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Who marries whom in Great Britain?

Tak Wing Chan
Department of Sociology
University of Surrey
Guildford GU2 5XH, UK
email: tw.chan@soc.surrey.ac.uk

Brendan Halpin
Institute for Social and Economic Research
University of Essex
Colchester CO4 3SQ, UK
email: brendan@essex.ac.uk

Non-technical Summary

Education is a very important economic and cultural asset in modern societies, and marriage patterns may intensify or compensate for inequalities in its distribution. In this paper we examine the factors affecting how quickly or slowly respondents to the BHPS marry for the first time, taking account of their and their ultimate spouse's education. Our concern is to understand what affects the tendency to marry and in particular whether there is change over time in the tendency for people to marry within educational groups.

Using life history data allows us to examine the effect of age and educational participation on the rate of first marriage, and how this varies over time. Using data on existing marriages is less informative because it loses information on the timing of marriage, and may be biased if there are significant numbers of marriages which break up before being observed. There are two important secular trends which can be expected to affect educational marriage patterns: the changing age of school-leaving and the changing age of first marriage. The former is rising through time and the latter shows a U-shaped trend. We can hypothesise that the greater the time gap between leaving education and marrying the weaker the tendency for educational homogamy, that is, like marrying like. Another important factor affecting the proportions marrying homogamously is the educational distribution: the more individual categories are large the greater the opportunity for people to marry at the same level.

We use two sets of models: the first separates three outcomes, marrying up, marrying homogamously and marrying downwards. This gives us direct estimates of the affect of, for instance, cohort, on homogamy. The second disaggregates the analysis and treats marriage to persons of each of four educational levels as four separate outcomes. Homogamy and heterogamy have to be estimated indirectly in these models, using the respondent's level of education as an independent variable. This indirectness is compensated for by the ability to separate the processes behind, for instance, graduates marrying graduates, from those behind homogamy among early school-leavers.

In many respects both sets of models tell us the same thing. Age has a very strong effect on the rate of marriage, and so does educational participation. Cohort has a strong effect, with the middle cohorts having a higher rate of marriage (and therefore a lower mean age of marriage). Time between leaving education and first marriage has an effect somewhat like that expected. In the first set of models we see an increasing tendency for men to marry upward or homogamously over time. For women we see a decreasing tendency to marry upwards for those with A-levels or O-levels.

The qualification-specific models are substantively more interesting, because they allow us to separate more things out, and we see, for instance, that the age patterns of the rate of marrying people of different educational levels are markedly variable. It also shows a clear 'gradient' in the chances of marrying a graduate or someone with low education, based on respondent's education.

It is difficult to address the issue of change in the tendency to homogamy over time with this sort of analysis, because of the change in the 'opportunity structure', that is the age- and time-specific distribution of possible partners. We have tried to control for this by including in the model measures of the relative educational distribution of members of the opposite sex in an appropriate age range. This has a significant impact on the models, and changes the observed effect of cohort and age. However, it is difficult to interpret this variable in its own right, as it is so closely correlated with the distribution of members of the same sex, and with other temporal changes.

Abstract

We investigate educational assortative mating, or homogamy, by modelling the hazard of entry to first marriage for a sample of residents of Great Britain. Using marital, and imputed educational, life-histories drawn from the British Household Panel Study, we estimate first a set of competing-risk models where the outcome variable is defined as respectively hypogamy, homogamy and hypergamy. Age, educational participation and cohort have very strong effects, in directions that may be expected. Time gap between leaving education and marrying shows some signs of the hypothesised effect (that a greater time-lag means less homogamy). Cohort differences suggest men are decreasingly likely to marry down and women decreasingly likely to marry up. We also fit a second set of competing-risk models, where the outcomes are marriage to individuals of specific educational levels. We argue that these are more stable models, and that they provide more insight into the actual marriage patterns. In both sets of models we attempt to control for the changing opportunity structure by including estimates of the educational distribution of single people in an appropriate age range. While this is obviously a necessary control variable, and while it has strong effects on other covariates, its own parameter estimates are hard to interpret. We cannot therefore claim that the effect of changing marginal distributions has been fully removed, but we feel our estimates are nonetheless improved.

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Department of Sociology
University of Surrey
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email: tw.chan@soc.surrey.ac.uk

Brendan Halpin
Institute for Social and Economic Research
University of Essex
Colchester CO4 3SQ, UK
email: brendan@essex.ac.uk*

1 Introduction

Does the UK differ in its patterns of educational intermarriage? Popularly Britain is perceived to be, on the one hand, more flexible in both economy and society than mainland countries, but also hidebound by class and having an economy in long-term relative decline. There may be only fragments of truth in these popular conceptions, but it is interesting to consider why, and whether, educational intermarriage has different characteristics in the UK.

The education system has had a role in the symbolic articulation of 'class' (more properly status differentiation) in British society through the years, especially with respect to accent. It has

*This paper is part of a multinational comparative project initiated by Hans-Peter Blossfeld and Andreas Timm, and is based on work first presented at a workshop at Sonderforschungsbereich 186, Bremen University, in December 1998. We are grateful for the comments and insights of the participants at that workshop.

provided as much a means of, as a barrier to, social mobility. On the other hand, it has also tended to produce relatively high numbers of poorly-qualified young men who moved into semi-skilled work in the one-time large manufacturing sector. However, this has changed in more recent years: as in most western countries, educational participation is growing strongly, such that acquisition of intermediate levels of education is becoming less socially discriminating in absolute terms. The manufacturing sector has all but disappeared with the consequence that leaving school with low or no qualifications is no longer an economically sensible thing to do. As in other European societies, female educational achievement has increased, overtaking male in many respects. Female labour force participation is relatively high in European terms, though much of it is part-time, suggesting that women's external contribution to the family economy is increasingly important, and thus also their level of education.

In this paper, we use data from the British Household Panel Study (BHPS) and event history models to study the process of educational homogamy from a dynamic life course perspective. The key questions we address are the following: First, has the level and the pattern of educational homogamy in Britain changed between cohorts? Instead of tracing how a single summary homogamy index has evolved over time, we will show how the hazard of homogamy varies over the life course and changes across cohort. Our second objective is to investigate the social forces which regulate the observed marriage pattern. Here we have found Mare's (1991) work particularly helpful.

2 Previous research – two social forces

Mare is primarily concerned with the trends of homogamy in the United States. He points to two social forces which may affect this trend – demographic trends, and the role of women in society.

2.1 Demographic trends

Mare argues that the probability of marrying someone with the same level of education should decrease with time since school departure. This is because if a person gets married soon after leaving school, his/her spouse is likely to be an old classmate, and school class is by definition homogeneous in educational attainment.

Also, 'because at successive levels of schooling, the ultimate educational status of students still in school is increasingly homogeneous' (Mare, 1991, p. 16), homogamy should be more prevalent among the better educated. Moreover, Mare argues that the effect of the time gap between school departure and marriage should be stronger among the better educated.

