"School's out for summer, school's out forever": the long-term health consequences of leaving school during a bad economy.*

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Preliminary version. Please do not quote.

Abstract

This paper investigates whether leaving school in a bad economy deteriorates health in the long-run. We focus on individuals in England and Wales who left full-time education in their last year of compulsory schooling immediately after the 1973 oil crisis. Our identification strategy builds on two sources. First, it relies on the comparison of very similar individuals – born the same year – whose school-leaving behaviour in different economic conditions was exogeneously induced by compulsory schooling laws. More specifically, within a same birth cohort, pupils born at the end of the calendar year (September to December) were forced to leave school almost a year later than pupils born earlier in the year (January to August). Second, we exploit the sharp increase in unemployment rates generated by the 1973 oil crisis. Between 1974 and 1976, each school cohort indeed faced worse economic conditions at labour-market entry than the previous one. Unlike school-leavers who did postpone their entry on the labour market during the 1980s and 1990s recessions, we provide evidence that pupils' decisions to leave school at compulsory age between 1974 and 1976 were not endogeneous to the contemporaneous economic conditions at labour market entry. We use a repeated cross section of individuals over 1983-2001 from the General Household Survey (GHS) and take a life-course perspective, from 7 to 26 years after school-leaving. Our results show that men who left school in a bad economy have a higher probability of smoking over the whole period (1983-2001) and of having ever smoked. Women who left school in a bad economy are more likely to report poorer health over the whole period under study. They also have a higher probability to restrict their activity due to illness or injury and to consult the General Practitioner. We do not find any significant effects of poor economic conditions at labour-market entry on subsequent labour-market, marriage and fertility outcomes.

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1 Introduction

The youth has suffered disproportionately during the Great Recession (Bell and Blanchflower, 2011). Young cohorts who left full-time education in the late 2000s faced historically high unemployment rates and have experienced difficulties accessing employment. Recent labour economic research shows that poor macroeconomic conditions at labour-market entry lead to persistent and negative career effects. As career and labour-market outcomes do influence health[1], one may expect that leaving school in a bad economy has a negative effect on long-term health. This is what we investigate in this paper.

Over the last few years, a growing strand of the literature has studied the long-term labour-market consequences of graduating in a recession. According to Kahn (2010), those who graduate in bad economies may suffer from underemployment and are more likely to experience job mismatching since they have fewer jobs from which to choose. There is indeed evidence that graduating from college in a recession has a large, negative and persistent effect on men's wages in the USA and Canada (Kahn, 2010; Oreopoulos et al., 2012). According to Kahn (2010), the catch-up process for wages is as long as 15 years. Similarly, Kondo (2007) finds a negative effect of a recession at labour-market entry on wages in the USA, although the effect is weaker for women than for men. Graduating in a recession also has adverse effects on men's probability of being employed, especially among the low-educated (Gaini et al., 2012; Genda et al., 2010; Stevens, 2007), although this negative effect generally fades out over the next few years (except in Japan – see (Genda et al., 2010)). Hershbein (2012) shows that high-school graduate women (but not men) who graduated in an adverse labour-market are less likely to be in the workforce in the next four years, but that longer effects are minimal. Workers who enter firms in economic downturns may also initially be placed in lower-level jobs with less important tasks and less promotions (Gibbons and Waldman, 2006). For instance, Schoar and Zuo (2011) show that economic conditions when CEOs enter the labor market have a long-lasting impact on their career paths and managerial styles – they are less promoted and are in less prestigious occupations than their luckier counterparts. Although there is not much work to date on these aspects, graduating in a recession may also have negative effects on other dimensions of job quality e.g. job stress, perceived job security, working hours or more generally working conditions.
Hence, a large consensus has been reached according to which adverse economic conditions when leaving full-time education have negative consequences on labour-market outcomes. In particular, highly-skilled workers and individuals with strong attachment to the labour force seem to suffer from larger penalties.

There is both theoretical and empirical evidence that career outcomes are linked to health. Income is generally thought to improve health (Currie, 2009; Duleep, 1986; Gardner and Oswald, 2007; Grossman, 1972), job loss is associated with lower health, adverse health behaviours and higher mortality rates (Browning and Heinesen, 2012; Deb et al., 2011; Eliason and Storrie, 2009a; Salm, 2009; Sullivan and Von Wachter, 2009), while other job dimensions – such as job stress, perceived job insecurity, long working hours, harmful working conditions, downward occupational mobility – have been shown to deteriorate health (Caroli and Godard, 2014; Fischer and Sousa-Poza, 2009; Fletcher et al., 2011; Llena-Nozal, 2009;

[1] See for instance Fischer and Sousa-Poza (2009); Grossman (1972); Llena-Nozal (2009); Sullivan and Von Wachter (2009).

Robone et al., 2011). Whether leaving school in a bad economy has long-term consequences on health is thus a relevant question.

Previous studies have investigated the short-term health effect of contemporaneous economic fluctuations. They point to health and health behaviours being countercyclical, at least in the short-run (Buchmueller et al., 2007; Gerdtham and Ruhm, 2006; Neumayer, 2004; Ruhm, 2000, 2003, 2004, 2005). In reality, researchers' findings on this issue are mixed, as other studies find opposite or zero effects (Charles and DeCicca, 2008; Dave and Kelly, 2012; Dee, 2001; Economou et al., 2008). The latter studies seem to suggest that the impact of contemporaneous macroeconomic conditions is highly heterogeneous across worker's ex-ante employment probabilites, with individuals predicted to be at higher risk of unemployment being more likely to adopt risky health behaviours in bad times.

Our focus, however, is decidedly different : rather than examining the short-term health consequences of contemporaneous economic fluctuations, we investigate whether experiencing a bad economy at the time of the school-to-work transition – which is a critical moment in the lifecourse – is harmful to health in the long-run.

In this paper, we examine the impact of leaving full-time education in a bad economy on middle and long-term health in England and Wales. We focus on individuals who left full-time education in their last year of compulsory schooling[2] immediately after the 1973 oil crisis. Our identification strategy builds on two sources. First, it relies on the comparison of very similar individuals – born the same year – whose school-leaving behaviour in different economic conditions was exogeneously induced by compulsory schooling laws (both school entry and exit rules). More specifically, within a same birth cohort, pupils born at the end of the calendar year (September to December) were forced to leave school almost a year later than pupils born earlier in the year (January to August). We then exploit the sharp increase in unemployment rates generated by the 1973 oil crisis. Between 1974 and 1976, each school cohort indeed faced worse economic conditions at labour-market entry than the previous one. As a consequence, unlucky pupils born in September-December faced higher unemployment rates at labour-market entry than pupils born in January-August of the same calendar year.

Prior research has linked schooling choice to decreased labour-market opportunities (Betts and McFarland, 1995; Card and Lemieux, 2001; Gustman and Steinmeier, 1981). More specifically, there is evidence that individuals tend to remain in school during economic downturns. This could be a problem if pupils in their last year of compulsory education tend to enroll in an additional year of schooling when the economy deteriorates. We show, however, that – within a birth cohort – pupils' decisions to leave school at compulsory age between 1974 and 1976 were not endogeneous to the contemporaneous economic conditions at labour-market entry – unlike school-leavers during the 1980s and 1990s recessions. We argue that the 1973 oil crisis was highly unexpected and that pupils who were in their last year of schooling at that time did not anticipate the adverse career effects of leaving school when unemployment rates were high.

[2]In the early 1970s only about 50 percent of pupils stayed on in full-time education after having reached the minimum school-leaving age. This proportion was remarquably low by OECD standards (Micklewright et al., 1989).

We use a repeated cross section of individuals over 1983-2001 from the General Household Survey (GHS) and take a life-course perspective (from 7 to 26 years after school-leaving). We investigate the middle to long-term impact of leaving school in a bad economy on health status, health care and health behaviours. We examine the plausible channels by which adverse economic conditions at school-leaving may affect later health, e.g. labour-market outcomes as well as marriage and fertility.[3] Our results show that men who left school in a bad economy have a higher probability of smoking over the whole period (1983-2001) and of having ever smoked. Women who left school in a bad economy are more likely to report poorer health over the whole period under study. They also have a higher probability to restrict their activity due to illness or injury and to consult the General Practitioner. We do not find any significant effects of poor economic conditions at labour-market entry on subsequent labour-market, marriage and fertility outcomes.

This paper relates to several strands of literature. First and foremost, it contributes to the emerging literature investigating the long-term health consequences of graduating in a bad economy. To our knowledge, only a very limited number of studies (Cutler et al., 2014; Hessel and Avendano, 2013; Maclean, 2013) have addressed this question. So far, results turn out to be mixed. Maclean (2013) uses US data – the National Longitudinal Survey of Youth 79 (NLSY79) – and exploits the variation in school-leaving state unemployment rates to identify the effect of leaving school in a bad economy on health at age 40. Members of her sample left school between 1976 and 1992. As time or location of school-leaving may be endogeneous to the contemporaneous unemployment rate, she uses instrumental-variable (IV) methods to deal with selection problems related to what she refers to as "endogeneous sorting". She finds that men who left school when the state unemployment rate was high have a higher probability to report poor or fair health as well as depressive symptoms and have lower physical functioning at age 40. Surprisingly, she finds that women leaving school in a bad economy tend to have fewer depressive symptoms at age 40. Hessel and Avendano (2013) use European data, namely the Survey of Health, Ageing and Retirement in Europe (SHARE). They consider individuals aged 50 and over who left school from 1957 onward. They use country-specific unemployment rates and trend deviations based on the reported year of leaving full-time education. According to their results, poor conditions at school-leaving predict worse health status among women and better health status among men. They provide evidence that highly-educated women are particularly affected. However, the authors acknowledge that both selection into higher education and causation mechanisms may explain this association. Finally, Cutler et al. (2014) use the Eurobarometer data and consider economic fluctuations over 50 years across 31 countries. They show that higher unemployment rates at graduation are associated with lower income, lower life satisfaction, greater obesity, more smoking and drinking later in life, for both men and women. According to their results, education seems to play a protective role, especially when unemployment rates are high. In a series of recent papers Maclean (2014a,b,c) specifically tests whether leaving school in an economic downturn persistently

[3] See Gustman and Steinmeier (1981); Gutiérrez-Domènech (2008); Hashimoto and Kondo (2012); Hofmann and Hohmeyer (2014); Kondo (2012); Wolbers (2007) on the relationship between economic conditions at labour-market entry and subsequent family formation.

affects drinking behaviour, body weight and the probability of access to an employer-sponsored health insurance. She uses the same methodology and data as in Maclean (2013) and finds that men, but not women, who leave school in a bad economy consume more drinks and are more likely to report heavy and binge drinking than otherwise similar men. Unlucky men also have lower bodyweight and are less likely to be overweight and obese at age 40. Finally, she finds that both men and women are less likely to have access to an employer-sponsored health insurance (up to 18 years after school-leaving).

