Full Title: Adapting the 'own children method' to allow comparison of fertility between populations with different marriage regimes

Short title: Adapting the OCM for marital fertility

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Abstract

The Own Children Method (OCM) is an indirect procedure for deriving age-specific and total fertility rates from children living with their mothers at a census or survey. The method was designed primarily for the calculation of overall fertility, although there are variants which allow the calculation of marital fertility. In this paper we argue that the standard variants for calculating marital fertility can produce misleading results and require strong assumptions, particularly when applied to social or spatial sub-groups. We present two new variants of the method for calculating marital fertility: the first of these allows for the presence of non-marital fertility and the second also permits the more robust calculation of rates for social sub-groups of the population. We illustrate and test these using complete count census data for England and Wales in 1911.

Keywords

Fertility

Total fertility rate

Age-specific fertility rate

Estimation

Own Children Method

Marital fertility

Historical demography

England and Wales

Introduction

In the absence of the recording of a mother's age at the birth of her child, it is very difficult to generate age-specific fertility rates. The Own Children Method (OCM) for fertility estimation was first developed in the 1960s as a way of deriving estimates of age-specific and total fertility from cross-sectional census data (Grabill and Cho 1965). The OCM exploits the facts that censuses and surveys record the age of each family member, and that young children are usually recorded in the same households as their mothers. This enables the age of mothers at the births of their children to be inferred, and age-specific child-woman ratios to be calculated. The method transforms these into age-specific fertility rates, using adjustments for child mortality, adult female mortality, and children living away from their parents.

Over the last forty years, the OCM has been frequently applied to historical census data. Full birth histories seldom exist for historic populations of women and it is impractical to create equivalent data through record linkage, so the OCM is frequently the most fruitful method for generating age-specific and total fertility rates in the past. The widening availability of 'big data' in historical demography, with the increased digitisation and harmonisation of full-count census data (such as I-CeM, IPUMS and NAPP), has opened up more avenues for the OCM to be used on large-scale populations (Ruggles 2014).

The first demographic transition was primarily a change from fertility control by marriage to fertility control within marriage and historical demographers are therefore particularly interested in separating out the contribution of changes in marital fertility during the fertility transition (Hinde 2003, p.219). Although the OCM was designed to calculate overall fertility, it has frequently been used to calculate marital fertility, using a variety of assumptions or additional data to adjust for the role of marriage in exposing women to the risk of conception. These variants of the basic OCM, however, have rarely been compared or tested for sensitivity to ways of calculating exposure. In this paper we argue that the commonly used variants have two main disadvantages: in the first place they require the strong assumption — known to be untrue in historical Europe — that all births occur within marriage (Laslett and Oosterveen 1973, Adair 1996, Muir 2018). Secondly, ways of calculating exposure to marriage are likely to produce distorted estimates when applied to social groups (Cho, Retherford and Choe 1986, pp. 30-32). This paper offers two new ways of calculating marital fertility, which overcome these issues. It illustrates and tests them using 'big data' from the 1911 census of England and Wales, the first British census to include questions on marital duration, and which therefore allows several different ways of calculating marital fertility to be compared and evaluated.

Background

The own-children method was pioneered by demographers at the East-West Center for demographic research in Hawaii. It was first set out clearly in the 1960s (Grabill and Cho 1965) and subsequently refined and tested by a series of applications to different censuses and surveys, and comparisons to other methods (Cho 1974; Rindfuss 1976, 2977; Retherford and Cho 1978; Retherford et al. 1980; Goldstein and Goldstein 1981; Retherford and Mirza 1982; Retherford et al. 1984). The East-West Center produced a computer programme and accompanying documentation to aid analysis (EASWESPOP, East-West Center 1992), although the method is not difficult to implement. Early applications calculated fertility rates for the mid-twentieth century United States, and in the 1960s and 1970s it was also widely applied to Asian countries such as Malaysia and Korea (Grabill and Cho 1965; Cho 1968, 1974; Rindfuss 1976, 1977). During the 1980s and 1990s the full birth histories (FBHs) collected by surveys such as World Fertility Survey (WFS) and Demographic and Health Surveys (DHS) became the main source of fertility information. Comparisons between FBHs and the OCM from accompanying household surveys showed that the OCM produced comparable estimates although age reporting was often better in FBHs (Retherford and Alam 1985). Some use of the OCM continued, however, with focus moving towards the calculation of fertility differences among migrant groups in wealthy countries (Abbasi-Shavazi 1997; Dubuc 2009; Coleman and Dubuc

2010; Krapf and Kreyenfeld 2015). The method received further validation with a thorough assessment of its performance when compared to DHS data which concluded that the OCM is generally at least as accurate as FBHs, at least for survey data, and that selection effects can distort fertility from FBHs (Avery et al. 2013). In his more general assessment of the reliability of reverse survival methods, Spoorenberg (2014) also tested some of the more important assumptions of the OCM against a simulated population. Most recently the method has been innovatively used to produce estimates of male fertility (Schoumaker 2017).

Since the 1970s, the OCM has been used by historical demographers who, of course, have not benefitted from the development of comparable sample surveys, and who study populations where vital registration was not yet established or did not request information on age of mother at birth of child. Although vital registration was operational in England and Wales from 1837, for example, the age of mother was not recorded on birth certificates until 1938 (Higgs 2004, p. 210). Some historical applications have used the EASWESPOP programme, some have used the APPLAUSI programme written by historical demographers (see collection of papers in Breschi et al. 2003), and others have performed the calculations themselves. The method has been applied to small area census populations in the USA (Hareven and Vinovskis 1975; Haines 1977, 1978), England and Wales (Woods and Smith 1983; Garrett et al. 2001; Boot 2017), and Germany (Gruber and Sholz 2016), to USA census samples (Tolnay 1981; Tolnay et al. 1982; Tolnay and Guest 1984), and to full count historic censuses for the USA and Sweden (Hacker 2003; Scalone and Dribe 2012). Further historical studies have applied the OCM to tax registers from Tibet (Childs 2004) and from 15th century Florence (Breschi and Serio 2003) and to Japanese population registers (Kurosu 2003).

As with contemporary data, many historical applications have compared fertility among sub-groups of the population, including by ethnic group, migrant status, socio-economic status, and urban/rural location, and there have been a number of studies which use unadjusted numbers of young children living with their mothers in multivariate and sometimes multilevel models of the determinants of fertility (Haines 1978; Scalone and Dribe 2012; Dribe and Scalone 2014; Klüsener et al. 2016; Hacker 2016; Dribe et al. 2017).

There is little corroboratory data for historical populations against which to test the results obtained via the OCM, but Swedish vital registration evidently asked questions about mother's age and Scalone and Dribe found that national level OCM estimates were very similar to age-specific rates calculated from vital registration data (Scalone and Dribe 2017). Haines compared OCM fertility estimates to those calculated using the 2-parity increment method and found the OCM to be more robust and accurate than parity statements, particularly for older women (Haines 1989). Derosas (2003) and Oris (2003) compared fertility calculated using the OCM with that from birth histories generated from population registers (for Venice and Belgium respectively). Both found that a failure to use the correct mortality schedule for different social or migrant groups could lead to over-estimation of fertility in some groups and under-estimation in others. The UN indirect estimation manual urges caution in interpreting OCM results for sub-populations which are not closed to migration, and the Belgian study demonstrated the compositional changes which can be produced by migration and highlighted the problems associated with population movement between areas with different mortality rates (UN 1983, p. 183; Oris 2003).

Despite these investigations, there has been little critical engagement with the way that the OCM calculates, and may distort, marital fertility. In this paper we discuss these potential issues and present new variants of the OCM for calculating marital fertility. The next section briefly describes the basic OCM for estimating overall, rather than just marital, fertility, noting the assumptions and adjustments which are commonly made. This is followed by a section in which the data we will use in this paper are detailed. Then comes a section in which we use the OCM to calculate overall fertility rates for England and Wales (1836–1911) and review the assumptions made when using the method. Finally we present and compare a number of different variants of the OCM for estimating marital fertility, including our two new variants.

The own-children method for overall fertility, assumptions and adjustments

In its classic form the OCM relies on matching children aged under 15 in a census or survey with their mother in the same household. Matched children and mothers are then cross-tabulated by single years of age and both children and mothers reverse-survived to yield annual births to woman by single years of age at birth of child. The number of women of each age in the population are also reverse survived, and used as population denominators to calculate age-specific fertility rates (ASFRs) for the 15 years leading up to the census or survey. These can be combined into 5-year age groups for added robustness. Early applications of the method were performed on censuses which had a pre-coded variable specifying the number of children under age 5 and therefore it was not possible to calculate annual rates and the precise ages of women at the birth of their children could not be ascertained by back projection (Grabill and Cho 1965). Instead the Sprague osculatory interpolation was used to redistribute births to the correct ages of women, and this procedure was also followed in some historical demography applications where small numbers of women may have made working with single year rates more problematic (Haines 1978, 1979; Woods and Smith 1983; Hinde and Woods 1984; Garrett et al. 2001).

ASFRs and TFRs (total fertility rates) can be produced easily for sub-groups, allowing comparison of fertility levels and trends over time among sections of the population for which age-specific fertility might not otherwise be easily obtained. These steps by themselves would produce accurate age-specific and total fertility rates under the following assumptions:

1. there is no mortality in the population;
2. enumeration is complete;
3. the ages of women and children are correctly reported;
4. all children under 15 live in the same household as their mother, and mother-child links have been correctly identified;
5. children are matched to their biological mothers;

6. group characteristics are constant.

The extent to which violation of any of these assumptions produces bias in the estimates depends on the time and place being studied but it is common to make adjustments to the numbers of children used in the cross-tabulations in order to take account of mortality in the years leading up to the census (inflating numbers of both women and children), to redistribute the children for whom a mother has not been identified, and to allow for the possible under-enumeration of very young children. These adjustments carry their own assumptions (the numbering below relates to the assumptions above and represents more realistic assumptions as a result of the adjustments):

1a. the age distribution of the mothers of children who died is the same as the age distribution of the mothers of children who survived;

1b. women who died had similar fertility to the rest of the population;

4a. the age distribution of the mothers of very young children not recorded with their mother at census is the same as that of mothers who are recorded with their very young children;

These assumptions and adjustments will be briefly discussed as we illustrate the method with data for nineteenth and early twentieth century England and Wales, which is described in the next section. In particular we will discuss age-misstatement and under-enumeration (assumptions 2 and 3), migration (assumption 6), mortality (assumptions 1a and 1b), and ‘non-own’ children (assumptions 4a and 5).

