Is it Enough to Increase Compulsory Education to Raise Earnings? Evidence from French and British Compulsory Schooling Laws*

Julien Grenet
Paris School of Economics

July 2011

Abstract

In this paper, we compare two reforms that raised the minimum school leaving age to 16 in France (1967) and in England & Wales (1972). Using an RDD estimation strategy, we find that while the British reform led to a 6 to 7% increase in hourly wages per additional year of compulsory schooling, the impact of the French law change was close to zero. Our results suggest that the major difference between the two reforms was that the fraction of individuals holding no qualifications dropped sharply after the introduction of the new minimum school leaving age in England & Wales whereas it remained unchanged in France.

JEL Classification I20, I21, I28, J24, J31

Keywords Returns to education, compulsory schooling, sheepskin effects

*Mailing address: Paris School of Economics, 48 boulevard Jourdan, 75014 Paris, France. E-mail: julien.grenet(at)ens.fr. I wish to thank Robert Gary-Bobo, Marc Gurgand, Francis Kramarz, Stephen Machin, Alan Manning, Éric Maurin, Philip Oreopoulos, Thomas Piketty, Steve Pischke as well as anonymous referees for their helpful comments. An earlier version of this paper has circulated under the title "Is it Enough to Increase Compulsory Education to Raise Earnings? Evidence from the French Berthoin Reform".

1 Introduction

Compulsory schooling laws have been extensively used as an exogenous source of variation in educational attainment to estimate the causal impact of education on earnings. Most papers find that instrumental variable estimates of the economic return to compulsory schooling are higher than the naive OLS returns and usually in the range of 10 to 15% per additional year of education. However, a number of recent studies have found the rather puzzling result that these returns are small and even close to zero in certain countries.

Most early compulsory schooling papers follow the instrumental variable strategy first proposed by Harmon and Walker (1995), which consists in estimating a typical Mincerian earnings equation in which schooling is instrumented by the minimum school leaving age faced by individuals. In this setting, the impact of increased educational attainment is identified by comparing average schooling and earnings across pre- and post-reform cohorts. Using this methodology and exploiting the reforms that raised the minimum school leaving age in England & Wales in 1947 and 1972, Harmon and Walker find that the Two Stage Least Squares (2SLS) estimates of the returns to schooling are higher than the OLS naive estimates (15% vs. 6%). The application of this approach to other countries leads to similar conclusions[1]. However, its main limitation is that, by not involving a proper control group, it fails to neutralize cohorts fixed effects[2]. To overcome this problem, subsequent papers have relied on two alternative empirical strategies to estimate the returns to compulsory schooling, either by exploiting variations in the timing of compulsory schooling laws across states or regions within a difference-in-difference (DID) framework, or by controlling smoothly for the evolution of schooling and earnings across cohorts in a Regression Discontinuity Design (RDD) framework. Papers that have used this approach in the US, Canada or the UK typically find returns to schooling in

[1]See in particular Callan and Harmon (1999) for Ireland, Brunello and Miniaci (1999) and Brandolini and Cipollone (2002) for Italy, Levine and Plug (1999) for the Netherlands, Vieira (1999) for Portugal and Pons and Gonzalo (2002) for Spain.

[2]This issue is discussed in Card (1999) and Oreopoulos (2006a).

the range of 7 to 14% (Acemoglu and Angrist, 2000; Oreopoulos, 2006b,a, 2007)[3].

The consensus view that compulsory schooling laws have large positive effects on wages has nevertheless been challenged by a number of recent studies carried out on European data. Meghir and Palme (2005) find that although the Swedish reform which raised compulsory schooling to nine years in the late 1940s led to a 3.4% increase in the average earnings of individuals with unskilled fathers, it had only a small and insignificant overall impact of 1.4% on average earnings[4]. Pischke and von Wachter (2008) find zero returns from the raising of the minimum school leaving age in West Germany during the period from 1947 to 1969. Oosterbeek and Webbink (2007) evaluate the impact of the extension from 3 to 4 years in the length of vocational training programs in the Netherlands and find no beneficial effect from the change. Devereux and Hart (2010) carry out an RDD analysis of the 1947 British compulsory schooling law previously studied by Harmon and Walker (1995) and Oreopoulos (2006a) and find a zero return for women and a modest return for men in the 4 to 7% range.

At the minimum, these results suggest that compulsory schooling laws do not systematically improve the labour market prospects of early school dropouts. With respect to this vast literature, the aim of the present study is to throw light into the black box of compulsory schooling laws and identify some of the institutional features that could drive the benefits of compulsory education. More precisely, I argue that in the European context, where examination-based diplomas that certify the successful completion of compulsory education are widespread, academic credentials are an important element to take into account when assessing the impact of increases in the minimum school leaving age – a dimension which has not received much emphasis in previous studies.

[3]Furthermore, a growing body of literature based on similar estimation strategies suggests that the returns to compulsory education are not restricted to the labour market and influence outcomes as different as subjective measures of well-being (Oreopoulos, 2007), health (Lleras-Muney, 2005), criminal behavior (Lochner and Moretti, 2004), political involvement (Milligan et al., 2004), teenage childbearing (Black et al., 2008) or the intergenerational transmission of inequality (Oreopoulos et al., 2006).

[4]If all the earnings changes were due to changes in the quantity of education, this would correspond to an average return of 4.8% and a return of 8.4% for individuals with unskilled fathers.

I explore this issue by carrying out a comparative analysis of two important compulsory schooling reforms that were adopted in the same period: the 1967 Berthoin edict which raised the minimum school leaving age from 14 to 16 in France and the 1972 Raising of the School Leaving Age order (hereafter referred to as ROSLA) which extended compulsory education from the age of 15 to the age of 16 in England & Wales. Using an RDD approach, which to the best of my knowledge has not been applied to either of these two reforms[5], I show that despite their similar impact on average educational attainment (about +0.3 year), both reforms had very different effects on earnings. While the 1967 Berthoin reform had no sizable impact on the wages of French women and men, the monetary returns to the 1972 ROSLA were positive and significant for both in England & Wales.

I argue that factors such as the heterogeneity of the returns to schooling, the role of wage-setting institutions or a hypothetical defective implementation of the Berthoin reform cannot credibly account for this discrepancy. Instead, I provide suggestive evidence in favour of an explanation based on the role of academic credentials. Indeed, I find that a major difference between the two reforms is that the fraction of individuals holding no qualifications dropped sharply after the introduction of the new minimum school leaving age in England & Wales, but remained unchanged in France. I link this observation to a number of institutional features of the French and British educational systems and discuss the reasons why in both countries, the actual quantity of education may be less important than credentials in determining the returns to compulsory schooling. I also present some evidence supporting the idea that the benefits derived from better credentials by early school dropouts in England & Wales were not entirely driven by signalling mechanisms, but also reflected a substantial improvement in their level of skills. I interpret my findings as suggesting that in countries where sub-degree qualifications play a key role in certifying

[5]The 1972 ROSLA in England & Wales has been previously examined by Harmon and Walker (1995) and Oreopoulos (2007), but together with other law changes and within a different estimation framework. The French 1967 Berthoin reform has not been previously studied in the literature to evaluate the monetary returns to compulsory education.

the knowledge and skills a person has achieved through study, compulsory schooling laws may be effective only to the extent that they induce a significant fraction of pupils to become enrolled in grades that lead to academic or vocational certification examinations.

The remainder of this paper is as follows. Section 2 provides some institutional background on the French and British compulsory schooling reforms of 1967 and 1972; section 3 presents the estimation strategy; section 4 describes the data and provides summary statistics; the main results are discussed in section 5; section 6 provides suggestive evidence consistent with an explanation based on certification effects; section 7 discusses alternative interpretations and section 8 concludes.

2 Institutional background

Although implemented in different institutional contexts, the 1967 Berthoin reform in France and the 1972 ROSLA in England & Wales share a number of characteristics which support and justify the comparison of their respective effects.

First, both reforms extended the minimum school leaving age to 16. In France, the Berthoin reform was adopted in 1959 and added two years of compulsory education by raising the minimum school leaving age from 14 to 16 in 1967 for pupils born from January 1, 1953 onwards[6]. In England & Wales, where the school cutoff date is September 1, the extension of the minimum school leaving age from 15 to 16 was decided upon in 1964 and enacted in March 1972 (Statutory Instrument No. 444). From September 1, 1972, onwards, pupils could not leave school until they turned 16. Since in England & Wales, the fraction of pupils who do not start school in the expected year or do not progress through the system in the usual manner is negligible, the actual date of birth cutoff for the 1972 ROSLA was September 1, 1957[7].

Second, the raising of the minimum school leaving age was envisaged in both countries

[6] In France, the school cutoff date is January 1.

[7] In the online appendix, I explain why deviations from perfect correspondence between date of birth and assignment to the pre- or post-compulsory schooling regimes are unlikely to be important in France and England & Wales.

as an important building block for the unification and expansion of secondary education. In France, the unification of secondary schooling started in 1959 and consisted in gradually replacing extended primary schooling by a comprehensive curriculum of junior secondary schooling. The organization of the French school system at the time of the 1967 reform is described in figure 1. After 5 years of primary schooling, about half the pupils would attend 3 years of extended primary schooling before continuing into vocational education. The other half would start a 4-year junior secondary schooling curriculum in one of the newly created *Collèges*. The most academically oriented would then spend 3 years in junior secondary school (*Lycées*) while the less able would enter vocational education.

In much a similar way as the French Berthoin edict, the 1972 ROSLA came with a series of reforms aimed at widening the access to secondary education in England & Wales. At the time of the reform, most Local Education Authorities had switched to the new comprehensive system which is described in figure 2. After 6 years of primary education, pupils would spend between 3 and 5 years in secondary education. The most able pupils would then continue their upper secondary studies in a secondary school with an attached Sixth form or transfer to a local Sixth form college, while the less academically inclined would continue into vocational education.

