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Shifting Childrearing to Single Mothers: Results from 17 Western Countries

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The final third of the twentieth century witnessed remarkable changes in patterns of family formation within most Western countries. During the two decades that followed World War II, marriage and childbearing occurred early in adulthood and were tightly linked. Then, beginning in the late 1960s in some countries and spreading to many others over the next few decades, the average age at first marriage increased as growing proportions of couples cohabited either as a prelude or an alternative to marriage (Bumpass and Sweet 1989b; Cherlin 1992; Casper and Cohen 2000; Kiernan 2000; Prinz 1995; Raley 2000; Smock 2000). Accompanying these changes in marriage practices were large increases in the prevalence of nonmarital childbearing (Sardon 2000; Ventura and Bachrach 2000), especially conspicuous in the face of rapid declines in marital fertility (Smith, Morgan, and Koropeckyj-Cox 1996).

Labeled by van de Kaa (1987) as the "second demographic transition," this multifaceted departure from the ordered sequence of marriage and childbearing has created challenges for researchers attempting to describe contemporary patterns of family formation with precision. The nonmarital fertility ratio (NMFR), perhaps the most closely watched indicator of changes in family structure, has become an increasingly blunt instrument in light of the share of nonmarital fertility accounted for by parental cohabitation. Moreover, most fertility studies cannot easily determine the number of children women have had by different partners (much less the number of childbearing partners that men have had). Different family forms are often confused in public rhetoric and sometimes even by those researchers who use marriage as a proxy for a nuclear household or nonmarriage as a proxy for single parenthood.

This article builds on a growing effort by some family demographers to fashion new techniques for capturing the ongoing changes in partnership formation and dissolution, marriage, and childbearing, and for examining how these changes translate into different familial experiences for children (Bracher and Santow 1990; Bumpass 1984; Bumpass and Lu 2000; Bumpass and Raley 1995; Bumpass and Rindfuss 1979; Bumpass and Sweet 1989a; Clarke 1992; Furstenberg et al. 1983; Graefe and Lichter 1999; Hoem and Hoem 1992). We extend previous research by analyzing the nationally representative Fertility and Family Surveys from 14 European countries and from Canada, New Zealand, and the United States. We employ multistate life table analysis to estimate children's total expected

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duration in selected family types and probabilities of transition between them for a series of national synthetic birth cohorts covering, in most cases, the late 1970s to the mid-1990s.

This exercise is instructive because previous research has shown that family structure—the set of residential arrangements of children's main caregivers—has important consequences for the welfare of children. Numerous studies have shown that individuals generally fare best both in childhood and in later life when they grow up with both of their biological parents (Amato and Booth 1997; Cherlin 1999; Cherlin, Chase-Lansdale, and McRae 1998; Furstenberg and Kiernan 2001; Jonsson and Gähler 1997; Kiernan and Cherlin 1999; McLanahan and Sandefur 1994). The reasons for this are largely related to the economic disadvantages faced by single and divorced mothers (Burkhauser et al. 1991; Duncan and Hoffman 1985; Garfinkel and McLanahan 1986; Jarvis and Jenkins 1999; Smock, Manning, and Gupta 1999) and the consequences of childhood poverty (Duncan et al. 1998; Guo and Harris 2000). Put simply, children benefit from the economic and emotional investment of parents who reside together continuously, and these investments are generally higher among biological than among surrogate parents.

While children may therefore be better off residing in a cohabiting union formed by two biological parents than in a married household where one of the parents is not a biological parent, the current focus in survey research on marital status makes changes in the former living arrangement more conspicuous than changes in the latter. Research has shown that in spite of the economic benefits of stepfamilies relative to single-parent families (Morrison and Ritualo 2000), stepfamilies also suffer disadvantages associated with disruption following divorce or competition with a nonresidential biological parent (Cherlin 1978; Cherlin and Furstenberg 1994; Kiernan 1992). We still know little about the consequences for children of growing up in a de facto marriage when both biological parents are present (Manning and Lichter 1996; Smock 2000). Until we have good techniques for charting children's experiences of different types of families at both the macro and micro level, our understanding of the effects of family structure on child well-being will be incomplete.

This article contributes to the development of such techniques by tracing the effects of family change on children's family structure experiences in a number of Western countries. Substantively, we focus on children growing up with only one parent, since the literature to date indicates that this is the living arrangement that most profoundly affects child well-being. We study both the incidence and the duration of living with only one parent, and the respective contributions of out-of-wedlock fertility and parental separation to children's exposure to a single-mother household. We take into account parental cohabitation at birth, which leads nonmarital fertility ratios to overstate the incidence of single parenthood at birth, and we estimate whether parental cohabitation is more likely than parental marriage to dissolve before the end of childhood. We also take into account parental "repartnering" and estimate the reduction it provides in the duration of life with only one parent.

While we aim primarily at addressing these descriptive challenges, our cross-national scope also provides insight into the underlying causes of the observed changes. As exemplified years ago by the European Fertility Project, the convergence in demographic behavior among countries that differ widely with respect to one alleged cause of that behavior calls for a reframing of extant theories of demographic change. Interestingly, the first demographic transition in Europe has been characterized by Watkins (1991) as one of demographic integration within countries, with the gradual fading of provincial idiosyncrasies between 1870 and 1960. Wilson (2001) also describes a global demographic convergence between countries for the second half of the twentieth century. Yet, many authors expect national demographic differences to persist in view of the deep historical roots of family patterns (e.g., Reher 1998) and the enduring differences in welfare systems (Esping-Andersen 1990) that likely affect family behavior.

Our findings indicate that the second demographic transition exhibits little sign of convergence because the decline in marriage, the increase in the prevalence of nonmarital cohabitation with children, and changes in family “reconstruction” have each proceeded at quite different paces across countries. Perhaps the only universal Western trend is that childrearing is being shifted from married parents to single mothers more than to cohabiting parents, stepfamilies, or single fathers.

Data and methods

Data

The Fertility and Family Surveys (FFS) is an international sample survey program focusing on fertility and family change in the member countries of the United Nations Economic Commission for Europe. The list of participating countries includes over 20 European nations, as well as Canada, New Zealand, and the United States. The program coordinated the sample and questionnaire design of nationally representative surveys carried out by national statistical offices (Macura and Klijzing 1992). The first countries participating in the program contributed existing family surveys (e.g., Norway’s 1988 Family and Occupation Survey), while countries joining later attempted to fit the model survey instruments into ongoing data collection ventures. A common strategy was to use a particular cycle of an existing survey with core topics most similar to those of the FFS, appending ad hoc modules if necessary (e.g., Cycle 5 of the General Social Survey in Canada; the 1994 Annual Employment Survey in France; Cycle 5 of the National Survey of Family Growth in the United States). Given these diverse strategies, the years of data collection range from 1988 to 1998, the sampling designs differ (e.g., with respect to age range, inclusion of a male sample), and the questionnaires vary in content. Although the FFS data are imperfectly standardized, they represent an unparalleled source of information about differences in fertility and family trends across a number of Western countries.

