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Francesco Cinnirella Ifo Institute, CESifo and CEPR
Marc P. B. Klemp Brown University and University of Copenhagen
Jacob L. Weisdorf University of Southern Denmark, Utrecht University and CEPR

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Malthus in the Bedroom: Birth Spacing as a Preventive Check Mechanism in Pre-Modern England*

Francesco Cinnirella†
Ifo Institute, CESifo and CEPR

Marc P. B. Klemp‡
Brown University and University of Copenhagen

Jacob L. Weisdorf§
University of Southern Denmark, Utrecht University and CEPR

Abstract

The role of demography in long-run economic growth has been subject to increasing attention. This paper questions the received wisdom that marital birth control was absent before the nineteenth century. Using an extensive individual-level dataset covering 270,000 births from 80,000 families we show that higher national and sector-specific real wages reduced spacing between births in England over more than three centuries, from 1540-1850. This effect is present among both poor and rich families and is robust to a wide range of control variables accounting for external factors influencing a couple's fertility such as malnutrition, climate shocks and the disease environment.

Keywords: Spacing, birth intervals, fertility limitation, natural fertility, preventive check

JEL classification: J11, J13, O12, N33

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† Corresponding author: Poschingerstr. 5, 81679 Munich, Germany. E-mail: cinnirella@ifo.de.
‡ Øster Farimagsgade 5, building 26, DK-1353 Copenhagen, Denmark. E-mail: marc.klemp@econ.ku.dk.
§ Campusvej 55, DK-5230 Odense M, Denmark. E-mail: jacobw@sam.sdu.dk.

1. Introduction

The industrial revolution and the subsequent fertility transition strongly influenced the economic development of Western economies (Galor, 2011; Murtin, 2013; Dalgaard and Strulik, 2013). England was the first country to make the transition from Malthusian stagnation to sustained economic growth. Recent studies have argued that birth limitation helped secure England's world leadership and that it was practised long before the Industrial Revolution (Voigtländer and Voth, 2013). The common view is that fertility limitation was achieved only by delaying marriages, effectively reducing the wife's childbearing period, and that England remained a *natural fertility* society characterised by uncontrolled marital fertility until the nineteenth century (Clark, 2007; Voigtländer and Voth, 2013; Wilson, 1984). This view is shared by the *European Fertility Project*, concluding that marital birth limitation first came into play in the late nineteenth century through the diffusion of knowledge about contraceptive methods including *coitus interruptus*, sexual abstention, and extended breastfeeding (Coale and Watkins, 1986; Coale, 1986).

This paper makes two key contributions that cast doubt on these views. We show that couples did indeed postpone the timing of their marriage in response to lower aggregate real wages. Yet, we also show that lower real wages were associated with longer intervals between births *within* marriage long before the nineteenth century. These findings are based on a comprehensive dataset based on family reconstitutions covering more than 270,000 births from more than 80,000 families spanning the three centuries leading up to England's fertility transition in the late nineteenth century. Using a wide range of duration model specifications, our analysis establishes a negative effect of national (and sector-specific) real wages on the spacing of births within families. This effect concerns not only poor families (labourers, servants, and husbandmen), but also their more affluent counterparts (gentry and merchants). We show that the effect is robust to the inclusion of factors that may interfere with the biological capacity to conceive, such as malnutrition caused by high food prices, extreme climatic conditions, and the disease environment.

Existing studies on historical England remain largely inconclusive about the existence of Malthusian *preventive checks*. The present study advances the research frontier along several dimensions.[1] First, we examine a larger sample than the existing studies on historical England, exploiting information from a total of 26 English parishes. Second, we employ duration models to study the effect of living standards on the *timing* of events (marriage, as well as first, last, and intermediate births). Third, we study a much richer sample than those used in the existing literature (e.g. Kelly and Ó Gráda, 2012). The nature of our data (family reconstitutions) allows us to account for a wide range of variables, including parental occupation, literacy, maternal marriage age, birth order, parish and time fixed effects.

The remainder of the paper is structured as follows: In Section 2 we describe the key features of the data and the potential problems related to their use; in Section 3 we analyse fertility patterns by estimating duration models; in Section 4 we perform several robustness checks to ensure that the effect of real wages on spacing reflects deliberate behaviour and that our results are not affected by migration and parish attrition; in Section 5 we conclude.

2. Data

The data used to explore the influence of real wages on the spacing of births come from Anglican parish registers (English church books). Collected over the past 40 years by the *Cambridge Group for the History of Population and Social Structure* and described by Wrigley and Schofield (1981), the original dataset provides yearly occurrences of births, deaths, and marriages for the period 1541–1871. These data have previously been used to explore the effect of real wages on the number of national or regional (aggregate) births.[2]

Meanwhile, inspired by Louis Henry's reconstitution of families in France (Henry, 1967), the Cambridge Group collated the ecclesiastical events in 26 parishes to reconstitute over 80,000 families responsible for more than 270,000 births. The 26 parishes, forming the *reconstitution data*, were chosen on the basis of the quality of the

[1] Recent studies in historical demography have shown that marital birth limitation was practiced in the Low Countries, Germany, and Sweden from the late eighteenth century onwards (Bengtsson and Dribe, 2006; Dribe and Scalone, 2010; Van Bavel and Kok, 2004; Van Bavel, 2004).

[2] Kelly and Ó Gráda (2012) find some support that lower real wages reduce birth rates, depending on the time period and region examined. Bailey and Chambers (1993), Crafts and Mills (2009), and Lee and Anderson (2001) have also looked at effects at the aggregate level finding limited support.

data and with the aim of selecting a set of parishes that would be representative of England. The sampled parishes range from market towns to remote rural villages, including proto-industrial, urban, and agricultural communities. The data is discussed in detail by Wrigley et al. (1997).

The reconstitution data are not without shortcomings. Among their limitations, Ruggles (1999) has pointed to migration as being of particular concern to scholars dealing with birth spacing intervals. When compared to the French families, studied by Louis Henry, English families were rather more mobile. In the English reconstitution data, about one-third of all births permanently left their parish of origin. This raises two concerns. One is that migrants and non-migrants had different tendencies when it came to spacing their offspring. Fortunately, permanent migrants can be identified in the data by their missing birth or death dates, allowing us to separate them from their non-migrating counterparts. Robustness analyses focusing on *completed families*, that is, including only those couples that exhausted their reproductive lifetime in the parish of origin, establish that our results are robust to the exclusion of permanent migrants.[3]

The second concern is that of *temporary* migration leading to unobserved births. A systematic association between living standards and giving birth in a parish other than the home parish may induce a selection bias. Robustness analyses discarding particularly large birth intervals, thus reducing the possibility of unobserved intermediate births, suggest that our results are not biased by temporary migration.[4]

Under-registration of births is another concern raised against the use of parish registers for the present purpose. As is the case with temporary migration leading to unobserved births, under-registration is an issue if it is correlated with living standards. Since couples paid a fee to the church to have their newborn baptized, one might suspect that the poor would have delayed a baptism, or skipped it entirely, during periods with depressed wages. There are three reasons why we do not believe this to be the case. First, poor couples were usually exempted from paying the church fees, because it was in the interest of the church to keep them within the Christian faith (Wrigley et al.,

[3] As will be explained further in section 4.4, a "completed family" is defined as a marriage in which both the wife and the husband survived (at least) until the wife reached the age of 50 years. It therefore consists of a couple that exhausted its reproductive lifetime in the parish of origin.

