The Instability of Divorce Risk Factors in the UK

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Abstract

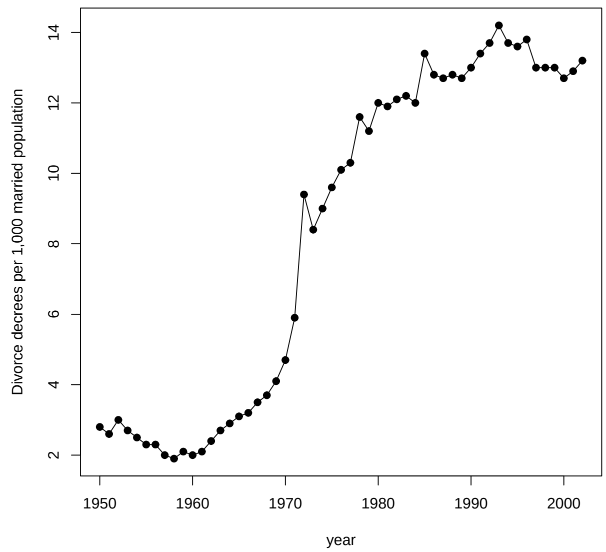
We examine the stability of divorce determinants in the UK between 1960 and 1989. Using retrospective marriage history data, we show that, over successive marriage cohorts, the associations between marital instability and many covariates have undergone substantial, systematic and quite interpretable changes. Thus, the parameter for marrying a divorcee has weakened over time, while that for premarital cohabitation follows a non-monotonic pattern: an initial decline is followed by a rise. We also report a reversal of the educational gradient in divorce. Better educated women used to have higher divorce risks, but now their marriage is more stable than that of women without qualifications. Finally, we confirm a recent finding that in the UK the familiar marriage-stabilising effect of children has been reversed. We interpret these changes in terms of self-selection and other social changes.

1 Introduction

It is well known that the divorce rate in the UK has increased very sharply since the 1960s. Figure 1 shows the trend for England and Wales, where the annual average between 1950 and 1964 was 2.4 divorces per 1,000 marriages. This increased sixfold to over fourteen divorces per thousand in 1993 (Office of Population Censuses and Surveys, 1985, 1990, 1993).[1] Crude divorce

[1] The crude divorce rate of recent years can be found on the web site of the UK Office for National Statistics: `www.statistics.gov.uk`.

rate in England and Wales has since declined slightly to about thirteen per thousand in the late 1990s, but it is still among the highest in the European Union (Office for National Statistics, 2004, Table 2.13).



Source: Office of Population Censuses and Surveys (1985, 1990, 1993)

Figure 1: Crude divorce rate in England and Wales.

What is also well established are some of the risk factors that are associated with divorce. For example, Murphy (1985a,b) analyses data from the 1976 Family Formation Survey and the 1980 General Household Survey and reports that young brides, those who were pregnant at the time of the wedding, those who got married in civil ceremonies, childless couples as well as couples with four or more children face significantly higher divorce risks. He also reports an association between divorce and social class, though its magnitude is rather modest.[2]

Murphy's findings are largely consistent with other studies done in the UK (e.g. Berrington and Diamond, 1999) and elsewhere (e.g. White, 1990; Bracher *et al.*, 1993; Jalovaara, 2001). What is not clear, however, is whether the associations between marital instability and various risk factors have changed over time. In this paper we address this question using retrospective

[2]'The maximum gross difference between the relative risks in Social Classes II to IV [of the UK Registrar General classification] is similar to that of a change of one year in age at marriage' (Murphy, 1985a, p.459).

marriage history data extracted from a series of nationally representative surveys.

2 Possible sources of change

Why might the 'effects' of the various divorce risk factors be changing? One possibility is that the direct and indirect costs of divorce have changed. When divorce was rare and the associated stigma strong, it took more resources to cope with divorce (Hoem, 1997, p.19). But as divorce becomes more common and accepted, more people would be able to do so. This would imply a flattening of the divorce gradient of those variables which index resources, such as income or educational attainment (Goode, 1993, p.vii). Other exogenenous social changes, for example those that are related to women's changing roles in society, might also be relevant. We shall discuss these further after reviewing the evidence. But given the dramatic changes in the relative frequency of many divorce risk factors themselves (see Table 1 below), we shall highlight the endogeneous process of self-selection.