Data

This paper has used the individual level census returns for England and Wales for the decennial censuses 1851–1911 (excluding 1871) published and enhanced by the I-CeM project. These amount to 17.5 million individual records for 1851 and over 36 million in 1911. Numbers of children and women used for the fertility calculations are therefore extremely robust and are not, at the national level, subject to small number fluctuations, even for single year age groups. For the calculation of overall fertility rates from 1836 to 1911 we use all available censuses, but for the exposition of our new variants we concentrate on the 1911 census which, as it asked women currently un a marriage a question on the duration of that marriage, allows a more extensive comparison of different variants of the OCM for calculating marital fertility. For these comparisons we use a substantial subset of the 1911 data consisting of those women with plausible reported marital duration and age, and whose husband was co-resident with them on census night. Married women with their husband present were excluded from the 1911 sample if their age or marital duration was missing, or if their age at marriage, calculated as age at census minus marital duration, was implausible. The 1911 census included questions on the numbers of children ever born, still alive and dead, and women with inconsistent answers to these questions were also excluded. Of 6,662,862 women reported to be aged 15–64 and married in the ICeM data for 1911, 3.7 per cent were excluded due to implausible answers either to the marital duration of one or more of the fertility questions. The majority of these instances are likely to be due to data transcription errors, which will be unbiased. A further 14.8 per cent of the women were excluded from calculations of marital fertility (variants B, C, D and E) because they could not be linked to a husband in the same household. These women (discussed in more detail later) tended to have fewer co-resident children than women with husbands present in the household, so their omission results in slightly higher fertility than if all women reported to be married are included. However excluding these women enabled us to make comparisons between different variants of the method using exactly the same sample. The final sample used for comparative analysis of marital fertility amounts to 5,425,682 women.

For national mortality data we used single year life-tables downloaded from the Human Mortality Database. Sub-national mortality estimates for infant and early child mortality were based on published mortality statistics from the Quarterly and Decennial Reports of the Registrar General (Jaadla and Reid 2017).

While many of the results we present are for England and Wales as a whole, we have also calculated sub-national estimates based on Registration Sub-Districts (RSDs), of which there were around 2,000 in each census year. We have classified RSDs into eight types of places, defined by their occupational structure and population density, each with a distinctive demographic regime (see https://www.populationspast.org/about/ under 'type of place' for more details of how these types of place were defined). In this paper we concentrate on three key types of place - professional, mining, and textile - which have very different marriage patterns, as illustrated in Figure 1 where the figures for England and Wales are also shown. Each panel in Figure 1 shows the percentages of women who were married and whose husbands were present in the same households on census night, 1911, for a different type of place. In England and Wales as a whole, and generally in the types of place not otherwise singled out in Figure 1, women married earlier than men but over a more extended range of ages, thus the slope of the 'E&W women' line in each panel is not as steep in its rapidly increasing phase (up until about age 30) as the line for 'E&W men'. More striking, however, is the fact that in the population as a whole a greater proportion of women than of men remained unmarried towards the end of the fertile age range. The maximum proportion of women 'married with spouse present' occurred at around age 40, and was lower than the male maximum which occurred in the mid-40s. The downturn of the proportions married at older ages predominantly reflects bereavement which was more common among women than men due to higher male mortality, women being usually younger than their husbands, and lower rates of female remarriage after widowhood.

Figure 1 about here

Table 1 about here

The different panels of Figure 1 show that this pattern differed between professional, mining and textile areas in 1911. In mining places women married particularly young, and were more likely to be reported as married at any age than men, a characteristic mainly due to significant in-migration of young, single men and out-migration of single women. These migration patterns are visible in the unusually high sex ratio amongst adults aged 20-49 in mining areas, shown in Table 1 along with certain demographic characteristics for England and Wales and the three selected types of place in 1911. High adult sex ratios are usually attributable to an imbalance of job opportunities for men and women. Professional areas, in contrast, attracted young women who migrated in to work as domestic servants. These servants swelled the ranks of unmarried women, leading to extremely low sex ratios, low proportions of women married, and high estimates of age at marriage. Textile areas also had plentiful work opportunities for women, but by 1911 these were mainly taken up by local girls, so that the proportions married were little distorted by in-migration. These different patterns of marriage are likely to have led to different amounts of exposure to marriage among married women of the same age. For example, a married women of age 25 in a mining area was likely to have been already married for substantially longer than a married woman of the same age in a professional area. However, if migration had a life-cycle element, i.e. single women migrated in to work but left to marry and live elsewhere, it is possible that those women who were married in professional areas actually did so at an earlier age than implied by the proportions married. It is a well-known issue that cross sectional measures such as SMAMs, based on a synthetic cohort, can be significantly distorted by migration, especially at sub-national levels (Schürer 1989). This paper considers the effect of these issues on exposure to marriage in the context of the OCM.

Using the OCM to calculate overall fertility in England and Wales

We used the individual level data from the censuses of 1851 to 1911 to derive the annual TFR series for England and Wales, 1836 to 1911, shown in Figure 2. Because the 1871 census is not available there is a gap between 1861 and 1866. These estimates were calculated using the data for all women aged 15–64, together with that for children aged 0–14 years. We adjusted the figures for 'non-own' (un-matched) children by multiplying the number of children in each age of child and age of mother combination by the reciprocal the proportion of children of that age who were matched to mothers. We also inflated numbers of both children and women to account for mortality. These single year TFRs allow greater elucidation of the method and an assessment of some of the OCM assumptions laid out below.

Figure 2 about here

Age-misstatement and under-enumeration (assumptions 2 and 3)

Each segment of line is derived from a different census. The rightmost end of each segment is contributed by children aged zero (born during the year before the census), and each previous annual point is produced by children a year older, up to age 14. Ideally we would see a smooth line with exactly overlapping segments. Jagged fluctuations usually indicate age-heaping or age-misstatement among children, and a lack of overlap can indicate differential under-enumeration by age and census year or faulty estimates of mortality. Figure 2 shows that the fit is relatively good. There is evidence of age-heaping among older children in the earlier censuses, but this was eliminated by the end of the century. However the overlap between segments is not perfect: in general older children from one census indicate higher fertility than that indicated by the youngest children in the next census. There appears to have been a strong deficit of zero and one year-olds and a bulge of two and three year-olds. To some extent this can be explained by the tendency of parents to report children as the next age up (i.e. to report an eleven month-old as one year), but it is also possible that some infants were omitted from the census forms.

We tested for age-misstatement by smoothing the age distributions and found only a very small difference in TFR. The possible under-enumeration of zero year-olds, a common issue with censuses, is a potentially larger problem (Myers 1993). We tested for this using Lee and Lam's age adjustment factors, obtaining revised estimates slightly higher than the raw ones, by 0.1 to 0.3 children. Lee and Lam argue that their adjustment factors for the 0–4 age group seem implausible, and the extent of underestimation is therefore unclear (Lee and Lam 1983). As we do not consider that the age distributions of the very young would be reliable enough to produce specific adjustment factors when working with small subgroups of the population, and because these would be further complicated by differences in migration, we decided not to adjust for possible under-enumeration. A uniform adjustment throughout would be of little benefit when comparing different groups. More detail on these tests is given in Online Appendix 1.

Migration (assumption 6)

Another potential issue for the accuracy of OCM rates is the migration of young people - this is not a major issue with national data, but it is potentially a problem with place-specific calculations. Calculations carried out for different types of place (not shown here) demonstrate that young people started to leave home for employment as life-cycle servants from the age of ten or even earlier (see Schürer et al. 2018). This movement among young people results in underestimation of fertility 10–14 years before the census in agricultural districts, and overestimation in professional areas where servant-keeping was common.

In view of these considerations, most of our analyses only use children aged zero to four, who were less likely to be found living apart from their parents, and provide estimates of fertility for the five years immediately preceding each census combined. While this is avoids the problem of migration of older children, we do have to be aware that the under-enumeration of infants weighs more heavily when analysis is restricted to younger children.

Mortality (assumptions 1a and 1b)

Various authors have assessed the impact of using inappropriate levels and trends in mortality on overall and sub-group fertility. There is a general consensus that at a national level the failure to account for mortality leads to underestimates of fertility, but that mis-specification of mortality levels makes little difference and trends over time are relatively little affected (Rindfuss 1976; Retherford and Cho 1978; Retherford and Alam 1985; Cho, Retherford and Choe 1986; de Santis

2003; Spoorenberg 2014). However, an inability to properly account for differences in mortality among sub-groups can lead to mis-specification of differences in fertility between groups (Rindfuss 1976, Retherford et al 1980, Goldstein and Goldstein 1981; Young 1992; Oris 2003), although some argue that this may not confound fertility differences if groups with high mortality also have high fertility (Scalone and Dribe 2012, Dribe and Scalone 2014).

Online Appendix 2 explains how we have examined the effects of possible mis-specification of mortality, both for the overall population and for sub-groups. In summary this shows that if no mortality adjustment is made then estimates of overall fertility are between ten and 23 per cent too low, with the larger underestimations for earlier census years when child mortality had not yet started to fall. Similarly, underestimation was larger in settings where mortality was higher, such as mining and textile areas. Our analyses demonstrated that a ten per cent overestimation in early age mortality resulted in fertility between five and six per cent too high, and a ten per cent underestimation in mortality resulted in fertility which was six to eight per cent too low. Because the differences in mortality between our geographic areas were considerably greater than ten per cent, and because high mortality was not always correlated with high fertility (textile areas, for example, had low fertility but high infant and child mortality, while the reverse was true in agricultural areas), we felt that it was necessary to use as accurate mortality adjustments as possible.

'Non-own' children (assumptions 4a and 5)

Perhaps the most problematic assumption for the application of the OCM to nineteenth century England and Wales is the assumption that 'non-own' children (i.e. those who *have not* been matched to their mother in the census) are representative of 'own' children (who *have* been matched to their mother), particularly in respect to their mothers' ages. In assessing this assumption it is worth considering which groups of children were likely to end up as 'non-own'. These include legitimate children living with their father but not their mother, and those living with neither parent: orphans, children living temporarily or permanently with relatives or other carers and those in institutions.

Some legitimate children living with their mothers in extended households may not have been matched to their mother, due to complexity of the households. However this is not likely to be a large problem as less than two per cent of households were extended in 1911, and even fewer will have contained children under the age of five (Schürer et al 2018). Transcription error may also have prevented some child-mother matches from being made. Such un-matched legitimate children might well have been reasonably representative of other legitimate children. However the majority of illegitimate children are also likely to be included in the 'non-own' category. Legitimacy status was not recorded on the census, so children of unmarried mothers can only be securely identified when their mother was head of her household. This was very uncommon, and it is likely that many illegitimate children were living, with or without their mother, as grandchildren of the head of household. Unless the grandchild had a different surname to the head (in which case it was likely to be the child of a married daughter), it is very difficult to tell whether the grandchild was illegitimate, which of the head’s own children was their parent, or indeed whether their parent was present in the household at all. In most cases where a (potentially illegitimate) child was living as a grandchild, she or he was not matched to any potential mother.

