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DEMOGRAPHIC DECOMPOSITION OF THE MARRIAGE MARKET IN ENGLAND AND WALES 1911–1991

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Summary. A method for decomposing partner availability into its demographic components (preferences, previous birth trends, migration and mortality, and structure by marital status) is presented and applied to marriage market estimates for selected census years 1911–91 in England and Wales. Preferences are a key component at the youngest ages. The role of other factors varies by age and time period. Contrary to widespread assumption, variation in cohort sizes resulting from past fertility trends is not the dominant contributor to partner availability during this period. Mortality and migration effects tend to be larger than the effect of birth trends and the two marital status components are generally the largest in size. Determinants of intercensal change are similar to the cross-sectional picture. Reasons for the modest contribution of trends in annual births are discussed. Cohort effects on partner supply are not necessarily absent but could arise through a number of mechanisms.

Introduction

While the determinants of the size and age structure of populations as a whole are very well understood, the determinants of the state of the marriage market are less well documented. An appreciation of the structure of the marriage market is hampered by two factors. First, little attention has been given to developing defensible measures of the supply of partners, a situation that remains substantially unchanged since Goldman *et al.* (1984) commented on the fact over two decades ago. In addition, few empirical attempts have been made to evaluate the demographic causes of variation in the marriage market. While the role of mortality and migration in determining the relative supply of partners at the national level is usually recognized along with fertility in the specialized literature on the marriage market, the role of previous marriage trends *per se* is less often acknowledged. Also, more general commentary sometimes assumes fertility trends to be a key determinant of marriage market conditions, though there are some exceptions (e.g. Coleman & Salt, 1992, Chapter 5). The assumption is sufficiently embedded in present day demography that it is encountered in at least one recent demography textbook (Yaukey & Anderton, 2001). However, is this really the case? How important are fertility trends and each of the other demographic factors in determining marriage market conditions at national level? The question is of substantial demographic interest in relation to marriage squeeze and to the consequences of rapid fertility declines combined with high sex ratios in China and other countries of east Asia (Tuljapurkar *et al*., 1995; Das Gupta & Li, 1999). This question cannot be answered without a method for evaluating the contribution of the various factors to the relative supply of partners. This paper presents such a method, and applies it for illustrative purposes to 20th century marriage markets in England and Wales.

The method to be presented decomposes measures of the marriage market – average partner availability by age and sex – into their demographic components. The marriage market can be thought of in a narrower or broader sense – either in pure demographic terms, with age as the sole matching criterion, or in a broader sense, with matching not only on age but also on social and economic characteristics. The paper measures partner availability in pure demographic terms – via numbers by age and sex, in combination with age preferences. This choice follows Dixon's (1978) classification of factors influencing the level and timing of marriage: the feasibility of marriage, its attractiveness and the availability of partners. It is useful to distinguish and measure potential marriage opportunities and constraints that result purely from demographic conditions – age–sex structure – from any that may be superimposed on these by economic conditions or by social customs such as assortative mating and endogamy. Marriage rates have long been known to be linked to economic prosperity (Yule, 1906; Glass, 1938; Galbraith & Thomas, 1941; Oppenheimer, 1994, 2000) and so marriageability in economic terms – whether a person is employed or has satisfactory income and prospects – will in many cases reflect transient local or national economic conditions. It seems more appropriate that any constraints on, or opportunities for marriage that result from economic and social factors should be identified separately from those due to age–sex structure. A second reason for the choice is that matching in the marriage market depends on preferences, and although age preference data are available, no solid information exists on preferences for such partner attributes as educational level, social class and religion, and especially on how flexible these are. Goldman *et al*. (1984) show that the precise matching rule adopted in relation to educational attainment strongly influences the resulting availability indicators and so, in the absence of explicit data on preferences for potential partners' social characteristics, estimates of partner supply would be heavily influenced by the assumptions made.

Two key points are emphasized. As noted above, demography does not currently have an empirically well-tested, conventionally accepted measure of conditions in the marriage market. A measure must therefore be chosen from among those available. But the decomposition method to be presented does not depend on the choice of measure – it can be used with any measure of partner supply. While the results of the decomposition are, as will be seen, to some extent measure-specific, the present paper does not attempt to assess the sensitivity of the decomposition results to the specification of the measure. This is not only for reasons of economy but also because the sensitivity of differing measures of the marriage market to variations in population age–sex structure is itself little understood, and little investigated (but see Lampard, 1993). A second important point is that the England and Wales case is illustrative, and the detailed findings cannot be generalized beyond that setting. Just as the factors giving rise to population change will vary with time and place, so also the relative size of the components contributing to marriage market variations can be expected to differ according to context. Nevertheless, some sharp points emerge from the analysis, as will be seen.

Measurement and data

Iterated availability ratio

The measure of partner supply used here is the most sophisticated currently available in the literature – a weighted version of Lampard's (1993) Iterated Availability Ratio (IAR), a development of the Availability Ratio originally proposed by Goldman *et al.* (1984). It measures the average number of potential partners available by age, taking account of age suitability and the competition within each sex for partners of the opposite sex. A particularly advantageous feature of the IAR is that, unlike the Goldman *et al.* measure, it sums to the total number of unmarried persons of each sex, in the age range covered. Preference data described in a later section are used to calculate the proportions of women of age i who would accept a man aged j, a_{ij}, and the corresponding proportions, β_{ji}, of men aged j who would accept a woman aged i. The product of these, λ_{ij}, representing the joint preference of women aged i and men aged j, is used as a set of age suitability weights in calculating the IAR. The IAR for woman i is specified as:

$$\mathrm{IAR}_i = \sum_j \frac{\lambda_{ij}/\mathrm{IAR}_i}{\sum_k \lambda_{kj}/\mathrm{IAR}_k}$$

where the outermost summation extends over all males in the marriage market, and the summation in the denominator term extends over female participants of all ages (k indexing women in the marriage market). Since IAR_i appears on both the right- and left-hand sides of the equation, the solution is arrived at iteratively. The expression above is in terms of individuals i and j, but since age is the only matching factor used here, all persons of a given age and sex have the same IAR. For computational purposes the formula can be expressed as follows, where I indexes female age groups and J male age groups:

$$\mathrm{IAR}_I = \sum_J \frac{\lambda_{IJ}/\mathrm{IAR}_I}{\sum_K (\lambda_{KJ}/\mathrm{IAR}_K)\ W_K} M_J\,.$$

The expression for males is analogous.

The element of the IAR relating to woman i and man j can be understood as attributing to woman i her 'share' of man j, her share being represented by her

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demand for man j (the numerator) as a fraction of the total demand for man j on the part of all women in the marriage market (the denominator). Woman i's demand for man j is proportional to their mutual interest (λ_{ij}) and inversely proportional to the number of her potential alternative partners. The total demand for man j (the denominator) is the sum of the demand for man j over all women in the marriage market. Woman i's overall partner supply is then the sum over all men in the marriage market of her (fractional) share of each of them (Lampard, 1993). Note that woman i's demand for man j is a function not only of her preference for man j and her potential alternatives but also of his preference for her, since the weighting factor, λ_{ij}, is a product of their preferences for each other. Thus the demand factor reflects the requirement in matching for some mutual interest: where at least one of α_{ij} and β_{ji} is zero then λ_{ij} is zero and hence the potential match contributes nothing to either i's or j's partner supply. The age range for which estimates are made for each sex is 17–60. The male IAR is specified correspondingly.