A third important common feature is that in both countries, the school curriculum is structured by a set of academic and vocational credentials. The main certification thresholds are displayed in figures 1 and 2. When the Berthoin reform was initiated, the certificate of primary schooling examination (CEP[8]) had lost its importance as was no longer required to enter French junior secondary schools. The two main credentials that could be obtained by the end of secondary schooling were the *Baccalauréat*, for the most academically inclined students, and the CAP/BEP[9] vocational qualifications, for the less able. These qualifications were gained through examinations that students took at the

[8] *Certificat d'Études Primaires.*

[9] CAP: *Certificat d'Aptitude Professionnelle.* BEP: *Brevet d'Études Professionnelles.*

age of 17 or 18. Alongside these two major qualifications, students had the opportunity of obtaining a lower qualification (BEPC[10]) at the end of junior secondary schooling. However, this credential was never explicitly designed to certify the completion of compulsory education and very few individuals would hold it as their highest qualification.

Academic and vocational credentials also play an important role in the English & Welsh education system. When the 1972 ROSLA came into effect, two types of qualifications could be obtained at the end of compulsory education. At the age of 16, the most academically inclined pupils would prepare the General Certificate of Education Ordinary Level (GCE O-Level) while the others would prepare the Certificate of Secondary Education (CSE). Both credentials were awarded after an examination that took place at the end of secondary school[11]. Depending on their grades on the O-level or CSE, pupils could choose either to follow the academic track leading to the GCE Advanced Level (GCE A-Level) that would qualify them for higher education, or undertake a variety of basic or more advanced vocational qualifications (GNVQ level 1 to 4) which were usually provided by private institutions.

3 RDD estimation

Since both the 1967 Berthoin reform in France and the 1972 ROSLA in England & Wales were implemented uniformly in both countries for individuals born after a given date, Regression Discontinuity Design appears as the most appropriate estimation strategy to evaluate their impact on educational and labour market outcomes.

In the case of compulsory schooling laws, the discontinuity is created by the raising of the minimum school leaving age for cohorts born after a certain cutoff date c (January 1, 1953 in the case of the 1967 Berthoin reform and September 1, 1957 in the case of the 1972 ROSLA). Following most RDD studies of compulsory schooling laws (Oreopoulos, 2006b;

[10] *Brevet d'Études du Premier Cycle.*

[11] Note that there was an overlap between these two types of certificates in that a CSE grade 1 result was regarded as equivalent to an O-level.

Devereux and Hart, 2010), I estimate the first stage and reduced form effect of the raising of the minimum school leaving age by the means of a global polynomial approximation. This methodology involves using the whole sample and choosing a flexible high-order polynomial to fit the relationship between an outcome Y_i (school leaving age, wages, etc.) and the forcing variable X_i (school cohort), allowing for an intercept shift at the cutoff.

More specifically, the baseline first stage specification regresses age left full-time education S_i on the treatment variable T_i (which takes the value 0 for individuals born before the cutoff date c and 1 for those born after) and a quartic function of school cohort X_i:

$$S_i = \alpha_0 + \alpha_1 T_i + g(X_i - c) + \epsilon_i \tag{1}$$

where $g(\cdot)$ is a quartic function of school cohort and ϵ_i is an error term. In equation (1), the estimated coefficient $\hat{\alpha}_1$ on the treatment variable measures the reform's average causal effect on schooling at the assignment threshold c.

The reduced form specification regresses the log of hourly wages ($\ln W_i$) on the treatment variable T_i and a quartic function of school cohort X_i:

$$\ln W_i = \beta_0 + \beta_1 T_i + g(X_i - c) + \epsilon_i \tag{2}$$

In equation (2), the estimated coefficient $\hat{\beta}_1$ on the treatment variable measures the reform's average causal effect on hourly wages at the assignment threshold c.

Finally, the returns to compulsory schooling are estimated by 2SLS using the following specification:

$$\ln W_i = \gamma_0 + \gamma_1 S_i + g(X_i - c) + \epsilon_i \tag{3}$$

where the assignment variable T_i is used as an instrument for schooling. If the returns to compulsory education are heterogeneous, $\hat{\gamma}_1$ can be interpreted as measuring the returns to compulsory education of those individuals who were compelled to increase their educational attainment as a result of the new compulsory schooling requirement (the "compliers") and who belong to the first cohort facing the new minimum school leaving age. This will be true as long as the following monotonicity condition holds (Imbens and Angrist, 1994): raising compulsory schooling induces some individuals to increase their own educational attainment but no one to drop out earlier from school, which seems a fairly reasonable assumption.

Because the forcing variable X_i (school cohort) is discrete by nature, the group structure of the random specification errors needs to be taken into account. Following Lee and Card (2008), I obtain robust standard errors by clustering at the school cohort level.

4 Data

To perform the analysis of the 1967 Berthoin compulsory schooling reform, I use the French *Enquête Emploi* (EE) survey which is carried out every year by the French National Statistical Office. The working sample is constructed from the 13 surveys carried out between 1990 and 2002[12]. Individuals' school cohorts are given by their year of birth since school cohorts and calendar years coincide in the French educational system. Information on educational attainment is given by age completed full-time education and highest qualification held, which I group into six mutually exclusive categories: no qualification; primary schooling certificate; low vocational qualification (CAP or BEP without the BEPC); junior secondary schooling certificate (BEPC); intermediate vocational certificate (CAP or BEP with the BEPC); senior secondary schooling certificate and above (which includes the *Baccalauréat*, advanced vocational training, higher education degrees, etc.). Hourly earnings are obtained by dividing gross monthly wages (converted into 2005 euros) by the usual number of weekly hours worked. I use information on employment status to create two binary variables for employment and self-employment. To ensure that all included school cohorts are surveyed each year between 1990 and 2002, the sample is restricted to female and male respondents who were born in France between 1944 and 1962, left full-time education between the ages of 6 and 25 and were aged 25 to 60 when

[12] I do not include later surveys because the *Enquête Emploi*'s sampling plan and definition of variables changed in 2003.

surveyed. To avoid using several observations for the same individual, I further restrict the sample to respondents who were interviewed in the first of the three survey waves. Descriptive statistics for this sample of 155,562 observations are displayed in table 1.

To study the consequences of the raising of the minimum school leaving age to 16 in England & Wales, I use the Quarterly Labour Force Survey (QLFS) which provides the same type of information as the French EE. I combine the 53 surveys conducted between Fall 1993 and Winter 2006 to construct a large dataset that contains information on individuals' education and labour market outcomes. Because the school cutoff date in England & Wales is September 1, school cohorts are retrieved using respondents' year and month of birth: individuals born between September 1 of year t and August 31 of year $t+1$ are assigned to school cohort $t+1$[13]. In the QLFS, educational attainment is measured by age completed full-time education and highest qualification held, which I group into five mutually exclusive categories: no qualification; low vocational qualification (GNVQ level 1 and 2 or equivalent); junior secondary schooling certificate (GCE O-Level or CSE); intermediate vocational qualification (GNVQ level 3 or equivalent); senior secondary schooling certificate and above (GCE A-Level, GNVQ level 3 and 4, higher education degrees, etc.). Information on individual hourly pay is available only for individuals interviewed in the first and fifth wave of the survey. The corresponding wages are expressed in 2005 British pounds using the quarterly consumer price index series. Finally, I use the QLFS to construct binary variables for employment and self-employment status. Imposing similar restrictions to those used in the French case, the final sample includes 247,440 males and females who were born in England & Wales between September 1, 1948 and August 31, 1967, left full-time education between the ages of 6 and 25, were aged 25 to 60 when surveyed and were interviewed in the last of the five survey waves[14]. Descriptive statistics for this sample are displayed in table 2.

[13]The school cohort of an individual born on October 5, 1960 is therefore labeled 1961.

[14]I use the fifth rather than the first survey wave because of its smaller average non-response rate.

5 Results

As a starting point, I use the constructed samples to compute the OLS estimates of the returns to education in both countries. The values of these naive estimates provide a natural benchmark for the causal returns to compulsory schooling that I evaluate using the 1967 Berthoin reform and the 1972 ROSLA. The standard OLS specification regresses the log of hourly wage on the age left full-time education and includes a quartic in age and a full-set of survey year dummies.

The naive estimates of the returns to education, which are reported separately for female and male workers in the first column of table 3, are in the range of 7 to 12% in both countries. The returns appear higher for women than for men (.087 vs. .073 in France and .119 vs. .095 in England & Wales) and slightly larger in England & Wales than in France.

Figures 3 and 4 provide graphical evidence on the impact of French and British compulsory schooling reforms on the average school leaving age, for female and male workers separately. The solid lines show the fitted values from a global fourth order polynomial allowing for an intercept shift at the cutoff point. Visual inspection suggests that the average school leaving age of women and men follows an almost linear trend and exhibits a substantial "jump" at the cutoff point in both countries. The first stage RDD estimates, which are reported in the second column of table 3, indicate that the raising of the minimum school leaving age to 16 in France increased the average number of years of education by .272 (.035) for women and by .264 (.025) for men. The corresponding estimates for England & Wales are of the same order of magnitude, i.e. .309 (.023) for women and .267 (.028) for men[15].

[15] A detailed examination of the reforms' impact on the distribution of school leaving ages is provided in the online appendix of this paper. The analysis indicates that in both countries, about a third of the individuals belonging to the last pre-reform school cohort left school before reaching the age of 16. In France, about half of the individuals "caught up" by the new minimum school leaving age in France complied with the new mandatory requirement, while the other half did not. In England & Wales, the fraction of compliers is slightly higher (about three quarters).

