Vive la Révolution!
Long term returns of 1968 to the angry students*

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Abstract

The famous events of May 1968, starting with student riots, threw France into a state of turmoil. The period of 'revolution' coincided with the time in which important examinations are undertaken. As a result, normal examination procedures were abandoned and the pass-rate for various qualifications increased enormously in that one year. These events were particularly important for students at an early (and highly selective) phase of higher education. They are shown to have pursued further years of education because thresholds were lowered at critical stages (i.e. at entry to university and in the early years of university).

These historic events provide a natural experiment to analyse the returns to years of higher education for the affected generation and to consider consequences for their children. Thus, we can contribute to the debate on two very controversial questions in the economics of education: What is the true causal relationship between educational attainment and its labour market value? Is there a causal relationship between the education of parents and that of their children? Much of the existing literature considers the effect of interventions altering an individual's years of compulsory schooling on the margin rather than an intervention which occurs at a later stage. Our results are based on the latter and show a high rate of return to an additional year of education.

The treatment group considered here is on the margin of the higher education system. This study suggests that expanding the university system to accommodate such people can yield very high private returns. There is also evidence of a strong causal relationship between obtaining an additional year of higher education and the educational outcomes of children. Hence our study suggests very positive effects of the '1968 events' for affected cohorts and is of contemporary relevance given the current debate in many countries about widening access to higher education.

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Introduction

In May 1968 a conflict between students and the university authorities in Paris precipitated a series of events that would lead to mass student protests, the biggest national strike in the history of France and the dissolution of Parliament. By mid-June, the revolution was over and normal life resumed. However there were important consequences for those taking examinations in that year. In the university context, exams became a central aspect of the bargaining process between students and the authorities, with the former successfully bargaining for 'light-touch' exams 'to avoid harming students who have spent a lot of time struggling for a better university'.[1] There are numerous examples of delays and modifications to university examinations taking place in that one year. The important examination taken for the *baccalauréat* (success at which guarantees access to university), only involved oral tests in that year. As a result the pass rate for various qualifications increased enormously. We show that the lowering of thresholds at an early (and highly selective stage) of the higher education system enabled a significant proportion of students born between 1947 and 1950 (particularly in 1948 and 1949) to pursue more years of higher education that would otherwise have been possible. This was followed by a significant increase in their subsequent wages and occupational attainment, which was particularly evident for persons coming from a middle-class family background. Finally, returns were transmitted to the next generation on account of the relationship between parental education and that of their children.

These findings are of interest in their own right. The 1968 events are shown to have had a positive and enduring impact on the affected cohorts – and their children.

[1] For example, see the debates reported in the French newspaper *Le Monde*, 17 May 1968

Although much has been written about the 1968 events, the literature has not picked up on this fact. Furthermore the '1968 events' are an important historical example of a situation where thresholds were lowered at critical stages of the higher education system, enabling those on the margin either to enter university or pursue further years of higher education. This is of strong contemporary relevance in many countries, where there is a big debate about widening access to higher education. The experience of 1968 suggests that enabling the 'marginal' person to enter and persist in higher education can result in high private returns in the labour market and, because this is also transmitted to the next generation, may have a long run effect on the stock of human capital in the economy.

One can use these '1968 events' as a 'natural experiment' to identify the causal effect of education, and hence contribute to important debates within the economics of education literature: What is the true causal impact of education on labour market outcomes such as the wage? Is there a causal relationship between the education of parents and that of their children? As is well known, it is normally very difficult to identify the causal effects of education on subsequent outcomes. Highly educated individuals are likely to have many other specific attributes that are unobserved or difficult to measure. It is unclear whether their good labour market outcomes and high achieving children should be interpreted as a consequence of their education or as an effect of these unobserved attributes. One approach to overcome this problem is to find an instrument for educational attainment which is uncorrelated with an individual's other attributes.[2] Many studies have used this Instrumental Variable approach to estimate the returns to education, especially following Angrist and Krueger's (1991) landmark study, which used an individual's quarter of birth as

[2] See Card (2001) for a review of the methodological issues in this literature and many of the important studies.

an instrument for schooling. In the present context, we are able to use birth cohort as an instrument for education. We also make use of the fact that within the most affected cohorts (1948 and 1949), there was a disproportionate effect on the education of those coming from a middle-class family background. Hence, we use the 1968 events as a quasi-experimental setting to analyse the effects of an exogenous shock to educational attainment on the subsequent outcomes of the affected generation and their children.

