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The Divorce Cycle in Cross-National Perspective: Results from the Fertility and Family Surveys

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The divorce cycle in cross-national perspective: results from the Fertility and Family Surveys[1]

Abstract

We use data on women's first marriages from the Fertility and Family Surveys to analyse the intergenerational transmission of divorce—the divorce cycle—across 18 countries and seek explanations for the cross-national variation from macro-level characteristics. Our results show that women whose parents have divorced have a significantly higher risk of divorce in 17 countries. Furthermore, using multilevel models, we find cross-national variation in these associations. This variation is negatively associated with the share of women in each cohort who experienced parental divorce and the levels of female labour market participation. We conclude that increases in the number of peers who experience parental divorce weakens its signal of low marital commitment that is passed to children of divorce, thus weakening the divorce cycle.

Keywords: divorce, parental divorce, cross-national comparison, event history analysis, multilevel analysis, Fertility and Family Surveys

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Introduction

Social inheritance takes many forms, one of which is the higher than average likelihood of those with divorced parents to divorce themselves, the divorce cycle (Wolfinger 2005). This empirical regularity has been established across a wide range of countries, including the United States (Glenn and Kramer 1987; Amato 1996; Wolfinger 1999; Wolfinger 2005), the Netherlands (Dronkers 1997), the former West (Diekmann and Engelhardt 1999) and East Germanies (Engelhardt et al. 2002), France (Traag et al. 2000), and several other nations (Diekmann and Schmidheiny 2004).

Researchers have given several explanations to this finding. Divorce can have negative psychological effects on children (Amato and Keith 1991; Cherlin 1999) and children of divorce generally show lower interpersonal skills than those from intact marriages (Amato 1996). These effects may translate into more problematic marriages. However, one would expect that high parental conflict―which need not always lead to divorce―would produce similar outcomes, but empirical results suggest that people from divorced backgrounds have a higher likelihood of divorce than those raised by unhappily married and conflicting parents, who did not divorce (Amato and Booth 1991; Amato and DeBoer 2001).

Alternatively, parental divorce can legitimize divorce in the eyes of children of divorce as a way of ending an unhappy relationship and reduce commitment to marriage (Amato and DeBoer 2001; Wolfinger 2005). Behavioural evidence for these mechanisms is hard to find, but attitudinal data shows that adult children of divorce hold more liberal attitudes towards divorce than others (Amato 1988; Trent and South 1992; Axinn and Thornton 1996). Parental divorce is sometimes followed by a decline in levels of living and an increase in geographical mobility. These factors account for part of the poorer long-term outcomes of single parenthood (McLanahan and Sandefur 1994) and may shape the success of the marriages of those from divorced backgrounds. People with divorced parents are also more likely to marry someone with a similar background (Wolfinger 2003), which further increases their chances of divorcing, or someone with lower attained education (Wolfinger 2005; Erola et al. 2007), which in many countries is a risk factor of divorce (Härkönen and Dronkers 2006a). Finally, since personality traits that increase the likelihood of divorce partly transmit socially and biologically to the offspring, the association between parental divorce and children's risk of divorce may not be causal.

While these factors may help understand how divorce transmits from one generation to the next, we need to examine these mechanisms in their social context to account for possible cross-national and cross-temporal variation in this association. For example, Diekmann and Schmidheiny (2004) reported wide cross-national variations while Wolfinger (1999; 2005) found that the intergenerational transmission of divorce had become weaker in the United States, and Engelhardt and associates (2002) found similar results for the former two Germanies.

The objective of this paper is to analyse the intergenerational transmission of divorce across 18 countries. We use data on the first marriages of 43,071 women from the Fertility and Family Surveys. We estimate discrete time event history models separately for each country and compare the effects across countries. We then examine macro-level correlates of the divorce cycle with multilevel discrete time event history models using 16 for which we have macro-level data. Before proceeding to the empirical analyses, we next discuss reasons to expect cross-national variation in the divorce cycle and after that, present the data and the methods.

Reasons to expect cross-national variation in the divorce cycle

Macro-level factors can affect the divorce cycle by shaping the circumstances in which the child experiences the divorce of her or his parents or the environment of her or his own marriage. However, due to heavy collinearity, including both factors into the same statistical model is all but straightforward, so practical reasons often lead researchers to focus on just one life stage. Previous research has concentrated on the context of the parental divorce (Wolfinger 1999; Engelhardt et al. 2002; Diekmann and Schmidheiny 2004), and we will follow suit in this paper.

