A New Look at the Separation Surge in Europe: Contrasting Adult and Child Perspectives

Matthijs Kalmijn[1] & Thomas Leopold[2]

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Abstract
This study contrasts adult and child perspectives on divorce and separation. Based on harmonized retrospective life history data from eight European countries, we study the risk of divorce and separation from the perspective of adult unions and the perspective of children born into these unions. The analysis connects adult and child perspectives, focusing on union cohort changes (1945 to 2005) in the associations between parenthood, education, and (parental) separation. Our findings show that trends differ substantially between adult and child perspectives. First, the cohort surge in divorce and separation is stronger in adults than in children. Second, inequality in the risk of divorce and separation grows faster in children than in adults. For both trends, disparities between adult and child perspectives grow across cohorts, due to increasingly negative associations between parenthood, education, and separation. In several countries, the separation surge has been trivial for children of higher-educated couples.

Keywords
Divorce, separation, demography, second demographic transition, diverging destinies

[1] Netherlands Interdisciplinary Demographic Institute (NIDI) and University of Groningen. Lange Houtstraat 19, 2511 CV The Hague, Netherlands
E-mail: matthijskalmijn@gmail.com

[2] University of Cologne.

The Second Demographic Transition (SDT) has changed the face of families. One intensely studied dimension of the SDT is the increasing instability of married and cohabiting unions, a trend we call the "separation surge." During the past two decades, this trend has raised concerns for research and policy. In several societies, the risk of union dissolution has become increasingly concentrated among less-educated people (De Graaf and Kalmijn 2006; Härkönen and Dronkers 2006; Martin 2006; Musick and Michelmore 2018). McLanahan (2004) argues that this gradient in union instability, along with other socially stratified shifts across the SDT, has intensified social inequality among children—a drift she labels "diverging destinies."

What we know about the separation surge and its social gradient is primarily based on trends found in adults. A focus on the adult perspective is intuitive because the separation surge is driven by changes in adults' norms and values, their preferences, and their decision-making. Explanations have focused on ideational and behavioral shifts, declining gains to specialization, women's increasing economic independence, declining legal and moral barriers to divorce, and, potentially, increasing financial strain on unions formed among the lower-educated (Härkönen 2014). The conversion to children appears intuitive: if adults separate more often, then more children will experience parental separation. Similarly, if adult separation risk is unequal, then the children affected will be predominantly those of less-educated parents.

This seemingly straightforward conversion from adults to children is common in research on divorce and separation. For example, in his overview work on marriage, divorce, and remarriage, Cherlin (1992:26) argues that "one of the most important effects of the rise in divorce . . . is to increase the proportion of parents whose marriage is dissolved while their children are still at home." Similarly, in a widely cited article on education and divorce, Martin (2006:538) argues that "growing class differences in marital dissolution rates would . . . perpetuate inequality across generations, to the extent that marital dissolutions contribute to . . . children's physical and emotional health, educational attainment and performance." Although these analysts recognize the difference between adult and child perspectives, conclusions drawn from one perspective are commonly extended to the other (McLanahan 2004:612).

In the current study, we demonstrate theoretically and empirically that adult and child perspectives can diverge to an extent that conclusions about the separation surge need reconsideration. What we labeled "a new look" in the title refers to the insight we obtain about the separation surge in particular, and the SDT more generally, when we contrast adult and child perspectives. We draw different conclusions about the scope of the cohort surge in separation and about social inequality in the risk of separation. This insight is important, as it is now widely accepted that separation entails negative consequences for adults and children in several domains, including well-being, mental health, school and career outcomes, income, and social relationships (Amato 2000; Andreß et al. 2006; Fomby and Cherlin 2007; Härkönen, Bernardi, and Boertien 2017; Leopold and Kalmijn 2016; Williams and Dunne-Bryant 2006).

Why would adult and child perspectives on union dissolution differ? First, only about half of all separations involve children (OECD 2015). If the separation surge was concentrated in childless couples, rather than in parents, then the trend viewed from the child perspective would be less pronounced than the trend found in population data on separation rates in adults. Conversely, if the separation surge was concentrated in couples with children, then it would be even stronger when viewed from the child perspective. Second, conclusions about social inequality also depend on whether the education gradient in separation is concentrated in adults with fewer or more children. If separation risk was stratified only in childless couples, then we would not observe diverging destinies from the child perspective. Conversely, if the education gradient in separation risk was concentrated in couples with more children, then the separation experience would be (even) more stratified when viewed from the child perspective.

These considerations demonstrate that although adult and child perspectives on separation are linked, they cannot be directly transposed into one another. Kennedy and Thomson (2010) addressed the child perspective, showing that in Sweden, children of higher-status parents experienced a smaller increase in family instability than did children of lower-status parents. This finding calls for a systematic comparison between adult and child perspectives for a larger number of countries. This is the aim of our analysis. In analyzing these two perspectives, we do not intend to compare how adults and children differentially perceive a separation, or how they are differentially affected by a separation. Instead, we compare adults and children with respect to trends and social disparities in the *risk* of separation.

We proceed in two steps. First, we examine changes in the risk of separation from the adult perspective. Our focus in these analyses is on the association between parenthood and the risk of separation. This association is the key link between the adult and child perspectives on separation because it determines the extent to which the separation surge has affected children. Moreover, if the association between parenthood and separation changed across cohorts and if it is socially stratified, this determines the extent to which the separation surge has resulted in diverging destinies among children. Second, we examine changes in the risk of separation from the child perspective, covering all children born in the unions selected in the first step of the analysis. We combine adults and children into one analytic framework for a direct statistical comparison of the two perspectives. Addressing the methodological challenges in comparing the two perspectives, we develop models showing increasing divergence between adult and child perspectives, especially when considering the social stratification of family instability.

The context of our study is Europe. In Northern, Western, and Eastern Europe, the separation surge has been ubiquitous. Northern European countries showed the earliest increases in separation rates, followed by Western European countries such as France and Belgium. Unions have remained relatively stable in Southern countries such as Italy and Spain (Andersson, Thomson, and Duntava 2017). In most European countries, separation rates plateaued in the late 1980s and early 1990s, and in some cases declined thereafter (Härkönen 2014). A special feature of the European context is that cohabitation—before and instead of marriage—is more common and more accepted than in the United States (Kiernan 2002). The rise of cohabitation may have undermined union stability, but the causal link between the separation surge and the rise of cohabitation is debated (Perelli-Harris et al. 2017; Sassler and Lichter 2020). We include cohabiting unions in our work and explore the role of cohabitation in the trends we observe.

We analyze data from eight European countries in which a separation surge occurred: Belgium, the Czech Republic (Czechia), Estonia, France, Hungary, Norway, Sweden, and the United Kingdom. Given the limited number of countries, our comparative goal is limited to studying one idea—the contrast between adult and child perspectives—in multiple national contexts. We use state-of-the-art life-history data (Vikat et al. 2007) that offer complete union and fertility histories for large representative samples and have been harmonized by a team of European and U.S. demographers (Perelli-Harris, Kreyenfeld, and Kubisch 2011). These data allow us to cover unions formed between 1945 and 2005, facilitating a comprehensive cohort design. Due to the large samples, the data offer sufficient statistical power to analyze stratified trends. In the eight countries combined, our data include about 75,000 marital and cohabiting unions, 100,000 children born in these unions, and 20,000 separations.

BACKGROUND

Although no work has systematically compared or connected the adult perspective to the child perspective on separation, related research in the fields of demography, family sociology, social stratification, and psychology is relevant to studying the linkage between these perspectives. To understand how adult and child perspectives are linked, we consider (1) the association between parenthood and separation, (2) the association between education and separation, and (3) the interaction effect of parenthood and education on separation. To understand if and how *trends* differ between adult and child perspectives, we additionally consider cohort change in these three associations. Finally, we discuss expectations about the role that cohabitation and unmarried parenthood played in the separation surge.

Parenthood and Separation

The first line of related research is on the association between parenthood and separation. A long history of research shows that this association is negative: parents are less likely than childless couples to separate (Diekmann and Schmidheiny 2004; Van Zanten and Van den Brink 1938; Waite and Lillard 1991).

This association has been interpreted in terms of investment, protection, and selection. According to the investment hypothesis, children protect parents from the risk of separation. Children are seen as union-specific capital that locks partners into their union, even if the union is unhappy (Kalmijn 1999). This applies to fathers who risk a decline in or even loss of contact with their children after separation. The mechanism also applies to mothers who specialized in domestic labor and thus increased their economic exit costs (Becker 1991). According to the protection hypothesis, unhappy couples are aware of the negative effects of separation on children, and decide, for the sake of their children, to work out their differences. Such couples may either postpone separation or improve their relationship (Klein and Rapp 2010). The selection hypothesis is based on a reverse causal argument: couples who are uncertain about their relationship are less likely to decide to have children (Lillard and Waite 1993).

Despite evidence for each hypothesis, it remains difficult to adjudicate between these different interpretations for the negative association between parenthood and separation. For our purposes, the association itself is more relevant than the underlying mechanisms or causal direction. Considering our linkage of interest, a negative association between parenthood and separation means separation is concentrated in unions without or with fewer children. Therefore, we expect separation is more common when viewed from the adult perspective than when viewed from the child perspective *(Hypothesis 1)*.

Education and Separation

The second line of related research is on the association between education and separation. Studies show that this association is negative in several Western countries (Härkönen and Dronkers 2006; Martin 2006) The negative education gradient in separation is a demographic backdrop to studies motivated by the diverging destinies hypothesis (Bernardi and Radl 2014). Explanations for why lower-educated people are more likely to separate include poorer marital matches (Herrnstein and Murray 1994), higher economic strain that spills over into marital life (Goode 1962), and their reduced capacity to cope with and successfully overcome marital conflict (Conger, Conger, and Martin 2010). Similar to research on the mechanisms behind the negative association between parenthood and separation, research on the mechanisms behind the negative association between education and separation is limited (for an exception, see Boertien and Härkönen 2018).

In the absence of further differences, an adult education gradient in separation translates into a parental education gradient when seen from the child perspective. However, if the relationship between education and separation depends on whether a couple has children, adult and child perspectives will diverge. If the negative association between education and separation is larger for parents than for childless couples, the educational gradient will be even larger from the perspective of children. If the association is smaller or absent for parents, the gradient will be smaller or absent for children. To our knowledge, there is no research on this interaction effect. Yet, related work is suggestive about its nature and direction.

Two arguments suggest the negative association between education and separation is stronger in the presence of children. First, the motive and capacity to shield children from negative experiences in various domains of life could be stronger among higher-educated couples, and this includes protection of their own union in the interest of their children. Research shows that higher-educated people know more about the causes of (mental) health problems (Cutler and Lleras-Muney 2008) and that higher-educated mothers know more about their children's emotional development (Huang et al. 2005; Marjanovič-Umek and Fekonja-Peklaj 2017). Such differences may result in better consideration of children's interests when deciding about a break-up. Education differences in knowledge and awareness may result in a stronger negative association between parenthood and separation among higher-educated couples.

Second, higher-educated people may be more cautious about having children if they are uncertain about their relationship. Such an effect is plausible given the classic observation that more unplanned births occur among lower-educated than higher-educated women (Musick et al. 2009; Wellings et al. 2013). This means the selection hypothesis stipulated above would apply more strongly to higher-educated than to lower-educated people.

If the negative association between education and separation is stronger in the presence of children, as these arguments suggest, then this has an obvious implication for our linkage of interest. From the adult perspective, the *overall* education gradient in separation is a combination of the weaker (or absent) gradient in childless couples and the stronger gradient in couples with children. From the child perspective, the gradient by parental education is *only* determined by the stronger gradient in parent couples. Accordingly, we expect the education gradient in separation to be stronger when viewed from the child perspective than when viewed from the adult perspective *(Hypothesis 2)*.

Cohort and Parenthood

The overall expectation of an increasing separation risk is clear not only from a cohort perspective on adult unions but also from a cohort perspective on the children of these unions. Studies show that children's experiences of single parenthood, step-parenthood, and other outcomes of separation have increased across cohorts (Thomson 2014). However, we expect the negative association between parenthood and separation will reduce the separation surge when viewed from the child perspective. If the association between parenthood and separation is constant across union cohorts, then the suppression effect will be similar across earlier and later stages of the separation surge. If the association changes, then the differences between adult and child perspectives on separation become a function of cohort.