An example of selection-induced parameter change concerns premarital cohabitation. It is a well established finding that marriages which began as cohabiting unions are less stable than those which did not (DeMaris and Rao, 1992; Axinn and Thornton, 1992; Lillard *et al.*, 1995). As many researchers have observed, this finding could plausibly be interpreted as arising from a process of self-selection. That is, individuals with non-traditional attitudes about marriage and the family are more likely to cohabit before marriage; but they are also more likely to divorce. Thus, the association between premarital cohabitation and marital instability is, under this view, a spurious one. To the extent that this is true, because the rate of premarital cohabitation in the UK has risen from 3% in the 1960s to about 70% in the 1990s (Berrington and Diamond, 2000),[3] the cohabitants should have become a much less self-selected group, and thus the association between premarital cohabitation and marital instability should attenuate over time. However, as the level of premarital cohabitation becomes *very* high, an opposite self-selection process might begin to operate: only those who are very conservative about marriage and the family would marry directly. If this group is especially unlikely to divorce, we would expect the association between cohabitation and divorce to become stronger again. In short, if a self-selection process is driving the association between marital instability and premarital cohabitation, we should

[3]See also Table 1 below. For trends of cohabitation in the US, see Smock (2000) and Bumpass and Lu (2000).

observe a non-monotonic trend in the relevant parameter across marriage cohorts.

A similar argument could be made in relation to marrying a divorcee.[4] When this was very rare, the individuals involved are likely to be a selected, unconventional group. However, as divorce and remarriage become more common, this group should become less self-selective. Accordingly, the association between remarriage and marital instability would attenuate over time. However, as with premarital cohabitation, if marrying a divoree is *causing* greater marital instability, then the rising level of remarriage should *not*, in general, be accompanied by a weakening of the relevant parameter.

3 Data and methods

Our data come from the General Household Survey (GHS), which is a repeating general purpose survey conducted by the UK Office for National Statistics. The GHS has been running on an annual basis, almost without interruption, since 1971. Each year, about 9,000 households are sampled, and face-to-face interviews are conducted with all individuals aged 16 or above in these households. The GHS is a cross-sectional survey, but since 1979 respondents were asked quite detailed retrospective questions on marriage and the family, allowing us to reconstruct their complete marriage and fertility histories. In this paper, we pool together relevant GHS data from 1989 to 2000, and focus on women's first marriage.[5]

Because of the retrospective nature of the GHS data, almost all covariates that are available to us, e.g. year of marriage, age at marriage, premarital cohabitation and educational attainment, are time-constant in nature. There is just one set of time-varying covariates, which measures parity.[6] This means that we are not able to examine the dynamic effect of those factors which might themselves be changing over the course of a marriage, such as women's employment or family income.

[4]Although the focus of this paper is women in their first marriage, their spouses could be divorcees.

[5]Before 1989 information on cohabitation before first marriage was not available for respondents who have been married twice or more. Since the multivariate analyses reported below include the covariate of premarital cohabitation, our analysis is based on a smaller data set of GHS 1989–2000.

[6]In this paper, we consider children born to the respondents only. There is some information on step, adopted and fostered children who are *currently* living with the respondents. But since step children from dissolved first marriages are probably not living with the respondents any more, such information is, for our purpose, incomplete, and therefore not used in this paper.

Table 1: Descriptive statistics of covariates by marriage cohort

	marriage cohort					
	1950s	1960s	1970s	1980s	1990s	overall
			mean			
age at marriage	20.6	21.4	22.1	23.5	26.0	22.5
number of children	2.6	2.2	1.9	1.5	0.9	1.8
			percentage			
premarital cohabitation	1.1	3.3	14.5	41.2	67.7	21.9
husband being a divorcee	2.0	4.0	8.7	12.7	15.9	8.2
premarital birth	2.3	4.7	6.6	10.5	20.1	7.9
premarital conception	16.2	17.8	13.4	11.2	7.5	13.7
degree	8.0	13.9	19.9	21.7	27.2	18.3
A-levels	1.8	5.1	9.0	12.6	17.4	8.9
O-levels	11.1	18.1	25.5	33.0	32.0	24.6
no qualifications	79.1	62.9	45.6	32.7	23.4	48.2

Also, we have only one piece of information (supplied by the respondents) about the husbands: whether they were previously married. Because the GHS is a household survey, there is additional information about the husband of those women who are currently married, such as their age or educational attainment. But such information is not available for women whose first marriage had already dissolved. This precludes us from including these covariates in the analysis. These limitations notwithstanding, the GHS is still an invaluable data source. It gives us information on the first marriage of 31,381 women who got married between 1960 and 1989. The size of this data set and its temporal coverage allow us to examine the stability and change of the effect of divorce risk factors over a period in which family behaviour has been changing most rapidly.