For this article we analyzed female samples only, permitting the inclusion of several countries that did not interview males. For idiosyncratic reasons a few of the surveys could not yield the desired national life tables.[1] For the remaining 17 countries, required items are missing for only a small proportion of children. Sample sizes net of item nonresponses and internal consistency checks are shown in the Appendix Table (for a fuller analysis of data quality see Kveder 2002). The samples’ nonmarital fertility rates were also compared to official birth registration statistics to verify the reliability of the data.[2]

Methods

Children’s family structure—We reconstructed children’s family structure experiences by combining the partnership and fertility histories of the female FFS respondents. For up to nine cohabiting partners, respondents were asked the dates of coresidence (beginning and end), whether and how the partnership ended, and the date of marriage, if applicable. In addition, for each of up to 13 live births, respondents reported the date of birth, whether the

[1]The survey for Lithuania does not include questions about the coresidence of children at the time of data collection. The survey for Norway does, but the date when coresidence ended is available only for deceased children. The surveys for Bulgaria and Portugal provide information about only the most recent partnership, so we cannot assess the birth status of children born before that relationship.

[2]There is a substantial discrepancy in the proportion born out of wedlock between the estimate in German vital statistics and the estimate from FFS data (also noted in the FFS standard report for Germany). Thus, the results for Germany discussed below must be regarded with caution.

child was currently a coresident, and the date of and reason for the child's departure if not coresiding at the time of the interview. As long as the child was living with the respondent (i.e., the child's mother), we knew whether or not he or she was living with the male partner of the respondent and, if so, whether the couple was married. We limit our analyses to living arrangements involving the mother, with all other arrangements lumped into a single residual state. While this state comprises several distinct family structures (e.g., living with a single father, in a paternal stepfamily, with grand-parents), we found that less than 5 percent of childhood years, defined here as years from birth to exact age 15, are lived without the mother, which reduces the utility of making further distinctions within the residual state.

Figure 1 depicts the states analyzed below. Research on the impact of family structure on child well-being suggests beginning with the distinction between living with both parents and not doing so. Although the effect of parental marital status is less clear, married unions tend to be more stable, so we also account for parents' marital status. This allows us to study the family structure trajectories of children born to married versus cohabiting parents. For children whose parents live apart, we distinguish between living with the mother and not. Finally, when the child lives with his or her mother only, we distinguish between living with a single mother and living with a mother and her cohabiting partner (irrespective of marital status). Unfortunately, as is true of commonly used measures of family structure, the conventional nomenclature of family demography is poorly suited to describe this kaleidoscope of family forms efficiently. Figure 1 also defines the terms we use to denote the five states and three additional combinations of states. Most notably, by "single" mother we always imply "not in a partnership"; and by "both" parents (as opposed to "two" parents in a "two-parent family"), we refer to the two biological parents.

Multistate life table construction—There are two principal methods for constructing a multistate life table, one based on rates of transition between states, and another based on probabilities of transition (Rogers 1995). In most instances, transition probabilities cannot be directly estimated, so the former is the more commonly used method. With retrospective data, however, transition probabilities can be estimated directly, which greatly simplifies the calculation of the tables. More precisely, we estimate conditional probabilities of transition —that is, conditional on the survival of the mother. Given the low mortality rate of women in the age range sampled in the FFS (on the order of 1 per 1,000 per year), this should not deter us from applying this more straightforward technique.

Under the typical stationary assumptions of life table construction, rendered acceptable by the use of short age intervals, the survivorship ratios are estimated as:

$$\frac{{}^{i}_{n}N^{j}_{x}\left(t\right)}{{}_{n}N^{i}_{x-n}\left(t-n\right)}=\frac{{}^{i}_{n}L^{j}_{x}\left[t-n,t\right]}{{}_{n}L^{i}_{x-n}\left[t-n,t\right]} \quad (1)$$

where ${}_{n}N^{i}_{x-n}\left(t-n\right)$ is the number of children aged *x*−*n* to *x* and in state *i* at time *t*–*n*;

${}^{i}_{n}N^{j}_{x}\left(t\right)$ is the number of children aged *x* to *x* + *n* and in state *j* at time *t* who were in state *i* at time *t* – *n*;

${}_{n}L^{i}_{x-n}\left[t-n,t\right]$ is the number of child-years lived in state *i*, between age *x*−*n* and *x* in the period [*t*–*n*, *t*];

${}^{i}_{n}L^{j}_{x}\left[t-n,t\right]$ is the number of child-years lived in state *j*, between ages *x*−*n* and *x* in the period [*t*–*n*,*t*] by children who were in state *i* at time *t*–*n*; and *i* is any state and *j* is any state unless *i* is the absorbing state (in which case *j*=*i*).

Using the above-mentioned assumptions, we reconstructed children's living arrangements from birth to the time of the survey and calculated the quantities ${}_nN^i_{x-n}[t-n]$ and ${}^i_nN^j_x(t)$ at any time *t* before the survey. We then obtained the distribution of child-years lived across states between ages *x* and *x* + *n* from the distribution of child-years lived across states between ages *x*–*n* and *x*, using the equation above and the following identity:

$$ {}_nL^j_x[t-n,t]=\sum_i {}^i_nL^j_x[t-n,t] \quad (2) $$

Starting from any distribution of birth statuses, we derived sequentially the distribution in each three-year age group. We calculated four sets of life tables—three corresponding to one of the three possible statuses at birth (with married parents, with cohabiting parents, with a single mother), and a fourth using the observed distribution at birth in that period. To analyze recent trends, we estimated life tables for three-year periods (and in three-year age groups), that is, for the three years before the survey [*t*–*3*, *t*] and for previous three-year intervals [*t*–6, *t*–*3*], [*t*–*9*, *t*–*6*], [*t*–*12*, *t*–*9*], and [*t*–*15*, *t*–*12*]. These period life tables provide the expected number of years in different states at the transition rates observed during the period. We subsequently use the term "childhood expectancy," analogous to life expectancy in standard mortality life tables, to refer to the expected total number of years an average member of a synthetic birth cohort would spend in a given living arrangement between birth and age 15.