This approach has similarities to several other studies that use a 'natural experiment' affecting particular birth cohorts as an Instrumental Variable strategy to estimate the labour market returns to schooling (e.g. Aakvik *et al.*, 2003; Harmon and Walker, 1995; Lemieux and Card, 1998; Meghir and Palme, 2004, Oreopoulos, 2003) and to identify the relationship between parental education and that of their children (Black *et al*, 2003; Chevalier, 2004; and Oreopoulos *et al*., 2003). Thus, the first part of the paper relates to the private return to education, whereas the second part is relevant to the debate on whether education of the current generation can affect the intergenerational transmission of human capital. Little is known about the second issue in particular, where there is far less empirical evidence. To our knowledge, this is the first study which estimates the returns to education of parents on their own outcomes and those of their children in the same framework.

Unlike all other papers in this literature, the intervention is 'one-off', unexpected and temporary – it has no consequences for cohorts coming after the 1968 events; the incentive structure of the education system is unchanged.[3] Furthermore, the focus of our paper is on the returns to higher education. In contrast, most other papers that use a similar identification strategy (cited above) focus on interventions

[3] It has been shown that reforms to the compulsory school-leaving age can affect the educational decisions of those not directly affected by the rule change (e.g. Lang and Kropp, 1986).

which affect years of compulsory schooling.[4] In studies using changes to compulsory school-leaving laws for identification, the ‘treated group’ (and hence those for whom returns to education are identified) are generally those who leave the education system with relatively low-level qualifications. However, the effect of an increase in years of education might be quite different depending on when the individual is exposed to it. The composition of the treatment group might also differ for a shock which affects higher education. The treatment group in our paper is composed of those on the margin of the higher education system. We show large private returns to an additional year of higher education for this group – comparable to the upper bound in studies which rely on an intervention affecting the length of compulsory schooling. This suggests that the return to an additional year of higher education could potentially have an effect on an individual’s earning capacity which is at least as large as an additional year of compulsory schooling. Furthermore, there is a strong causal relationship between an additional year of higher education and the educational performance of children.

This paper has the following structure: in Section 1, our descriptive section, we provide further details of the consequences of the 1968 events for the examinations undertaken in that year; and we use our data to provide descriptive evidence of the cohorts affected. In Section 2, we develop a simple model to show how different cohorts will be affected by the relaxation of examination standards in 1968 and then we estimate the returns to education in the labour market using two identification strategies - the first using the fact that particular birth cohorts were

[4] Lemieux and Card (1998) focus on the effect of a program designed to facilitate the further education of post World War II veterans in Canada. Hence this intervention also directly affects years of post-compulsory schooling. Aakvik *et al.* (2003) use multiple reforms in Norway to estimate returns to different levels of education (including higher education). There is a literature using college proximity as an IV approach to estimate returns to years of education (Card, 1995). The main problem is that living near a college is potentially endogeneous (Card, 2001).

affected by the relaxation of examination standards in 1968; and the second using the fact that the impact was strongly linked to the family background of the affected cohorts. In Section 3, we consider the consequences for the intergenerational transmission of human capital when we use the ‘natural experiment’ to estimate the relationship between parental education and that of their children. Finally, we draw together conclusions from the analyses.