Our paper contributes to the existing literature in several ways. First, our identification strategy is different in spirit from the ones previously used in the literature. Rather than considering long periods of economic fluctuations and exploiting the variation in country (or state) school-leaving unemployment rates, we focus on two birth cohorts only – the 1958 and 1959 cohorts. Our strategy relies on the comparison of similar individuals – born the same year – whose school-leaving behaviour in different economic conditions was exogeneously induced by compulsory schooling laws. As a consequence, our results cannot possibly be biased by country-specific (or state-specific) cohort effects. Second, we show that – within each birth cohort – pupils' decisions to leave school at compulsory age between 1974 and 1976 were not endogeneous to the contemporaneous economic conditions at labour-market entry – unlike school-leavers during the 1980s and 1990s recessions. There is no need, then, to deal with problems related to endogeneous sorting of school-leaving and our results do not rely on the usual assumptions when implementing instrumental-variables models. Third, our data allow us to adopt a lifecourse perspective, which is only present in the paper by Cutler et al. (2014). Finally, we focus on low-educated individuals. There are some good reasons to focus on pupils leaving school at binding age : first, they represent a sizeable proportion of pupils in England and Wales in the mid-70s (approximately 50%). Second, whether they should be more affected than highly-educated individuals by high unemployment rates at labour-market entry remains an open question. On the one hand, economic theory predicts less persistence of poor economic conditions at school-leaving for low-skilled workers and those with weak attachment to the labour force. On the other hand, education has been hypothesised to increase one's ability to cope with negative shocks and uncertainty (Cutler et al., 2014; Cutler and Lleras-Muney, 2006). Our paper contributes to provide a better understanding of the long-term health consequences of leaving school in a bad economy among the low-educated.

The rest of the paper is organised as follows. Sections 2 presents the institutional framework and Section 3 the empirical approach. Section 4 describes the data that we use. Section 5 reports our results and Section 6 concludes.

2 Institutional framework

We investigate the long-term health consequences of leaving full-time education in a bad economy. We focus on low-educated individuals. More specifically, we focus on pupils who left full-time education as soon as they were eligible to do so – i.e. pupils who left school in their last year of compulsory schooling.

Our identification strategy builds on two sources. First, we exploit the presence of compulsory schooling laws (both school entry and school exit rules) in England and Wales. Second, we take advantage of the sharp and unexpected increase in unemployment rates immediately after the 1973 oil crisis.
Section 2.1 describes the compulsory schooling laws in England and Wales and section 2.2 provides graphical evidence of the sharp increase in unemployment rates after the 1973 oil crisis.

2.1 Compulsory schooling in England and Wales

The British compulsory schooling laws specify the maximum age at which pupils have to start school and the minimum age at which pupils are allowed to leave school.
The official school-starting age is the beginning of the term starting after the child's fifth birthday. Hence, entry rules determine that a school cohort consists of children born between the first day of September and the last day of August in the following calendar year (Del Bono and Galinda-Rueda, 2007). In other words, due to the discontinuity introduced by the school-entry rule, students within a same birth cohort belong to different school cohorts. There is evidence that compliance with school-entry requirement is almost perfect and that grade repetition (or grade skipping) is almost non-existent in England and Wales (Grenet, 2013; Sharp et al., 2002).

The current school leaving age of 16 was increased twice in England and Wales[4], from age 14 to 15 in 1947 and from age 15 to 16 in 1972.[5] The proportion of children leaving education at the first legal opportunity in the UK is high by the standards of other industrialised countries (Micklewright et al., 1989). In the early 1960s, only about 20% of pupils stayed in full-time education after having reached the minimum school-leaving age (Del Bono and Galinda-Rueda, 2007; McVicar and Rice, 2001). In our data, this proportion amounts to 50% in the mid-1970s. After the 1972 Raising Of the School-Leaving Age (ROSLA), students in their last year of compulsory schooling were normally attending secondary school (Year 11) while the less academically inclined were in vocational training. Two types of qualifications could be obtained at the end of Year 11 : the General Certificate of Education Ordinary Level (GCE O level) or the Certificate of Secondary Education (CSE). Both credentials were awarded at the end of junior secondary school, after an examination (Grenet, 2013).
Unlike other countries – and unlike the USA –, the implementation of compulsory schooling in England and Wales differs in that a student is not allowed to leave school on the exact date (birthday) in which she reaches the school-leaving age. Between school years 1963-1964 and 1996-1997, (see the Education Act of 1962, Appendix D), the rules governing school exit implied that pupils who reached age 16 between the 1^{st} of September and the 31^{st} of January had to complete their education until the following Easter. Students who reached the age of 16 between the 1^{st} of February and the end of August were forced to leave school at the end of the summer term, typically in May/June. Pupils born between the end of the summer term and August – i.e. pupils born in July or August – were thus allowed to leave school before

[4]The education system in Scotland is different and is not considered here.
[5]Several studies use these changes in minimum school-leaving age to identify the returns to education on labour market outcomes and health (Clark and Royer, 2013; Devereux and Hart, 2010; Grenet, 2013; Harmon and Walker, 1995; Oreopoulos, 2006). Overall, results tend to show that while an additional year of schooling induced by the reform has a positive impact on educational attainment and earnings, it does not seem to have any effect on health.

their 16^{th} birthday, i.e. at age 15.

To show how these exit rules support our identification strategy, we present in Figure 1 the authorised school-leaving date with respect to students' month-year of birth. It makes it clear that students born in the same calendar year belonged to different school cohorts due to the discontinuity introduced by the school entry rule (see column 3). Figure 1 provides evidence that, within the same birth cohort, the oldest pupils – born between January and August – were allowed to leave school at Easter or in May/June of year t whereas the youngest – born between September and December – were not allowed to leave school until the following Easter of year *t+1*. Figure 2 provides an illustration of how the compulsory schooling rules operate by taking the 1958 birth cohort as an example. Note that, due to the discontinuities introduced by both school-entry and school-exit rules, pupils who left school at binding age had a fairly similar quantity of schooling (in months) at the end of full-time education regardless of their month of birth.

2.2 Unemployment rates

The sharp and unprecedented increase in the oil price from three to ten dollars a barrel in October 1973 had serious effects on the balance of payments of the industrial nations, which were oil-importer countries. This first world-wide recession had strong effects on unemployment rates in a number of industrialised countries, including the UK (Bhattarai, 2011).[6]
Figure 3 provides evidence of the sharp increase in unemployment rates after the 1973 oil crisis. The 1973 oil crisis – which occured in october 1973 – is symbolised by the vertical dark blue line on the left-hand side. The blue line shows the unemployment rates for all individuals aged 16 and over[7] and the red and green lines show the unemployment rates for men and women respectively. As shown in Figure 3, unemployment rates gradually increased starting from 1974 until 1978 – when the economy recovered – with the sharpest increase between 1974 and 1976. As made clear by the graph, each school cohort faced significantly higher unemployment rates than the previous school cohort.[8]

[6] Thus, it can reasonably be argued that the 1973 crisis was not endogenous to health in the UK.

[7] Unemployment rates for individuals aged 16 and over are provided by the Office for National Statistics (ONS). Unfortunately, we cannot produce unemployment rates for the 16-25 on a monthly basis for the period under study as they are not available from the ONS. Rather, we compute unemployment rates for the 16-25 using the 1975 and 1977 waves of the UK Labour Force Survey (LFS). Unemployment rates among the 16-25 are high and increase from 7.34% in 1975 to 9.05% in 1977, corresponding to a 23% increase within a two-year period. This increase is in the same range of magnitude as the increase in unemployment rates experienced by individuals aged 16 and over – from 4.5% in 1975 to 5.6% in 1977, corresponding to a 24% increase.

[8] The unemployment rate for all individuals aged 16 and over increased by 0.7 percentage point between Easter/May/June 1974 and Easter/May/June 1975. It increased by 1.1 percentage point between Easter/May/June 1975 and Easter/May/June 1976. This increase was somewhat milder between 1976 and 1977 as well as between 1977 and 1978 (a 0.1 percentage point increase in both cases).

3 Empirical approach

Section 3.1 presents our identification strategy as well as the model we estimate and Section 3.2 discusses the validity of the identification strategy.