Despite their similar impact on schooling, the 1967 Berthoin reform and the 1972 ROSLA had very different effects on wages. Figures 5 and 6 plot the evolution of log of hourly wages across school cohorts, separately for female and male workers[16]. While their is no graphical evidence of a wage discontinuity around the 1953 cohort in France, the hourly wage of female and male workers in England & Wales appears to have been positively impacted by the raising of the minimum school leaving age to 16. The econometric analysis confirms this visual impression. The reduced form RDD estimates of the reforms' impact on hourly wages are reported in column 4 of table 3. In the French case, the estimates are statistically insignificant and close to zero for both female (-.000) and male (.001) workers, which suggests very small wage effects from the change in compulsory schooling. On the contrary, the estimates for England & Wales are positive and significant with a coefficient of .019 (significant at the 1% level) for women and of .019 (significant at the 5% level) for men.

The corresponding 2SLS estimates of the returns to compulsory education in France and England & Wales are reported in column 8. In line with the reduced form estimates, the economic return to compulsory schooling is insignificant and close to zero in France for women (-.001) and for men (.005). In England & Wales, the estimated returns are all positive and significant with a value of .060 (.011) for women and .069 (.028) for men.

I performed a number of robustness checks to assess the stability of these estimates to alternative specifications. Because earnings typically follow a life-cycle pattern, one may wonder whether including age controls might affect the results. To address this issue, I added a quartic in age to the previously specified global polynomial regressions. The first stage, reduced form and 2SLS estimates reported in columns 3, 5 and 9 are very close to

[16]Cross-sectional age-wage profiles indicate that in the UK, women's hourly earnings reach their maximum in their early thirties, while for men the peak is later, i.e. in their early forties (NEP, 2010). This phenomenon explains why contrary to what is observed for males in England & Wales and for both females and males in France, the female profile of hourly wages by cohort in England and Wales (figure 6(a)) is increasing by year of birth (in the sample, the average age of women in school cohort 1949 is 32 vs. 50 for those in school cohort 1967).

those obtained without controlling for age effects[17].

A second set of robustness checks involved estimating jumps at points where there should be no discontinuities. Following the guidelines set by Imbens and Lemieux (2007), I tested for jumps in average educational attainment and hourly wages at the median of the two subsamples of pre- and post-reform school cohorts in each country. These placebo cutoff points are 1948 and 1958 in France and 1953 and 1963 in England & Wales. The results, which are reported in the online appendix, indicate that the corresponding first stage, reduced form and 2SLS estimates are all statistically insignificant and close to zero.

Finally, I assessed the stability of my estimates by using a non-parametric RDD estimator based on local linear regression (LLR) rather than on a parametric global polynomial approximation. I used a triangular kernel to weight the observations and selected the bandwidth (i.e. the school cohort range) based on the cross-validation developed by Ludwig and Miller (2007). The estimated returns to compulsory schooling, which are included in the online appendix, are very close to those obtained using the parametric approach.

6 An interpretation based on certification effects

While this paper is not the first to challenge the standard view that raising the minimum school leaving age systematically improves the labour market prospects of early school dropouts[18], I believe than none of the interpretations that have been offered in the literature to explain variations in the returns to compulsory schooling can convincingly account for the fact that the 1967 Berthoin reform failed to increase earnings in France while the 1972 ROSLA succeeded in doing so in England & Wales. In this section, I

[17] To ensure that my results are not sensitive to differences in the average age at which individuals are observed in the French and British samples (i.e. 43 vs. 41), I estimated the impact of the French Berthoin reform on the subsample of respondents interviewed between 1990 and 1999. The average age of individuals included in this subsample is the same as in the full British sample (i.e. 41). The corresponding estimates, which can be found in the online appendix of the paper, are very close to those obtained with the full sample.

[18] As emphasized in the introduction, a number of recent papers find small and possibly zero returns to compulsory schooling in the European context (Oosterbeek and Webbink, 2007; Pischke and von Wachter, 2008; Devereux and Hart, 2010).

present suggestive evidence consistent with an explanation based on certification effects, before attempting to rule out alternative interpretations in the next section.

While the actual number of years of education is certainly an important outcome to consider when assessing the effect of increases in the minimum school leaving age, it is not necessarily the most adequate way of measuring the educational benefits of compulsory schooling laws. This is especially true in many European countries where the school curricula are often structured to allow each level of study to lead to a specific academic or vocational credential. These credentials play a significant role in the job hiring process, by providing employers with a way of verifying that an applicant has attained a certain level of competency in a specific field.

The empirical analysis reveals that despite their similar impact on average educational attainment, the Berthoin reform in France and the 1972 ROSLA in England & Wales had very different effects on the distribution of credentials. Figures 7 and 8 show how the distribution of academic and vocational qualifications was affected by the raising of the minimum school leaving age in both countries, with the corresponding reduced form RDD estimates reported in table 4. In the French case (columns 1 and 2), there is no clear sign of a discontinuity around the 1953 school cohort for both women and men and the RDD estimates indicate only a very small and marginally significant increase (less than 1 point) in the fraction of individuals holding a certificate of junior secondary schooling (BEPC). On the contrary, a strong discontinuity can be observed in the case of England & Wales (columns 3 and 4) around the 1958 school cohort: the fraction of respondents holding no qualification drops sharply (by 11.4 points for women and 5.7 points for men) while the fraction of respondents holding a certificate of junior secondary schooling (O-Level or CSE) goes up by a similar percentage (12.6 points for females and 5.4 points for men). These results indicate that the improvement in the level of academic credentials induced by the 1972 ROSLA was large and more important for women than for men[19].

[19]However, because women were twice more likely than men to hold an O-Level or CSE as their highest

Three institutional features of the French and British school systems are likely to explain the contrasting impact of the Berthoin reform and the 1972 ROSLA on the distribution of academic credentials.

First, junior secondary schooling certification is less central in the French schooling system than it is in England & Wales. In France, the BEPC was never explicitly meant to manifest the fulfilment of compulsory education. The examinations, which took place at the end of the final year of junior secondary schooling for pupils theoretically aged between 14 and 15, were not mandatory and many junior secondary schooling pupils would simply not sit for them[20]. On the contrary, the GCE O-Level and CSE certificates in England & Wales were explicitly designed to serve as a standard of qualification for pupils who reached the end of compulsory education. Although not mandatory, most pupils would prepare the O-Level/CSE examinations during the two final years of junior secondary schooling (i.e. from the age of 13-14) and sit for the examinations at the age of 15-16. In this context, the most likely explanation for the sudden jump in the fraction of English and Welsh respondents holding a junior secondary schooling certificate as their highest qualification is that the raising of the minimum school leaving age to 16 exogenously induced a number early school dropouts to sit for the examinations and, for some of them, to be awarded the credential[21].

The prevalence of grade repetition in France is a second factor explaining why the Berthoin reform failed to improve the distribution of qualifications in France. While

qualification before the implementation of the 1972 ROSLA, the proportional increase in the fraction of men holding a junior secondary schooling certification is roughly equivalent to the corresponding increase for women (i.e. about 30%).

[20]The fairly marginal status of BEPC in the French school system can be clearly inferred from the graphs displayed in figure 7, which show that among the respondents who were born before 1953, very few (about 10%) declare holding the BEPC as their highest academic qualification. The fact that this fraction remained unchanged after the introduction of the new minimum school leaving age should not therefore be viewed as completely surprising.

[21]In the absence of a separate source of identification, the exact contribution of changes in certification levels to the overall wage increase in England & Wales cannot be directly estimated. However, the results in tables 3 and 4 can be combined to yield an upper bound for the returns to holding a junior high school certificate (as opposed to no qualifications at all). The results, which are discussed in the online appendix, yield estimated returns to the O-Level/CSE of 16% for women and 34% for men. These returns are of the same order of magnitude as existing estimates derived from OLS specifications (McIntosh, 2006).

hardly any children are held back in England & Wales if they failed to meet the academic targets, grade repetition has always been a very common practice in France. It can be inferred from retrospective statistics released in 1969 and 1978 by the French Ministry of Education that most pupils born in 1952-1953 repeated at least once in primary school: by the end of primary schooling, 60% were one or two years behind. Hence many of the pupils who had to comply with the new minimum school leaving age requirement reached the age of 16 well before the period of the junior secondary schooling certification examinations, enabling them to drop out of school without even sitting for the exam. On the contrary, when the 1972 ROSLA came into effect, virtually all 15 to 16-year-old English & Welsh pupils were enrolled in year 11, at the end of which the O-Level/CSE examinations took place.

A third explanatory factor is related to the organization of vocational tracking. In the French context, the raising of the minimum school leaving age had little chances of increasing the number of vocational qualifications awarded to early school dropouts because the CAP/BEP vocational credentials were typically awarded at the end of a 3-year curriculum, i.e. beyond the minimum school leaving age.

These observations raise the question of whether the positive wage returns to better academic qualifications in England & Wales were driven by a signalling mechanism or by an improvement in productivity. According to the first explanation, the productivity of early school dropouts in England & Wales may not have increased after the introduction of the new minimum school leaving age, but their earnings were higher simply because the signaling value (or "sheepskin" effect) of junior secondary schooling credentials remained unchanged. By contrast, the second interpretation would argue that their exists some degree of "lumpiness" in the learning process, which leads to more skill acquisition in examination years than in preceding years. In support of this view, several empirical studies have found that curriculum-based external exit examinations are associated with higher student learning outcomes (Bishop, 1997; Jürges and Schneider, 2010). Hence a plausible channel for the positive wage effect of the 1972 ROSLA is that the preparation of the O-Level/CSE examinations in year 11 induced pupils in England & Wales to work harder and to learn more labour market relevant skills than would have been the case without these examinations.