Limitations

International survey research always raises issues of data comparability. The information culled from the FFS (age, dates of birth, coresidence) is reasonably objective, and therefore less prone to the different meanings and interpretations that can hamper comparative research on attitudes, for example. However, these data also have several limitations. First, for each child information was provided on only one departure from the maternal household, if one had occurred before the interview. Even when the survey included a male sample, that sample was independent from the female sample, precluding the complete reconstruction of a child's living arrangements for children who did not continuously stay with one of the respondents. Heuveline and Timberlake (2002) provide a method to "splice" together information obtained from the male and the female samples and to estimate aggregate life tables with transitions between maternal and paternal households. This more complex approach yields results that are not numerically different enough from those presented here to justify this added complexity. The reporting of, at most, only one move also requires us to assume that children continuously coresided with their mother from birth until either the time of the survey—if they coresided then—or the date reported as the end of coresidence. Since reentries into households are not reported, we also had to assume that a child who leaves the maternal household before age 15 remains out of that household through age 15; that is, we treat the residual state as an "absorbing" state. Under stable conditions, the two biases would exactly balance out, and in any event the total bias is likely to be small given the low incidence of leaving the maternal household in the first place.[3]

[3]To confirm this, we used the available male samples to estimate childhood expectancy of living in a paternal (i.e., nonmaternal) household. Were there any systematic bias in our estimation from the female samples of childhood expectancy of not living with the mother, we should find the opposite bias when estimates are derived from the male samples. On the contrary, we obtained results that were nearly identical and consistently low (less than one year by age 15 being spent away from the maternal household).

A more nettlesome limitation of the maternal data is the absence of positive identification of a child's father. As with most surveys, the FFS were designed to measure marital rather than parental status; therefore we were compelled to develop rules to distinguish children living with both biological parents from children living with an unrelated male who cohabits with the biological mother. If a child was born while the biological mother was in a cohabiting partnership (married or not), we assumed that the partner fathered the child. For children born outside of a partnership, we used the timing of the birth and the next union formation to distinguish between a parental union and another partnership. If the next partnership was formed within six months of the birth or if the mother married within a year of the birth, we coded this partner or spouse as the child's father. Although we could not find data to externally validate this rule, it is well documented that the likelihood of forming a new partnership increases sharply after an out-of-partnership birth (e.g., Brien, Lillard, and Waite 1999). Any rule based on timing will necessarily create some false assignments; however, the numerical impact of these false assignments is likely to be low because the vast majority of children are born within a partnership, even in recent years. Simulations from the United States, the country with the highest proportion of out-of-partnership births, suggested that the proportion of recent birth cohorts experiencing a postnatal parental union varied between 2.3 percent and 3.4 percent depending on the identifying rule applied.[4]

More general concerns associated with the use of retrospective data must also be addressed. First, retrospective data are subject to recall errors, although the more salient the reported events are to the respondent, the lower the chances of recall errors. Dates of birth (of self and own children) and marriage are among the most accurately reported items in retrospective surveys, especially by women (Poulain, Riandey, and Firdion 1991). Retrospective reports on the incidence and timing of cohabitation are less reliable, so it is possible that some early and short-lived partnerships might have gone unreported (Casper and Cohen 2000; Murphy 2000). Their omission would tend to bias estimates of the incidence of children's transitions between various family structures. On the other hand, if such partnerships ended before a child's birth, their omission would not affect his or her family structure experience. Furthermore, as long as respondents tend to forget the shortest partnerships, their omission should not contribute much bias to duration measures.

Finally, retrospective data on children are subject to selectivity biases with respect to maternal age at birth (Rindfuss, Palmore, and Bumpass 1982). As shown in the Appendix Table, the upper age limit of women interviewed across national samples varies appreciably: nine countries had 50 years or older as their upper limit, but the other eight had upper limits ranging from 40 to 49 years. In calculating a three-year period life table up to age 15, the last survivorship ratios estimated with equation (1) above include 9- to 12-year-olds at the beginning of the period, becoming 12- to 15-year-olds at the end of the period. Hence, the youngest children contributing to the estimates were born 12 years before the end of the period. With an upper age limit for female respondents of 40 years, for instance, the last survivorship ratio is estimated only from children born to mothers under age 28 in the most recent period life table, under age 25 for the previous one (three to six years before the survey), and so on.

[4]The US data include 2,421 births three to six years before the survey. Of these births, 521 were not born in a partnership and, of these, 205 experienced at least one partnership formation by their mother before the survey, three to six years later. The problem of identifying whether the mother's first postnatal partner was in fact the child's father thus concerned 8.5 percent of the birth cohort. Given our allocation rule that combines timing and marital status, we estimated that 59 children (2.4 percent) experienced their parents' forming a partnership. Had we used a stricter timing rule of six months regardless of marital status, the estimate would be 56 children (2.3 percent). With a more liberal timing rule of one full year regardless of marital status, the estimate would be 78 children (3.2 percent), and with an additional six months in the case of marriage the estimate would be 82 children (3.4 percent). Although the uncertainty about the exact value is unfortunate, the numerical effect on the average estimates for a birth cohort is limited even in a country where out-of-partnership births and new partnership formations are prevalent.

This causes selection problems because younger mothers are more likely to give birth out of wedlock (Morgan and Rindfuss 1999) or in unstable partnerships. For the United States, Bumpass and Lu (2000) found that children born to mothers under age 24 can expect to spend much less time with a married mother than children of mothers aged 24 to 26. We therefore used age 25 as the maternal age at birth threshold below which we considered the estimates too biased. Since 40 years is the lowest upper age limit of respondents across countries, we could compute at least one period life table up to age 15 in each country. When we computed change over time, however, we gradually lowered the last age group of the life table of children's living arrangements, so that the data did not come only from children born to mothers under age 25. When the upper age limit of respondents is 40, as in Germany, we estimated only one period life table up to age 15, with the previous three-year period table ending at age 12, the one before at age 9, and so on. When the age limit is 55 or higher, as in Austria, Canada, and New Zealand, we computed four period life tables up to age 15 without risking substantial selectivity biases.

Results

Exposure to single parenting: The predominance of parental separation

We begin by analyzing the two main childhood routes to single parenting: parental separation and birth to a single mother. Although children experiencing parental separation may transit rapidly from the parental household to a stepfamily, we assume that these children experience a transitory period, however brief, during which they live with only one of their parents, typically the mother. Countries are ranked in Table 1 by childhood exposure to single parenting at early 1990s rates (column 6).[5] The nonmarital fertility ratio is presented in column 1.