Table 1 reports some basic descriptive statistics of our covariates, which all show significant change over time. For example, the mean age at marriage of all women in our sample is 22.5. But for the 1950s marriage cohort, the figure was 20.6, rising to 26.0 for those marrying in the 1990s. This is consistent with the well-known trend of marriage postponement. A trend towards smaller family size is also evident, although it should be noted that because many women of the more recent marriage cohorts have not yet completed childbearing, the actual decline in fertility is less steep than that shown in Table 1.[7]

[7]Indeed, from a period perspective, much of the decline in total fertility rate in the UK

As for premarital cohabitation, 22% of all women in our sample lived with their future husband as a couple before marriage.[8] But this overall average masks a great deal of variation across marriage cohorts: from 1% of those who married in the 1950s to over two thirds of the 1990s cohort. Similarly, the proportion of women marrying a divorcee rose from 2% to about 16%.

The decline of 'shotgun marriage' is also evident in Table 1. We see an upward trend in the incidence of premarital birth, from 2% among those who married in the 1950s to about one fifth of the 1990s cohort. At the same time, the percentage of women with premarital conception declines from 16% to 8%.[9]

Finally, Table 1 also reports the trends in educational attainment.[10] There is an almost linear decline in the proportion of brides with no qualifications (from almost 80% to 23%), while the shares of brides with A-levels or university degree increase significantly, reaching 17% and 27% respectively in the 1990s cohort. These trends are, of course, consistent with what we know about the general expansion of education and the closing of the gender gap in educational attainment.

4 Results

In the analysis that follows, we focus on first marriages formed between 1960 and 1989. We censor all marriage spells at ten years if they were still intact by then.[11] Figure 2 shows the proportion of marriages that are still intact by duration across marriage cohorts. Reflecting the accelerating divorce rate over this period, the survival curves for recent marriage cohorts are invariably

took place between the early 1960s and the mid-1970s, and variation in TFR since the mid-1970s has been largely trendless within a relatively narrow range (Office for National Statistics, 2007, p.1).

[8]Because cohabitation with someone other than future husband is not counted in the GHS, this is an underestimation of the extent of premarital cohabitation.

[9]Premarital conception is defined as giving birth to the first child within eight months of marriage.

[10]The minimum school-leaving age in England and Wales has been set at 16 since 1972. This is the end of secondary education for most young people, and O-levels are formal qualifications typically gained at this point. A-levels are matriculation qualifications for university entrance, typically taken at the age of 18.

[11]Censoring the spells at ten years implies that there is relatively little difference between the period perspective and the cohort perspective that we adopt in this paper. We have repeated our analysis without censoring at ten years, and the results are largely the same. Details are available from the authors on request. It should be noted that in our analyses, marriages ended in death of spouse are coded as censored, and marital dissolution is measured at time of separation.

steeper than those for previous cohorts. For example, just over 10% of those who married between 1960 and 1964 had divorced within ten years. But for the 1985-9 marriage cohort, this figure rose to 26%.

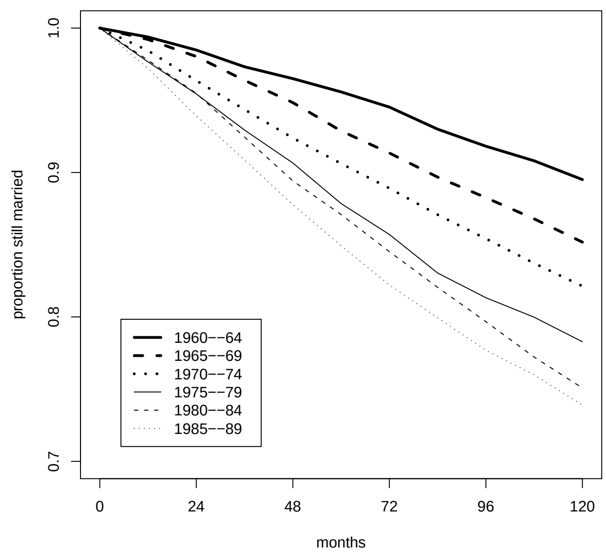


Figure 2: Survival function by marriage cohort.

We then fit the Cox proportional hazard model to the data. Our baseline model contains the following covariates: a linear term and a quadratic term for age at marriage (centred at age 21), year of marriage (centred at 1960), and dummies for having one child, two children, three or more children (hereafter as the children dummies), whether the youngest child is less than four years old (hereafter as the preschooler parameter), whether the husband was a divorcee, premarital cohabitation, and educational attainment. This model is fitted to all first marriages in our data, and the parameter estimates are reported in the first column of Table 2.