Table 2 shows the numbers of children under five in the 1911 census and, for those who were matched to a mother, the marital status and average age of those mothers. It also provides the number of unmatched children, and an estimate of the number of illegitimate children aged under five in England and Wales in 1911. It is notable that the average age of the unmarried mothers who could be identified as such was significantly younger than that of mothers who were married with their spouse present on census night (28.9 years compared to 32.3 years). If these unmarried mothers were representative of all unmarried mothers in terms of age, or were older, then the similarity assumption (4a above) would be violated. It is plausible that some illegitimate children were passed off as the children of older, married females which would result in a partially compensatory distortion of the age-structure of fertility (Rindfuss 1976). Interestingly, the average age of married mothers whose spouse was not present in the household was closer to that of unmarried mothers than to that of other married mothers, and we suspect that a sizable proportion of such women were not actually married, but were passing themselves off as such in order to make themselves look more respectable. We will return later to the implications of this for the calculation of marital fertility.

Table 2 about here

To test the implications of the possibility that the mothers of 'non-own' children were in fact younger than the mothers of 'own' children, we recalculated overall age-specific fertility rates assuming that the mothers of 'non-own' children followed the age distribution of the mothers of illegitimate children. To do this we calculated ASFRs for single women only, using children of identifiable single mothers and assuming that unattached children belonged to this group. We then calculated overall ASFRs by weighting the ASFRs for single and ever-married women. The proportionate effect on ASFRs was relatively large at very young ages - with a difference of 30 per cent in the 15–19 age group and six per cent in the 20–24 age group. However the absolute effect was very small even among these young age-groups because fertility was low for these women, and the proportionate effect was small at older ages. Of course this is a maximum effect, on the assumption that all un-matched children were illegitimate. In fact the number unmatched children and estimate of illegitimate births surviving to age five, both given in Table 2, suggest that a maximum of 60 per cent of unmatched children were illegitimate, and less if a proportion of the children of women with spouse absent were also illegitimate. Therefore the distortion of ASFRs is likely to be less than half the effect given above.

Despite this, the consideration of 'non-own' children is still important as the possibility that a large proportion of them were illegitimate raises issues for the calculation of marital fertility. It is very difficult to tell whether illegitimate children really did make up a large proportion of 'non-own' children. However our estimates of the number of illegitimate children born in the five years leading up to 1911 who survived to be enumerated in the census suggest that up to 60 per cent of unmatched children may have been illegitimate. It is also notable that there was striking similarity in the geographic patterns of the percentage of births that were illegitimate, and the percentage of children under the age five who were not living with their mother. Figure 3 shows that the correlation between these two variables, at a county level, was high (R^2 of 0.78).

Figure 3 about here

Of course we are not arguing that all illegitimate children will have been 'non-own' (living apart from their mother): we have already mentioned the very small group identified as children of an unmarried mother, and we suspect that there is also a larger group who were 'own' children of mothers who reported themselves to be married. Nor are we arguing that all 'non-own' children were illegitimate; there will have been 'non-own' legitimate children as mentioned above. Nevertheless, we are arguing that illegitimate children made up a large proportion of 'non-own' children, and thus for the calculation of marital fertility in nineteenth century England and Wales 'non-own' children should not be redistributed among married women. We also argue that women who reported themselves as currently married but had no husband in the household were likely to have been unmarried and therefore that including them in marital fertility is misleading. Of course this might not have been the case in all communities: many husbands may have been temporarily absent from fishing villages, port cities and military bases, and in forthcoming papers we will compare versions of marital fertility including these women, and widows, to investigate the effects of spousal separation and widowhood on changes in fertility.

Marital fertility using the OCM

The OCM can be used to calculate marital fertility, but to do this it is necessary to take duration of marriage, or exposure to the risk of pregnancy, into account. In the past, three variants for doing this have been employed:

Variant A: ASFRs multiplied by inverse of proportions married

This variant calculates ASFRs using the standard OCM and then multiplies these by the inverse of the proportion of women married in that age group (Haines 1978, 1978; Retherford and Mirza 1982; Cho, Retherford and Cho 1986). It generally carries the assumption that all fertility occurs within marriage, although some scholars (e.g. Breschi and Serio 2003) adjust for illegitimacy by multiplying by the proportion legitimate as well as dividing by the proportion married. In our comparisons we have used ASFRs and proportions married for single years of age.

Variant B: Duration of marriage

This variant takes married women only and uses data on duration of marriage, provided by the census or survey, to calculate the number of women exposed at each age (Cho 1968; Tolnay 1981; Tolnay et al. 1982; Retherford, Cho and Kim 1984). Here we have used the following formula to calculate the number of woman-years of exposure (see Online Appendix 3 for explanation of the formula):

$$N_x = \frac{M_x^0}{8} + \frac{M_x^{1-34}}{2} + \frac{M_{x+1}^1}{2} + M_{x+1}^{2-34} + \frac{M_{x+2}^2}{2} + M_{x+2}^{3-34} + \frac{M_{x+3}^3}{2} + M_{x+3}^{4-34} + \frac{M_{x+4}^4}{2} + M_{x+4}^{5-34} + \frac{3M_{x+5}^5}{8} + \frac{M_{x+5}^{6-34}}{2}$$

where M_x^i is the number of married women age x with duration of marriage i.

Variant C: No adjustment for exposure

This variant uses married woman only and recognises that back-projecting the number of women married at later ages overestimates exposure to marriage among young women, and therefore underestimates fertility in these ages (Garrett et al. 2001). If children not matched to their mother are adjusted for, this variant also assumes that all fertility occurs within marriage. To avoid having to make this assumption, some researchers consider only currently-married women with their spouse present in the household on census night. This carries the alternative assumption that all children living apart from the mother were the products of unmarried motherhood, orphanhood or abandonment. Those who follow this route generally only consider children under five years of age as the assumption becomes less plausible at older ages (Woods and Smith 1983; Hinde and Woods 1984; Garrett et al. 2001).

Most historical data lack information on duration of marriage, so variant A is the most common variant deployed. It has been used to calculate marital fertility for countries, geographical areas, and also for social groups (Breschi, Derosas and Rerraroli 2003; De Santis 2003; Oris 2003; Scalone and Dribe 2012; Gruber and Scholz 2016; Scalone and Dribe 2017). Using this variant to calculate fertility for social groups involves calculating the percentage of women of each social group who were married, and therefore being able to identify the social status of both married and unmarried women, and this is highly problematic - at least for the British context - because of the link between female occupational status and marital status. Social status or class may be assigned to a woman on the basis of her own occupation, but this is difficult when not all women work and particularly when they are likely to have given up paid work after marriage or child-bearing. In such circumstances it is common to use father's occupation for daughters still living at home and husband's occupation for wives, but this leaves a gap in societies, such as nineteenth and early twentieth century England and Wales, where women had a lengthy period during their teens and twenties where they were with neither father nor husband. Their own occupation during this period was constrained to certain types of job, such as domestic or agricultural servants, which in many cases reflected neither the household they came from nor that they would eventually marry into. Neither can occupation of the head of the household in which they were living between leaving home and marriage be used as a proxy, as many women were domestic servants in a wealthier household than that from which they originated or in which they ended up. Certain occupational groups were therefore dominated by single women, but this is not a good guide to the marriage age or duration of any group of married women.

We now present two new variants to calculate marital fertility. The first, variant D, allows relaxation of the assumption that all fertility occurs within marriage, and the second, variant E, also allows the more robust calculation of marital fertility by social group.

Variant D (new): Female proportions married applied to female populations

This variant uses married women only and back-projects exposure based on the proportions of women married at each age in the five years leading up to the census. This variant allows 'non-own' children to be treated as illegitimate, and also allows researchers to exclude women (and their children) without their spouse present if this is felt to be appropriate. In our application of this variant we used the children aged zero to four years of married women only, and we calculated fertility rates by single years of women's age. We calculated rates for the five years leading up to the census, so our calculations needed to include both children and exposure contributed by women who had moved into an older age by the time of the census.

The estimation procedure of age-specific marital fertility rates (ASMFRs) follows a set of computational steps. Firstly, enumerated children were reverse projected to the time of their birth so that the number of children born to women aged x in the five years before the census could be estimated:

$$B^x = 0.5 * \left(C_0^x + C_0^{x+1} + C_1^{x+1} + C_1^{x+2} + C_2^{x+2} + C_2^{x+3} + C_3^{x+3} + C_3^{x+4} + C_4^{x+4} + C_4^{x+5}\right)$$

where C_a^x is the number of children aged a to women aged x in the census. Adjusting for mortality, gives:

$$B^x = \left(\frac{(C_0^x + C_0^{x+1})}{{}_1L_0} + \frac{(C_1^x + C_1^{x+1})}{{}_1L_1} + \frac{(C_2^x + C_2^{x+1})}{{}_1L_2} + \frac{(C_3^x + C_3^{x+1})}{{}_1L_3} + \frac{(C_4^x + C_4^{x+1})}{{}_1L_4}\right) * 0.5$$

where ${}_1L_a$ values are single year survivorship probabilities from an appropriate life table. The number of children born in the five years leading up to the census and the years of exposure are illustrated in the lexis diagrams in Figure 4, where the box outlined thickly in black indicates the experience we want to capture. The next step involves estimating woman years of exposure for women aged x. The equation used in this case is:

$$N_x = (W_x * m_x) * 0.5 + (W_{x+1} * m_x) + (W_{x+2} * m_x) + (W_{x+3} * m_x) + (W_{x+4} * m_x) + (W_{x+5} * m_x) * 0.5$$

where W_x represents the number of women aged x at census year and m_x is the proportion of women aged x who are married at the time of enumeration.

Figure 4 about here

The calculation of age-specific marital fertility rates for single years or five-year age groups is then straightforward:

$$ASMFR = \frac{B^x}{N_x} * 1000$$

Variant E (new): Male or standard proportions married applied to female populations

This variant, which enables the calculation of fertility for social groups based on husbands' occupations, uses married women with resident husbands only and back-projects exposure based on the relative proportions of men married at each age in the five years leading up to the census, adjusting for the mean spousal age gap.