3.1 Identification strategy and model

We consider pupils who left school at minimum school-leaving age and who entered the labour market between Easter 1974 and Easter 1976, i.e. the 1958 and 1959 birth cohorts. We do not consider older individuals so as to abstract from the effect of the increase in school-leaving age from 15 to 16 from September 1972 on.[9]

Our identification strategy relies on the comparison of similar individuals – born the same year – whose school-leaving behaviour in different economic conditions was exogeneously induced by compulsory schooling laws (both school entry and exit rules). More specifically, within a same birth cohort, pupils born at the end of the calendar year (September to December) were forced to leave school almost a year later than pupils born earlier in the year (January to August). We exploit the fact that between 1974 and 1976, each school cohort faced worse economic conditions at labour market entry than the previous one (due to the sharp increase in unemployment rates generated by the 1973 oil crisis). Thus, within each birth cohort, pupils born between September and December faced higher unemployment rates at labour-market entry than pupils born between January and August. Note that our identification strategy does not rely on the comparison on individuals who left school *before* and *after* 1973.[10] In our setup, all individuals are affected by the 1973 oil crisis. However, some pupils (the "treated") left school in worse conditions than otherwise similar pupils.[11]

We use a repeated cross-section of individuals over 1983-2001 to estimate the following equation by probit, for men and women separately :

$$H_i = \alpha + \gamma T_i + BirthYear_i + InterviewYear_i + \epsilon_i \quad (1)$$

where H_i denotes the health status of individual i and is measured as a dummy variable. T_i is a dummy variable taking value 1 if individual i is treated, i.e. born between the 1^{st} of September and the 31^{st} of December and value 0 if non-treated, i.e. born between the 1^{st} of January and the 31^{st} of August. $BirthYear_i$ is a dummy variable for individual i's year of birth. $InterviewYear_i$ is a dummy variable for individual i's interview year[12] and ϵ_i is the error term.

[9] In our setup, all individuals are affected by the 1972 ROSLA reform. Our identification strategy does not rely on the comparison on pre-reform cohorts and post-reform cohorts.

[10] This is because pupils who left school at minimum school-leaving age before and after the 1973 oil crisis are not comparable. Those who left school at binding age *before* the crisis were 15 years old, while those who left school at binding age *after* the crisis were 16 years old (due to the 1972 ROSLA reform). This difference in years of education makes it difficult to attribute the differences in health outcomes to economic conditions at labour-market entry.

[11] The 1973 oil crisis had a disproportionate impact on some regions – typically in Wales and the North of England. However, we do not use this additional regional variation. This is because (i) we do not have reliable data at a sufficiently disaggregated level on UK regional unemployment rates in the 1970s (ii) we do not have information on the actual region in which the individual lived at age 16.

[12] We control for *InterviewYear* to account for the fact that we observe individuals at different points in time. We choose

Within each birth cohort, treated individuals (born September-December) are younger than non-treated pupils (born January-August). As age and health are negatively correlated, not taking into account this age difference − which is a difference in months within a birth cohort − may lead us to underestimate the negative impact of leaving school in a bad economy. To account for it, we introduce in Equation (2) $f(BirthMonth_i)$, a function of age in months. We then estimate the following equation by probit, for men and women separately :

$$H_i = \alpha + \gamma T_i + BirthYear_i + f(BirthMonth_i) + InterviewYear_i + \epsilon_i \tag{2}$$

As including a quadratic function of age in the model seems to fit our data better (according to the Akaike Information Criterion), we define $f(BirthMonth_i)$ as a quadratic function of age in months within a birth year. It is equal to $(12 - BirthMonth_i) + (12 - BirthMonth_i)^2$, where $BirthMonth_i$ denotes the month of birth of respondent i and varies from 1 to 12.[13] The results shown later on in Section 5 are estimated following Equation (2), where $f(BirthMonth_i)$ is defined as a quadratic function of age.[14]

Equation (2) estimates the average effect of leaving school in a bad economy on health over the whole period ($\widehat{\gamma}$). However, our empirical approach also allows us to take a life-course perspective. More specifically, for each year of interview, we can compute the marginal effect of the treatment. This allows us to investigate whether the impact of poor economic conditions at labour-market entry on health is driven by middle or long-term effects.

A key assumption is that pupils in their last year of compulsory schooling do not stay strategically in school when the economy deteriorates, i.e. do not engage in what we refer to as "endogeneous timing". More specifically, if pupils anticipate the adverse effects of leaving school in a bad economy and enroll in an additional year of schooling, our estimates will be biased. We discuss this point in section 3.2.1.
A second identifying assumption is that if there are no other institutional differences within each birth cohort generating differences in health among the treated and the non-treated apart from school-exit rules (see sections 3.2.2, 3.2.3 and 3.2.4 for a discussion on school-entry rules and section 3.2.5 for a discussion on the differential incentives to take GCE O-level/CSE examinations induced by the January/February discontinuity), we can safely attribute observed differences in health to the impact of labour-market conditions at labour-market entry. To the extent that individuals born between January and August and individuals born between September and December are identical in all observable and

to include a dummy indicating the year in which an individual is interviewed rather a dummy indicating the survey wave in which she is interviewed. This is because a survey wave can be conducted over several years – usually two.

[13] One may worry that introducing simultaneously variables T_i, $(12-BirthMonth_i)$ and $(12-BirthMonth_i)^2$ in Equation (2) may lead to multicollinearity issues. When estimating Equation (2) with T_i, $BirthYear_i$, $InterviewYear_i$ and $(12-BirthMonth_i)$ as control variables, the VIF (Variance Inflation Factor) criterion is lower than 10 for all variables, thus suggesting inconsequential multicollinearity (see the rule of thumb provided by Hair et al. (1995)). The inclusion of $(12 - BirthMonth_i)^2$ as an additional variable leads to VIFs higher than 10 for variables $(12 - BirthMonth_i)$ and $(12 - BirthMonth_i)^2$. This is expected, as these two variables are highly likely to be correlated, as $(12 - BirthMonth_i)^2$ is a power of $(12 - BirthMonth_i)$. According to Allison (2012) this type of multicollinearity can be safely ignored and has no adverse consequences on the estimated results.

[14] As expected, estimating Equation (1) or Equation (2) where $f(BirthMonth_i)$ is simply defined as a linear function of age yields very similar estimates, although of lower magnitude and less significant. Results are not shown but available upon request.

unobservable characteristics (see section 3.2.6 for a discussion of the potential effects of season of birth) the differences in health status will be driven only by school-exit rules and hence different unemployment rates at labour-market entry, thus allowing us to identify the health consequences of leaving school in a bad economy.

3.2 Validity of the identification strategy

3.2.1 Endogenous timing of school-leaving

Time of school-leaving may be endogeneous to the contemporaneous economic conditions. The sign of the bias arising from endogeneous timing is difficult to predict. On the one hand, school-leavers who avoid leaving school in a bad economy may have unobserved characteristics (e.g. financial resources) that allow them to postpone their entry on the labour market. On the other hand, it is likely that only the most capable and hardworking are able to leave school during a bad economy since their abilities allow them to secure desirable jobs regardless of the economic conditions. These characteristics may be correlated with subsequent health, in which case our estimates will be biased.

Whether pupils in their last year of compulsory schooling stay strategically in school when the economy deteriorates is an empirical question. For each birth cohort, Figure 4 shows the proportion of pupils who left school at age 16 among the treated and non-treated group. It also pictures the one-year growth in school-leaving unemployment rates (calculated for the March-June period) faced by the youngest school cohort (treated) – compared to the previous school cohort (non-treated). When considering the 1958 and 1959 birth cohorts, Figure 4 shows that within each birth cohort, the proportion of pupils who left school at age 16 among the treated and the non-treated group is equal, indicating that school-leaving behaviours in last year of compulsory schooling were not shaped by the sharp increase in unemployment rates generated by the 1973 oil crisis. Although treated pupils from the 1958 (1959) birth cohort faced a 21% (resp. 23%) increase in unemployment rates as compared to luckier pupils born January-August, they did not enroll in an additional year of schooling.[15] When considering younger birth cohorts, however, we do find that a sharp growth in unemployment rate (e.g. the 1980s and 1990s recessions) is associated with a significant decrease in the proportion of treated pupils leaving school at age 16.[16]

Overall, we find no evidence that school-leavers born in 1958-1959 – the cohorts that we consider – did engage in endogeneous timing of school-leaving. When considering younger birth cohorts, however, we do find that a sharp growth in unemployment rate (e.g the 1980s and 1990s recessions) is associated with a decrease in the proportion of treated pupils leaving school at age 16. It can be hypothesised that pupils in their last year of compulsory schooling in 1974-1976 did not anticipate the adverse consequences of

[15]One may argue that even if the proportion of pupils who left school at compulsory age is equal in the treated and non-treated group, the composition of each group might be different. Due to the lack of information on individual characteristics at age 16, we cannot test this assumption in a proper way. However, we use information on father's occupation and show that among pupils born the same year who left school at compulsory age, the proportion of pupils whose father was in a manual occupation was the same whether they were treated or not.

[16]The two proportions are significantly different for the 1963-1964 and 1973-1974 birth cohorts. Interestingly, this result seems to suggest than rather than high unemployment rates or even increasing unemployment rates, it is a sharp increase in unemployment rates – typically occuring during recessions – that induces endogeneous timing among pupils in their last year of education.

high unemployment rates at labour market entry – contrary to school-leavers in the 1980s and 1990s recessions. It may be due to the fact that the 1973 oil crisis was highly unexpected and was the first post-war crisis to generate such a sharp increase in unemployment rates.

3.2.2 Age relative rank

School-entry rules introduce a discontinuity between August-born and September-born children; they determine that a school cohort consists of children born between the first day of September and the last day of August in the following calendar year. These rules imply that students within a same birth cohort belong to different school cohorts. As a consequence, treated individuals (born September-December) are the oldest pupils in their *school cohort*, while untreated pupils are the youngest.[17] According to Plug (2001), month of birth influences school performance not only in terms of compulsory schooling effects but by means of relative age effects. Older people in a given school cohort tend to have higher wages than younger individuals in the same school cohort. The author interprets this as an indication of the persistence of maturity effects related to age differences between students in the same class.