[4] It should be noted that, while times of hardship may have induced husbands to travel in search of work, hence delaying the birth of the next child, such temporary migration does not pose a problem. Indeed, such behaviour is a type of fertility control that affects the interval between births in response to changing living standards.

1997). Second, according to the doctrine of the Church of England, the sacrament of baptism was necessary for salvation; hence, in light of the high infant mortality throughout this time period, there was a strong religious incentive not to delay the baptism of a child. Last, the association between aggregate wages and birth spacing is found among both rich and poor families.

The use of parish registers for the purpose of examining the spacing of births is not new. In a descriptive analysis of the parish of Colyton (one of the 26 parishes in the reconstitution data) Wrigley (1996) found limited support for birth control in around 1700. This was attained, he argued, through late marriages, extended birth intervals, and low stopping ages. However, after adding 12 parishes to the sample (totalling 13 of the 26 parishes) Wilson (1984) did not find similar results, concluding instead that pre-industrial England was a *natural fertility* society. We build on the work of Wrigley (1966) and Wilson (1984) and include all the 26 reconstituted parishes. Furthermore, using a flexible duration model approach, we estimate the impact of real wages on birth intervals accounting for a wide range of potentially confounding factors.

In contrast to the aggregate birth and marriage rates used in recent work on English birth patterns (Kelly and Ó Gráda, 2012), family reconstitution data enable us to account for individual-level and family-level covariates. The reconstituted families are built around a marriage, providing information about the birth and death dates of the spouses, as well as the gender and birth and death dates of their offspring. In fact, the church would usually record the date of the baptism rather than that of the birth, but we use birth dates where available. In our econometric analysis we use the date of conception, obtained by subtracting 280 days from the birth date variable.[5] To assess the validity of the implied birth dates, Figure 1 and 2 illustrate the distribution of births by the month and the day of the month, respectively. The distribution by month does not show any tendency of heaping. However, Figure 2 suggests some heaping, especially in the months of January and December. The spike on December 25th can be explained by the preference of couples to baptize their children on Christmas Day. The spike on January 1st relates to missing (unreadable) dates, imputed by the transcribers as the first day of the year. England switched from the Julian to the Gregorian calendar in 1752, by which

[5] This corresponds to 38 weeks of gestation and an average duration of two weeks from the onset of intercourse until ovulation. Our results are not sensitive to the use of alternative intervals, e.g. intervals of 38 or 42 weeks.

time it was necessary to correct by 11 days. Hence, the spike on January 11th is the result of similar reasons to the spike on January 1st. In our analysis, we include dummy variables indicating the following dates: December 25th, January 1st, and January 11th.[6]

[Figure 1: Distribution of births by month]

[Figure 2: Distribution of births by day of the month]

The dataset provides ample information about the socio-economic background of the family. The clergy frequently reported the occupation of the husband (and, very rarely, of the wife). The occupations were recorded at the time of the husband's marriage and burial, but also on the occasion of the baptisms or burials of offspring. The occupational titles allow us to classify the families according to their wealth or income potential. Using will records from historical England, Clark and Cummins (2010) have constructed seven socio-economic groups ordered according to the wealth information found in the wills. From the poorest to the richest these are: labourers, husbandmen, craftsmen, traders, farmers, merchants, and gentry.[7] Relying on their coding system, we have used the earliest recorded occupation of the husband to classify the sampled families (and a binary variable when the occupation is missing). Moreover, the literacy status of the spouses can be inferred from their wedding certificates. Literate spouses left a signature on the certificate (as opposed to a mark). Literacy has been widely used as an indicator of human capital before public schooling became widespread (Clark, 2008).

As is common in historical demography (Wrigley et al., 1997) we can use the church book statistics to estimate the fecundability of the sampled couples by measuring the time span from the date of their marriage to that of their first birth (Klemp and Weisdorf, 2012). Finally, we can also account for couples in which the first child was prenuptially conceived. In particular we constructed a binary variable which takes on value one if the difference between the date of marriage and of the first birth is less than 40 weeks, the average length of the gestation period.

[6] Removing these observations entirely, or omitting the dummy variables, does not change our qualitative results.

[7] We are grateful to Gregory Clark for providing us with a mapping procedure for this.

2.1 Fertility Outcomes and Control Variables

We estimate the effect of real wages on the hazards of four different events: (*i*) marriage, (*ii*) having the first birth (denoted "starting"), (*iii*) having a birth following a preceding birth (denoted "spacing"), and (*iv*) having the last birth (denoted "stopping"). In the "marriage", "starting" and "stopping" analyses, the wife is included once and the outcome variables are the points in time when the relevant event occurred. We assume that the wife becomes *at risk* of conception from the age of 15.[8] In the "spacing" analysis, the event analysed is the timing of the conception of a child and the wife is considered to be at risk of conceiving at the time of the previous birth. Each of the four events are regressed on real wages (described below) as well as a set of control variables, including the birth order of the child, the family-level child mortality up until the point in time of the conception of the child, the father's income class, the mother's age at marriage, the mother's literacy status, and the couple's estimated fecundity.[9]

[Table 1: Summary statistics]

The summary statistics are reported in Table 1. The average age at (first) marriage of the wives is 23.7 years, and the average age at "starting" is 25 years. The average time span between the marriage and the first birth (a proxy for the couples' fecundity) is thus slightly more than a year. The average length of birth intervals is 929 days (2.5 years) with a standard deviation of 475 days (1.3 years). The occurrence of twin births constitutes approximately 2% of all births and is treated as a single event. The relatively few cases (0.5%) where the birth intervals are less than 40 weeks (stemming either from premature births or data errors) have been removed from the sample.[10] The most common occupations in the data are labourers (32% of the observable occupational titles), craftsmen (21%), and husbandmen (17%). In roughly 51% of all families in the sample we have no information about the father's occupation. Information about the literacy of women appears only after 1750 (19% of all mothers). Within this subsample, about 33% of the women were able to sign the register.

[8] The Church of England did not permit women under 15 years old to marry. Our results are robust to different specifications.
[9] Since parental literacy is highly correlated, we include only the wife's literacy.
[10] Their inclusion has no impact on our conclusions.

The last birth spacing interval is substantially longer, on average, than preceding intervals (see Table 2). While this can be attributed, in part, to the fact that fertility declines with age (Baird et al., 2005), some demographers have argued that it could reflect failed attempts to end the childbearing period (Van Bavel, 2004; Okun, 1995; Knodel, 1987; Anderton, 1989). For these reasons we estimate the model for birth spacing both including and excluding the last birth intervals. It should be noted that the sample in the stopping regressions consists only of completed families, ensuring that permanent migration does not affect the observation of the stopping event.

[Table 2: Average birth intervals by birth order and period]

2.2 Real wages, Prices and Death Rates

Our key explanatory variable is living standards measured by national real wages. The real wage series used in the main analysis is provided by Clark (2007). The series is constructed by dividing a national index of nominal wages for unskilled farm workers by a national cost-of-living index.[11] The wage series combines observed wages across all of England, as discussed in Clark (2007).[12] Since all households face the same annual real wages in the analysis, we introduce some cross-sectional variation by using Clark's wage series for building workers (Clark, 2005). Since builders were skilled, and farm workers unskilled, we assigned the unskilled wages to labourers and husbandmen, and the skilled wages to craftsmen, traders, farmers, merchants, and gentry.