Given these, Mare argues that there are good reasons to believe that the level of homogamy in the U.S. has changed over time. This is due to historical trends of the timing of first marriage and school departure. As is well known, the average age of first marriage in the U.S. (and in other industrial countries) follows a U-shaped pattern: a substantial decline between the 1930s and the 1970s was followed by a rebound. At the same time, the average level of educational attainment has increased throughout the century, which means that young people leave school at progressively later ages. The combined effect of these trends is that the time gap between school departure and marriage has shortened between the 1930s and the 1970s, which should lead to higher level of homogamy. Since the mean age of first marriage has risen since the 1970s, while the school-leaving age has also increased, the time gap between school departure and marriage may have lengthened. Nevertheless, Mare (1991, p. 17) argues 'that marriages are now occurring at much later ages suggests that, in the absence of other causes of change, the trend toward greater educational homogamy may have weakened.'

2.2 Changing role of women in society

Furthermore, Mare argues that increased labour market participation of women affects what men and women expect from marriage. Specifically, quite apart from the roles of mother and homemaker, women are also expected to be breadwinners. This may lead to stro-

nger competition for better-educated women because they have higher earning power. This in turn may increase the level of educational homogamy.

By and large, Mare has found confirming evidence for his arguments. Using loglinear models, he analyses a set of mobility tables of recent marriages. He shows that the level of educational homogamy has increased between the 1930s and the 1970s. Then the association between husband's and wife's educational attainment has declined or remained stable during the 1980s. He also shows that the time gap between school departure and marriage is partly related to the level of homogamy. It is worth noting that Mare's interest is in trends over time, rather than, say, explicating the process generating the trends, and therefore loglinear models are particularly appropriate, having *inter alia* the strong advantage of controlling for changing marginal distributions.

3 Social changes relevant to homogamy in GB

Are the demographic trends and the trends related to women's role in society in Britain similar to those in the United States?

3.1 Timing of first marriage

In Britain, the timing of first marriage followed the well-known U-shaped pattern. Between 1910 and the late 1930s, the mean age of marriage was about 27 for men, and 25 for women. Then there was a trend towards early marriage such that by the late 1960s, the mean ages dropped to about 24.5 and 22 for men and women respectively. Since the 1970s this trend has been reversed. In 1990 the mean age of marriage was 26 for men and about 24 for women (see Coleman and Salt 1992, p. 181). This curvilinear pattern can also be found among the BHPS respondents. In Figure 1, we plot the percentage of the never-married by age for four birth cohorts. Readers can see that for both men and women the curves for the oldest and the youngest cohorts lie mostly above those of the middle cohorts.

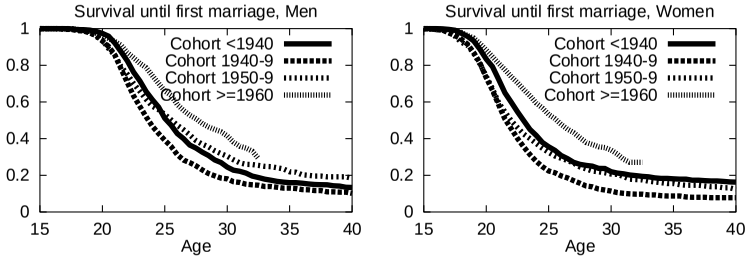

Figure 1: Survival to first marriage by cohort and sex

3.2 Educational attainment

As in other industrial countries, younger cohorts in Britain stayed in school longer and received progressively higher level of education (Kiernan and Lelièvre, 1995, p. 136). At the time when our oldest respondents went to school, the minimum school leaving age was 14. This was raised to 15 in 1947, and then to 16 in 1974. This trend can also be found among the BHPS respondents. Figure 2 shows a steady cohort-wise progression in numbers remaining in school.

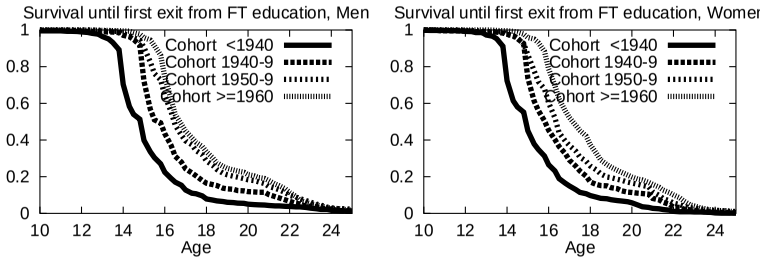

Figure 2: Survival in full-time education by cohort and sex

There is also a hint in Figure 2 that this progression may be slowing down for men and speeding up for women. This reflects the narrowing of the gender gap in educational attainment, which is evi-

Table 1: Educational achievement by 1992, by cohort and sex, percentages.

	Men				
	University	A-level, other sub-Univ	O-level, commercial	Sub-O-level, none	Total
Born					
before 1924	4.1	24.5	21.7	49.7	290
1924—1933	4.6	20.2	25.1	50.1	351
1934—1943	7.3	33.9	21.8	37.0	422
1944—1953	13.1	39.4	21.2	26.3	579
1954—1963	17.5	38.9	21.5	22.1	741
1964—1973	13.1	41.9	24.4	20.6	807
Total	11.6	35.7	22.6	30.0	3190
	Women				
	University	A-level, other sub-Univ	O-level, commercial	Sub-O-level, none	Total
Born					
before 1924	1.3	19.9	18.6	60.2	236
1924—1933	3.5	15.0	24.0	57.5	346
1934—1943	3.8	21.8	26.2	48.2	390
1944—1953	9.1	25.5	34.0	31.4	561
1954—1963	14.8	28.2	34.6	22.4	709
1964—1973	10.0	41.5	33.0	15.6	873
Total	8.8	28.5	30.6	32.1	3115

Source: BHPS

dent from Table 1. Among those born before 1924, 60 percent of the women and 50 percent of the men had no qualifications. However, for the 1964–73 birth cohort, there are fewer unqualified women (16 percent) than men (21 percent). Indeed, women of recent cohorts have surpassed men at O-level, and caught up with men at A-level, though there is still a gender gap at the degree level.[1]

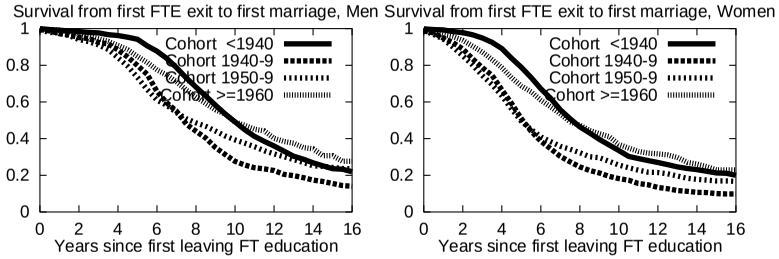

Figure 3: Time-gap between leaving education and first marriage, by cohort and sex

3.3 Time gap between school departure and first marriage

The combined effect of the above trends can be seen Figure 3, in which we plot the proportion of men and women who had stayed single by time since leaving school. Similar to Figure 1, the survival function of the oldest and the youngest cohorts are above those of the middle cohorts, though in this case the youngest cohort is closer to the middle ones for low duration (which, because of the greater

[1] We do not have systematic long term time series data on the extent of gender segregation by school in the U.K. But in the 1990s, less than 10 percent of all schools are single-sex schools, which tend to be concentrated in urban areas (Ofsted, 1998). West and Hunter (1993) claim that 'almost all children are educated together until the age of 11', but the pattern changes a lot afterwards, with much local variation. They claim that in London in 1985, 50 percent of the girls and 46 percent of the boys attended single-sex schools, and that 38 percent of secondary school students in Kent attended single-sex schools.

