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Fertility postponement is largely due to rising educational enrolment

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The rise in educational enrolment is often cited as a possible cause of the trend to later childbearing in developed societies but direct evidence of its contribution to the aggregate change in fertility tempo is scarce. We show that rising enrolment, resulting in later ages at the end of education, accounts for a substantial part of the upward shift in the mean age at first birth in the 1980s and 1990s in Britain and in France. The postponement of first birth over that period has two components: a longer average period of enrolment and a post-enrolment component that is also related to educational level. The relationship between rising educational participation and the move to later fertility timing is almost certainly causal. Our findings therefore suggest that fertility tempo change is rooted in macro-economic and structural forces rather than in the cultural domain.

Keywords: fertility tempo; fertility postponement; delayed childbearing; aggregate change; educational expansion; educational enrolment; compulsory school-leaving age; post-compulsory education; fertility trends

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Background

While a later start of childbearing is a prominent feature of recent fertility trends in developed countries, a clear explanation for this systematic and pervasive change is still lacking. The growth in educational participation is often mentioned as a potential cause, but explicit evidence of its role as a driver of timing change is sparse. We show in this paper that the rise in educational enrolment has been a substantial contributor to the trend to later childbearing in Britain and in France, by providing evidence that the lengthening of time spent in education accounts for a sizeable part of the aggregate change in the timing of first birth from the early 1980s to the late 1990s. The study we report breaks new ground in two respects: it focuses on the age at completing education rather than on educational attainment, and we make a quantitative assessment of the empirical contribution of rising enrolment to aggregate change in fertility tempo.

Although rising education is probably the most frequently cited candidate cause, a wide array of other influences on the trend to later childbearing have been proposed: increases in women's participation in the labour force; the increasing importance they attach to a career; their growing economic independence; the rising opportunity costs of childbearing; the cost and availability of childcare; globalization; labour market uncertainty; differential change in living standards; housing factors; near-perfect control over fertility through contraception and abortion; instability of unions; social interaction processes and institutional factors, together with cultural factors central to the concept of second demographic transition such as secularization, a retreat from family values, individualization and value change, rising aspirations, 'post-materialist' values, and a retreat from personal commitment (Lesthaeghe and Willems 1999; Martin 2000; Lesthaeghe 2001; Santow and Bracher 2001; Kohler et al. 2002; Mills et al. 2005, 2011; Billari et al. 2006; Toulemon et al. 2008; Neels and De Wachter 2010). Several of these factors are, of course, related to, and possibly the result of, the growth in educational participation. However, as in the case of enrolment, little quantitative evidence is available on the contribution of these potential causes to macro-level trends in fertility tempo.

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Studies going back several decades have demonstrated a clear association at the micro level between education, measured in various ways, and the frequency and timing of transitions to adulthood, particularly marriage and first birth. Early analyses used regression in various forms with age at first birth, an explicit measure of timing, as dependent variable. Education as the independent variable was usually represented by attainment, for example, years of schooling or highest qualification achieved. These studies generally found an inverse relationship between educational attainment and the timing of first birth (Rindfuss et al. 1980; Rindfuss and St John 1983; Marini 1984, 1985); they also demonstrated the strong relationship between activity status, including educational participation, and the occurrence of first birth (Rindfuss et al. 1988, chap. 8). The advent of survival analysis brought a re-specification of both dependent and independent variables. In the event-history approach, the dependent variable represents both the occurrence and pace of family transitions and is therefore a measure of level and timing combined. Education as the independent variable is transformed in many of these analyses into two covariates: time-varying enrolment and educational attainment. As a predictor, enrolment is by far the most consistent in its effects, having a strong negative association with rates of transition to marriage, to cohabitation, and to first birth. By contrast, the net effect of educational attainment is less clear when enrolment is present in the model (Lappegard and Ronsen 2005): in some studies, transition rates are significantly associated with educational level, either positively or negatively, net of enrolment (Kravdal 1994; Oppenheimer et al. 1995; Pinnelli and De Rose 1995; Robert and Blossfeld 1995; Liefbroer and Corijn 1999; Santow and Bracher 2001; Gustafsson et al. 2002; Winkler-Dworak and Toulemon 2007) but in others such a net effect is absent (Hoem 1986; Blossfeld and Huinink 1991, Table 1; Blossfeld and Jaenichen 1992; Blossfeld and Rohwer 1995).

Although the inference is often made from the individual-level relationship between enrolment and the level and timing of first birth that educational expansion has resulted in later first births, it is an inference that goes beyond the evidence. Where a variable is a significant micro-level predictor of an outcome variable, aggregate change in the outcome may be attributable either to compositional change in respect of the predictor variable, or to change over time in the coefficient of the predictor, or to some combination of the two factors (Wooldridge 2008). Most studies that comment explicitly on the link between aggregate change in education and birth timing have relied for evidence on the net effects of period or cohort in micro-level models of the occurrence of first birth. These provide useful but partial evidence on the causes of aggregate change in tempo. In such analyses, the estimated effects of calendar time—period or cohort—on level and timing of first birth tend to be modified but retain statistical significance when a full range of covariates, including education variables, is included, though cohort effects become non-significant in a few cases (Blossfeld and Huinink 1991; Blossfeld and Jaenichen 1992; Kravdal 1994; Oppenheimer et al. 1995; Pinnelli and De Rose 1995; Santow and Bracher 2001; Billari and Kohler 2002; Gustafsson et al. 2002). Results for the net effect of calendar time are, thus, not consistent. Furthermore, it is unclear from these analyses, as reported, how far it is specifically the education variables—enrolment or attainment, or both—that reduce or remove the calendar-time effects.

Several studies have used standardization to address directly the question how far changing educational composition has contributed to later family formation. Among these, the evidence of Oppenheimer et al. (1995) for the period 1965–90, Rindfuss et al. (1996) for the 1970s and 1980s, and Rendall et al. (2010) covering cohorts of the 1950s and 1960s, did not support the proposition that rising educational attainment explains the trend to later family transitions. By contrast, recent analyses of Belgian census data from both a period and a cohort perspective (Neels 2009; Neels and De Wachter 2010) found that compositional effects could explain a significant part of fertility change; their findings are in accord with those reported in the study presented here.

Skirbekk et al. (2004) carried out a further type of investigation, exploiting a natural experiment based on birth month and school-leaving regulations. They demonstrated that school-leaving age has a causal effect on the timing of demographic events in Sweden. They also showed that in Sweden in the period 1975–2000, the average years spent in education by women at age 45 had risen by about a year less than the increase in their mean age at first birth. Several econometric analyses using instrumental variables have also presented evidence of a causal link between compulsory school-leaving age and later first-birth timing (Black et al. 2008; Monstad et al. 2008; Silles 2011). However, these studies were primarily concerned with establishing causation, and none of them examined the empirical contribution of changing school-leaving age to observed trends in the tempo of demographic events.