In our setup, the positive impact of relative maturity would be in favour of treated pupils. To the extent that maturity effects and age at school entry positively affect adult labour market and health outcomes, we should measure a lower bound, i.e. underestimate the negative impact of poor economic conditions at labour market entry on long-term labour market and health outcomes.

3.2.3 Job-experience effects

School-entry rules imply that within a birth cohort, pupils born at the end of the calendar year (the treated) start school one year later than pupils born earlier in the year (the non-treated). Starting school later entails the opportunity cost of entering the labour market later, and treated pupils lack one year of job experience as compared to non-treated pupils. However, whether an additional year of job experience has a positive or negative impact on health is not clear. On the one hand, it may lead to higher wages and turn into better health. On the other hand, it implies a longer exposure to adverse working conditions. To the extent that we are considering low-skilled individuals, this possibility is not to be discarded. Overall, the direction of the effect of an additional year of job experience on health is not clear.

However, if an additional year of job experience was beneficial to health, it would lead us to overestimate the negative impact of economic conditions at labour-market entry on subsequent health outcomes. We check in the robustness section that this is not the case. To investigate this, we run a placebo test on the 1953-1954 birth cohorts – for which the same schooling rules applied[18] (see the 1962 Education Act, Appendix D) – but which faced very similar school-leaving unemployment rates at the end of compulsory schooling.

[17] Conversely, treated individuals (born September-December) are the youngest pupils in their *birth cohort*, while untreated pupils are the oldest.

[18] Except that the minimum school-leaving age was then 15.

3.2.4 School-cohort-size effects

Since treated and non-treated pupils belong to different school cohorts, another concern has to do with school-cohort sizes. School-cohort size has been shown to have a negative impact on labor market outcomes due to an excess of supply on the labour market (e.g. Berger (1985, 1989); Korenman and Neumark (1997); Macunovich (1999); Morin (2011); Welch (1979)). We focus on three school cohorts only, which are not likely to be different in size.[19] To the extent that cohort-size effects exist, however, we should measure a lower bound : the fertility rate peaked in 1957 and declined after that, so that for a given birth year the youngest school cohort (treated) would have higher wages on average than the previous school cohort.

3.2.5 Differential incentives to take GCE O-level/CSE examinations

Depending on their date of birth, individuals within the same school cohort were allowed to leave school only after one of two specific dates (Easter or the end of the summer term) upon reaching their final year of schooling. Pupils who reached age 16 between the 1^{st} of September and the 31^{st} of January had to complete their education until the following Easter. Students who reached the age of 16 between the 1^{st} of February and the end of August were forced to leave school at the end of the summer term, typically in May/June. Pupils who left school at the end of the summer term, however, had higher incentives to take the exam at the end of Year 11 (O-Level/CSE qualifications) in which they could be awarded nationally-recognized qualifications.[20] In this context, the January/February discontinuity might introduce a bias in our analysis : treated pupils (born between September and December) are allowed to leave at Easter, and have less incentives to take the exam at the end of the year. It might impact their educational achievement as well as their adult labour market and health outcomes.

We check in the robustness section that this differential incentive in taking the exams at the end of Year 11 is not likely to bias our results.

3.2.6 Season-of-birth effects

We assume that individuals born between January and August and individuals born between September and December are identical in all observable and unobservable characteristics. Yet, a growing body of literature has shown the importance of season-of-birth effects on subsequent labour and health outcomes (Bound and Jaeger, 1996). As the seasonality of births varies from one social group to another, the season of birth is likely to be correlated with parental sociodemographic characteristics. On US data, Kestenbaum (1987) reports that children born to high-income families are more likely to be born in spring. If this is the case in the UK context – and to the extent that children born to high-income families are in better health on average – untreated pupils (born January-August) would be in better health and we

[19]Cohort sizes do not vary substantially from one year to the next. This is the reason why studies have focused on long term (typically 8-25 years) variations in cohort size (Morin (2011)).

[20]Del Bono and Galinda-Rueda (2007) implement a regression discontinuity design around the January/February discontinuity to estimate the impact of compulsory schooling on educational attainment and longer labour market outcomes. In this paper, we do not exploit the January/February discontinuity as in Del Bono and Galinda-Rueda (2007). First, because unemployment rates do not vary enough between Easter and the end of the summer term. Second, because our sample would probably be too small to detect any effect.

would overestimate the impact of economic conditions at entry on adult health outcomes.
We check in the robustness section that our results are not likely to be biased by season-of-birth effects related to parental characteristics. To investigate this, we run a placebo test on the 1953-1954 birth cohorts as in section 3.2.3.

4 Data

We use data from the General Household Survey (GHS). The GHS is an annual survey of over 13,000 households and a nationally representative survey of private households in Great-Britain.[21] It ran from 1972 to 2011 as a repeated cross-sectional survey.

Among other things, it includes information on demographics (including month-year of birth from 1983 to 2001[22], the survey waves that we use), education (including the age at which the individual left full-time education and the highest qualification obtained), labour market characteristics (including earnings and employment status) and health (including health status, health care and health behaviours). Importantly, a number of the GHS respondents left full-time education immediately after the 1973 oil crisis, a period that experienced thriving unemployment rates.

We restrict our sample to all individuals born in 1958 and 1959 and who left full-time education in their last year of compulsory education. By doing so, we consider individuals who entered the labour market between Easter 1974 and Easter 1976. We further exclude truands, i.e. pupils who left full-time education before compulsory age, and pupils born in July/August.[23] We also exclude individuals reporting that they never went to school or individuals whose highest qualification was equivalent to Year 12 or more.[24] Finally, we restrict our sample to individuals who lived in England and Wales at the time the survey was carried out because we examine school-leaving rules operating in these countries.[25] As the outcomes of interest are not collected consistently over the period (see Data Appendix C.2), we include all possible observations for each outcome to maximise sample size. Overall, our sample consists of a maximum of 1096 men over the 1986-2001 period and 1921 women over 1983-2001.

[21] See Data Appendix C.1 for more information on the sampling procedures and sample sizes in GHS.

[22] Month and year of birth in 1983-1985 are only available for women who completed the Family Information section. They are available for all respondents over 1986-2001.

[23] Under the 1962 Education Act, legislation allowed pupils born between the summer-leaving date (May/June) and the 31^{st} of August to leave school at the summer-leaving date preceding their 16^{th} birthday – i.e. at age 15. In this context, our sample should (i) include pupils born in July/August who left school at age 15 (ii) exclude pupils born in July/August who left full-time education at age 16, as the minimum school leaving age was not binding for them. Yet, patterns of school leaving seem to be quite different for pupils born in July/August. In particular, the proportion of pupils who left full-time education at binding age (i.e. at age 15) among pupils born in July/August is significantly lower than the proportion of pupils who left full-time education at binding age (i.e at age 16) among other pupils. It may be due to the fact that employers are reluctant to hire individuals under age 16 – being 15 years old at labour-market entry can be interpreted as a negative signal. We exclude all individuals born in July/August. By doing so, leaving school at age 16 in our sample becomes equivalent to leaving school at binding age.

[24] Obviously, individuals who left full-time education at age 16 could have engaged in further education and obtain higher degrees in adulthood (e.g A-Level equivalent degrees or higher degrees etc.). However, the proportion of individuals reporting Year 12-equivalent or higher degrees amounts to 20% in our sample, which is surprisingly high. One concern could be that the reported age at which individuals left full-time education suffers from measurement error. To minimise measurement error in this variable, we exclude individuals whose highest qualification was equivalent to Year 12 or more. Results are robust to re-introducing these observations (although results for men become marginally significant).

[25] Our data do not allow us to take into account migration patterns from Scotland or Northern Ireland, which is likely to generate some noise.

In addition to the variables mentioned above – month-year of birth, the age at which the individual left full-time education, the highest degree obtained and the region in which she lives – we use some information on health status, health care and health behaviours. The GHS contains several health indicators that are comparable over the 1983-2001 period. They include self-reported health status (dichotomised as poor (fair or bad health) versus good health), the presence of a longstanding illness or disability and whether the respondent restricted his activity during the two weeks preceding the interview due to illness or injury. In addition, we create several dummy variables indicating whether the respondent consulted a General Practitioner (GP) during the two weeks preceding the interview or whether she went to hospital as an outpatient (inpatient) during the 3 months (12 months) preceding the interview. The GHS also includes self-reported health behaviours such as smoking and drinking (which are measured in alternate years).

It also provides information on the economic status of the individual (employed or self-employed, unemployed according to the ILO definition, keeping house, and other statuses (e.g. other unemployed, unable permanently, student, retired or other)). For those currently employed or self-employed, it indicates the time with the present employer. We use the usual gross weekly earnings from main job (available and measured in a consistent way over the 1984-2001 period) as our main measure of earnings.[26]

Finally, we use some information regarding marriage history and fertility : marital status, age at first marriage (conditional on having been married), number of children (only for women) and age at first birth (only for women, and conditional on having at least one child).

Summary statistics for all the variables mentioned are shown separately by gender in Tables 1, 2 and 3. We also provide a breakdown of our sample by survey wave and birth cohort in Table 4.