It can be seen that, considered over the entire sample, women who marry late face lower divorce risks. For example, compared to those who married at age 21, the divorce hazard of those marrying at age 22 is, on average, 15% lower ($1 - e^{-.170+.006}$). The positive coefficient of year of marriage reflects the rising divorce rate over time. The divorce harzard of those women who marry a divorcee is 29% ($e^{.257} - 1$) higher, and premarital cohabitation raises divorce rate by 71% ($e^{.539} - 1$). The relationship between divorce risks and educational attainment is non-monotonic. Compared with the reference category of women with no qualifications, university graduates have lower divorce risks, but women with A-levels have higher divorce risks (though the A-levels parameter is marginally insignificant with $p = .08$). Having a preschooler is associated with substantially lower divorce risks. Controlling for all of the above, the dummies for having one child, two children, three or more children are all positive, though only the last of these is statistically significant at the 5% level. These estimates are not all that exceptional. But are they stable over marriage cohorts?

To answer this question, we first report in Table 3 the fit statistics of our baseline model (model A) and those with an additional term postulating a *linear* interaction effect between year of marriage and one covariate (or one set of covariates, see models B to I). On the basis of the likelihood ratio test, it is clear that models B, C, D, E, H and I fit the data better than does the baseline model, while for model G the improvement in fit is just outside the conventional 5% level. When we allow all covariates to interact with year of marriage, the fit of model J is again significantly better than that of the baseline model.[12]

Thus, there is rather clear evidence of significant shifts in the determinants of divorce in the UK.[13] To represent these changes, and allowing for non-linearities, we divide our sample into five-year marriage cohorts and fit the Cox model to each cohort separately. The results are reported in columns 2 to 7 of Table 2.

It can be seen that the magnitude of the two age-at-marriage terms generally become smaller across cohorts. But the linear term remains negative and the quadratic term positive throughout, and they have retained their statistical significance. As expected, year of marriage becomes insignificant within the five-year bands, except for 1960-64. But we see much greater changes in other parameters. For example, the parameter for marrying a divorcee generally weakens across marriage cohorts,[14] such that by the 1985-9 cohort, the last cohort for which we have a full ten-year observation window, this parameter became insignificant. This shift could be explained in terms of the self-selection process discussed above.

Likewise, the parameter for premarital cohabitation also weakens across

[12]We would arrive at the same conclusion if BIC is used as the criterion of model selection, except that models F and G would also be regarded as fitting the data better than model A.

[13]Our results are therefore quite different from those of Teachman (2002), who finds a pattern of overall stability in divorce risk factors in the US. It is beyond the scope of this paper to investigate why the two countries should be so different.

[14]In relation to this and some other parameters, the 1970-4 cohort is an outlier. This might be a result of the fact that the 1970-4 cohort is the one most immediately affected by the 1969 Divorce Reform Act, which came into effect in 1971 (see the spike in Figure 1).

Table 2: Cox proportional hazard model fitted to episode data of first marriage (baseline model).

		marriage cohort					
	1960-89	1960-4	1965-9	1970-4	1975-9	1980-4	1985-9
age at marriage	−.170**	−.198**	−.216**	−.162**	−.166**	−.155**	−.166**
	(.006)	(.018)	(.015)	(.013)	(.013)	(.013)	(.014)
age at marr. sq.	.006**	.022**	.012**	.007**	.006**	.005**	.005**
	(.000)	(.004)	(.002)	(.001)	(.001)	(.001)	(.001)
year of marriage	.043**	.075*	.034	.041	−.003	.030	.038
	(.002)	(.032)	(.024)	(.022)	(.021)	(.021)	(.026)
husband a divorcee	.257**	.593**	.307	.095	.335**	.189*	.178
	(.048)	(.197)	(.171)	(.126)	(.102)	(.094)	(.107)
pre. cohabitation	.539**	1.014**	.786**	.660**	.423**	.467**	.524**
	(.035)	(.205)	(.138)	(.088)	(.074)	(.064)	(.074)
O-levels	−.043	−.177	.013	.013	−.088	−.121	.023
	(.033)	(.140)	(.090)	(.078)	(.071)	(.070)	(.083)
A-levels	.091	.724**	.070	.019	.095	.039	.037
	(.051)	(.194)	(.169)	(.125)	(.109)	(.104)	(.117)
university	−.097*	.011	.281**	.101	−.228*	−.336**	−.272*
	(.044)	(.151)	(.104)	(.091)	(.098)	(.103)	(.118)
one child	.095	−.173	−.395**	.100	.065	.174	.499**
	(.052)	(.172)	(.130)	(.113)	(.116)	(.116)	(.140)
two children	.008	−.496**	−.435**	−.250	−.136	.246	.744**
	(.058)	(.189)	(.140)	(.130)	(.133)	(.126)	(.149)
three + children	.231**	−.586*	−.289	−.057	.264	.509**	1.018**
	(.073)	(.228)	(.176)	(.172)	(.167)	(.157)	(.189)
youngest child < 4	−.369**	−.016	−.200	−.287**	−.462**	−.414**	−.585**
	(.045)	(.145)	(.109)	(.099)	(.104)	(.099)	(.124)
# respondents	31381	4669	5796	5903	5182	4888	4943
# divorce	5388	478	845	1030	1098	1121	816
Log-likelihood	-53915.56	-3929.15	-7128.66	-8718.99	-9115.61	-9160.67	-6424.15