Countries with low nonmarital fertility ratios—Italy, Spain, and Belgium—tend to have relatively stable parental unions, and therefore low overall childhood exposure to single parenting. At medium to high levels, the association between the ratio and childhood exposure to single parenting is attenuated by the large variance in the share of nonmarital fertility accounted for by parental cohabitation. At one extreme is Sweden, where 41.2 percent of all births are to cohabiting parents, compared to only 5.5 percent to single mothers (columns 2 and 3). Parental cohabitation also accounts for much of nonmarital fertility in several other European countries (Slovenia, Finland, France) and Canada. By contrast, in the United States more births are to single mothers (16.2 percent) than to cohabiting parents (10.7 percent). Single mothers also account for a substantial proportion of all births in New Zealand (12.6 percent), Austria (13.6 percent), and Germany (15.2 percent). While the Austrian exception within Europe has been documented previously (Prinz 1995), the estimates for Germany are inflated by the above-mentioned overestimation of nonmarital births in the FFS.

Once status at birth is adjusted to account for parental cohabitation, it becomes clear that parental separation is a more frequent route to single parenting than birth to a single mother. Two exceptions are Slovenia and Poland, where parental separation and birth to a single mother are both rare. Birth cohorts in the United States and New Zealand have the highest

[5]As mentioned above, the surveys were fielded in different years across countries (see Appendix Table). To make the results more comparable, we present either the most recent set of period tables (i.e., three years before the survey) or the previous set (three to six years before the survey). The most recent set is used when the national survey was fielded in 1991, 1992, or 1993. The previous set is used when the survey was fielded in 1994, 1995, or 1996, thus scaling the reference period back three years. All but two surveys fell within one of the two three-year windows: Finland's survey was fielded in 1989–90 (the most recent set is presented), and the Czech Republic's survey was completed in 1997 (we present the tables referring to three to six years before the survey). Except for these two outliers, the cross-national comparisons below all refer to a three-year period that includes January 1991, to which we refer for convenience as "the early 1990s."

combined proportions of children born to a single mother and children born to married or cohabiting parents who separate during childhood. Combining these two routes, at early 1990s rates, 51.3 percent of a birth cohort is expected to experience living with a single parent during childhood in the United States. The proportion is similar in New Zealand (49.0 percent, column 6). In both countries parental separation accounts for more than two-thirds of this childhood exposure. In the majority of European countries and Canada, the percentages of children expected to experience parental separation range from the low 20s to the low 30s, often four to five times higher than the percentage born to a single mother.

Cohabitation and marriage from the perspective of children

It is clear that failing to account for parental cohabitation creates a distorted picture of the exposure at birth to living with a single mother. If parental cohabitations are highly unstable, however, the overestimation of the total childhood exposure to single parenting would not be large. We find in most countries that children born to cohabiting parents are two to four times more likely to see their parents separate than are children of parents married at the time of birth (column 7). Sweden again stands out: the likelihood that children born in a cohabitation experience the separation of their parents during childhood is only 30 percent greater than that of children born to married parents. The Swedish exception is only one of degree, however, since parental cohabitation is less stable than parental marriage in every country. Nevertheless, variation in the degree to which marriage relative to parental cohabitation “protects” children from parental separation complicates certain cross-country comparisons. Our results indicate, for example, that children born to cohabiting parents in Sweden are less likely to experience a parental break-up (column 4÷column 2 = 0.347) than children born to married parents in the United States (column 5÷[100–column 1] = 0.369).

Birth to cohabiting parents therefore has quite different effects across countries on the childhood probability of experiencing a parental separation. In Sweden, most children born to cohabiting parents never experience single parenting. Whereas 41.2 percent of Swedish children are born to cohabiting parents, we estimate that about a third of them (14.3 percent) experience parental separation, at early 1990s rates (see columns 2 and 4). The proportion of a Swedish birth cohort that is born out of wedlock and yet is expected to remain with both parents from birth to age 15 is the difference, or 26.9 percent. The corresponding percentages are markedly smaller in other countries, though not trivial in Finland (7.1), Austria (7.8), Canada (8.3), France (8.4), and Slovenia (10.6). However, this pattern of longstanding de facto marriages is not universal. In the United States, for example, the stability of cohabiting unions is far lower even when children are involved. Of the 10.7 percent of American children born to cohabiting parents in the early 1990s, a very large majority is expected to see their parents separate by age 15 (8.1 percent of a birth cohort). Thus, in the United States, parental cohabitation merely postpones the experience of single parenting to later childhood years. This expectation is similar for children born to cohabiting parents in Latvia and, to a lesser extent, in New Zealand and the Czech Republic.

Beyond exposure: Other partnerships and childhood expectancy of living with a single mother

The duration of single parenthood is related to public costs and perhaps private costs to parents, children, and extended kin. Hence, it is crucial to look beyond incidence and analyze the duration of children’s coresidence with a single parent, overwhelmingly the mother. Table 2 first compares childhood expectancy of living with a single mother across birth statuses (columns 1 to 3). For each country, the first two columns indicate a longer childhood expectancy (two to four times longer in most countries) of living with a single mother for children born to cohabiting rather than to married parents. Even in Sweden, where the parental cohabitation-to-marriage ratio of exposure was smallest (1.30), the ratio of duration is nearly two (2.17÷1.09). In the United States, with a more typical exposure ratio (2.05), the ratio of duration exceeds three (3.95÷1.28). Within countries, the duration ratios are higher than the exposure ratios, indicating that children born to cohabiting parents are more likely both to see their parents separate and to see them separate sooner than children born to married parents.

Because exposure starts at birth, children born to single mothers can expect to live longer with single mothers than children of other birth statuses—in a majority of countries spending more than half of their childhood with single mothers (column 3). Childhood expectancy of living with a single mother for children born to single mothers is shortest in countries where these children are more rare (Spain 4.47 years, Italy 4.50 years, and Slovenia 4.52 years), suggesting social pressure to raise children within partnerships even if they were conceived outside of a partnership. In contrast, the childhood expectancy of living with a single mother for children born to single mothers exceeds two-thirds of childhood years in Germany (11.67 years), Belgium (11.06 years[6]), and Poland (10.24 years). In each of these countries, childhood expectancy of living with a single mother is less than one year for children born to married parents; thus, childhood living arrangement experiences are highly conditioned by birth status. Overall, out-of-partnership fertility accounts for a larger share of a birth cohort's average expected duration in single-mother households than it does for the percentage of a birth cohort ever exposed to living in such households. Nevertheless, children born to single mothers contribute more than half of the years that children spend with a single mother in only three countries, Germany, Poland, and the United States.[7]

The country rankings in Table 2 are based on the total childhood expectancy of living with parents apart (column 7). Columns 4 through 6 decompose this total into childhood expectancy of living with a single mother, in a stepfamily, and not with the child's mother. New Zealand and the United States again stand out with more than a third of childhood years expected to be with parents apart (5.08 and 5.12 respectively). The difference between the expected childhood exposure to single parenting (about 50 percent in these two countries) and the expected proportion of childhood years spent with a single mother reflects the fact that exposure frequently occurs several years after birth, through parental separation.