Wolfinger (1999, p. 415) expected that the association between parents' and children's divorce in the United States would weaken due to increasingly relaxed attitudes toward divorce. He discussed two mechanisms that should translate a more liberal social environment toward divorce into a weaker divorce cycle. Firstly, adding to the micro-level explanations outlined above, he argued that when divorce is more accepted, divorce carries less stigma and stress which may then reduce its negative consequences for children with divorced parents. Amato and Keith (1991), Engelhardt and associates (2002), and Diekmann and Schmidheiny (2004) made similar arguments concerning cross-national variation. However, Sigle-Rushton and colleagues (2005) did not find that the effects of parental divorce on educational, psychological, and economic outcomes weakened across two British cohorts, unlike one might expect given an increase in British divorce rates. Secondly, Wolfinger argued that a more liberal divorce regime in terms of attitudes and legislation makes divorcing easier. Couples who divorce under socially and legally strict divorce regimes are more conflicting and since divorce is often only the ultimate option, such a regime is likely to keep quarrelling couples together longer than they would wish. Furthermore, under a strict divorce regime couples may deliberately intensify their conflicts as a way to demonstrate the need for marital dissolution. However, as we discussed above, stress and stigma caused by parental divorce may not be the most important mechanism behind the divorce cycle (cf. Wolfinger 2005). Alternatively, it may be that divorcing under a strict divorce regime sends a stronger signal of low commitment and nonconformity, which may pass on to the children. Regardless of the mechanism, we expect that *the frequency of parental divorce, liberal attitudes towards divorce, and liberal divorce legislation are negatively correlated with the strength of intergenerational transmission of divorce*.

Other macro-level factors may shape the intergenerational transmission of divorce. In their comparison of the former West and East Germanies, Engelhardt and associates (2002) expected that East German family policies—such as income transfers and childcare—would produce a weaker divorce cycle. Firstly, generous social transfers can reduce the financial consequences of divorce and thus mitigate the effects of parental divorce on children. Secondly, high female labour market participation—supported by active policies such as public childcare—may decrease the effects of parental divorce by providing financial support for single mothers, by creating more balanced power relationships between the spouses, and by exposing children less to their possibly conflict-ridden parents (Diekmann and Schmidheiny 2004). We would thus expect that *generous income transfer policies and levels of female labour force participation are negatively correlated with the intergenerational transmission of divorce*.

Data and Methods

In the first step of the analysis, we use data for 18 countries from the Fertility and Family Surveys (FFS), collected by the Population Activities Unit of the United Nations Economic Commission for Europe (see Andersson and Philipov 2002). The countries are Austria, Flanders (Belgium), the Czech Republic, Estonia, Finland, France, East Germany, West Germany, Greece, Hungary, Italy, Latvia, Lithuania, Poland, Spain, Sweden, Switzerland, and the United States. The data were collected in different years in the different countries (see Table 1). Since men are underrepresented in the samples, we use data on the first marriages of women. We focus on marriages instead of all unions (including cohabitations) because marriages continue to send a stronger signal of commitment. Given the increases in long-term consensual unions in many countries, we miss many marriage-like unions. However, since the increase in parental divorce is unlikely to have been a major factor increasing nation-wide levels of cohabitation, the possible bias introduced to our results should be negligible.

Table 1: Year of data collection, number of cases, person-years units and events (dissolution) by country

	Year collected	Marriages	Person-years	Dissolutions
Austria	1995-96	3,080	33,250	452
Czech	1997	1,193	11,655	235
Estonia	1994	871	7,548	227
Finland	1989-90	2,675	28,544	382
Flanders	1991-92	2,314	21,196	231
France	1994	1,734	18,271	580
East Germany	1992	1,797	16,041	354
West Germany	1992	1,368	12,235	242
Greece	1999	1,989	21,737	120
Hungary	1992-93	2,678	25,713	409
Italy	1995-96	3,021	33,387	124
Latvia	1995	1,996	19,458	534
Lithuania	1994-95	2,082	20,010	294
Poland	1991	2,914	31,319	161
Spain	1994-95	2,513	26,911	122
Sweden	1992-93	1,721	14,931	294
Switzerland	1994-95	2,942	28,787	383
USA	1995	6,179	52,956	1,966
Total		43,071	423,949	7,110

Source: Fertility and Families Surveys

We transformed the data into discrete time event history format, with person-years as the basic unit of analysis (Yamaguchi 1991). We limited the maximum number of person-years to 15. After rather considerable data cleaning, we ended up with a total sample of 423,949 person-years from 43,071 marriages, of which 7,110 (16.5 %) ended in divorce during the observation period.