5 Results

5.1 The impact of leaving school in a bad economy on health status, health care and health behaviours

Probit estimates of Equation (2) are presented in Table 5 for men and women respectively. Each line presents the marginal effect[27] (resp. standard error and number of observations used in the model) of having left school in a bad economy (i.e. being treated) for a different health outcome. All our models control for interview and birth year as well as for a quadratic function of age in months – see Equation (2). Our results show that over the whole period (1983-2001), men who left school in a bad economy have a 22% higher probability to smoke (at the 10% significance level) and a 27% higher probability to have ever smoked (at the 1% significance level). As regards the other health outcomes, the marginal effects for men are not statistically significant at conventional levels. However, leaving school in a bad economy seem to be particulary health-damaging for women. Marginal effects in Table 5 imply that women who

[26] Usual gross weekly earnings are measured in pounds. When running our regressions, we take the logarithm of earnings as our outcome variable.

[27] Marginal effects are computed as marginal probability effects at the sample mean value of the regressors.

left school in a bad economy have a 17% higher probability of reporting poor self-rated health (at the 1% significance level) than otherwise similar women over the whole period (1983-2001). Poor economic conditions at labour-market entry increase by 11% their probability of restricting their activity due to illness or injury during the two weeks preceding the interview (at the 5% significance level). It also raises the frequency at which they consult a GP (at the 1% significance level), since they are 17% more likely to consult the GP during the two weeks preceding the interview. In contrast, leaving school in a bad economy does not seem to affect women's propensity to suffer from a long-standing illness, nor to go to the hospital as an outpatient (inpatient) during the 3 months (12 months) preceding the interview. It does not seem to be particularly harmful to women's health behaviours such as smoking and drinking either.

Figures 5 to 7 present the impact of having left school in a bad economy on health outcomes in a life-course perspective. While estimates in Table 5 provide the average impact of poor economic conditions at labour-market entry over the whole period (1983-2001), Figures 5 to 7 allow us to investigate whether this impact is driven by middle or long-term effects. Each figure pictures the marginal effect of having left school in a bad economy (i.e. of being treated) by interview year for a different health outcome, for men and women separately.[28] For the sake of conciseness, these figures are presented only for health outcomes previously found to be significant in Table 5. For instance, Figure 5a (Figure 5b) shows – for each interview year – the marginal effect of poor economic conditions at labour-market entry on the probability of currently smoking (having ever smoked) for men. Figure 6a (resp. Figure 6b and Figure 7a) shows – for each interview year – the marginal effect of poor economic conditions at labour-market entry on the probability of reporting poorer health (resp. restricting one's activity and consulting the GP) for women.

Overall, these figures show that the average impact of leaving school in a bad economy on health does not seem to be particularly driven by middle or long-term effects. For each figure, the majority of marginal effects lie above the zero line for all interview years – and not particularly at the beginning or at the end of the interview period. Thus, men and women seem to negatively affected by poor economic conditions at labour-market entry over the whole period under study.

5.1.1 The impact of leaving school on labour-market outcomes, marriage and fertility

Labour market, marriage and fertility can be viewed as mechanisms by which leaving school in a bad economy affects long-term health. To investigate this further, we regress labour market, marriage and fertility proxies on the treatment variable T_i as well as on birth-year, the quadratic function of age and interview dummies (as in section 3.1, see Equation (2)). Models are estimated by OLS or probit – depending on the nature of the dependent variable (continuous or dichotomous)).

Table 6 presents the effect of having left school in a bad economy on labour-market outcomes for men and women respectively. We find no effect on unemployment, inactivity patterns and earnings, neither

[28] We compute marginal effects by interview year. This is why we lose statistical power compared to the main analysis presented in Table 5.

for men, nor for women. Men and women who left school in a bad economy do not seem to have been in their current job for a shorter period of time.

Table 7 presents the effect of having left school in a bad economy on marriage and fertility outcomes for men and women respectively. Our results show that leaving school in a bad economy does not seem to have a significant impact on marital status and age at first marriage, neither for men, nor for women. Women's number of children and age at first birth do not seem to be significantly affected by poor economic conditions at labour-market entry.

Overall, we do not find that leaving school in a bad economy has a lasting impact on labour-market, marriage and fertility outcomes 7 to 26 years after school-leaving. This is not really suprising, though, as we do consider a very specific population, i.e. low-educated individuals. Economic theory indeed predicts less persistence of poor economic conditions at school-leaving for low-skilled workers subsequent labour-market outcomes. Stevens (2007), Gaini et al. (2012) and Genda et al. (2010) provide evidence that the negative effect of graduating in a bad economy on labour-market outcomes vanishes after a few years when considering low-educated individuals in Germany, France and the USA. As regards marriage and fertility outcomes, Kondo (2012) shows that labor market fluctuations in the USA induce only intertemporal adjustments for marriage timing without affecting reservation match quality or total fertility, which is consistent with our results.

5.2 Robustness Checks

5.2.1 Placebo test

Seasonality of birth linked to parental socio-demographic characteristics may be beneficial to students born January-August (non-treated pupils). It might lead to an overestimation of the negative impact of economic conditions at labour-market entry on adult outcomes (see section 3.2). To investigate whether being born between January and August influences health, labour, marriage and fertility outcomes not in terms of economic conditions at labour market entry but by means of experience and parental socio-demographic characteristics, we run a placebo test on the 1953-1954 birth cohorts. The 1953 and 1954 birth cohorts faced very similar school-leaving unemployment rates at the end of compulsory schooling.[29] School-leaving unemployment rates (averaged over March-June[30]) varied from 2.475 to 2.675 over a three-year period (1968-1970).

We re-run our probit models on the 1953-1954 cohorts. Results are presented in Table 8. As expected, we find no significant effect of being born between September and December (as compared to being born between January and August) on any health outcome (all coefficients are unsignificant at conventional levels), indicating that our results are not likely to be biased by season-of-birth effects.

[29] The same school exit rules applied for the 1953 and 1954 birth cohorts (see the 1962 Education Act, Appendix D) except that the school compulsory age was 15.

[30] Unemployment rates are computed using administrative data, namely the monthly "registrant count" and borrowed from Denman and McDonald (1996).

5.2.2 Differential incentives to take GCE O-level/CSE examinations

One may worry that treated pupils have fewer incentives to take examinations at the end of Year 11. It might impact their educational achievement and later health outcomes. To control for this potential bias, we re-run our regressions controlling by an dummy variable indicating whether the individual holds a Year-11-equivalent qualification (O-level, CSE etc.). Our results are virtually unchanged.

5.2.3 Alternative empirical approach

Up to now, our treatment variable has been a dummy variable indicating whether an individual was born at the end of the calendar year or earlier in the year. We have considered the 1958 and 1959 cohorts. A possible drawback of this approach is that is linearises the impact of the treatment across the two birth cohorts – which may be problematic to the extent that treated pupils do not face the same increase in school-leaving unemployment rates as compared to non-treated pupils across the two birth cohorts (a 0.7 and a 1.1 point increase respectively).
To deal with this potential problem, we estimate the following equation :

$$H_i = \lambda + \pi UR_i + BirthYear_i + f(BirthMonth_i) + InterviewYear_i + \eta_i \tag{3}$$

where UR_i denotes the school-leaving unemployment rate faced by individual i and the other variables are presented in section 3.
Probit estimates of Equation (3) are presented in Table 9 and are very similar to the ones presented in the main analysis – see Table 5. In particular, estimates in Table 9 imply that a one-point increase in school-leaving unemployment rates leads to a 16% higher probability to report poor health and a 23% higher probability to have ever smoked among men (at the 10% and 5% significance level respectively). Similarly, it leads to a 16% higher probability to report poor health (at the 5% significance level), a 9% higher probability to restrict one's activity due to illness or injury (at the 5% significance level) and a 11% higher probability to consult the GP (at the 10% level) among women.

6 Conclusion

In this paper, we provide evidence of the impact of leaving school in a bad economy on long-term health status, health care consumption and health behaviours. We consider pupils in England and Wales who left school in their last year of compulsory schooling immediately after the 1973 oil crisis and whose school-leaving behaviour in worse economic conditions was exogeneously induced by compulsory schooling laws. We provide evidence that these pupils did not engage in endogeneous timing. We use a repeated cross section of individuals over 1983-2001 from the General Household Survey (GHS) and take a life-course perspective. We show that men who left school in a bad economy have a higher probability of smoking and of having smoked. Women having left school in a bad economy also have a higher probability to report poorer health, to restrict their activity due to illness or injury and to consult the GP over the whole period under study. We find no significant results on labour-market, marriage and fertility outcomes.

Overall, our results provide evidence that leaving school in a bad economy has a large, lasting and negative impact on health and health behaviours among low-educated individuals. Cutler et al. (2014) documents that the impact of poor economic conditions of subsequent health is likely to be larger for lower-educated individuals, i.e. that education plays a protective role. If this is the case, economic conditions at labour-market entry would be an important determinant of education-based inequalities in health.

Our results are consistent with the idea that poor economic conditions have an impact – at least in the short-run – on labour-market, marriage and fertility outcomes, which cumulate over the lifecourse and later translate into worse health. We cannot totally discard the hypothesis, however, that leaving school in a bad economy generates an initial and negative health shock (e.g. through addictive behaviours) which is not compensated over the life-course. A promising avenue for future research would consist in investigating which hypothesis is most likely to hold in the data.

There are some limitations to our study. The most notable is the small sample size, which generates quite imprecise results. In particular, the subsample of men is rather small, so that our results on this population cannot be interpreted as ruling out any damaging impact of poor economic conditions at labour-market entry on health outcomes. We plan to use data from the Health Survey for England (HSE) to increase sample size and consider a broader set of health outcomes.

A potential extrapolation of our findings is that the Great recession will have lasting and negative health effects among lower-educated individuals. However, the external validity of our findings depends on the similarity between the 1958 and 1959 GLS cohorts and current cohorts of school-leavers. It also depends on the similarity between the 1973 crisis and the current Great recession. To the extent that the 1973 oil crisis and the current Great recession do not seem to have had the same effects on unemployment rates, wages and working conditions, the external validity of our study is still an open issue.