Obviously, these two lines of interpretations have very different policy implications. Under the first scenario, the extent to which compulsory schooling laws can improve the labour market prospects of early school dropouts appears fairly limited, since the signalling value of credentials remains unaffected only as long as the distribution of academic qualifications does not shift too much. Under the second scenario, raising the minimum school leaving age could serve as an efficient productivity-enhancing tool if it induces more pupils to acquire the skills that are needed to succeed on certification exams and that are ultimately rewarded in the labour market.

While it is difficult with the available data to distinguish between these two competing explanations, I believe that the signalling theory is too extreme to fully account for the positive returns enjoyed by early school dropouts after the introduction of the new minimum school leaving age in England & Wales. First, the shift in the distribution of academic credentials seems too large to have been completely unnoticed by employers when the post-reform cohorts entered the labour market. Second, I provide some empirical evidence supporting the view that the 1972 ROSLA actually improved the cognitive skills of early school dropouts, using data from the British sample of the 1994 International Adult Literacy Survey (IALS), which was designed to assess the literacy skills of adults in a number of OECD countries[22]. The specific literacy tasks designed for IALS were scaled by difficulty from 0 to 500 points and each participant received a global score for prose, document, and quantitative literacy. While the relatively small size of the sample (about 1,500 individuals) inevitably limits the statistical power of the RDD approach,

[22] Unfortunately, data on the literacy skills of French participants to the IALS are not available because France decided to withdraw from the program after the publication of the preliminary results.

the reduced form estimates presented in table 5 suggest that the raising of the minimum school leaving age to 16 in England & Wales led to a positive and significant improvement in literacy skills of about 13 points (i.e. about one sixth of a s.d.) on all three scales for women, and a positive but only marginally significant impact on the quantitative literacy skills of men. These estimates are consistent with the view that in England & Wales, the benefits derived by early school dropouts from better credentials were not entirely driven by signalling but also reflected an actual improvement in labour market relevant skills.

While certifications effects are certainly not the only explanation for observed cross-country differences in returns to compulsory schooling, I believe that they deserve more attention than they have received in the literature, since they can be viewed as a plausible channel to reconcile some of the divergent patterns that have been found in the literature on compulsory schooling laws in Europe, where contrary to the US or Canada, most countries certify the successful completion of compulsory education by the means of examination-based diplomas[23]. For instance, the lack of wage gains in the cases of the Dutch compulsory education reform of 1975 (Oosterbeek and Webbink, 2007) and the German schooling reforms of 1947-1969 (Pischke and von Wachter, 2008) is at least qualitatively consistent with the certification hypothesis. Indeed, a striking feature of these reforms is that certification thresholds were shifted by one year when the extra compulsory schooling year was introduced, thereby preventing substantial improvements in graduation rates. Conversely, the fact that according to Brandolini and Cipollone (2002)'s estimates, the raising of the minimum school leaving age to 16 in Italy had a positive impact on wages could derive from the fact that this reform increased the fraction of individuals holding a junior high school diploma.

[23]e.g. the *Diploma di scuola media inferiore* in Italy, the *Graduado en educación secundaria obligatoria* in Spain, the *Diploma do ensino básico* in Portugal, the *Folkeskolens afgangsprøve* in Denmark, the *Slutbetyg från grundskolan* in Sweden, the *Avgångsbetyg från grundskola* in Finland, etc.

7 Discussion of alternative explanations

The evidence discussed so far is consistent with an interpretation based on certification effects. However, several alternative explanations could be proposed to account for the divergent wage effects of the French and British compulsory schooling laws. In this section, I present suggestive evidence that attempts to rule out some of these explanations.

A defective implementation of the Berthoin reform? An obvious explanation for the lack of wage gains in France is that the Berthoin reform might not have been effectively implemented. If, for instance, low achieving pupils who were compelled to stay in school longer were in fact kept in the same grade until they reached the new minimum school leaving age, it would be hardly surprising to find zero returns to the extension of compulsory education.

Precise statistical information on the educational careers of individuals born in 1952 and 1953 can be found in the retrospective statistics released by the French Ministry of Education in 1969 and 1978. The distribution of pupils across the different school levels by cohort and age, which is shown in table 6, indicates that the fraction attending school at the age of 15-16 went up by about 13.6 ppt between cohorts 1952 and 1953. Most of this increase was driven by junior and senior high school enrollment (+8.2 ppt) and to a lesser extent, by vocational education (+3.5 ppt), while extended primary schooling and special education account for only a small share of this increase (+1.9 ppt.). Compared to the previous cohort, pupils born in 1953 were more likely to be enrolled in a junior secondary school form the age of 11-12 onwards than to spend these years attending extended primary schooling[24]. This analysis suggests that when the Berthoin reform was implemented, the French schooling system had made the necessary investments (including the construction of new junior secondary schools) to accommodate the sudden increase in

[24]The fraction of pupils enrolled in junior high schools at the age of 12-13 increased went up from 49.2 to 54.4% between cohorts 1952 and 1953, while the fraction enrolled in primary or extended primary schools fell from 49 to 43.6%.

total enrollment and to avoid early school dropouts born after 1953 from being artificially kept in the same grade until they reached the age of 16. It seems therefore difficult to attribute the lack of positive returns to compulsory schooling in France to a defective implementation of the Berthoin reform.

Heterogeneous returns to compulsory schooling? A common interpretation of cross-country variations in the estimated returns to compulsory schooling is that these returns might be fundamentally heterogeneous across individuals. According to this view, the size of the local average treatment effect that is measured using changes in the minimum school leaving age is a function of the characteristics of the compliers, which are likely to differ from one reform to another[25]. I believe, however, that heterogeneous returns to schooling cannot credibly explain the divergence between France and England & Wales. First, the fact that my estimates for the 1972 ROSLA are of the same order of magnitude as those reported by Devereux and Hart (2010) for the 1947 reform suggest that modest returns can arise even when the fraction of compliers is relatively small (the proportion of individuals impacted by the 1972 law change was roughly 20% versus 50% for the 1947 reform). Second, the discrepancy between French and British estimates seems unlikely to be driven by the characteristics of compliers, since both laws raised the minimum school leaving age to 16 at about the same period and impacted the schooling of pupils belonging to the bottom 30% of the distribution of school leaving ages in each country. Moreover, the comparison of compliance rates across a number of pre-determined characteristics suggests that compliers in both countries share reasonably similar attributes (see online appendix). Finally, differences in the relative supply of high vs. low skills workers in both countries seem too small to explain the divergent returns[26].

[25]For instance, Devereux and Hart (2010) claim that a possible reason why Canadian or US compulsory schooling laws seem to yield stronger returns than the 1947 Butler Act in the UK is that the latter impacted a much larger fraction of the population than the former. Their estimates for the UK would then be picking up something closer to the average treatment effect in the population, whereas US and Canadian estimates would reflect the high returns of a small number of compliers.

[26]A higher relative supply of skilled workers could induce lower returns to skills. However, the relative supply of high vs. low skill workers among pre-reform cohorts was not unambiguously higher in France

Wage rigidity? A third potential explanation for my findings pertains to the role of wage rigidity. The influence of trade unions on the setting of wages and the size of the public sector are two commonly cited determinants of wage rigidity that might be considered as possible explanations for the lack of wage gains in the French case. However, the empirical evidence suggests that both these factors are unlikely to account for the observed discrepancy between France and England & Wales. Although the fraction of workers who are covered by collective bargaining agreements in much higher in France than in the UK (according to OECD figures, the coverage rate in 2000 was 90% in France vs. 30% in the UK), these agreements have only limited importance in determining individual wages in France. Most of them are not focused on wage negotiations but rather on issues such that training, retirement, employment contracts, gender equality, hours of work and work organization. Moreover, a recent report by the French ministry of Labour (DARES, 2006) indicates that for the majority of French employees working in the private sector, pay rises are negotiated on a individual basis[27]. These facts suggest that the structure of collective bargaining is not the most likely explanation for the zero return to compulsory schooling in France.

Wage rigidity in the public sector is another factor that could be invoked to explain the lack of wages gains in France. However, two empirical observations seem to rule out this possibility. First, the share of public sector employment was similar in both countries over the periods covered by my study. In the French sample, 23% of workers are employed in the public sector whereas in the British sample, the fraction is 24%. Second, separate regressions for public and private sector employees in France yield zero returns

than in England & Wales. On the one hand, the average school leaving age was slightly higher in France than in England & Wales (16.2 vs. 16) but on the other hand, the fraction of individuals holding a qualification above primary school certification was lower (62% vs. 69%).

[27]Based on administrative records that cover all workers employed in private firms with more than 10 employees, the report shows that only 34% of workers whose pay went up in 2004 (representing 86.4% of all workers) depended on a general increase (at the firm level), whereas 49% saw their pay increase through the combination of a general increase and an individual pay rise, and 17% depended entirely on an individual pay rise. According to the same study, the individualization of wage increases was even more prevalent in 2000, since less than 26% of workers whose pay went up depended on a general increase only.

to compulsory schooling[28]. These results suggest that the French vs. British discrepancy is not directly related to public sector wage rigidity.

In the French and British contexts, the minimum wage is probably the most important source of wage rigidity at the bottom of the earnings distribution. The reason why this labour market institution could play a key role is explaining my findings is that France has a fairly high minimum wage since 1950 whereas Britain did not introduce a national minimum wage before 1999 at a fairly low initial level, before gradually increasing it[29]. A potential explanation for the zero returns to the French Berthoin is that since early school dropouts are overrepresented at the bottom end of the wage distribution, where the minimum wage is known to have a large "bite", the minimum wage might have prevented their higher productivity from being reflected in their earnings.

