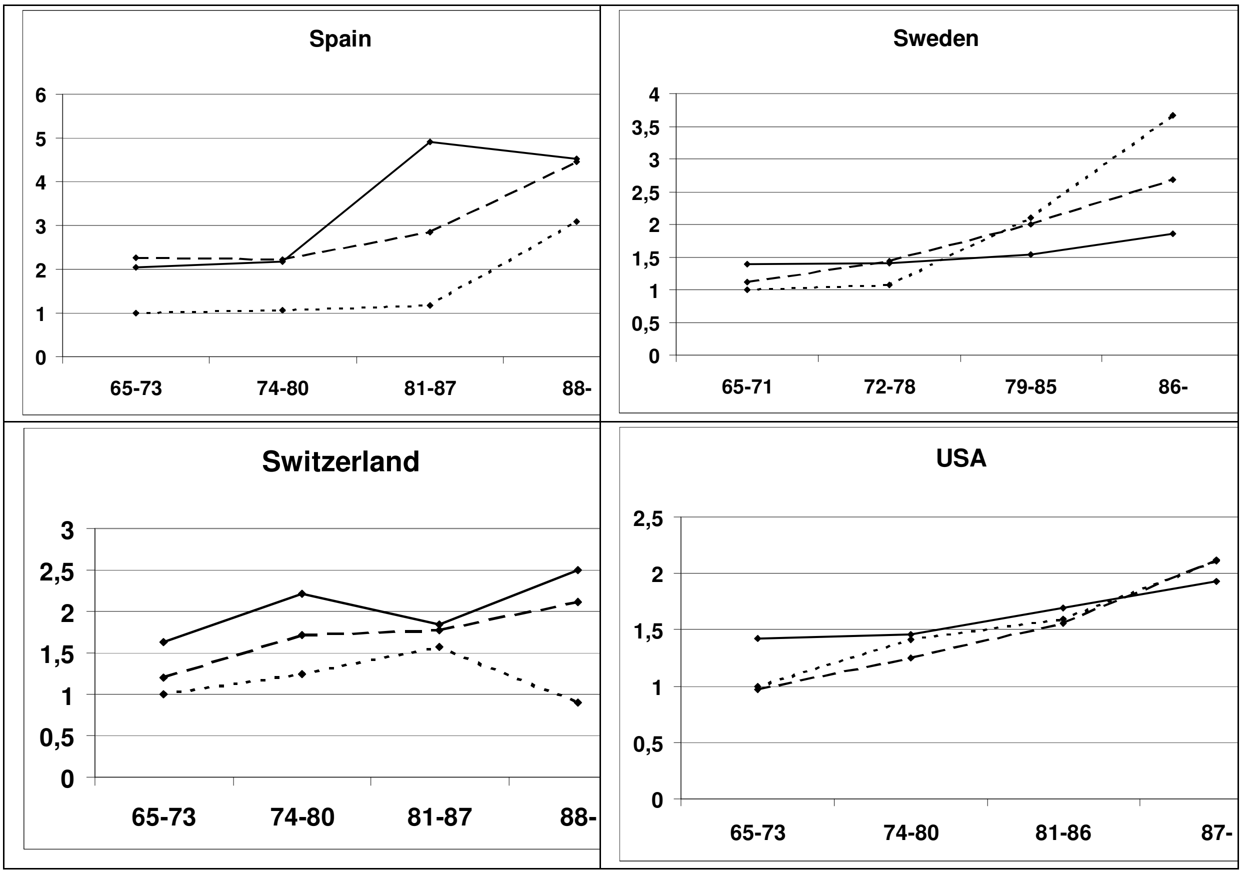


Source: UNECE-PAU: Fertility and Family Surveys

Figure 2a to q. Changes across union-cohorts in the effects of educational attainment on the risk of union dissolution, odds ratios (reference category: lowest education, oldest cohort)




Western-Germany
4,5
4
3,5
3
2,5
2
1,5
1
0,5
0
70-76
77-83
84-
Greece
3,5
3
2,5
2
1,5
1
0,5
0
70-75
76-81
82-86
87-
Hungary
1,8
1,6
1,4
1,2
1
0,8
0,6
0,4
0,2
0
70-76
77-83
84-
Italy
16
14
12
10
8
6
4
2
0
65-72
73-79
80-86
87-
Latvia
3,5
3
2,5
2
1,5
1
0,5
0
65-71
72-78
79-85
86-
Lithuania
3,5
3
2,5
2
1,5
1
0,5
0
65-71
72-78
79-85
86-
Norway
5
4,5
4
3,5
3
2,5
2
1,5
1
0,5
0
61-67
68-74
75-81
82-
Poland
2,5
2
1,5
1
0,5
0
60-67
68-74
74-81
82-


Spain
6
5
4
3
2
1
0
65-73
74-80
81-87
88-
Sweden
4
3,5
3
2,5
2
1,5
1
0,5
0
65-71
72-78
79-85
86-
Switzerland
3
2,5
2
1,5
1
0,5
0
65-73
74-80
81-87
88-
USA
2,5
2
1,5
1
0,5
0
65-73
74-80
81-86
87-