Our main independent variable is parental divorce. The exact question concerning parental divorce varied somewhat across the countries, although in most countries the respondents were asked whether her parents ever divorced and how old she was when that occurred (Festy and Prioux 2002). Festy and Prioux expressed most concern for the parental divorce variable in Finland (where situation at age 14 was reported), France, and Poland (where the present situation is recorded). We decided, however, to keep these countries. The Finnish data probably underestimate parental divorce to some extent, and the French and the Polish data do not report the age at which the divorce happened, and might overestimate parental divorce during childhood. We discuss possible consequences of these decisions in the results section. We excluded respondents whose parents had not divorced but who did not live most of their childhood with two parents. Thus, our comparison group is formed of women who lived with both parents who did not divorce.

Our other variables include the year of birth, marriage duration, duration squared, number of siblings the mother gave birth to, type of locality in which the respondent grew up (less than 10,000 inhabitants, 10,000 to 100,000 inhabitants, more than 100,000 inhabitants), completed education (Low: ISCED 0-2; Middle ISCED 3; High ISCED 4-6) at the time of interview (because of inconsistencies in the education histories), cohabitation before marriage (dummy), and age at marriage. Some of these variables (such as type of locality in childhood and number of siblings) are control variables. Others (age at marriage, completed education, premarital cohabitation) are better seen as intermediary variables. Unfortunately, the data do not allow us to include a fuller set of controls, which would also include such variables as the education or socio-economic status of the parents. Moreover, the type of locality variable is missing for Belgium, Finland, France, and the United States, and the French data do not have information on the number of siblings. We discuss these issues in the results section of the paper. Table A1 in the appendix presents descriptive information on the variables.

We first estimated three models, separately for each country using logit regression, as is standard in event history analysis with discrete time data (Yamaguchi 1991). The first model includes parental divorce, marriage duration, duration squared, and year of birth. The second includes the control variables, when available, and the third includes the intermediary variables. For a more explicit cross-national comparison, we also estimated a full interaction model.

In the second step we analysed the macro-level correlates of the intergenerational transmission of divorce. As discussed, we focused on contexts during childhood and how they shape the experience of parental divorce. We use macro level data collected earlier for analysis of cross-national differences in the educational gradient of divorce (Härkönen and Dronkers 2006a; 2006b). To the extent possible, we tried to collect data that not only vary between countries, but also across time. Therefore, we use time- and country-varying macro variables. We focus on seven contextual factors, which can be roughly divided into two groups.

The first group consists of variables that measure social and legal aspects of the divorce regime. They include divorce legislation, the share of children of each (ten-year) cohort that experienced parental divorce in each country, and two indicators of attitudes towards divorce. In the above discussion, we expected these factors to be negatively correlated with the intergenerational transmission of divorce.

We measure strictness of divorce legislation with a three-level categorical variable. In the first category, we include regimes in which divorce is either not permitted or is permitted on grounds of fault or other major disruption of marital life. In this category, getting a divorce is very difficult. We included divorce prohibition in this group for practical reasons having to do with the low number of divorces (or often, annulments) in these regimes. In the second category, divorce is permitted on the grounds of fault, mutual consent of the spouses, prolonged separation, or other indications of a factual breakdown of marriage. Finally, in the third category there are no or minor legal grounds to deny divorce, and divorce can be granted with very short waiting times. In some cases, classifying a country into one of these categories was not very straightforward. The United States is the hardest case, as there states have independent divorce legislations. We could not differentiate between the states, so we treat it as a single case.[2] We refer the reader to Härkönen and Dronkers (2006a; 2006b) for more information on the divorce law variable and the sources.

We estimated the share of children who experienced parental divorce in each ten-year cohort by country from the FFS files. We also include two indicators for the attitudinal environment toward divorce: per cent atheists or non-religious persons (Barrett et al. 2001) and per cent respondents who consider divorce justifiable. This variable was constructed based on country (and period, if possible) means from the World Values Surveys and European Values Study (for Greece).