The study presented in this paper quantifies the contribution of the growth in educational participation to the rising age at first birth. Our study adds to existing work in several ways. We adopt a period approach in contrast to the cohort perspective of most previous studies in this area (but see Kravdal 1994; Oppenheimer et al. 1995; Rindfuss et al. 1996; Neels 2009). While previous related investigations have employed indicators that reflect both level and tempo of fertility, we measure period timing explicitly, net of period differences in level. We focus on terminal education age, rather than educational attainment. We also examine first-birth rates by time since leaving education, and how these schedules have changed over time. Finally, we address explicitly, and quantify, the contribution made by time trends in the age at completing education to aggregate change in fertility tempo.

Data and methods

We adopted a period approach and confined analysis to the first birth. This was because in both Britain and France later childbearing in recent decades was due entirely to later first births. For example, in England and Wales the age-standardized period mean age at childbearing rose by 2.5 years, from 26.4 to 28.9 between 1974 and 2004. The age-standardized period mean age at first birth increased over the same period by 2.9 years, from 24.2 to 27.1, and therefore accounts for all of the change in birth timing (Office for National Statistics 2011, Table 4b). The change of timing in France is also attributable to a later start of childbearing. The overall age-standardized period mean age at birth in France was 26.7 in 1975 and 29.4 in 2000, an increase of 2.7 years; the age-standardized period mean age at first birth rose by 3.3 years, from 24.1 to 27.4 over the same period (Toulemon and Mazuy 2001; Prioux and Mazuy 2009, Table 4).

We employed two sources of data. For Britain, we used a pooled series of General Household Survey (GHS) rounds from 2000 to 2007—a subset of a larger time-series data-file of GHS surveys from 1979 to 2007 compiled by the Centre for Population Change. The GHS is an annual continuous survey of the population in private households in Great Britain, carried out by the Office for National Statistics. It collects a range of socio-economic information on household members, including the fertility and union histories of women aged 16–59. The original fertility histories in recent rounds of the GHS were biased by excess reporting of childlessness among older women (Murphy 2009). The biases have, however, been substantially corrected. Period fertility estimates based on the revised histories approximate well to national vital registration sources (Ní Bhrolcháin et al. 2011). Two issues of data quality limited the range of available data that could be used in our study. First, the variable representing age at leaving education was found to be defective in GHS rounds before 2000; the data analysed were therefore confined to survey rounds 2000–07. Second, although our correction of the GHS fertility histories was fairly successful, it did not remove all errors. One issue that remains is that the estimated change in fertility timing is biased, relative to vital registration, when the period 2000 onwards is the endpoint. However, when analysis is confined to the period 1980–84 to 1995–99 the estimated timing change is unbiased, assessed against vital registration (see Appendix). Hence, we restricted our study to the period 1980–84 to 1995–99.

All estimates were weighted by a set of weights constructed specifically for use with the Family Information section of the GHS (Beaujouan et al. 2011). For France, we used the Family History Survey (FHS) linked with the French census of 1999 (Cassan et al. 2000). Estimates of period fertility from the FHS have been found to correspond well with vital registration (Mazuy and Toulemon 2001). Because fertility histories are available in the GHS for women only, men were omitted from the analysis.

Period indicators of timing

Three aggregate measures of timing were used for each quinquennial period 1980–84 to 1995–99:

(1) The mean age at first birth estimated from a period life table, in which elapsed time was indexed by years since the 13th birthday, and births were confined to those occurring by exact age 40; this therefore represents the period mean age among those experiencing the event by the 40th birthday in the life table; 13 was chosen as an initial age to accommodate some early school leavers in the French sample and to allow for some very early births.
(2) The mean age at leaving continuous education, estimated from a period life table, again indexing time by years since 13th birthday, and again

confined to those experiencing the event by the 40th birthday.

(3) The mean duration from the end of continuous education to first birth, also derived from a period life table; the period mean duration to first birth was estimated for those experiencing a first birth by the 20th anniversary of leaving continuous education (for further details, see Appendix).

Each of these measures is independent of the distribution of exposure to the event in question, whether by age or by duration, in each period. The period covered extended back to 1980–84. Estimates could not be obtained in the GHS for earlier periods because the oldest women in the GHS sample were aged 59 in 2000–01, which means that 1980 is the earliest year for which exposure and events to age 39 are observed in the data-set. Since there is no age limit on the FHS sample, estimates could, in principle, be made for France for earlier periods. However, for consistency between the two data sources, and ease of comparison, the French data used were restricted to the period 1980–84 onwards. Person years at risk at ages 13–39 by period in the two data-sets are given in Table 1.

Measures of educational participation

In both the GHS and the FHS, a question was asked about the age at which the respondent finished her continuous education and this age was coded in completed years. It reflects the first age at which a person ceased to be in education or training, and not the age at which people later returning to education ultimately finished. We used this information to infer whether a respondent was in education at specified ages and dates. Estimates of age-specific participation in the GHS were based on denominators in August and numerators in December of the calendar year in question, to correspond with the methods used in official English education statistics.

Table 1 Person years at risk at ages 13–39. Britain and France 1980–84 to 1995–99

	Period			
	1980–84	1985–89	1990–94	1995–99
Britain	104,039	112,958	112,042	110,963
France	467,043	486,814	475,183	354,083

Source: Centre for Population Change GHS time-series data-file and Family History Survey associated with the French census of 1999.

Estimates derived in this way are intermediate between official estimates of the proportions in education and the proportions in education or training in England (Department for Education 2011 and associated historical data, together with Labour Force Survey estimates for those aged 19+ provided by the Department for Education). For France, the FHS estimates correspond well, in both levels and trends, to official sources in education at ages under 20, but are somewhat higher than official estimates at ages above 20 (Durier and Poulet-Coulibando 2004). The FHS figures at ages 21+ are likely to reflect participation in training as well as in education. In both countries, the participation figures thus reflect training and forms of apprenticeship as well as education. However, for economy we refer throughout to participation in education.