Figure 3 illustrates the relationship between the average birth intervals and standardized real wages in percentiles. It clearly shows that high real wages are associated with short birth spacing intervals. We obtain a similar gradient when looking at average birth spacing intervals by occupational group (Figure 4): more affluent social groups (traders, merchants, and gentry) have shorter birth intervals.

[Figure 3: Spacing by real wages]

[11] The data are available from the author's website: http://www.econ.ucdavis.edu/faculty/gclark/data.html.

[12] The same real wages series was used in a recent work analysing the impact on parish-level marriages and births (Kelly and Ó Gráda, 2012).

[Figure 4: Spacing by occupation]

A central concern is the possibility that the observed association between living standards and birth intervals is the result of biological effects and not the result of deliberate choice. In particular, the ability to conceive can be adversely affected by malnutrition (Bongaarts, 1980) and the response of birth intervals to living standards may therefore reflect a higher risk of malnutrition in periods with low real wages. Likewise, it is possible that climatic conditions, revealed by air temperatures, may affect fertility (Lam and Miron, 1996). We attempt to control for such potentially confounding factors by using the following three variables: wheat prices, temperature, and mortality rates.

Wheat was a main staple in historical England, so high wheat prices are indicative of years with a high risk of malnutrition. The annual prices of wheat are provided by Clark (2007). The longest available record of *measured* surface air temperatures for any country in the world is for England for the period 1650–1850, provided by Manley (1953). Lastly, the national crude death rate given by Wrigley et al. (1997) is a useful control variable that additionally proxies for years of famine, war and thus malnutrition. Furthermore, we exploit the reconstitution data to calculate yearly parish-specific stillbirth rates using the share of births baptised and buried on the same date. Since these rates are computed annually and at the parish level, they provide a useful measure of local conditions that were associated with infertility and miscarriages. The descriptive statistics of these control variables are presented in Table 3.

[Table 3: Summary statistics of aggregate variables]

3. Duration Analysis

In this section, we explore the effect of real wages on "marriage", "starting", "spacing", and "stopping". We use the Cox Proportional Hazard model treating the real wage as a time-varying covariate (Cox, 1972). The model is specified as follows:

$$h(t) = h_0(t)\exp(\beta_1 x_1 + \cdots + \beta_k x_k + g(t)(\gamma Z)) \qquad (1)$$

The term $h_0(t)$ is the baseline hazard function; $(x_1, \ldots, x_k)$ are socio-economic and demographic covariates; and Z is the standardized time-varying national real wage. The model is stratified by parish and quarter century. Thus, each stratum has its own baseline hazard function. Since the demographic events are measured at the individual level, whereas the real wage is measured annually at the national level, standard errors are clustered by years.

3.1 Effect of Aggregate Real Wages

Table 4 reports the results of the duration models for the period 1540–1850. The wage series used in the baseline analysis is that of unskilled farm workers. To ease interpretation, the real wages are standardized with a mean of zero and a standard deviation of one. The coefficients are reported as semi-elasticities, with more positive coefficients corresponding to a higher risk of occurrence of the events studied, i.e. a shorter duration to the event.

Table 4 establishes that the real wage has a significant, positive impact on the risk of marriage and starting (columns 1 and 2). A one-standard deviation increase in the real wage increases the probabilities of marriage and first conception by roughly 25%. The negative effect of real wages on the marriage age, never before documented at the family level, is first-hand evidence of a deliberate Malthusian *preventive check* operating in historical England.[13] Consistent with the social rule of the time that expected women to conceive immediately after marriage, we find that the magnitude of the effects on "marriage" and "starting" are roughly identical.

[Table 4: Marriage, starting, spacing, and stopping]

Malthus (1789) held that deliberate birth limitation existed *prior* to marriage. This makes sense from the perspective that couples had no access to modern contraceptives,

[13] Results from regressions of the duration between the marriage and the first birth (not shown) establish that there is no significant effect of the real wage, suggesting that the "starting" effect is deliberate. Furthermore, since the time span from the marriage to the first birth is a proxy of fecundity (Wrigley et al., 1997; Klemp and Weisdorf, 2012), these findings indicate that fecundity is not affected by the aggregate real wage over this time period.

and that marital births usually continued until the end of the wife's reproductive period (age 40-45). Yet, the fertility decline of the nineteenth century was partly achieved by parental prudency *within* the marriage, achieved by means of *coitus interruptus* (withdrawal), abstention, or extended breastfeeding (Coale and Watkins, 1986). The coefficients of columns 3 and 4 suggest that such behaviour was also common before the fertility decline. The real wage has a significant positive effect on the risk of a successive birth, meaning that low real wages increased the time elapsed between consecutive births. While column 3 reports the average effect of the real wage on all birth intervals, column 4 reports the effect on all but the last birth interval. By minimizing the possibility of intervals reflecting failed attempts at stopping, the estimate in column 4 implies that a one-standard deviation increase in the real wage increased the risk of a birth by 10%, corresponding to approximately three-months on average in the regression sample. In all subsequent "spacing" regressions, the last birth interval is excluded.

Turning to the "stopping" model (Table 4, column 5), we find no significant effect of the real wage on the risk of a last conception. Since the "stopping" time interval often cover periods of more than 25 years, the lack of a significant effect may not be surprising. However, this conclusion is robust to alternative ages in which the wife is considered as entering the risk of stopping.

Our covariates help to shed light on the role of socio-economic rank for historic birth patterns. The reference group in the specifications of Table 4 is 'labourers' (the poorest group). We find that, on average, the poor had longer birth intervals than the rich (farmers, merchants and gentry) but also, interestingly, that the poor stopped later than their more affluent counterparts (Table 4, columns 3 to 5). The fact that the risk of a further birth generally increases with family wealth, which was evident in Figure 4, is also sustained by the regression analysis.

The fact that the rich had more offspring than the poor, as also demonstrated by Clark and Hamilton (2006) and Boberg-Fazlic et al. (2011), is partly ascribable to shorter birth intervals (Table 4, columns 3 and 4). The early "stopping" among the rich may suggest that wealthy families had a target number of offspring (Table 4, column 5).[14]

[14] For a discussion on this see Van Bavel (2004).

Among the remaining covariates, it should be noted that female literacy is associated with comparatively shorter birth intervals and an early stopping age, even after controlling for socio-economic status. Not surprisingly, couples of higher fecundity, as proxied by a shorter time from marriage to first birth, have significantly shorter birth intervals than couples of low fecundity. Couples giving birth to prenuptially conceived children have a reduced risk for subsequent births. Also in line with our expectations, the death of sibling *n* during infancy (age 0 to 1) or in early childhood (age 1 to 3) substantially raises the risk of conception of sibling *n*+1, indicating that parents attempted to replace the deceased child. Finally, it appears that birth order has a significantly negative effect on the risk of a subsequent birth, meaning that birth intervals increase with the birth order of the child consistent with an age-related decline of fecundity (Baird et al., 2005).[15]

4. Robustness Checks

This section is devoted to testing the robustness of our findings. This involves estimating models with alternative wage series, testing for compositional effects, accounting for biological effects, allowing for migration, and, finally, exploring effect heterogeneity.