Age–sex distributions

Age–sex distributions by marital status and single year of age were obtained from the England and Wales decennial censuses of 1911–81, with the exception of 1941 when there was no census. Mid-year population estimates for England and Wales 1991 were used in preference to census data because of a sizeable undercount at ages 15–35 that was substantially larger among men than women (Heady *et al.*, 1994). At earlier censuses, the count was relatively accurate, and so no adjustments have been applied. Estimates of partner supply are obtained only for those unmarried by legal marital status, i.e. those declaring themselves to be single, widowed or divorced. Up to 1971, this is likely to be an accurate representation of the marriage pool. Beyond 1971, the rise of informal cohabitation makes the picture less clear, since those of the unmarried who are cohabiting have found a partner and so are in principle seeking partners less actively than those who are unpartnered. The effect of cohabitation has been examined using the Labour Force Survey (LFS) of 1991; availability ratios estimated from the LFS have been found to be fairly insensitive to the specification of availables as either all the legally unmarried or the legally unmarried who are not cohabiting (for further details see Ní Bhrolcháin & Sigle-Rushton, 2005). It is arbitrary whether the definition of the pool of available partners is based on all the legally unmarried or confined to that subset who are not currently cohabiting. Beyond the clear advantage that the unmarried are in a position to marry formally, several arguments favour the use of the unmarried as a population base in the present case. Cohabiters are more active in the partner market than are married people since informal relationships are less stable than formal marriages (Prinz, 1995; Haskey, 1999; Ermisch & Francesconi, 2000) and cohabiters have more sexual partners than do married people (Johnson *et al.*, 2001). Also, unmarried cohabiters of each sex offset each other to some extent. Consistency across the historical period covered in this paper is a further advantage. There is, importantly, sizeable disagreement between British data sources in the measured levels of cohabitation (Murphy, 2000a, 2000b), thus adding further uncertainty. Finally, a comprehensive assessment of the sensitivity of marriage market estimates to varying definitions of those available as partners is beyond the scope of the present paper, whose purpose is to present a method of decomposition, rather than to examine the measurement of partner supply *per se*.

Annual births

Annual totals of births for the years 1841–1974 (cohorts reaching age 70 in 1911 and age 17 in 1991) are used in the decomposition and were obtained from Table 3·1 of Office of Population Censuses and Surveys (1987). The sex ratio at birth was assumed throughout to be 1·05.

Age preferences

Any estimate of the relative numbers of men and women available for marriage must, for realism, incorporate an allowance for age preferences. It has long been known that the pairing of ages in marriage is not random but displays a preference for matches in which the groom is somewhat older than the bride, with the average gap being fairly narrow in England and Wales and in other West European countries. Previous studies of the marriage market in the aggregate have had to rely on assumption for the preference component of their marriage market estimates. Assumptions range from an arbitrary fixed preference of two or three years to paired age ranges derived by a variety of methods based on observed marriages (Akers, 1967; Henry, 1969; Hirschman & Matras, 1971; Muhsam, 1974; Schoen, 1983; Goldman *et al.*, 1984; Lampard, 1993). All of the assumptions and procedures adopted hitherto are known to be problematic and are usually acknowledged to be so. Assuming a fixed gap of two or three years in favour of the groom is unrealistic both because age differences vary substantially by age at marriage and because the variance in age differences is substantial (Levy & Sardon, 1982; Goldman *et al.*, 1984; Casterline *et al.*, 1986; Oppenheimer, 1988). Besides, the majority of marriages do not feature age gaps of two or three years: in England and Wales the proportion of marriages in which the groom was two or three years older than the bride is well below 50%, whether evaluated on a period or cohort basis (Ní Bhrolcháin, 2001, 2005). Deriving age preferences from the age differences observed in marriages of a given period is known to be an unsatisfactory solution because these result from a mixture of preferences and the opportunities (i.e. age–sex structure) available at a particular time (see e.g. Hirschman & Matras, 1971; Schoen, 1977; Goldman *et al.*, 1984). In the approach adopted by Qian & Preston (1993) age preferences do not need to be specified explicitly.

For these reasons, direct evidence on age preferences is highly desirable and can be expected to improve our knowledge in this area. Ideally, age preferences from a general population sample are required but to date preference information from a representative sample sufficiently large to be usable for demographic purposes has not been available, though explicit age preference information has been collected for small specialized samples of, for example, students and newspaper advertisers (see Kenrick & Keefe, 1992; Buss, 1994). The present study uses data obtained from a UK dating agency on the partner age preferences of 32,326 clients aged 18–60+ in 1996. Clients were asked to state the minimum and maximum age of partner that they would be prepared to accept. From these data, the proportions of men and women were obtained by single year of age 18–59 and 60+ who would accept a partner of each single year of age 18–59 and 60+. These were used as weights in calculating the iterated availability ratio. Some attribution was necessary to extend the preference data beyond that directly measured; details are given in the Appendix. Additional information together with evidence of a close correspondence between mean preferred and observed age differences by age and sex in England and Wales 1991 are given in Ní Bhrolcháin & Sigle-Rushton (2005). Note that the preference function by age and sex is not uniform. That is, although each individual respondent was asked to specify an acceptable age range but not to indicate relative preferences within that range, the aggregation of individuals' preferences across the sample results in preferences expressed as proportions. That is, the preferences of respondents of a given age and sex take the form of an asymmetric curve of the proportion who would accept a partner of each age within a given age range. The curve tapers to zero at each end and reaches a maximum at age gaps where the man is a few years older than the woman, with the most preferred age difference, and the overall shape of the curve, varying by age for each sex (for examples, see Fig. 8 of Ní Bhrolcháin, 2001). Thus, the weighting scheme employed here differs from those used by Goldman *et al.* (1984) and by Lampard (1993), in which all ages within the range of partner ages considered suitable for a person of a given age were regarded as equally acceptable.

Since the dating agency clients are not a probability sample, the extent to which their preferences represent those of the general population of unmarrieds is unknown. The validity of the dating agency preferences is investigated in detail in Ní Bhrolcháin (2004); validation is of necessity indirect since there is no external source against which the preferences can be evaluated. This study establishes both that the Dateline preferences can be considered reasonably representative of the preferences of the general British population in the 1990s and that they are superior, as preference weights, to any available alternatives in making marriage market estimates throughout the twentieth century in Britain. The central points are as follows:

(1) The profile of differentials by age in mean preferred age gaps corresponds well with actual age gaps in England and Wales, declining with female and rising with male age (see Ní Bhrolcháin & Sigle-Rushton, 2005: Fig. 1). The correlations between the mean preferred age differences and those observed in 1991 marriages are 1·0 (male) and 0·7 (female), and the figures are 0·99 and 0·4 in 1951 (corresponding data are not available for 1911). While the relationship between preferred and observed age differences is linear in the male case, there is substantial curvilinearity in the case of women, and a correspondingly lower correlation coefficient.

(2) Expected age differences based on the dating agency preferences provide a closer fit to observed age differences (evaluated by the root mean square error) than either fixed preferences of two or three years or preferences derived from observed age differences at each date by the method devised by Goldman *et al.* (1984) in most comparisons at census dates from 1921 to 1991 (the comparison cannot be carried out for 1911).

(3) Relative gender differentials in the IAR, to which the age preferences are a key input, correspond fairly well with gender differentials in marriage rates by age at census dates 1911–91, though the correspondence is less good in 1961 and 1971.

(4) A very distinctive feature of the comparative preferences of men and women – that women are less interested in younger men than men are in slightly older women – reproduces the findings of a survey of French couples carried out in 1984, the only existing demographic account of explicit age preferences based on a general population sample (Bozon, 1991).

Note that only the first of these points relies on comparisons of mean values, the other three depending on the profiles of age preferences by age and sex. The dating agency preferences are, in all, the most realistic currently available for England and Wales. Since the objective of the present paper is to present a method of decomposition, that is all that is required. The method can be applied with any set of preference weights and does not rely on those used here. Note that if a fixed preference is assumed at all ages – that men and women of all ages will accept only matches in which the man is, say, 3 years older than the woman – the weighted IAR reduces to the conventional sex ratio with the corresponding gap.

Age preferences are assumed in this paper to be constant through time, as in previous studies of time trends in the marriage market over shorter or longer periods (Akers, 1967; Henry, 1969; Hirschman & Matras, 1971; Heer & Grossbard-Schechtman, 1981; Schoen, 1983; Goldman *et al.*, 1984). Given that age preference information is so scanty, this is the only realistic assumption that can be made. The use of a single set of preferences across time can be seen as a standardizing procedure, allowing change in those aspects of population structure by age, sex, and marital status that are relevant to the marriage market to be assessed net of preferences. In addition, detailed evidence supporting the assumption is set out in Ní Bhrolcháin (2004). The key points are these:

(1) If preferences varied substantially through time, we would expect observed age differences to do so also; however, the mean age difference fluctuated within a fairly narrow range in England and Wales through the twentieth century. Nevertheless, the standard deviation of the age difference has been rising in recent decades (from 4·7 in 1969 to 6·3 in 2001) and this may, though need not, reflect changing preferences.