mean school-leaving age, are mostly at higher ages than for those cohorts)

3.4 Female labour force participation and gender wage gap

In the late 1960s, just over half (55 percent) of all women were in the labour force. This has risen to over two thirds (69 percent) in the 1990s. The increase was most impressive among women with pre-school children (from 22 percent to 42 percent). However, the increase is concentrated in the part-time work force. Accompanying this trend is a slow closing of the gender wage gap. In 1968, women's hourly wage rate was 60 percent of men's, rising to about 67 percent in 1990 (see Davis and Joshi, 1998)

3.5 The growth in cohabitation

Alongside the decline in marriage rates in recent decades there has been a dramatic rise in the incidence of cohabitation. Between 1980 and 1995 the proportion of women aged 20–24 in co-residential partnerships who were not in legal marriages rose from 11 percent to 55 percent (General Household Survey data quoted by Ermisch and Francesconi 1998). As Table 2 shows, this process is accelerating, to the extent that the rise in cohabitation more or less offsets the fall in marriage, such that the numbers in co-residential partnerships remain more or less stable.

Ermisch and Francesconi (1998) show that this new cohabitation phenomenon is predominantly serving as a precursor to legal marriage: first partnerships which begin as cohabitations tend to be relatively short (median duration about 2 years) and about two thirds end in marriage.

This has a number of implications for marriage patterns. First, the nature of marriage has now changed, representing a more formal option for partnership rather than the only acceptable form. Thus some couples who would have married will now cohabit instead, but also some couples who would not have married will now cohabit. Second, the processes by which couples arrive at first marriage have changed, in that the practice of cohabitation prior to marriage can be seen as a matching process that may have fundamental differences to that operating where marriage occurs without prior co-residence. This could mean that marriages are now more likely to be good matches, if cohabitation filters out bad matches, but it could also mean the opposite, if couples form more casually (and then persist) when the initial commitment is less.

Table 2: Proportions (per 1,000) ever in partnership by age 24

Cohort	Men		
	Marriage	Cohabitation	Total
1930–49	473	29	502
1950–62	327	163	490
1963–76	110	387	497
	Women		
	Marriage	Cohabitation	Total
1930–49	700	33	733
1950–62	544	207	751
1963–76	210	448	658

Source: extracted from Ermisch and Francesconi (1998).

It is hard to predict the net effect on educational marriage patterns: the 'good' matches which endure may have more to do with personal characteristics that affect compatibility than with educational credentials, allowing the subsequent marriages to be less educationally homogamous than they would otherwise be. On the other hand, cohabitation may mean that the social networks relevant to a subsequent marriage are those the individuals were in at the commencement of the cohabitation rather than at the time of marriage. This will 'preserve' the effect of the homogeneity of the school-based social world for several years after it has been left, and thereby increase homogamy arising from Mare-style factors.

4 The BHPS data and its limitations

Researchers studying homogamy are interested in the process of family formation. However, estimates based on prevailing marriages would be biased if homogamous and heterogamous marriages dissolve at different rates. The typical solution to this long recognised problem is to analyse recent rather than prevailing marriages (Kalmijn, 1991).

Since the BHPS is a longitudinal study, we should in principle be able to focus on the marriage formation process alone, and thus avoid the above bias. Indeed, the BHPS collects data on the timing of all marriages of the respondents. However, there is no information on the spouse except for current marriages (as of 1992) or marriages formed during the panel period. Since we need to know the spouse's qualification in order to determine whether a marriage is homogamous, this means that we cannot consider marriages dissolved before 1992. In effect, our data pertain to prevailing marriages and suffer from the bias discussed above.

5 Modelling strategy

5.1 The event history data set

The event of interest is getting married for the first time. Since cohabitation as a prelude to marriage is becoming a 'majority practice' (Kiernan and Lelièvre 1995, p. 130, Ermisch and Francesconi 1998) a case could be made for treating the beginning of cohabitation as the event of interest. At very least, the date of marriage will on average be later than the true start of partnership. However, this may lead to two problems. First, cohabitation is itself a diverse phenomenon. Some cohabitation spells are very short, suggesting that, even from the point of view of the respondent, they may be quite different from those which lead to eventual legal marriage. It is of course very difficult to differentiate the various forms of cohabitation. Some transient cohabitations will also be unreported, thus biasing our sample. The second problem is that the meaning of cohabitation has changed across cohort. Given these problems, we have decided to

persist with using legal marriage as the event of interest, though it is clear that our results have to be read in the knowledge of this striking change in behaviour.

We follow the life history of the respondents from the age of 15 until their first marriage, or for those who had remained single, age 60 or the interview year, whichever came first. We update the following time-varying covariates each year: age, whether the respondent was in full-time education, time since leaving full-time education, time spent in school since age 15, qualification achieved, and the opportunity structure of marriage. We also have a set of time-constant covariates – year of birth, sex, and spouse's qualification.

We describe these covariates and our classification scheme for educational attainment in Table 3, which should be fairly self-explanatory. Here we wish to highlight three points. First, we distinguish four levels of educational attainment. The highest level is university degree. This is followed by 'A-level', which includes post-secondary qualifications such as nursing, teaching and other higher qualifications.[2] The third level includes O-levels, certain commercial qualifications and apprenticeship. The fourth level is no or sub-O-level qualification.

The second point is that the qualification history of each respondent is imputed from three sources: his/her educational participation history, educational attainment in 1992, and typical age of acquisition of antecedent qualifications.

Thirdly, as we have seen in Table 1, the distribution of educational attainment by sex varies substantially across cohort. This directly affects the pattern of homogamy and heterogamy that is possible. If 90 percent of both men and women have no qualification (as in Spain around 1900, see Gonzáles, forthcoming), there will be far more homogamy than in a society where educational attainment is more evenly distributed. This effect of the 'margins' is controlled for in a loglinear analysis of marriage pattern. But in a hazard modelling approach such as the present, cohort difference in the educational distribution by sex may be confounded with other secular

[2] For older cohorts, teaching qualifications would not necessarily have involved university education, and therefore we have included them in this category. At least some respondents from younger cohorts with degree-level teaching qualifications are coded as having degrees.

Table 3: Covariates used in the analysis

Variable	Description	
Log(current age – 15yrs)	Age measured in months	
Log(60yrs – current age)	Age measured in months	
Not in school	Not-in-School = 1 In-School = 0	
Time since leaving school	Length measured in years	
Duration in school	Duration measured in years since age 15, reverts to zero on leaving education	
Year of birth	1900=0, 1901=1, and so on	
Education	1 = Degree level	Higher degree, First degree
	2 = A-level	Teaching qualification, Other higher qualification, Nursing qualification, GCE A level
	3 = O-level	GCE O level or equivalent, Commercial qualification, Apprenticeship
	4 = Sub-O level	CSE grade 2-5, Scot G, Other qualification, No qualification

changes. In an attempt to control for this, we introduce a measure of the opportunity structure, which is based on the proportion of single persons of the opposite sex, in a defined the age-range, in each of the four qualification levels, for each calendar year.