Because maternal repartnering (column 5) is more prevalent in the United States than elsewhere,[8] childhood expectancy of living with a single mother (2.70 years, column 4) is shorter than in New Zealand (2.96 years) and nearly the same as in Germany (2.69 years), despite a longer childhood expectancy of living with parents apart. Not coresiding with the mother is quite rare. The longest childhood expectancy not with mother is 0.71 years in New Zealand (column 6), less than 5 percent of the first 15 years. In all countries, living with a single mother accounts for the largest share of childhood expectancy of living with parents apart. Across countries, the ratio of childhood expectancy of living with a single mother to total childhood expectancy of living with parents apart varies between 43 percent and 70 percent (column 8). In seven of the 17 countries, childhood expectancy of living with a single mother reaches two to three years at early 1990s rates. In sum, at the time of the survey, living with a single mother was the most common alternative to living with married

[6]Because of a smaller sample size and lower proportion of births to single mothers, the estimates for Belgium are relatively unstable.
[7]These results are not shown here but can be obtained by weighting the estimates in Table 2, columns 1 to 3, by the corresponding proportions of a birth cohort in each birth status from Table 1, columns 1 to 3 (note that in the latter table the percentage born to married parents is obtained by subtracting column 1 from 100).
[8]In particular, high numbers of sequential transitions appear rare outside the United States. By age 15, 11.7 percent of American children in the FFS had lived in three or more parental partnerships. The second-highest proportion in the FFS was 3.1 percent in Sweden. These proportions are estimated directly on all uncensored observations, i.e., children over age 15 at the time of the interview and still living with their mother at age 15 (results not shown). Life table estimates might differ from these direct estimates, which nevertheless suffice to illustrate the uniqueness of the United States in this respect.

parents. This could be due to the relatively recent emergence of other two-parent families, such as cohabiting parents and stepfamilies.

What offsets the declining proportion of childhood spent with married parents?

To analyze within-country trends over time, we compare the most recent three-year period life table with the table corresponding to an earlier three-year period. We focus on childhood expectancy across the four states that we believe best reflect underlying family structure transitions. Because of the variable severity of the selectivity concerns discussed above, in Table 3 we used somewhat different time intervals (column 1) and upper age limits (column 2) to generate the within-country trends. Although necessary to reduce selectivity bias, this strategy complicates cross-national comparisons. Columns 3 to 6 in Table 3 hence provide annualized rates of change in childhood expectancy of living in each family structure. These rates standardize the pace of change, independent of the age limit or time interval used in the comparison.[9] Absolute changes (in childhood years) during the period are presented in columns 7 to 10.

Countries are ranked in Table 3 by the pace of the decline in childhood expectancy of living with married parents (column 3). Quite rapid declines, between 1 percent and 3 percent annually, are found in Latvia, France, Canada, New Zealand, and Austria. These rates of change reflect absolute declines in childhood expectancy of living with married parents of 2.17 years over a 15-year period in Austria to 2.40 years over a nine-year period in Latvia (column 7). In eight other countries, annual rates of decline averaged between 0.4 percent and 0.6 percent. In the remaining four countries (Spain, Italy, Switzerland, and Sweden), childhood expectancy of living with married parents was more stable.

What changes in other family structures have been concurrent with the decline in childhood expectancy of living with married parents? In the five countries where these declines were most rapid, childhood expectancy of living with cohabiting parents increased rapidly, from 6 percent per year in New Zealand to 14 percent per year in Canada (column 4). Overall, childhood expectancy of living with cohabiting parents was on the rise in nearly every country (columns 4 and 8), with France and Canada experiencing the largest absolute increases (a little over one year). However, these increases were not large enough to substitute fully for declines in childhood expectancy of living with married parents, resulting in overall declines in childhood expectancy of living with both parents (i.e., irrespective of marital status). Even in France and Canada, the increase in childhood expectancy of living with cohabiting parents represents only one-half of the decline in childhood expectancy of living with married parents.

Childhood expectancy of living with cohabiting parents declined slightly in the years before the survey in Sweden, where it had reached its record duration in the early 1990s. The Swedish trend seems to be linked to a change in pension policies in 1990 that induced cohabiting parents to marry, causing a temporary increase in marriages. The impact of this policy change is also visible in the annual nonmarital fertility rate, which dropped in 1990 and did not return to its 1989 level until 1994 (Sardon 2000). The decrease in childhood expectancy of living with cohabiting parents between the pre- and the post-1990 period partially reflects this temporary marriage surge. However, the decrease in childhood expectancy of living with cohabiting parents still exceeds the increase in childhood

[9]The following illustrates how this standardization operates: we computed the estimates of changes over the past 15 years with censoring at ages 3, 6, 9, 12, 15, and 18 in New Zealand, the country with the highest upper age limit and thus the fewest selectivity concerns. At these different ages, the annual rates of decline in time spent with married parents are 1.51 percent (at age 3), 1.53 percent (at ages 6, 9, 12, and 15), and 1.41 percent (at age 18). In spite of the standardization, the very young upper age limit may slightly bias international comparisons because parental separation is relatively less likely to occur in the first few years after birth.

expectancy of living with married parents. Thus, even in Sweden, childhood expectancy of living with both parents declined between the two periods. In sum, across countries there appear to be limits to the extent to which parental cohabitation is substituting for parental marriage.

Stepfamilies appear to constitute, on average, an even less substantial alternative. In the early decades of rising divorce rates, family sociologists speculated that stepfamilies would become more prevalent, knitting different households in complex networks resembling “new extended families.” The actual changes of the past several decades proved this sanguine vision to have been mistaken. In the years before the FFS, increases in childhood expectancy of living in stepfamilies were visible in only two countries (Canada, 0.47 years and Austria, 0.64 years, column 10), and amounted to only a fraction of the decreases in childhood expectancy of living with married parents. Also, like parental cohabitation in Sweden, childhood expectancy of living in stepfamilies declined slightly in the years before the survey in the United States, where it had reached its highest level in the early 1990s. The results indicate that the prevalence of stepfamilies was even less able than parental cohabitation to expand when childhood expectancy of living with married parents declined.