4.1 Alternative Wage Series

One potential concern is that the relationship between national real wages and birth intervals is a spurious correlation. To exclude this possibility, we performed placebo tests, shifting the real wage series forward by 3, 5, and 7 years. As can be seen in Table 5, this renders the coefficient of real wages on birth intervals small and insignificant, indicating that the baseline effect is not spurious.

[Table 5: Placebo test]

We also estimated our model with alternative wage series to ensure that our results are robust to the use of different or independently collected series. To ease comparison,

[15] We have also checked if there is an effect of child gender on birth intervals, but the coefficient was always insignificant. Hence, we proceeded to drop this variable.

column 1 of Table 6 replicates the baseline results. In column 2 we introduce cross-sectional variation by assigning skilled wages of building workers (Clark, 2007) to craftsmen, traders, farmers, merchants and gentry. The estimated coefficient is highly significant and the magnitude of the effect is almost identical to our baseline estimates. In column 3 we use an independently collected wage series of skilled workers (i.e., building workers) constructed by Robert Allen.[16] Although this wage series has considerably less variation than Clark's wage series, the coefficient estimate is still highly significant.

[Table 6: Alternative real wages]

4.2 Time Trend and Compositional Effects

Another potential concern is that the stratification of the hazard model by quarter century is not sufficient to completely account for secular changes that might affect both real wages and birth outcomes, thus resulting in a correlation which has no bearing on the existence of a causal relationship. To address this concern we estimate a model that adds a cubic time trend and a model with stratification by decade, respectively. The coefficient estimates reported in Table 7 (columns 1 and 2) show that the effect of the real wage on spacing is robust towards the use of a polynomial time trend and a more restrictive period stratification.

The long period under consideration causes sample attrition of some parishes which are thus not observable across the entire time-span. To address any potential compositional effect, we estimate a model with parish-quarter century fixed effects. The estimates of this model presented in column 3 show that the effect of real wages on spacing is basically unaffected by any compositional effects.[17] Consistently, estimating our baseline model for a subsample of 12 parishes covered for the whole period yields the same results (column 4).[18]

[16] This data is available at http://www.nuffield.ox.ac.uk/users/allen/data/labweb.xls.

[17] The use of parish-decade fixed effects was computationally unfeasible due to the vast number of interaction terms.

[18] We also estimated the model on a smaller random subset of the sample including family fixed effects (results available upon request). This model led to similar conclusions as the baseline regression. Given the size of the dataset, regressions involving family fixed effects are extremely computationally demanding, and as such we have not been able to estimate the model on the

4.3 Accounting for Biological Effects

A key concern is that our effect may be entirely, or to a large extent, due to biological causes, generating a correlation between wages and birth intervals that has no bearing on the existence of a deliberate choice. Wages close to a subsistence level can result in malnutrition and hence temporary infertility, leading to longer birth intervals. Furthermore, climatic conditions may potentially have effects on the fecundability of couples, leading to variation in birth spacing. In this subsection we try to account for factors that might influence our results through a biological channel controlling for years of potential malnutrition, high temperatures, and periods of high-mortality. To this end, we include three control variables: (*i*) wheat prices (Clark, 2007) which approximate food prices and therefore episodes of malnutrition; (*ii*) average yearly temperatures (Manley, 1953) to capture extreme climatic conditions and potential crop failures; and (*iii*) mortality rates to capture the disease environment. In the latter case, we use both crude death rates from Wrigley and Schofield (1981) covering the whole of England as well as stillbirth rates from our sampled parishes to capture local effects. We identify stillbirths in our sample as children who are baptized and buried on the same day. The data contains approximately 9,600 cases of stillbirths. We construct a parish-level stillbirth rate by dividing the number of stillbirths in a given parish in a given year by the total number of births, approximating the incidence of local and year specific miscarriages.

The results of this analysis are reported in Table 8. As expected, wheat prices exert a negative effect on the risk of a birth. This means that higher wheat prices increases the spacing between successive offspring (Table 8, column 1). Although malnutrition (here proxied by the price of wheat) may lead to temporary infertility (due to nutritional amenorrhea) and/or miscarriages, hence expanding the spacing between two consecutive births, the effect of the real wage is robust to the inclusion of wheat prices.[19] Furthermore, we find that higher average annual temperatures reduce the risk of a birth

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complete sample.

[19] To overcome the limitation of having national time series, we also have estimated "local" effects using quarterly wheat prices from Winchester on a sub-sample of parishes situated close to Winchester (namely Odiham, Reigate, Ash, Hartland, Morchard Bishop, Colyton, Bridford, Ipplepen, Dawlish). The estimates are consistent with the results presented above (available upon request). The sources of the quarterly wheat prices are provided by Nicholas Poynder, "Grain storage in theory and history", Table 1, a paper presented at the Third Conference of the European Historical Economics Society, Lisbon, October 29–30, 1999 (`www.iisg.nl/hpw/poynder.pdf`).

(column 2), although the coefficient is not significant. The point estimate is consistent with the findings of Lam and Miron (1996) showing that high temperatures lead to lower birth rates.[20] In our case the control for temperature, if anything, makes the effect of the real wage on birth intervals larger. In column 3, we use annual crude death rates to proxy for years of famine and high mortality, finding that high death rates reduce the risk of births, implying longer birth intervals. Also in this case the effect of real wages on spacing remains highly significant and of similar magnitude. In column 4, we control for the annual stillbirth rate at the parish level and we find a strong, negative effect on the risk of birth.[21] This suggests, reassuringly, that the stillbirth rate is also an accurate proxy for periods of malnutrition which might have led to miscarriages. The effect of the real wage on the spacing of births, however, is still highly significant and virtually unchanged in magnitude. Lastly, column 5 includes all the control variables used above. While the effect of wheat prices is now no longer significant (possibly due to the correlation with temperatures) the effects of mortality remain. If the effect of real wages on the spacing of births were entirely (or to a large extent) attributable to biological circumstances, the inclusion of these proxies should have weakened substantially the real wage effect. The set of results presented in Table 8, instead, strongly suggests that the effect of real wages on birth intervals was the result of a deliberate choice.

[Table 8: Accounting for food prices, climate, and mortality]

4.4 Accounting for Migration

As mentioned earlier, one weakness of the reconstitution data is the impossibility of following migrants. This can be problematic if the decision to migrate is correlated with real wages. We tackle this issue in a number of ways. First, we control for permanent migration by using binary variables indicating parents who have missing birth or death dates. Column 1 of Table 9 shows the estimates when controlling for permanent

[20] We also have experimented with different seasonal patterns in temperatures, but did not find any seasonal effects.

[21] Stillbirth rates require information about birth *as well as* death dates, causing us to use a sub-sample in Columns 4 and 5.

migrants.[22] Couples for which the husbands' birthday is unknown (e.g. immigrants) tend to have longer birth intervals, whereas the remaining migrants do not have significantly different birth intervals. Importantly, the effect of real wages on the spacing of births remains unaffected. In column 2, we estimate our model restricting the sample to couples whose dates of birth are unknown for both members. Similarly, column 3 reports the effects when we constrain the sample to couples whose dates of death are unknown for both members. The estimates show that, if anything, permanent movers are more responsive to changes in real wages than stayers.