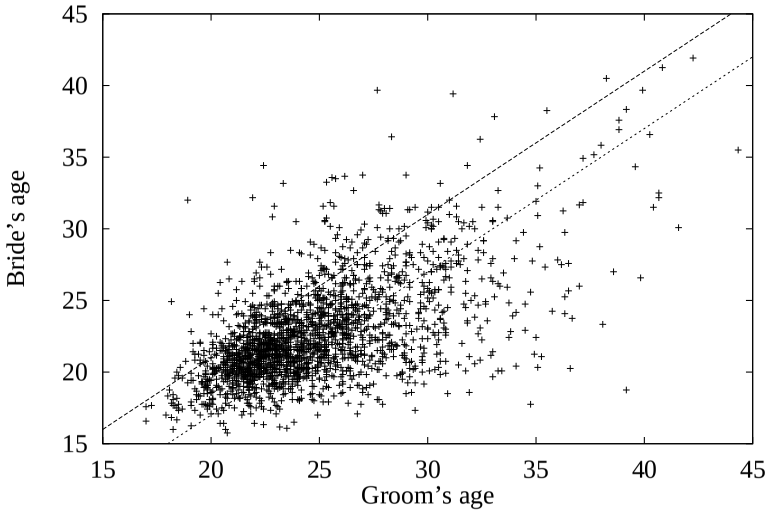

Figure 4: Age differences in marriage

This measure is also based on the BHPS data. For each male respondent, we consider the number of single women who were between three years younger and one year older than him, for each calendar year.[3] For women, we consider the reverse age range of one year younger to three years older. This approximates the joint distribution of age at marriage. In our data, husbands were on average 2.15 years older than their wives, and the band we use accounts for almost 64 percent of the cases (see Figure 4).

Figure 5 represents the opportunity structure visually. Looking vertically down the years we can see the enormous growth in the

[3] In practice, in order to smooth the data, the calculations were done on five-calendar-year bands, with subsequent interpolation.

proportions holding higher qualifications. We can also see some cohort effects, in the form of bulges which shift five units to the right for each unit down (which may be due to sampling variation).

6 Defining the transitions

The dependent variable of our analysis is time until first marriage. But what are the origin and destination states which define the transition? We approach this question in two ways. First, we follow Blossfeld, Timm and Dasko (forthcoming): All respondents started as being single. If they got married, we distinguish three destinations: upwards (if spouse has a higher level of education than respondent), homogamy (spouse has the same level of education), and downwards (spouse has a lower level of education). Because people at the top educational category are not at risk of marrying up, and those at the bottom category cannot marry down, the risk set is different for the three transitions. We estimate our model for the three transitions separately, with the risk set suitably adjusted. We shall refer to this as the BTD model.

However, upward, homogamous and downward marriages, as defined above, are all very diverse and heterogeneous categories. Consider two graduates, one marrying an unqualified person, and the other marrying someone with A-level. Under the BTD model, both instances would be considered as downward marriages. We believe this may obscure meaningful differences. Also, the BTD model assumes that all homogamous marriages are in some sense similar. However, the social forces which regulate the marriage of two graduates may be quite different from those for two unqualified persons. To model these transitions more flexibly, we have carried out a second set of analyses. We refer to this as the qualification-specific model, using spouse's qualification to define destination, and respondent's qualification as a predictor (see below). We use the piecewise exponential model in our analysis. Because there are multiple destinations in both set of analyses, ours are competing risk models.

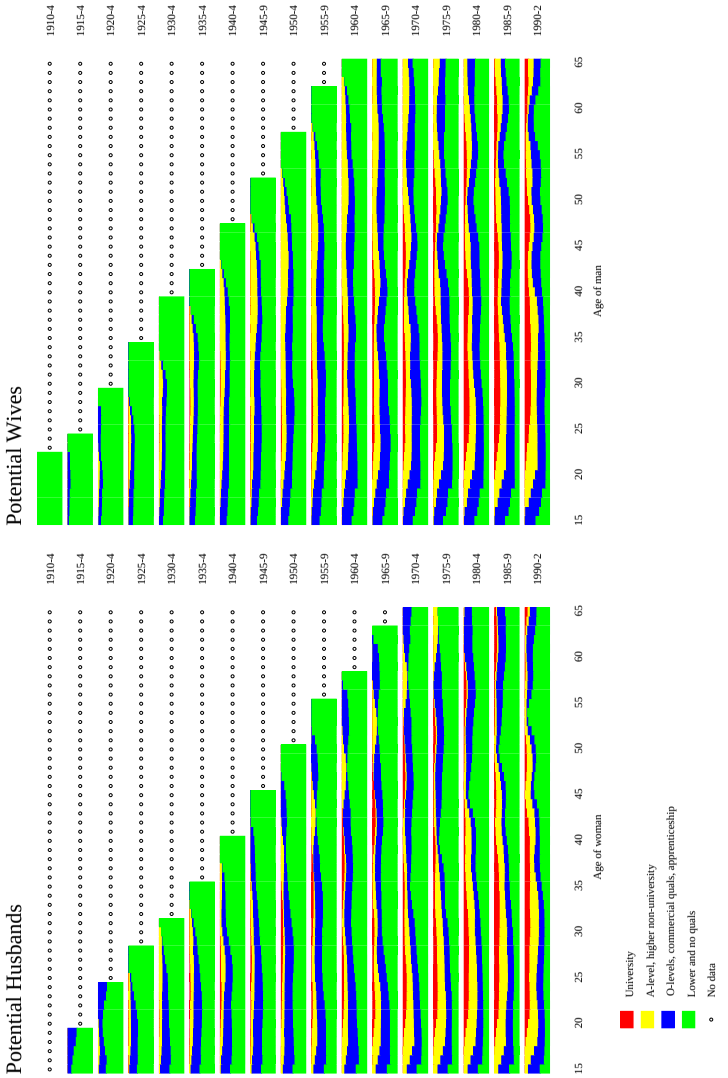

Figure 5: The qualification distribution of single people, by year of age and five-year calendar band

Table 4: Educational marriage patterns by sex and cohort, observed and expected rates under independence

	Men					
	Up		Homogamous		Down	
Born	Observed	Expected	Observed	Expected	Observed	Expected
before 1924	13.67	20.78	47.66	39.29	38.67	39.93
1924—1933	20.77	25.32	45.05	38.21	34.19	36.47
1934—1943	19.44	26.03	37.12	31.01	43.43	42.95
1944—1953	19.13	27.56	39.64	27.43	41.22	45.01
1954—1963	24.71	32.05	38.02	26.97	37.26	40.99
1964—1973	31.72	35.16	36.56	28.54	31.72	36.30
	Women					
	Up		Homogamous		Down	
Born	Observed	Expected	Observed	Expected	Observed	Expected
before 1924	39.89	41.12	45.36	37.44	14.75	21.44
1924—1933	36.84	38.59	43.75	37.47	19.41	23.94
1934—1943	42.19	42.07	38.36	31.88	19.45	26.05
1944—1953	36.96	42.32	42.02	27.78	21.01	29.90
1954—1963	39.40	42.40	37.28	27.13	23.32	30.47
1964—1973	30.77	37.83	42.69	28.48	26.54	33.69

7 Results

7.1 Observed trends

We report the observed trends of homogamy and heterogamy in Table 4. As readers can see, for men there is an upward trend for upward marriage and a downward trend for downward marriage. For example, for men born before 1924, 14 percent married up, 39 percent married down. These changed to 32 percent for both upward and downward marriage for the 1964–73 birth cohort. The trend for women is just the opposite – more women are marrying down, and fewer of them are marrying up. This corresponds to the closing of the gender gap in educational attainment. We also observe an uneven decline in the proportion of homogamous marriage for men across cohort. To analyse the social forces, such as age, cohort, qualification, and opportunity structure, which produce this observed pattern, we now turn to a multivariate analysis of the hazard of marriage.