Notes: * $p < .05$, ** $p < .01$, standard errors in parentheses.

Table 3: Model fit for the baseline additive model and models involving an interaction between each predictor variable and year married.

model	term interacting with year of marriage	log-likelihood	χ^2	*df*	*p*	BIC
A	baseline	-53915.56	1917.79	12	—	-2020.89
B	age at marriage	-53913.34	1922.23	13	.035	-2033.93
C	age at marriage squared	-53908.18	1932.56	13	.000	-2044.26
D	age at marriage & age at marriage squared	-53908.17	1932.58	14	.000	-2052.87
E	children dummies	-53906.71	1935.49	15	.000	-2064.37
F	youngest child < 4	-53914.43	1920.05	13	.133	-2031.75
G	husband was a divorcee	-53913.91	1921.09	13	.069	-2032.79
H	premarital cohabitation	-53911.68	1925.56	13	.005	-2037.26
I	education	-53897.21	1954.48	15	.000	-2083.36
J	all	-53878.04	1992.84	23	.000	-2190.45

cohorts. For the 1960-4 cohort premarital cohabitation was associated with a 176% ($e^{1.014} - 1$) increase in divorce risks, but by 1985-9, when over half of the couples cohabited before marriage, this dropped to a still large but much reduced 69% ($e^{.524} - 1$). Again, this decline could be explained in terms of selection effects. Further, while the cohabitation parameter weakens over the 1960s and 1970s, it seems to be strengthening in the 1980s. This non-monotonic pattern is consistent with the view that an opposite selection process begins to operate when cohabitation becomes very common. This lends further support to the self-selection interpretation.

The changing association between educational attainment and marital instability is intriguing. Recall that taken over the whole sample, the pattern was rather complex: university graduates are *less* likely to divorce than women with no qualifications, but women with A-levels are *more* likely to do so (see first column of Table 2). When split into marriage cohorts, the following pattern emerges: compared to women with no qualifications, better educated women used to have higher divorce risks, but the education gradient has now reversed. For example, in the 1960-4 cohort, the divorce rate of women with A-levels was twice ($e^{.724}$) as high as that for women with no qualifications. But for subsequent cohorts, women with A-levels were *not* more likely to divorce. As for university graduates, the divorce rate of graduates who got married in 1965-9 were 32% ($e^{.281} - 1$) higher that than for women with no qualifications. This parameter became insignificant for the 1970-4 cohort, and then turned negative and significant for subsequent cohorts, such that the divorce hazard of graduates who married between 1985 and 1989 were 24% ($1 - e^{-.272}$) lower than that for women with no qualifications.[15]

Coming to the parameters for children, Table 2 shows that for marriages formed in the 1960s, the children dummies were all negative in sign, and four of the six parameters were statistically significant at the 5% level. Further, the preschooler parameter was not significant. Thus, for the first two marriage cohorts, children generally reduced the divorce hazard, and the marriage-stabilising effect of children applied regardless of their age.

For marriages formed in the 1970s, the children dummies were insignificant, while the preschooler parameter turned significant. Thus, children were still associated with marital stability, but only when they were under the age of four. When we get to the 1985-9 cohort, the children dummies were positive and significant, and their effects were large. Although the preschooler parameter was also significant and its magnitude had in fact increased, the net association for having two children, or three or more children, even when they were under four, was marriage-destabilising. For example, the divorce hazard of those couples with two children (at least one child being a preschooler) was 17% ($e^{.744-.585} - 1$) higher than that for otherwise similar but childless couples.

This result is consistent with the finding of Böheim and Ermisch (2001) and Chan and Halpin (2002), who use panel data from the 1990s to show that children are associated with greater marital instability in the UK. Using retrospective GHS data, we now see that this shift was well underway in the 1980s.