In countries where childhood expectancy of living with married parents declined fastest, increases in childhood expectancy of living in alternative forms of two-parent families were not sufficient to compensate fully for the decline in marriage. As a result, childhood expectancy of living with a single mother increased in these countries (column 9). In Latvia and New Zealand, for example, the bulk of the decline in childhood expectancy of living with married parents was translated into an increase in childhood expectancy of living with a single mother. In France, where childhood expectancy of living with cohabiting parents increased but that in stepfamilies did not, about one-half of the decline in childhood expectancy of living with married parents was transferred into increased childhood expectancy of living with a single mother. In both Latvia and France, childhood expectancy of living with a single mother increased at an annualized rate of more than 7 percent per year (column 5), equivalent to a doubling time of less than ten years. Had that pace continued to the present, childhood expectancy of living with a single mother would now be longer in these two countries than in the United States. Even in Canada and Austria, where childhood expectancies of living with cohabiting parents and in stepfamilies both increased, about one-third of the reduction in expected time with married parents was converted into childhood expectancy of living with a single mother.

While these trends are most visible in the five countries where childhood expectancy of living with married parents declined quickly, similar observations apply in the eight countries where the annualized rate of decline was more moderate (between 0.6 percent and 0.4 percent per year). In these countries, absolute changes in childhood expectancy of living in stepfamilies were negligible. Increases in childhood expectancy of living with cohabiting parents largely offset decreases in living with married parents in three central European countries (Czech Republic, Hungary, and Slovenia) and Belgium, whereas childhood expectancy of living with a single mother increased most in the United States, Finland, Poland, and Germany.

Discussion

Although it is widely acknowledged that, at least since the 1970s, marriage and divorce statistics have become increasingly flawed indicators of family structure, more appropriate data have not been collected frequently enough to trace the quickly changing contours of children’s family environments. The retrospective data from the Fertility and Family Surveys provide a unique opportunity to compare children’s family structure experiences during a time of transition in Western countries. At the rates occurring in the early 1990s, we estimate the proportion of a birth cohort expected to experience single parenting by age 15 to reach one-half in New Zealand and the United States. The expected proportions are lower elsewhere, but still exceed one-third in Canada and five European countries.

Parental separation, regardless of marital status, contributes more to childhood exposure to single parenting than does birth to a single mother. For that reason, and to a lesser extent because of the relative importance of stepfamilies, the share of childhood years expected to be spent with a single mother is substantially lower than the proportion of a birth cohort expected to experience single parenting during childhood. Nevertheless, the expected duration of the former approaches 20 percent (three years by age 15) in a few countries. For the United States, our estimate (2.70 years, or 18 percent, by age 15; see Table 2, col. 4) replicates Bumpass and Lu's (2000: 38) estimate of 20 percent by age 16, derived by similar techniques from the same data. Bumpass and Lu define different states of interest, however, dividing the remaining childhood years into 9 percent spent in cohabiting unions and 71 percent spent in marriage. We separate the remaining years before age 15 into 66 percent with both biological parents (since 34 percent—5.12 years by age 15—are with parents apart; Table 2, col. 7), nearly all of which occurs within marriage, and 12 percent in a maternal stepfamily (1.87 years by age 15; Table 2, col. 5), nearly half of which occurs within cohabitation. At early 1990s rates, childhood expectancy of living with parents apart reached five years in New Zealand and the United States, and up to four years in three of the other countries examined here.

A compelling reason to track the changing patterns of family formation is that they are likely to exert economic pressures on families and require policy interventions to help children and parents who may require added support. To the extent that instability in families creates greater hazards for children's development, it is essential to develop ways of discerning whether changes in patterns of family formation are relatively nominal (e.g., from official to de facto marriage) or potentially more consequential for children's welfare.

Even though the nonmarital fertility ratio misrepresents children's exposure to growing up with only one parent, its variation across countries and over time thus far has captured reasonably well the *direction* of temporal changes and cross-national differences in childhood exposure to single parenting (the correlation coefficient between our estimates of incidence in column 6, Table 1 and nonmarital fertility rates is 0.67.) This is true in part because, while some nonmarital births are to cohabiting parents, the incidence of separation for parents who were cohabiting at the time of birth is greater than if they were married at the time of birth. Another reason is less obvious, but equally important. We also find a strong association (a correlation coefficient of 0.59) between the nonmarital fertility rate and the risk of parental divorce before the child reaches age 15 among married parents, which, as we have shown, is still the most frequent route to single parenting. It appears that the social conditions that lead individuals to be hesitant about entering marriage before having children are also associated with greater levels of marital instability among couples who do enter matrimony. The cultural and institutional accommodation to the expansion of single parenting hardly discriminates between divorced custodial parents and single (at birth) mothers.

We have not attempted to explain fully the variations that we have identified among the 17 countries examined here. Consistently standing out, New Zealand and the United States are two of only three English-speaking countries included in our analyses. At the other end of the distributions of single childbearing and likelihood of parental separation stand three Mediterranean countries: Italy, Spain, and Slovenia. We suspect it is no coincidence that these three countries are also those with some of the lowest fertility levels at the time of the

FFS. Their total fertility rates were 1.27, 1.27, and 1.36 children per woman (United Nations 2001), while New Zealand (2.06) and the United States (2.05) were the two countries with the highest total fertility rates among the 17 countries examined in this article. To the extent that increases in out-of-partnership childbearing and parental divorce reflect divergence from traditional family living arrangements, it seems likely that the countries maintaining traditional practices of family formation do so by postponing fertility to later ages. In fact, the correlation between our estimate of the average duration spent without two parents in the early 1990s (column 7 of Table 2) and the 1990–95 total fertility rates is 0.65.

This is but one plausible possibility in accounting for the large variation in both the pace and the pattern of change that we observed in our analysis. There are many other possible explanations for why different countries are characterized by different family formation strategies. Longstanding historical differences related to cultural preferences undoubtedly play a part in the process (Reher 1998). Similarly, we expect that public policies designed to support these differing cultural values also affect the tempo of change and the type of family formation patterns that emerge across countries. The three countries that appear here as exemplars of a traditional family structure (Italy), an expansion of parental cohabitation (Sweden), and an increase in single parenting (the United States) also typify Esping-Andersen’s (1990) three categories of capitalist welfare states: the conservative, the social democratic, and the liberal. Yet, much diversity in family behavior remains to be explained within the “conservative” welfare states of continental Europe—for example, between Italy, France, and Germany.