1: Descriptive section

1.1 The 1968 events and the examinations

In 1967 and 1968, a wave of student protest movements broke out across Europe, Japan and the United States. This was the era of mass protest against politics and social values, about which much has been written. However, pre-May 1968, there was no mass student movement in France (Touraine, 1971) and the events of that month took everyone by surprise. This is well encapsulated by the now famous editorial written in a French national newspaper in March of that year: 'the French are bored. They are not taking part either directly or indirectly in the great convulsions which are shaking the world….'[5] The spark that set off the dramatic sequence of events in France was the decision to close the University of Nanterre in Paris on 2 May (following months of protest between students and the authorities at that institution) and the heavy-handed response of the police to the subsequent protest at the Sorbonne. The protests quickly snowballed in university towns across the whole of France. Then workers became involved: 'it was only in France that the revolt of the students got a response from the workers…that precipitated the biggest general strike in French history, paralyzing the economy…' (Singer, 2000). Over 10 million French workers were involved in the strikes – roughly two-thirds of the French workforce. De Gaulle dissolved the national assembly and called for new parliamentary elections in June. These actions effectively ended the brief, dramatic 'revolution'.

In the midst of these events, university examinations became a central aspect of the bargaining process between the administration and the students. The former exerted considerable pressure for exams to be 'light' in that year. There was extensive

[5] *Le Monde*, 14 March 1968.

coverage in the French newspaper *Le Monde* of general controversy surrounding the examination procedures of that year and notice of the numerous delays and modifications to procedures in different institutions throughout the country. The following quotations are among those of student representatives: 'exams do not have to be abolished: it would harm students who need their diploma'; 'exams do not have to be as hard as usual: it would damage the students who have spent a lot of time struggling for a better university'.[6] The authorities seem to have given way to student demands for 'light-touch' exams in that year, so as not to deny the value of the student movement.[7]

The examination for the *baccalauréat* in 1968 provides one very interesting and important example of the consequences of these decisions. The *baccalauréat* is 'the symbolic national diploma both crowning the successful completion of secondary education and providing a passport for entry into higher education'.[8] It can be seen as the first examination within the French system of higher education. It has a long and distinguished history from its foundation by Napoleon in 1808. Normally it involves several oral and written examinations, which consists of open-ended questions, a commentary on documents etc…, and takes place over several days. Students must obtain the *baccalauréat* to advance within the system of higher education. Indeed, this qualification gives recipients an automatic right to attend university. In May 1968, it

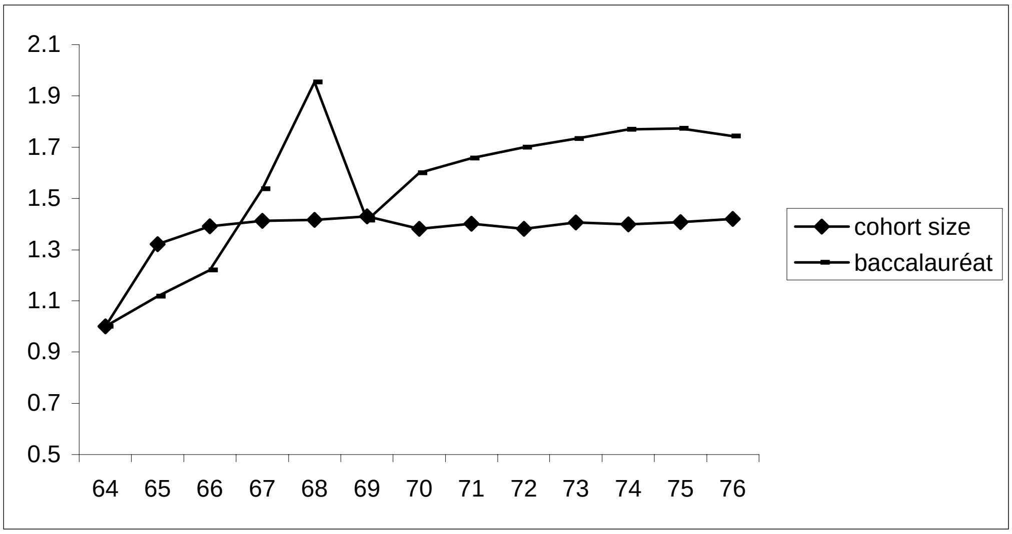
[6] Students attending the '*commission des examens*' at the Sorbonne, as reported in *Le Monde*, 17 May 1968.