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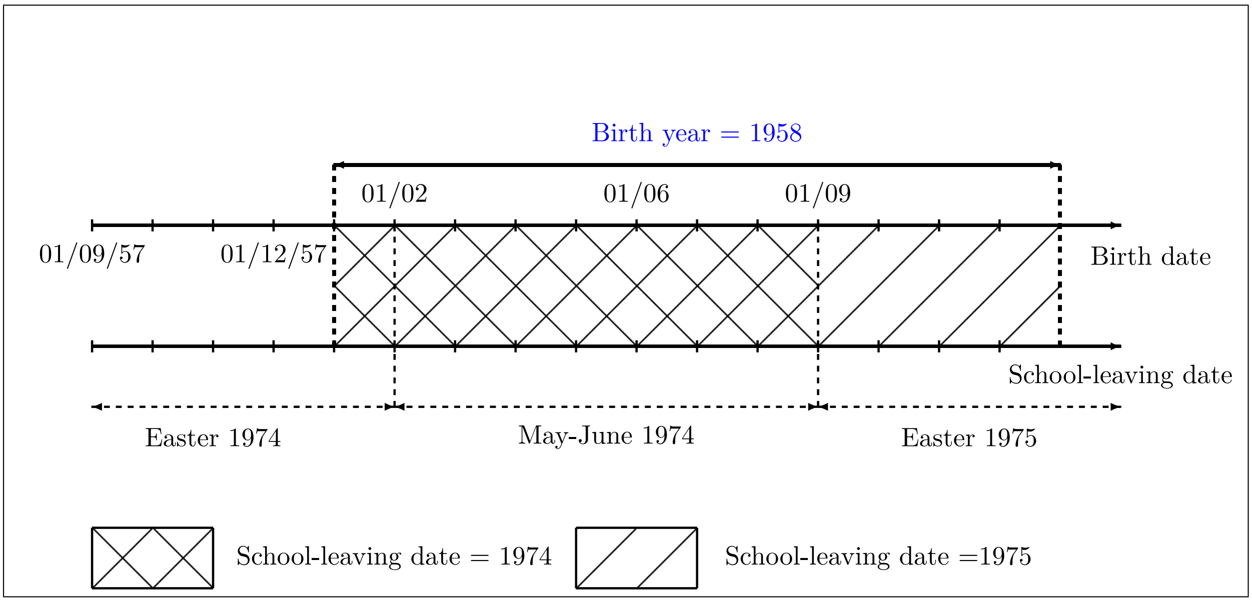
A Figures

Figure 1: Compulsory schooling rules by month-year of birth

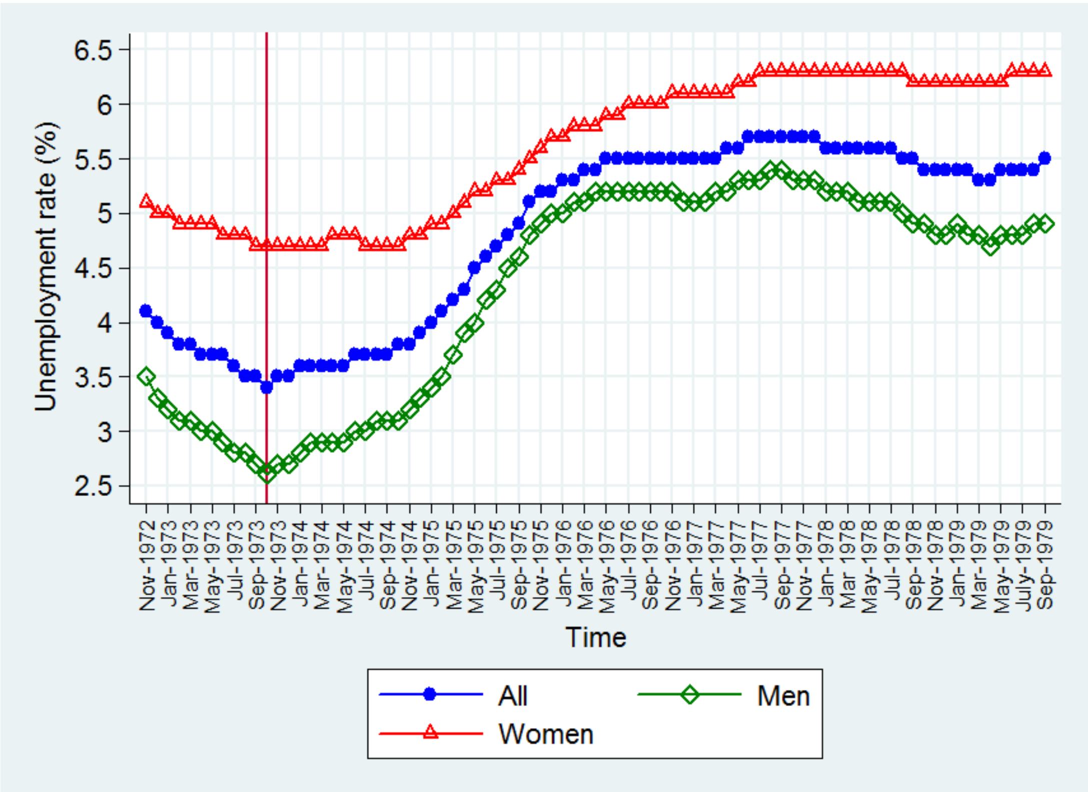
Birth year (1)	Month of birth (2)	School starting date (3)	Allowed to leave school (4)
1958	January	Sept. 1963	Easter 1974
1958	February	Sept. 1963	May/June 1974
1958	March	Sept. 1963	May/June 1974
1958	April	Sept. 1963	May/June 1974
1958	May	Sept. 1963	May/June 1974
1958	June	Sept. 1963	May/June 1974
1958	July	Sept. 1963	May/June 1974
1958	August	Sept. 1963	May/June 1974
1958	September	Sept. 1964	Easter 1975
1958	October	Sept. 1964	Easter 1975
1958	November	Sept. 1964	Easter 1975
1958	December	Sept. 1964	Easter 1975
1959	January	Sept. 1964	Easter 1975
1959	February to August	Sept. 1964	May/June 1975
1959	September to December	Sept. 1965	Easter 1976

Figure 2: Focus on the 1958 birth cohort



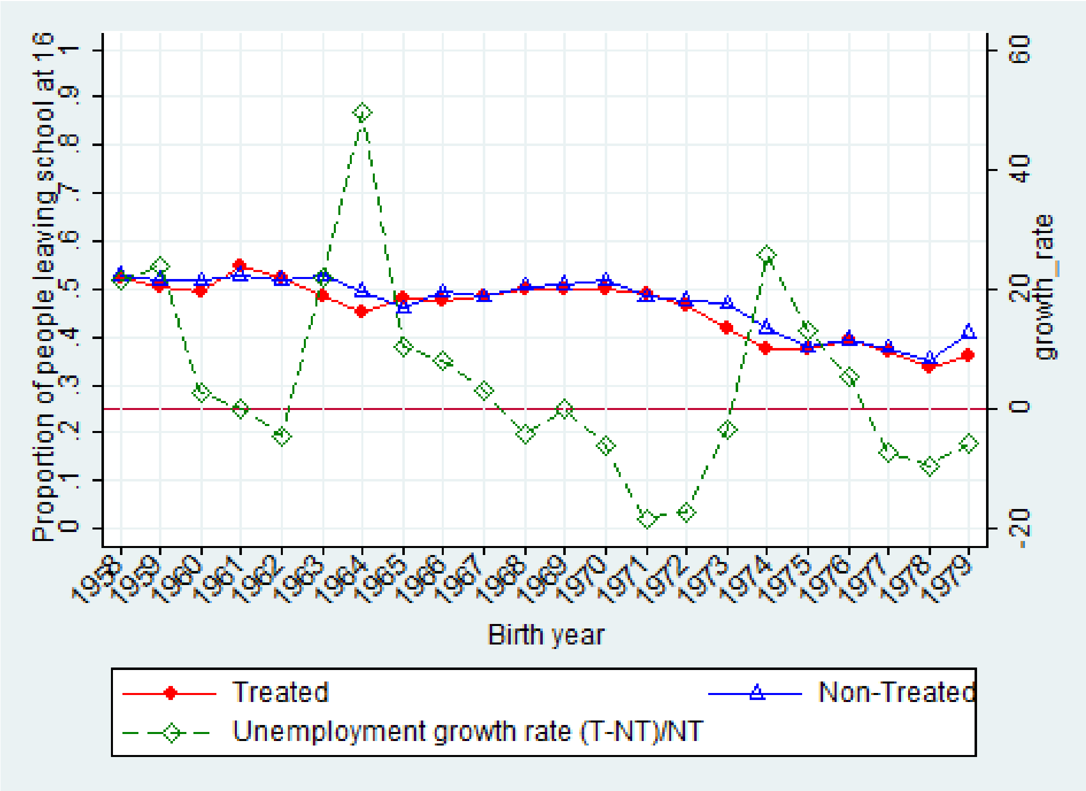
Reading: A pupil born between the 1^{st} of September 1958 and the 31^{st} of December 1958 is allowed to leave school at Easter 1975.

Figure 3: Unemployment rates for all individuals aged 16 and over the 1973-1979 period, seasonally adjusted.



Source: Labour Force Survey (LFS), provided by the Office for National Statistics (ONS).