Complicating matters further, our estimates of change during the past decade indicate that we are still in the midst of the second demographic transition. Some countries continue to experience rising levels of nonmarital childrearing both within and outside of de facto marriage. Divorce and remarriage rates continue to result in considerable flux in children’s living arrangements. It is still too early to tell whether the end points of the second demographic transition are in sight; thus, it is still too soon to tell whether countries will eventually converge or whether they will cluster in different cultural or economic categories (Kuijsten 1996). At this point, the evidence points toward the latter alternative, with the possible exception that—at paces that depend on the stability of marriage, the expansion of parental cohabitation, and the prevalence of family reconstruction—childrearing is increasingly being shifted to single mothers. In other words, while children who do not live with married biological parents could in principle live in other two-adult families, most do not or do so only temporarily. Childhood expectancy of living with a single mother remained just under three years at early 1990s rates; but in a few countries, if the increases observed in the years just before the survey were to continue unabated, this expectancy would double within a decade.

It is abundantly clear from this and related research that we cannot continue to cling to the traditional categories for measuring change in marriage and childbearing. Accordingly, surveys must begin to produce data that are amenable to the family living arrangements that currently exist, rather than to the forms observed in the past. In so doing, we will advance our understanding of these demographic changes and be in a better position to evaluate policy options aimed at promoting children’s welfare.

APPENDIX TABLE: Survey dates, upper age limits of women interviewed, sample sizes, and data quality checks, by FFS survey

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Country	Survey dates			Number of own children reported			NMFRs[a] (percent)	
	From	To	Upper age limit	Total	With complete data	Percent complete	Observed from FFS sample	Recorded in vital statistics
Austria	Dec-95	Apr-96	55	7,043	7,032	99.8	33.8	27.4
Belgium[b]	Mar-91	Dec-92	42	3,781	3,710	98.1	5.9	7.1
Bulgaria[c]	Dec-97	Dec-97	49	2,794	—	—	—	—
Canada	Jan-95	Dec-95	55	5,490	5,112	93.1	27.8	24.4
Czech Republic	Nov-97	Dec-97	45	2,364	2,363	100.0	14.2	n/a
Finland	Aug-89	Jan-90	53	6,223	6,213	99.8	16.3	20.8
France	Jan-94	May-94	51	4,525	4,525	100.0	31.8	33.0
Germany	Jul-92	Jul-92	40	6,024	5,748	95.4	26.0	15.3
Hungary	Dec-92	Dec-93	42	4,935	4,933	100.0	10.2	13.1
Italy	Nov-95	Jan-96	50	5,501	5,479	99.6	6.3	7.4
Latvia	Sep-95	Oct-95	50	3,903	3,894	99.8	26.1	29.9
Lithuania[c]	Oct-94	Oct-95	50	3,801	—	—	9.3	12.6
New Zealand	Sep-95	Nov-95	60	5,978	5,847	97.8	34.9	n/a
Norway[c]	Oct-88	Jun-89	43	4,914	—	—	33.6	32.4
Poland	Dec-91	Dec-91	49	7,024	6,947	98.9	12.1	6.2
Portugal[c]	Apr-97	Jul-97	50	4,920	—	—	—	—
Slovenia	Nov-94	Nov-95	46	4,005	3,997	99.8	24.2	29.8
Spain	Nov-94	Nov-95	50	5,065	5,065	100.0	7.3	10.8
Sweden	Oct-92	May-93	45	4,689	4,685	99.9	48.3	47.0
Switzerland	Oct-94	Jun-95	50	5,441	5,427	99.7	7.0	6.4
United States	Jan-95	Oct-95	44	14,847	14,749	99.3	28.9	32.6

[a] Nonmarital fertility ratios (NMFRs) computed from FFS data include all live births reported by female respondents to have occurred in the year(s) the survey was undertaken and in the last three full years before the survey started. In Austria, for instance, the survey was fielded in 1995–96, so all 1992–96 births are included. They are compared with the closest calendar year statistic available from national birth registration (Source: FFS Standard Reports, Table 1).

[b] The survey for Belgium covers only the Flanders region.

[c] In these countries, data to estimate childhood trajectories were not available.

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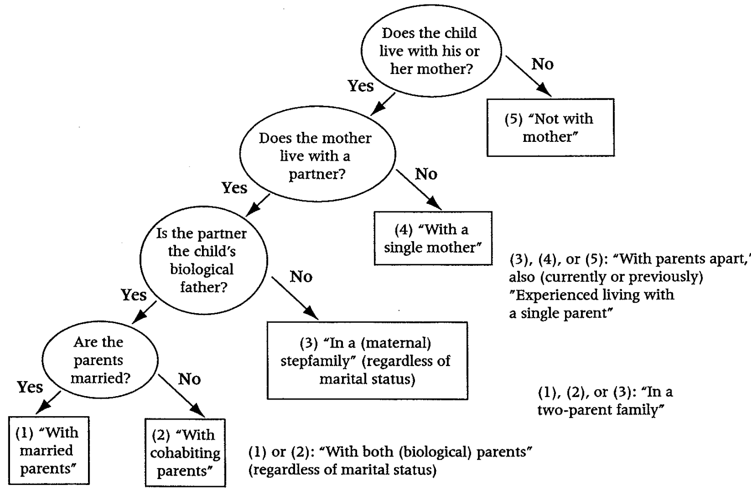

FIGURE 1.

Definition of childhood living arrangements and shorthand descriptive terms

TABLE 1

Childhood exposure to single parenting (from birth to age 15), by child's birth status: Children of the FFS female respondents (in percent)

	Status at birth		Childhood exposure to single parenting				Relative risk of parental seperation: cohabitation vs. marriage
			Born to a single mother	Born to a two-parent family			
	Out of wedlock	Cohabiting parents		Cohabiting parents	Married parents	Total exposure	
Country	**(1)**	**(2)**	**(3)=(1)–(2)**	**(4)**	**(5)**	**(6)=(3)+(4)+(5)**	**(7)[a]**
Italy	6.3	4.1	2.2	1.1	7.6	10.9	3.50
Slovenia	18.9	12.1	6.8	1.5	5.3	13.6	1.93
Spain	6.5	3.4	3.1	2.6	9.2	14.9	7.75
Belgium	6.4	4.9	1.5	3.1	12.4	17.0	4.77
Poland	12.1	2.4	9.7	0.5	8.2	18.4	2.14
Switzerland	7.4	4.4	3.0	2.4	17.3	22.7	2.98
Finland	16.9	13.8	3.1	6.7	16.0	25.8	2.50
Hungary	11.5	7.1	4.4	5.1	17.8	27.3	3.55
France	25.6	21.3	4.3	12.9	11.7	28.9	3.85
Sweden	46.7	41.2	5.5	14.3	14.2	34.0	1.30
Canada	24.1	15.8	8.3	7.5	18.7	34.5	1.92
Czech Republic	13.2	7.8	5.4	5.0	24.4	34.8	2.28
Germany	25.9	10.7	15.2	5.5	18.6	39.3	2.05
Austria	30.7	17.1	13.6	9.3	16.9	39.8	2.22
Latvia	19.3	10.5	8.8	7.8	24.3	40.9	2.47
New Zealand	31.0	18.4	12.6	13.9	22.5	49.0	2.40
United States	26.9	10.7	16.2	8.1	27.0	51.3	2.05

NOTES: Countries are listed in ascending order according to total exposure to single parenting (shown in col. 6). See Figure 1 for definitions of the labels for columns 3 to 6. Columns 1 to 3 are observed from FFS data. Columns 4 and 5 are derived from synthetic cohorts at early 1990s rates.