The second group of variables measures welfare state practices and female labour market activity. As discussed, these we expect these to correlate negatively with the intergenerational transmission of divorce. We use two measures for welfare state practices: social expenditure per GDP to measure overall welfare state generosity, and family cash benefits per GDP to measure income transfers targeted at families. With data from the ILO (1967, 1988) and OECD (1997) we managed to build

[2] In 1970, California was the first state to enact no-fault divorce legislation and in 1985, South Dakota was the last state to do so. This solution is obviously not an optimal one, but with the common “divorce tourism” across state borders and the liberal attitudes of the judges, the law was often a dead letter (Castles and Flood 1993). Therefore, we use this as a second-best solution.

rather good time series that mapped the developments in these areas. The exception is Poland, for which we found data for only one point in time. Because of unsatisfactory labour market histories in the FFS, we use measure female labour market activity at the macro-level simply as the share of women active in the labour market, with data from the ILO Labour Statistics (http://laborsta.ilo.org).

Except for the share of each cohort experiencing parental divorce, we measure these variables at the time the child was 15 years old. At the time of data collection, we did not manage to collect data for all variables for each country. Therefore, although we examine the Czech Republic and East Germany in the first phase, we do not include these countries in the second phase. Therefore, in the second phase our sample consists of sixteen countries. Often, we also could not go back in time as far as we wanted and so we exclude the eldest birth cohorts.

Results

The divorce cycle in 18 countries

Table 2 shows the effects of parental divorce on offspring divorce in logged odds. We show the results from four models. The first model includes parental divorce, marriage duration, duration squared, and year of birth. This simple model is the most comprehensive in terms of controls found for all countries. In the second model we control for type of locality the respondents grew up in and the number of siblings. France lacked both of these control variables, and we do not estimate this model for France. The Belgian, Finnish, and U.S. data lack information on the type of locality one grew up in, and therefore we only control for the number of siblings in the case of these three countries. In the third model we further include the highest level of attained education, the age at marriage, and experience of pre-marital cohabitation to examine whether intermediating factors explain the observed associations. The full results for Model 3 are included in the appendix (Table A2).

Because of neglected heterogeneity (see Winship and Mare 1984, p. 517; Wooldridge 2002, pp. 470-472), these estimates cannot be compared directly as showing that, say, the variables added to Model 2 control for so and so much of the estimated effect in Model 1. Therefore, while we present the familiar logged odds ratios from each model in Table 2, we checked these results against more appropriate estimates using y-standardised coefficients with STATA's listcoef command (Long and Freese 2001, p. 74). The results were closely in line with the ones presented here.

Finally, the fourth model tests for cross-national differences in the intergenerational transmission of divorce explicitly by estimating a full interaction model between countries and the independent variables in Model 3, but without locality and number of siblings, either or both of which were missing from France, Belgium, Finland, and the United States. This model should give conservative estimates of the country differences (although they are very much the same when Model 1 was used as the baseline). This interaction model equals running separate models for each country, but enables a better comparison between them. The United States forms the reference category for this comparison.

Table 2. The intergenerational transmission of divorce in eighteen countries, discrete time event history models.

	Model 1	Model 2	Model 3	Model 4
Austria	0.726**	0.596**	0.561**	0.321*
Czech	0.524**	0.431**	0.309 +	0.037
Estonia	0.380*	0.265	0.185	-0.045
Finland [1]	0.751**	0.725**	0.676**	0.360*
Flanders [1]	1.035**	1.033**	0.844**	0.502**
France [2]	0.553**	-	0.492**	0.132
E-Germany	0.582**	0.545**	0.496**	0.165
W-Germany	0.784**	0.656**	0.639**	0.394*
Greece	1.118**	0.852**	0.861**	0.744*
Hungary	0.362**	0.284*	0.265*	-0.021
Italy	1.287**	1.171*	1.062**	0.832**
Latvia	0.393**	0.267*	0.233*	-0.014
Lithuania	0.596**	0.465**	0.430**	0.226
Poland	0.237	0.078	0.079	-0.160
Spain	0.875**	0.772*	0.648*	0.403
Sweden	0.674**	0.593**	0.526**	0.263
Switzerland	0.813**	0.717**	0.676**	0.396*
United States [1]	0.465**	0.458**	0.342**	Ref.