Results

Change over time in first-birth timing

Figure 1(a) and (b) show the period schedules of age-specific first-birth rates for each quinquennial period 1980–84 to 1995–99. The schedules were standardized to sum to 1.0 in each period to reveal the shifting age profile of first birth independently of changes in level. Both graphs display the steady shift to a later age at first birth that is characteristic of the fertility schedules of developed countries in recent decades. The displacement is more pronounced for France (see also Toulemon 2005; Toulemon et al. 2008), and so the age profiles of first-birth propensity show distinctive trends in each country. The French schedules are of the familiar unimodal form, and move in successive periods up the age range. By contrast, the British first-birth schedules are static at ages under 20 and shift rightwards at older ages, thus showing signs of the bimodal pattern that emerged in English-speaking countries in recent decades (Chandola et al. 1999, 2002; Rendall et al. 2005; Sullivan 2005).

Table 2 summarizes these trends. Over the two decades, the mean age at first birth in Britain is estimated from the GHS to have risen by 1.4 years, from 25.5 to 26.9, while the increase in France was more substantial, at 2.4 years. We see in Figure 1(a) and (b) that in both countries, variability in the age at first birth increased—that is, women were becoming mothers at a wider range of ages than in the recent past. However, the dispersion in the first-birth schedule increased more in Britain than in France:

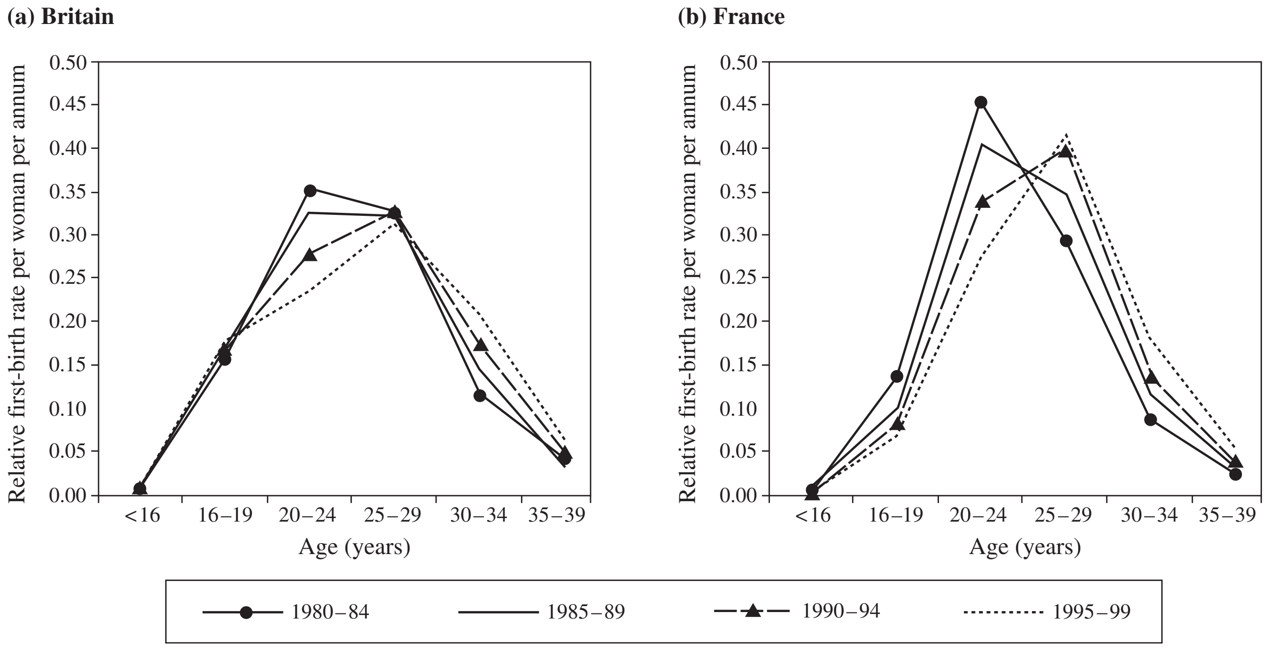

Figure 1 Age-specific first-birth rates by period (standardized to sum to 1.0 in each period). Britain and France 1980–84 to 1995–99
Source: As for Table 1.

the interquartile range in the age at first birth in the life table, among those having a first birth by age 40, rose by over a quarter in Britain, from 6.8 to 8.6 years, and by just under 10 per cent, from 5.6 to 6.1 years, in France.

Change over time in educational participation

Educational participation has increased substantially in both Britain and France in recent decades. This is shown in Figure 2(a) and (b), which plot time trends in age-specific participation rates in each country (the time span is chosen to correspond to the period for which fertility information is available). Educational participation among 17-year-olds increased from 47 per cent in 1980–84 to 69 per cent in 1995–99 in Britain, and among 20-year-olds from 17 to 39 per cent over the same period. In France, enrolment at each age is substantially higher but sharp upward trends are also seen: 80 per cent of 17-year-olds were in education in 1975–79, rising to 90 per cent in 1995–99, and participation among 20-year-olds more than doubled from 30 to 65 per cent over the period.

The result of these upward trends in enrolment rates is a substantial shift in the distribution of the age at leaving education (Table 3). In Britain between 1980–84 and 1995–99 the proportions leaving education at ages 13–17 almost halved, falling from 65 to 37 per cent, while the proportion leaving at ages 18–20 rose from 21 to 25 per cent, and the proportion leaving at 21 and above came close to tripling, rising from 14 to 38 per cent. In France, terminal education ages shifted upwards even more. Over half, 54 per cent, of those leaving education in 1980–84 left at ages 13–18, a proportion that had fallen to 17 per cent by 1995–99. A small increase occurred from 30 to 36 per cent in leaving ages 19–21, but the proportion leaving at ages 22 and above tripled, rising from 16 to 47 per cent. Compositional change in terminal education

Table 2 Period mean age at first birth (life table calculation). Exposure and births at ages 13–39. Britain and France 1980–84 to 1995–99

	Period				
	1980–84	1985–89	1990–94	1995–99	Change 1980–84 to 1995–99
Britain	25.5	25.6	26.4	26.9	*1.4*
France	25.1	26.1	26.7	27.5	*2.4*

Source: As for Table 1.