What might account for such a change? Previous research suggests that the association between children and marital stability is multidimensional in nature. Not only are the *number* of children and their *age* important, the *timing* of childbirth also matters (e.g. Murphy, 1985a; Waite and Lillard, 1991). To disentangle the effect of timing from those of number and age, we add premarital conception and premarital birth to the model, and report the results in Table 4.

The parameters of Table 4 are mostly similar to their counterparts in Table 2.[16] But with the timing of childbirth controlled for, the children dummies generally become more negative (i.e. more marriage-stabilising). Nonetheless, the overall trend is still for children to become less of a marriage-stabilising factor across cohorts. This can be seen from Figure 3: the lines for our model which controls for premarital conception and premarital birth

[15] Hoem (1997) reports a similar reversal of the educational gradient in divorce risks for Sweden.

[16] We have also repeated our analyses with a subsample of respondents who did not have premarital birth. This gives largely the same results as those reported here. Details are available from the authors on request.

Table 4: Cox proportional hazard model fitted to episode data of first marriage (extended model: baseline model plus control for premarital birth and premarital conception).

	1960-89	1960-4	1965-9	1970-4	1975-9	1980-4	1985-9
age at marriage	−.167**	−.189**	−.211**	−.165**	−.161**	−.147**	−.165**
	(.006)	(.019)	(.016)	(.013)	(.013)	(.013)	(.015)
age at marr. sq.	.006**	.020**	.011**	.007**	.006**	.005**	.005**
	(.000)	(.004)	(.002)	(.001)	(.001)	(.001)	(.001)
year of marriage	.042**	.073*	.030	.039	−.001	.034	.029
	(.002)	(.032)	(.024)	(.022)	(.021)	(.021)	(.026)
husband a divorcee	.198**	.540**	.252	.038	.292**	.160	.142
	(.049)	(.198)	(.173)	(.126)	(.102)	(.095)	(.107)
pre. cohabitation	.440**	.755**	.608**	.558**	.360**	.407**	.409**
	(.036)	(.214)	(.142)	(.090)	(.075)	(.066)	(.076)
O-levels	−.017	−.151	.005	.025	−.070	−.091	.061
	(.033)	(.140)	(.090)	(.078)	(.072)	(.070)	(.083)
A-levels	.129*	.784**	.098	.055	.114	.073	.099
	(.051)	(.195)	(.169)	(.126)	(.109)	(.104)	(.118)
university	−.047	.040	.322**	.140	−.187	−.298**	−.195
	(.044)	(.151)	(.104)	(.092)	(.099)	(.103)	(.118)
one child	−.309**	−.439*	−.670**	−.154	−.288*	−.198	−.243
	(.058)	(.181)	(.139)	(.127)	(.130)	(.129)	(.165)
two children	−.527**	−.862**	−.817**	−.584**	−.587**	−.237	−.273
	(.066)	(.204)	(.155)	(.149)	(.153)	(.147)	(.193)
three + children	−.507**	−1.134**	−.889**	−.547**	−.377	−.117	−.287
	(.085)	(.255)	(.201)	(.199)	(.198)	(.187)	(.246)
youngest child < 4	−.114*	.137	−.012	−.115	−.236*	−.193	−.132
	(.047)	(.148)	(.112)	(.104)	(.110	(.105)	(.133)
pre. birth	.908**	1.035**	1.060**	.764**	.750**	.604**	1.090**
	(.053)	(.197)	(.145)	(.124)	(.131)	(.117)	(.128)
pre. conception	.358**	.392**	.274**	.104	.398**	.535**	.425**
	(.042)	(.128)	(.100)	(.099)	(.097)	(.091)	(.120)
# respondents	31381	4669	5796	5903	5182	4888	4943
# divorce	5388	478	845	1030	1098	1122	816
Log-likelihood	-53776.82	-3915.65	-7105.20	-8701.27	-9097.59	-9138.75	-6388.95

Notes: * $p < .05$, ** $p < .01$, standard errors in parentheses.

(model 3) are shifted downwards from the lines for the baseline model (model 1). At the same time, all lines of Figure 3 generally slope upward. Overall, for marriages formed in the 1980s, children no longer raise or reduce the divorce hazard.[17]

Given the importance of premarital birth, the question arises as to who are more likely to have premarital birth? To answer this question, Figure 4 shows the association between educational attainment and premarital birth. It is striking how that association has strengthened over time. For example, in the 1960s, the odds of a woman with no qualifications having a premarital birth was about 2 to 3 times that of a female graduate. By the late 1980s, the odds ratio has risen to almost 6, which gets even higher in the 1990s.[18]

5 Summary and discussion

In this paper, we use retrospective marriage history data to show that in the UK the 'effects' of many divorce risk factors have changed over successive marriage cohorts. These changes are substantial, systematic and quite interpretable.