Two studies have examined cohort changes in the association between parenthood and separation. Using retrospective survey data on Norway—the same we use in our analysis—Hart, Lyngstad, and Vinberg (2017) used a set of dummy variables based on children's age and number and interacted these with a series of dummy variables for period (not cohort). They found no evidence for change in the association between parenthood and separation. Andersson (1997) used register data on marriages in Sweden and found parallel period trends in separation risks for married couples with and without children. Unfortunately, the data did not include cohabiting couples.

Theoretically, there are still good reasons to expect a change in the association between parenthood and separation, as also argued by Hart and colleagues (2017:319). The investment hypothesis implies that the association between parenthood and separation has become less negative because the separation-related loss of investments in children as marital-specific capital has declined across cohorts. Most notably, the role of fathers after separation has changed radically. After separation, fathers more often see their children, are more often part of a co-parenting arrangement, and their involvement in their children's lives after separation has become normalized (Westphal, Poortman, and van der Lippe 2014). Conversely, separated mothers have become less stigmatized, facilitating family relationships after separation. As a result, less marital-specific capital is lost after separation, suggesting it has become "easier" for parents to separate. The negative association between parenthood and separation would then weaken across cohorts and, as a result, the adult and child perspectives would converge across cohorts *(Hypothesis 3a)*.

The protection and selection hypotheses, in contrast, suggest a different trend. As separation has become more common, knowledge about the causes and consequences of separation has expanded, along with the availability of preventive services, such as relationship counseling, family therapy, and psychotherapy (Olfson et al. 2002). Applied to the protection hypothesis, this suggests parents can more effectively shield their children from the risk of separation. For the same reason, childless couples may deliberate more carefully about having children when their union is uncertain, so the selection effect would become more salient. These arguments suggest the negative association between children and the risk of separation would intensify across cohorts. If this hypothesis is supported, then adult and child perspectives will increasingly diverge, resulting in a less pronounced separation surge when viewed from the child perspective *(Hypothesis 3b)*.

Cohort and Education

In several European countries (including Belgium, France, Hungary, and Sweden), the initially positive association between education and divorce has reversed across cohorts, similar to what was found for the United States (Garriga and Cortina 2017; De Graaf and Kalmijn 2006; Härkönen and Dronkers 2006; Martin 2006; Matysiak, Styrc, and Vignoli 2014). We note that the empirical picture is not consistent across Europe. In some countries, the education gradient in separation risk is absent or has not reversed across cohorts (Matysiak et al. 2014).

The cohort trend in education-specific separation risks supports theoretical accounts of the higher-educated as spearheads of social change who overcame legal and normative barriers to separation earlier than the lower-educated (Lesthaeghe 2014). As barriers vanish across cohorts, lower-educated couples' higher economic strain will increasingly translate into higher separation risk (Goode 1962). An alternative but related explanation is that the group of lower-educated people is shrinking across cohorts and increasingly represents a negative selection in terms of marital matches, marital strain, cognitive ability, and coping resources (Gesthuizen, De Graaf, and Kraaykamp 2005).

Cohort, Parenthood, and Education

An increasingly negative education gradient in separation can affect the linkage between adult and child perspectives if fertility differs by education. Apart from such compositional effects, the linkage is also affected if parenthood interacts with cohort change in the education gradient. A closer look at the protection and selection hypotheses suggests a three-way interaction of cohort, parenthood, and education may be negative. Research shows that education promotes problem-solving capacities in various areas, including family conflict related to marital problems and children. Higher-educated people are more confident in their ability to solve problems, show more agency and sense of personal control, are less vulnerable to stress, more often draw on professional support, and can more effectively translate professional support into behavioral changes (Ross and Mirowsky 1999). This means changing motives and capacities related to the mechanisms of protection and selection may have promoted the protective role of (prospective) parents, particularly among later cohorts of higher-educated people. These considerations suggest the negative association between parenthood and separation has intensified across cohorts, especially among the higher-educated. If this is true, then the adult–child contrast in the separation surge will be stronger for higher-educated than for lower-educated individuals *(Hypothesis 4)*.

Cohabitation

Cohabitation is now an integral part of the life course in Europe (Perelli-Harris et al. 2017; Sassler and Lichter 2020), but its meaning is heterogeneous. Some couples view cohabitation as a common and socially accepted means of testing a relationship; for others, cohabitation serves an alternative to marriage that can be legally recognized as equivalent (Hiekel and Keizer 2015; Perelli-Harris et al. 2014). Because cohabiting unions are less stable than married unions (Liefbroer and Dourleijn 2006), part of the increase in union instability is linked to the rise of cohabitation. Competing theoretical perspectives exist on this link. A supply-side perspective argues that the option of cohabitation became available as a result of exogenous ideological changes such as secularization, and the availability of this option in itself contributed to union instability (Waite and Gallagher 2000). A demand-side perspective holds that couple relationships eroded first—for both cultural and economic reasons—and the demand for cohabitation increased as a consequence (Giddens 1992; Mills and Blossfeld 2005; Perelli-Harris et al. 2017). From a supply-side perspective, cohabitation is a pertinent predictor in models of separation; from a demand-side perspective, cohabitation is an endogenous covariate that cannot be meaningfully interpreted as a predictor of separation.

The rise in cohabitation was stratified. In Western European countries, the link between cohabitation and education seems to have reversed, changing from a positive to a negative gradient (Bhrolcháin and Beaujouan 2013). In several Eastern European countries, however, the rise of cohabitation seems to be a continuation of the traditional pattern of a “poor man’s marriage” (Sobotka 2008). Fertility in cohabiting unions has increased over time in virtually all Western countries (see Table 1). Nonmarital childbearing has been more common among the lower-educated (Perelli-Harris et al. 2010; Sobotka 2008). Children’s risk of experiencing parental separation is higher when they are born in a cohabiting union (Andersson et al. 2017; Manning, Smock, and Majumdar 2004). Many unmarried parents marry after having children, and when cohabiting parents marry, their unions become as stable as couples who married before parenthood (Musick and Michelmore 2018; Perelli-Harris et al. 2012). The meaning of having children outside of marriage is apparently heterogeneous. Some couples jointly decide on marriage and parenthood without marriage necessarily preceding parenthood, whereas other unmarried couples may have children despite the fact that their union is not yet secure (Carlson, McLanahan, and England 2004).

Because cohabitation is heterogeneous and endogenous, (interaction) effects of cohabitation on separation must be interpreted with caution. We consider four expectations suggested in previous work: (1) Assuming cohabitation is an indicator for commitment to a union, we expect a positive association between cohabitation and the risk of separation (Brown and Booth 1996). (2) Assuming cohabitation has become normalized and less socially selective, we expect the association between cohabitation and separation has declined across cohorts (Liefbroer and Dourleijn 2006). (3) Assuming cohabiting parents have fewer resources and receive less support from their social network, we expect children to reduce the risk of separation less in cohabitation than in marriage (Manning et al. 2004; Waite and Gallagher 2000). (4) In higher-educated couples, marriage is considered a stronger symbol of achieved status, and cohabitation is more often seen as a trial marriage (Cherlin 2009). Therefore, we expect the association between cohabitation and separation is stronger for higher-educated than for lower-educated couples.

The European Context

We analyze eight European countries in which a separation surge took place. Table 1 summarizes key demographic, social, and economic indicators for these countries. Current separation rates are higher in Estonia, Sweden, and the United Kingdom, and lower in France and Belgium. Cohabitation is more common in the Northern countries than elsewhere, but unmarried parenthood is common in all countries. Demographic differences between the eight European countries are not large when compared to the United States (presented at the bottom of Table 1). All the European countries have considerably lower marriage and divorce rates than does the United States, as well as substantially higher cohabitation rates.

<Table 1 about here>

Differences in social and economic indicators between the European countries are large. Our study countries include poorer nations with weaker welfare states (e.g., the Czech Republic) and richer nations with extensive welfare states (e.g., Sweden). The countries also differ in terms of inequality. We look at the income ratio of the top versus bottom 10 percent of the income distribution, commonly used to describe inequality at the macro-level. Given that we use education as a key measure of stratification, we also look at earnings differences between education groups. Economic returns to schooling are one dimension of educational inequality in society (Hout 2012). Educational inequality and income inequality are lowest in Norway and Sweden. Income inequality is highest in the United Kingdom, and educational inequality is highest in former socialist societies.

Given the limited number of countries, we cannot systematically examine how macro-level mechanisms affect adult and child perspectives on separation. Our comparative goal is therefore limited to studying one specific idea—the contrast between adult and child perspectives—in multiple national contexts. This enhances the scope of our study in terms of population coverage and allows us to assess whether the main patterns we found are specific to a few countries or general across several countries. If patterns are general despite contextual differences, this may suggest the underlying mechanisms are also of a general nature.

DATA AND METHOD

We analyze data from the Harmonized Histories Data (HHD) file (Perelli-Harris et al. 2011; Vikat et al. 2007).[1] The HHD harmonizes childbearing and marital histories from 14 countries in the Generations and Gender Programme (GGP) with data from Spain (Spanish Fertility Survey), the United Kingdom (British Household Panel Study), and the United States (National Survey for Family Growth). All datasets offer large samples (approximately 10,000 respondents each) and elaborate retrospective information on partnerships and fertility. The data include men and women but no information on former spouses' characteristics. We considered using register data but these data were limited in their measures of education and cohabitation.

We selected countries using the following inclusion criteria: (a) the occurrence of a separation surge, (b) a sufficient number of separations or separations in the data (more than 1,000), (c) a broad range of union cohorts covering the SDT, (d) the availability of retrospective cohabitation, marriage, and fertility histories, (e) the presence of both sexes in the data, and (f) an absence of quality concerns about the data. Applying these criteria led to selection of eight European countries: Belgium, the Czech Republic, Estonia, France, Hungary, Norway, Sweden, and the United Kingdom. We dropped Germany because of concerns about the measurement of fertility histories (Kreyenfeld, Hornung, and Kubisch 2013).

We use information on all married and unmarried cohabiting unions. With the term separation, we refer to the breakup of both union types. We exclude very short cohabiting unions from the data as these could reflect non-committed trial unions, and their inclusion could lead to an artificial increase in the separation risk of nonparents. Specifically, we exclude cohabiting unions that began and ended in the same calendar year and unions that ended in the calendar year after (5.1 percent of all unions). We further exclude respondents who entered a union after age 50 or before age 12; respondents with missing or impossible information on dates (the year of birth, the year of union, the year of a child's birth, or the year of separation); respondents who entered a union before 1945, after 2005, or in the year of the survey;

and respondents with missing or uncodable education. These exclusions leave us with 75,461 unions and 21,418 dissolutions (see Table 2 for country details).

<Table 2 about here>

Measures

Union cohort is the central time variable in all analyses. Because of the wide cohort range in the HHD, we are able to examine unions formed over six decades, between 1945 and 2005. We divided the cohort variable by 10 and centered the variables (within each country) so that the main effects apply to the average union cohort (approximately 1975). We include a quadratic cohort term to capture the slowdown of the separation surge. For descriptive purposes, we divide cohorts into five-year groups. Table 2 presents the means of all variables by country along with descriptive information about the separations in our data.

We use *education* as an indicator of social position. Income is only available at the time of the survey, and occupational status is not an ideal measure of status at the start of a union due to nonemployment and career entry processes. Because education is usually completed before union and family formation, it is used as the central inequality measure in most demographic analyses (Härkönen and Dronkers 2006; Kravdal and Rindfuss 2008; Martin 2006; Matysiak et al. 2014). Education is also a classic indicator of inequality in studies of the consequences of divorce for children (Astone and McLanahan 1991; Sigle-Rushton, Hobcraft, and Kiernan 2005). Our education measure pertains to the respondent's education. Data on (ex-)partners' education are too incomplete to be useful. Education was coded into the International Standard Classification of Education (ISCED), which is comparable across countries. The ISCED categories are ranked but not interval-scaled. Because we estimate two-way and three-way interactions, we constructed a linear measure of education by recoding ISCED categories into the length of schooling (Schneider 2011; Schröder and Ganzeboom 2014). The GGS data provide information on the age at leaving school. This is not a useful individual measure, but the *average age* at leaving school (minus 6) within an educational category is a good proxy for the length of schooling. We coded ISCED 1 and 2 as 9.1 years, ISCED 3 as 12.7 years, ISCED 4 as 15 years, and ISCED 5 and 6 as 17.1 years.