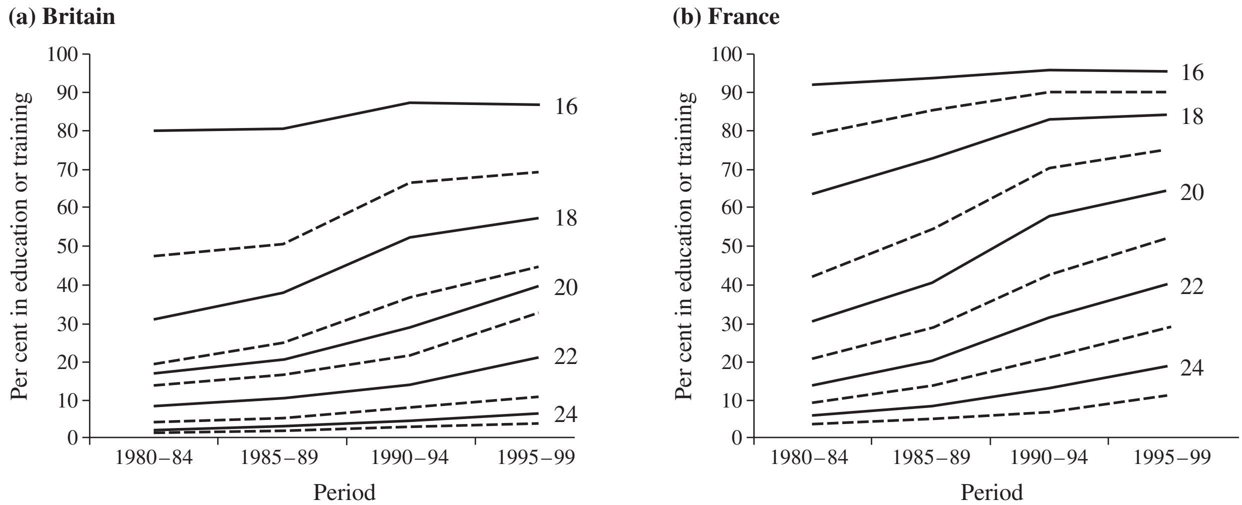

Figure 2 Proportion of women enrolled in education and training by age. Britain and France 1980–84 to 1995–99
Source: As for Table 1.

ages has clearly been substantial and of a magnitude sufficient to have had a sizeable effect, in principle, on the timing of family formation. Similar trends have occurred among men, but are not shown here.

Trends in educational participation can be summarized by the period mean age at leaving continuous education, also derived from a period life table. This indicator rose in Britain by 1.4 years from 18.3 in 1980–84 to 19.7 in 1995–99. In France, the corresponding change was from 19.8 to 21.6, a slightly larger increase of 1.8 years. Thus, in both countries, time spent in education lengthened during these decades. As a result, women of a given age in more recent periods have been out of education for a shorter time than in the past. For example, in Britain the median length of time since leaving education among women aged 20–24 was 1 year shorter in 1995–99 than in 1980–84; in France the median length in this age group was 3 years shorter, reflecting the more substantial rises in enrolment in France over these two decades (those still in education in each period were included in this calculation, grouped at a duration of <0). In other words, owing to a longer time spent in education, women aged 20–24 were of what Skirbekk et al. (2004) have termed a younger social age in the late 1990s than in the early 1980s.

Change in the duration from the end of education to first birth

The question arises therefore whether women have been having their first birth later in recent years simply because they have been spending more time in education. To answer that question, we began by examining the profile of first-birth rates by time since finishing education. The data are plotted in Figure 3, again standardizing the duration-specific rates to sum to 1.0 in each period. We see that first-birth rates are closely associated with time since leaving education in both countries, displaying a unimodal curve that is truncated at the left. In both

Table 3 Distribution of age at finishing education among women leaving education at ages 13–39 in each period. Britain and France 1980–84 to 1995–99

		Period			
		1980–84	1985–89	1990–94	1995–99
	Age at finishing education	%	%	%	%
Britain	13–17	65	58	42	37
	18–20	21	24	30	25
	21 +	14	19	28	38
France	13–18	54	41	25	17
	19–21	30	35	39	36
	22 +	16	24	36	47

Source: As for Table 1.

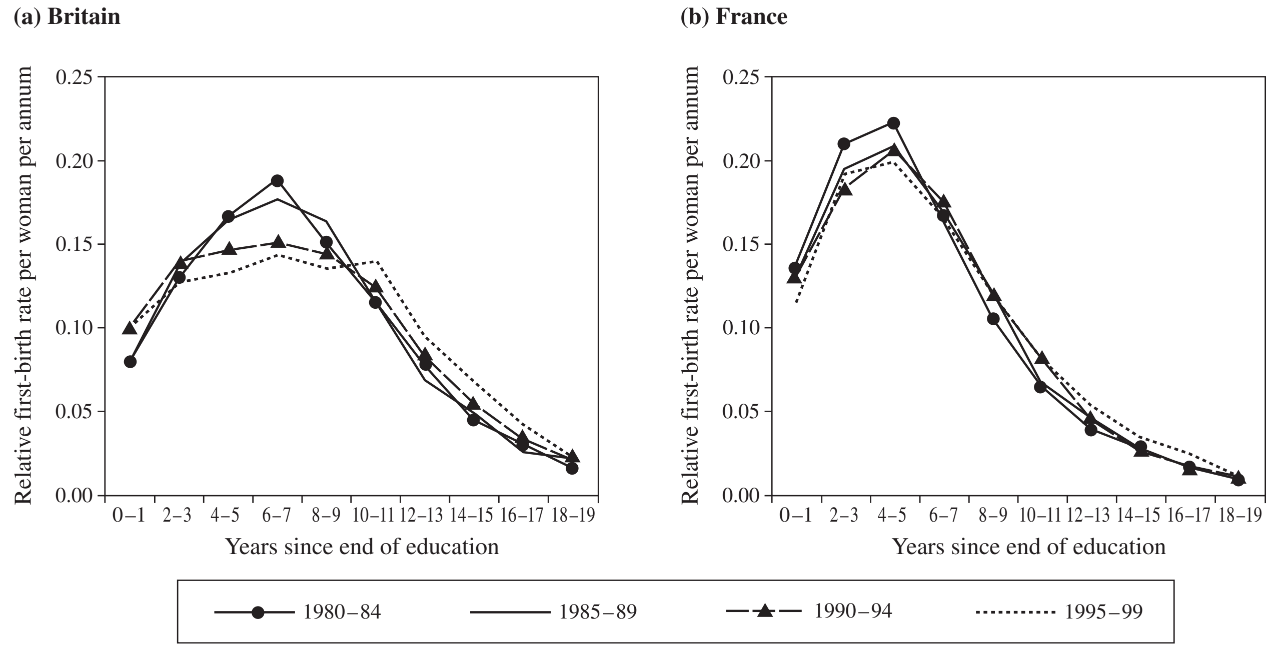

Figure 3 First-birth rates by duration since the completion of education, 1980–84 to 1995–99. Women aged 13–39 in each period, who have completed their education. Britain and France 1980–84 to 1995–99
Note: First-birth rates at durations < 0, while in education, are very low and are not shown here.
Source: As for Table 1.

countries the distribution shifts to the right, but the displacement of the curve is much less pronounced on the duration scale than on the age scale. For example, while the peak age of first birth moves up from 20–24 to 25–29 in both British and French first-birth schedules over the period covered (Figure 1), the peak duration of first-birth rates, measured relative to terminal education age, remains static at 6–7 years in Britain and 4–5 years in France (Figure 3). As is true of the age-specific schedules, the profiles of first-birth rates by time since leaving education become somewhat flatter and broader in spread over time. This is especially true in Britain, where the interquartile range of the time to first birth in the period life table rose by a quarter, from 6.1 to 7.6 years, compared with a rise of 10 per cent in France (from 5.3 to 5.8 years).