[7] After the 1968 events, various university reforms were proposed (by the new Minister for Education, Edgar Faure). The aim was to allow some re-structuring of universities, give them more autonomy and ensure a greater representation of academics and students on university and departmental councils. According to Ball (1979), nothing was done administratively and financially to facilitate the application of reforms, although he notes some success with regard to the decentralisation of decision making.

[8] French Embassy in the UK, http://www.ambafrance-uk.org/asp/service.asp?SERVID=100&LNG=en&PAGID=92

was announced that the examination for the *baccalauréat* would be postponed[9] but this was superseded by a later announcement, which stipulated that there would only be oral examinations, to take place on the same day. Candidates were informed of their results on the day of taking the examination, thus implying that examiners from different places did not consult with one another before making a final decision over the grades (normally there is a national commission to harmonize marking procedures). As a result, the number of persons obtaining the *baccalauréat* in 1968 was about 30 per cent larger than in the adjacent years, whereas the size of the corresponding birth cohorts was fairly stable over this period (see Figure 1).[10]

Figure 1: Trends in the number of *bacheliers* and in cohort size.



Source: French Ministry of Education (number of *bacheliers*) and the French Statistical Office (cohort size). The size of the cohort for year t corresponds to the number of persons born at t-19 (19 is the median age of candidates). The two series are normalised to 1 in 1945.

The lax examination procedures of 1968 were also important at the most selective stages of the university system, typically for students in the first two years

[9] This decision was taken on 24 May, the day the Minister of Education, Alain Peyrrefites, resigned from his post.

[10] An examination for entry to *terminal* (i.e. the year of preparation for the *baccalauréat*) was abolished in 1966. This led to an unusually high number of candidates for the *baccalauréat* in 1967, which explains the increase in the number of *bacheliers* in that year. Other changes to the *baccalauréat* are discussed in Appendix 1, with reference to the Labour Force Survey.

(*premier cycle*). These first two years make it possible to obtain a university diploma (generally, an upper-level 'technician' diploma) and are essential in order to pursue a university degree (1 year) or a masters (2 years). We have made a systematic investigation of newspaper archives and they are full of examples showing the difficulties faced by university administrators in organising exams to be held at that time. In almost every university, it was not possible to organise regular examinations without protracted delays and adaptations reflecting the bargaining process between the students and the administration. For example, the *Brevets de Techniciens Supérieurs* (upper-level 'technician' diploma) was granted to students without any specific examination in 1968. Instead the decision to confer this qualification was made on the basis of the candidate's work throughout the two years.[11]

1.2 The 1968 events and educational credentials

The potential consequence of easier examinations is an increase in the final educational outcomes of the beneficiaries. Each year, a certain proportion of students (generally those who have been held back a grade already) do not pass examinations in the early stages of higher education and drop out of the education system, whereas they would have stayed if the examinations had been easier. The effect of a significant relaxation of examinations on any given birth cohort is important if it contains a high proportion of such students at relevant stages of education in that year. Hence we can speculate that the relaxation of the examinations in 1968 have primarily affected students born between 1947 and 1950, i.e. students who were at an early stage of higher education at this time.

[11] *Le Monde*, 25 June 1968.

In our analysis we choose 1946 and 1952 as control cohorts as they are less likely to be affected by the relaxation of examination standards in 1968.[12] Cohorts born in 1945 and earlier are less suitable as controls since cohort sizes during the period of World War II were much smaller (whereas they were stable between 1946 and 1952). Cohorts born in 1953 and after were potentially affected by the 'Berthoin' reform, which increased the compulsory school-leaving age by 2 years. However, in Appendix 5, we show that our basic findings are robust to alternative choices of control cohorts.