In supplementary analyses (see Tables S1a, S1b, and S1c in the online supplement), we examined three alternative measures of education: (a) a dichotomous measure of education, contrasting ISCED 1, 2, and 3 to ISCED 4, 5, and 6; (b) a linear measure of education that uses country-specific conversions to years of schooling (for countries where this is possible); and (c) a linear measure of education that converts detailed ISCED categories into ISLED scores, a recently suggested metric for comparative research on education (Schröder and Ganzeboom 2014).[2] The results for all interactions are robust to using these alternative scalings. Because of the complexity of our models, in which education is entered in several two-way interactions and one three-way interaction, we present results for a linear instead of a categorical measure.

The surveys provide detailed *fertility* histories. We created a time-varying variable coded 1 if a (first) child was born during the union or one year before the union began (to allow for pre-union births). This variable captures the effect of parenthood in a couple regardless of their number of children. We also explored a linear variable representing the number of children in a union, but we found no consistent effects beyond parenthood.

We capture *cohabitation* by a time-varying variable (0 married union, 1 unmarried cohabiting union) and by its interactions with parenthood, education, and cohort. For the child perspective, we additionally use a time-constant variable for whether a child was born in a cohabiting union.

Models for the Adult Perspective

We created an adult-year file starting with the year of union formation and ending in the year of separation or the year of interview if the union was intact at the time of the survey. We include all unions for each respondent. Unions ending due to widowhood are censored in that year. We estimate the risk of separation using discrete-time event-history models (Yamaguchi 1991). By using event-history models, we are able to adjust for the time dependency in the risk of dissolution as well as for the fact that we cannot observe the unions in all cohorts for the same amount of time (right censoring).

We model *process time* using a quadratic specification (union time and union time squared) to allow for the nonlinear pattern of duration-dependency. A single-year specification provides a better fit to the data but is less parsimonious. Diagnostic plots show that our parametrization matches non-parametric year-to-year risks reasonably well. Separation risks first increase for a union but quickly decline. This temporal shape mainly reflects increasing selection toward stable relationships in the pool of surviving unions.

Table 3 presents event-history models that add interactions step-by-step. The full model is defined as follows:

$$\ln (S_i / (1- S_i)) = \beta_1 + \beta_2 T_i + \beta_3 T_i^2 + \beta_4 K_i + \beta_5 (C_i - \overline{C}_i) + \beta_6 (E_i - \overline{E}_i) + \beta_7 K_i (E_i - \overline{E}_i) + \beta_8 K_i (C_i - \overline{C}_i) + \beta_9 K_i (C_i - \overline{C}_i) (E_i - \overline{E}_i) + \varepsilon_i$$

where K_i is a time-varying variable indicating whether adults have children (1 for adult-years with children, 0 otherwise), T_i is *union time*, C_i is *union* *cohort*, and E_i is *education*. We expect parenthood to be negatively associated with separation (Hypothesis 1; H_1: $\beta_4 < 0$), that this association is stronger for higher-educated than for lower-educated individuals (Hypothesis 2; H_2: $\beta_7 < 0$), that the separation surge was weaker for couples with children (Hypothesis 3; H_3: $\beta_8 < 0$), and that this last difference was more pronounced for the higher-educated (Hypothesis 4; H_4: $\beta_9 < 0$). Table 3 presents the estimates (Model 3).

To ease the interpretation of interaction effects in nonlinear models (Mize 2019), we present all our results in terms of average marginal effects. Specifically, we use Model 2 to calculate the annual separation risks (i.e., predicted probabilities) for each union cohort, for higher and lower-educated adults, and for adults with and without children. Figure 2 presents these probabilities and their confidence intervals. We evaluated the robustness of our findings using linear probability models that yield similar findings (see Table S2 in the online supplement).

We corrected the standard errors for the clustering of unions within persons. No correction is needed for the clustering of years within unions (Allison 1984; Rabe-Hesketh and Skrondal 2005; Singer and Willett 1993). We centered all variables to define the main effects of education and parenthood for the mean values of the interacting variables. To keep the presentation parsimonious, we estimate interactions with the linear cohort variable.

Comparing Adults' and Children's Risks

To compare adult and child perspectives, we first constructed a person-period file including all children born to our sample of adult unions (N = 109,360). We expanded these data to a child-year file starting with the child's year of birth and ending with (a) the year of parental separation (event), (b) the year in which a parent died (censored), (c) the year of the survey (censored), or (d) the year in which a child turned 18 (censored).[3] We appended the child-year file to the adult-year file described earlier and estimated models for the pooled data to compare the two perspectives.

A demographic comparison between adult and child perspectives on separation risk poses several challenges. First, the *historical time scale* differs between the perspectives because the year of union formation is usually not the year in which a child is born to a union. In practice, this discrepancy is minor because the correspondence between adults' union cohort and children's birth cohort is close (r = .96). On average, a parent's union cohort preceded a child's birth cohort by 4.5 years (e.g., a union formed in 1980 would correspond to a child's birth cohort of 1984). In our analysis, we use union cohort as a common historical time scale for comparing adult and child perspectives.

Second, *process time* differs between the perspectives. For children, our interest is in experiencing parental separation before reaching adulthood, that is, before age 18. For adult unions, process time can, in principle, be longer but should align with the childrearing period. Hence, we follow adults until the moment the average child turns 18, which is about 23 years in the union. In supplementary analyses (see Tables S3a, S3b, and S3c in the online supplement), we assessed whether results are sensitive to using different observation windows. All interactions of interest are insensitive to the choice of the time window. Adult separation rates tended to increase at very short durations. This trend reflects an increasing sample selectivity toward childless couples at shorter durations.

Third, a demographic comparison of the two perspectives requires linking *samples* of adults and children. Prospective studies of intergenerational mobility show the difficulties of linking a sample of children to identifiable and representative cohorts of adults as potential parents of those children (Skopek and Leopold 2020). The GGS data allow us to solve this issue by selecting adult samples representative of adult unions in each country and obtaining child samples drawn from these reference samples of adult unions. Because all children born to adult unions were reported, we could obtain the child samples as indirect probability samples representative of children born to our study cohorts of adult unions. Note that the child samples are not representative of children born outside of unions.

Fourth, the models estimated for adults and children require comparable *outcome measures.* Conditional *annual* probabilities of experiencing (parental) separation, as estimated by our models, are directly comparable between adults and children. Substantively, however, this measure is less meaningful than the cumulative probability of experiencing (parental) separation up to a certain duration. To solve this issue, we use life tables translating annual separation probabilities into cumulative probabilities as an additional outcome measure. To ensure comparability between adult and child life tables, the *length* of the risk period should be the same. We define this period as 18 years, in line with our focus on children's risk of experiencing separation before adulthood. For adults, we calculate cumulative separation probabilities for the first 18 years of the union; for children, our risk period of 18 years starts at the median union year in which a child was born (approximately five years after union formation). Additional analyses show this average is consistent across study countries.[4]

To analyze the pooled data, we estimate event-history models for annual (parental) separation risk (Table 5). We created a variable called *child perspective* P_i, coded 0 for adults and 1 for children. The other independent variables are identical in the two datasets. The model is defined in the same way as described earlier:

$$\ln (S_i / (1- S_i)) = \quad \beta_1 + \beta_2 T_i + \beta_3 T_i^2 + \beta_4 P_i + \beta_5 (C_i - \overline{C}_i) + \beta_6 (E_i - \overline{E}_i) +$$

$$\beta_7 P_i (E_i - \overline{E}_i) + \beta_8 P_i (C_i - \overline{C}_i) + \beta_9 P_i (C_i - \overline{C}_i) (E_i - \overline{E}_i) + \varepsilon_i.$$

The coefficient for perspective tests Hypothesis 1 (H_1: $\beta_4 < 0$), the interaction between education and perspective tests Hypothesis 2 (H_2: $\beta_7 < 0$), the interaction between cohort and perspective tests Hypothesis 3 (H_3: $\beta_8 < 0$), and the three-way interaction between perspective, cohort, and education tests Hypothesis 4 (H_4: $\beta_9 < 0$).

For ease of interpretation, we again present average marginal effects. Figure 3 presents the predicted annual separation probabilities of Model 3. Figure 4 presents cumulative probabilities of separation for adult unions and cumulative probabilities of parental separation for children. These probabilities are calculated from life tables based on annual separation risks estimated by the event-history model. We present cumulative probabilities of experiencing (parental) separation up to a duration of 18 years calculated for two cohorts: 1950 and 1995.

The pooled dataset has a complex structure (see Table S4 in the online supplement). Adults could have multiple unions and unions could have multiple children. We corrected the standard errors for the *multiway* clustering of (a) unions in adults/parents and (b) the clustering of both children and adults in unions using the *vcemway* module in Stata (Gu and Yoo 2019). We centered all variables within countries.

Including Cohabitation

In the analyses of adult separation risks, we use cohabitation as a time-varying variable and add interactions of cohabitation with cohort, education, and parenthood to Models 2 and 3. We test for cohabitation interactions (Table 4) and we assess the extent to which our main results change when the cohabitation interactions are included in the models (Table 3 versus Table 4). The analyses of children's risk of parental separation allow us to treat cohabitation as a static variable, indicating whether children were born in a married union or in a cohabiting union (cf. Manning et al. 2004). We examine the interaction of education and cohort in a sample of all children and in a subsample of children born to married parents to assess if parental separation risk also diverges by education if the sample is limited to children born to married parents (Table 6).

FINDINGS

Figure 1 shows how annual separation probabilities of adult unions changed across cohorts. These are predicted probabilities obtained from an event-history model evaluated at the mean union time. The numbers are similar to the well-known net divorce rate, which is the annual number of divorces per 1,000 married couples. Our estimates differ from these rates in the sense that they are expressed as probabilities and apply to marriages and cohabitations.

Figure 1 shows a strong increase in separation probability in all countries, consistent with other research. Note, however, that the increase did not level off in the late 1980s and the 1990s, as has been documented for a number of countries (Härkönen 2014). This difference most likely reflects the fact that our data include cohabiting unions. Because cohabiting unions are increasingly common and are less stable, this has resulted in a continuous increase of overall union instability.

<Figure 1 about here>

Analyses of Adult Separation Risks

Our baseline model for adults contains the three main variables of interest without cohort interactions (Model 1, Table 3). In all countries, we found a strong negative association between parenthood and the risk of separation. This association is often substantial in magnitude but it varies across countries. In three countries, the association is modest (the Czech Republic, Estonia, and the United Kingdom), with odds ratios between 1.25 and 1.35. In the other five countries, the association is stronger. In these countries, annual separation odds of nonparents were 1.98 to 2.8 times the odds of parents. Note that our interest in this association concerns only the relation between adult and child perspectives, not the causal effect of parenthood on separation.

<Table 3 about here>

In the absence of cohort interactions, we see weak negative main effects of education on the risk of separation in most countries. The interaction between education and parenthood is negative in seven of the eight countries, indicating that the gap in separation risk between parents and nonparents grows with education. These findings provide initial support for Hypothesis 2. Although we can only speculate about underlying mechanisms (e.g., education promoting selection, protection, and fathers' investment), we note that the interaction is strong and present in many countries. Moreover, this interaction has implications for "diverging destinies" observed from a child perspective on parental separation.

In Model 2 of Table 3, we examine cohort change in these effects. In seven of the eight countries, the interaction between parenthood and cohort is negative, and in five of these countries, this interaction is statistically significant. A negative interaction implies that the separation surge was stronger for childless unions than for parents, inconsistent with Hypothesis 3a and consistent with Hypothesis 3b. In two countries (Estonia, United Kingdom), however, the interaction is absent. The second cohort interaction of interest is the education gradient in separation. Interactions between education and union cohort are negative and significant in all countries except the United Kingdom (and marginally significant in Sweden). Given the numerical range of the cohort variable, the results imply that the education gradient in separation changed from positive in early cohorts to negative in later cohorts.[5] The reversal of the education gradient in separation risk confirms earlier findings (Härkönen and Dronkers 2006). The conventional interpretation of this trend is that initially, the higher-educated overcame legal and normative barriers to separation, and financial strain or psychological coping skills started to matter more for union stability after these barriers had vanished (De Graaf and Kalmijn 2006; Härkönen and Dronkers 2006).