7.2 The BTD model

As noted above, we have carried out two sets of analyses. In the first set, we follow Blossfeld *et al.* and define the transitions as from the origin state of never married to one of three destination states – upward, homogamous, or downward marriage. The results are reported in Table 5.

The two parameters, $\log(\text{age} - 15)$ and $\log(60 - \text{age})$, are used to approximate the non-monotonic pattern of age dependence. For all three transitions, and for both sexes, the second parameter is much stronger than the first. This implies that the distributions of the hazard rates by age are skewed to the right – most of our respondents got married when they were relatively young. We have also included two interaction terms (age-dependence by qualification). But they are significant for homogamous marriage only. The sign of the interaction terms suggests that the skewness in age dependence is more pronounced for the less qualified.

The covariate 'not-in-school' is positive and significant for all but one transition. Its effect is also very strong. For example, other

Table 5: Modelling homogamy: parameter estimates

	Men		
Parameter estimate	Upward	Homog.	Downward
Constant	-71.4496*	-66.1893*	-86.3718*
Age effects			
Log(age–15)	2.6401*	2.6978*	2.2431*
Log(60–age)	7.4536*	6.9288*	10.8242*
Log(age–15) by qual	-0.1310	-0.2043*	-0.1628
Log(60–age) by qual	0.1583	0.1871*	0.0650
Not in school	3.3923	1.6569*	1.8956*
Months in current educational spell	0.1017	0.0679	0.0708
Time since leaving education			
Linear months	-0.0099*	-0.0048†	0.0109*
Quadratic[a]	0.0200*	0.0089†	-0.0208*
Cubic[a]	-0.0003†	-0.0001‡	0.0003*
Cohort effect			
Linear year of birth	0.0592†	0.0769*	0.0924*
Quadratic[a]	-0.0759*	-0.0955*	-0.1339*
Opportunity structure	2.0233†	2.0502*	-0.0080
	Women		
Parameter estimate	Upward	Homog.	Downward
Constant	-58.9621*	-60.7200*	-79.9888*
Age effects			
Log(age–15)	1.6645*	1.8598*	0.3568
Log(60–age)	6.5375*	6.2570*	10.3977*
Log(age–15) by qual	-0.1509‡	-0.3373*	0.0664
Log(60–age) by qual	0.0987	0.2682*	-0.0563
Not in school	1.9102*	4.1950*	3.1870*
Months in current educational spell	0.0095	0.4008*	0.2815†
Time since leaving education			
Linear months	-0.0070*	-0.0036	0.0144*
Quadratic[a]	0.0117†	0.0084‡	-0.0265*
Cubic[a]	-0.0002‡	-0.0002†	0.0004*
Cohort effect			
Linear year of birth	0.1240*	0.1012*	0.1328*
Quadratic[a]	-0.1641*	-0.1168*	-0.1447*
Opportunity structure	2.6822*	1.4631*	1.4111

Notes: *: Significant at 1%; †: significant at 5%; ‡: significant at 10%; (a) Squared time gap in months divided by 900, cubic in months divided by 27,000; squared year of birth divided by 100.

things being equal, being out of school increases the homogamy hazard by a factor of five ($e^{1.6569}$) for men and a factor of sixty-six ($e^{4.1950}$) for women.

The variable 'duration in school' is significant for women's homogamous and downward marriage only. For each additional year a woman stays in school since age 15, homogamous hazard increases by 49 percent ($e^{0.4008}-1$) and downward hazard increases by 33 percent ($e^{0.2815}-1$). We would stress that this is a very specific effect, as the variable measures duration since age 15 for those currently in education (who, as we have just seen, have a substantially lower hazard of marrying), and this covariate reverts to zero on leaving education. Thus, given that a woman is in education, over time she becomes more likely to marry homogamously or downward. This is an attenuation of the strong opposite effect of being in school.

We use a set of linear, quadratic and cubic terms to capture the effects of the time gap between school departure and first marriage. For men, time since leaving school depresses the upward and homogamy hazards, but raises the downward hazard. The quadratic terms are of the opposite sign to the linear terms, suggesting that these trends level off (or bottom out) as time gap increases. Recall Mare's argument that the longer the time gap, the lower the probability of homogamy. Our results confirm his argument in part. For men, time gap does reduce the homogamy hazard, and raises the hazard of marrying downward. But staying single also reduces upward hazard. The pattern for women is similar to that for men, with the exception that the linear time gap parameter is not significant for homogamy.

Cohort difference in marriage timing is represented by a pair of linear and quadratic 'year of birth' variables. For all destinations and for both sexes, the linear term is positive while the quadratic term is negative. This leads to a curvilinear pattern of the hazard of marriage which peaks for those born between 1930 and 1945, and declines thereafter. The magnitude of the cohort effect is generally greater for women than for men; but within gender, the estimates are quite similar across destinations. In general, the cohort effects are quite substantial, and depending on the values of other variables, particularly age and qualification, they change the rank order of the hazards, as we will see later.

As noted above, the opportunity structure variable is a simple measure of the proportion of single people of the opposite sex in the defined age-range who have the appropriate qualifications.[4] It is positive and significant for upward and homogamous marriage, and the effects are very strong. To illustrate, for men as the proportion of single women with better qualification increases by 1 percent, the hazard for upward marriage increases by 2 percent ($e^{2.0233/100} - 1$). However, this variable is not significant for downward marriage. Dropping this variable from the model affects mainly the estimates of the cohort trend, but also the age effects, particularly for upward marriage.

To help visualise the combined effect of the covariates, we use the estimates reported in Table 5 to generate a set of predicted hazard rates for men and women of two birth cohorts, 1940 and 1960 (see Figure 6).[5] Several things are notable. First, for both men and women, the hazards of marriage are generally lower for the 1960 cohort than for the 1940 cohort.

Secondly, for men, over the prime marriage years of the life course, downward hazard is greater than the hazard of homogamy, which in turn is greater than upward hazard. This is true for all except the unqualified (for whom upward hazard dominates). This pattern holds for men of both cohorts. Within the stable rank order, we also observe a relative rise for upward and homogamous hazards. For women, the rank order of the hazards is not so clear cut, and it changes between cohorts. Specifically, relative to homogamous or downward hazard, upward hazard declines for women with A-levels or O-levels. Upward hazard also declines in comparison with homogamous hazard for unqualified women, but not to an extent

[4] Appropriate, that is, as defined by the transition concerned. For example, under the BTD model, the opportunity structure for upward mobility for a man with O-levels is the sum of the proportion of single women in the defined age-range for that particular calendar year who have a degree or A-level qualification.

[5] In generating the predicted hazards, we make three assumptions: (a) that O-level, A-level, and university-type qualifications were obtained at age 16, 18 and 21 respectively; (b) that our respondents left school when they acquired their highest qualification, or at age 16 if they have no qualification; and (c) that people are uniformly distributed across the four educational groups. Thus, for instance, a respondent with no qualification will have an opportunity structure of upward marriage of 75 percent, and all respondents have an opportunity structure of homogamy of 25 percent.