(2) If preferences were changing through the twentieth century, it is likely that there would have been a unidirectional trend towards, for example, smaller age differences reflecting more companionate marriages. However, the time series of mean age differences in England and Wales displays no secular trend – rather it fluctuates and the fluctuations appear to be associated largely with the age structure of available partners, and also with the distribution of marriage ages (Ní Bhrolcháin, 2001).

(3) Expected age differences earlier in the century can be predicted reasonably well from age structure together with the 1996 dating agency preferences.

(4) In 1921, just after the start of the period analysed here, the profile by age of relative gender differentials in the IAR, which has preferences as a key input, is similar to that of gender differentials in marriage rates though the agreement in the absolute level of these two measures is not as close as later in the century.

Finally, the causes of non-random assortment of brides and grooms by age are not well understood. It is generally assumed to be due to age preferences, but it (and they) may have other origins such as enduring age differentials between men and women in psychological maturation (Cohn, 1991), the continuing importance of male earning power in securing a stable base for family formation and social processes that

Table 1. Decomposition of the availability ratio for the unmarried: specification of components

Component	Expression
Age preferences	$IAR_r - 1$
Overall population structure	$IAR_u - IAR_r$
Age–sex structure	$IAR_a - IAR_r$
Sex ratio at birth	$IAR_{rsr} - IAR_r$
Birth trends	$IAR_b - IAR_{rsr}$
Mortality and migration	$IAR_a - IAR_b$
Structure by marital status	$IAR_u - IAR_a$
Never married	$IAR_n - IAR_a$
Previously married	$IAR_u - IAR_n$

The population base for each IAR is as follows:
IAR_u, the unmarried; IAR_a, all marital statuses; IAR_n, the never married; IAR_r, rectangular population with equal numbers of men and women at each age; IAR_{rsr}, rectangular population with 105 men per 100 women at each age; IAR_b, the original birth cohorts of men and women of each age at each census year t; that is men aged $x = 1{\cdot}05 \times B_{t-x}/2{\cdot}05$ and women aged $x = B_{t-x}/2{\cdot}05$, where B_{t-x} = births in the calendar year $t - x$.

structure the encounters of men and women by age (cf. the marriage circles of Henry, 1972). With increases in educational enrolment on the part of both sexes, and increasing labour force participation by women, the structure of social interaction by age and sex may have been altering, but solid information is lacking on the subject. To assume constant preferences does not, of course, imply that the actual pattern of age matching in marriages has been constant through time. Observed distributions of age differences can alter in response to the age distribution of available partners, as can mean age gaps, even though underlying preferences remain the same.

Methods

The method described here separates the Iterated Availability Ratio into its demographic components and thus enables the demographic determinants of partner supply to be quantified. The estimate of partner availability, the IAR, has two inputs: the age preference weights and the age–sex distribution of the unmarried population. The age–sex distribution of the unmarried at a point in time results, in turn, from past trends in annual birth numbers combined with cumulative sex differentials in previous mortality, in net migration and in the proportions unmarried. The contribution of these components to the IAR can be quantified, with the mortality and net migration components being obtained jointly as a residual.

The method proceeds by calculating age- and sex-specific IARs for a series of population bases, specified so as to separate out the various components. The decomposition is additive, the components being obtained by subtraction. For clarity, subscripts for age and sex are omitted in this exposition, and are implicit throughout.

Let IAR_u denote the IAR for the unmarried (overall availability) and IAR_r the availability ratio for a rectangular population having equal numbers of men and women in each age group. IAR_r gives us the component of the overall availability ratio, IAR_u, that is attributable purely to age preferences combined with a population structure that incorporates one potential partner per person. $IAR_r - 1$ then represents the pure effect of preferences, i.e. the extent to which preferences act to alter the 1:1 balance present in a rectangular population. The difference, $IAR_u - IAR_r$ is the total effect on the availability ratio of population structure (everything except preferences). The IAR estimated for people of all marital statuses, IAR_a, allows the overall population structure effect ($IAR_u - IAR_r$) to be separated into a component due to age–sex structure ($IAR_a - IAR_r$) and one due to structure by marital status ($IAR_u - IAR_a$). Further components are obtained by calculating IARs for a population with a rectangular age structure but 1·05 men per woman in each age group (IAR_{rsr}), a population comprising at each age x in census year t the number of births occurring in year $t - x$ distributed by sex by assuming a sex ratio at birth of 1·05 (IAR_b), and finally the population of those never married (IAR_n). The component due to fertility ($IAR_b - IAR_r$) can be separated into that attributable to the sex ratio at birth ($IAR_{rsr} - IAR_r$), and one due to trends in birth numbers *per se* ($IAR_b - IAR_{rsr}$). (While there have been secular trends in the sex ratio at birth in England and Wales, with a decline through the nineteenth century, followed by a rise to the 1940s, and a further easing beginning in the 1970s, these will not have had a substantial effect on partner supply.) The full details are set out in Table 1. Note that the effect of marital status structure reflects mortality to some degree, since at older ages the higher proportions of women than men widowed are a sizeable element of this component. Apart from higher male mortality, two further factors contribute to the higher proportion of women than men widowed: the higher remarriage rates of previously married men and the tendency for women to marry on average a somewhat older man. (For example, in 1991 in England and Wales the remarriage rate of previously married men was one-third higher than that of women in the age group 35–44, 68% higher at ages 45–54 and six times as high at ages 55+: Office of Population Censuses and Surveys, 1993, Table 3·3.) With an average age difference of around three years, the (former) husband of a married woman of a given age will be six years older on average than the (former) wife of a man of the same age: hence the greater risk of widowhood of a married woman of a given age compared with a married man of the same age comprises both the cumulative sex differential in mortality to that age and an average six additional years of cumulative sex differentials in mortality resulting from the disparities in their partners' ages.

Results

England and Wales 1911, 1951 and 1991

While the decomposition has been applied to all census years 1911–91, for economy the results are illustrated by three selected years: 1911, 1951 and 1991. The contribution of age preferences and of overall population structure to the age- and sex-specific IARs for England and Wales for these years is shown in Fig. 1 (the age