In this variant the numerator of age-specific marital fertility equation is exactly the same as above in variant D. However, the woman years of exposure for women aged x is estimated by:

$$N_x = \frac{h_{x+y}}{h_{x+y}} * M_x * 0.5 + \frac{h_{x+y}}{h_{x+y+1}} * M_{x+1} + \frac{h_{x+y}}{h_{x+y+2}} * M_{x+2} + \frac{h_{x+y}}{h_{x+y+3}} * M_{x+3} + \frac{h_{x+y}}{h_{x+y+4}} * M_{x+4} + \frac{h_{x+y}}{h_{x+y+5}} * M_{x+5} * 0.5$$

Where M_x denotes the number of married women aged x at census year, for x to $x+4$; h_x is the proportion of men aged x who are married at census; and y is the average spousal age gap, rounded to a whole number (note that in the first term $\frac{h_{x+y}}{h_{x+y}}$ cancels out). The rationale for this is as follows: in variant D each cell of exposure is

$$(W_{x+n} * m_x) \text{ or } W_{x+n} * \frac{M_x}{W_x} \text{ or } \frac{W_{x+n} * M_x}{W_x}$$

Where, as before, W_x represents the number of women aged x at census year and m_x is the proportion of women aged x who are married at the time of enumeration.

Multiplying this through by the number of married women (M_{x+n}) gives:

$$\frac{W_{x+n} * M_x * M_{x+n}}{W_x * M_{x+n}} \text{ or } \frac{W_{x+n} * M_x}{W_x * M_{x+n}} * M_{x+n}$$

which can be re-written as:

$$\frac{M_x / W_x}{M_{x+n} / W_{x+n}} * M_{x+n} \text{ or } \frac{m_x}{m_{x+n}} * M_{x+n}$$

In other words, we are adjusting the number of married women age $x+n$ at the census by the ratio of the proportion married at that age relative to the proportion married one or more years older. A woman's marital duration will, by definition, be the same as that of her husband, so if the spousal age gap is relatively constant by age, and if the change in proportions married by age is similar among men and women, then using the male ratios of proportions married, shifted by the average spousal age gap, should give a similar result. We will consider the plausibility of these assumptions as we discuss the results below.

Comparing variants for calculating marital fertility - for England and Wales

Figure 5 shows the age-specific marital fertility rates (for women aged 20 to 49) for England and Wales in 1911, calculated using the five variants of the OCM laid out above, and Table 3 summarises these in terms of the Total Marital Fertility Rates (TMFRs). Two versions of TMFR are shown in Table 3: $TMFR_{20}$ shows the number of children a woman would have had if she married at age 20 and experienced the age-specific marital fertility rates for each age in turn, and $TMFR_{25}$ shows the number of children she would have had if she married at age 25.

Figure 5 about here

Table 3 about here

Variant A produces the highest fertility rates. When single years of age are used, this variant differs from variant D primarily in the fact that the numerator in variant A includes 'non-own' children. Not unsurprisingly, assuming all children are legitimate results in a higher estimation of marital fertility, but the effect is not very large. Variant C consistently gives the lowest fertility rates, showing implausibly low fertility among young married women. This is attributable to the lack of adjustment for marital duration: in essence the variant assumes married women had been married for the full five years before the census. We strongly urge against using this variant.

It is worth remembering that the numbers of children used for variants B, C, D and E are exactly the same, therefore the differences between them are due entirely to differences in calculated exposure. Differences in exposure are small, with variant D estimating slightly higher exposure for older women, and slightly lower exposure for younger women. Because the number of married women is high at older ages, differences in exposure make very little difference to fertility. However relatively low numbers of married women in their early twenties mean that small differences in exposure can produce large differences in fertility. Therefore these variants result in very similar fertility rates among women over the age of 30, but more discrepancy among women in their early twenties, resulting in noticeable differences in $TMFR_{20}$ (5.89, 6.17 and 5.93 using variants B, D and E respectively), although differences in $TMFR_{25}$ are minimal (3.64, 3.68 and 3.65 respectively). Before considering how these variants perform in situations with different marriage patterns, it is worth considering why they produce different answers.

As noted in the data section above, the preceding comparisons have all (excepting variant A) been carried out using only those married women in 1911 who gave consistent age, duration of marriage and fertility information and who could be linked to a husband in the same household. This selection was used in order to allow comparisons of the different variants, but calculations of proportions married (for variants A, D and E) have to use both married and unmarried women, as exclusion of the 20 per cent of married women who gave inconsistent information or whose husband was not in the household would alter the proportions. Systematic differences in durations married between excluded and included married women may have affected the calculations. It is important, therefore, to consider whether women giving inconsistent fertility or marital duration information in 1911 were an unbiased subset of the married population.

Figure 6 shows the distribution of marital durations across couples with consistent information in 1911. Unsurprisingly there is heaping on durations of 10 and 20 years, but the most noticeable feature is a deficit of couples with very short marital durations, particularly zero years.

While it is possible this could have been produced by a sudden drop in marriages in the years leading up to 1911, the number of marriages registered in England and Wales was actually slightly higher in 1908-10 than it had been in previous years. The deficit of couples reported as recently married in the 1911 census was also noticed by Thomas Stevenson, who compiled the official reports on fertility arising from the 1911 census. He argued that this was partly due to ‘accidentally’ reporting the duration at the next wedding anniversary (1911 Census of England and Wales 1923, pp.viii-x). Stevenson did not discuss why couples may have given such ‘accidental’ responses, beyond stating that he expected it to be more likely among less educated couples. We speculate that some couples may have worked out their marital duration from the age of their oldest child, or by subtracting 1911 from the year of their marriage. Figure 7, showing the distribution of couples giving particular marital durations by social class, indexed against zero years, demonstrates that the deficit was indeed strongest among the manual and un-skilled social classes. Stevenson also noted that the deficit of marriages of less than one year was particularly acute among those with young ages at marriage, who were also most likely to have been pregnant on marriage, and he argued that such couples were systematically over-stating their marital duration to hide a pre-marital conception (1911 Census of England and Wales 1923, pp.viii-x).

Figures 6 and 7 about here

The systematic overstatement of marital duration could explain the lower levels of fertility calculated using variant B. However we should also consider whether the deficit of short marital durations could be the result of missing data. The census instructions stated that marital duration and fertility information was to be filled out for married women, but it was not uncommon for a man to have entered the information against his own rather than his wife’s name. In such cases the census clerks would strike this out and re-write the information in the row for the wife. However, in so doing they would sometimes use '-' to indicate a zero. The transcriptions do not include the crossed out records and so it is impossible to tell whether the '-' indicates a real missing entry (for example if they were not married or they did not know their marital duration) or a checking clerk's zero. If the absence of very short marital durations is because they were omitted, rather than because marital durations were overstated, then fertility calculated by variant B will be correct for those included in the selection, but exposures calculated using variant D will overestimate the durations of marriage, and therefore underestimate fertility. This recording practice might have been more likely among less educated couples, but it should also have been more likely among the young, where short marital durations were concentrated. However, although very young women were particularly likely to have been missing a marital duration (perhaps due to some unmarried mothers reporting themselves as married), the proportion missing a marital duration increased steadily from about age 25. Overall, therefore, it is likely that the deficit of people married for less than a year was due more to overstatement of marital durations leading to understatement of fertility in variant B, but this may have been partially offset by omission of some of those married for less than a year.

For England and Wales as a whole, variant E gives slightly lower fertility than variant D, particularly among younger women. This is due to the fact that the marriage curve for men is slightly steeper than that for women, so the area under the curve for a five year age span is larger than that for women. However differences are small at the national level.

Comparing variants for calculating marital fertility - for types of place

We developed variant E for use with social groups where it is not possible to calculate proportions of married women, and our first tests compare variants D and E within places characterised by very different male and female marriage patterns. Figure 8 shows the ASMFRs for the three selected types of place using the five different variants, and Table 4 shows the corresponding values of $TMFR_{20}$ and $TMFR_{25}$. Table 4 also shows variant E.std, which uses a standard female marriage schedule (here that for all England and Wales in 1911) instead of the male marriage schedule for the type of place.

Table 4 about here

Figure 8 about here

Table 4 shows that differences in fertility estimated using the different variants for places with extreme marriage patterns are larger than for England and Wales as a whole. Variant C shows particularly large understatement for professional areas, where marriage was late, because the over-estimation of exposure is more severe. For England and Wales as a whole variant E estimates slightly lower fertility than variant D, and this pattern is exacerbated in professional areas where the male marriage curve was steeper than the female. In places where women married young relative to men, such as mining areas, the male marriage pattern is less steep than the female and variant E generates slightly higher fertility. In general, however, it seems that different marriage patterns in different places do not make much difference to estimates of fertility, particularly for $TMFR_{25}$. Our final comparison for different types of place tests this by using the same marriage pattern (here that for all England and Wales in 1911) for all types of place. Results, however, show that the male marriage patterns (variant E) for particular types of place represent their respective female patterns better than a standard marriage pattern (variant E.std) does. Online Appendix 4 demonstrates that a further variation, excluding servants from the marital fertility calculations, significantly underestimates fertility in places where servant keeping was common.

This section has tested variants of the OCM for estimating marital fertility for different types of place with different marriage patterns, to allow comparison of variants D and E, which respectively use female and male proportions married to estimate exposure. The marriage patterns in these places were driven by the behaviour of particular social groups – the middle classes in professional areas and miners in mining areas. Yet these social groups are still generally a minority or a small majority in the places they typify, and their marriage behaviour is likely to be more extreme. The next section therefore compares variants B and E for social classes.

Comparing variants for calculating marital fertility - for social classes

Figure 9 shows age specific marital fertility rates for social classes using variants B, E and E.st, and Table 5 shows the corresponding values of $TMFR_{20}$ and $TMFR_{25}$. In general variant E generates lower fertility compared to variant B (which we have already established is likely to underestimate fertility a little). The deficit in $TMFR_{20}$ is equal to between 0.37 and 0.93 children per woman, but this is concentrated among younger women, so the deficit in $TMFR_{25}$ (0.18 to 0.28 children per woman) is both smaller and less variable. The differences between variants B and E are largest for social classes 1 and 2, and also for agricultural labourers, as spousal age gaps in these groups are larger. Variant E.std results differed less from variant B among young ages for social class 2 and for agricultural labourers, but results for social class 1 were just as low as for variant E, and results for social class 5 and miners, whose wives married at younger ages than the national average, were considerably higher.

Figure 9 about here

Table 5 about here

Conclusions

This paper has presented two new variants of the OCM to calculate marital fertility, calculating exposure to marriage by adjusting numbers of married women using ratios of female and male proportions married. We have compared the results from these new variants against the results obtained using older variants of OCM, including the variant which uses exposure calculated from reported durations of marriage. We have tested them using 'big data' from the whole of England and Wales in 1911, for specific types of place with different marriage regimes, and for social groups.