Time trends in the duration from the end of education to first birth are summarized in Table 4 (see Appendix for further details). We see that the gap between education and first birth has changed much less than has the mean age at first birth. In Britain the rise of 1.4 years in the mean age at first birth contrasts with an increase of just 0.6 years in the time from completing education to first birth, over the two decades analysed. In France age at first birth increased by 2.4 years but the gap between education and first birth increased by just 0.5 years. In other words, the timing of parenthood has changed much less if we start the clock at the end of education than if we measure in years of age. Assuming that the direction of causation is largely from enrolment to lower fertility, an issue considered further in the discussion section, the clear implication is that the growth in educational participation, and the resulting later terminal education ages, are an important part of the explanation of the later childbearing observed in recent decades. On these figures, the increase in continuous educational enrolment explains 57 per cent ((1.4 − 0.6)/1.4) of

Table 4 Mean duration from end of education to first birth (period life table calculation). Women aged 13–39 in each period who had completed their education. Britain and France 1980–84 to 1995–99

	Period				
	1980–84	1985–89	1990–94	1995–99	Change 1980–84 to 1995–99
Britain	7.8	7.8	8.2	8.4	0.6
France	5.8	6.1	6.1	6.2	0.5

Source: As for Table 1.

the rise in age at first birth in Britain, and 79 per cent ((2.4 − 0.5)/2.4) in France. The upward shift in educational participation is, on these estimates, a major contributor to the aggregate increase in the age at first birth in these two countries, via its impact on the age at finishing education.

Differentials by educational level

On the present estimates, the increase in enrolment accounts for around three-fifths of the timing shift in Britain and about four-fifths in France, over the two decades covered. However, that leaves a sizeable part of tempo change to be explained. The time to first birth can be separated into two phases—time until leaving school and the time from the end of education to first birth. Change in first-birth timing can, correspondingly, be divided into two components: one attributable to changing enrolment and one to a lengthening of the post-enrolment phase (Marini 1984, 1985; Blossfeld and Huinink 1991).

A natural question is whether educational level is associated with the tempo of fertility after the end of education. This question is in two parts, relating to the individual level and to aggregate change. (i) Does the time from the end of education to first birth differ between women of different educational levels? (ii) Do educational groups differ in the extent to which they have lengthened the time to first birth following the end of education, in recent decades?

To address these questions, we estimated the time from completing education to first birth for three educational groups, defined by their terminal education ages, for each of the periods examined. The comparison was confined to women known, in each period, to have already completed their education; the calculation thus omits those in education at the time of the survey (see Appendix). The age groups distinguished were chosen to be reasonably homogeneous internally and heterogeneous between groups in respect of the highest educational qualification achieved. For this reason, we classified terminal education ages differently in Britain and France. In the British data, we used the following groupings of terminal education age: 13–17 comprising those who either had no qualifications or the lowest level of qualification; 18–20 identifying those with qualifications above the minimum level but below the third level; and 21+ covering those with third-level qualifications. In France, the corresponding categories were 13–18, 19–21, and 22+.

Results are given in Table 5 and Figure 4. We look first at the cross-sectional differentials between educational groups. In France these are straightforward—in all periods, the top education group had their first child sooner after leaving education than did the less well educated. The best educated women seem in other words to be 'catching up' following their lengthier spell in education (Blossfeld and Huinink 1991). The picture in Britain is a little more complex. At the start of the period, the most qualified women were, as in France, having their first birth sooner than the less well qualified after leaving education. However, by 1995–99 well-educated British women took 1 year longer after leaving education to have a first birth than did the least well educated. Note however that, as elsewhere, and consistent with long-standing patterns, the best educated women in both countries had their first birth at later average ages, at all points in time (not shown). In France, and in Britain at the start of the period covered, this resulted from a combination of later terminal education ages and shorter time to

Table 5 Mean duration from end of education to first birth, by age at finishing education (period life table calculation). Women finishing their continuous education at ages 13–39. Britain and France, 1980–84 to 1995–99

		Period				
		1980–84	1985–89	1990–94	1995–99	Change 1980–84 to 1995–99
	Age at finishing education					
Britain	13–17	7.8	7.6	7.7	7.6	−0.3
	18–20	7.9	8.1	8.8	8.8	0.8
	21+	7.0	7.6	8.1	8.6	1.6
France	13–18	6.5	7.0	6.8	7.1	0.6
	19–21	6.2	6.7	6.6	6.8	0.6
	22+	5.1	5.7	6.2	6.4	1.2

Note: The figures within each group reflect the timing of birth relative to the age at completing education, and therefore within-group change in time to first birth is not attributable to within-group change over time in terminal education age.
Source: As Table 1.

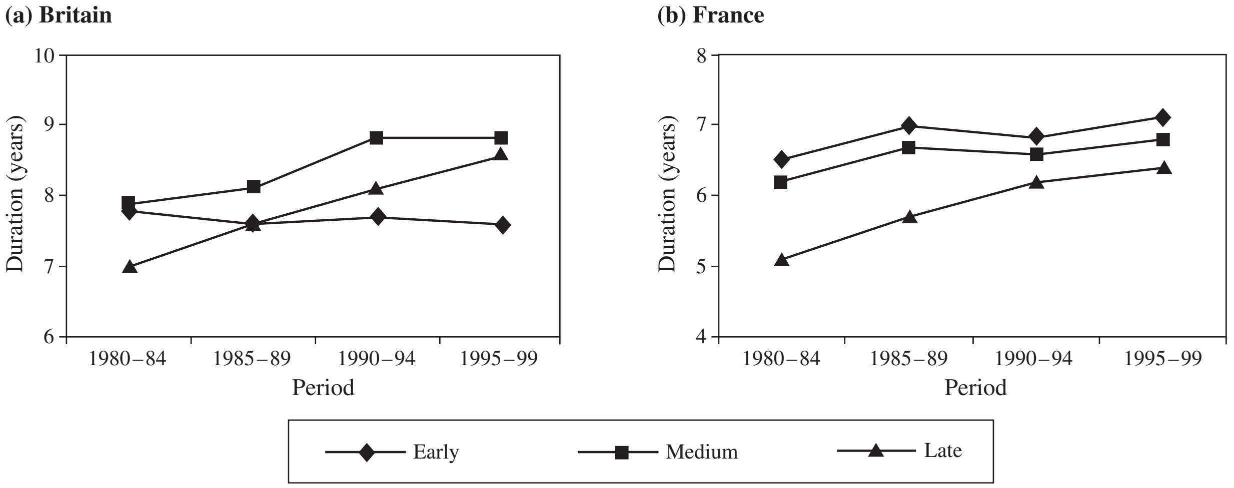

Figure 4 Mean duration from end of education to first birth (period life table calculation) by age at finishing education. Women aged 13–39 in each period who had completed their education. Britain and France 1980–84 to 1995–99
Source: As for Table 1.