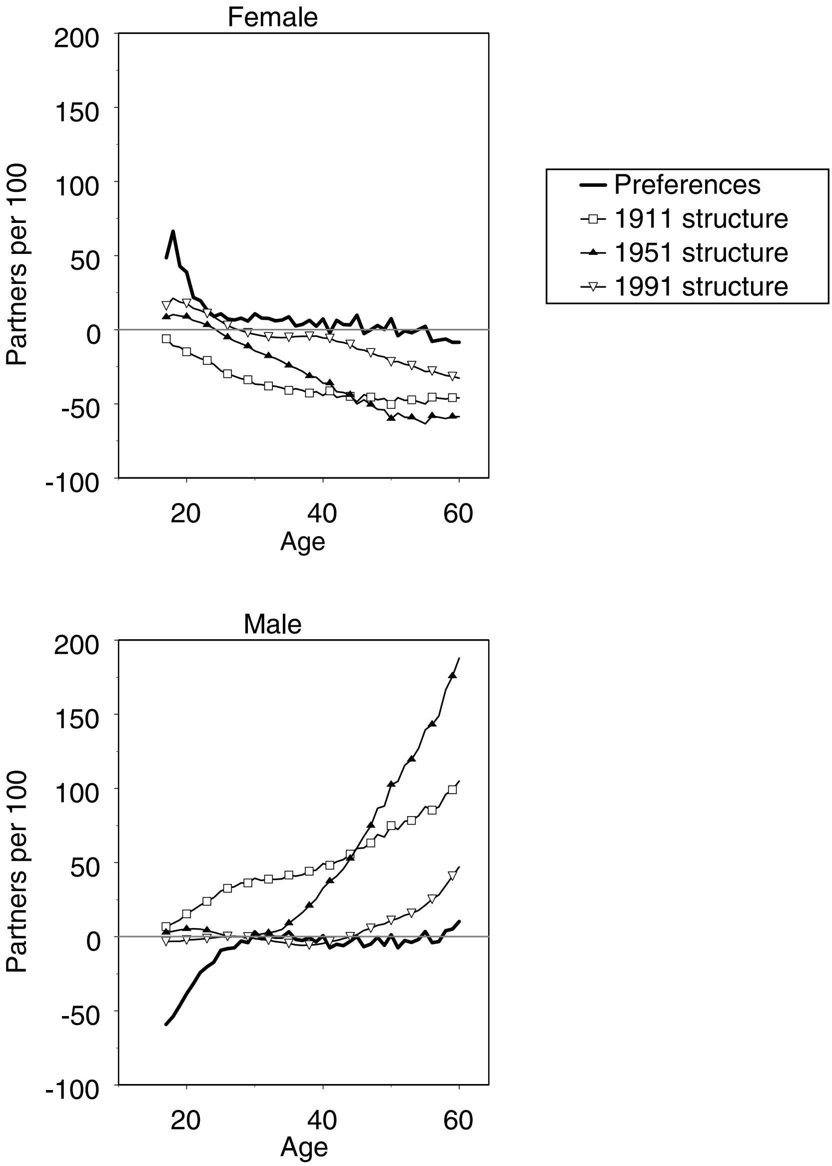

Fig. 1. Components of the availability ratio due to preferences and to the age–sex structure of the unmarried, by age and sex, England and Wales, 1911, 1951, 1991.

preferences effect is, by assumption, constant through time). At the youngest ages, the largest component tends to be that due to age preferences, which are mainly responsible for the age pattern of IARs at ages under about 30. From the early to mid-30s the preference component is approximately balanced as between the sexes, and structural factors increase in importance in generating the characteristic decline in female and improvement in male partner supply with rising age. The effect of

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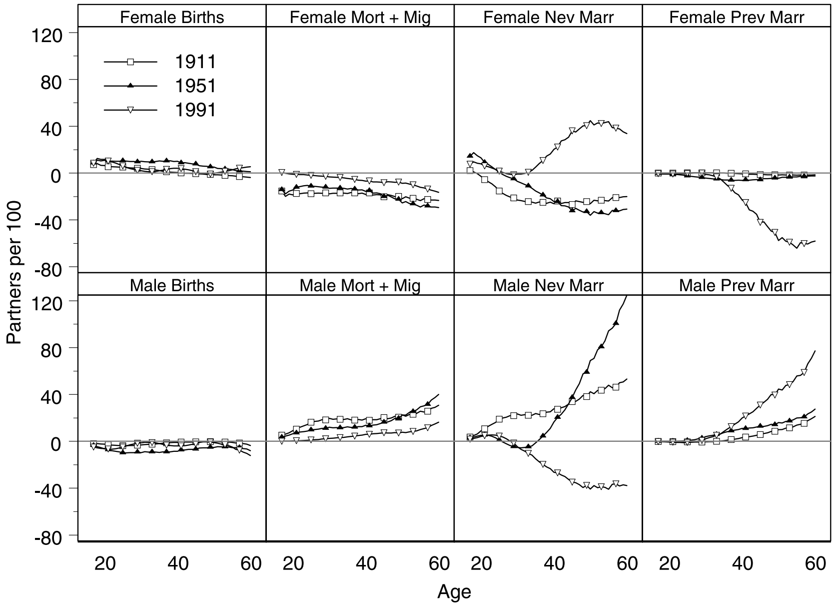

Fig. 2. Components of the availability ratio by age and sex, England and Wales 1911, 1951, 1991.

overall population structure with rising age tends to be negative for women and positive for men. The effect of population structure is fairly strong throughout the age range in 1911, favouring men at all ages. In 1951, men in their 50s are particularly advantaged by the structural effect, for reasons that will be discussed presently.

Figure 2 separates the overall effect of population structure into its components: birth trends (total effect including the sex ratio at birth), migration and mortality, and marital status (proportions never married and previously married). The effect of the sex ratio at birth (constant by assumption across all three dates) has not been plotted. It has, as would be expected, a small positive effect of 5–8 per 100 for women and ranges between −2 and −5 per 100 for men (throughout, effects are expressed per 100 – that is the number of potential partners available per 100 persons). The effect of the sex ratio at birth varies by age, since it interacts with the structure of age preferences, which is not constant by age.

It can be seen that the total effect of births is small relative to other factors at these three dates; across all ages at the eight census years the minimum value for women is −1 and the maximum +13 and for men the range is −10 to +1 per 100. When the effect of the sex ratio at birth is omitted, thus confining the effect to the pure influence of time trends in fertility, these ranges are narrowed further to −6 to +7 per 100 (female) and −5 to +5 (male). Nevertheless, the overall fertility effect can, in some instances, be non-negligible, benefiting women by an additional 9–10 potential partners per 100 at ages under 43 in 1951, and by an additional 9–13 partners per 100 at ages 18–23 in 1991, with in each case a comparable deficit in male partner supply due to births.

But it is clear that the births component tends to be the smallest of all the structural components, a finding that contrasts sharply with the emphasis in some sections of the demographic literature on fertility trends as a prime determinant of partner supply. Across all eight census years for which the decomposition has been carried out, the births component of female partner supply (including the sex ratio at birth) is the largest of the four age-specific structural components in absolute terms only in 1991, at ages under 29. Excluding the sex ratio at birth, this occurs only at ages 23 in 1981 and at ages 29–33 in 1991. Among men, fertility is of somewhat greater importance, with the overall births effect being the largest single structural component, in an absolute sense, at ages under 29 in 1991 and also at most ages under 24 in 1951 and 1971. Again, this effect owes much to the sex ratio at birth rather than reflecting fertility trends *per se*; excluding this, it features much less frequently. Birth trends are of greater importance, in relative terms, at younger than at older ages. Across the eight census years covered here, the births effect is never, age-specifically, the largest in absolute terms at ages over 33. In all, the impact of birth trends on cross-sectional partner supply is not negligible but is probably smaller in relative terms than many demographers would have supposed *a priori*.

The combined effects of mortality and migration are, as is seen in the second column of Fig. 2, somewhat more substantial than the fertility component. The mortality and migration component declines across the century. This is what would be expected both from the decline in absolute sex differentials in mortality across the century, particularly at younger ages, and from the gradual elimination from the age–sex structure of the effects of the heavy excess male emigration in the late nineteenth century and early decades of the twentieth century (Carrier & Jeffery, 1953; Charlton, 1997; Dobson & McLaughlan, 2001). In combination these sources accounted for a loss of between 15 and 23 potential partners per 100 women across female ages in 1911 compared with a range across female ages of − 16 to +1 potential partners in 1991. Male marriage market gains from migration and mortality are in the range 5 to 31 per 100 in 1911 and 0 to 16 per 100 in 1991. Male gains and female losses from this factor rise with age, as would be expected from cumulative sex differentials in mortality, and are particularly pronounced at the top of the age range considered in 1951, reflecting both war losses and excess male migration in the early decades of the century (Carrier & Jeffery, 1953; Winter, 1985).