We argue that there is good reason to think that durations of marriage are slightly over-reported and therefore fertility rates calculated using durations married to generate exposure are likely to be a little too low, an effect also found in similar comparisons elsewhere (Tolnay 1982). However, it is worth bearing in mind that inflation of marital duration in order to disguise a pre-nuptial pregnancy or illegitimate child may actually be a better indicator than actual marital duration of exposure to the risk of pregnancy. It has been estimated that around half of all first births in 1800 in Britain were either illegitimate or conceived prenuptially (Levene et al 2005, p. 6). Illegitimacy dropped over the course of the nineteenth century, and it is plausible that there was a concomitant increase in pre-nuptial pregnancy. Marital fertility rates of young women were likely to have been pushed upwards by both underestimated exposure and high fecundity, with those who fell pregnant selected into marriage.

Our new variants of the OCM which use ratios of the proportions of women or men married in successive age groups generate substantially similar age-specific fertility rates, particularly for ages over 25, suggesting the calculation of exposure to marriage using proportions married is not particularly sensitive to marriage patterns. This insensitivity to marriage patterns is perhaps surprising, given the large differences in proportions married between men and women, but this can be explained by the fact that only five years' worth of exposure are being considered. The variant adjusts the number of married women of a particular age by the ratio of proportions of men married at the same age (adjusted for the spousal age gap) to the proportions married at 1, 2, 3, or 4 years of age older. These ratios will become increasingly different to the equivalent ratios for women as the number of years increases, but each of these years only contributes a small amount to the overall calculation. In general using male proportions married produces slightly lower estimates of fertility, particularly at young ages. These are likely to be underestimates because the male marriage curve tends to be slightly steeper than the female, and are exacerbated where there are larger differences in the marriage patterns of men and women, particularly among the higher social groups. Nevertheless we feel that it is reasonable to use ratios of male proportions married when calculating fertility for social groups. If desired, a standard female marriage schedule may be used instead, but our findings indicate that this increases the range of the discrepancies.

The TMFR is a very strange beast. Unlike many demographic measures, even those calculated from synthetic cohorts or cross-sectional surveys (such as singulate mean age at marriage), and unlike the TFR, it cannot be thought of as capturing the experience of a 'typical' woman. Instead it measures the fertility of a hypothetical women who married at a particular age. This is usually 15, but here we have used 20 or 25, and we argue that $TMFR_{25}$ is preferable for comparing sub-groups in England and Wales. Even in societies with relatively complete and early marriage for women, the TMFR often indicates implausibly high fertility (Hoem and Mureşan 2011) and this is particularly the case in societies with late marriage such as England and Wales. In 1911 the singulate mean ages at marriage were 26.2 for women and 27.6 for men, and only seven per cent of 20 year old women were already married. In this society where pre-marital sexual intercourse and conception were common, the ASMFRs for those who married in their teens and early twenties are inflated by selection of the already pregnant into marriage (see also Wrigley et al. 1997: 378-9). Rates for young married women are also quite susceptible to variations in proxy marriage patterns used when it is not possible to calculate proportions married, such as for social classes, and this constitutes another reason for recommending $TMFR_{25}$ as the preferred summary measure of marital fertility in historical populations.

In summary, we have presented two new variants of the OCM, which we have dubbed 'Variant D' and 'Variant E' for the calculation of age-specific marital fertility. These have two distinct advantages over the standard variants of the OCM for calculating marital fertility. Both variants allow relaxation of the assumption that all children are legitimate, and the second also allows the more accurate calculation of marital fertility for social groups when reported durations of marriage are unavailable. We have tested them extensively and demonstrated that they work well, despite some underestimation among late marrying social groups. There has been a recent and continuing growth in the availability of complete count census data for historic populations, opening possibilities for the comparison of age specific fertility (both within marriage and overall) for population sub-groups, and our new variants of the OCM will facilitate and improve such comparisons and will therefore produce more nuanced analyses of historic trends in fertility during the first demographic transition.

Notes

This work is based on I-CeM, a standardised, integrated dataset of most of the censuses of Great Britain for the period 1851 to 1911: K. Schürer and E. Higgs, *Integrated Census Microdata (I-CeM); 1851-1911* [computer file]. Colchester, Essex: UK Data Archive [distributor], April 2014. SN: 7481, http://dx.doi.org/10.5255/UKDA-SN-7481-1. A user guide and manual to the I-CeM data is available as E. Higgs, C. Jones, K. Schürer and A. Wilkinson, *The Integrated Census Microdata (I-CeM) Guide,* (Colchester, 2013). Further details on the I-CeM database together with a number of related resources are available from the I-CeM website at: https://www.essex.ac.uk/history/research/icem/. The creation of the I-CeM database was made possible through funding from the UK Economic and Social Research Council (ESRC), grant number RES-062-23-1629. The version of the I-CeM data used here has been enhanced as the result of work by K. Schürer, H. Jaadla, A. Reid and E. Garrett as part of the ESRC-funded An Atlas of Victorian Fertility Decline project (ES/L015463/1) at the Cambridge Group for the History of Population and Social Structure, Department of Geography, University of Cambridge, and will be deposited with the UK Data Archive at a future date. For further details, see http://www.geog.cam.ac.uk/research/projects/victorianfertilitydecline/.
Mortality data for England and Wales were obtained from the Human Mortality Database, http://www.mortality.org/, and from the Quarterly and Decennial Reports of the Registrar General of England and Wales.

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Table 1: Selected demographic characteristics of England and Wales, and three selected types of place, 1911.

	England and Wales	Professional	Mining	Textile
Sex Ratio (20–49 years)	0.92	0.69	1.15	0.86
Population Density	0.97	1.24	1.43	3.58
IMR	117.23	87.07	126.54	122.07
SMAM women	26.23	28.49	24.2	26.72
SMAM men	27.58	28.49	26.87	27.05
Population	36,070,492	3,222,409	3,597,665	3,270,122

Notes:

Population Density is the number of people per acre.

IMR is the Infant Mortality Rate, infant deaths per 1,000 births in the 5 years leading up to the census (April 1906–March 1911).

SMAM is the Singulate Mean Age at Marriage, calculated in the standard way from proportions married in the census.

Table 2: Children under age 5 at 1911 census: marital status and average age of mothers

marital status of mothers	number of children	average age of mothers
matched to a mother*		
single	5975	28.9
married with spouse present	3302019	32.3
married with spouse absent	202450	30.5
widowed	33891	36.7
divorced	23	33.9
unknown	764	34.1
not matched to a mother	242105	
estimated illegitimate children**	147804	

Note: *Excludes children of mothers reported to be aged less than 15 and more than 49.

**The number of illegitimate children present in 1911 has been estimated by taking the number of illegitimate births in England and Wales (ONS Birth Summary Tables) in each of the years 1906 to 1910 and surviving them on, year by year, to 1911, using mortality rates given in the Human Mortality Database, with infant mortality multiplied by 1.5 to allow for higher mortality among illegitimate infants.

Table 3: Total Marital Fertility Rates for women marrying at age 20 and age 25, England and Wales 1911, calculated using different variants of the OCM

	$TMFR_{20}$	$TMFR_{25}$
variant A	7.09	4.16
variant B	5.89	3.64
variant C	4.65	3.40
variant D	6.17	3.68
variant E	5.93	3.65

Note: See text for details of variants

Table 4: Total Marital Fertility Rates for women marrying at age 20 and age 25, types of place in England and Wales 1911, calculated using different variants

$TMFR_{20}$	mining	professional	textile	E&W
variant A	7.47	6.31	6.12	7.09
variant B	6.88	5.57	5.42	5.89
variant C	4.80	2.25	2.98	4.65
variant D	6.77	5.79	5.41	6.17
variant E	6.82	5.42	5.47	5.93
variant E.std	7.35	5.33	5.20	5.93
$TMFR_{25}$	mining	professional	textile	E&W
variant A	4.67	3.60	3.46	6.26
variant B	4.38	3.29	3.17	3.64
variant C	3.37	1.76	2.27	3.40
variant D	4.26	3.32	3.11	3.68
variant E	4.34	3.22	3.10	3.65
variant E.std	4.40	3.17	3.09	3.65

Note: See text for details of variants

Table 5: Total Marital Fertility Rates for women marrying at age 20 and age 25, social classes in England and Wales 1911, calculated using different variants of the OCM

$TMFR_{20}$	**Variant B**	**Variant E**	**Variant E.std**
SC1	5.33	4.41	4.41
SC2	5.81	4.95	5.36
SC3	6.21	5.82	5.99
SC4	6.30	5.89	6.11
SC5	6.85	6.48	7.11
Textile workers	5.43	4.83	5.15
Miners	7.51	7.13	7.95
Agricultural Labourers	7.09	6.34	6.98
$TMFR_{25}$	**Variant B**	**Variant E**	**Variant E.std**
SC1	3.05	2.77	2.77
SC2	3.47	3.19	3.24
SC3	3.75	3.56	3.60
SC4	3.84	3.65	3.66
SC5	4.38	4.17	4.24
Textile workers	3.17	2.99	3.03
Miners	4.91	4.70	4.83
Agricultural Labourers	4.47	4.20	4.25

Note: See text for details of variants

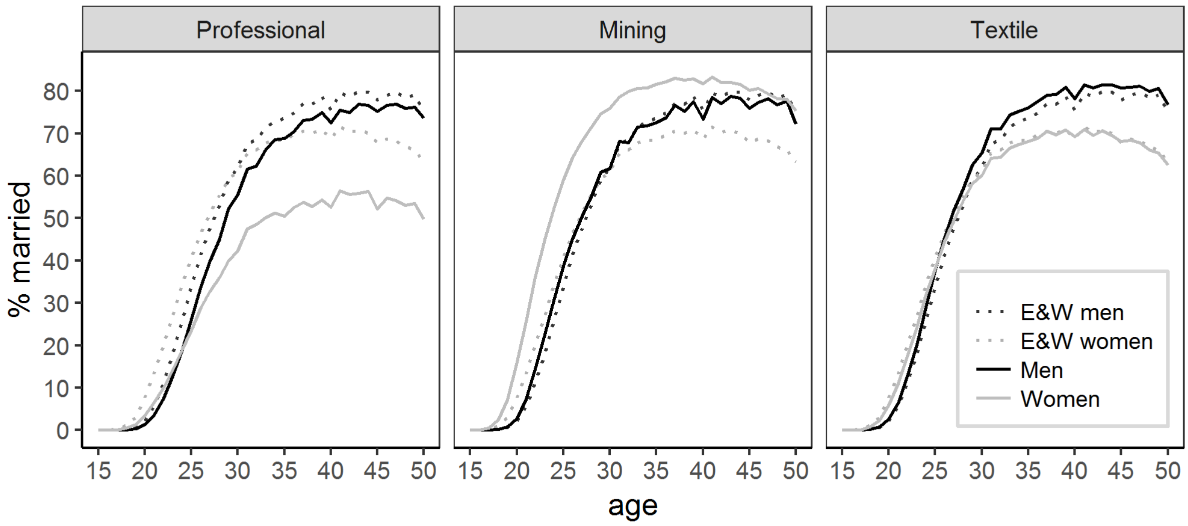


Figure 1: Percentages of men and women who were married and whose spouse was present in the same household in three different types of place, England and Wales, 1911.