Educational and labour market outcomes for cohorts considered here (1946 to 1952) can be illustrated using the French Labor Force Survey (LFS) for the years 1990, 1993, 1996 and 1999. These four years were selected as the sample is rotating every three years, and we are interested to observe the same person only at one point. We do not use surveys before 1990 as they do not contain information on wages. The LFS is a large representative sample of the French population of age 15 and above. The number of respondents per cohort is about 10,000 for our pooled sample. We focus on outcomes for male workers, leaving an analysis of the impact of the 1968 events on female outcomes for a separate paper. In Appendix 1, we provide a description of the LFS with relevance to this paper, in particular in relation to educational qualifications and destinations. We could have chosen the Population Census to do a more limited analysis of the effects of 1968. The Census does not have information on wages or the social background of individuals. Furthermore, it is a lot less reliable than the LFS for measuring individual characteristics (notably education and date of birth). However, in Appendix 2, we show that a similar story is revealed

[12]In contrast, broader effects of 1968 (for example, being part of the student culture in which 'revolutionary' activities occurred) might be expected to affect the control cohorts as much as the treated cohorts.

with respect to available measures using the Census. The larger sample sizes in the Census allow much greater precision.

A detailed breakdown of the effect of 1968 is provided in Table 1, which shows the percentage of male wage earners in each birth cohort within the following categories of educational attainment: the *baccalauréat*, but no higher; any university qualification; a two-year university diploma '*bac+2*'; and at least a degree-level university qualification '*>bac+2*'.

Table 1 : The distribution of education across male workers, by year of birth

	Bac, but no higher	>Bac		
		All	*Bac+2 (diploma)*	*>Bac+2 (degree+)*
1946	**9.9**	**17.4**	***6.3***	***11.0***
	(0.4)	**(0.6)**	***(0.3)***	***(0.5)***
1947	9.0	18.8	*7.1*	*11.8*
	(0.4)	(0.6)	*(0.4)*	*(0.5)*
1948	9.1	19.0	*6.9*	*12.1*
	(0.4)	(0.6)	*(0.4)*	*(0.5)*
1949	**9.8**	**20.3**	***8.6***	***11.6***
	(0.4)	**(0.6)**	***(0.4)***	***(0.5)***
1950	9.8	18.4	*8.0*	*10.4*
	(0.4)	(0.6)	*(0.4)*	*(0.5)*
1951	9.8	17.9	*7.5*	*10.4*
	(0.4)	(0.6)	*(0.4)*	*(0.5)*
1952	**9.8**	**17.8**	***7.3***	***10.4***
	(0.4)	**(0.6)**	***(0.4)***	***(0.5)***

Source: Labor Force Survey, 1990, 1993, 1996, 1999. Male wage earners.

A number of interesting facts emerge from this table. Firstly, the percentage of workers with the *baccalauréat only* remains stable across cohorts. In particular, it is very similar in 1946, 1949 and 1952. In contrast, the percentage of workers with at least a university qualification is significantly higher for students born in 1949 than for students born in 1946 or 1952. Generally speaking, the majority of workers in the LFS with the *baccalauréat* subsequently achieve a university qualification, but the percentage is significantly higher for students born between 1947 and 1950 than for students in adjacent years. Taking the trend between 1946 and 1952 as a reference, the increase in the percentage of students with a university qualification is as follows: 1.3 in 1947; 1.5 in 1948; 2.7 in 1949; 0.8 in 1950 and 0.2 in 1951. Cumulatively, this is about 6.5 per cent of a cohort (which is consistent with the increase in the number of *bacheliers* implied by Figure 1, taking the trend between the early Sixties and the early Seventies as a reference).

Hence, there is a significant shift in the percentage of workers with at least a university diploma (*bac+2*), with a peak for the 1949 cohort (i.e. those of age 19 at the time of the 1968 events). For those with a university degree or higher (>*bac+2*), the peak is for the 1948 cohort. Taken together, the most affected cohort is 1949, i.e. persons who were 19 years of age in 1968.