Independent variables (not shown):
Model 1: parental divorce, marriage duration, duration squared, year of birth
Model 2: Model 1 + type of location during childhood, number of siblings
Model 3: Model 2 + completed education, age at marriage, premarital cohabitation
Model 4: Full interaction model with country dummies based on Model 3 but without locality and number of siblings
Notes:
[1] No type of location during childhood for Models 2 and 3
[2] No type of location during childhood and number of siblings for Models 2 and 3
+ $p<0.1$, * $p<0.05$, ** $p<0.01$

Turning to the results, we find that in Poland, the estimates for parental divorce are not significant in any model. However, one should keep in mind that the question concerning parental divorce is different in Poland from the other countries, so the result is not fully comparable. In Estonia, the estimate becomes non-significant in the second model and remains so in the third one. In all other countries, the estimates remain significant in each model, although in the Czech Republic the estimate of parental divorce in the third model is significant only at the ten percent level. As mentioned above, when we checked these estimates against y-standardised coefficients, the reduction in their size was very similar.

In most cases, controlling for type of locality in childhood and number of siblings does not decrease the size of the estimates in any major way. The y-standardised coefficients decrease by over 30 percent only in Latvia, and additionally, by over 20 percent in Austria, Estonia, Greece, and Lithuania. One must keep in mind that one or both of these controls are not available for Belgium, Finland, France, and the United States. In any case, we find the divorce cycle in most countries after including our set of control variables.

With the exceptions of Estonia and Poland, the association between parents' and their daughters' divorce remains significant also after we add completed education, age at marriage, and premarital cohabitation into the models. Again, the reductions in coefficient size are very similar when we looked at the y-standardised coefficients. According to them, these three variables explain over 20 percent of the intergenerational transmission of divorce only in Belgium, the Czech Republic, Estonia, and the United States.

We can conclude from the first three models that with the exception of Poland, we find a significant intergenerational transmission effect of divorce in each country, which does not disappear (except for Estonia) when we control for type of locality and the number of siblings. Obviously, this does not yet account for a causal relationship. Neither do differences in educational attainment, age at marriage, and premarital cohabitation fully explain the intergenerational transmission of divorce.

We also tested whether the association has changed across the birth cohorts, but did not find any significant interaction terms (not shown). Thus, we were not able to replicate Wolfinger's (1999) results for the United States and Engelhardt et al.'s (2002) finding for the former two Germanies.
We are also interested in cross-national differences in this intergenerational association. For this, we turn to Model 4, which presents the interactions between parental divorce and the country dummies from the full interaction model. We use the United States as the reference category. We can see that although eyeballing through the results from Models 1-3 suggests plenty of cross-national variation, in many countries the effects do not differ in a statistically significant way from the American ones. The estimates are significantly higher in Austria, Belgium, Finland, West-Germany, Greece, Italy,

and Switzerland. Possibly due to low levels of parental divorce, the seemingly much stronger association between the intergenerational transmission of divorce in Spain than in the United States is not statistically significant.

Multilevel analysis of macro-level correlates

Model 4 in Table 2 showed that in seven countries, the intergenerational transmission of divorce was significantly stronger than in the United States. Country dummies are, however, rather crude proxies of social context since they do not differentiate between macro-level factors. Additionally, the full interaction models fitted in Model 2 are rather inefficient as they use many degrees of freedom.

To examine the role of specific macro-level factors, we continue our analysis with multilevel discrete time event history models with sixteen countries, as discussed in the methods section of the paper. We use the same model as in Model 4 of Table 1, but without the country dummies. The first model (Model A) includes the individual level variables. Parental divorce has a positive effect on divorce-risk, and there is still country variance in this divorce-risk unexplained by the individual characteristics. In the next models, we add each macro-variable separately together with the interaction of the macro-variable and parental divorce. This was necessary as models with all macro-level variables and their interactions with parental divorce failed to converge. The interaction term is the most interesting estimate of the models given our interest in whether the divorce cycle is correlated with macro-level indicators. We decline from interpreting the main effects of the macro-level variables as they are measured when the respondent was around 15 years old. We also estimated the equations based on Model 1 from Table 2 without changes to our substantive results (not shown).

Table 3: Results from multilevel discrete time event history models with macro-level variables measured around age 15 introduced separately.