first birth after completing education than other education groups. However, at the end of the period in Britain, the later first-birth timing of the best educated women was due both to a later age at completing education and to a longer wait to first birth, after leaving education.

Turning now to change over time, we see from Figure 4 and Table 5 that education groups differ in the extent to which they postponed the first birth after leaving education, over the two decades. In both countries, the best educated increased the time to first birth, after completing their education, by a greater margin than did the less well educated. In Britain, women who had stayed longest in education waited 1.6 years more after leaving education to have a first birth in 1995–99 than in 1980–84; this contrasts with a slight decline, of 0.3 years, among the early school leavers and an increased wait of 0.8 years in the middle group. In France also, the best educated postponed their first birth by longer, an additional 1.2 years, after completing their education than did the other two groups, each of whom postponed by a further 0.6 years. Thus, tempo change in the post-enrolment phase of the life-course was positively associated with educational level in both countries.

In all, education was associated in two distinct ways with the lengthening time to first birth during the 1980s and 1990s. The major effect, as we saw earlier, was the impact of the trend to later school-leaving ages. A secondary factor was that the best educated women postponed their first birth, after completing their education, by more than the less well educated.

Britain and France compared

We comment here briefly on the contrasts between Britain and France. We saw above that early school leaving was more common, and declined less, in Britain than in France over the period analysed. Table 3 above shows that in the period 1995–99 over twice as many—37 per cent—left school early (at ages 13–17) in Britain as in France, where 17 per cent left at ages 13–18. As a result, many more teenage women in Britain than in France were not in education. Furthermore, the disparity between the two countries grew over the 20-year period. One outcome of these contrasts is shown in Figure 5, which plots the ratio of the proportion of women not in education at each age from 16 to 29 in Britain to the equivalent in France, by period. The proportion of 16-year-olds not in education in Britain was two and a half times that in France in 1980–84, and the ratio increased to over three by 1995–99. In other words, proportionately more teenage women in Britain were free of the competition between the roles of student and prospective mother than were their counterparts in France. The relative levels of enrolment may, thus, partly explain the higher fertility rates among teenagers in Britain than in France. In addition, the smaller increase in enrolment among teenagers in Britain may partly explain the relatively static fertility rates at young ages in Britain—these contrast sharply with the shift up the age range of the entire schedule of *both* enrolment *and* fertility rates in France. More generally, differential trends in educational enrolment may be a significant part of the mechanism that gives rise to

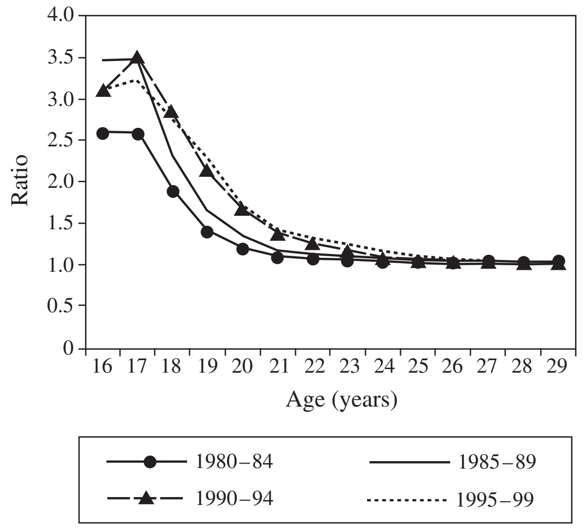

Figure 5 Ratio of proportion of women of each age who were not in education in Britain to the equivalent in France, by period. Women aged 16–29. Britain and France 1980–84 to 1995–99
Source: As for Table 1.

the 'hump' at early ages in the fertility schedules of the English-speaking countries (Chandola et al. 1999, 2002; Sullivan 2005).

Discussion

Educational attainment has long been known to be associated with later childbearing at the individual level, and educational expansion has often been cited as an influence on the changing timing of fertility. However, direct evidence of the role of rising educational levels in influencing aggregate change in fertility tempo has been scarce. Our study provides such evidence. We have demonstrated quantitatively the major contribution made by changing educational participation to aggregate trends in fertility timing. On present estimates, around three-fifths of the shift to later first-birth timing in the 1980s and 1990s in Britain, and four-fifths in France, is attributable to the rise in educational participation. We have shown also that part of the additional lengthening of the time to first birth not explained by rising educational enrolment is related to educational level, with the best educated women postponing their births for longer following the end of education than other groups. Our analysis highlights the importance of age at completing education, as distinct from educational attainment, both for the timing of first birth and for an understanding of trends in the tempo of fertility.

Previous investigations that have addressed substantially the same question are not numerous and have reported mixed results. Two studies based on US data employed standardization to evaluate how far period change in marriage (Oppenheimer et al. 1995) and in fertility (Rindfuss et al. 1996) could be attributed to compositional change resulting from the growth in educational attainment. Both concluded that educational expansion was not a primary cause of the trend to later marriage and childbearing. Rendall et al. (2010, pp. 218–9) found that composition by educational attainment accounted for little of the change in the cohort age pattern of first birth in seven countries. However, a recent study of the fertility of Belgian cohorts of 1946–50 and later, that also addressed the question directly, is closer to the results we report here. Using standardization, Neels and De Wachter (2010) found that 37–54 per cent of cohort change in cumulative first births to age 25 was attributable to the changing distribution of educational attainment. While these findings reflect both quantum and tempo, the age limit of 25 means that timing is a substantial part of the effect explained. Neels (2009) standardized period fertility rates for composition by educational attainment, and reported substantively similar results. Skirbekk et al. (2004) showed that the link between later school-leaving age and the timing of demographic events in Sweden was causal, but they did not address the empirical question of how far changing enrolment contributed to change over time in childbearing ages.