As the rate of premarital cohabitation rose from under 5% in the 1960s cohort to over 40% in the 1980s cohort, the coefficient of our cohabitation parameter almost halves in magnitude (.409/.755, see Table 4). The estimate for marrying a divorcee also weakens over time, such that it became insignificant for the 1985-9 marriage cohort. We have argued that the change of these two parameters could be interpreted in terms of self-selection, i.e. as cohabitation and remarriage become more common, they become less selective and tell us less about the individuals involved.

The plausibility of this interpretation is bolstered further by the observation that as premarital cohabitation change from a rare choice to a *very* popular one, there seems to a selection process operating from the opposite direction, leading to a non-monotonic trend in the relevant parameter.

Among women who got married in the 1960s, it was the better educated who faced higher divorce risks. Over time, the education gradient has not only flattened, it has in fact been reversed, such that for those who got married in the 1980s, university graduates were less likely to divorce than

[17]Figure 3 also shows that it is premarital birth rather than premarital conception which affects the association between children and marital stability. This can be seen from the fact that the lines for model 2 which controls for premarital conception only are very close with those for model 1.

[18]We have carried out some data smoothing for Figure 4 by averaging the contingency table of any year with those of the immediately preceding and subsequent years.

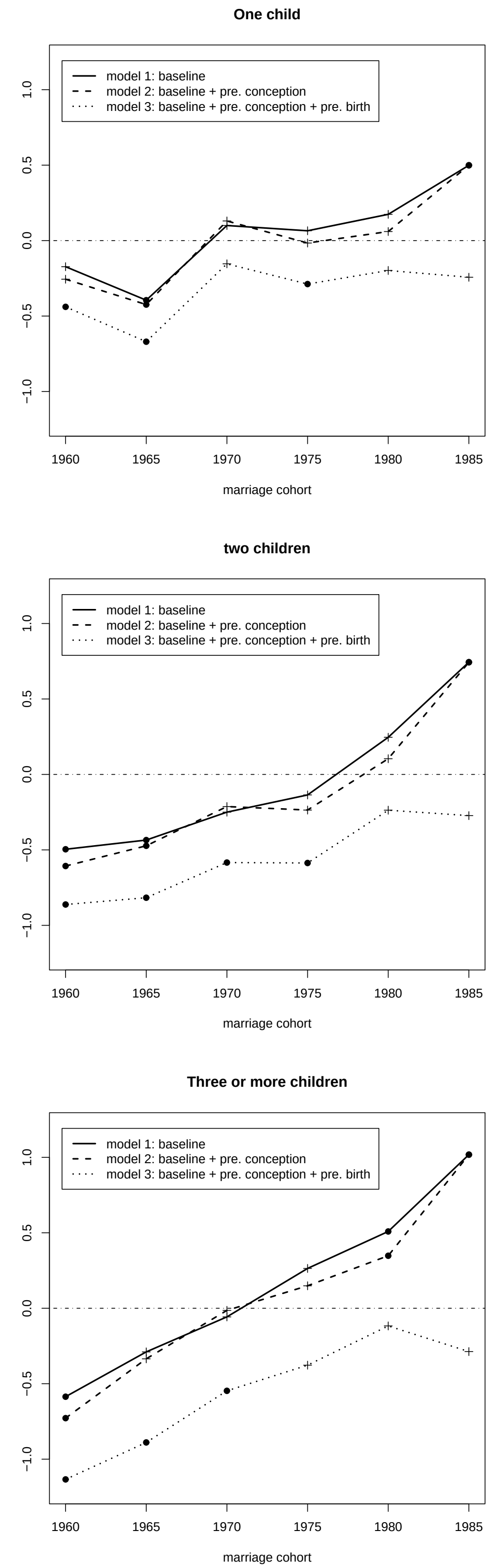


Figure 3: Parameter estimates of having one child (top panel), two children (middle panel) and three or more children (bottom panel) by marriage cohort under three models.

Note: Model 1 is reported in Table 2, Model 3 in Table 4, while Model 2 is an intermediate model. "•" denotes that the parameter estimate is significantly different from zero at the 5% level, while "+" denotes that it is not.