		Social and legal context of divorce					Welfare state practices and female labour market		
Micro & Macro variables	Model A: Individual model	B: A + divorce laws (Ref: no & strict)		C: A + % atheists or non-religious	D: A + % divorce justifiable	E: A + Level of parental divorce (/100)	F: A + Female labour market participation	G: A + Family cash benefits / 10	H: A + Social expenditure
		Less strict	Unilateral						
Parental divorce	0.529**	0.727 **		0.594 **	0.290	0.913 **	1.172 **	0.476**	0.369**
Year of birth	0.031**	0.035**		0.030**	0.031**	0.029**	0.032**	0.031**	0.028**
Duration	0.192**	0.193**		0.200**	0.191**	0.193**	0.192**	0.192**	0.193**
Duration ^{2}	-0.013**	-0.013**		-0.013**	-0.013 **	-0.013**	-0.013**	-0.013 **	-0.013 **
Education	-0.050*	-0.050*		-0.053*	-0.049*	-0.051*	-0.049*	-0.050 *	-0.050*
Age at marriage	-0.035 **	-.034**		-0.035 **	-0.035 **	-0.035**	-0.035**	-0.035**	-0.035**
Cohabited	0.368 **	0.370 **		0.376 **	0.367 **	0.373 **	0.366**	0.368**	0.369**
Macro-variable		0.186 *	-0.115	0.010 *	-0.004	0.039 **	0.002	-0.001	0.006
*parental divorce		-0.115	-0.223	-0.004	0.005	-0.017**	-0.011**	0.006	0.012
Constant	-5.645**	-6.004**		-5.764 **	-5.473 **	-6.206**	-5.784**	-5.637 **	-5.601**
Country variance	0.313**	0.299 *		0.300**	0.316**	0.173**	0.310**	0.314**	0.323**
Individual variance	0.691**	0.701**		1.039**	0.659**	0.604 **	0.669**	0.692**	0.738**
- Log-likelihood	414410	416386		416947	414322	456153	415663	414546	412215

+ p<0.1, * p<0.05, ** p<0.01

Model B shows that the strictness of divorce laws is not significantly related to the strength of the divorce cycle. The same holds for the percentage of atheist or non-religious persons in the country. Neither is acceptance of divorce (% of the population finding divorce acceptable) correlated with the divorce cycle. Therefore, liberal attitudes and liberal divorce legislation are contrary to our expectations not related with the strength of the intergenerational transmission of divorce.

Turning to the frequency of parental divorce within a cohort in each country, the estimates of Model E support our hypothesis that the divorce cycle is weaker when parental divorce is more common. This result is also in line with previous studies and interpretations by Amato and Keith (1991), Wolfinger (1999; 2005), Engelhardt and associates (2002), and Diekmann and Schmidheiny (2004).

Only one of the interactions between parental divorce and the three welfare state related factors is statistically significant: higher female labour market participation levels during childhood are negatively related to the strength of the divorce cycle, whereas welfare state spending patterns do not show a significant association. As discussed, female labour market participation levels can proxy at least three social patterns, more balanced power relationships between the spouses, the extensiveness of the child care system (which exposes children less to conflicting parents), and single mothers' chances of supporting themselves economically.

Finally, we combine the two macro-level variables that were significantly related to the divorce cycle to the same model, Model I shown in Table 4 below.

Table 4: Results from multilevel discrete-time event history models with all significant macro variables. Standard errors between parentheses.

	Model I
Parental divorce	1.379 **
Year of birth	0.029 **
Duration	0.193 **
Duration^2	-0.013 **
Education	-0.051 *
Age at marriage	-0.035 **
Premarital cohabitated	0.371 **
Level of parental divorce	0.039 **
Level of parental divorce * Parental divorce	-0.015 **
Level of female labour market participation	0.003
Level of female labour market participation* Parental divorce	-0.009 **
Constant	-6.337 **
Country variance	0.169 **
Individual variance	0.599 **
Log-likelihood	-457554

+ p<0.1, * p<0.05, ** p<0.01

The substantive conclusions remain fundamentally the same and both the level of female labour force participation and the frequency of parental divorce correlate negatively with the divorce cycle. The estimate of the interaction term between the frequency of parental divorce and the divorce cycle is stronger than the interaction between female labour force activity and parental divorce, pointing to the importance of the "normalisation" of divorce in reducing its intergenerational transmission. When looking at Model E in Table 3, this term also explains more of the country variance than any other macro-variable. In Model I this variance is only marginally reduced compared to Model E.

Discussion

Our analysis of the first marriages of 43,071 women from 18 countries with data from the Fertility and Family Surveys shows that with the exception of Poland, women whose parents have divorced have a higher divorce risk than those with an intact family background. Furthermore, in most cases this association remains after inclusion of various controlling and intermediate variables. The relationship also varies across countries: when compared to the United States, we found seven countries where the association was significantly stronger. These results are in concordance with the earlier results of Diekmann and Schmidheiny (2004) and Engelhardt and associates (2002), who also found cross-national variations in the intergenerational transmission of divorce.