The most complex aspect of our model is the three-way interaction between education, parenthood, and cohort, which is included in Model 3 of Table 3. The estimates show negative three-way interactions in seven of the eight countries and significantly negative interactions in six countries. Substantively, this means the negative education gradient in the association between parenthood and separation intensified across cohorts, supporting Hypothesis 4. Possible interpretations for this change are that higher-educated couples have become increasingly cautious in deciding whether to have children (selection), or higher-educated couples have become increasingly aware of the negative consequences of separation for children (protection).

Figure 2 presents average marginal effects. The figure is based on the predictive margins of Model 3 in Table 3 and depicts trends in separation risks for different subgroups evaluated at the mean union time. For each country, the left-hand graph pertains to lower educated couples (nine years of education) and the right-hand graph to higher educated couples (17 years of education). To evaluate differences between adult and child perspectives, we present cohort trends in the probabilities for childless couples (blue circles) and parents (purple triangles). Quadratic main effects of cohort are included.

<Figure 2 about here>

We see a stronger separation surge in childless couples than in parent couples. Moreover, these differences are strongly stratified by education. A comparison of childless couples (represented by the blue circles) indicates a similar separation surge for the lower- and higher-educated. A comparison of parent couples (represented by the purple/dashed lines) shows a clear divergence by education, with a steeper trend for the lower-educated. In several countries across Europe (Belgium, Hungary, Norway, Sweden), the separation surge has not only been less pronounced for higher-educated parents, but it is almost absent. In sharp contrast, separation probabilities for lower-educated parents have surged in all countries, with increases that resemble and sometimes exceed those found among childless couples.

The Role of Cohabitation

To assess the role of cohabitation in the separation surge, we re-estimated our models with added cohabitation variables and their interactions (Table 4). All variables are mean-centered to ease the interpretation of main effects. In line with expectations—and all previous studies—we found positive associations between cohabitation and separation in all countries. The negative parenthood effect weakens when cohabitation is controlled (compare Tables 3 and 4), suggesting parenthood and marriage are closely connected decisions. We also found a consistent pattern for interactions between cohabitation and education. In all countries, the gap in separation risk between married and cohabiting couples widened with education, suggesting cohabitation is more often used as a trial union by the higher-educated.

<Table 4 about here>

For the remaining expectations, evidence is less consistent. The expectation that cohabiting unions have become more stable vis-à-vis married couples is supported in three of our eight countries, as demonstrated by the negative interactions of cohort and cohabitation in Belgium, Estonia, and the United Kingdom. The expectation that children are less protective of cohabiting unions is supported in four countries (Estonia, France, Hungary, and the United Kingdom), but significant findings in the opposite direction emerged in Belgium and Sweden. Overall, our data are broadly in line with pertinent hypotheses advanced by previous research on cohabitation, but the evidence is not consistent across countries.

How do our main interactions of interest change when cohabitation variables are added? All of the two-way interactions between education and cohort remain significantly negative. Similarly, we still observe significantly negative two-way interactions between parenthood and cohort in four countries. The three-way interactions slightly intensify in Belgium, France, and Sweden, and they weaken in Estonia, Hungary, and Norway (losing statistical significance in the latter two countries). These findings suggest that in some countries, the rise of cohabitation coincided with changes in the joint influence of education and parenthood on separation risks.

Adding the Child Perspective

In the models shown in Table 5, we pool adults and children to examine differences between the two perspectives. These results are the first demographic estimates directly comparing adult and child perspectives on separation. In each country, the indicator variable for the child perspective shows a negative effect (Model 1, Table 5). The negative association between parenthood and separation in our analyses of adults translates into a lower separation risk when viewed from the child perspective, in line with Hypothesis 1.

<Table 5 about here>

Model 1 in the top panel of Table 5 shows that the two-way interaction between education and child perspective is significant and negative in six countries. Because the main effect of education is negative, this means the risk of parental separation, viewed from the child perspective, is *more* stratified than the risk of separation viewed from the adult perspective. This evidence from the pooled adult–child dataset corroborates our earlier evidence in support of Hypothesis 2.

How has the contrast between adult and child perspectives changed across cohorts? The interaction between child perspective and cohort is negative and significant in seven of the eight countries (Model 1). In Belgium, the Czech Republic, France, Hungary, Norway, Sweden, and the United Kingdom, the separation surge has been stronger from the adult perspective than from the child perspective. This finding supports Hypothesis 3b, showing that adult and child perspectives increasingly diverge across cohorts.

Model 2 of Table 5 presents results for the three-way interaction between cohort, education, and child perspective. In six of the eight countries, we see negative and significant three-way interactions (including one marginally significant effect). In line with our previous results, adult and child perspectives differ most when we stratify the analysis by education. The three-way interaction found here parallels and validates the three-way interaction found in the adult sample between cohort, education, and parenthood (Hypothesis 4).

Figure 3 illustrates the findings of this last model. In each subgraph, we compare adults and children broken down by education, similar to Figure 2. The left-hand graphs pertain to (children of) lower-educated people, the right-hand graphs to (children of) higher-educated people. The graphs in Figure 3 are in line with those from Figure 2 and provide a further illustration of our main findings. The separation surge was stronger for adults than for children (compare adult and child curves), the surge was weaker for those of higher (parental) education (compare left- and right-hand plots), and the adult–child contrast grows across cohorts and especially among the higher-educated. One implication is that for children of higher-educated parents, the separation surge has been weak, and in some cases, absent. This means diverging destinies in children emanate from the changing interactions of education and parenthood in adult couples. We found this pattern in six of our eight study countries, the United Kingdom being one notable exception.

<Figure 3 about here>

A limitation of Figures 2 and 3 is that conditional annual separation probabilities are not intuitively interpretable in terms of our outcome of interest—separation risk. A meaningful measure from the child perspective is the percentage experiencing parental separation before reaching adulthood (age 18); a meaningful comparable measure for adults is the percentage separating across a risk period of the same union duration. In Figure 4, we present findings obtained from life tables with equivalent exposure times for adults and children. We constructed these life tables from the models on annual separation probabilities shown in Table 5 (Model 2).

Figure 4 summarizes our main findings for two reference cohorts—1950 and 1995—pertaining to union cohorts of adults and birth cohorts of children, respectively. The light blue bars, representing early-SDT cohorts of 1950, indicate low separation risks, with proportions separated after 18 years ranging between 5 and 18 percent. A comparison of the adult and child perspectives shows a pattern of similarity: separation percentages hardly differ between adult and child perspectives, for the lower- and higher-educated. The dark blue bars, representing late-SDT cohorts of 1995, illustrate three main changes across the SDT. First, separation risk surged, often to values that approached or even exceeded 50 percent. Second, separation risk diverged between adult and child perspectives—the proportion experiencing separation surged more strongly in adults than in children. Third, this divergence was stronger among the higher-educated. When comparing the dark blue bars in the right-hand panel of each country plot, we see a clear gradient that is steepest in Belgium, France, Norway, Sweden, and Hungary. For example, only 18.5 percent of Belgian children born in 1995 to a higher-educated parent experienced parental separation. In contrast, 47 percent of unions formed in 1995 among higher-educated Belgians ended in separation—2.5 times the risk found in children. Among the corresponding groups in Sweden and Norway, risk in adults was approximately double the risk in children. Again, the United Kingdom is a notable exception to the general picture, showing a separation surge that was least stratified between the higher- and lower-educated, and between adult and child perspectives.

<Figure 4 about here>

Table 6 shows the extent to which the findings for children are associated with the rise of unmarried parenthood. We estimated the interaction of cohort and education for all children and the same interaction for children born in married unions. In both models, we see negative interactions between education and cohort. The significance levels vary as a result of a change in sample size, but the direction of these interactions does not change when limiting the sample to children born in married unions. In three countries, the magnitude of the education–cohort interactions is somewhat reduced when limiting the sample to children born to married couples (Belgium, France, Norway). In two countries, the interaction slightly increases (Estonia, Sweden), and in one country the interaction remains unchanged (Hungary).

<Table 6 about here>

CONCLUSIONS

Our study demonstrates that adult and child perspectives on the separation surge differ, and these differences have grown across cohorts. The contrast between adult and child perspectives matters for conclusions about the strength of the separation surge and for conclusions about its consequences for social inequality. Both of these themes are key areas of investigation in demography, family sociology, and social stratification research, and both matter for our general understanding of the SDT (Lesthaeghe 2010), the current era of family complexity (Thomson 2014), and concerns about increasing divergence in the destinies of children (McLanahan 2004). Adult and child perspectives differ most where it concerns the unequal impact of the SDT on the lives of adults and children.

The conventional view from the adult perspective shows a separation surge of massive scale across the SDT. It also shows an increasing socioeconomic gradient in many societies, as separation has become increasingly concentrated among less-educated people (De Graaf and Kalmijn 2006; Härkönen and Dronkers 2006; Martin 2006; Musick and Michelmore 2018). On the surface, the implications of these trends for children appear obvious: if adults are more likely to separate, then more children will experience parental separation. If separation is concentrated among less-educated adults, then the "children of separation" will be those of less-educated parents.

As we have shown, this conversion from the adult to the child perspective may be misleading. For example, if the separation surge and the increasing education gradient in separation were concentrated in childless unions, then both trends would be absent when viewed from the child perspective. Conversely, if concentrated in parents, both trends would be even stronger when viewed from the child perspective. In contrasting adult and child perspectives, we considered factors that influence their linkage, most notably the associations between parenthood, education, and separation, as well as cohort change in these associations.

Our analyses of retrospective life history data (HHD) from eight European countries yield various new findings, of which two are central. First, the negative association between parenthood and separation intensified across cohorts. As a consequence, the separation surge has been less radical for children than what data for adults suggests. This difference was statistically significant in six countries, and the differences between adult and child perspectives were substantial in scope. Moreover, the differences have grown across our study cohorts (1945 to 2005) and may grow further in future cohorts. Second, the association between parenthood and separation was stratified by education, and in most countries, this stratification intensified across cohorts. As a consequence, inequality in the separation experience has increased more for children than for adults. This corroborates the thesis of "diverging destinies" in a novel way (McLanahan 2004). In some countries, the separation surge across the SDT has been almost absent for higher-educated parents.

In cross-national comparison, we found more similarities than differences, but the picture was not fully consistent across the eight European countries we studied. Our hypothesis that children's risk of parental separation is lower than adult separation risk (Hypothesis 1) was supported in all eight countries, but the magnitude of this difference varied considerably, just like the association between parenthood and separation in the adult data. Our hypothesis that the education gradient in separation risk is more negative for children than for adults (Hypothesis 2) was supported in six countries but not in the United Kingdom and Estonia, both of which are high-divorce societies. Our hypothesis that the separation surge has been weaker for children (Hypothesis 3) was supported in all countries except Estonia. Finally, our hypothesis that the adult–child contrast would grow across cohorts, and especially among the higher-educated (Hypothesis 4), was supported in all countries except the United Kingdom and the Czech Republic. Overall, this shows our main conclusions were well supported by the cross-national evidence but were not universal or similarly pronounced across societies. The United Kingdom showed the poorest fit to our expectations, suggesting high-divorce societies with weak welfare states and high levels of inequality may exhibit different demographic dynamics. This possibility calls for a replication of our analysis in the United States.

For cohabitation, we found the expected positive association with separation. In three countries, this association weakened across cohorts, pointing to a normalization of this union type (Liefbroer and Dourleijn 2006). In four countries, lower separation risks in the presence of children pertained more strongly to married than to cohabiting unions. Regarding education, the negative gradient in separation risk was stronger in married than in cohabiting unions. Our main conclusions about the contrast between adult and child perspectives were not overturned by the inclusion of cohabitation and its interactions with parenthood, cohort, and education, despite changes in the size and statistical significance of parameters in some countries. Overall, these results indicate that the rise of cohabitation is associated with the main trends we found.