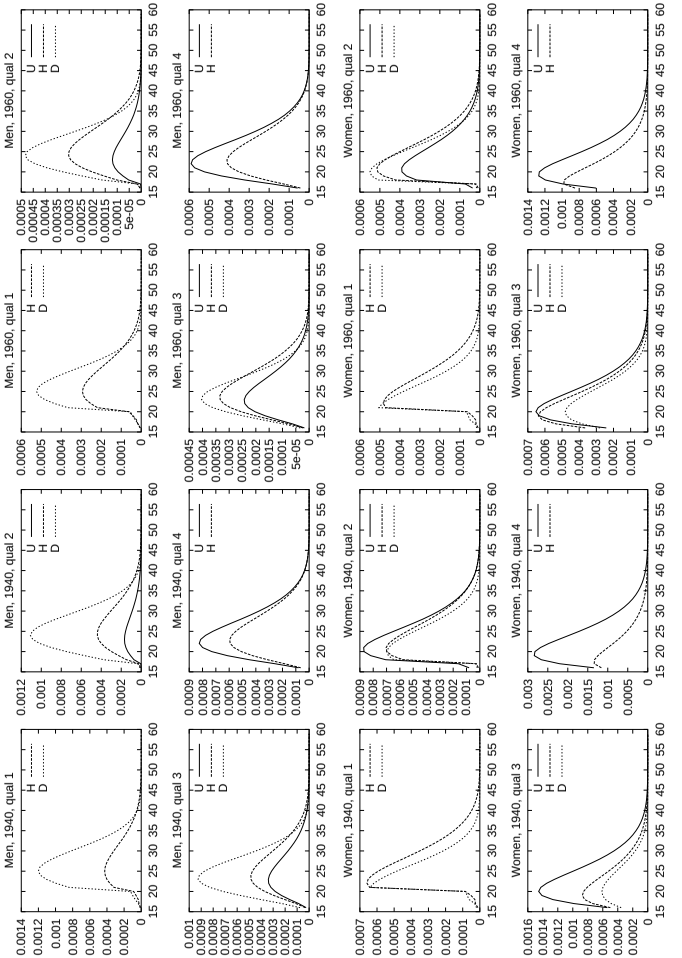

Figure 6: Predicted hazard rates from the BTD model

which changes their rank order.

Thirdly, the age profile of male and female graduates provides an interesting contrast. In the case of female graduates, downward and homogamous hazards shoot up very sharply as they leave university, followed by a steep decline. For male graduates, hazard rates also rise as they finish university, but the rise is not quite as sharp. In fact, there is a gradual climb before the hazards peak, suggesting that most male graduates work for several years before they get married. Similarly, the subsequent decline in hazard rates is somewhat less abrupt for male than for female graduates. Also, downward hazard of male graduates clearly dominates over their homogamous hazard. In contrast, the two hazards are very similar in the case of female graduates.

7.3 The qualification-specific model

While the results of the BTD model are informative, we believe they are limited for reasons explained in section 6. We now report the results of a second set of analyses which uses spouse's qualification to define the destination state. The respondent's own qualification enters into the model as a covariate. This allows us to estimate the hazard of moving from any origin state to any destination state.[6] We have tested for interaction between respondent's own qualification and many predictors in the model. Most turned out to be non-significant and are dropped. The model reported in Table 6 is our preferred model.

There are similarities between the BTD model and the qualification-specific model. For example, 'not-in-school' is again found to be positive and significant for all transitions and for both sexes. Its effect is also very strong. We have tested for interactions between this predictor and respondent's own qualification, but none are significant.

The general shape of age dependence is the same across destinations, and similar to what we saw under the BTD model, the

[6] Ideally speaking, we would also like to use respondent's qualification to define the origin state. This would imply modelling 16 distinct transitions separately. In practice, this leads to numerical problems in estimation.

Table 6: Modelling selective marriage: parameter estimates, men and women

	Qualification of bride				Qualification of groom			
	Univ	A-Level plus	O-Level etc	Low/no quals	Univ	A-Level plus	O-Level etc	Low/no quals
Constant	-122.5978*	-36.7802*	-74.6716*	-101.8597*	-94.6412*	-94.1400*	-96.7797*	-92.0637*
Age effects								
Log(age–15)	4.6276*	0.5813	1.7308*	3.4690*	2.8281*	3.0323*	1.2535*	2.1955*
Log(60–age)	14.6898*	4.6702*	9.1087*	14.1906*	12.2680*	14.2679*	11.8878*	13.2536*
Log(age–15) by qual				-0.8376*		-0.5649†		-0.5423*
Log(60–age) by qual				-2.2086*		-2.1827†		-2.0819†
Not in school	1.1629*	1.3288*	1.7990*	1.6181*	0.6373†	2.3783*	1.7658*	2.2067*
Time since leaving education								
Linear months			0.0068*	0.0128*		-0.0042‡	0.0174*	0.0098*
Quadratic[a]			-0.0014†	-0.0030*		0.0011†	-0.0037*	-0.0022*
Cubic[a]			0.0008†	0.0017*		-0.0007†	0.0020*	0.0012*
Cohort effect								
Linear year of birth	0.1195†	0.0053	0.0207	0.0634*	0.1395*	0.0722*	0.1153*	0.0668*
Quadratic[a]	-0.1271†	-0.1027*	-0.1025*	-0.1228*	-0.1774*	-0.1634*	-0.1346*	-0.1061*
interaction, qual 2		0.0272*	0.0227			0.0465*		
interaction, qual 3		0.0307*	0.0378*			0.0310†		
interaction, qual 4		0.0414*	0.0530*			0.0331†		
Opportunity structure								
A-Level plus		3.8825*	4.9974*	9.5820*	-4.5322	2.8473‡	6.9795*	5.6588*
O-Level etc.		-2.0523	3.4396†	4.9765*	-1.9148	0.2768	5.9208*	5.5503*
Low/no quals		-4.9360†	-0.9914	4.2001†	-5.6487‡	-2.8431†	3.6703†	1.4961
Qualification effect								
A-Level plus	-1.4844*	-1.4829*	-0.3470	18.0659*	-0.7243*	14.0711†	0.9466†	16.5404*
O-Level etc.	-2.8857*	-1.8115*	-1.2636‡	35.8129*	-1.6817*	30.7863†	0.8926†	31.8509†
Low/no quals	-3.9971*	-2.8361*	-2.2259*	52.9557*	-2.5472*	46.0731†	1.1410*	47.5453†

Notes: * : Significant at 1%; † : significant at 5%; ‡ : significant at 10%; (a) Squared time gap in months divided by 100, cubic in months divided by 100,000; squared year of birth divided by 100.

second term is stronger than the first. However, the qualification-specific model reveals considerable variation in the magnitude of these terms by destination. As we shall see (Figure 9), this leads to substantial and important differences in the overall life course pattern of hazard rates for the various origin and destination combinations. There is significant interaction between the age dependence terms and respondent's qualification for three destinations, namely for men marrying unqualified women, and for women marrying A-level or unqualified men. The signs of these interaction terms suggest that the age-dependence pattern is generally flatter for the less qualified.[7]

The general pattern of the cohort effects is also similar to what we saw earlier – a positive linear term moderated by a negative quadratic term, leading to a curvilinear pattern. But here we see considerable variation between destination in the magnitude of the linear term. As a result, unlike the BTD model there is within gender much variation by destination in the magnitude of the net effect of cohort, and in when it peaks. For example, for men marrying graduates, the curvilinear pattern is most pronounced, and peaks for those born in the mid-1940s. In contrast, for men marrying women with A-levels or no qualification, the net cohort effect is weaker and peaks for the 1920 birth cohort. For women there is also variation by destination in the net cohort effect, the magnitude of which is greatest for marriage to graduates and O-level men.