A contribution of this paper is that we try to explain cross-national variation in the divorce cycle with macro-level variables that measure the social and legal aspects of divorce and welfare state and labour market patterns in the 16 countries, for which we had data on these factors. We expected that high parental divorce rates, more liberal attitudes towards divorce, and more liberal divorce laws contribute to normalising divorce, to making parental divorce less stigmatizing and marriages that end in divorce less conflict-ridden, and to weakening the signal of low marital commitment of parental divorce. As a result, we expected that these measures are negatively correlated with the strength of the divorce cycle. We also expected that welfare state spending—both overall and spending targeted at families—and high levels of female labour force participation reduce the financial consequences of parental divorce. We also made the argument that high levels of female labour force participation indicate more power balance between the husband and the wife. Furthermore, they indicate that children spend less time with their parents and more time in childcare, for instance, and are thus less exposed to their possibly conflict-ridden parents. Both these mechanisms would produce a negative correlation between welfare state spending, female labour force participation, and the divorce cycle.

Out of the interactions between these macro-level factors and parental divorce, only the interactions with the frequency of parental divorce and the levels of female labour force participation were statistically significant and negative. These interactions remained negative and statistically significant when included in the same model. This suggests that instead of the legal and attitudinal environment surrounding children with divorcing parents, the actual behaviour of parents is the most important factor. Regarding female employment, the result that welfare state spending was not correlated with the divorce cycle suggests that the non-financial aspects of female labour market participation play a stronger role.

Of these two interactions, the association between the frequency of parental divorce and the divorce cycle proved stronger. This association has been suggested as an explanation to the decreases in the divorce cycle found in the United States (Wolfinger 1999) and the weakening trends in and overall differences between the former East and West Germanies (Engelhardt et al. 2002). This has been often explained as reflecting either decreasing stigma associated with parental divorce or the changing composition of divorcing parents in the sense that under strict divorce regimes (legally and well as socially), dissolving marriages are more likely to have a longer and more serious history of conflict. These explanations have been questioned by results that show no change in the effects of divorce on other relevant outcomes (Sigle-Rushton et al. 2005). One would also have expected to find a significant interaction between the divorce cycle and other macro-level indicators such as divorce laws.

Optionally, the negative association between the frequency of parental divorce and the divorce cycle can reflect weaker signalling of low marital commitment when divorce is common. Recent research has linked the divorce cycle to low marital commitment (Amato and DeBoer 2001; Wolfinger 2005) and more positive attitudes toward divorce (Amato 1988; Trent and South 1992; Axinn and Thornton 1996) among children of divorce, presumably due to role modelling the decisions parents take when confronted with marital difficulties. When parental divorce is rare, it may signal lower levels of marital commitment relative to the overall population than when children of divorce share similar experiences with more peers. The relative strength of parental divorce in shaping commitment to marriage is likely to weaken further as marital commitment decreases and divorcing increases even among those whose parents did not divorce.

We invite researchers to replicate our results with new data, such as the Gender and Generations Survey, which also collects data on macro-level variables. Our results are in many ways tentative, not least because of the gaps in our macro-level time series. Nevertheless, the importance of our analysis is that it suggests that the divorce cycle is shaped by measurable contextual factors. We also invite researchers to further test our interpretation of the negative correlation between the frequency of parental divorce and the divorce cycle.

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Appendix

Table A1. Descriptives of the variables

	Mean	Standard deviation	Min	Max
Divorce	0.02	0.13	0	1
Parental divorce	0.11	0.31	0	1
Marriage duration	5.63	4.13	0	14
Year of birth (19--)	55.75	6.63	38	81
Locality in childhood: <10,000	0.52	0.50	0	1
Locality in childhood: 10,000-100,000	0.26	0.44	0	1
Locality in childhood: >100,000	0.22	0.42	0	1
Number of siblings	3.71	2.24	0	20
Age at marriage	21.68	3.47	15	57
Completed education: low (ISCED 0-2)	0.35	0.48	0	1
Completed education: middle (ISCED 0-2)	0.43	0.50	0	1
Completed education: high (ISCED 0-2)	0.22	0.42	0	1
Cohabited before marriage	0.26	0.44	0	1
Divorce laws: Prohibited or strict	0.34	0.47	0	1
Divorce laws: Breakdown, other less strict	0.65	0.48	0	1
Divorce laws: Unilateral no-fault	0.01	0.11	0	1
Share of atheists (%)	10.73	10.08	1.1	53.3
Agree divorce justified (%)	49.19	6.49	39	72
Share of parental divorce (%)	13.65	8.59	0	29.60
Female labour force participation (%)	52.85	13.91	27.5	78.9
Family cash benefits (% of GDP)	1.15	0.91	0	4.9
Social expenditure (% of GDP)	14.54	4.07	9.5	31.9