The causal interpretation of this finding remains ambiguous. A supply-side interpretation suggests unmarried cohabitation is a novel option that has independently destabilized unions (Waite et al. 2000). A demand-side interpretation suggests declines in the commitment to long-term unions had other causes, such as broader cultural and economic changes that increased not only separation but also the prevalence of cohabitation (Blossfeld et al. 2005). According to the former interpretation, our findings on diverging destinies were (at least partly) due to the rise of cohabitation. According to the latter interpretation, trends in cohabitation and unmarried parenthood were a parallel outcome to the changing pattern of separation we observed. Regardless of which interpretation is favored, rising inequality in children's experience of parental separation is an important phenomenon—and relevant to parents, policymakers, and researchers.

The distinction between cohabitation and marriage is also relevant for the consequences of the trends we found. Studies show that the dissolution of a cohabiting union has a weaker negative effect on ex-partners' well-being than does the dissolution of a marriage (Blekesaune 2008; Dush 2013; Kalmijn 2017). This may partly result from the fact that such unions are shorter. Yet, these differences do suggest that the rapid increase in union instability, especially among childless couples, may not diminish individual well-being. At the same time, union dissolutions that involve children appear more harmful for ex-partners than dissolutions that do not involve children (Leopold and Kalmijn 2016; Williams and Dunne-Bryant 2006). The trends we found may thus contribute to increasing inequality in mental health among adults, especially in high-divorce societies.

We considered different mechanisms underlying the adult–child contrast as well as its cohort change and social stratification, including changes in the selectivity of parenthood, changes in parents and children "protecting" each other from the risk of separation, and changes in how these aspects are stratified by education. The presence and relative importance of these mechanisms remain to be tested, but their implication is clear: the social demography of separation differs for adults and children, and the differences between the two perspectives have grown across cohorts. Future work could add to our study in at least four ways.

First, given our focus on the associations between parenthood, education, and separation, we did not explore several other factors that can influence the linkage between adult and child perspectives. These include general trends in fertility as well as specific factors such as the timing of union formation, the timing of fertility in (marital and cohabiting) unions, and the increase in the prevalence of higher-order unions. Our models were deliberately parsimonious, as we used no controls for other sources of demographic change. We leave it to future analyses to examine the role of these demographic processes in the contrast between adult and child perspectives on divorce and separation.

Second, our profile of the "children of separation" is incomplete. We based our analyses on changes in the relative risk of separation and disregarded absolute numbers as well as compositional changes. For example, even if the risk of parental separation surged among children of the lower-educated, the size of this risk group has diminished with educational expansion. In addition, rising economic inequality could contribute to cohort increases in the diverging family experiences of children of higher- and lower-educated parents. To gain better insight into the socially stratified separation experience and the associated notion of diverging destinies, future studies on the children of separation should address the changing size and socioeconomic composition of education groups as well as macro-level trends of rising inequality.

Third, our comparative scope is necessarily limited. We examined eight countries representing a variety of national contexts across Europe, but we did not examine societies that differed substantially from our study countries in terms of separation trends and possible factors underlying the linkage between adult and child perspectives. Our findings do not suggest an obvious link between the degree of educational and income inequality in a society and the strength of the key interactions of interest. We also found no suggestive evidence for a link with the prevalence of divorce, as high-divorce countries such as Norway and the United Kingdom showed very different patterns. A larger number of countries is needed to examine these or other potentially relevant macro-level factors.

Fourth, further micro-level research is needed on the link between parenthood and separation. It is important to further evaluate the mechanisms of investment, protection, and selection to understand which processes drive the changing interactions between parenthood and education we observed.

These directions for future research could complement our demographic comparison of adult and child perspectives on divorce and separation. Linking both perspectives will also benefit studies focusing on outcomes in adults and children. For example, analysts commonly assume that negative effects of parental separation on child outcomes such as well-being, mental health, deviance, and education are partly mediated by the personal crises mothers and fathers experience throughout the process of separation. Conversely, negative outcomes in adults may be mediated by children's problems following separation. Long-running genealogical panel studies and multi-actor studies now offer possibilities to observe adults and children jointly across the separation process, thus linking their perspectives in studies on the consequences of separation for adults and children.

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Notes

1. The Harmonized Histories data file was created by the Non-Marital Childbearing Network (http://www.nonmarital.org) (see Perelli-Harris, Kreyenfeld, and Kubisch 2011).

2. In the HHD file, the British Household Panel Survey (BHPS) does not include the age at leaving school. The United Kingdom also has a more complex educational classification system. For this reason, we constructed an alternative education variable for the United Kingdom, consisting of six education categories, and we scaled these to the average years of schooling (i.e., no qualification [9.3], O-levels [10.8], A-levels [12.3], other higher qualification [12.5], first degree [14.8], and higher degree [15.6]). We obtained these averages from the General Household Survey of 2005. The regression results were highly similar in terms of point estimates and significance levels. The alternative educational variable was highly correlated to the ISCED-approximation (r = .85). For this reason, we decided to keep the ISCED-rescaling consistent for all countries.

3. Deceased children are excluded. We do not transfer children from earlier unions to new unions.

4. The length of the risk period for calculating cumulative proportions is different from the observation time window used for estimating event-history models.

5. The year in which the effect of education becomes negative is $\mu_y + 10\ (-\beta_1 / \beta_2)$, where μ_y is the average year in a country, β_1 is the main effect, and β_2 is the interaction effect.

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Matthijs Kalmijn is a full Professor of Sociology at the University of Groningen and senior researcher at the Netherlands Interdisciplinary Demographic Institute (NIDI-KNAW) in The Hague. His main research fields are family, life courses, and intergenerational relationships. He was also co-director of several large-scale surveys in the Netherlands, including the survey Parents and Children in the Netherlands (OKiN), the Netherlands Kinship Panel Study (NKPS), and the Netherlands Longitudinal Lifecourse Study (NELLS). Kalmijn has a PhD from UCLA (1991). He is principal investigator of the ERC Advanced Grant project Intergenerational Reproduction and Solidarity in an Era of Family Complexity.

Thomas Leopold is a full Professor of Methods of Empirical Social Research at the Institute of Sociology and Social Psychology at the University of Cologne. He completed a PhD at the University of Bamberg and was a Max Weber Postdoctoral Fellow at the European University Institute. His interests are in social demography, family relations, and life course research. He is principal investigator of the ERC Starting Grant project KINMATRIX: Uncovering the Kinship Matrix: A New Study of Solidarity and Transmission in European Families. He is also principal investigator of the NORFACE project CRITEVENTS (2017–2020), funded by the European Commission via an ERA-NET Cofund grant.

Figure 1. Cohort trends in adult separation risks

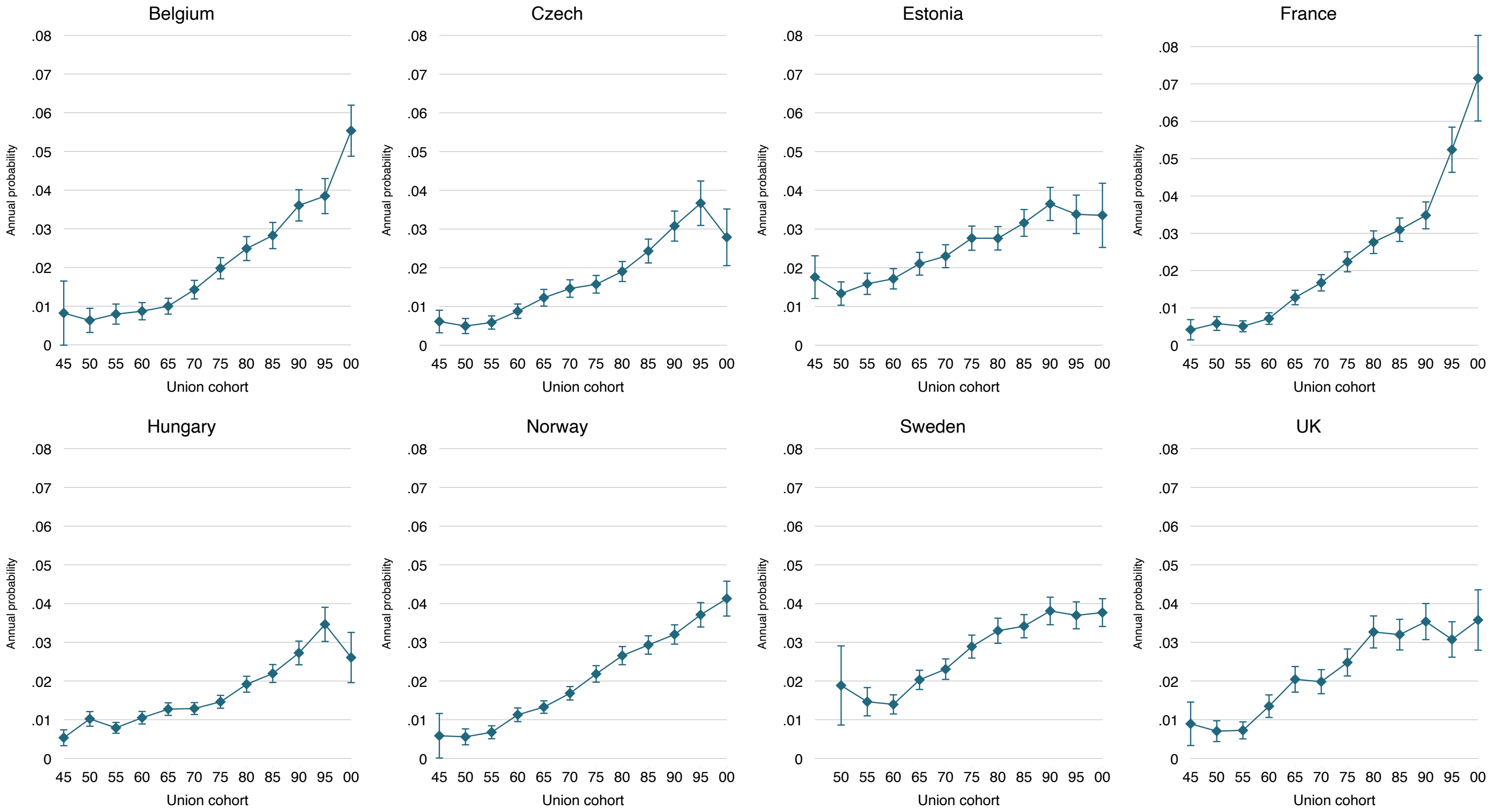
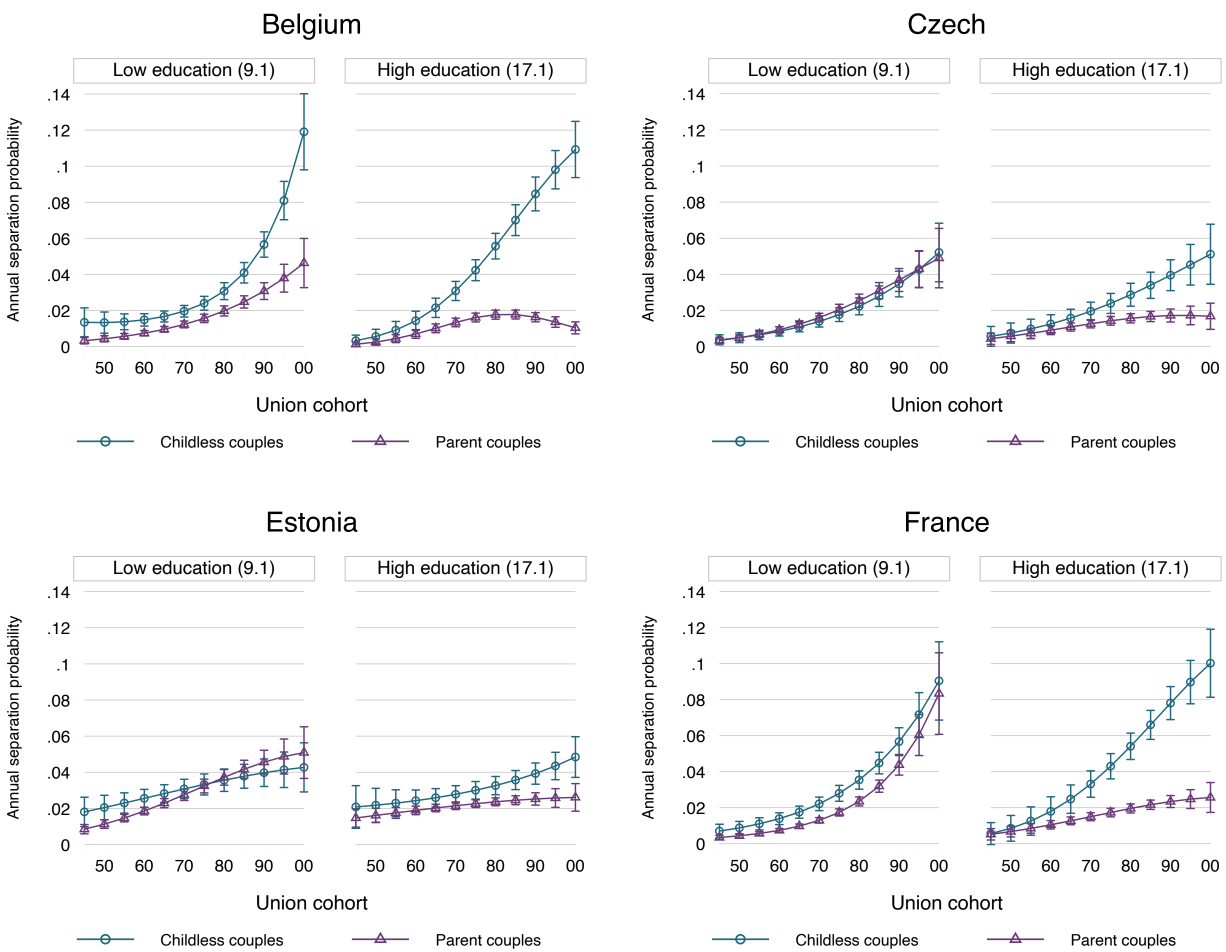