Figure 4: Proportion of pupils leaving school at age 16; Growth in school-leaving unemployment rate

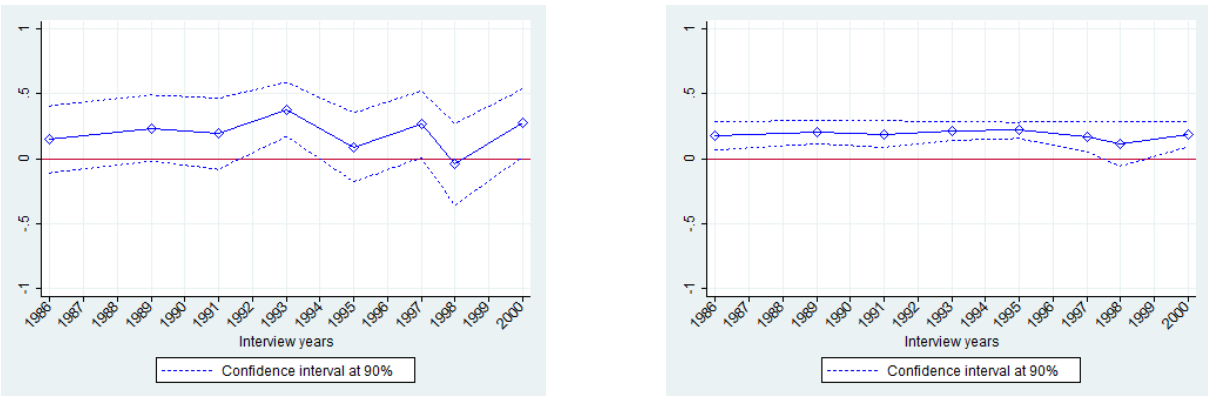


Reading: Figure 4 presents the proportion of pupils leaving school at age 16 and the growth in school-leaving unemployment rate (calculated for the March-June period) for pupils born the same year but leaving school one year apart (i.e. belonging to different school cohorts).

Figure 5: Health behaviours for men

(a) Currently smokes

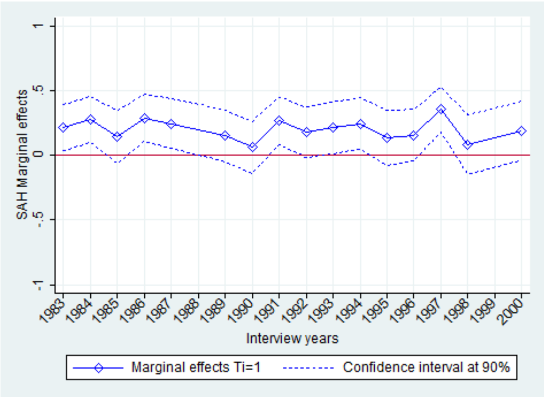
(b) Ever smoked



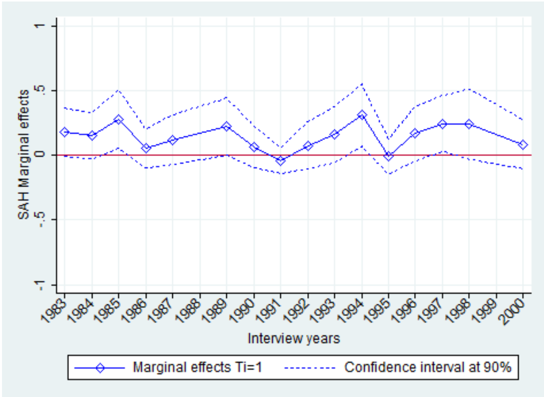
Notes for Figures (a) and (b): we compute marginal effects of having left school in a bad economy (i.e. of being treated) for each interview year (1983-2001).

Figure 6: Health status for women

(a) Poor health



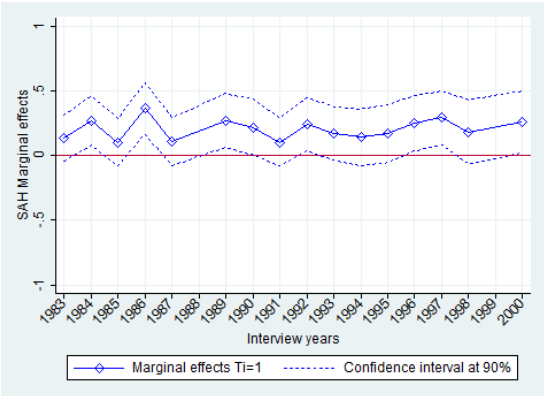
(b) Restrict activity



Notes for Figures (a) and (b): we compute marginal effects of having left school in a bad economy (i.e. of being treated) for each interview year (1983-2001).

Figure 7: Health care for women

(a) GP consultations



Notes: we compute marginal effects of having left school in a bad economy (i.e. of being treated) for each interview year (1983-2001).

B Tables

Table 1: Summary statistics of demographic and health variables – pooled sample

	Men			Women		
	Mean	s.e	N	Mean	s.e	N
	(1)	(2)	(3)	(4)	(5)	(6)
Demographics						
Age	33.20	(4.20)	*1096*	31.19[a]	(5.00)	*1921*
Health status						
Poor self-rated health	0.30	(0.46)	*1044*	0.34	(0.47)	*1909*
Longstanding illness/disability (yes/no)	0.26	(0.44)	*1096*	0.23	(0.42)	*1917*
Restricts activity due to longstanding illness/injury	0.08	(0.27)	*1095*	0.13	(0.33)	*1920*
Health care						
GP consultations last 2 weeks (yes/no)	0.12	(0.32)	*1094*	0.21	(0.41)	*1920*
Outpatient spell last 3 months (yes/no)	0.13	(0.34)	*1095*	0.11	(0.32)	*1918*
Inpatient spell last year (yes/no)	0.05	(0.22)	*1095*	0.16	(0.37)	*1920*
Health behaviours						
Smoking status			*619*			*1029*
Currently smokes	0.43	(0.50)		0.42	(0.49)	
Has smoked but does not anymore	0.33	(0.47)		0.27	(0.44)	
Ever smoked	0.76	(0.43)		0.69	(0.46)	
Heavy smoker – more than 19 cig./day – (yes/no)[b]	0.59	(0.49)	*262*	0.54	(0.50)	*421*
Self-reported drinking behaviour			*597*			*945*
High to moderate alcohol consumption (yes/no)	0.52	(0.50)		(0.34)	0.47	

Notes : [a] : Women are on average younger than men because they are observed over the whole period (1983-1985) while men are only observed from 1986 on (see Table 4). [b] : conditional on currently smoking.

Table 2: Summary statistics of labour-market characteristics – pooled sample

	Men			Women		
	Mean	s.e	*N*	Mean	s.e	*N*
	(1)	(2)	(3)	(4)	(5)	(6)
Economic status			*1096*			*1920*
Employed or self-employed	0.84	(0.37)		0.58	(0.49)	
Unemployed	0.10	(0.31)		0.06	(0.23)	
Keeping house	0.01	(0.09)		0.34	(0.47)	
Other	0.05	(0.21)		0.02	(0.15)	
For those currently employed or self-employed						
Usual gross weekly earnings from main job (in pounds)	283.72	(880.68)	*819*	109.92	(99.03)	*970*
Time with present employer			*724*			*861*
Less than 1 month	0.02	(0.13)		0.03	(0.17)	
Between 1 and 3 month	0.04	(0.20)		0.06	(0.24)	
Between 4 and 6 month	0.04	(0.20)		0.06	(0.24)	
Between 7 and 12 months	0.08	(0.27)		0.11	(0.31)	
Between 1 and 5 years	0.20	(0.40)		0.34	(0.47)	
Five years or more	0.61	(0.49)		0.38	(0.49)	

Table 3: Summary statistics of marriage and fertility variables – pooled sample

	Men			Women		
	Mean	s.e	*N*	Mean	s.e	*N*
	(1)	(2)	(3)	(4)	(5)	(6)
Marriage						
Marital status			*1096*			*1921*
Married	0.70	(0.46)		0.71	(0.45)	
Single	0.25	(0.43)		0.16	(0.37)	
Widowed	0.00	(0.04)		0.00	(0.05)	
Divorced	0.04	(0.19)		0.08	(0.28)	
Separated	0.02	(0.12)		0.04	(0.20)	
Age at first marriage$^{(a)}$	23.20	(3.75)	*819*	20.86	(3.36)	*1617*
Fertility						
Number of children	-	-		1.68	(1.24)	*1903*
Age at first birth$^{(b)}$	-	-		22.79	(4.11)	*1124*

Notes : $^{(a)}$: conditional on having been married. $^{(b)}$: conditional on having at least one child.

Table 4: Number of observations by survey wave and birth cohort

	Men	Women	All
	(1)	(2)	(3)
Survey wave			
1983	-	159	159
1984	-	153	153
1985	-	127	127
1986	100	130	230
1987	92	140	232
1988-1989	83	140	223
1989-1990	82	102	184
1990-1991	74	126	200
1991-1992	107	124	231
1992-1993	76	97	173
1993-1994	85	109	194
1994-1995	93	98	191
1995-1996	71	121	192
1996-1997	92	118	210
1998-1999	62	85	147
2000-2001	79	92	171
Birth cohort			
1958	544	972	1516
1959	552	949	1501
Total number of observations	1096	1921	3017

Notes: (1) The GHS was conducted annually, except for breaks in 1997-1998 when the survey was reviewed, and 1999-2000 when the survey was redeveloped. (2) Month and year of birth in 1983-1985 are only available for women who completed the Family Information section. They are available for all respondents from 1986-2001 on.