I explore this possibility by testing two implications of the minimum wage hypothesis. First, if this was the main explanation for the lack of monetary gains from the French reform, one would expect the raising of the minimum school leaving age to have translated into higher employment rates for early school dropouts, by reducing the fraction of individuals whose productivity fell below the minimum wage threshold. The empirical evidence does not, however, support this hypothesis. The reduced form RDD estimates of the impact of the raising of the minimum school leaving age in France on the probability of being in employment, which are reported in column 6 of table 3, are insignificant and close to zero[30]. The same analysis for England & Wales indicates that if anything, the 1972 ROSLA may have induced marginal employment gains for women of +0.7 ppt in

[28]The regression tables can be found in the online appendix of the paper.

[29]In my samples, the level of the minimum hourly wage in 2002 was 57% of the median wage in France vs. 42% in England & Wales.

[30]Since the employment effects of the raising of the minimum school leaving age are likely to have been be larger for young workers than for older ones, I estimated similar regressions using the French *Enquête Emploi* that were conducted between 1975 and 1989. Although these pre-1990 surveys do not include information on earnings, they do provide information on employment status. By construction, individuals in that sample are observed at a younger age (31) than those included in the 1990-2002 sample (43). The reduced form effects of the impact of the Berthoin reform on the employment rates of these younger women and men are not significantly different from zero (see online appendix), which rules out the existence of "hidden" labour market returns from the Berthoin reform, in the form of higher employment rates rather than higher wages.

addition to the previously estimated monetary gains.

A second implication of the wage rigidity argument is that the individuals whose educational attainment was positively affected by the raising of the minimum school leaving age would be more likely to become self-employed if their higher productivity was not rewarded in the wage and salary sector. Following Pischke and von Wachter (2008), I test this hypothesis by focusing on the evolution of the fraction of self-employed workers across school cohorts. As in the case of employment rates, the reduced form RDD estimates reported in column 7 of table 3 provide no support to the wage rigidity argument, since self-employment rates were unaffected by the French and British reforms.

8 Conclusion

The evidence presented in this paper is consistent with the view that the 1972 ROSLA in England & Wales yielded positive wage returns because it induced a number of early school dropouts, and especially female pupils, to complete junior secondary schooling and to learn the skills that are needed to succeed in the final examinations and are ultimately rewarded in the labour market. On the contrary, because the French 1967 Berthoin reform was not explicitly designed to reduce the fraction of pupils dropping out of school without any academic qualification, it failed to raise the labour market outcomes of those who complied with the new mandatory requirements.

Although the results of this study must be appreciated in light of the limitations of any country-level assessment of educational reforms, they suggest that raising the minimum school leaving age is not necessarily an efficient way of improving the labour market prospects of early school dropouts and that the particular organization of national school systems is an important dimension to take into account when assessing the benefits of compulsory education.

References

Acemoglu, D. and Angrist, J. (2000). "How Large are Human Capital Externalities? Evidence from Compulsory Schooling Laws". In *NBER Macroeconomics Annual 2000*, volume 15, pp. 9–59. MIT Press, Cambridge, MA.

Bishop, J. (1997). "The Effect of National Standards and Curriculum-Based Exams on Achievementl". *American Economic Review, Papers and Proceedings*, 87(2), pp. 260–264.

Black, S., Devereux, P., and Salvanes, K. (2008). "Staying in the Classroom and Out of the Maternity Ward? The Effect of Compulsory Schooling Laws on Teenage Births". *Economic Journal*, 118(530), pp. 1025–1054.

Brandolini, A. and Cipollone, P. (2002). "Return to education in Italy: 1992-1997". unpublished manuscript, Bank of Italy.

Brunello, G. and Miniaci, R. (1999). "The Economic Return to Schooling for Italian Men. An Evaluation Based on Instrumental Variables". *Labour Economics*, 6(4), pp. 509–519.

Callan, T. and Harmon, C. (1999). "The Economic Return to Schooling in Ireland". *Labour Economics*, 6(4), pp. 543–550.

Card, D. (1999). "The Causal Effect of Education on Earnings". In Ashenfelter, O. and Card, D., editors, *Handbook of Labour Economics*, volume 3a. North Holland, Amsterdam.

DARES (2006). "L'individualisation des hausses de salaires de base: près de six salariés sur dix concernés". *Premières Informations*, No 09.1.

Devereux, P. and Hart, R. (2010). "Forced to be Rich? Returns to Compulsory Schooling in Britain". *Economic Journal*, 120(549), pp. 1345–1364.

Harmon, C. and Walker, I. (1995). "Estimates of the Economic Return to Schooling for the United Kingdom". *American Economic Review*, 85(5), pp. 1278–1286.

Imbens, G. and Angrist, J. (1994). "Identification and Estimation of Local Average Treatment Effects". *Econometrica*, 62(2), pp. 467–475.

Imbens, G. and Lemieux, T. (2007). "Regression Discontinuity Designs: A Guide to Practice". *Journal of Econometrics*, 127(2), pp. 615–635.

Jürges, H. and Schneider, K. (2010). "Central Exit Examinations Increase Performance... but Take the Fun out of Mathematics". *Journal of Population Economics*, 23(2), pp. 497–517.

Lee, D. and Card, D. (2008). "Regression Discontinuity Inference with Specification Error". *Journal of Econometrics*, 142(2), pp. 655–674.

Levine, J. and Plug, E. (1999). "Instrumenting Education and the Returns to Schooling in the Netherlands". *Labour Economics*, 6(4), pp. 521–534.

Lleras-Muney, A. (2005). "The Relationship Between Education and Adult Mortality in the United States". *Review of Economic Studies*, 72(1), pp. 189–221.

Lochner, L. and Moretti, E. (2004). “The Effect of Education on Crime: Evidence from Prison Inmates, Arrests, and Self-Reports”. *American Economic Review*, 94(1), pp. 155–189.

Ludwig, J. and Miller, D. (2007). “Does Head Start Improve Children’s Life Chances? Evidence from a Regression Discontinuity Design”. *Quarterly Journal of Economics*, 122(1), pp. 159–208.

McIntosh, S. (2006). “Further Analysis of the Returns to Academic and Vocational Qualifications”. *Oxford Bulletin of Economics and Statistics*, 68, pp. 225–251.

Meghir, C. and Palme, M. (2005). “Educational reform, Ability and Parental Background”. *American Economic Review*, 95(1), pp. 414–424.

Milligan, K., Moretti, E., and Oreopoulos, P. (2004). “Does Education Improve Citizenship? Evidence from the United States and the United Kingdom”. *Journal of Public Economics*, 88(9-10), pp. 1667–1695.

Ministère de l’Éducation Nationale (1969). *Tableaux de l’Éducation Nationale. Statistiques rétrospectives 1958-1968.* Service Central des Statistiques, Paris.

Ministère de l’Éducation Nationale (1978). *Tableaux des Enseignements et de la Formation. Statistiques rétrospectives 1967-1977.* Service Central des Statistiques, Paris.

NEP (2010). *An Anatomy of Economic Inequality in the UK.* National Equality Panel, London.

Oosterbeek, H. and Webbink, D. (2007). “Wage Effects of an Extra Year of Basic Vocational Education”. *Economics of Education Review*, 26, pp. 408–419.

Oreopoulos, P. (2006a). “Estimating Average and Local Average Treatment Effects of Education when Compulsory Schooling Laws Really Matter”. *American Economic Review*, 96(1), pp. 152–175.

Oreopoulos, P. (2006b). “The Compelling Effects of Compulsory Schooling: Evidence from Canada”. *Canadian Journal of Economics*, 39(1), pp. 22–52.

Oreopoulos, P. (2007). “Do Dropouts Drop out too Soon? Wealth, Health and Happiness from Compulsory Schooling”. *Journal of Public Economics*, 91(11-12), pp. 2213–2229.

Oreopoulos, P., Page, M., and Stevens, A. (2006). “The Intergenerational Effects of Compulsory Schooling”. *Journal of Labor Economics*, 24(4), pp. 729–760.

Pischke, J.-S. and von Wachter, T. (2008). “Zero Returns to Compulsory Schooling in Germany: Evidence and Interpretation”. *Review of Economics and Statistics*, 90(3), pp. 592–598.

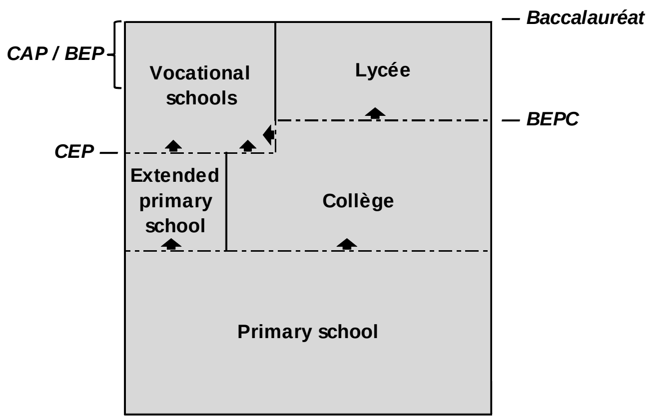
Pons, E. and Gonzalo, M. (2002). “Returns to Schooling in Spain: How Reliable are Instrumental Variable Estimates”. *Labour*, 16(4), pp. 747–770.

Vieira, J. (1999). “Returns to Education in Portugal”. *Labour Economics*, 6(4), pp. 535–541.