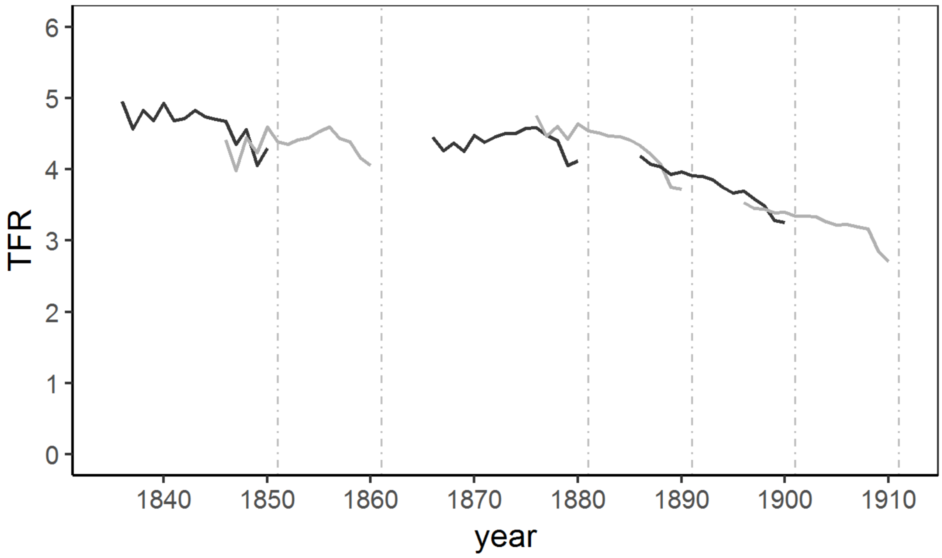


Figure 2: Total Fertility Rates (TFRs), England and Wales, 1836–1911, calculated using the OCM

Note: Vertical lines indicate the years of the censuses used to calculate fertility.

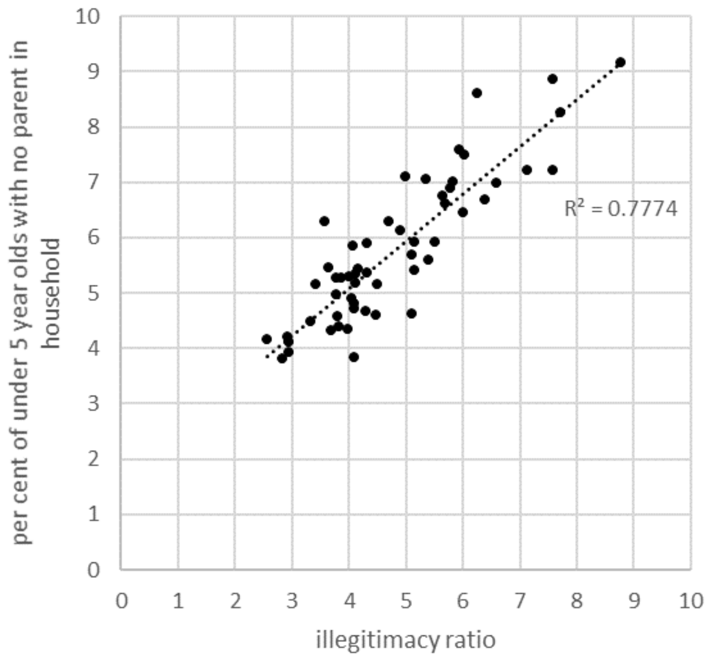


Figure 3: Relationship between percentage of children under 5 years old with no parent in the household (1911) and the illegitimacy ratio (1910), for the counties of England and Wales.

Note: the illegitimacy ratio is the number of illegitimate births as a percentage of all births, and is derived from the Registrar General's Annual Report for 1910.

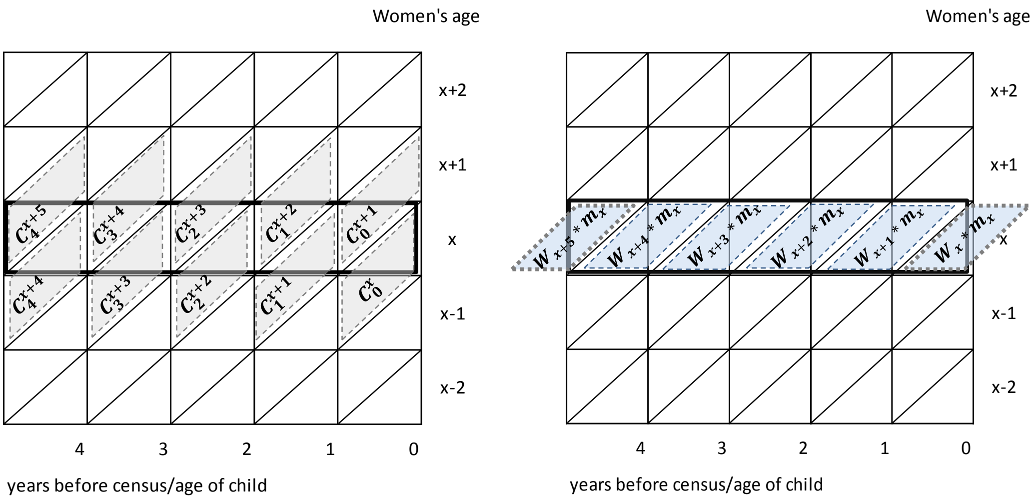


Figure 4: Lexis diagrams to illustrate numerators (left) and denominators (right) for the calculation of fertility

Note: See text and equations for explanation of terms.

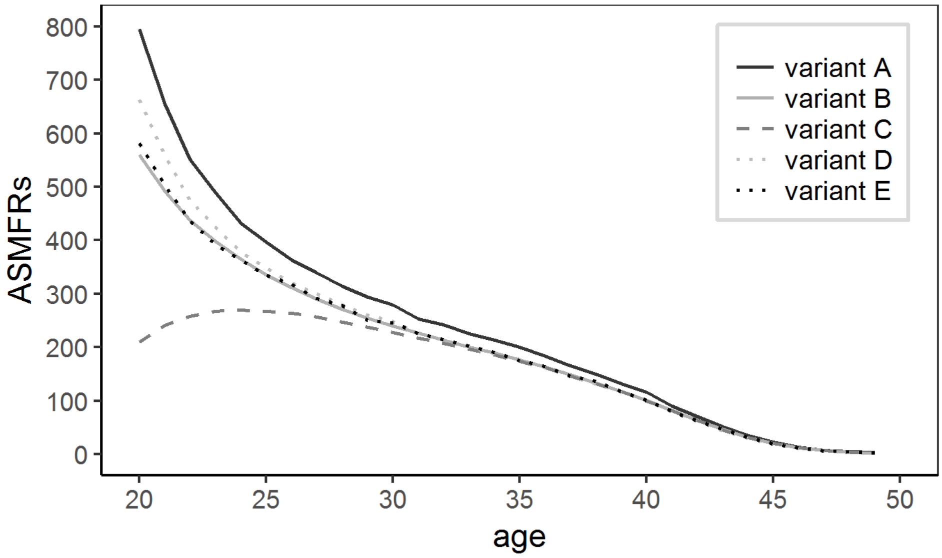


Figure 5: Age-specific marital fertility rates (ASMFRs) calculated using different variants of the OCM, England and Wales, 1911

Note: See text for details of variants.

Figure 6: Number of couples by duration of marriage (in years), England and Wales, 1911

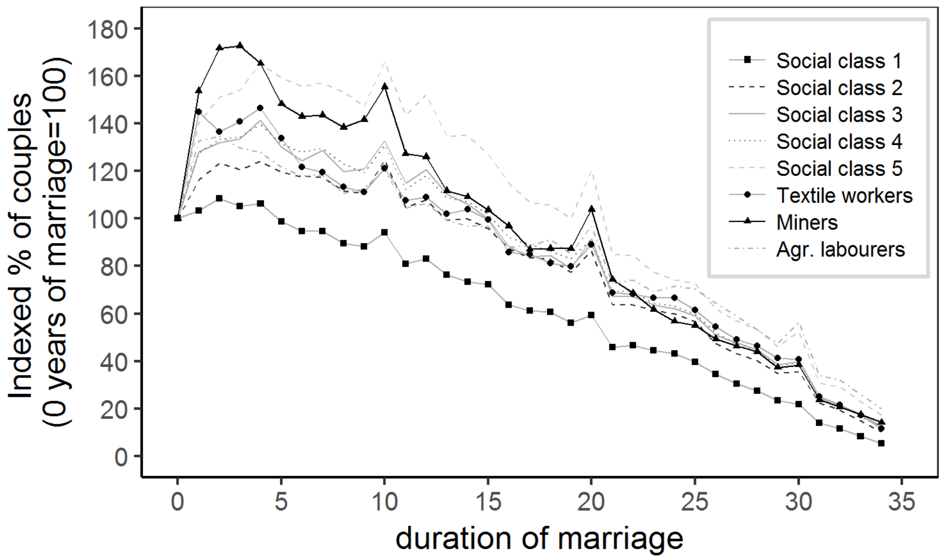


Figure 7: Distributions of couples by duration of marriage (in years), indexed at 0 years married, by social class, England and Wales, 1911

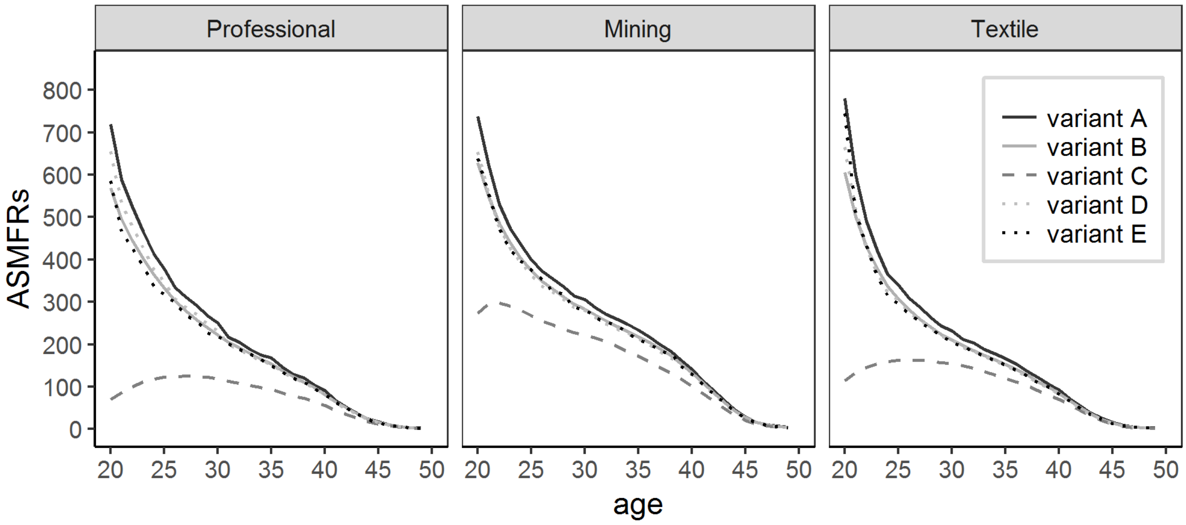


Figure 8: Age-specific marital fertility rates (ASMFRs) calculated using different variants of the OCM, professional, mining and textile places within England and Wales, 1911

Note: See text for details of variants.