Figure 2. Cohort trends in adult separation risks



[figure continues]

Figure 2. Continued

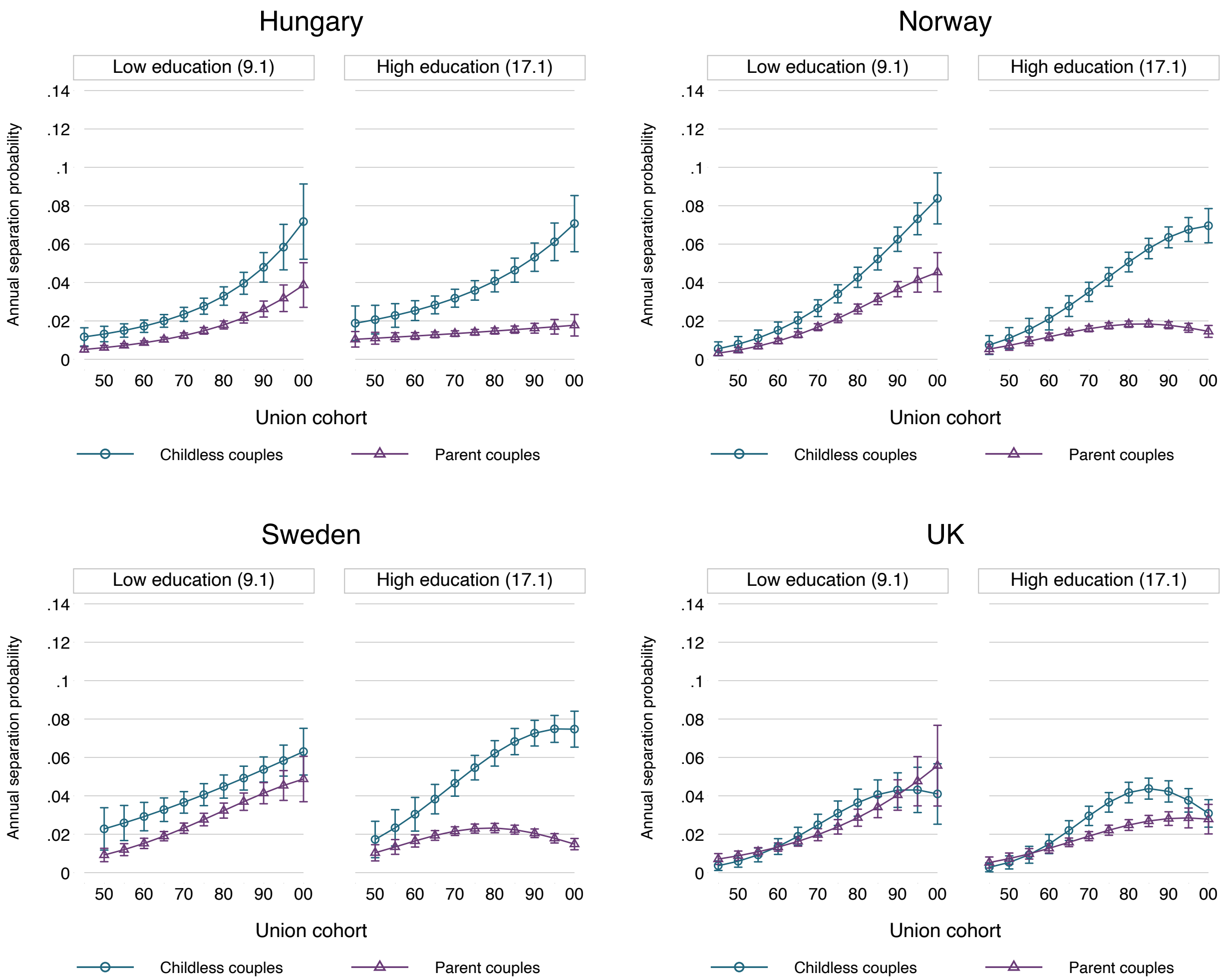
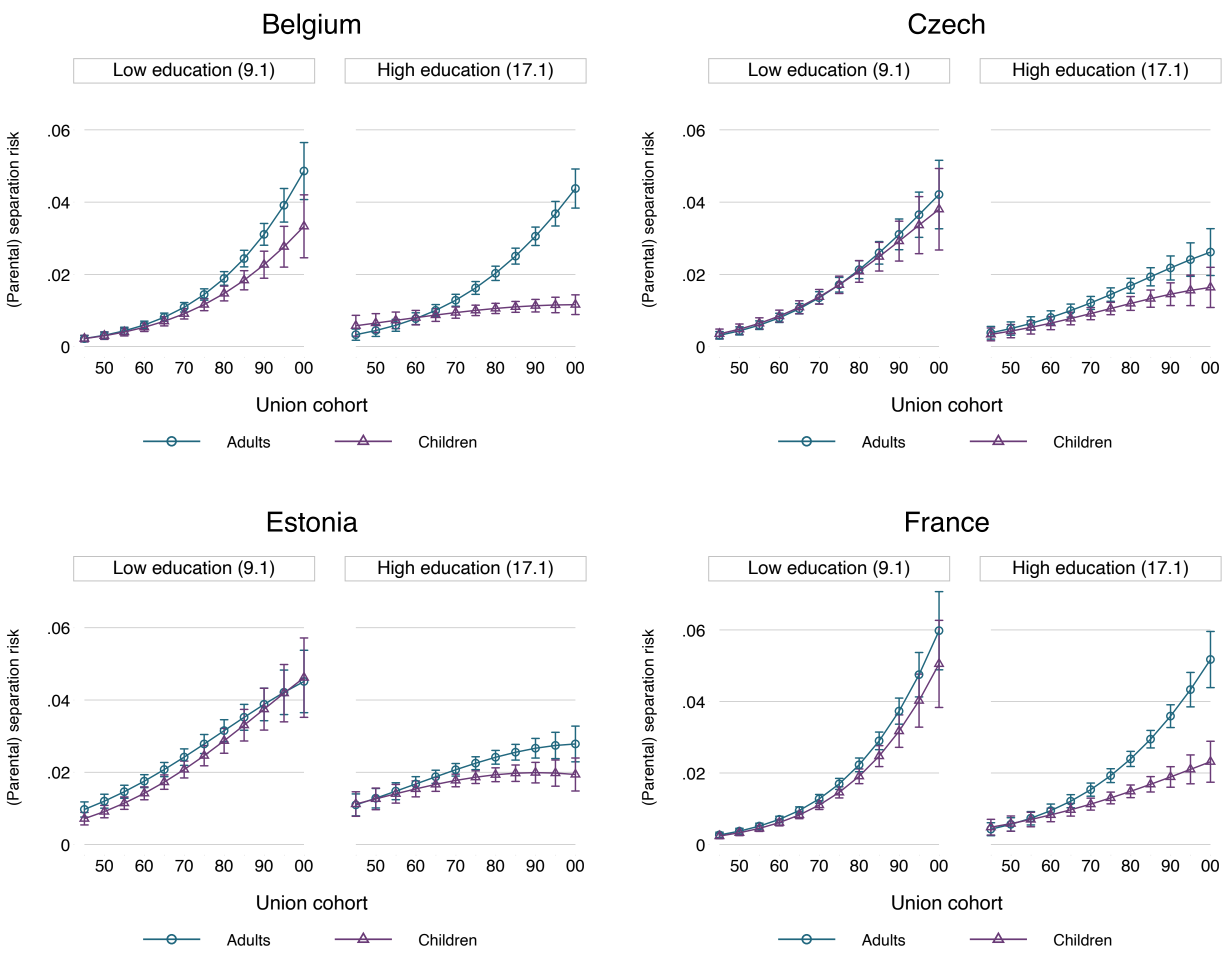


Figure 3. Cohort trends in separation risks of adults and children by education



[figure continues]