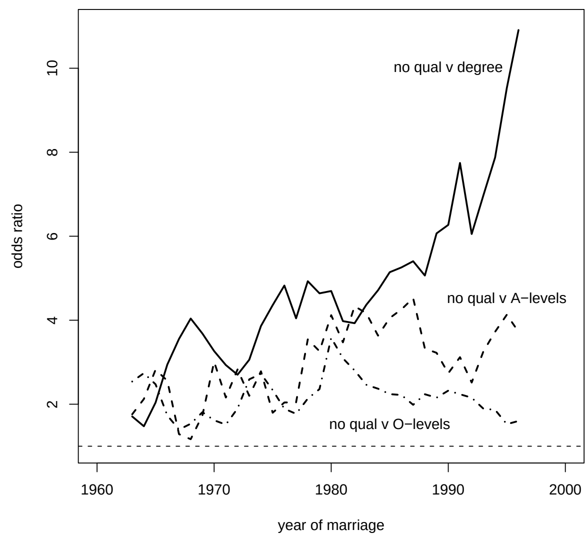


Figure 4: Trends in the association between premarital birth and educational attainment.

women with no qualifications. What explains this reversal? Without further investigation, we could only offer some speculative remarks here.

First, as noted in Section 2, when divorce was rare and the associated stigma strong, more resouces would be needed to cope with divorce. Since education is a good index for resource, we would expect better educated women to have higher divorce risks. As divorce becomes more common and accepted, more women would be able to do so, and this would imply a flattening of the educational gradient. Secondly, it is likely that with greater female access to the labour market, and pressure (from the housing market and elsewhere) to have two incomes, British society has become more accepting of educated and economically successful women. Thirdly, the expansion of higher education and the closing of the gender gap in educational attainment might have made universities and colleges more efficient institutions for sorting and matching potential spouses. Fourthly, as women's education improves, the *meaning* of the various educational categories would have changed too. In particular, having no qualifications has become increasingly uncommon (see Table 1), and this educational category might have become more strongly associated with unobserved characteristics that relate to both marital instability and low educational attainment (Hoem, 1997, p.26).

As for our results regarding children, it should first be noted that childbearing decisions and the decision to divorce are endogenous (Weiss, 1997). Couples with greater investment in their marriage (and children are a major marriage-specific investment) are more likely to stay together, and those couples who think that they might divorce are less inclined to have children. Empirically, it has been shown that although in a reduced-form (i.e. single equation) framework, children are associated with lower divorce risks, when childbearing and marital disruption are modelled jointly in a simultaneous equation framework, children could be associated with greater marital instability (see e.g. Koo and Janowitz, 1983; Lillard and Waite, 1993; Svarer and Verner, 2008). These results are indeed very insightful, but they are quite different from those reported in this paper. What we have shown is that, within the reduced-form framework, the gross association between children and marital stability has changed across marriage cohorts.

Why should there be a reversal of the familiar protective effect of children on marriage? This is a question for further research. For example, has there been a decline in the UK in the overall level of parental investment in children (cf. Sayer *et al.*, 2004)? But we could rule out a plausible argument, which goes as follows. Children are a marriage-stabilising factor because they are a marriage-specific investment (Becker *et al.*, 1977). And since step children receive less parental investment than biological children (e.g. Biblarz and Raftery, 1999; Case *et al.*, 2000), the reversal in question might be accounted for by the growing prevalence of step families in British society. We are not convinced by this argument, firstly because in this paper we consider children born to the respondent only (see note 6). Of course, children born to the respondent could be step children from the husband's point of view. We do not have information on the children's paternity. But step relationship might be thought more likely if the husband was previously married.[19] Thus, in analyses not reported here, we have tested for interaction effects between the children parameters and husband's previous marital status. It turns out that only one of the 24 interaction terms (4×6 cohorts) are significant at the 5% level. Furthermore, this significant interaction term pertains to the 1965-9 cohort, and so cannot explain the change in the 1980s.[20]

We have shown that the change in the children parameters is related to premarital birth, and that premarital birth is becoming ever more strongly associated with low educational attainment. This is important because, as we have shown, the less qualified have become more likely to divorce than graduates anyway. But they are also much more likely to have premarital

[19]For example, the husband could bring his own children from a previous marriage into the marriage in question, though this is unlikely as custody of children usually goes to the mother in the event of divorce.

[20]Details are available from the authors on request.

births, which means that their children are more likely to have a marriage-*destabilising* effect. The upshot is that children born to mothers with no qualifications are much more likely to go through a parental divorce. Compounding this is the link between low wage and no qualifications. Overall, the welfare implications for individuals, especially the children, is considerable. Thus, our results clearly echo some of the concerns for the 'diverging destinies' of US children (McLanahan, 2004).

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