The effect of structure by marital status operates primarily through sex ratios by marital status. At a given date, these result essentially from the preceding history of age- and sex-specific marriage, marital dissolution and remarriage rates in interaction with overall age–sex structure. The third and fourth columns of Fig. 2 show that both marital structure components can be very substantial indeed. Their impact can be particularly large at older ages, but may be sizeable at younger ages also. In 1911, the impact of sex differentials in the proportions single and sex ratios among the single results in a loss of between 3 and 21 potential partners per 100 women in their 20s, and in gains of between 9 and 22 partners per 100 men of this age. In 1951, non-marriage had a much smaller effect among the young, but had a very substantial effect indeed on partner supply at ages 45+, particularly among men. It is this that explains most of the exceptionally large structural effect at older male ages in 1951. Sex ratios among single people in their early 40s and older were lower, and the excess

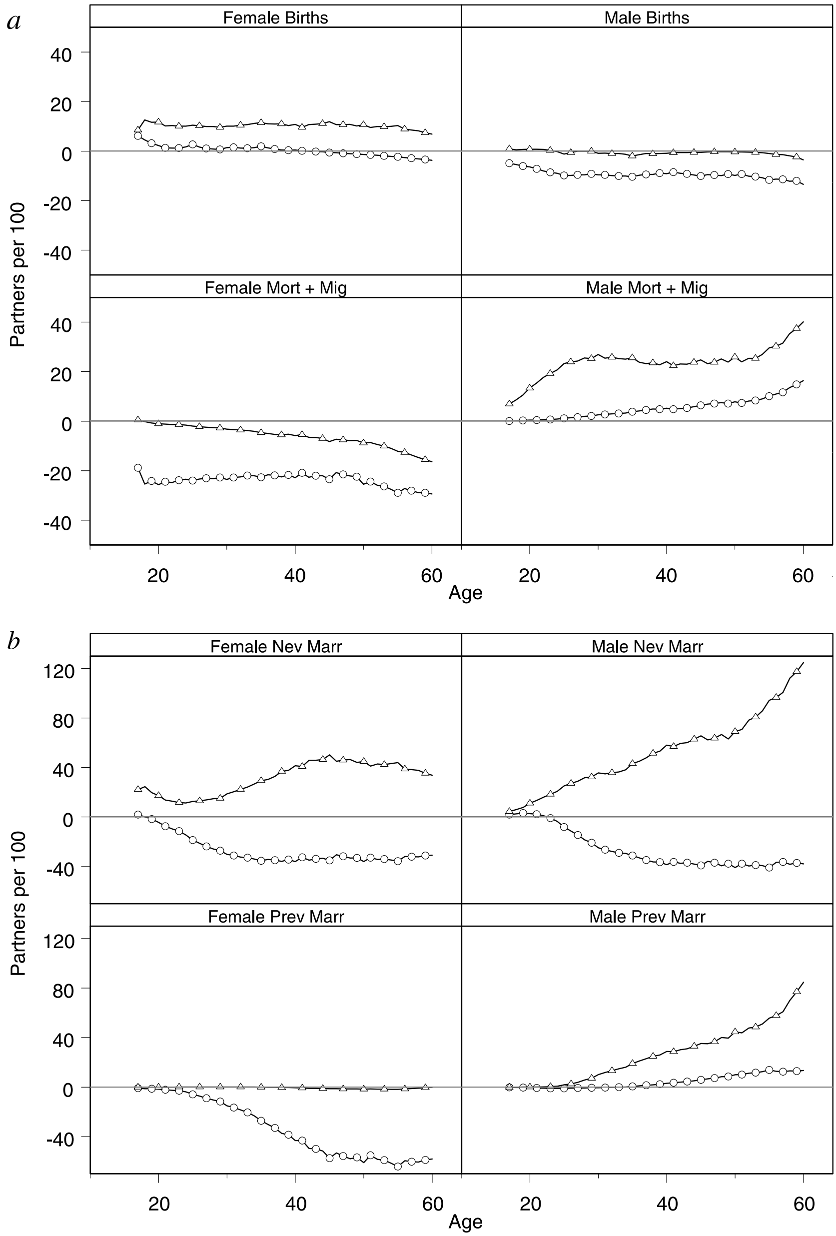

Fig. 3. *a.* Minimum and maximum values of the age-structure components of the availability ratio, by age and sex, England and Wales 1911–1991. *b.* Minimum and maximum values of the marital status components of the availability ratio, by age and sex, England and Wales 1911–91.

female proportions single at ages 41–59 higher in 1951 than at any of the three dates considered here. By 1991, the non-marriage component is mainly to women's

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Table 2. Absolute and percentage change in partner availability attributable to each component, by sex and age group, England and Wales, 1911–91, 1911–51, 1951–91

	Female				Male			
	17–24	25–34	35–49	50–60	17–24	25–34	35–49	50–60
Absolute change in partners per 100[a]								
1911–1991	31	33	35	20	− 19	− 38	− 53	− 62
1911–1951	22	22	4	− 12	− 12	− 35	− 8	53
1951–1991	9	11	31	32	− 6	− 3	− 45	− 115
Percentage change								
1911–1991								
Births	12	− 4	5	30	20	− 2	3	7
Mortality and migration	55	43	34	51	50	43	25	24
Never married	35	65	148	320	28	65	108	138
Previously married	− 2	− 5	− 87	− 301	2	− 6	− 36	− 69
1911–1951								
Births	18	28	61	-44	36	18	− 127	− 9
Mortality and migration	18	23	34	42	28	21	− 76	7
Never married	68	69	45	89	38	72	155	90
Previously married	− 4	− 20	− 40	13	− 2	− 11	148	12
1951–1991								
Births	− 3	− 67	− 22	1	− 15	1328	− 15	0
Mortality and migration	144	90	34	47	103	− 1641	24	16
Never married	− 45	50	161	228	− 2	795	117	115
Previously married	4	26	− 72	− 176	13	− 382	− 25	− 32

[a]Unweighted average of change in availability ratios at single years of age within each age group.

advantage, and is very substantial at ages 40+, with a correspondingly large disadvantage to older men. The effect of previous marriage on female partner supply was not very substantial in 1911 and 1951. Among men of 40+ it is somewhat larger in these years, adding between 3 and 21 potential partners per 100 men in 1911 and between 11 and 28 per 100 in 1951. The previous-marriage effect is much larger for both sexes in 1991, accounting for a shortfall of between 23 and 64 potential partners per 100 women aged 40+ and for gains of between 20 and 77 potential partners per 100 men of this age. The reason for the substantially larger effect in 1991 is that the absolute differences in sex ratios between marital status groups are much wider in 1991 than in the two earlier years. Sex ratios are certainly lower among the unmarried as a whole in earlier years than in 1991, but in 1911 and 1951 they do not differ as much from those of the never-married as they do in 1991. In 1991 the two marital status effects offset each other to some extent, but this is not true in earlier years.

The variation in each of the four structural components across all eight census years 1911 to 1991 is summarized in Fig. 3a and b. Figure 3a plots the minimum and

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maximum values of the births and mortality and migration components and Fig. 3b the corresponding values for the marital status components; for greater clarity, the scale of Fig. 3a is four times as large as that of Fig. 3b. Figure 3 underlines the findings presented thus far for the three selected years. By comparison with other factors, fertility has not been a major component of partner supply for either sex in the period 1911–91, and in size and range of variation it is eclipsed by the mortality and migration factor. Both of these components are well below the two marital status components in the size of their effects as well as in their range of variation, though, as would be expected, the effect of previous marriage at the youngest ages is negligible throughout. Note that these ranges relate to a particular historical period in England and Wales and would not necessarily obtain in other settings or other time periods.

Intercensal change

Beyond the role of each demographic factor in determining partner supply at a point in time, the present section looks at how far each component contributes to change in partner availability through time. The absolute and relative contribution of each factor to change in male and female availability ratios by age is summarized in Table 2 for the period 1911–91 and also for two sub-periods 1911–51 and 1951–91. While marriage markets were changing within these 40-year periods, the division probably gives a reasonable representation of long-run change in British marriage markets during the twentieth century. In the 40 years prior to 1951 British marriage markets were very unfavourable to women and in the four decades following 1951 female marriage markets initially improved substantially (to 1971) and then stabilized. The figures presented are unweighted averages across single years of age of the overall change in the IAR that is accounted for by each factor. The average absolute change for each age group is given in the top bank of figures in Table 2, followed by a bank of percentage change figures for each period. Note that where the percentage given for a factor is negative, the contribution of that factor is opposite in sign to the overall change. For example between 1911 and 1991 average supply at ages 35–49 declined by 53 potential partners per 100 men, made up of an increase of 19 per 100 (-36% of -53 per 100) due to change in sex differentials in previous marriage, and declines of 2, 13 and 57 per 100 due to fertility, mortality + migration and non-marriage (3%, 25% and 108% of -53 per 100, respectively).