Appendices

Table A1: Divorce legislation and social policies

	Divorce legislation	Social exp. % GDP	Family cash ben. % GDP
Austria	Breakdown or other less restrictive (2)	1980: 23.3 1990: 25.0	1980: 2.73 1990: 2.06
Flanders	Until 1975: Fault or other restrictive (1) Since 1975: Breakdown or other less restrictive (2)	1980: 24.2 1990: 24.6	1980: 2.93 1990: 2.19
Estonia	Until 1966: Fault or other restrictive (1) [1] Since 1966: Breakdown or other less restrictive (2) [1]	1965: 11.6[1] 1983: 13.8[1]	1980: 0.3[1] 1990: 0.3[1]
Finland	Until 1987: Breakdown or other less restrictive (2) Since 1987: No judicial ground to deny (3)	1980: 18.5 1990: 24.8	1980: 1.07 1990: 1.88
France	Until 1975: Fault or other restrictive (1) Since 1975: Breakdown or other less restrictive (2)	1980: 21.1 1990: 26.5	1980: 2.18 1990: 2.10
West Germany	Breakdown or other less restrictive (2)	1980: 20.3 1990: 20.3	1980: 1.60 1990: 1.09
Greece	Until 1983: Fault or other restrictive (1) Since 1983: Breakdown or other less restrictive (2)	1980: 11.5 1990: 21.6	1980: 0.34 1990: 0.84
Hungary	Until 1964: Fault or other restrictive (1) Since 1964: Breakdown or other less restrictive (2)	1965: 10.9 1983: 18.7	1980: 3.0 1990: 2.6
Italy	Until 1971: Not permitted (0) 1971-1987: Fault or other restrictive (1) Since 1987: Breakdown or other less restrictive (2)	1980: 18.4 1990: 23.9	1980: 0.99 1990: 0.72
Latvia	Until 1966: Fault or other restrictive (1) [1] Since 1966: Breakdown or other less restrictive (2) [1]	1965: 11.6[1] 1983: 13.8[1]	1980: 0.3[1] 1990: 0.3[1]
Lithuania	Until 1966: Fault or other restrictive (1) [1] Since 1966: Breakdown or other less restrictive (2) [1]	1965: 11.6[1] 1983: 13.8[1]	1980: 0.3[1] 1990: 0.3[1]
Norway	Until 1991: Breakdown or other less restrictive (2) Since 1991: Unilateral non-fault (3)	1980: 18.6 1990: 26.0	1980: 1.25 1990: 1.91
Poland	Until 1964: Fault or other restrictive (1) Since 1964: Breakdown or other less restrictive (2)	1990: 16.2	1990: 1.84 1998: 0.93
Spain	Until 1981: Not permitted (0) Since 1981: Fault or other restrictive (1)	1980: 15.8 1990: 19.3	1980: 0.45 1990: 0.19
Sweden	Until 1974: Breakdown or other less restrictive (2) Since 1974: No judicial ground to deny (3)	1980: 29.0 1990: 31.0	1980: 1.75 1990: 2.15
Switzerland	Breakdown or other less restrictive (2)	1980: 15.2 1990: 19.8	1980: 1.04 1990: 1.02
USA	Until 1970: Fault or other more restrictive (1) 1970-1985: Breakdown or other less restrictive (2) Since 1985: Unilateral non-fault (3)	1980: 13.1 1990: 13.4	1980: 0.46 1990: 0.22

Sources: Chester (1977); Castles and Flood (1993); Glendon (1987; 1989); Goode (1993); Hamilton and Perry (2002); ILO (1967; 1988); OECD (1987-1998) Social Expenditure Database; Fajth and Zimakova (1997: 124).
[1] Former USSR. [2] Not available.

Table A2: Economy

	Urbanisation	Women's economic activity	Employment in services
Austria	1960: 66.8 1990: 67.0	1960: 53.0 1990: 55.2	1960: 29.89 1990: 54.71
Flanders	1960: 92.5 1990: 96.5	1960: 30.5 1990: 47.8	1960: 44.52 1990: 69.6
Estonia	1960: 57.5 1990: 71.1	1960: 67.3 1990: 75.9	1960: 35.6 1990: 44.3
Finland	1960: 38.1 1990: 61.4	1960: 55.5 1990: 72.4	1960: 37.1 1990: 61.0
France	1960: 62.4 1990: 74.0	1960: 43.6 1990: 57.0	1960: 39.2 1990: 55.9
West Germany	1960: 76.1 [1] 1990: 85.3 [1]	1970: 48.1 [2] 1990: 57.0 [2]	1960: 37.3 1990: 57.9
Greece	1960: 42.9 1990: 58.8	1960: 26.3 1990: 42.4	1960: 26.6 1990: 49.7
Hungary	1960: 42.6 1990: 62.0	1960: 46.9 1990: 59.3	1960: 27.0 1990: 46.9
Italy	1960: 59.4 1990: 66.7	1960: 30.4 1990: 45.0	1960: 29.8 1990: 60.0
Latvia	1960: 56.9 1990: 70.3	1960: 64.3 1990: 75.3	1960: 32.4 1990: 44.5
Lithuania	1960: 40.0 1990: 67.8	1960: 61.2 1990: 70.4	1960: 25.4 1990: 40.9
Norway	1960: 49.9 1990: 72.0	1960: 26.8 1990: 69.1	1960: 43.5 1990: 68.4
Poland	1960: 47.9 1990: 60.7	1960: 62.1 1990: 65.1	1960: 23.1 1990: 36.7
Spain	1960: 56.6 1990: 75.4	1960: 20.3 1990: 41.5	1960: 26.3 1990: 55.5
Sweden	1960: 72.6 1990: 83.1	1960: 37.9 1990: 80.6	1960: 40.7 1990: 65.9
Switzerland	1960: 51.0 1990: 59.7	1960: 41.0 1990: 60.7	1960: 38.7 1990: 59.8
USA	1960: 70.0 1990: 75.2	1960: 40.6 1990: 66.5	1960: 56.8 1990: 71.2