That the increase in educational participation is linked with slower aggregate fertility tempo via both rising enrolment and a lengthening of the post-enrolment phase fits well with Marini's suggestion that, at the individual level, educational attainment influences transitions to adulthood by a two-step process (1984, 1985): the impact on school-leaving age, and effects net of school-leaving age (see also Hogan 1978; Blossfeld and Huinink 1991; Blossfeld 1995). The greater upward shift in the length of the post-enrolment phase among the best educated is broadly consistent with evidence of greater temporal shifts in the level and timing of fertility among women with higher educational attainment than others reported for the USA (Rindfuss et al. 1996; Ellwood and Jencks 2004; Sullivan 2005), the Nordic countries (Kravdal and Rindfuss 2008; Andersson et al. 2009), Britain (Rendall et al. 2005), and France (Robert-Bobée 2005); nevertheless, these studies do not distinguish enrolment and post-enrolment effects. The result is, however, at variance with Blossfeld and Huinink's conclusions (1991), in relation to

Germany, that later childbearing was due entirely to longer enrolment, and that educational level was not associated with tempo change, net of enrolment. Of course, the size of the two components may differ cross-nationally.

The growth in educational participation is generally seen as having influenced the trend to later birth timing in two principal ways. Micro-economic theory is the source of the most commonly proposed mechanism—opportunity costs. The rise in educational attainment is suggested as having resulted in later fertility tempo via opportunity costs for a combination of reasons. Women in general now participate in the labour force for a much longer portion of their lives than formerly; consequently, many more than in the recent past will expect to return to the labour force after starting a family. In addition, better educated women experience greater opportunity costs from leaving the labour force around the time of childbearing. Together, these two factors mean that the rise in educational attainment will have increased the opportunity costs of childbearing for the average woman: proportionately more women will have an incentive to postpone childbearing and establish their earning power, to ensure they can later return to the labour force on favourable terms (Liefbroer and Corijn 1999; Billari and Kohler 2002; Gustafsson et al. 2002; Rendall et al. 2005). This mechanism may well account for our finding that post-enrolment increases in the time to first birth have been greater among the best educated women than others. Note, however, that this explanation is relevant only to the rise in the *post-education* postponement of childbearing. This is because the opportunity costs associated with better educational qualifications are encountered only *after* these have been attained. The component of the rise in the age at first birth resulting from enrolment itself—the major part of the tempo shift on our estimates—cannot be explained via opportunity costs. The same applies to the expectation that the link between education and later childbearing may weaken, on the grounds that economic activity and childbearing are becoming easier to combine (Liefbroer and Corijn 1999; Lappegard and Ronsen 2005; Winkler-Dworak and Toulemon 2007). For example, the increasing provision of formal childcare has reduced the employment–fertility conflict, as suggested by evidence that the availability of childcare is positively associated with first-birth rates (Rindfuss et al. 2007, 2010). On this argument, the opportunity costs of childbearing are expected to diminish accordingly. Again, any weakening of the education–tempo link for this reason can affect only the post-enrolment element of the move to later childbearing. Greater compatibility of childbearing and employment will have no impact on the tempo shift attributable to educational participation itself, though childcare provision might well reduce the impact of enrolment on fertility (Cohen et al. 2011).

The second hypothetical mechanism linking educational expansion and later childbearing is cultural and is associated with the theory of the second demographic transition. From this perspective, rising education promotes later childbearing through the impact of education on individuals' ideas, values, aspirations and preferences in relation to self-actualization, career, leisure, and family life (Rindfuss et al. 1980; Lesthaeghe and Surkyn 1988; Skirbekk et al. 2004). Again, such a process may have contributed to the greater increase in the length of the post-enrolment phase among better educated women found in the present study. However, as in the case of the opportunity-costs argument, cultural change resulting from educational participation cannot explain the larger part of the increase in the age at first birth, the part that is due to enrolment. Any changes in values and preferences brought about by exposure to education will influence *post-education* experience and behaviour. Thus, if educational expansion has had a role in slowing the speed of family formation via change in values or preferences, it is the post-enrolment component of tempo change that is influenced by this pathway.

The age at leaving education is revealed by our results to be crucial to the analysis of fertility timing at both individual and aggregate levels. The greater regularity and stability of the distribution of first-birth rates by time since leaving education, seen in Figure 3 above, suggests that this duration is more meaningful than age as a scale of personal time on which to locate demographic and social events (Skirbekk et al. 2004). Several previous authors have drawn attention to the age at leaving school as a key feature of the timing of adult transitions (Hogan 1978; Marini 1985; Oppenheimer 1994) but its importance appears to have been underappreciated. We suggest speculatively two possible reasons for this neglect. While educational attainment rose in many developed countries over the second half of the twentieth century, ages at marriage and first birth initially declined and then rose again. This U-shape has been interpreted by some authors as indicating that, in the more recent period, rises in education could not be the cause of the increase in ages at marriage and first birth since the 1970s (Blossfeld and Jaenichen 1992; Blossfeld 1995;

Oppenheimer et al. 1995; Prioux and Mandelbaum 2003; Robert-Bobée 2005). A further possibility is that contemporary social science focuses particularly on the assumed economic, ideational, and cultural impact of educational attainment, and views the enrolment effect as a merely temporary, and theoretically less interesting, influence on the adult life course. However, the straightforward, mechanical effect of educational enrolment in strongly curtailing birth rates, together with major trends in enrolment, means that it has been a powerful driver of change.

Our findings on Franco-British contrasts, and the differentiation in timing of first birth by education in each country, are in some respects parallel to those of Rendall et al. (2005, 2009, 2010). However, these authors focus on educational attainment rather than terminal education age, and consider the interaction between educational attainment, socio-economic heterogeneity, and family policy, with policy having a central conditioning role. Andersson et al. (2009) also see a strong link between fertility patterns and welfare policy. Our approach via terminal education age provides an alternative perspective on the same basic landscape. Terminal education age is clearly a product of differentiation in institutional framework and policy between countries, since it reflects, at least in part, public provision and policy in relation to education, training, and employment. Our findings on later childbearing may well be accommodated within an analysis that sees family and welfare policy as a key determinant. Another possibility is that national patterns of fertility timing are, at least in part, the incidental outcome of the histories of education systems and educational policies specific to each national context (Breen and Buchmann 2002; Rindfuss and Brauner-Otto 2008). Cross-national patterns of fertility timing could, thus, result from traditions and measures with unrecognized but non-negligible demographic effects.