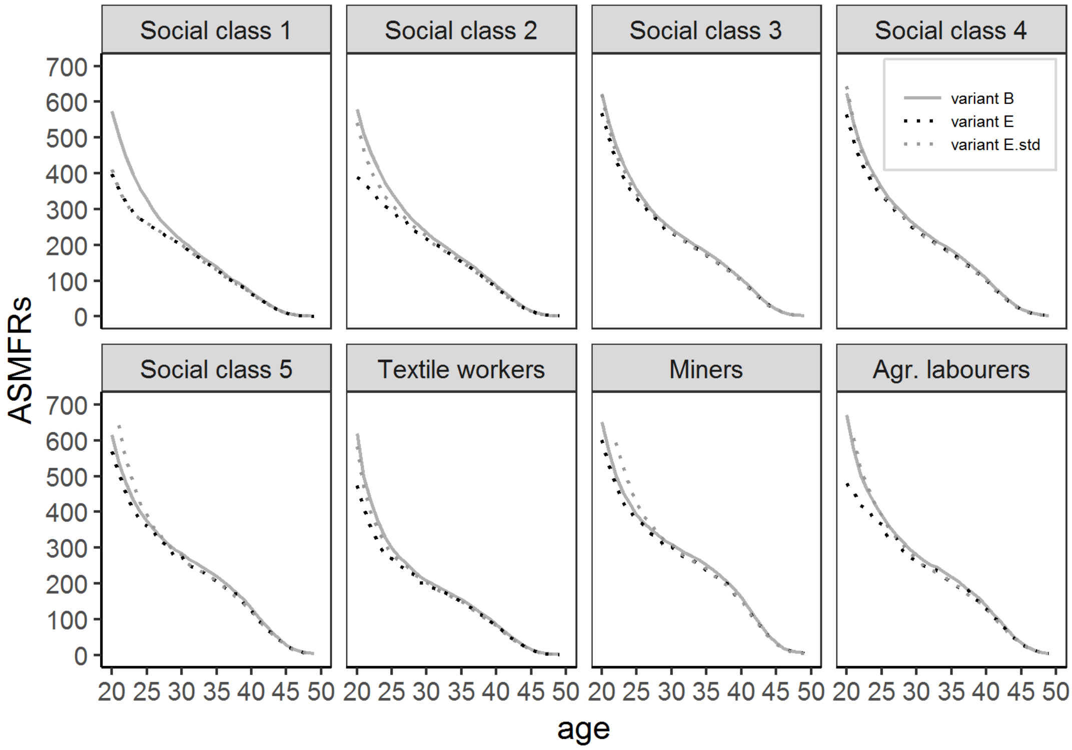


Figure 9: Age-specific marital fertility rates (ASMFRs) calculated using three different OCM variants, for different social classes, England and Wales, 1911.

Online Supplementary Material to

Adapting the 'own children method' to allow comparison of fertility between populations with different marriage regimes.

Alice Reid, Hannaliis Jaadla, Eilidh Garrett, Kevin Schürer.

Population Studies 2019

Appendix 1: Age distribution adjustments for own child method

Under-enumeration and age-misstatement, particularly age-heaping, are often highlighted when dealing with historical census data. The accuracy of enumeration varies across different age groups: young children are often under-enumerated and there is tendency to report children's ages in even numbers (Myers 1993). Age-misstatement in older ages is often due to heaping on ages ending in zero or five or the exaggeration of age by the elderly. Women in their teens or thirties are often reported to be in their twenties (Lee and Lam, 1983).

Under-enumeration of children and age-misstatement among both women and children affect the estimation of fertility using the OCM. Previous studies have highlighted this as especially problematic when calculating yearly fertility rates (Guilmoto and Rajan 2002). Rindfuss (1976) notes that if the ages of children are misreported or under-enumerated, then all the estimates of fertility for some years will be artificially inflated and all the estimates for other years will be artificially depressed; and if the ages of women are mis-stated, then some age-specific estimates for every year will be over- or underestimated. A common way of minimising errors from age misreporting is to estimate age-specific rates in five-year age groups over one or more calendar years (Retherford and Cho 1978).

Previous studies considering fertility decline using historical census data are not always very clear whether and, if so, how they have dealt with deficiencies in census coverage and underreporting. One exception is Hacker (2000, 2013) who estimated age- and sex-specific net census under-enumeration of the native-born white population in the United States for the census years 1850–70 and revised existing estimates for the census years 1880–1930. His method has four main assumptions: i) there was negligible out-migration of native-born whites from the US, ii) age was accurately reported in the census, iii) yearly estimates of mortality were reliable and iv) white women aged 15–29 experienced a known and unchanging net undercount. The age assumption is the most problematic, in order to adjust for age-heaping he used an approximately linear trend line based on Zelnik's work (1961). He argues that researchers relying on nineteenth-century censuses for fertility estimation need to apply large correction factors to properly estimate fertility levels (Hacker used adjustments in his OCM calculations: 2003, 2016, and Haines and Hacker 2011).

To investigate the effect of age-misstatement and under-enumeration in the censuses of England and Wales, 1851-1911, we used two approaches: age-smoothing and age adjustment coefficients.

Age smoothing

In line with King's graduation of ages (1913), the smoothing of the 1911 census data allowed us to measure how age-heaping influences our fertility estimation. We used local polynomial regression to smooth the 1911 age distribution of the whole population aged between zero and 70, which fits well with King's graduated ages, and the resulting numbers of observed and adjusted children in 1911 are shown in Table A1.1.

Table A1.1: Observed and adjusted (based on smoothed ages) number of children in England and Wales, 1911.

	Age of child					
	0	1	2	3	4	Total
Observed	758255	740129	791288	779117	768230	3837019
Adjusted	774569	769330	763997	758577	753080	3819554
Adjustment factor	1.0215	1.0394	0.9655	0.9736	0.9803	

The TFR calculated using the smoothed ages was 2.97, only 0.03 lower than unadjusted one. The predicted values of the local polynomial regression depend on the length of the age distribution used, but the changes in TFR are still minor even if only the age distribution between zero and 15 is smoothed. We conclude that our estimated fertility rates, when aggregated across five-year age groups, are not very sensitive to age-misreporting or age-shifting amongst children under five.

Age-adjustment coefficients

It is possible that a proportion of infants were omitted from the census returns entirely. To investigate whether this was the case, we used age-adjustment coefficients which had previously been estimated for historic English censuses 1821-1931 by Lee and Lam (1983), who modified Demeny and Shorter's (1968) approach to the correction of distortions in age distributions. Lee and Lam's intercensal mortality assumptions were based on mortality patterns taken from Coale and Demeny's North and West model life tables. Overall they found that CD-West mortality fitted the census survival ratios better, particularly for the younger age groups.

We used both North and West adjustment coefficients to measure the combined effect of under-enumeration and age-misstatement on age-specific fertility rates and total fertility rates. As they only present adjustment factors for females, we assumed that for male and female children (age group 0–4) had similar levels of under-enumeration. As there are two sets of age adjustment coefficients for most censuses, we used an average estimate of the two, except for 1851 because Lee and Lam deemed adjustment factors for 1841–1851 unreliable due to errors in the enumeration of the 1841 census. Table A1.2 shows that the TFR estimated with Model West age adjustment factors differs only slightly from the unadjusted TFR.

Table A1.2: Total fertility rates, England and Wales, 1851–1911, with adjustment for age mis-reporting following Lee and Lam's age adjustment coefficients.

	1851	1861	1881	1891	1901	1911
TFR	4.3	4.3	4.4	4.0	3.5	3.0
TFR West	4.4	4.4	4.6	4.3	3.7	3.1
TFR North	4.6	4.6	4.9	4.6	4.0	3.3

These adjustments have a drawback in that they are only provided for five-year age groups so do not allow specific treatment of the different patterns among children under the age of five. It is also important to note that Lee and Lam argue that the results for the age group 0–4 seem implausible. We definitely see that the differences between adjusted and unadjusted TFR are largest for the years 1881, 1891 and 1901. Under-enumeration among children seems to increase towards the end of the century, suggesting that the English and Welsh child mortality patterns may not have conformed to either the North or West model pattern (see Jaadla and Reid 2017 for more information on age patterns of child mortality). The extent of the effect of under-enumeration is therefore unclear, but it is likely to have only a minimal effect on the fertility rates estimated using the OCM.

Appendix 2: Testing mortality assumptions

To assess the possible bias in our fertility estimates introduced by inappropriate mortality assumptions we examined the sensitivity of reverse survival methods both at the national-level and by different types of place using two different approaches. The first of our approaches tested the effect of not adjusting for child mortality. The second examined the effect of assuming lower or higher child mortality levels represented by national life table estimates where life expectancies were five years lower or higher than those of the census year which is used to produce the fertility estimates. For different types of place in England and Wales we then calculated fertility using mortality rates which were ten per cent higher and lower mortality than the actual rates. Here we present all these tests for Total Marital Fertility, where we adjust only for child mortality, but the results for TFR, where we also adjust for adult female mortality, are very similar.

No mortality adjustments:

The fertility estimates produced without any mortality adjustments at the national-level (TMFR^) are, as expected, substantially underestimated in comparison to those where we adjust using national mortality figures (TMFR) (Table A2.1). The TMFR^ differs from the TMFR by more than 15 per cent for every census year shown in the table. This margin is equivalent, on average, to more than one child per woman.