Availability ratios for women of all ages increased over the 80 years from 1911 to 1991. The change in estimated partner supply for women over the period as a whole is quite sizeable – an additional 31–35 partners per 100 at ages under 50 and 20 per 100 at older ages. The marriage market position of men deteriorated across the century, particularly at older ages; again, the change is substantial, amounting to a loss of 19 potential partners per 100 at ages under 25, 38 per 100 for men aged 25–34 and as many as 62 per 100 for men aged 50–60. Change is not uniform across the entire period by age. The main improvement for women under 35 occurred by 1951 and that for older women between 1951 and 1991; most of the decline at younger male ages took place during the earlier part of the century and that at older ages in the more recent period. As would be expected from the findings of the preceding section, birth trends are not the predominant cause of change in the marriage market over this period. Fertility accounts for no more than 30% of the average change in any female age group and for no more than 20% among males by age. This is true also of the two sub-periods considered. Birth trends were responsible for no more than 28% of the improvement in the partner supply of women under 35 during 1911–51 and for at most 36% of the decline in male partner supply at these ages. Fertility trends account for less than half of the change observed in 20 of the 24 age–sex groups set out in Table 2. In those cases where fertility is responsible for a third or more of the change occurring, either the absolute change involved is small or the fertility effect is offset by other factors.

A larger share of marriage market change during the 80-year period is attributable to mortality and migration trends than to fertility. Between 34% and 55% of the improvement in women's marriage markets over the century was due to declining sex differentials in mortality and migration, with a corresponding and broadly similar role in the deterioration of male marriage markets. The impact of changing sex differentials in mortality and migration is most in evidence during the second half of the period. The mortality and migration experience in question is, of course, cumulative. Thus the increase between 1951 and 1991 of eleven potential partners per 100 women aged 35–49 (34% of 31) that is due to mortality and migration is attributable to a reduction in the absolute size of sex differentials in mortality and migration from about 1891–1951 to 1941–1991 (Carrier & Jeffery, 1953; Charlton, 1997).

In general, structure by marital status played a larger part in changing marriage markets than did age-structure effects, particularly at ages 25+, over this period in England and Wales. A substantial part of the improvement in the marriage market position of women of all ages over the century is due to higher sex ratios among the single in 1991 than in 1911, though this improvement was offset to some extent by a deterioration due to the changing age–sex structure of the previously married. Essentially, sex ratios are more differentiated by marital status in 1991 than earlier in the century, with particularly high sex ratios among the single at ages 30+.

Further investigation of components

Discussion of the marriage market, and especially of marriage squeeze, often assumes that trends in birth numbers are the key determinant both of the cross-sectional state of the marriage market and of time-trends in partner supply. Since the findings reported here are therefore in some ways unexpected, it is worth examining the components of partner supply in more detail. This section looks first at the fertility component, and goes on to consider the marital status effect.

There are three reasons why the fertility component is small in absolute and relative terms in England and Wales in the period covered. First, fertility trends in England and Wales have not been such as to have a sizeable effect on the relative numbers of the sexes. Although annual birth numbers have been quite variable over the period of interest, the time series of births has not been characterized in general by rises or declines that are either steep enough or sufficiently sustained over lengthy periods to have a major impact. This is seen in Fig. 4, which plots annual birth numbers in England and Wales from 1841 to 1974, the earliest and latest cohorts

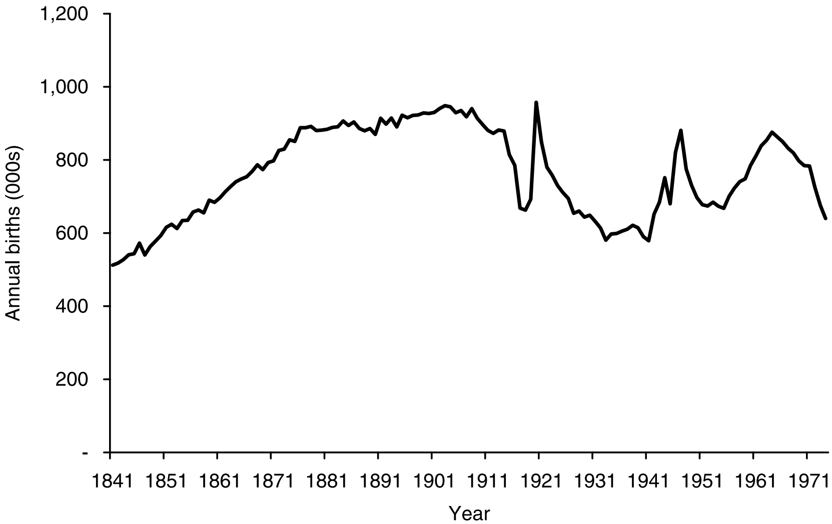

Fig. 4. Annual number of births in England and Wales 1841–1974.

contributing to the estimates of partner supply in the years considered (the birth cohort of 1841 feature as potential partners in 1911 and those born in 1974 are aged 17 in 1991). There are some periods of rapid declines/rises in birth numbers in the series, and these do indeed advantage one or other sex in the marriage market. For example, women born in the mid- to late 1920s were particularly advantaged by the very steep decline in births from 1920 to 1933. The female cohort in the middle of this decline – that of 1926 – was, of all birth generations 1900–59 in England and Wales, the most favoured by birth trends, with a fertility effect of an additional eleven partners per 100. The fertility-based advantage to this female cohort stems not only from the larger male cohorts of the preceding six years but also from the lack of competition from smaller female cohorts up to seven years younger than them. But rapid and sustained growth/decline such as occurred during the 1920s is not typical of the England and Wales births series. The period 1921–27 was in fact the longest for which births were declining continuously at a rate of 2% or more per annum in the 150 years to 1991. There is no period during which births were rising this rapidly for more than three years continuously, and hence no instance in which men were advantaged by fertility trends to the same extent. Although births grew at an annual rate of at least 5% from 1942 to 1944, and in 1946–47, they declined by 10% in 1945. The upward trend in numbers reduced the male disadvantage in partner supply, the male cohort that benefited most being circa 1943, for whom the average fertility effect across all ages is zero – thus, for men of this generation the fertility effect was large enough to offset the marriage market penalty resulting from the sex ratio at birth.

Second, the preferences weights used to weight the availability ratio reflect acceptability of partners across a broad rather than a narrow range of ages, and this has the result of smoothing any large disparities in size between neighbouring cohorts that occur. The general expectation that fertility trends should be a key determinant

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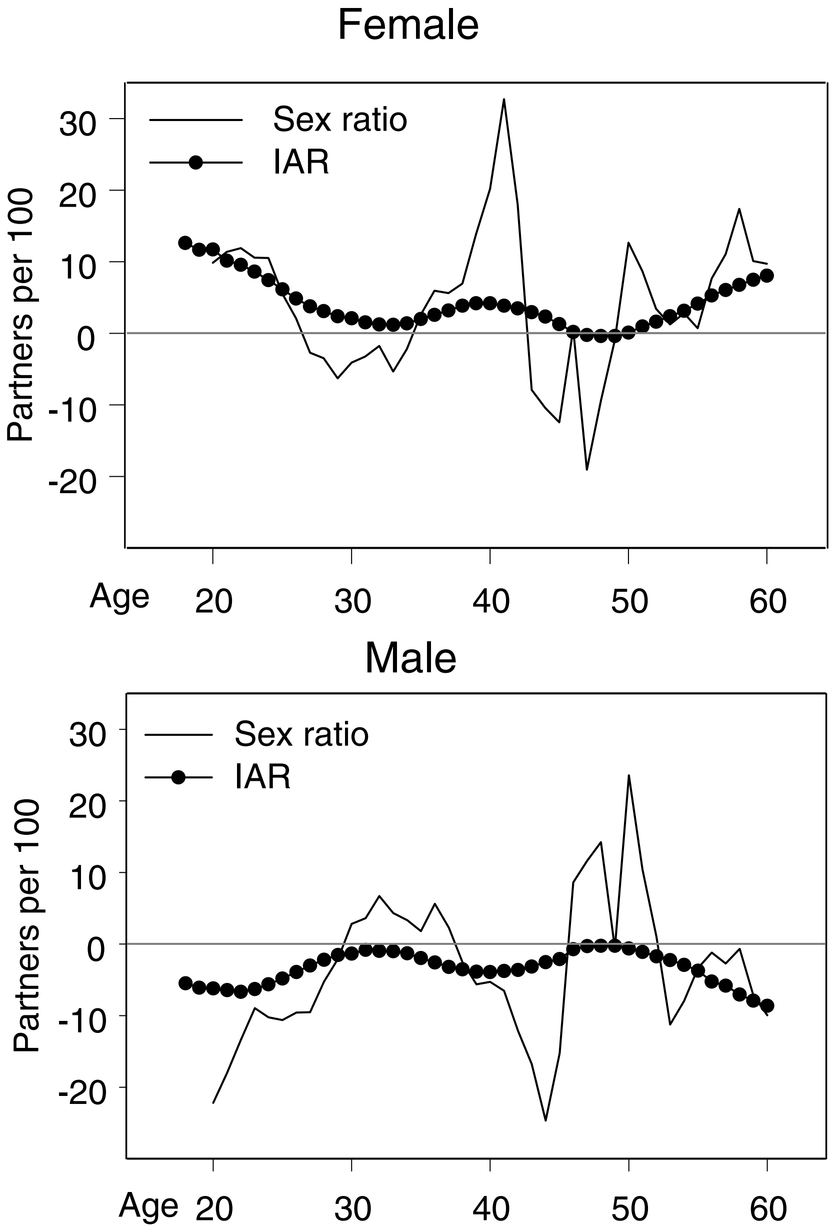