[a]The formula for column 7 is: (column 4÷column 2)÷(coIumn 5÷[100-column 1]).

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TABLE 2

Childhood expectancy of the duration of living in selected family structures (in years, from birth to age 15) by child's birth status: Children of the FFS female respondents

	With a single mother (no partner)							
Country	**Born to married parents (1)**	**Born to cohabiting parents (2)**	**Bom to a single mother (3)**	**All births (weighted average) (4)**	**In a maternal stepfamily (5)**	**Not with mother (6)**	**Total with parents apart (7)=(4)÷(5)÷(6)**	**Ratio with a single mother to total with parents apart (8)=(4)÷(7)**
Italy	0.43	0.60	4.50	0.52	0.16	0.13	0.81	0.64
Spain	0.56	3.11	4.47	0.72	0.35	0.07	1.14	0.63
Slovenia	0.32	0.63	4.52	0.61	0.55	0.09	1.25	0.49
Belgium	0.58	3.82	11.06	0.82	0.53	0.06	1.41	0.58
Switzerland	0.78	1.43	8.04	1.03	0.36	0.31	1.70	0.60
Poland	0.45	1.00	10.24	1.41	0.34	0.28	2.03	0.69
Hungary	1.04	2.87	6.86	1.46	0.68	0.26	2.40	0.61
France	0.75	3.02	8.82	1.55	0.76	0.13	2.44	0.64
Finland	0.93	2.85	8.14	1.44	0.76	0.31	2.50	0.57
Sweden	1.09	2.17	9.40	2.08	0.75	0.33	3.16	0.66
Czech Republic	1.04	2.11	5.05	1.35	1.71	0.12	3.18	0.43
Canada	1.31	4.03	9.20	2.38	0.93	0.08	3.39	0.70
Austria	1.32	2.14	7.76	2.32	1.36	0.26	3.94	0.59
Latvia	1.47	5.50	4.89	2.14	1.57	0.26	3.97	0.54
Germany	0.89	3.00	11.67	2.69	1.20	0.10	3.99	0.67
New Zealand	1.44	4.29	9.78	2.96	1.41	0.71	5.08	0.58
United States	1.28	3.95	8.56	2.70	1.87	0.56	5.12	0.53

NOTES: Countries are listed in ascending order according to the expected duration of living (in years) with parents apart (shown in col. 7). See Figure 1 for definitions of the labels for columns 1 to 6. Columns 1 to 6 are derived from synthetic cohorts at early 1990s rates. The difference between 15 years and the value in column 7 is the childhood expectancy of the duration of living with both biological parents. In the United States, e.g., that expectancy is 9.88 years.

TABLE 3

Annualized rate of change and absolute change in childhood expectancy of living in selected family structures: Children of the FFS female respondents

Country	Time interval (1)	Upper age limit (2)	Annualized rate of change (percent) of living: With married parents (3)	With cohabiting parents (4)	With a single mother (5)	In a step-family (6)	Absolute change (in years) of living: With married parents (7)	With cohabiting parents (8)	With a single mother (9)	In a stepfamily (10)
Latvia	9	15	−2.7	9.1	7.4	−0.3	−2.40	0.35	2.05	−0.03
France	9	15	−2.3	11.8	7.7	−2.6	−2.35	1.16	1.23	−0.16
Canada	15	15	−1.6	13.7	3.0	3.4	−2.60	1.12	0.91	0.47
New Zealand	15	15	−1.5	5.9	5.1	−1.0	−2.30	0.58	1.78	−0.18
Austria	15	15	−1.4	10.0	1.5	3.7	−2.17	0.83	0.54	0.64
Slovenia	9	12	−0.6	6.1	2.8	0.4	−0.49	0.51	0.16	0.02
United States	9	9	−0.6	0.2	2.0	−1.2	−0.33	0.01	0.34	−0.08
Finland	12	15	−0.5	2.7	4.1	−0.6	−0.66	0.25	0.56	−0.06
Poland	9	15	−0.5	17.0	2.4	0.4	−0.52	0.22	0.28	0.01
Czech Republic	9	9	−0.5	18.2	1.9	−2.1	−0.38	0.33	0.09	−0.07
Hungary	9	6	−0.5	13.1	2.9	0.1	−0.23	0.18	0.09	0.00
Belgium	9	6	−0.4	13.6	2.7	15.2	−0.20	0.13	0.04	0.06
Germany	9	6	−0.4	0.0	3.7	−6.6	−0.16	0.00	0.29	−0.12
Spain	9	15	−0.2	12.7	−3.5	−5.7	−0.28	0.30	−0.24	−0.16
Italy	9	15	−0.1	0.6	4.7	−4.9	−0.10	0.01	0.21	−0.06
Switzerland	9	15	0.0	−0.5	2.2	−1.6	−0.03	−0.01	0.16	−0.08
Sweden	9	9	0.3	−2.5	3.2	−2.4	0.14	−0.42	0.27	−0.06

NOTES: Countries are listed in ascending order according to the annualized rate of change of living with married parents (shown in col. 3). See Figure 1 for definitions of the labels for columns 3 to 10. Estimates are based on different synthetic cohorts. In each country, a synthetic cohort to which are applied the period rates of the three years before the survey—see Appendix Table for survey dates—is compared with a synthetic cohort to which are applied the rates of an earlier three-year period. The time difference is shown in Column 1 above. Since we are comparing three-year period rates, a nine-year comparison is between the rates of the period three years before the interview date and the rates of the period nine to 12 years before the interview date. When the age range of respondents permitted, we went back to 12 years earlier (in one country) or 15 years earlier (in three countries) without having to censor the life table below age 15. In some other countries, however, selectivity concerns led us to limit the comparison to up to age 12 (one country), age nine (three countries), or even age six (three countries) in order to maintain a nine-year time interval (as shown in column 2).