Temporary migration, that is, families who move only provisionally to an unobservable parish, may also pose a problem. Periods of low wages may induce families to move outside the sampled parishes where subsequent births are not observed. We address this issue by excluding excessive birth intervals. Even in the absence of temporary migration, restricting our sample to short birth intervals will bias our estimates towards zero. Still, the significance of the estimates will inform us about the robustness of our findings towards the exclusion of temporary migrants. In column 4 we have constrained the sample to birth intervals strictly less than three years (approximately the 75th percentile). The coefficient on real wages remains highly significant at about half the magnitude of the baseline estimate. The effect, however, is sizable: a standard-deviation increase in the real wage reduces the spacing of birth by about 0.16 standard deviations. Furthermore, we take an even more restrictive approach (column 5), by constraining the sample to intervals strictly less than 2.5 years (close to the sample mean), still finding a significant effect. Note that all the specifications of Table 9 include the controls for wheat prices, temperature, and mortality as discussed above.[23]

Finally, we have followed Ruggles (1999) who proposed to restrict the sample to mothers who are born, married, and buried in the sampled parishes and who remain in the marriage until the age of 50. Together with their husbands these women constitute a "completed marriage". The estimates for completed marriages are reported in column 6.

[22] The effect of mothers whose date of birth is unknown is already captured by the set of controls of mother's age at marriage which includes a category for unknown age. That explains why the binary variable for immigrant mothers is not included in column 1 of Table 9.

[23] In this set of robustness checks stillbirths are excluded to avoid a further reduction of the number of observations. However, including the control for stillbirth in column 1 does not change the results.

The effect of real wages on spacing remains highly significant and is practically identical to the baseline estimate.

[Table 9: Accounting for migration]

4.5 Spacing by Occupational Group

The richness of our data allows us to estimate the effect of real wages by income groups inferred from the occupational status. Figure 4 and the baseline estimates in Table 4 establish a clear socio-economic gradient in birth spacing: affluent families tend to have shorter birth intervals than their less affluent counterparts. In Table 10 we present estimates by six socio-economic groups as categorized by Clark and Cummins (2010): labourers, husbandmen, craftsmen, traders, farmers, merchants and gentry.[24] The effect of the real wage on spacing is large and significant across almost all occupational groups. Importantly, the effect is also present among the most affluent groups: the merchants and the gentry. The fact that the effect of real wages is also present among the most affluent groups, even after controlling for variables capturing potential biological effects (wheat prices, temperatures, national and local death rates), suggests once more that the detected effect is likely to reflect a deliberate choice.

[Table 10: Estimates by socio-economic group]

Interestingly, farmers seem to react differently to changes in real wages as captured by the negative point estimate in column 5. Contrary to other groups, farmers may have benefited from higher wheat prices even though this meant lower real wages. If this is true, we would see an increasing risk of birth among farmers when wheat prices increase. In Table 11 we have interacted the price of wheat with the different occupational groups, finding that farmers did indeed respond differently to price changes than the other occupational groups. While the main effect of the wheat price is negative,

[24] Due to the low number of observations, we merge merchants and gentry.

the coefficient of the interaction term indicates that farmers have comparatively shorter birth intervals when the price of wheat increases.

[Table 11: Effect heterogeneity]

5. Concluding Remarks

Britain was the first country to escape Malthusian stagnation and to enter into a regime of modern economic growth. Late marriages and a high celibacy rate have been pointed to as viable explanations for Britain's low population-pressure, high-wage economy, as well as for its leadership in the Industrial Revolution. The consensus among scholars has been that marital birth control was absent in pre-modern England, and that it emerged only towards the end of the nineteenth century, when the fertility transition swept across Western Europe. In this paper we can reject this hypothesis by estimating duration models on detailed family-level data. Using a large sample of family reconstitutions data we show that couples adjusted birth intervals in response to changes in national real wages, suggesting that spacing was used as a means of fertility control throughout the period 1540-1850.

By using several specifications and accounting for wheat prices, temperatures, national and local mortality, we argue that the effect of the real wage on birth intervals is due to a deliberate choice and not to infertility or miscarriages caused by, for example, malnutrition or the disease environment. Although migration and parish sample attrition are potential threats to studies centred on family reconstitutions, we show that these issues are highly unlikely to have affected our results.

By demonstrating that it was not only unmarried individuals but also married couples that responded to changing economic conditions by means of postponing births, we offer strong support for Malthus' notion that England was a *preventive check* society, and by extension to the idea that this led to Britain's world leadership in the run up to the Industrial Revolution.

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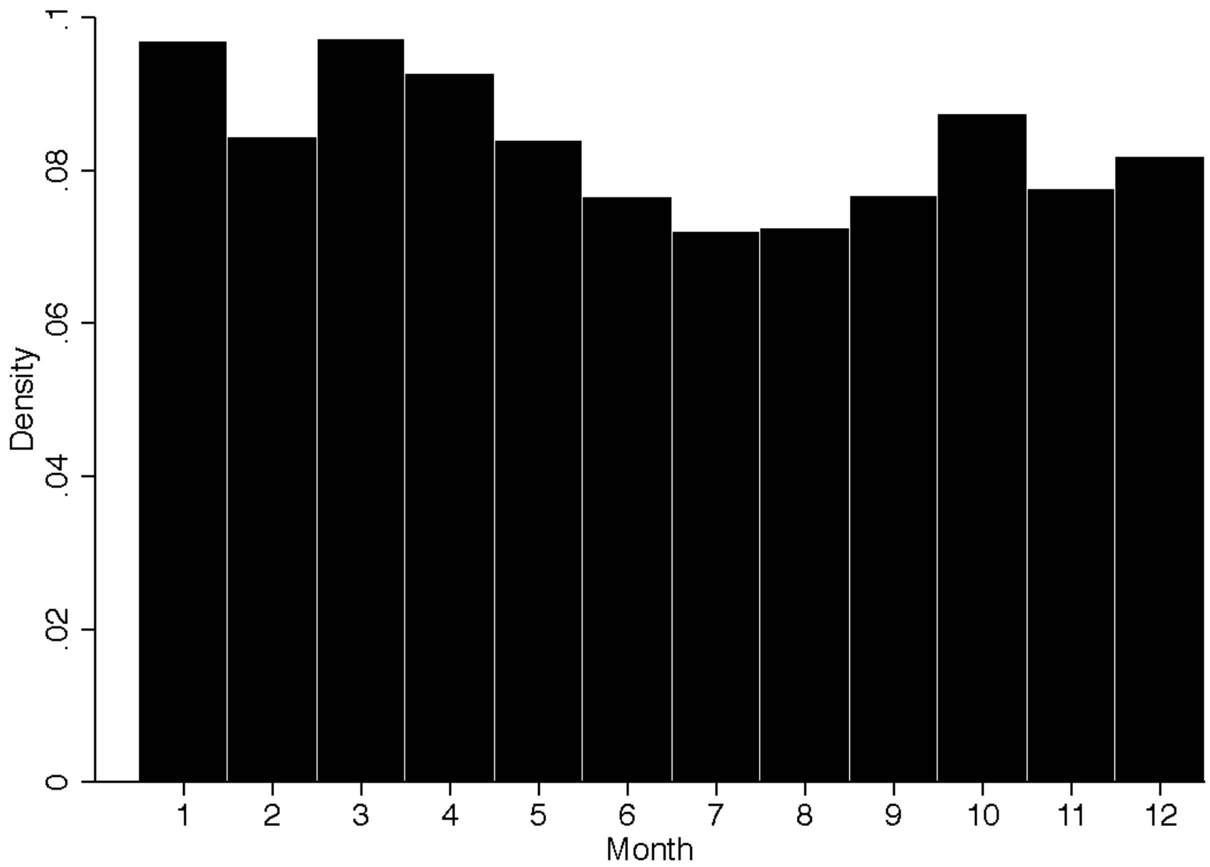