The interaction terms between respondent's qualification and the linear cohort trend are significant for three destinations. Thus, there is variation by 'origin' as well as by destination.

We use three predictors to capture the effect of the opportunity structure, each measures the relative size of a potential spouse pool. Transition to a particular destination is affected not only by the size of that destination, but also by the size of other destinations. For example, if A-level people and O-level people are substitutes in the marriage market, then the parameter for the A-level spouse pool will be negative for the O-level destination. This is the main reason why

[7] A cautionary note is in order: because the model contains interaction terms between qualification and age, and between qualification and cohort, the main effects of these predictors cannot be interpreted in isolation from their interactions. We will illustrate the complex interaction pattern graphically below.

we have included three predictors instead of just one.

Having said that, we hasten to add that because the relative size of qualification groups covaries with each other, the effects picked up by our opportunity structure terms may partly reflect such incidental covariation, rather than the substitution effects that we are interested in. This is an identification problem, for which we have no satisfactory solution. As readers will see, this means that in practice we find it difficult to interpret some parameter estimates.[8] Nonetheless, we would expect the term for the corresponding destination to be the strongest. Our findings are as follows.

First, for both men and women, in relation to the three destinations of A-level, O-level, and no qualification, the size of the destination category has the expected positive effect. Thus, for example, the higher the proportion of single women with A-level, the greater the hazard of men marrying an A-level woman. Secondly, again for both sexes, opportunity structure has no effect on the hazard of marrying graduates. The exception here is that the more unqualified men there are, the lower the hazard of women marrying a graduate. We believe this simply reflects the negative correlation between the relative size of the two destinations.

Thirdly, in relation to the destinations of O-level and no qualification, the proportion of people with more qualifications has a positive effect on the hazard rates. This may have to do with positive covariation in the relative size of the categories concerned.

Fourthly, the hazard of men marrying A-level or O-level women is negatively affected by the proportion of women with less qualifications. For example, as the proportion of unqualified single women increases by 1 percent, the hazard of men marrying A-level women dropped by 5 percent ($e^{-4.936/100} - 1$). Thus, it seems that unqualified women are substitutes for A-level and O-level women in the marriage market.

We observe an interesting pattern for the main effect of qualification, which can be described as a positive qualification gradient

[8] We also have interpretation problems with the model with the opportunity structure parameter for the relevant destination only. In that model, the opportunity structure terms for marriage to unqualified people are large and negative, implying that the more unqualified people there are, the lower the hazard of marrying them.

– the better qualified you are, the higher your hazard.[9] For men, this applies to the hazard of marrying women with a degree, A-level or O-level. For women, this applies only to the hazard of marrying graduates.

However, as we note in footnote 7 above, because qualification interacts with cohort and/or age for five of the eight destinations, the net effect of qualification varies with age and cohort. To illustrate the complex interaction pattern, we have used the parameter estimates of Table 6 to generate a set of predicted hazards (see Figure 7). Each panel in this figure corresponds to a particular destination (*i.e.*, a column in Table 6). These figures reveal a slightly modified picture of the effect of respondent's qualification.

Let us first consider the 1940 birth cohort: The positive qualification gradient applies to men marrying women with degrees or A-levels, and to women marrying graduates. For these transitions, education improves a person's position in the marriage market. In contrast, a negative qualification gradient applies to women marrying O-level or unqualified men. For these transitions, the higher the qualification, the lower the hazard. For the remaining transitions, there is no simple clear-cut pattern.[10]

What is notable is the asymmetry between men and women in where the positive qualification gradient applies. Qualification is a good predictor for future earnings, and is probably considered partly in such terms in spouse selection. Our result is, therefore, consistent with the view that in the marriage market grooms, far more often than brides, are valued for their earning prospects: competition among women is mostly for the highest-educated men, while competition among men extends further down the scale.

We also observe some interesting cross-cohort changes, which is most noticeable for the marriage market for A-level men and women. Specifically, the hazard of male graduates marrying A-level

[9] Given the way we order the four qualification categories (the highest level being 1, and the lowest level being 4), the positive gradient is indicated by negative parameter estimates, and vice versa.

[10] If one accepts the argument that male graduates generally do not compete in the marriage market for women with O-levels or no qualifications, and that A-level men do not compete for women with no qualification, then a positive qualification gradient can be said to apply for these transitions as well for men with the remaining qualifications.

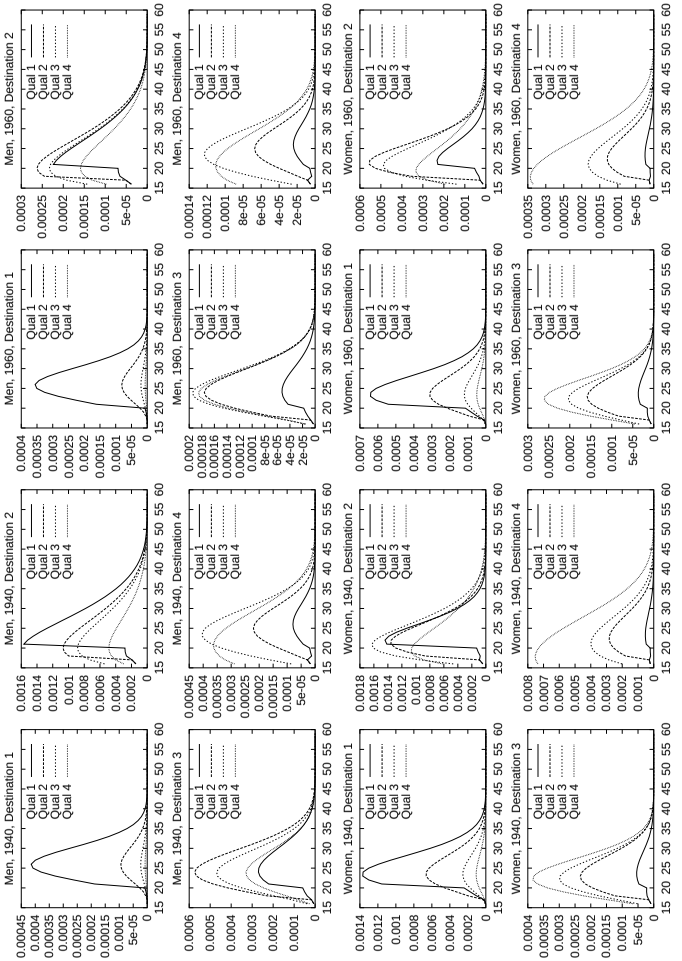

Figure 7: Predicted hazard rates from the qualification-specific model, grouped by destination

women has declined relative to the hazard of men with less qualifications. There is a relative decline across cohort in the hazard of female graduates marrying A-level men.