Figure 3. Continued

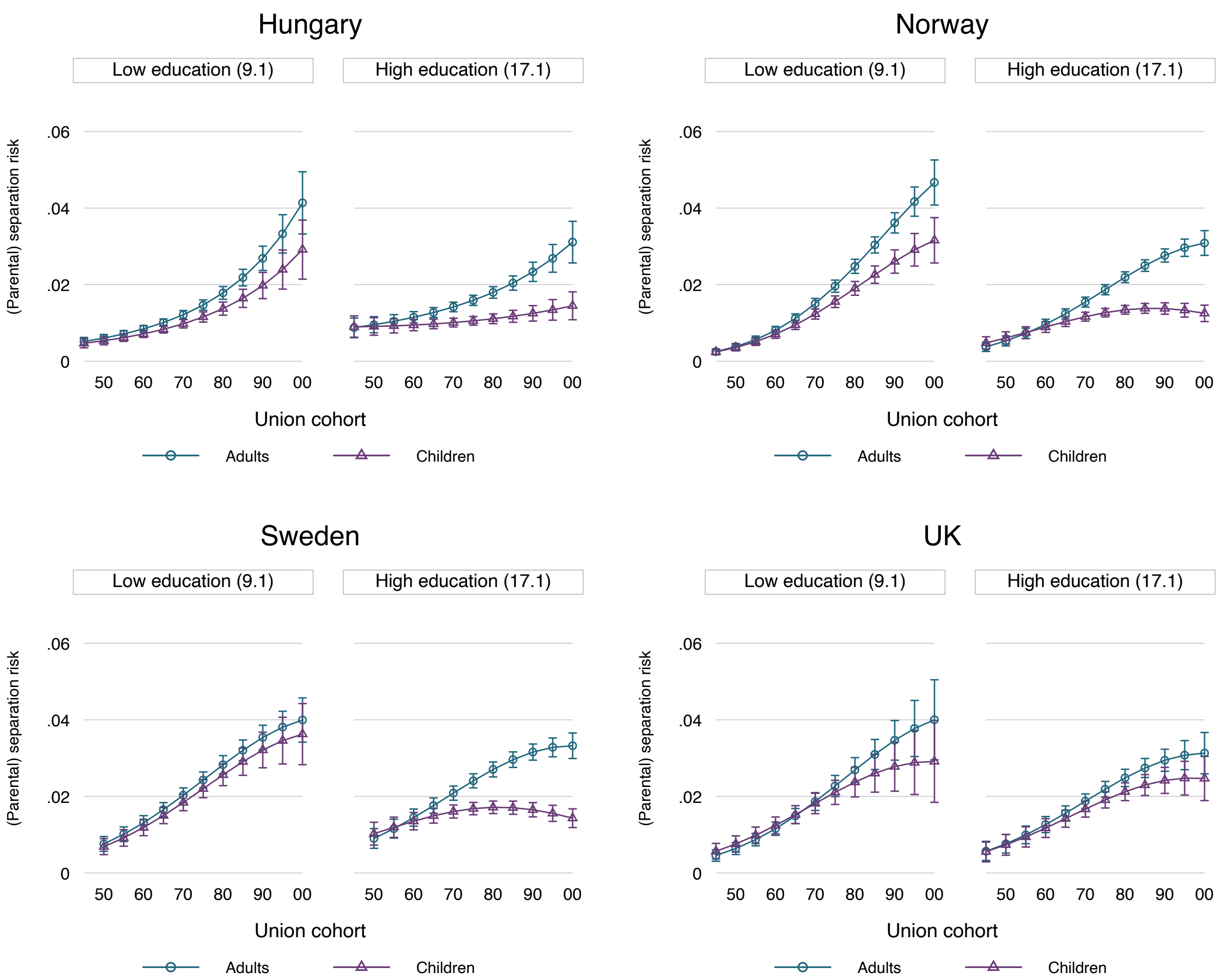
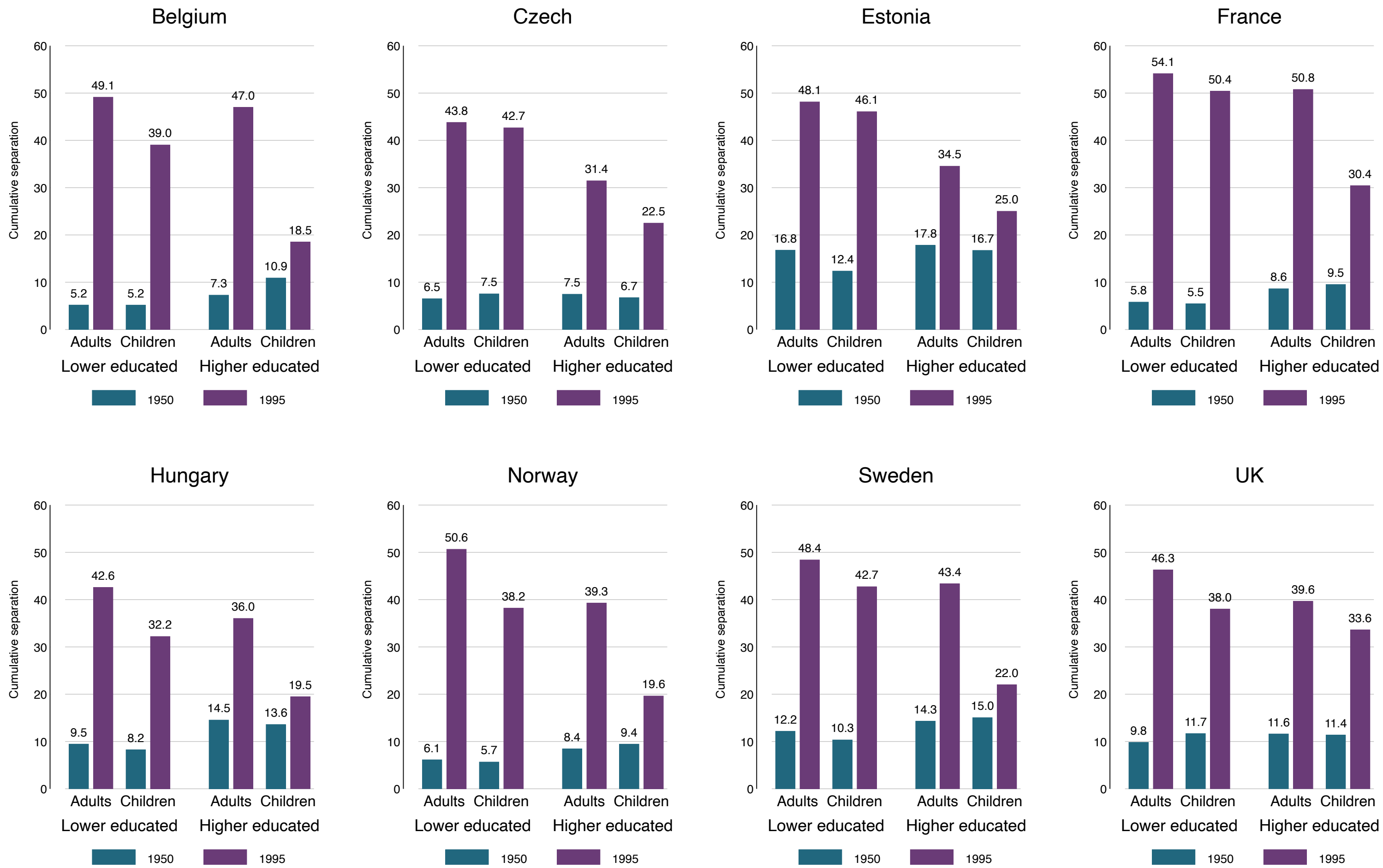


Figure 4. Cumulative (parental) separation percentages



Note: Percentages based on predicted annual risks of the model in Table 3.

Table 1. Macro-Characteristics of Countries in the Analysis

	Crude Divorce Rate[a]	Crude Marriage Rate[a]	Total Fertility Rate[a]	Cohabitation[b]	Children in Cohabitation[c]	GDP per Capita[d]	Social Spending[e]	Income Inequality[f]	Educational Inequality[g]
Belgium	2.5	5.2	1.60	13.9	47.2	36.6	23.5	3.3	130
Czech Republic	2.9	6.1	1.42	11.3	46.2	21.1	17.9	3.2	169
Estonia	3.8	5.0	1.50	30.8	56.9	15.6	13.8	4.9	130
France	2.0	4.7	1.75	23.0	53.2	34.7	27.6	3.4	157
Hungary	2.3	5.1	1.56	18.4	52.6	17.9	20.1	3.4	179
Norway	2.3	5.0	1.87	24.4	60.2	58.0	20.4	2.8	126
Sweden	2.4	4.1	1.79	30.6	50.4	36.8	26.8	2.8	122
United Kingdom	2.8	5.7	1.74	25.0	43.3	33.4	16.2	4.1	142
United States[h]	4.4	8.8	2.02	11.9	39.5	45.7	15.6	5.9	168

[a]OECD Family Database (1990 to 2000 averaged).
[b]OECD Family Database (percentage of cohabitors of all couples in 2011).
[c]OECD Family Database (percentage of cohabitors who had children in 2011).
[d]World Bank Database (2000).
[e]Social spending as percentage of gross national income (OECD, 2000 or closest).
[f]Income ratio of top and bottom 10 percent (p90/p10) (OECD, 2004 or closest).
[g]Earnings ratio of people with tertiary versus upper-secondary education (OECD, most recent data).
[h]Not included in the analysis but presented as a reference.

Table 2. Means by Country in Person-Period File in All Years

	Age	Education	Union Time	Union Cohort	Separation Year	Age at Union	Cohabiting	Separated	Children	*N*
All Years										
Belgium	33.3	12.9	10.1	1979.1	1995.5	24.2	19.9	2	60.5	109,154
Czech Republic	33.1	12.6	10.3	1974.8	1992.7	23.8	8.8	1.4	72.2	124,596
Estonia	33.7	13.5	10	1974.4	1990.2	24.7	18.2	1.9	76.4	128,005
France	33.3	12.2	10.1	1975.1	1992.7	24.2	21.9	1.8	71.1	144,183
Hungary	32.9	12.9	10.4	1973	1991	23.5	7.4	1.4	77.3	198,288
Norway	34.1	13.5	9.9	1978.8	1993.7	25.2	24.7	2	71	214,850
Sweden	34.6	13.8	10	1981.2	1995.9	25.6	37.1	2.5	66.9	142,294
United Kingdom	34.1	14	9.7	1976.7	1991.2	25.4	14.5	2	64.8	91,039
First Year										
Belgium	24.4	13.2	1	1983.5	1994.6	24.4	57.5		4.5	7,051
Czech Republic	24.3	12.7	1	1980	1991.2	24.3	28.4		16.4	7,898
Estonia	24.8	13.5	1	1978.7	1987.9	24.8	48.8		11.7	8,553
France	24.9	12.6	1	1980.4	1991.8	24.9	55.5		7.9	9,431
Hungary	23.9	13.1	1	1977.5	1988.7	23.9	21.1		11.5	12,303
Norway	25.6	13.6	1	1983.4	1992.9	25.6	63.3		11.1	14,367
Sweden	25.6	13.9	1	1984.1	1993.7	25.6	80.1		8.1	9,285
United Kingdom	26.1	14.3	1	1982.7	1990.2	26.1	46.4		9.5	6,557
Separation Year										
Belgium	30.1	13.5	8.5	1986.8	1994.3	22.6	56.1		37.2	2,155
Czech Republic	32	12.7	9.6	1981.3	1989.9	23.4	18.9		68.3	1,697
Estonia	31.5	13.5	8.8	1978.5	1986.3	23.6	27.8		71.6	2,472
France	33	12.9	9.1	1983.3	1991.5	24.8	54.1		54.4	2,583
Hungary	31.8	13.3	9	1978.7	1986.8	23.7	22.5		59.5	2,766
Norway	32.1	13.6	8.6	1984.8	1992.4	24.5	53.1		50.6	4,299
Sweden	31.8	13.9	8.4	1985.3	1992.7	24.4	63.3		46.5	3,623
United Kingdom	32.4	14.3	8.8	1982	1989.8	24.6	29.6		55.7	1,799

Table 3. Event-History Models of Adults Experiencing Separation: Coefficients and *p*-Values

	Belgium	Czech Republic	Estonia	France	Hungary	Norway	Sweden	United Kingdom
Model 1								
Union time	2.136**	1.674**	1.659**	2.022**	1.440**	1.705**	1.615**	1.371**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union time (sq.)	–.878**	–.695**	–.850**	–.764**	–.635**	–.737**	–.799**	–.631**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort	.465**	.425**	.229**	.494**	.278**	.338**	.188**	.317**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort (sq.)	–.025	–.045**	–.037**	–.030*	.002	–.061**	–.057**	–.097**
	(.086)	(.002)	(.001)	(.020)	(.864)	(.000)	(.000)	(.000)
Education	–.017*	–.027*	–.030**	–.006	.006	–.034**	–.023**	–.010
	(.035)	(.019)	(.000)	(.409)	(.442)	(.000)	(.001)	(.238)
Children	–1.047**	–.259**	–.229**	–.687**	–.840**	–.871**	–.833**	–.300**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Education x Children	–.097**	–.083**	–.022	–.067**	–.032*	–.059**	–.087**	–.032*
	(.000)	(.000)	(.140)	(.000)	(.029)	(.000)	(.000)	(.047)
Chi-2	995.2	457.3	354.6	1062.6	843.9	1293.9	796.6	326.3
Model 2								
Union time	2.174**	1.697**	1.666**	2.001**	1.441**	1.724**	1.593**	1.366**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union time (sq.)	–.904**	–.709**	–.852**	–.759**	–.639**	–.753**	–.800**	–.629**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort	.404**	.409**	.216**	.474**	.255**	.309**	.163**	.309**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort (sq.)	–.033*	–.045**	–.026*	–.018	.009	–.072**	–.064**	–.089**
	(.045)	(.003)	(.030)	(.212)	(.396)	(.000)	(.000)	(.000)
Education	.016*	–.009	–.025**	.020*	.021**	–.008	–.003	–.005
	(.044)	(.456)	(.000)	(.019)	(.006)	(.207)	(.684)	(.575)
x Union cohort	–.013*	–.020*	–.018**	–.018**	–.023**	–.022**	–.009	–.008
	(.028)	(.025)	(.001)	(.004)	(.000)	(.000)	(.086)	(.221)
Children	–.917**	–.186**	–.240**	–.684**	–.803**	–.770**	–.775**	–.310**
	(.000)	(.007)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
x Union cohort	–.265**	–.112*	.015	–.055	–.077*	–.170**	–.146**	–.004
	(.000)	(.016)	(.682)	(.171)	(.017)	(.000)	(.000)	(.925)
Chi-2	1067.9	460.4	382.9	1042.3	895.4	1439.6	826.3	323.5
Model 3								
Union time	2.228**	1.722**	1.677**	2.051**	1.460**	1.751**	1.634**	1.376**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union time (sq.)	–.922**	–.717**	–.855**	–.776**	–.646**	–.762**	–.812**	–.631**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort	.407**	.408**	.214**	.475**	.256**	.307**	.163**	.313**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)