Table A2.1: TMFRs with and without child mortality assumptions

Year	*TMFR*	*TMFR^*	*Percentage discrepancy*
1881	7.23	6.06	16.2
1891	7.07	5.76	18.5
1901	6.41	5.29	17.5
1911	6.00	5.07	15.5

Note: *TMFR* calculated using mortality estimates for every census year from Human Mortality Database (HMD). *TMFR^* calculated without any mortality assumptions

Figure A2.1 illustrates the underestimation of fertility as a result of not adjusting for child mortality, as percentages of the TMFR calculated with mortality adjustment. This shows that the failure to adjust for mortality has a varied effect on fertility across these different types of place. As expected the under-estimation of fertility is the largest for the RSDs grouped into the transport, textile and mining categories which, in general, had higher levels of child mortality. In contrast, the estimated TMFRs and TMFR^s differ least in agricultural areas which were generally healthy. The substantial decline observed in 1911 for all places is mainly driven by overall secular decline in child mortality rates and the onset of infant mortality decline in the early years of the new century.

Figure A2.1: The percentage underestimation of TMFR as a result of not adjusting for child mortality, by type of place, England and Wales 1881-1911.

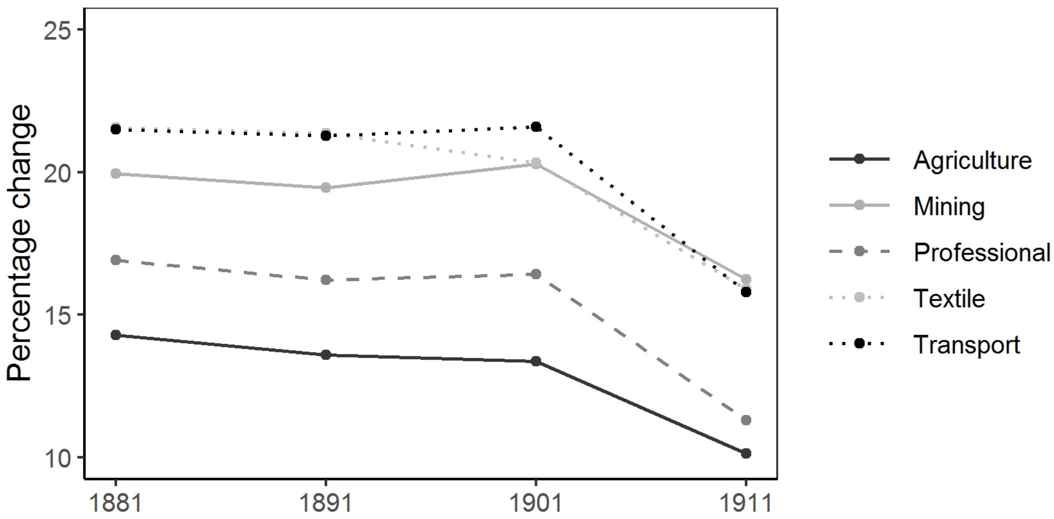
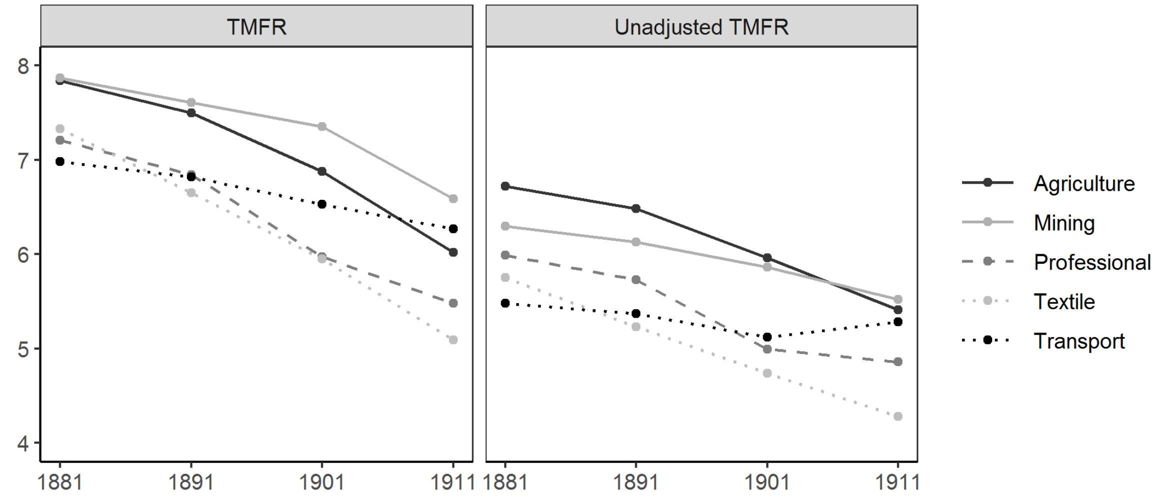


Figure A2.2 demonstrates how, compared to agricultural places, the relative positions of fertility levels for different types of place change according to whether or not a mortality adjustment is made. This indicates the possible discrepancies that adjusted and unadjusted mortality rates introduce when examining fertility differences by typology. For example, TMFRs (with mortality adjustments) for mining districts (shown with solid grey lines in the left hand panel of Figure A2.2) are always higher than the estimates for agricultural districts. The unadjusted rates (solid grey lines in the right hand panel of Figure A2.2), as expected, consistently underestimate fertility levels in the mining districts relative to agricultural areas prior to 1911, obscuring the true fertility differences. This is true for all types of places with high mortality. Holding mortality constant across sub-populations or different regions would produce similar distortions in the relative differences between groups.

Figure A2.2: Relative differences in fertility by typology (compared to agriculture) with and without mortality adjustments



Note: The left hand panel shows the differences between TMFRs with mortality adjustment. The right hand panel shows the differences between TMFR^s (without mortality adjustment).

Lower and higher mortality

At the national level the effect of using an incorrect level of mortality for adjustments has only a small impact on fertility estimates (Table A2.2). When estimates using life tables with levels of life expectancy five years lower or higher than the levels observed in the actual census year are used, the difference reaches around five per cent.

To examine the potential errors introduced by adopting inaccurate mortality assumptions for different types of places in England and Wales we took a slightly different approach, comparing TMFRs calculated with the observed mortality adjusted for with TMFR^s calculated with mortality ten per cent higher or lower than that observed. This resulted in average TMFR differences of around five per cent when mortality was overestimated by ten per cent, and six to seven per cent when mortality was underestimated by ten percent (table not shown).

Table A2.2: TMFRs and TFRs under different mortality adjustments

Census	*Year of Life Table*	e_0	*TMFR*	*TFR*
1881	1871	41.31	7.56	4.58
1881	1881	46.11	7.23	4.40
1881	1905	50.86	7.13	4.38
1891	1854	39.34	7.32	4.12
1891	1891	44.29	7.07	4.01
1891	1902	49.08	6.85	3.91
1901	1878	43.01	6.58	3.54
1901	1901	47.91	6.41	3.48
1901	1909	52.52	6.08	3.30
1911	1900	46.32	6.21	3.11
1911	1911	51.54	6.00	3.02
1911	1920	56.48	5.68	2.86

Note: Life tables are from HMD.

Appendix 3: Explaining the formula for calculating number of women-years of exposure from durations married (variant B)

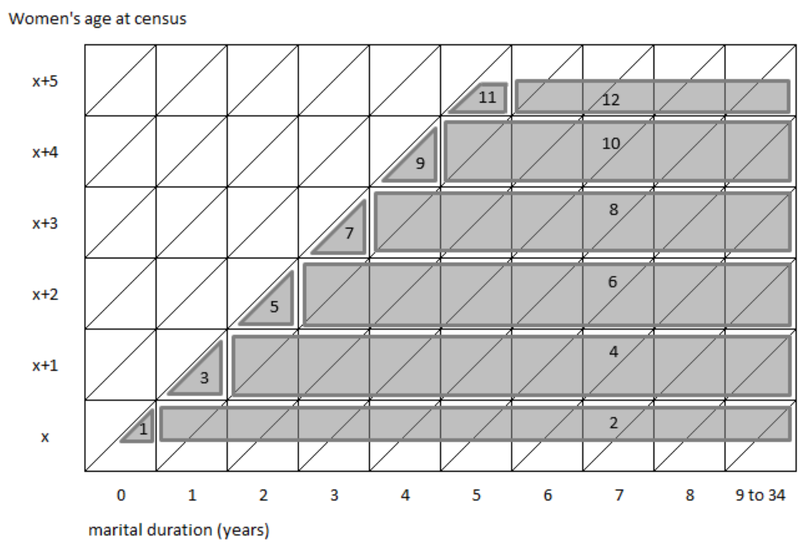
We have used the following formula to calculate the number of woman-years of exposure to women aged x during the five years leading up to the census:

$$N_x = \frac{M_x^0}{8} + \frac{M_x^{1-34}}{2} + \frac{M_{x+1}^1}{2} + M_{x+1}^{2-34} + \frac{M_{x+2}^2}{2} + M_{x+2}^{3-34} + \frac{M_{x+3}^3}{2} + M_{x+3}^{4-34} + \frac{M_{x+4}^4}{2} + M_{x+4}^{5-34} + \frac{3M_{x+5}^5}{8} + \frac{M_{x+5}^{6-34}}{2}$$

where M_x^i is the number of married women age x with duration of marriage i.

This is illustrated in Figure A3.1, in which the shaded shape labelled 1 represents the first term of the formula, the shape labelled 2 represents the second term, and so on. Only half the exposure for women aged x is used, as these women spend on average half their time aged x-1. Similarly, only half the exposure for women aged x+5 is used, as these women on average spent half their time aged x+6. Only half the exposure is used for women who were age x when they married (i.e. squares coinciding age x and marital duration 0, age x+1 and marital duration 1, and so on) as these women on average only spent half of that year married.

Figure A3.1 Lexis diagrams to explain formula for calculating the number of women years of exposure for variant B.



Appendix 4: Marital fertility eliminating servants from proportions married

This Appendix shows a comparison with a version of variant A where we exclude unmarried servants from the exposure calculations on the basis that they are temporary and not participating in the marriage market (following Kurosu 2003 and Breschi, Derosas and Rettaroli 2003). Variant A excluding servants from the exposure calculation shows a considerable reduction of fertility in professional areas: proportions married at each age are higher, exposure is longer and fertility lower. If servants really were not involved in the marriage markets in places they were living, then this might be a reasonable adjustment. However comparisons with variant B shows that removing servants from the calculations reduces fertility too much, both in professional areas where there were many servants, but also in other places and in England and Wales overall. Some servants did stay and get married in the places they were working, and it seems that including them in the calculation of marital duration using proportions married does not overly distort fertility rates.

Table A4.1: Total Marital Fertility Rates for women marrying at age 20 and age 25, types of place in England and Wales 1911, variant A and variant A excluding servants

$TMFR_{20}$	mining	professional	textile	E&W
variant A	7.47	6.31	6.12	7.09
variant A excluding servants	7.24	4.85	6.05	4.16
$TMFR_{25}$				
variant A	4.67	3.60	3.46	6.26
variant A excluding servants	4.67	3.07	3.45	3.87

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