How do we interpret these changes? We note earlier the argument that women's changing role in society may have affected the dynamic of spouse selection and hence the pattern of educational homogamy. In particular, with the closing of the gender gap in educational attainment, and the gradual (and slow) narrowing of the gender wage gap, brides may increasingly be valued for their contribution to the family income. If this is true, we would expect those women whose wage rate has increased the most to have become more attractive in the labour market. This may lead to a higher level of homogamy among graduates (which is indeed evident in our data, see the relative rise in homogamy hazard for graduates reported in the middle panel of Figure 8.) As a result, there could be a downward trend for male and female graduates to marry a spouse with A-levels. Hence the pattern reported in the last paragraph. We hasten to add that this is a speculative account, and the cross-cohort change described above may have other causes. We will leave this issue for future investigation.

Figure 9 shows the predicted hazards under the qualification-specific model again, but here the hazards are re-grouped by respondent's qualification. We present these figures because of two reasons. First, they are comparable to Figure 6 for the BTD model, and thus facilitate a direct comparison between the two models. Secondly, grouping the hazards by respondent's qualification allows us to think from the point of view of the respondent: given the qualification of a person, what are competing hazards facing him/her. The following are notable.

Like the BTD model, readers can see that the predicted hazards are uniformly lower for the 1960 cohort than for the 1940 cohort. However, unlike the BTD model, we see in these figures as much change in the rank order of men's hazard rates as there is for women. In the case of women, one noticeable change is at the top end of the education hierarchy. Specifically, women graduates of the 1940 cohort had about the same hazard of marrying a male graduate or a A-level man. For women graduates of the 1960 cohort, the hazard of the marrying a graduate clearly dominates the hazard of marry-

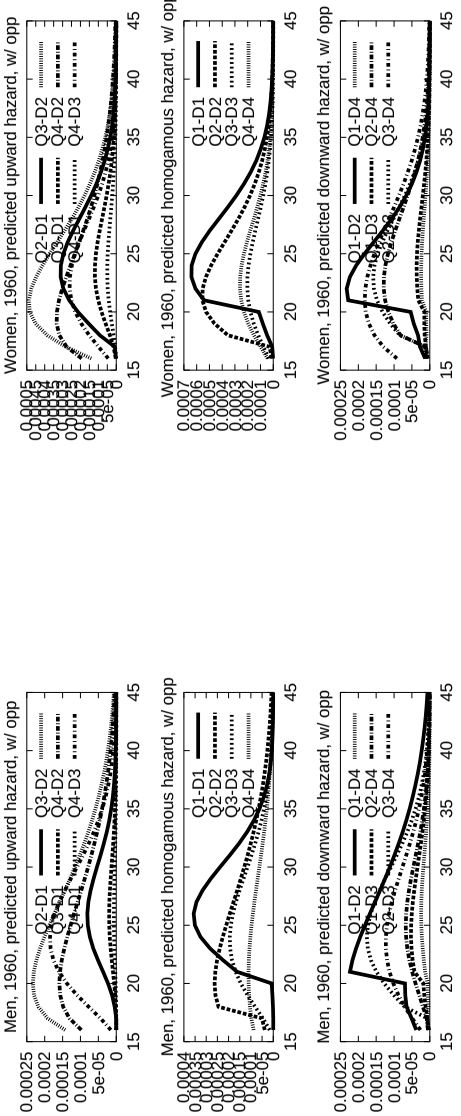

Figure 8: Hazards of upward, homogamous and downward marriage, qualification-specific model, men and women, 1960 cohort

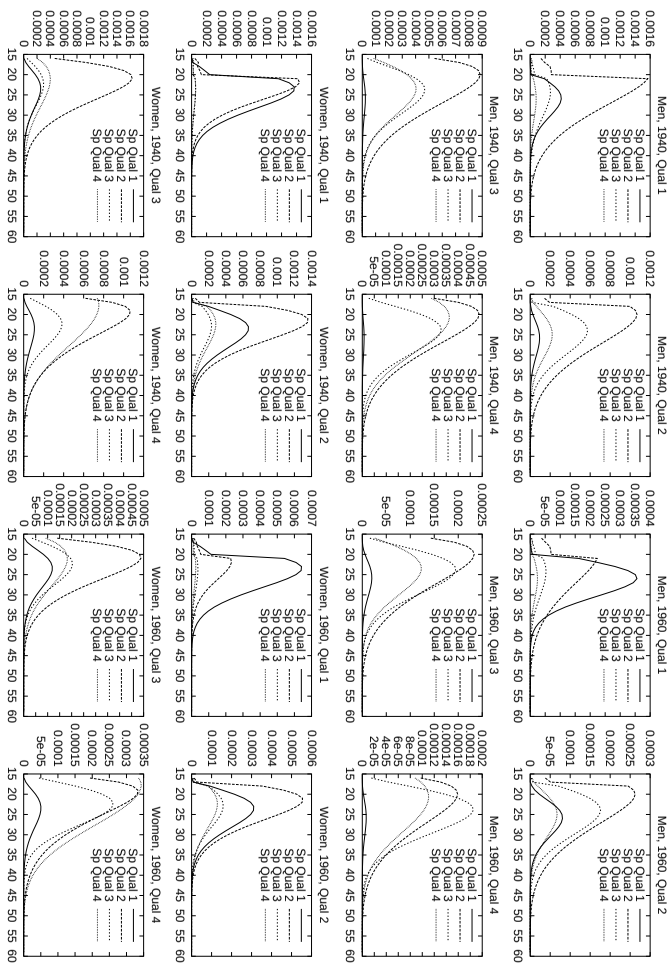

Figure 9: Predicted hazard rates from the qualification-specific model

ing down. We observe a similar if greater change for male graduates. The hazard of marrying A-level women was by far the most dominant hazard for the 1940 cohort. This has declined over time such that for the 1960 cohort, the hazard of homogamy had overtaken that of marrying A-level women. These are consistent with the cross-cohort change discussed above.

8 Summary

In this paper, we use two hazard models to analyse the life course dynamics of educational homogamy. The results of these two models are quite similar in some respects, but they also differ from each other in significant ways.

For example, both models show that being in education reduces the hazard of marriage of all types, for both sexes. Similarly, both models reveal a curvilinear pattern in the cohort trend in the hazards, and that all hazards of marriage are lower for the 1960 cohort than for the 1940 cohort. The general shape of age dependence is also found to be roughly comparable across models.

However, the qualification-specific model reveals much more variation in hazard by destination than does the BTD model. Under the qualification-specific model, for example, we see as much cross-cohort change in the rank order of the competing hazards facing male graduates as those for female graduates. Also, the qualification-specific model shows that by far the largest increase in the level of homogamy is found among the graduates.

Under the qualification-specific model, we see a positive qualification gradient for several destinations. We have noted the asymmetry of where this gradient applies for male and female respondents. We also noted some interesting cross-cohort change in this regard for marrying A-level men or women. Our interpretation is that this is related to the dynamic of spouse-selection, which in turn is rooted in the changing role of women in society.

The changing educational distribution by sex presents us with a problem of opportunity structure. We try to control for this with a set of simple opportunity structure variables. We think our effort is a partial success, but further work is needed here.

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