Fig. 5. Estimated effect of birth trends on partner supply based on the IAR with stated preferences versus a 3-year gap (sex ratio), England and Wales, 1991.

of individuals' marriage market position derives in part from the supposition that a small fixed age gap (about 2–3 years) is both generally preferred and typical of the majority of marriages. This two-fold assumption is mistaken. Stated age preferences reveal that there is substantial indifference regarding partner ages within broad age

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ranges (around 5–7 years), and this inherent flexibility in preferences is reflected in an apparent responsiveness of the distribution of age differences to changes in the age distribution of available partners (Ní Bhrolcháin, 2001). Furthermore, as noted earlier, variability in age difference distributions is substantial (Levy & Sardon 1982; Goldman *et al.*, 1984; Casterline *et al.*, 1986; Lampard, 1993). Partner supply estimates are necessarily influenced less by the relative size of surrounding cohorts when the assumption is made of diffuse rather than fixed or narrow age preferences.

This is seen in Fig. 5, which presents contrasting estimates of the fertility component resulting from a decomposition of the IAR based on stated preferences versus a preference for a 3-year age gap (note that the latter reduces to a simple sex ratio with a 3-year gap). Men who were aged 44 in 1991, and thus born in 1947, would have experienced a sizeable fertility-based shortfall of 25 potential partners per 100 when assessing partner supply by the sex ratio with a 3-year gap, since they are paired with the much smaller female cohort of 1950, but the fertility effect using the IAR based on the much less rigid, stated preferences is a reduction of just 3 per 100. Thus, assuming a fixed 3-year gap results in a fertility effect for this male cohort that is eight times that obtained using the dating agency preferences. Similarly, women aged 41 in 1991, and so born in 1950 and paired by the sex ratio measure to men born in 1947, would have experienced a fertility-based boost of 33 potential partners per 100 assuming a 3-year gap, again eight times the fertility effect of just 4 per 100 evaluated via the IAR using stated preferences.

Thus, assuming rigid preferences focused on a single-year age gap gives rise to much larger estimates of the fertility effect in these cases than does the use of the more realistic explicitly stated preferences. Since the sex ratio is simply the IAR with a fixed preference for a 3-year gap, the relatively modest size, relative to standard expectation, of the present study's estimates of the role of fertility trends in determining partner supply reflects in part the diffuseness of the age preference weights used, and is not due to the decomposition method employed. Note also that the births component plotted in Fig. 5 includes the effect of the sex ratio at birth, and that the pure effect of year-on-year variations in birth numbers for both sexes is smaller in absolute size than the total effect by about 0·05.

A third and related point is that fertility trends were responsible for only a minor part of the change through time in the age–sex structure component of the availability ratio in England and Wales 1911–91. Across the eight census years from 1911 to 1991 (there was no census in 1941) most of the variance in the overall availability ratio is accounted for by marital status, though there is non-negligible covariation between it and the age-structure component, particularly for women. This is shown in the first row of Fig. 6 (note that the scale for the female plot is four times that of the male plot). Furthermore, migration and mortality account for by far the largest part of the variance in the age–sex structure component of the availability ratio (second row Fig. 6, scale of both plots five times that of the overall female plot). Several points about these findings are worth noting. First, it seems reasonable to suppose that there may be an interaction between the age-structure and marital status effects. Any tendency for the supply of partners to promote or constrain marriage rates – and evidence on the subject is not clear (Keilman, 1985; Fortier 1988) – would mean that the age-structural and marital status effects would amplify each other. There is some

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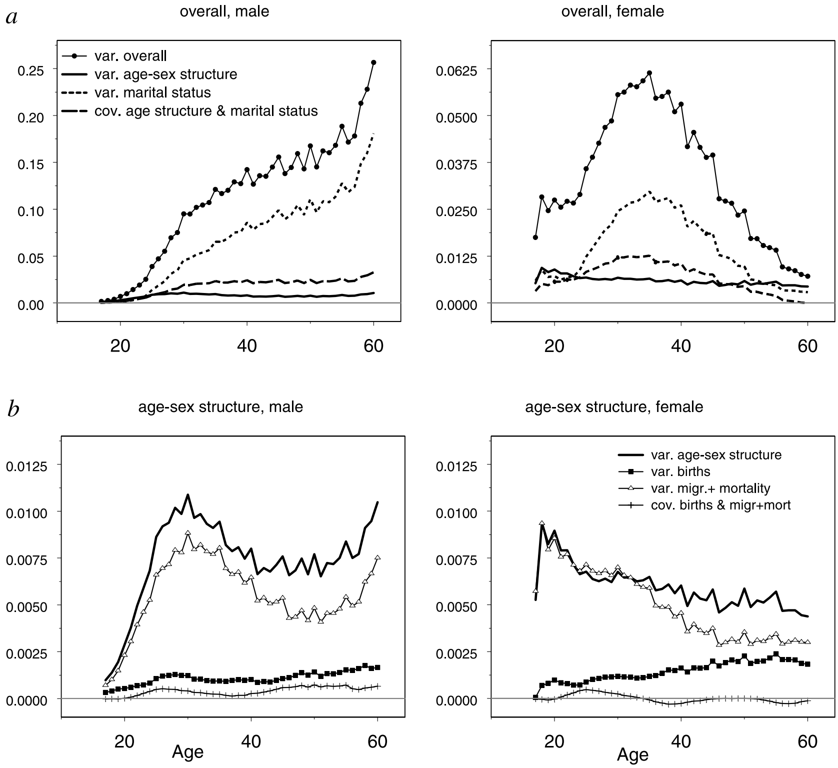

Fig. 6. *a*. Components of the overall variance in the availability ratio by age and sex, England and Wales, 1911–91. Note that the scale of the female plot is four times that of the male plot. *b*. Components of the variance in the age-structure component of the availability ratio by age and sex, England and Wales, 1911–91. Note that the scale is five times that of the female plot in (*a*) above.

evidence of this in the England and Wales case, in that the covariation between the age-structure and marital status components is positive; however, the covariance term makes a relatively small contribution to the overall variance in the availability ratio. Hence, the interaction between these two factors appears not to be sizeable in the present case. A more detailed decomposition that allows interaction effects to be identified directly also leads to the same conclusion: the interaction between age structure and marital status is small relative to the marital status main effect. (This is done by calculating IARs for four further populations beyond those set out in Table 1: a rectangular population to which the proportions unmarried or never married by age and sex in each census year are applied, and populations constructed from the original birth cohorts to which proportions unmarried or never married in each census year have been applied. Using these and the original estimates, interactions can be estimated if it is assumed that the interaction between preferences and demographic factors can be ignored.) A second point is that the estimated contribution of fertility trends to variation in the age-structure component refers to