Figure 1: *The French school system at the time of the 1967 Berthoin reform.*

Minimum school leaving age	Age on Dec 31*	School year level	Phase of education
	17	Y12	Upper secondary education
After the reform: 16	16	Y11	
	15	Y10	
Before the reform: 14	14	Y9	
	13	Y8	Lower secondary education
	12	Y7	
	11	Y6	
	10	Y5	Primary education
	9	Y4	
	8	Y3	
	7	Y2	
	6	Y1	

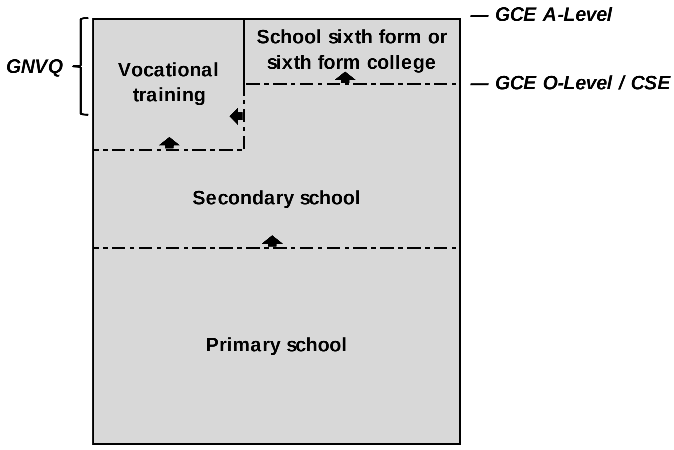
*excluding grade repeaters



Notes: The left hand side columns show age and school year levels running along the main phases of education. Two thick lines show the position of the minimum school leaving age before and after the 1967 Berthoin reform. The right hand side cells show the different types of schooling corresponding to each school year level, with arrows indicating where the tracking of students occurs. the diagram shows the placement of the main certification thresholds: CEP (*Certificat d'Études Primaires*), BEPC (*Brevet d'Études du Premier Cycle*), *Baccalauréat* and CAP/BEP (*Certificat d'Aptitude Professionnelle / Brevet d'Études Professionnelles*).

Figure 2: *The English & Welsh school system at the time of the 1972 ROSLA.*

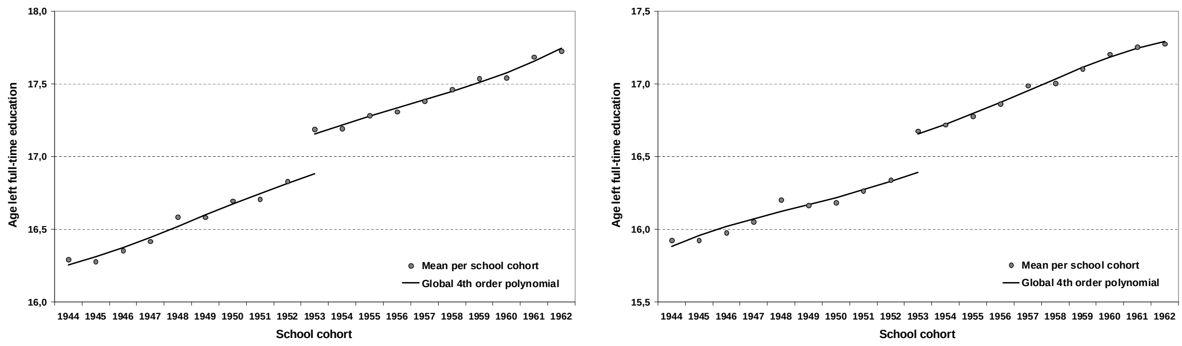
Minimum school leaving age	Age on Aug 31	School year level	Phase of education
	17	Y13	Upper secondary education
After the reform: 16	16	Y12	
Before the reform: 15	15	Y11	
	14	Y10	
	13	Y9	Lower secondary education
	12	Y8	
	11	Y7	
	10	Y6	Primary education
	9	Y5	
	8	Y4	
	7	Y3	
	6	Y2	
	5	Y1	



Notes: The left hand side columns show age and school year levels running along the main phases of education. Two thick lines show the position of the minimum school leaving age before and after the 1972 ROSLA. The right hand side cells show the different types of schooling corresponding to each school year level, with arrows indicating where the tracking of students occurs. the diagram shows the placement of the main certification thresholds: GCE O-Level (General Certificate of Education - O-Level), CSE (Certificate of Secondary Education), GCE A-Level (General Certificate of Education - A-Level) and GNVQ (General National Vocational Qualification).

Figure 3: *Impact of the 1967 Berthoin reform in France on the average age left full-time education, for female and male workers separately. School cohorts 1944-1962. Source:* Enquête Emploi *(1990-2002).*

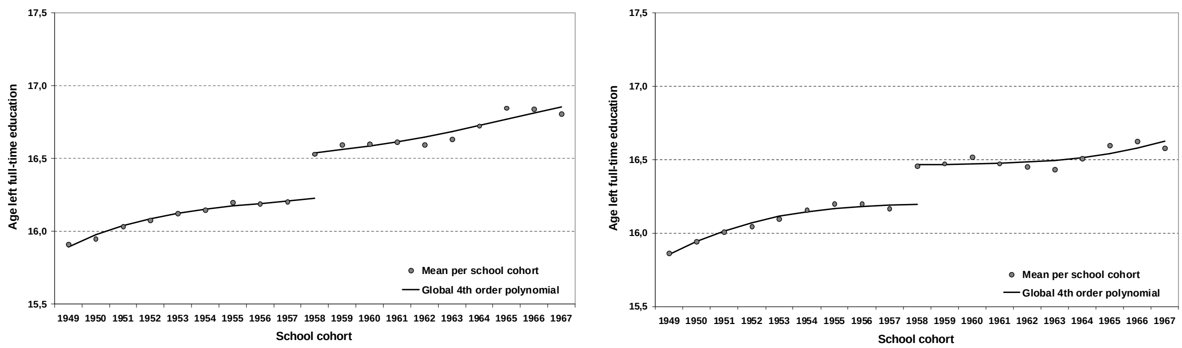
(a) Women (b) Men



Notes: The dots show the average age left full-time education grouped at the school cohort cell for the subsample of French female and male wage earners who were born between 1944 and 1962. The solid line represents the fitted values from a global fourth order polynomial regression allowing for an intercept shift at the 1953 school cohort.

Figure 4: *Impact of the 1972 ROSLA in England & Wales on the average age left full-time education, calculated for female and male workers separately. School cohorts 1949-1967. Source: Quarterly Labour Force Survey (1993-2006).*

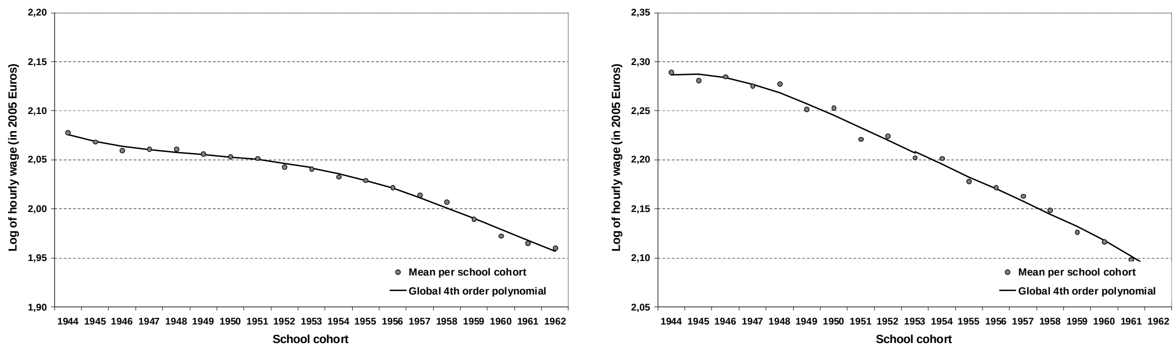
(a) Women (b) Men



Notes: The dots show the average age left full-time education grouped at the cohort cell for the subsample of English and Welsh female and male wage earners who belong to school cohorts 1949 to 1967 (i.e. born between September 1, 1948 and August 31, 1967). The solid line represents the fitted values from a global fourth order polynomial regression allowing for an intercept shift at the 1958 school cohort.

Figure 5: *Impact of the 1967 Berthoin reform in France on hourly wages (in 2005 euros), calculated for women and men separately. School cohorts 1944-1962. Source:* Enquête Emploi *(1990-2002).*

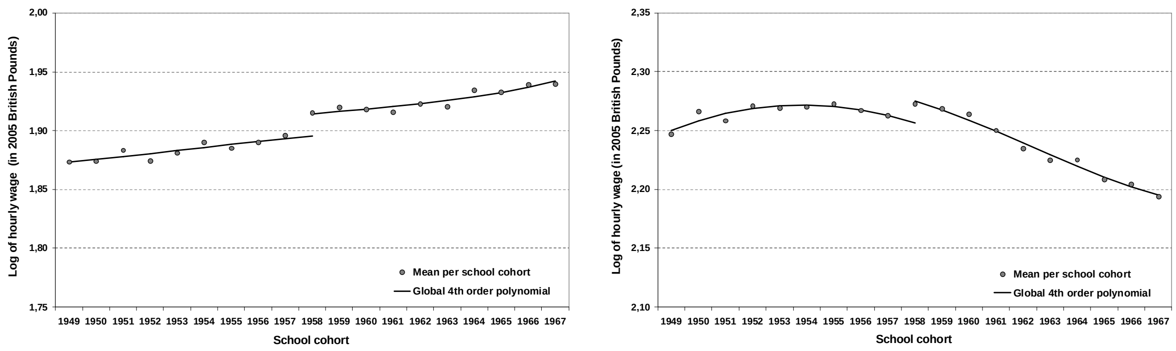
(a) Women (b) Men



Notes: The dots show the average log of hourly wage grouped at the school cohort cell for the subsample of French female and male wage earners who were born between 1944 and 1962. The solid line represents the fitted values from a global fourth order polynomial regression allowing for an shift at the 1953 school cohort.