Abundant evidence from event-history studies shows that enrolment is associated with very low fertility rates. The reasons generally given for the restricted fertility of women in education are that role-conflict, economic dependence, and normative expectation are strong inhibitors of fertility during this phase of the life course. Throughout the present paper we have assumed that, at the individual level, the link between enrolment and fertility timing is primarily due to a causal effect of enrolment on fertility. Two issues arise here. First is the possibility that educational participation has no impact on fertility timing but that the link between them is due to prior factors, particularly preferences and aspirations stemming from the family of origin. However, longitudinal evidence suggests that the association between enrolment and a slower transition to adulthood remains strong, net of family background and parents' characteristics (Marini 1978, 1985; Rindfuss et al. 1988; Thornton et al. 1995). Reverse causation is a further possibility. Our assumption of a direct effect from enrolment to fertility is in line with existing evidence of a much stronger effect of education on age at first birth than in the reverse direction (Rindfuss et al. 1980, 1984; Upchurch and McCarthy 1990). The findings of Skirbekk et al. (2004) and of related econometric analyses (Black et al. 2008; Monstad et al. 2008; Silles 2011) provide further empirical grounds for our assumed causal direction. On the other hand, a recent, sophisticated analysis has argued that the direction of causation between educational attainment and final family size in Norway is from fertility to education (Cohen et al. 2011). However, we believe that the findings of this remarkable study cannot be extended to our inquiry for several reasons: it analyses final family size, rather than first-birth timing; it classifies education as attainment at age 39, rather than age at first leaving education; and the findings refer to an educational system that is much more open to entries and exits throughout the life course than is the case in Britain and France.

Our results strongly imply that the shift in first-birth rates up the age range over time is largely due to the progressively later average ages at completing education. To explain the move to later childbearing the determinants of changing enrolment therefore need to be identified. These include macro-economic forces driving up demand for a more skilled workforce. Governments throughout the developed world have responded to such forces by raising compulsory school-leaving ages, and also with policies designed to encourage post-compulsory education and training (Clark 2002; Murtin and Viarengo 2011). In turn, the changing demand in the labour market, together with policy measures, has provided the economic rationale for individuals to invest further in education (Machin and Vignoles 2005). Trends in unemployment may also be implicated, by increasing the attractiveness of schooling over job search or precarious employment (Neels and De Wachter 2010; Clark 2011). In this perspective, postponed childbearing is a transformation of the life course that has originated primarily in broad macro-economic, educational, and labour market trends (Neels 2006, 2009).

In summary, our results show that changing educational enrolment explains a sizeable part of the increase in the age at first birth. They have, in addition, several further implications. First, they strengthen the rationale for analysing later childbearing, at both micro and macro levels, in two parts: a component attributable to enrolment and a second due to the post-education phase. Beyond enrolment itself, factors influencing fertility tempo, at the individual level, will, in virtually all cases, have their effect on the post-enrolment timing component. Second, our findings suggest that duration since the end of education is a more natural timescale for life-course transitions in young adulthood than is age or time since 16th birthday; it could therefore potentially give better results in modelling the frequency and timing of life-course transitions such as partnership, marriage, and first birth. Third, enrolment patterns may explain the emergence in recent decades of a 'hump' in the fertility schedules of English-speaking countries. Fourth, enrolment rates and the age at the end of education and training emerge from these findings as useful potential lagged indicators of fertility, in short-term to medium-term projections. Fifth, our findings reinforce the insight that if designing policies to target the tempo of childbearing, interventions directed to the structure and flexibility of educational systems are a potentially productive avenue of exploration (Lutz and Skirbekk 2005; Rindfuss and Brauner-Otto 2008; Cohen et al. 2011). Finally, the major role of educational expansion as a driver of the trend to later childbearing suggests that the root causes of fertility tempo change are to be found in fundamental macro-economic and structural factors rather than in the cultural domain.

Notes

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Appendix

Confining GHS analysis to 1980–84 to 1995–99

The period covered in the GHS is confined to 1980–84 to 1995–99 because of a bias, relative to vital registration, in the estimated change in the timing of period fertility when the period from 2000 onward is used as the end-point. Relative to estimates from vital registration, the standardized period mean age at first birth based on GHS rounds 2000–07 is upwardly biased in the four 5-year periods from 1980–84 onwards, by +0.4, +0.3, +0.4, and +0.5 years, respectively, but is unbiased in the period 2000–04. The result is that the estimate of change between 1980–84 and 2000–04 in the standardized period mean age underestimates the actual change by 0.4 years, a relatively large bias of 25 per cent. However, the change in the standardized period mean age at first birth between 1980–84 and 1995–99 is estimated at +1.2 years in both GHS and ONS sources, and is thus unbiased in the GHS. For this reason, we confined analysis of trends in the mean age at first birth to the periods 1980–84 to 1995–99. The upward bias in these decades in the GHS may result from the overstatement of the level of childlessness by older women in the GHS in recent years, and from the incomplete recovery of under-reported births in GHS rounds from 2000 onwards (Ní Bhrolcháin et al. 2011).

Mean duration from end of education to first birth

In calculating the mean length of time from the end of continuous education or training to first birth, women having a first birth while still in education are included as a group, for comparability with the estimated trend in the mean age at first birth. We did this by estimating the average duration at first births occurring before the end of education. For births occurring in 1980–84 and 1985–89, this is a straightforward mean number of years, since virtually all those interviewed in 2000–07 and having a birth in 1980–89 while in education, had completed their continuous education by the date of interview. Therefore, the (negative) length of time to completing education of any first birth in these periods is known. However, in 1990–94 and 1995–99, the age at leaving education, and therefore the (negative) length of time to first birth, was unknown for a minority of women who had not left education by the time of interview in 2000–07. Therefore, for births to women in education in 1990–94 and 1995–99, we estimated the mean length of time before age of ending education by assuming that the average (negative) length of time between finishing education and first birth of those whose education was censored by the occurrence of interview was equal to that of uncensored women in the corresponding period. The assumption has a very minor effect on the estimated mean because births before the end of continuous education are rare in these populations, accounting for only 2–4 per cent of births across the period examined.

Analysing timing change by age at completing education

We classify educational groups by terminal education age because a bias would be introduced by classifying according to reported educational attainment at survey. This is because attainment at survey includes qualifications reached on returning to education, after a spell out of it, and is therefore only an imperfect reflection of attainment at leaving continuous education (Hoem and Kreyenfeld 2006; Cohen et al. 2011). In general, this would be expected to bias downwards the mean age at first birth of women classified as better qualified since that average includes births occurring *before* they reached their highest level of attainment and when they were, therefore, less well qualified. As a result, the change over time in the mean age at first birth among better qualified women would probably be underestimated if classification by attainment were used.