Union cohort (sq.)	–.028	–.045**	–.029*	–.012	.009	–.069**	–.061**	–.087**
	(.078)	(.003)	(.015)	(.415)	(.411)	(.000)	(.000)	(.000)
Education	.001	–.013	–.030**	.013	.014	–.019**	–.024**	–.005
	(.923)	(.312)	(.000)	(.134)	(.080)	(.003)	(.001)	(.587)
x Union cohort	–.030**	–.030**	–.023**	–.033**	–.032**	–.037**	–.026**	–.012
	(.000)	(.001)	(.000)	(.000)	(.000)	(.000)	(.000)	(.086)
Children	–.892**	–.183**	–.218**	–.659**	–.783**	–.770**	–.768**	–.313**
	(.000)	(.008)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
x Union cohort	–.229**	–.103*	.009	–.004	–.063	–.162**	–.120**	.014
	(.000)	(.028)	(.809)	(.929)	(.053)	(.000)	(.000)	(.739)
Education x Children	–.070**	–.074**	–.022	–.059**	–.030	–.062**	–.083**	–.047**
	(.000)	(.007)	(.171)	(.002)	(.061)	(.000)	(.000)	(.010)
x Union cohort	–.039**	–.023	–.035**	–.031*	–.025*	–.024*	–.041**	.011
	(.001)	(.206)	(.002)	(.012)	(.023)	(.020)	(.000)	(.381)
Chi-2	1134.6	475.6	394.5	1076.1	901.2	1454.3	878.7	327.6
N union-years	109,154	124,596	128,005	144,183	198,288	214,850	142,294	91,039
N clusters	5,735	7,041	6,909	8,062	10,650	11,794	7,080	5,308

Note: Variables centered. Cohort divided by 10. *P*-values corrected for clustering of unions in persons (clusters are persons).
**p* < .05; ***p* < .01 (two-tailed tests).

Table 4. Event-History Models of Adults Experiencing Separation: Including Cohabitation Effects

	Belgium	Czech Republic	Estonia	France	Hungary	Norway	Sweden	United Kingdom
Model 1								
Union time	2.790**	1.826**	1.703**	2.366**	1.469**	2.219**	2.125**	1.745**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union time (sq.)	–1.073**	–.755**	–.856**	–.855**	–.647**	–.901**	–.949**	–.755**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort	.277**	.376**	.175**	.329**	.206**	.171**	.080**	.225**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort (sq.)	–.014	–.065**	–.036**	–.028	–.012	–.092**	–.039**	–.098**
	(.369)	(.000)	(.007)	(.090)	(.329)	(.000)	(.003)	(.000)
Cohabitation	1.430**	.591**	.553**	1.148**	.860**	.867**	.838**	1.024**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
x Union cohort	–.140**	.069	–.098*	.036	.022	.184**	.010	–.111
	(.008)	(.339)	(.040)	(.512)	(.711)	(.000)	(.808)	(.123)
x Children	–.350**	.088	.707**	.260*	.318*	–.100	–.527**	.306*
	(.006)	(.572)	(.000)	(.011)	(.010)	(.235)	(.000)	(.021)
x Education	.050**	.077*	.094**	.045**	.042*	.037**	.076**	.075**
	(.006)	(.027)	(.000)	(.009)	(.046)	(.008)	(.000)	(.001)
Education	–.008	–.009	–.025**	.008	.021**	–.012	–.024**	–.005
	(.348)	(.477)	(.000)	(.380)	(.008)	(.081)	(.001)	(.603)
x Union cohort	–.035**	–.032**	–.031**	–.033**	–.029**	–.041**	–.034**	–.020*
	(.000)	(.001)	(.000)	(.000)	(.000)	(.000)	(.000)	(.011)
Children	–.448**	–.115	–.196**	–.428**	–.690**	–.553**	–.430**	–.208**
	(.000)	(.097)	(.001)	(.000)	(.000)	(.000)	(.000)	(.001)
x Union cohort	–.176**	–.027	–.093*	–.036	–.008	–.097*	–.097*	.007
	(.000)	(.594)	(.027)	(.438)	(.819)	(.012)	(.011)	(.879)
Education x Children	–.032	–.063*	.012	–.021	–.021	–.055**	–.037*	–.027
	(.085)	(.021)	(.470)	(.302)	(.208)	(.000)	(.019)	(.151)
x Union cohort	–.044**	–.006	–.029*	–.035**	–.014	–.016	–.047**	.016
	(.000)	(.736)	(.012)	(.006)	(.204)	(.137)	(.000)	(.243)
Chi-2	1683.4	553.0	503.0	1606.0	1112.5	2153.9	1294.4	507.6
N unions-years	109,154	124,596	128,005	144,183	198,288	214,850	142,294	91,039
N clusters	5,735	7,041	6,909	8,062	10,650	11,794	7,080	5,308

Note: Variables centered. Cohort divided by 10. *P*-values corrected for clustering of unions in persons (clusters are persons).
$^{*}p < .05$; $^{**}p < .01$ (two-tailed tests).

Table 5. Event-History Models of Children and Adults Experiencing (Parental) Separation

	Belgium	Czech Republic	Estonia	France	Hungary	Norway	Sweden	United Kingdom
Model 1								
Union time	.066	.883**	.911**	.588**	.297**	.199*	–.031	.444**
	(.562)	(.000)	(.000)	(.000)	(.010)	(.015)	(.718)	(.002)
Union time (sq.)	.019	–.336**	–.496**	–.186**	–.162**	–.112**	–.072	–.188**
	(.693)	(.000)	(.000)	(.000)	(.001)	(.002)	(.055)	(.002)
Union cohort	.525**	.433**	.231**	.549**	.308**	.437**	.273**	.356**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort (sq.)	–.028	–.040*	–.031*	–.022	.018	–.064**	–.054**	–.058**
	(.077)	(.012)	(.013)	(.156)	(.140)	(.000)	(.000)	(.001)
Education	.025**	–.017	–.021**	.031**	.022**	.002	.001	–.004
	(.004)	(.186)	(.004)	(.002)	(.005)	(.790)	(.881)	(.657)
x Union cohort	–.025**	–.017	–.023**	–.024**	–.025**	–.028**	–.021**	–.008
	(.000)	(.078)	(.000)	(.000)	(.000)	(.000)	(.000)	(.329)
Child perspective	–.380**	–.125**	–.169**	–.227**	–.287**	–.321**	–.297**	–.088**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
x Union cohort	–.197**	–.053*	–.008	–.073**	–.090**	–.144**	–.110**	–.068**
	(.000)	(.011)	(.610)	(.000)	(.000)	(.000)	(.000)	(.001)
x Education	–.055**	–.040**	–.005	–.038**	–.020*	–.030**	–.045**	–.008
	(.000)	(.000)	(.422)	(.000)	(.010)	(.000)	(.000)	(.359)
Chi-2	1022.9	449.0	414.5	1106.4	877.2	1626.7	789.1	307.1
Model 2								
Union time	.069	.883**	.913**	.594**	.299**	.201*	–.028	.443**
	(.548)	(.000)	(.000)	(.000)	(.009)	(.014)	(.749)	(.002)
Union time (sq.)	.018	–.336**	–.496**	–.188**	–.162**	–.113**	–.074	–.187**
	(.705)	(.000)	(.000)	(.000)	(.001)	(.002)	(.050)	(.002)
Union cohort	.522**	.434**	.235**	.546**	.307**	.437**	.274**	.356**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort (sq.)	–.029	–.040*	–.032*	–.022	.017	–.064**	–.054**	–.058**
	(.070)	(.012)	(.012)	(.148)	(.146)	(.000)	(.000)	(.001)
Education	.013	–.018	–.023**	.022*	.020*	–.005	–.004	–.002
	(.120)	(.162)	(.001)	(.024)	(.012)	(.481)	(.549)	(.801)
x Union cohort	–.011	–.016	–.015**	–.015*	–.020**	–.020**	–.009	–.011
	(.066)	(.074)	(.005)	(.019)	(.000)	(.000)	(.068)	(.098)
Child perspective	–.362**	–.124**	–.157**	–.216**	–.279**	–.315**	–.284**	–.095**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
x Union cohort	–.203**	–.054*	–.018	–.071**	–.092**	–.147**	–.116**	–.067**
	(.000)	(.011)	(.257)	(.000)	(.000)	(.000)	(.000)	(.002)
Child perspective x Education	–.037**	–.039**	–.002	–.023**	–.016*	–.020**	–.040**	–.010
	(.000)	(.000)	(.715)	(.006)	(.032)	(.002)	(.000)	(.241)
x Union cohort	–.035**	–.002	–.017**	–.019**	–.012	–.018**	–.025**	.007
	(.000)	(.803)	(.001)	(.006)	(.076)	(.002)	(.000)	(.332)

Chi-2	1027.8	454.1	407.4	1120.3	884.3	1612.2	779.4	308.4
N	23,1802	272,178	285,863	344,225	452,395	520,565	326,187	208,747
N clusters 1	7,063	7,970	8,554	9,446	12,373	14,436	9,334	6,567
N clusters 2	5,740	7,087	6,909	8,071	10,693	11,831	7,101	5,312

Note: Variables centered. Cohort divided by 10. *P*-values corrected for multiway clustering of unions and adults/parents. Clusters 1 are unions, clusters 2 are adults/parents.
**p* < .05; ***p* < .01 (two-tailed tests).

Table 6. Event-History Models of Children Experiencing (Parental) Separation

	Belgium	Czech Republic	Estonia	France	Hungary	Norway	Sweden	United Kingdom
Children born in married and unmarried unions (Model 1)								
Union cohort	.415**	.404**	.223**	.491**	.233**	.311**	.173**	.289**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort (sq.)	–.083**	–.055*	–.039*	–.026	.002	–.075**	–.044*	–.021
	(.003)	(.012)	(.017)	(.235)	(.902)	(.000)	(.016)	(.388)
Education	–.033*	–.058**	–.023*	.004	.003	–.024**	–.046**	–.007
	(.016)	(.001)	(.014)	(.768)	(.753)	(.009)	(.000)	(.571)
x Union cohort	–.040**	–.018	–.031**	–.036**	–.029**	–.037**	–.035**	–.006
	(.000)	(.164)	(.000)	(.000)	(.001)	(.000)	(.000)	(.545)
Chi-2	166.5	181.9	174.9	313.0	110.4	254.7	99.8	103.3
N	121,522	147,280	158,782	199,130	255,816	303,325	181,131	117,336
N clusters 1	4,484	5,803	6,571	6,976	9,653	10,389	6,488	4,348
N clusters 2	4,318	5,552	6,023	6,681	9,213	9,829	6,071	4,097
Children born in marriage (Model 2)								
Union cohort	.327**	.391**	.199**	.393**	.204**	.183**	.139**	.238**
	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)	(.000)
Union cohort (sq.)	–.154**	–.070**	–.037	–.093**	–.021	–.120**	–.010	–.036
	(.000)	(.004)	(.057)	(.002)	(.278)	(.000)	(.673)	(.211)
Education	–.023	–.062**	–.025*	–.002	.012	–.016	–.056**	–.009
	(.107)	(.000)	(.018)	(.857)	(.272)	(.127)	(.000)	(.518)
x Union cohort	–.020	–.018	–.039**	–.026*	–.025**	–.029**	–.036**	–.007
	(.134)	(.205)	(.000)	(.032)	(.008)	(.004)	(.001)	(.578)
Chi-2	130.2	148.4	110.7	124.3	71.2	128.5	50.5	64.2
N	111,380	138,890	138,795	175,347	244,862	253,606	122,898	109,396
N clusters 1	3,802	5,355	5,381	5,586	9,069	8,062	4,351	3,782
N clusters 2	3,751	5,188	5,107	5,490	8,754	7,851	4,241	3,633

Note: Variables centered. Cohort divided by 10. Duration included. Standard errors (in parentheses) corrected for multiway clustering of unions and adults/parents. Clusters 1 are unions, clusters 2 are adults/parents.
**p* < .05; ***p* < .01 (two-tailed tests).