Figure 1: The distribution of births by month

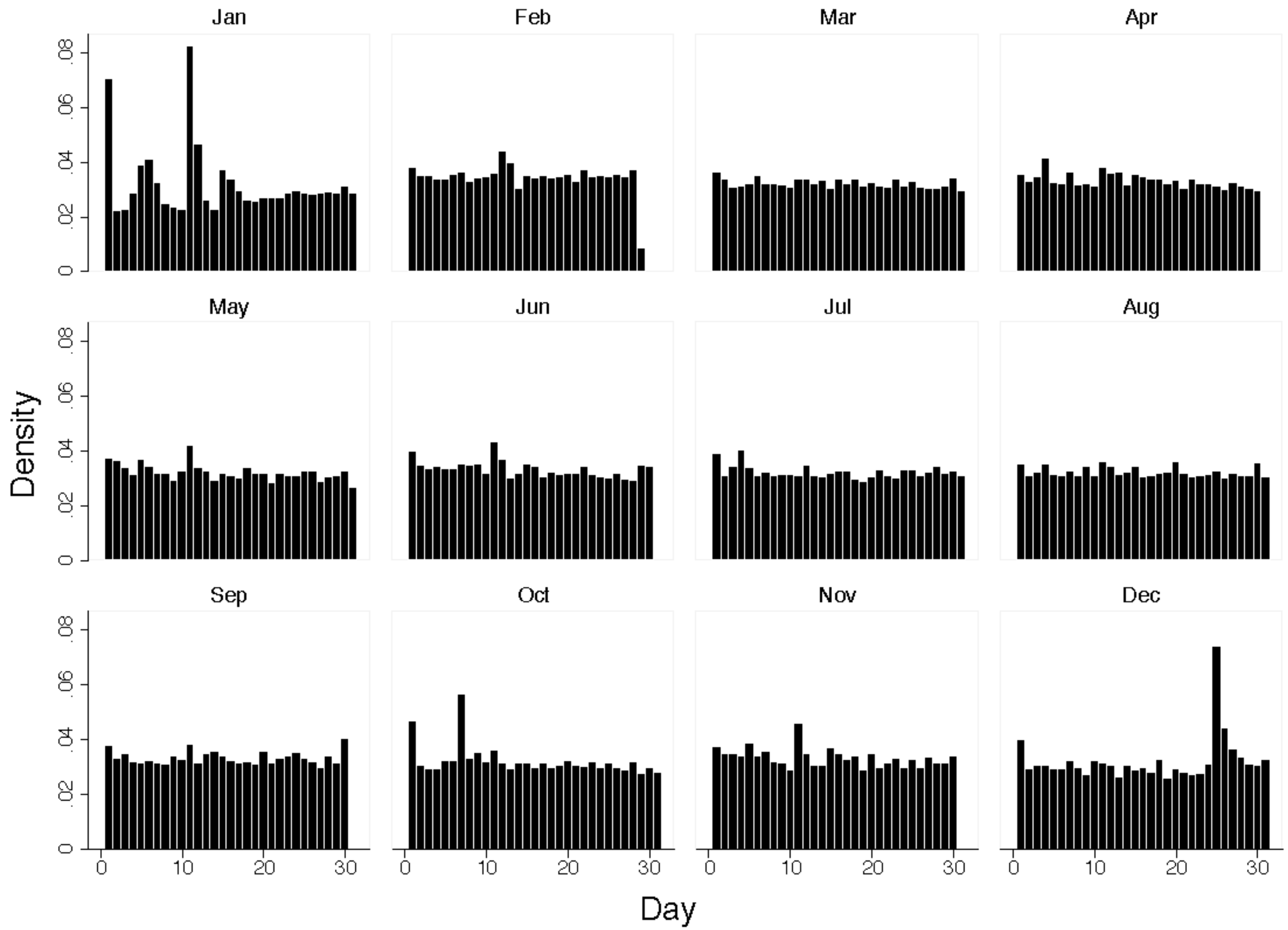

Figure 2: Distribution of births within the twelve months of the year

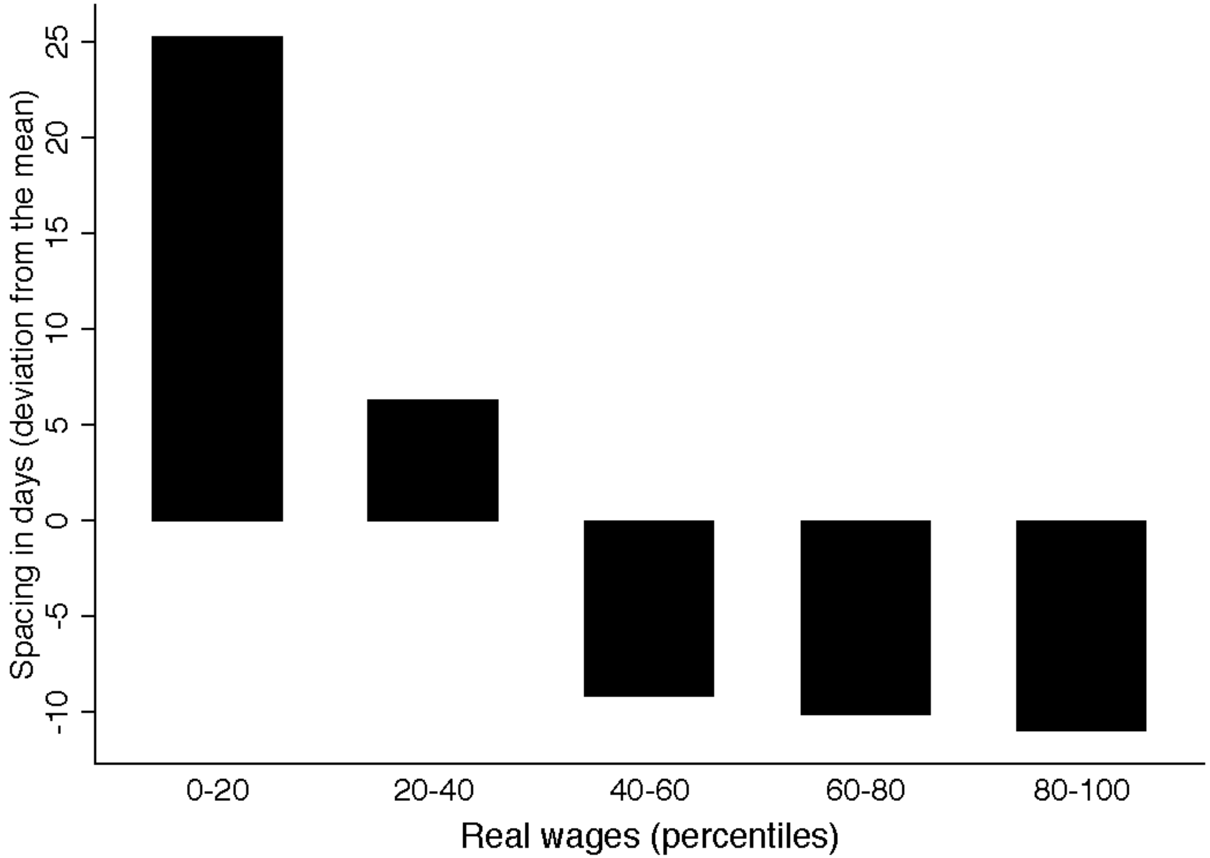

Figure 3: Average spacing by real wage percentiles

Figure 4: Average spacing by socio-economic group

Table 1: Summary statistics

Variable	Mean	Std. deviation	Min	Max	N
Spacing (days)	929.24	475.06	260	4368	191892
Mother’s age at marriage (years)	23.67	4.27	15.00	46.67	62515
Mother’s age at starting (years)	24.97	4.51	15.11	47.61	71556
Time to first birth (years)	1.19	1.13	-0.08	11.97	116220
Prenuptially conceived (share)	0.21	0.41	0	1	191892
Mother’s age at stopping (years)	38.41	5.86	16.79	49.99	71556
Labourers	0.15	0.36	0	1	191892
Husbandmen	0.08	0.28	0	1	191892
Craftsmen	0.10	0.30	0	1	191892
Traders	0.05	0.21	0	1	191892
Farmers	0.03	0.17	0	1	191892
Merchants	0.06	0.23	0	1	191892
Gentry	0.01	0.12	0	1	191892
Occupation unknown	0.51	0.50	0	1	191892
Mother’s age when giving birth (years)	30.01	5.87	15.11	49.00	71556
Mother literate	0.33	0.47	0	1	36126
Mother’s literacy unknown	0.81	0.39	0	1	191892
Birth order	3.08	2.14	1	19	191892
Number of siblings	6.17	2.70	2	21	191892
Child deceased age 0-1	0.14	0.34	0	1	191892
Child deceased age 1-3	0.06	0.23	0	1	191892
Child deceased unknown	0.59	0.49	0	1	191892

Source: Cambridge family reconstitution data.

Table 2: Average birth intervals (days) by birth order and period

Period	First interval	Second last interval	Last interval
1540–1699	830.4	936.0	1066.3
1700–1749	803.3	926.4	1076.6
1750–1799	798.2	922.9	1053.0
1800–1850	805.9	916.4	1005.3

Source: Cambridge family reconstitution data

Table 3: Summary statistics of aggregate variables

Variable	Mean	Standard deviation	Min	Max
Agricultural real wage	0.20	0.06	0.08	0.42
Wheat price	2.89	2.62	0.22	14.84
Mean temperature	9.21	0.66	6.84	10.82
Crude death rate	26.63	4.48	19.20	53.90
Stillbirth rate	0.04	0.05	0	1

Source: See text.

Table 4: Marriage, Starting, Spacing and Stopping

	(1) Marriage	(2) Starting	(3) Spacing	(4) Spacing w/o	(5) Stopping
Agri real wage	0.228*	0.221**	0.077***	0.095***	0.024
	(0.129)	(0.099)	(0.009)	(0.011)	(0.096)
Husbandmen	-0.031	-0.036	0.066***	0.074***	0.223***
	(0.032)	(0.032)	(0.014)	(0.017)	(0.078)
Craftsmen	-0.076***	-0.079***	0.072***	0.087***	0.111
	(0.027)	(0.028)	(0.013)	(0.016)	(0.068)
Traders	-0.039	-0.050	0.163***	0.186***	0.184*
	(0.042)	(0.036)	(0.020)	(0.022)	(0.104)
Farmers	-0.042	-0.072*	0.145***	0.222***	0.237**
	(0.045)	(0.039)	(0.019)	(0.022)	(0.101)
Merchant	-0.013	-0.039	0.197***	0.206***	0.216**
	(0.039)	(0.038)	(0.018)	(0.020)	(0.094)
Gentry	0.128	0.083	0.266***	0.306***	0.833***
	(0.086)	(0.071)	(0.031)	(0.035)	(0.223)
Occupation unknown	-0.105***	-0.126***	0.031***	0.068***	0.298***
	(0.024)	(0.025)	(0.012)	(0.013)	(0.067)
Mother literacy	-0.004	-0.009	0.067***	0.068***	0.212***
	(0.023)	(0.022)	(0.014)	(0.014)	(0.073)
Mother literacy unknown	-0.121***	-0.304***	-0.004	-0.008	0.107
	(0.037)	(0.026)	(0.017)	(0.020)	(0.083)
Time to first birth (years)			-0.058***	-0.052***	0.010
			(0.003)	(0.004)	(0.014)
Prenuptially conceived			-0.021***	-0.018*	0.019
			(0.008)	(0.009)	(0.041)
Birth order			-0.028***	-0.011***	.