Table 5: The impact of leaving school in a bad economy on health outcomes : Marginal effects

	Men			Women		
	m.e.	s.e.	*N*	m.e.	s.e.	*N*
Probit regressions						
Health status						
Poor health	0.103	(0.089)	*1043*	0.174***	(0.067)	*1907*
Longstanding illness	0.059	(0.081)	*1095*	0.062	(0.060)	*1915*
Restricts act	-0.025	(0.046)	*1094*	0.110**	(0.051)	*1918*
Health care						
GP consultations	-0.059	(0.053)	*1093*	0.166***	(0.060)	*1918*
Hospital outpatient consultation	-0.008	(0.060)	*1094*	-0.002	(0.043)	*1916*
Hospital inpatient consultation	0.042	(0.044)	*1033*	0.053	(0.049)	*1918*
Health behaviour						
Currently smokes	0.215*	(0.119)	*618*	0.025	(0.094)	*1027*
Heavy smoker$^{(a)}$	-0.009	(0.182)	*262*	-0.055	(0.146)	*420*
Ever smoked	0.266***	(0.087)	*618*	0.091	(0.083)	*1027*
Moderate to heavy drinking	0.023	(0.125)	*596*	-0.007	(0.092)	*943*

Notes : *** p-value<0.01, ** p-value<0.05, * p-value<0.1. Marginal effects (m.e.) are presented (computed as marginal probability effects at the sample mean value of the regressors). Robust standard errors in parentheses (s.e.). Our models include dummy variables for interview and birth year as well as a quadratic function of age in months – see Equation (2). $^{(a)}$: conditional on currently smoking.

Table 6: The impact of leaving school in a bad economy on labour-market outcomes

	Men			Women		
	m.e.	s.e.	*N*	m.e.	s.e.	*N*
Probit regressions						
Economic status						
Keeping house	0.012	(0.031)	*495*	0.073	(0.067)	*1918*
Unemployed	0.058	(0.061)	*1095*	-0.020	(0.028)	*1918*
For those currently employed						
Less than 1 month	0.036	(0.048)	*512*	0.049	(0.044)	*805*
Less than 3 months	-0.022	(0.051)	*613*	0.062	(0.067)	*861*
Less than 6 months	-0.039	(0.063)	*723*	0.028	(0.080)	*861*
Less than 1 year	0.012	(0.088)	*723*	-0.046	(0.093)	*861*
Less than 5 years	-0.002	(0.114)	*723*	-0.063	(0.104)	*861*
More than 5 years	0.002	(0.114)	*723*	0.063	(0.104)	*861*
Linear regressions						
Earnings (log)	-0.033	(0.109)	*799*	-0.169	(0.169)	*957*

Notes : *** p-value<0.01, ** p-value<0.05, * p-value<0.1. Marginal effects (m.e.) are presented (computed as marginal probability effects at the sample mean value of the regressors). Robust standard errors in parentheses (s.e.). Our models include dummy variables for interview and birth year as well as a quadratic function of age in months – see Equation (2).

Table 7: The impact of leaving school in a bad economy on fertility and marriage outcomes

	Men			Women		
	m.e.	s.e.	*N*	m.e.	s.e.	*N*
Probit regressions						
Marriage						
Married	0.073	(0.081)	*1095*	0.008	(0.063)	*1919*
Divorced	0.044	(0.042)	*1021*	-0.045	(0.033)	*1919*
Single	-0.110	(0.073)	*1095*	0.068	(0.051)	*1919*
Linear regressions						
Age at first marriage$^{(a)}$	-0.562	(0.839)	*818*	0.376	(0.471)	*1615*
Fertility						
Age at first birth$^{(b)}$	-	-	-	0.581	(0.729)	*1122*
Nb of children	-	-	-	0.233	(0.159)	*1901*

Notes : *** p-value<0.01, ** p-value<0.05, * p-value<0.1. Marginal effects (m.e.) are presented (computed as marginal probability effects at the sample mean value of the regressors). Robust standard errors in parentheses (s.e.). Our models include dummy variables for interview and birth year as well as a quadratic function of age in months – see Equation (2).

Table 8: Placebo test : the impact of leaving school in a bad economy on health outcomes for the 1953-54 cohorts

	Men			Women		
	m.e.	s.e.	*N*	m.e.	s.e.	*N*
Probit regressions						
Health status						
Poor health	-0.045	(0.111)	*631*	-0.128	(0.084)	*1204*
Longstanding ill	0.062	(0.111)	*664*	-0.015	(0.079)	*1210*
Restricts act	0.002	(0.071)	*663*	-0.022	(0.055)	*1213*
Health care						
GP consultations	-0.036	(0.062)	*664*	0.036	(0.070)	*1211*
Hospital outpatient consultation	-0.098	(0.062)	*664*	-0.075	(0.052)	*1212*
Hospital inpatient consultation	-0.018	(0.040)	*619*	-0.067	(0.046)	*1212*
Health behaviour						
Currently smokes	-0.094	(0.148)	*390*	0.178	(0.115)	*653*
Heavy smoker	0.043	(0.213)	*173*	-0.264	(0.161)	*297*
Ever smoked	0.028	(0.110)	*362*	0.098	(0.094)	*653*
Moderate to heavy drinking	-0.243	(0.153)	*372*	-0.036	(0.114)	*617*

Notes : *** p-value<0.01, ** p-value<0.05, * p-value<0.1.. Marginal effects (m.e.) are presented (computed as marginal probability effects at the sample mean value of the regressors). Robust standard errors in parentheses (s.e.). Our models include dummy variables for interview and birth year as well as a quadratic function of age in months – see Equation (2). $^{(a)}$: conditional on currently smoking.

Table 9: An alternative empirical approach. The impact of school-leaving unemployment rates on health outcomes : Marginal effects

	Men			Women		
	m.e.	s.e.	*N*	m.e.	s.e.	*N*
Probit regressions						
Health status						
Poor health	0.160*	(0.092)	*1043*	0.160**	(0.068)	*1907*
Longstanding illness	0.067	(0.084)	*1095*	0.080	(0.060)	*1915*
Restricts act	0.002	(0.052)	*1094*	0.092**	(0.046)	*1918*
Health care						
GP consultations	-0.026	(0.061)	*1093*	0.105*	(0.058)	*1918*
Hospital outpatient consultation	-0.040	(0.067)	*1094*	0.002	(0.045)	*1916*
Hospital inpatient consultation	0.048	(0.038)	*1033*	0.042	(0.048)	*1918*
Health behaviour						
Currently smokes	0.188	(0.131)	*618*	0.023	(0.099)	*1027*
Heavy smoker$^{(a)}$	0.056	(0.193)	*262*	-0.057	(0.155)	*420*
Ever smoked	0.234**	(0.113)	*618*	0.143	(0.091)	*1027*
Moderate to heavy drinking	0.090	(0.134)	*596*	0.007	(0.098)	*943*

Notes : *** p-value<0.01, ** p-value<0.05, * p-value<0.1. Marginal effects (m.e.) are presented (computed as marginal probability effects at the sample mean value of the regressors). Robust standard errors in parentheses (s.e.). Our models include dummy variables for interview and birth year as well as a quadratic function of age in months – see Equation (3). $^{(a)}$: conditional on currently smoking.

C Data appendix : Sample and variable construction

C.1 Changes to sampling procedures and sample sizes over time

According to the GHS Time Series Dataset User Guide (2007), "the sampling procedure used on the GHS has changed over time, resulting in different sample sizes between survey years. However, the changes to the GHS sample procedures and sample size were relatively small. As a result it was decided by ONS that these changes were likely to have little impact on the reliability of the estimates. Particularly as a representative sample of the population has been achieved for each survey year."

Non-response weights are only available in the GHS after 2000. As a consequence, all our estimates are unweighted.

C.2 Inconsistencies in variables over time

According to the GHS Time Series Dataset User Guide (2007), "in general variables in the GHS have remained fairly consistent over time. However as the GHS has been revised and research interests have changed, some variables have been modified over the past 30 years to reflect this. For example the marital status variable was revised in the 1986 survey to include a category for cohabitation. Similarly, some questions were only included on a few survey years, or in more recent rounds of the survey series, which limits analysis over time."

Those variables that were only available for a few years, or had substantially changed over time were not used in the analysis. The construction of each variable used in our analysis is depicted in greater detail below.

D School leaving age legislation in England and Wales

Relevant extracts of the 1962 Education Act are borrowed from Del Bono and Galinda-Rueda (2007).

Education Act 1962: relevant extracts from Section 9

Applies to 15 year old individuals in 1963, i.e. people born in 1947 or afterwards.

(2) If he attains that age on any date from the beginning of September to the end of January, he shall be deemed not to have attained that age until the end of the appropriate spring term at this school.

(3) If he attains that age on any date on or after the beginning of February but before the end of the appropriate summer term at his school, he shall be deemed not to have attained that age until the end of that summer term.

(4) If he attains that age on any date between the end of the appropriate summer term at this school and the beginning of September next following the end of that summer term (whether another term has then begun or not) he shall be deemed to have attained that age at the end of that summer term. [...]

(7) In this section, "the appropriate spring term", in relation to a person, means the last term at this school which ends before the month of May next following the date on which he attains the age in question, and "the appropriate summer term" [...] means the last term at this school which ends before the month of September next following that date [...].

Education School leaving Act 1976: relevant extracts from Section 1

Subsections (3) and (4) in Section 9 of the Education Act of 1962 were substituted by the following subsections of Section 1 of the Education School leaving Act 1976 in order to give a more precise meaning to the notion of school leaving dates, particularly for those born after the end of January.

(3) If he attains that age after the end of January but before the next May school leaving date, he shall be deemed not to have attained that age until that date.

(4) If he attains that age after the May school leaving date and before the beginning of September next following that date, he shall be deemed to have attained that age on that date. A new subsection was added at the end of Section 9 of the Education Act of 1962, according to which:

(8) In this section the May school leaving date means the Friday before the last Monday in May.