In theory, persons who are of 'normal' age pass the *baccalauréat* at the age of 18. In practice, the majority repeats at least one grade when in primary or secondary school and passes the *baccalauréat* at the age of 19. This means that a significant proportion of the 1949 cohort was actually studying for the *baccalauréat* in 1968.[13] Specifically, about 10 per cent of the 1949 cohort was in this position in that year. (see Appendix 3). Given that the pass rate for the *baccalauréat* increased by about 30 per cent, it is clear that the events of 1968 enabled a significant fraction of these students to cross the threshold for the *baccalauréat*, allowing the pursuit of further years of education and the eventual attainment of a university diploma or above.[14]

[13] The survey "*Formation et la Qualification Professionnelle*" (FQP, 1993) confirms that the median age of candidates for the *baccalauréat* is 19 and shows that students studying for the *baccalauréat* in 1968 comprised about 36 per cent from the 1949 cohort; 26 per cent from 1950; 18 per cent from 1948; 7 per cent from 1947 and 8 per cent from 1951. The FQP is a small representative survey of French adults, which contains detailed, retrospective questions about the educational career of participants.

[14] A pass rate of about 30 per cent applied to 10 percent of the 1949 cohort suggests that about 3 per cent of the cohort would have been affected. This order of magnitude is consistent with that implied by column 1 of Table 1.

There is also a group of persons born in 1949 who were already at university in 1968. This smaller group, who had passed the *baccalauréat* at 18 years or younger, constitutes about 7 per cent of the cohort. Students who enter university without having to repeat a grade at school are more likely to pursue their studies in more prestigious institutions and the vast majority (over 90 per cent) are successful in obtaining a university qualification. The impact of the events on the final educational qualification obtained by this group can only have been marginal.

Thus, among those born in 1949, the only group whose educational career was substantively affected by the events of May 1968 comprised students at the end of their secondary education who were taking examinations for the *baccalauréat*, having already repeated a grade at secondary school. As shown by Table 1, the events not only enabled an important fraction (about 2.5 per cent of the cohort) to pass examinations for the *baccalauréat*, but also to pursue further studies and subsequently obtain a university qualification. Had they been born one year earlier or one year later, they would have left school without the *baccalauréat*.

Hence, in our econometric strategy we largely focus on outcomes for the 1949 cohort as compared to 'control' cohorts of 1946 and 1952, who are unlikely to have been affected by the impact of the 1968 events on examinations. This allows us to identify the average effect of the years spent in higher education for the group of students who are just below the minimum level of scholastic ability required to pass the *baccalauréat* in a normal year and who would *not* have obtained it if the minimum level had remained constant in 1968.

1.3. Within cohort effects of 1968

The cohorts born between 1947 and 1950 have been more affected by the 1968 events because they were more likely to take important examinations at this time and hence more likely to be at a highly selective stage of the education system, where the probability of failure is high. However, it is not the case that the members of these birth cohorts have been equally exposed to the events of 1968. Specifically, only marginal students have been affected – i.e. people who would not have been able to pass the examinations if they had been taken one year before (or after) the 1968 events and who would not have been able to pass if they had repeated the exam year. It is possible to better identify these marginal students by exploring the distribution of educational credentials by social background.

In the Sixties and early Seventies, most children from lower socio-economic groups left school before taking the exam for the *baccalauréat*, whereas children from higher socio-economic groups were likely to pass the *baccalauréat* and enter university.[15] Middle-class children were more likely than the former group to persist in the secondary school system but less likely than the latter group to pass the exams for entry into tertiary education.[16] For example, using the LFS, one can observe that the sons of middle-class fathers (*employés* or *artisans-commerçants*) were the most likely to pass the diploma taken at the end of junior school (*BEPC*) without subsequently passing the *baccalauréat* (see Appendix 4). Hence, it follows that the 'marginal' student is less likely to be from a lower socio-economic group (who leave school without taking the *baccalauréat*) or a higher socio-economic group (who have less difficulty in passing the exams). The 'marginal' student is more likely to be

[15] Only 10% of the sons of manual workers passed the *baccalauréat*, compared to about two-thirds of the sons of *cadres* (i.e. an upper-level white collar position).

[16] As in other countries, there is evidence that educational attainment is strongly linked to family background in France (see for example Givord and Goux, 2004).

middle-class. Table 2 provides support for this hypothesis, since the effect of the 1968 events is more pronounced for male workers from a middle class family background than for workers from other social groups. The difference in the log odds of holding a university diploma between workers from a middle-class family background and workers from any other type of family background increased by about .40 between 1946 and 1949 and then decreased by about .28 between 1949 and 1952 (see the last column of Table 2). In other words, there is a notable peak in the log odds differential for the cohorts most affected by the 1968 events, which is consistent with these events having been particularly important for students from a middle-class family background. Thus, our second econometric strategy will focus on variations in the educational gap between the sons of middle class fathers and the sons of fathers from other socio-economic groups across birth cohorts, and test whether these variations have been accompanied by similar variations in their labour market outcomes.