			(0.002)	(0.002)	.
Child mortality (0-1)			0.661***	0.738***	-0.049
			(0.013)	(0.015)	(0.062)
Child mortality (1-3)			0.178***	0.161***	-0.153*
			(0.014)	(0.017)	(0.090)
Child mortality unknown			0.034***	0.028***	-0.085*
			(0.007)	(0.008)	(0.043)
Mother's age at marriage	No	No	Yes	Yes	Yes
Subjects	20040	22622	116030	85147	3795

Note: Cox proportional hazard model with time-varying real wages. Real wages are standardized with mean zero and unit standard deviation. Occupation refers to the father. Coefficients (semi-elasticities) reported. Standard errors in parenthesis are clustered by the year of the demographic outcome. Estimates are stratified by parish and quarter century. In Column 4 the last birth interval is not considered. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 5: Placebo test

	(1) Shifted 3 years	(2) Shifted 5 years	(3) Shifted 7 years
Agri real wage	-0.005	-0.006	-0.001
	(0.011)	(0.011)	(0.010)
Control variables	Yes	Yes	Yes
Subjects	85087	84942	84659

Note: Cox proportional hazard model with time-varying real wages. Real wages are standardized with mean zero and unit standard deviation. Coefficients (semi-elasticities) reported. Standard errors in parenthesis are clustered by the year of birth. Estimates are stratified by parish and quarter century. Control variables: father's occupation, mother's age at marriage, mother's literacy, time to first birth, prenuptially conceived, child mortality within household, and birth order. * $p < 0.1$, ** $p<0.05$, *** $p < 0.01$.

Table 6: Effect on spacing using alternative real wage series

	(1)	(2)	(3)
Agri real wage	0.095***		
	(0.011)		
Building real wage		0.107***	
		(0.012)	
Agri real wage (Allen)			0.057***
			(0.010)
Husbandmen	0.074***	0.075***	0.067***
	(0.017)	(0.017)	(0.017)
Craftsmen	0.087***	0.082***	0.085***
	(0.016)	(0.016)	(0.016)
Traders	0.186***	0.182***	0.169***
	(0.022)	(0.022)	(0.023)
Farmers	0.222***	0.220***	0.209***
	(0.022)	(0.023)	(0.023)
Merchants	0.206***	0.202***	0.201***
	(0.020)	(0.020)	(0.021)
Gentry	0.306***	0.302***	0.300***
	(0.035)	(0.035)	(0.035)
Occupation unknown	0.068***	0.070***	0.066***
	(0.013)	(0.013)	(0.014)
Mother literacy	0.068***	0.070***	0.080***
	(0.014)	(0.014)	(0.015)
Mother literacy unknown	-0.008	-0.007	-0.007
	(0.020)	(0.020)	(0.021)
Time to first birth (years)	-0.052***	-0.052***	-0.052***
	(0.004)	(0.004)	(0.004)
Prenuptially conceived (share)	-0.018*	-0.017*	-0.015
	(0.009)	(0.009)	(0.010)
Child deceased age 0-1	0.738***	0.737***	0.737***
	(0.015)	(0.015)	(0.016)
Child deceased age 1-3	0.161***	0.162***	0.168***
	(0.017)	(0.017)	(0.017)
Child deceased unknown	0.028***	0.028***	0.026***
	(0.008)	(0.008)	(0.008)
Birth order	-0.011***	-0.011***	-0.011***
	(0.002)	(0.002)	(0.002)
Mother's age at marriage	Yes	Yes	Yes
Subjects	85147	85147	81315

Note: Cox proportional hazard model with time-varying real wages. Real wages are standardized with mean zero and unit standard deviation. Occupation refers to the father. Coefficients (semi-elasticities) reported. Standard errors in parenthesis are clustered by the year of birth. Estimates are stratified by parish and quarter century. * p < 0.1, ** p < 0.05, *** p < 0.01.