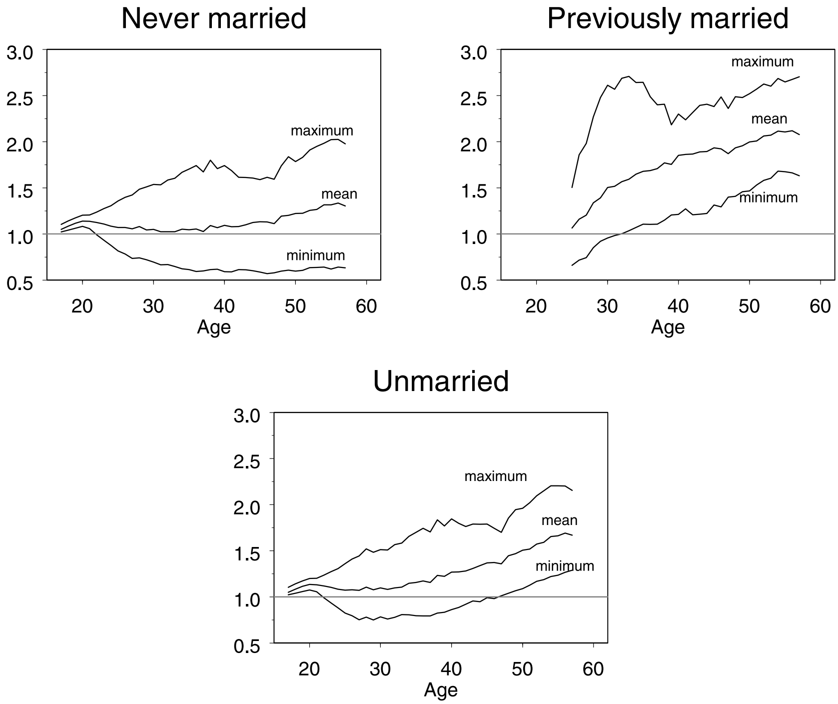

Fig. 7. Sex differentials in the proportions unmarried: ratio of female to male proportions unmarried, never married and previously married with a 3-year age gap, female ages 17–57, minimum, maximum and mean, England and Wales 1911–91.

change in that aspect of age–sex structure that is relevant to the marriage market. Age structure *per se* is of course strongly influenced by fertility trends.

Note that these results are specific to a particular country and historical period and cannot be taken to apply more widely. Findings of a decomposition of the US marriage market 1940–90 (to be reported elsewhere) are similar in broad outline to the England and Wales case but there are some differences also. The marital status effect also accounts for most of the variation through time in overall availability in the US, but the covariance between age structure and marital status is negative at some ages and fertility makes a larger proportionate contribution to the age-structure effect.

The effect of marital status results from past trends in marriage, divorce, widowhood and remarriage. It is not always seen as a key determinant of relative numbers in the marriage market. Its strength as a determinant is due to two related factors. First, estimates of potential marriage partners are based on the unmarried population, and small absolute differences between the sexes in the proportions unmarried/never married/previously married can result in a large proportionate differential, especially at older ages as the proportions unmarried decline. The scale of sex differentials in marital status is seen in Fig. 7, which plots the minimum, maximum and mean ratio of the proportions of women to men 3 years older who

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were never married, previously married or unmarried at each age, over the eight census years 1911–91. (Note that these differentials tend to be larger when based on the same age – the 3-year gap is used to approximate the ages of likely potential partners. The plots for the previously married start at age 25 rather than 17 since the very small proportions of men previously married at the youngest ages give rise to large ratios that are not particularly informative.) Two points are evident from these plots. Relative differentials in proportions unmarried, or in each sub-category, can be very sizeable, in each direction. For example, the proportion never married among 30-year-old women has been as low as 70% of the figure for males 3 years older and as high as 1·5 times the male proportion. At older ages, the disparities are larger, and this is conspicuously true of previous marriage, with the proportion of women in their late 50s who are previously married ranging from 1·6 times to 2·7 times the male proportion. The size of these differentials accounts for the importance of the marital status effect in the decomposition, and their variability through time in England and Wales is the reason why it explains so much of the overall variation in availability through time.

Part of the previous marriage effect results, as noted earlier, from cumulative sex differentials in mortality beyond the age range covered in the present study, though in 1991 the higher proportion of women than men currently divorced was a significant contributor to this component. The higher proportion of women previously married is also due in part to higher male remarriage rates and this in turn may be influenced by more advantageous supply conditions for older men.

Discussion

Particular attention has been given here to the modest size of fertility trends on estimates of partner supply in England and Wales, largely because they are often seen as the key determinant of the relative numbers of the sexes. Clearly the contribution made by fertility as a determinant depends on that of the other components and it is noteworthy that its relative importance at younger ages was greater, particularly for men, later in the series considered here, when absolute sex differentials in mortality and migration had narrowed. But the modest fertility effect does not imply that cohort effects on partner supply are either absent or insignificant, for two reasons. First, although the effect of fertility trends is not sizeable in general, the births effect for individual cohorts can be far from negligible. Second, cohort effects that result from mortality or migration could be substantial, and such effects might arise also from either period- or cohort-based trends in marriage, divorce, widow(er)hood and remarriage. The presence, size and origin of cohort effects on relative numbers in the marriage market need to be established directly via age–period–cohort analysis in any particular context and will almost certainly vary by time and place.

In the light of the decomposition findings, the absence of evidence of either elevated marriage rates or of marriage squeeze among cohorts at troughs and peaks in annual births, reported in Ní Bhrolcháin (2001), is unsurprising, both because such cohorts do not necessarily experience an exceptional fertility-based advantage/disadvantage and because fertility is only one of several determinants. This means that attempts to evaluate the hypothetical impact of the relative numbers of the sexes on marriage rates and on other demographic, social and economic phenomena, including marriage squeeze effects, would be more appropriately based on direct measures of partner supply by age than on the original sizes of birth cohorts as an indicator of demographic marriage market opportunities and constraints.

An implication of the sizeable influence of marital status distributions on partner supply, particularly at older ages, is that a fuller understanding of the marriage market both in cross-section and through time requires investigation of the determinants of time trends in marital status distributions, and of sex differentials in these. In situations where cohabitation is widespread, the origins of the distribution by partnership status would be of interest. As noted earlier, availability estimates based on the unpartnered and on the unmarried are very close, but the contribution of the various demographic components might differ somewhat.

To the author's knowledge, this is the first method to be proposed for estimating empirically the demographic components of partner supply. It is presented as a first step towards quantifying the demographic dynamics of marriage markets and to stimulate further research on the marriage process at the aggregate level. At minimum, the method and the illustrative findings reported here raise questions about whether the dynamics of marriage markets are altogether as is often assumed in the demographic literature. The caveat must be mentioned that the decomposition results reported are based on a particular measure of partner supply. Results may vary according to the indicator used and to the configuration of the preference weights. It has been seen that there was a particularly large disparity between the fertility effects estimated assuming a fixed age difference compared with more diffuse preferences. The components of partner supply may thus be measure-specific. Uncertainty in this respect stems primarily from the relative paucity of research on measures of the marriage market rather than from the properties of any particular decomposition method. However, it would be surprising if this dependence was substantial and further research would be required to evaluate its extent. Finally results will, of course, be driven by the structure of each population and by the history of its demographic flows. Just as the relative roles of fertility, mortality and migration in determining population change vary with time and place, so also the relative roles of the demographic determinants of partner supply can be expected to vary in different geographical and historical contexts.

Appendix

Further details of age preference data

Age preference data were available only for ages 18–60+ but since some men specified 17 as an acceptable partner age, weights were attributed to 17-year-olds of each sex by setting them equal to those of 18-years-olds, lagged by one year of age. The preference data were top-coded at age 60 and so single year of age preferences are lacking at ages over 59. Experimental assignment of preferences to ages 60–70 showed that availability ratios at ages 50–60 that were based on preferences and populations truncated at 60 were heavily biased, particularly for women. The preferences were therefore extended forward by attributing preferences to ages 60–70.

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Since the structure of age preferences, specific by age difference, was very similar in the age range 50–59, the average preference at these ages for age differences in the range − 41 to + 10 years was assigned to ages and partner ages 60–70 (the extremes of observed preferences in this age range were of persons aged 59 for partners aged 18 and persons aged 50 for partners aged 60). Preferences for age differences of 11–15 years were assigned by averaging the preferences of those aged 45–49, setting any unobserved preferences for age differences of 16 and above and below − 41 to zero. The attribution of preferences affects largely the age group 50+, and so their availability ratios are rather less firmly grounded than are those at younger ages.

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