Table A2. Full results of Model 3 in Table 2

	AT	CZ	EST	FIN	FLA	FRA	GDR	FRG	GRE
Parental divorce	0.561**	-0.309+	0.185	0.676**	0.844**	0.492**	0.496**	0.639**	0.861**
Duration	0.243**	0.119*	0.180**	0.325**	0.347**	0.222**	0.368**	0.167**	-0.004
Duration^2	-0.015**	-0.006	-0.015**	-0.023**	-0.021**	-0.009*	-0.027**	-0.007	-0.003
Year of birth	0.047**	0.051**	-0.025*	0.030**	0.042*	0.078**	0.041**	0.094**	-0.001
Mid educ.	-0.155	-0.315*	-0.0338	-0.024	-0.033	0.149	-0.384*	0.050	0.264
High educ.	-0.367*	-0.739*	-0.427	-0.010	-0.091	0.117	-0.387*	0.129	0.405
Age at marriage	-0.050**	-0.057+	-0.091**	-0.067**	-0.139**	0.028+	0.040	0.008	-0.033
Pre-marital cohabit.	0.275*	0.427**	-	0.304*	1.163**	0.432**	0.287*	0.132	0.164
Number of siblings	-0.039	-0.043	-0.111*	-0.038+	-0.024	-	-0.004	0.042	0.069
10,000-100,000	0.143	0.601**	0.226	-	-	-	0.248*	0.074	0.562+
>100,000	0.645**	0.863**	0.582**	-	-	-	0.304+	0.459**	1.202**
Constant	-6.586**	-6.575**	-0.093	-5.146**	-5.150**	-9,535**	-7.920**	-10.572**	-5.405**
Log-likelihood	-2210,79	-1,083.05	-807.98	-1,945.00	-1,209.57	-2,409.73	-1,567.17	-1,097.79	-682.63
N person-years	32,495	11,600	6,374	28,368	21,117	17,925	15,273	11,640	20,757
	HUN	**IT**	**LAT**	**LIT**	**POL**	**SP**	**SWE**	**SWI**	**USA**
Parental divorce	0.265*	1.062**	0.233*	0.430**	0.079	0.648*	0.527**	0.676**	0.342**
Duration	0.102*	0.077	0.172**	0.327**	0.064	0.345**	0.205**	0.295**	0.107**
Duration^2	-0.007*	-0.001	-0.012**	-0.022**	-0.003	-0.021**	-0.012**	-0.019**	-0.011**
Year of birth	0.022*	0.067**	0.026**	0.037**	0.046**	0.066**	0.034**	0.001	0.012***
Mid educ.	-0.135	0.413+	-0.156	-0.674**	0.168	0.618*	-0.032	-0.391*	-0.168**
High educ.	-0.285+	0.805*	-0.206	-0.716**	-0.013	0.249	-0.002	-0.173	-0.181**
Age at marriage	0.016*	-0.033	-0.048*	0.012**	0.036	-0.044	-0.027	-0.075**	-0.084**
Pre-marital cohabit.	0.214	1.084**	0.105	0.395**	0.785*	1.258**	0.339*	0.741**	0.425***
Number of siblings	0.030	-0.041	-0.074*	-0.076+	-0.050	-0.006	0.035	-0.036	-0.029**
10,000-100,000	0.425**	0.483*	0.080	0.286+	0.386+	0.078	0.351*	0.307*	-
>100,000	0.712**	0.687**	0.326**	0.501**	0.664**	0.775**	0.428**	0.413*	-
Constant	-6.329**	-9.714**	-4.312**	-6.946**	-8.871**	-9.833**	-6.349**	-3.715**	-2.281
Log-likelihood	-2,025.44	-762.93	-1,947.45	-1,369.53	-973.67	-716.48	-1,415.69	-1,548.80	-8,175.11
N person-years	25,348	32,439	15,415	18,723	31,178	26,709	14,881	23,874	52,860

+ p<0.1, * p<0.05, ** p<0.01

Reference categories: Level of education: low education: Size of childhood locality: <10,000 inhabitants.