Figure 6: *Impact of the 1972 ROSLA in England & Wales on hourly wages (in 2005 pounds), calculated for women and men separately. School cohorts 1949-1967. Source: Quarterly Labour Force Survey (1993-2006).*

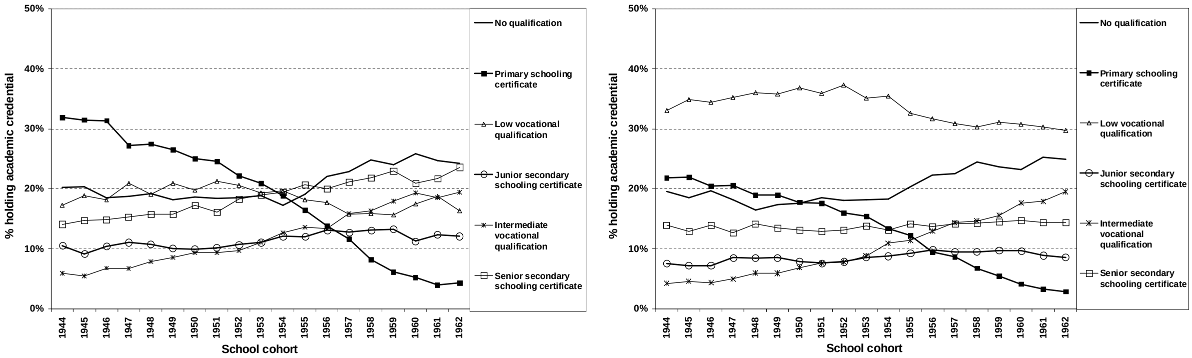
(a) Women (b) Men



Notes: The dots show the average log of hourly wage grouped at the school cohort cell for the subsample of English and Welsh female and male wage earners who belong to school cohorts 1949 to 1967 (i.e. born between September 1, 1948 and August 31, 1967). The solid line represents the fitted values from a global fourth order polynomial regression allowing for an intercept shift at the 1958 school cohort.

Figure 7: *Impact of the 1967 Berthoin reform in France on the distribution of academic and vocational credentials, calculated for women and men separately. School cohorts 1944-1962. Source:* Enquête Emploi *(1990-2002).*

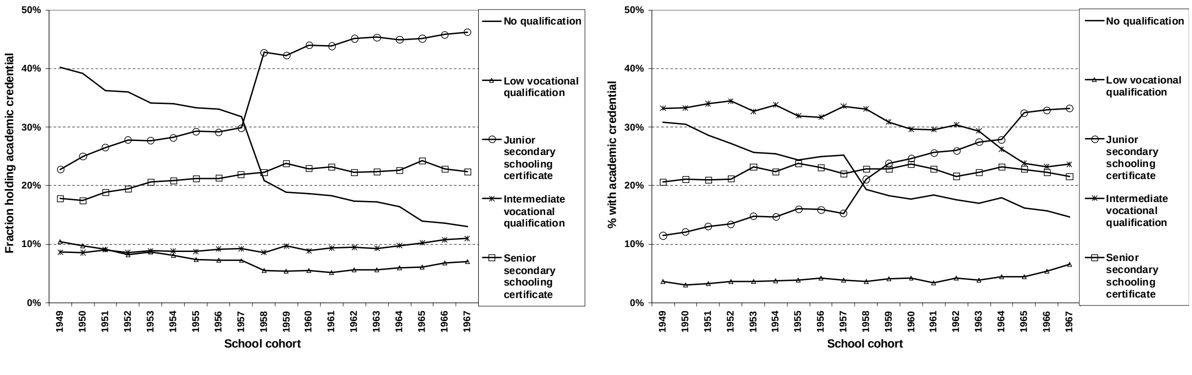
(a) Women (b) Men



Notes: This graph displays the evolution of the distribution of academic and vocational qualifications across school cohorts in France for the subsample of women and men born between 1944 and 1962. Academic and vocational credentials are grouped into six categories: no qualification; primary schooling certificate (*Certificat d'Études Primaires Élémentaires*); low vocational qualification (CAP or BEP without the BEPC); junior secondary schooling certificate (BEPC); intermediate vocational qualification (CAP or BEP with the BEPC); senior secondary schooling certificate or above (includes the *Baccalauréat*, advanced vocational training, higher education degrees, etc.).

Figure 8: *Impact of the 1972 ROSLA in England & Wales on the distribution of academic and vocational credentials, calculated for women and men separately. School cohorts 1949-1967. Source: Quarterly Labour Force Survey (1993-2006).*

(a) Women (b) Men



Notes: This graph displays the evolution of the distribution of academic and vocational qualifications across school cohorts in the England & Wales for the subsample of women and men who belong to school cohorts 1949 to 1967 (i.e. born between September 1, 1948 and August 31, 1967). Academic and vocational credentials are grouped into five categories: no qualification; low vocational qualification (NVQ level 1 and 2 or equivalent); junior secondary schooling certificate (GCE O-Level or CSE); intermediate vocational qualification (NVQ level 3 or equivalent); senior secondary schooling certificate or above (includes GCE A-Level, NVQ level 3 and 4, higher education degrees, etc.).

Table 1: *Summary statistics (France). Source:* Enquête Emploi *(1990-2002).*

Variable	Mean	s.d.	Min	Max	Number of obs
INDIVIDUAL CHARACTERISTICS:					
School cohort	1953.2	(5.34)	1944	1962	155,562
% female	0.51	(0.50)	0	1	155,562
Age	42.7	(6.45)	28	58	155,562
% in employment (ILO definition)	0.80	(0.40)	0	1	155,562
% self-employed (ILO definition)	0.13	(0.34)	0	1	124,189
Log of hourly wage (in 2005 euros)	2.12	(0.41)	0.00	7.27	98,597
SCHOOLING:					
Schooled after the 1967 Berthoin reform	0.54	(0.50)	0	1	155,562
Age left school	16.69	(1.93)	6	25	155,562
CREDENTIALS:					
No qualification	0.21	(0.40)			155,562
Primary schooling certificate	0.16	(0.37)			155,562
Low vocational qualification	0.26	(0.44)			155,562
Junior secondary schooling certificate	0.10	(0.30)			155,562
Intermediate vocational qualification	0.11	(0.32)			155,562
Senior secondary schooling certificate or above	0.16	(0.37)			155,562

Notes: The sample is constructed from the 13 *Enquête Emploi* surveys that were conducted by the French Office for National Statistics (Insee) each year between 1990 and 2002. It includes all respondents who were born in France, belong to the 1944 to 1962 school cohorts, left full-time education between the ages of 6 and 25 and were aged 25-60 when surveyed. Individuals born between January 1 of year t and December 31 of the same year are assigned to school cohort t. Each individual is only included once in the sample, as we keep only the information collected during the first of the three successive years during which a respondent is interviewed. Details on the grouping of academic and vocational credentials are given in section 4.

Table 2: *Summary statistics (England & Wales). Source: Quarterly Labour Force Survey (1993-2006).*

Variable	Mean	s.d.	Min	Max	Number of obs
INDIVIDUAL CHARACTERISTICS:					
School cohort	1958.4	(5.51)	1949	1967	247,440
% female	0.52	(0.50)	0	1	247,440
Age	40.65	(6.65)	26	58	247,440
% in employment (ILO definition)	0.81	(0.40)	0	1	247,440
% self-employed (ILO definition)	0.13	(0.34)	0	1	237,935
Log of hourly wage (in 2005 pounds)	2.07	(0.52)	0.85	7.11	199,241
SCHOOLING:					
Schooled after the 1972 ROSLA	0.57	(0.50)	0	1	247,440
Age left school	16.30	(1.17)	6	25	247,440
CREDENTIALS:					
No qualification	0.23	(0.42)			247,440
Low vocational qualification	0.06	(0.23)			247,440
Junior secondary schooling certificate	0.30	(0.46)			247,440
Intermediate vocational qualification	0.19	(0.39)			247,440
Senior secondary schooling certificate or above	0.22	(0.41)			247,440

Notes: The sample is constructed from the 53 Quarterly Labour Force Surveys that were conducted by the UK Office of National Statistics between Fall 1993 and Winter 2006. It includes all respondents who were born in England & Wales, belong to the 1949 to 1967 school cohorts (i.e. born between September 1, 1948 and August 31, 1967), left full-time education between the ages of 6 and 25 and were aged 25-60 when surveyed. Individuals born between September 1 of year t and August 31 of year $t+1$ are assigned to school cohort $t+1$. Each individual is only included once in the sample, as we keep only the information collected during the fifth of the five successive quarters during which a respondent is interviewed. Details on the grouping of academic and vocational credentials are given in section 4.