Table 7: Testing for time trend and compositional effects

	(1) Polynomial time trend	(2) Decade FE	(3) Parish x quarter century FE	(4) W/o parish attrition
Agri real wage	0.092***	-0.085***	0.095***	0.101***
	(0.011)	(0.011)	(0.011)	(0.013)
Control variables	Yes	Yes	Yes	Yes
Subjects	85147	85147	85147	55541

Note: Cox proportional hazard model with time-varying real wages. Real wages are standardized with mean zero and unit standard deviation. Coefficients (semi-elasticities) reported. Standard errors in parenthesis are clustered by the year of birth. Control variables: father's occupation, mother's age at marriage, mother's literacy, time to first birth, prenuptially conceived, child mortality within household, and birth order. Column 1 includes a third order polynomial time trend. In column 2 the hazard model is stratified by decade. In column 3 we include parish-quarter century fixed effects. In column 4 we restrict the sample to 12 parishes observable throughout the whole period 1540-1850. * p < 0.1, ** p < 0.05, *** p < 0.01.

Table 8: Accounting for food prices, climate and mortality

	(1)	(2)	(3)	(4)	(5)
Agri real wage	0.081***	0.103***	0.088***	0.102***	0.098***
	(0.013)	(0.014)	(0.010)	(0.013)	(0.019)
Wheat price	-0.020*				0.002
	(0.012)				(0.017)
Avg yearly temperature		-0.002			-0.004
		(0.009)			(0.011)
Crude death rate			-0.007***		-0.007**
			(0.002)		(0.003)
Stillbirth rate				-0.435***	-0.572***
				(0.168)	(0.216)
Control variables	Yes	Yes	Yes	Yes	Yes
Subjects	85147	66135	85146	35142	26169

Note: Cox proportional hazard model with time-varying real wages, wheat prices, yearly temperatures, and crude death rates. Stillbirth rates are computed at the parish level. Real wages and wheat prices are standardized with mean zero and unit standard deviation. Coefficients (semi-elasticities) reported. Standard errors in parenthesis are clustered by the year of birth. Estimates are stratified by parish and quarter century. Control variables: father's occupation, mother's age at marriage, mother's literacy, time to first birth, prenuptially conceived, child mortality within household, and birth order. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 9: Accounting for migration

	(1) Migration	(2) Immigrants	(3) Emigrants	(4) Spacing < 3 years	(5) Spacing < 2.5 years	(6) Compl. marriage
Agri real wage	0.086***	0.165***	0.129**	0.039***	0.023*	0.092***
	(0.015)	(0.052)	(0.050)	(0.015)	(0.014)	(0.028)
Immigrant father	-0.031***					
	(0.009)					
Emigrant mother	-0.013					
	(0.009)					
Emigrant father	0.014					
	(0.010)					
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Subjects	66135	4346	8083	54122	44032	11390

Note: Cox proportional hazard model with time-varying real wages, wheat prices, yearly temperatures, and crude death rates. Real wages and wheat prices are standardized with mean zero and unit standard deviation. Coefficients (semi-elasticities) reported. Standard errors in parenthesis are clustered by the year of birth. Estimates are stratified by parish and quarter century. Control variables: father's occupation, mother's age at marriage, mother's literacy, time to first birth, prenuptially conceived, child mortality within household, and birth order. The column "Immigrants" considers mothers and fathers whose date of birth is unknown but with a known date of death. The column "Emigrants" considers mothers and fathers whose date of death is unknown but with a known date of birth. * p < 0.1, ** p < 0.05, *** p < 0.01.

Table 10: Effect of real wages on spacing by socio-economic group

	(1) Labourers	(2) Husbandmen	(3) Craftsmen	(4) Traders	(5) Farmers	(6) Merchants & Gentry
Agri real wage	0.084***	0.095**	0.092***	0.050	-0.063	0.152***
	(0.031)	(0.041)	(0.033)	(0.053)	(0.075)	(0.053)
Wheat price	-0.024*	0.000	-0.017	-0.017	-0.050	0.072
	(0.014)	(0.036)	(0.026)	(0.045)	(0.047)	(0.054)
Avg yearly temperature	0.007	-0.010	0.022	0.005	-0.008	0.009
	(0.015)	(0.021)	(0.020)	(0.033)	(0.043)	(0.021)
Crude death rate	-0.006	-0.014**	-0.008	-0.008	-0.022**	-0.015**
	(0.005)	(0.006)	(0.006)	(0.007)	(0.009)	(0.007)
Control variables	Yes	Yes	Yes	Yes	Yes	Yes
Subjects	14121	6637	7997	2954	2035	5032

Note: Cox proportional hazard model with time-varying real wages, wheat prices, temperatures, and crude death rates. Real wages and wheat prices are standardized with mean zero and unit standard deviation. Coefficients (semi-elasticities) reported. Standard errors in parenthesis are clustered by the year of birth. Estimates are stratified by parish and quarter century.
Control variables: mother's age at marriage, mother's literacy, time to first birth, prenuptially conceived, child mortality within household, and birth order.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 11: The effect of food prices - Interaction terms

	(1) Interaction term
Wheat price	-0.050***
	(0.010)
Husbandmen	0.081***
	(0.021)
Craftsmen	0.105***
	(0.019)
Traders	0.203***
	(0.028)
Farmers	0.198***
	(0.031)
Merchants	0.192***
	(0.027)
Gentry	0.236***
	(0.047)
Occupation unknown	0.059***
	(0.016)
Husbandmen x wheat	-0.007
	(0.018)
Craftsmen x wheat	-0.025
	(0.017)
Traders x wheat	0.000
	(0.021)
Farmers x wheat	0.045*
	(0.025)
Merchants x wheat	-0.027
	(0.030)
Gentry x wheat	-0.020
	(0.055)
Occupation unknown x wheat	0.003
	(0.012)
Control variables	Yes
Subjects	66135

Note: Cox proportional hazard model with time-varying wheat prices, temperatures, and crude death rates. Wheat prices are standardized with mean zero and unit standard deviation. Coefficients (semi-elasticities) reported. Standard errors in parenthesis are clustered by the year of birth. Estimates are stratified by parish and quarter century. Control variables: mother's age at marriage, mother's literacy, time to first birth, prenuptially conceived, child mortality within household, and birth order.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.