Sources: World Bank World Development Indicators; ILO LABORSTA EAPEP Data http://laborsta.ilo.org/; Scharpf and Schmidt (2000: 349); United Nations Common Database (UNCDB) http://unstats.un.org/unsd/cdb/cdb_help/cdb_quick_start.asp

[1] One Germany. [2] Different dataset.

Table A3: Values and demographical practices

	% non-religious or atheist	Divorce justifiable[1]	Divorces / 100 marriages	% Extra-marital births	% 25-years old lived in consensual union
Austria	1970: 2.6 1990: 7.2	1990: 4.87	1970: 19.6 1990: 36.0	1960: 13.3 1990: 23.6	1970: 13.4 1990: 60.1
Flanders	1970: 5.9 1990: 7.2	1981: 4.0 1990: 4.90	1970: 8.7 1990: 31.5	1960: 2.1 1990: 11.6	1980: 6.6 1990: 16.9
Estonia	1970: 53.3 1990: 39.8	1990: 5.44	1975: 47.0 1985: 49.0	1960: 13.7 1990: 27.1	1970: 30.5 1985: 66.4
Finland	1970: 3.6 1990: 5.5	1981: 5.81 1990: 7.23	1970: 14.8 1990: 52.6	1960: 4.0 1990: 25.2	1970: 9.9 1985: 70.9
France	1970: 12.0 1990: 19.2	1981: 5.34 1990: 5.65	1970: 9.9 1990: 36.9	1960: 6.1 1990: 30.1	1970: 15.4 1990: 60.1
West-Germany	1970: 3.9 1995: 7.6	1981: 4.91 1990: 5.70	1970: 18.1 1990: 30.0	1960: 6.3 1990: 10.5	43.8
Greece	1970: 0.2 1990: 1.8	1999: 6.26	1970: 5.2 1998: 14.1	1965: 1.1 1990: 2.2	1975: 10.1 1990: 29.0
Hungary	1970: 14.3 1990: 12.7	1981: 4.51	1970: 23.6 1990: 37.5	1960: 5.5 1990: 13.1	1975: 6.8 1985: 18.1
Italy	1970: 11.4 1990: 15.6	1981: 5.07 1990: 5.37	1975: 2.8 1990: 8.7	1960: 2.4 1990: 6.5	1975: 1.4 1990: 4.8
Latvia	1970: 47.6 1990: 35.8	1990: 5.38	1975: 51.0 1990: 49.0	1960: 11.9 1990: 16.9	1975: 17.9 1990: 40.0
Lithuania	1970: 29.2 1990: 14.5	1990: 4.03	1975: 35.0 1990: 55.0	1960: 7.3 1990: 7.0	1975: 6.4 1990: 16.4
Norway	1970: 1.1 1990: 2.2	1981: 4.81 1990: 5.26	1970: 11.7 1990: 46.4	1960: 3.7 1990: 38.6	1975: 10.3 1985: 61.0
Poland	1970: 8.8 1990: 3.1	1990: 3.85	1970: 12.3 1990: 16.6	1960: 4.5 1990: 6.2	1970: 3.6 1985: 4.3
Spain	1970: 2.2 1990: 5.4	1981: 4.69 1990: 5.59	1985: 9.2 1990: 10.5	1970: 1.4 1990: 9.6	1975: 1.9 1990: 9.9
Sweden	1970: 24.7 1990: 29.4	1981: 6.09 1990: 6.33	1970: 29.9 1990: 47.8	1960: 11.3 1990: 47.0	1975: 71.0 1990: 79.0
Switzerland	1970: 1.1 1990: 7.2	1990: 4.77	1970: 13.7 1990: 28.3	1960: 3.9 1990: 6.1	1975: 18.4 1990: 50.9
USA	1970: 4.9 1990: 8.7	1981: 4.66 1990: 4.89	1970: 32.8 1990: 48.4	1980: 18.4 1990: 28.0	1975: 20.0 1990: 67.0

Sources: World Values Survey, waves 1981-1990-1995, weighted cases; European Values Survey, wave 1999, weighted cases; Council of Europe (2002) Recent Demographic Developments in Europe. Council of Europe, Strasbourg; OECD (2002) Social Indicators.
[1] Society mean of a scale from 1 to 10.