Table 3: *OLS, First stage, Reduced form and 2SLS effects of the 1967 Berthoin reform in France and the 1972 ROSLA in England & Wales on school leaving age, hourly wages, employment and self-employment. Sources:* Enquête Emploi *(1990-2002) and Quarterly Labour Force Survey (1993-2006).*

	OLS	First stage		Reduced form				2SLS	
Dependent variable:	Log of hourly wage	School leaving age		Log of hourly wage		Employed	Self-employed	Log of hourly wage	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Panel A: France – School cohorts 1944-1962 – Enquête Emploi (1990-2002)									
Women									
Coefficient	.087***	.272***	.274***	-.000	-.002	.001	.007	-.001	-.007
(s.e.)	(.001)	(.035)	(.035)	(.003)	.004)	(.013)	(.005)	(.013)	(.013)
Number of observations	47,670	47,670	47,670	47,670	47,670	78,586	55,874	47,670	47,670
Men									
Coefficient	.073***	.264***	.263***	.001	-.001	-.000	.002	.005	-.004
(s.e.)	(.001)	(.025)	(.025)	(.008)	(.008)	(.005)	(.004)	(.031)	(.029)
Number of observations	50,927	50,927	50,927	50,927	50,927	76,976	68,315	50,927	50,927
Panel B: England & Wales – School cohorts 1949-1967 – Quarterly Labour Force Survey (1993-2006)									
Women									
Coefficient	.119***	.309***	.310***	.019***	.021***	.007**	.005	.060***	.067***
(s.e.)	(.001)	(.023)	(.022)	(.003)	(.003)	(.003)	(.004)	(.011)	(.011)
Number of observations	75,304	75,304	75,304	75,304	75,304	129,813	96,321	75,304	75,304
Men									
Coefficient	.095***	.267***	.267***	.019**	.019**	.002	.001	.069***	.069***
(s.e.)	(.001)	(.028)	(.028)	(.008)	(.008)	(.004)	(.004)	(.028)	(.029)
Number of observations	68,535	68,535	68,535	68,535	68,535	117,627	102,920	68,535	68,535
Age controls	Quartic	None	Quartic	None	Quartic	Quartic	Quartic	None	Quartic

Notes : *: significant at the 10% level; **: significant at the 5% level; ***: significant at the 1% level. Each coefficient comes from a separate regression. The dependent variables are log of hourly wage (columns 1, 4, 5, 8 and 9), age left full-time education (columns 2 and 3), an indicator variable for being employed (column 6) and an indicator variable for being self-employed (column 7). OLS regressions in column 1 include a quartic in age and a full-set of survey year dummies. RDD regressions in columns 2-8 include a fourth order polynomial in school cohort allowing for an intercept shift at the cutoff point (i.e. 1953 in France and 1958 in England & Wales). The French sample (panel A) is constructed from the first wave of the 1990 to 2002 French *Enquête Emploi* and includes individuals who were born in France, belong to school cohorts 1944 to 1962, left full-time education between the ages of 6 and 25 and were aged 25-60 when surveyed. The English and Welsh sample (panel B) is constructed from the fifth wave of the 1993 to 2006 Quarterly Labour Force Survey and includes individuals who were born in England & Wales, belong to school cohorts 1949 to 1967 (i.e. born between September 1, 1948 and August 31, 1967), left full-time education between the ages of 6 and 25 and were aged 25-60 when surveyed. The samples used in column 6 include all individuals. The subsamples used in column 7 are restricted to workers. The subsamples used in column 1-5 and 8-9 include wage earners only. Standard errors are clustered at the school cohort level.

Table 4: *Reduced form effects of the 1967 Berthoin reform in France and the 1972 ROSLA in England & Wales on the distribution of the highest academic or vocational credential held by women and men. Sources:* Enquête Emploi *(1990-2002) and Quarterly Labour Force Survey (1993-2006).*

	France		England & Wales	
	Women	Men	Women	Men
	(1)	(2)	(3)	(4)
Highest Held Qualification:				
% No qualification	-.020	-.009	-.114***	-.057***
	(.012)	(.007)	(.007)	(.008)
% Primary schooling certificate	.008	.005	n/a	n/a
	(.006)	(.004)	n/a	n/a
% Low vocational qualification	-.006	-.012	-.016***	-.002
	(.007)	(.010)	(.001)	(.003)
% Junior secondary schooling certificate	.009*	.008**	.126***	.054***
	(.005)	(.003)	(.004)	(.011)
% Intermediate vocational qualification	.002	.004	-.003	.005
	(.005)	(.006)	(.004)	(.013)
% Senior secondary schooling certificate or above	.007	.004	.006	.001
	(.007)	(.004)	(.006)	(.008)
Number of observations	78,586	76,976	129,752	117,543

Notes: *: significant at the 10% level; **: significant at the 5% level; ***: significant at the 1% level. Each coefficient comes from a separate regression. The dependent variables are the fraction of individuals who hold a particular academic or vocational credential as their highest qualification and the regressions include a fourth order polynomial in school cohort allowing for an intercept shift at the cutoff point (school cohorts 1953 in France and 1958 in England & Wales). The French sample is constructed from the 1990 to 2002 French *Enquête Emploi* and includes individuals who were born in France, belong to school cohorts 1944 to 1962, left full-time education between the ages of 6 and 25 and were aged 25-60 when surveyed. The English and Welsh sample is constructed from the 1993 to 2006 Quarterly Labour Force Surveys and includes individuals who were born in England & Wales, belong to school cohorts 1949 to 1967 (i.e. born between September 1, 1948 and August 31, 1967), left full-time education between the ages of 6 and 25 and were aged 25-60 when surveyed. Standard errors are clustered at the school cohort level.

Table 5: *Reduced form effect of the 1972 ROSLA on the literacy scores of British women and men. Source: 1994 International Adult Literacy Survey (IALS).*

	Women	Men
	(1)	(2)
Literacy Score:		
Prose literacy	12.52*	13.83*
	(6.81)	(7.95)
Document literacy	13.60**	13.93
	(4.95)	(10.45)
Quantitative literacy	13.71**	17.83*
	(5.53)	(10.56)
Number of observations	830	662

Notes: *: significant at the 10% level; **: significant at the 5% level; ***: significant at the 1% level. Each coefficient comes from a separate regression. The dependent variables are the scores obtained by British participants to the 1994 International Adult Literacy Survey on each of the three literacy scales (prose, document and quantitative literacy), separately for women and men who were born in the UK between 1949 and 1967 and left school between the ages of 6 and 25. Scores are in the range 0-500. The regressions include a fourth order polynomial in school cohort allowing for an intercept shift at the cutoff point (school cohorts 1953 in France and 1958 in England & Wales). Individuals born in 1957 are excluded from the sample because some of them (born between January 1 and August 31) faced a minimum school leaving age of 15 while the others (born between September 1 and December 31) faced a minimum school leaving age of 16. Regressions include a fourth order polynomial in birth cohort allowing for an intercept shift at the cutoff point. Standard errors are clustered at the year of birth level.

Table 6: *Educational career of French school cohorts 1952 and 1953 between the age of 6-7 and the age of 18-19. Source:* Tableaux de l'Éducation Nationale, Ministère de l'Éducation nationale *(1969, 1978).*

SCHOOL COHORT 1952	← **Compulsory schooling** →												
School Year	1958-59	1959-60	1960-61	1961-62	1962-63	1963-64	1964-65	1965-66	1966-67	1967-68	1968-69	1969-70	1970-71
Age	6-7	7-8	8-9	9-10	10-11	11-12	12-13	13-14	14-15	15-16	16-17	17-18	18-19
Not enrolled/dropout	2.1%	3.9%	3.4%	2.9%	0.9%	1.0%	0.0%	1.5%	25.6%	37.5%	43.1%	58.1%	75.0%
Preschools and special education	4.7%	0.7%	1.0%	1.3%	1.6%	1.7%	1.8%	1.7%	0.9%	0.5%	1.1%	0.3%	
Primary schools	93.2%	95.4%	95.5%	95.6%	91.3%	58.7%	22.3%	7.9%	1.4%				
Extended primary schools				0.2%	0.8%	7.9%	26.7%	38.8%	9.9%	1.1%	0.2%		
Junior secondary schools					5.4%	30.6%	49.2%	50.1%	45.6%	29.4%	9.9%	1.1%	0.1%
Vocational training									14.1%	18.8%	21.9%	14.1%	5.2%
Senior secondary schools									2.5%	12.7%	24.3%	26.4%	19.7%

SCHOOL COHORT 1953	← **Compulsory schooling** →												
School Year	1959-60	1960-61	1961-62	1962-63	1963-64	1964-65	1965-66	1966-67	1967-68	1968-69	1969-70	1970-71	1971-72
Age	6-7	7-8	8-9	9-10	10-11	11-12	12-13	13-14	14-15	15-16	16-17	17-18	18-19
Not enrolled/dropout	3.7%	3.3%	2.3%	0.5%	0.7%	0.6%	0.0%	1.1%	14.8%	23.9%	38.1%	55.5%	73.5%
Preschools and special education	2.5%	0.7%	1.1%	1.5%	1.8%	2.0%	2.0%	1.9%	1.3%	1.3%	0.8%	0.3%	
Primary schools	93.8%	96.0%	96.6%	98.0%	91.1%	58.3%	21.0%	7.6%	1.8%				
Extended primary schools				0.1%	0.7%	6.4%	22.6%	33.6%	11.4%	2.2%	0.2%		
Junior secondary schools					5.7%	32.7%	54.4%	55.8%	52.3%	36.6%	10.5%	1.0%	0.1%
Vocational training									15.7%	22.3%	23.9%	15.2%	5.5%
Senior secondary schools									2.7%	13.7%	26.5%	28.1%	20.9%

Notes : This table is constructed from the retrospective statistics released in 1969 and 1978 by the French Ministry of Education. It follows the educational careers of the two school cohorts that surround the implementation of the Berthoin reform in 1967. Pupils born in 1952 faced a minimum school leaving age of 14 whereas those born in 1953 faced a minimum school leaving age of 16. Total enrollment is broken up by school type (lines) and school year (columns) between the age of 6-7 and the age of 18-19. School levels include preschooling and special education (pupils with special needs), primary schooling (five years from the age of 6-7 to the age of 10-11), extended primary schooling (*Classes de fin d'études*, three years from the age of 11-12 to the age of 13-14), junior secondary schooling (*Collège*, four years from the age of 11-12 to the age of 14-15), vocational training (*Lycée professionnel*, four years from the age of 14-15 to the age of 17-18 or three years from the age of 15-16 to the age of 17-18) and senior secondary schooling (*Lycée général et technologique*, three years from the age of 15-16 to the age of 17-18).