Table 2 : Percentage of male workers in each birth cohort with at least a university diploma, by social background.

Birth cohorts	Social Background: Middle Class (*)			Social Background: Other Social Classes			Relative Odd ratios (**)
	Upper-level middle	*Lower-level middle*	All	*Low (manual worker, farmer)*	*High (white collar)*	All	
1946	*21.8*	*15.7*	18.2	*8.8*	*47.2*	17.1	1
1947	*25.9*	*20.9*	22.9	*9.1*	*49.8*	17.4	1.30
1948	*25.6*	*22.5*	23.8	*10.1*	*44.1*	17.4	1.37
1949	***29.1***	***23.1***	**25.7**	***10.2***	***49.2***	**18.5**	**1.40**
1950	*27.4*	*17.9*	21.7	*9.6*	*45.4*	17.4	1.21
1951	*26.4*	*19.1*	22.1	*8.6*	*47.3*	16.6	1.31
1952	*22.9*	*20.2*	21.3	*9.3*	*44.9*	16.6	1.12

*Upper-level middle class corresponds to *Artisans* and *Commerçants* (i.e. shopkeepers, craftmen, self-employed…) whereas lower-level middle class corresponds to *Employés* (i.e., non-manual routine workers, clerks).

**If q_t denotes the probability of holding a university diploma for workers born in t and coming from a middle class family and p_t denotes this probability for workers from other social classes, then the relative odds under consideration can be written $[q_t(1- p_t)/ p_t(1-q_t)]/ [q_{46}(1- p_{46})/ p_{46}(1-q_{46})]$. The last column shows the changes in these relative odds across cohorts.

2: Labour market returns to education

2.1. Modelling framework

Let y_i, the earning capacity of worker *i* at the entry into the labour market, be written as follows:

$$y_i = \alpha n_i + u_i \qquad (2.1)$$

where n_i denotes the number of grades that worker *i* has passed in higher education;[17] u_i is the unobserved ability of the worker; and α represents the return to years of higher education. Let F_u denote the cumulative distribution function of u_i and assume it is constant across cohorts. If n_i were uncorrelated with u_i, the estimation of α would represent no difficulty – it would simply require comparing the wages earned by workers with different levels of education. However, a problem arises if high ability persons are more likely to pass exams which give access to higher education. In that case n_i will be correlated with u_i and the wage gap between persons with different years of higher education will also reflect their ability. The 1968 events provide us with an opportunity to overcome this problem on account of the exogenous shift in years of higher education, which enables an instrumental variable strategy.

To keep things as simple as possible, let us assume that the system of higher education consists of N basic grades and denote $c_1<c_2<\ldots<c_N$ as the levels of ability that are required to pass entry exams into each grade. Without any grade repetition, each student *i* will reach the level n_i, where $c_{n_i} \leq u_i < c_{n_i+1}$ (and we set $c_o = -\infty$ and $c_{N+1}=\infty$ as a normalisation).[18] He/she can only obtain access to the next grade by repeating a year and will only gain access to this higher grade (n_i+1) if

[17] Higher education is defined as any education above the first stage of secondary education. In France, the second stage of secondary education is at '*le Lycée*' and consists of preparation for the *baccalauréat*.

[18] For example, c_1 represents the minimum level of ability required to pass the *baccalauréat* without grade repetition.

$c_{n_i+1} \leq u_i + r < c_{n_i+2}$, where r represents the impact of grade repetition on the capacity to pass an exam. For simplicity, we assume that $c_k + r < c_{k+1}$, so that repetition only makes it possible to persist in education for one additional grade.

In this simple setting, the number of grades is a function of u_i and can be written as follows:

$$n_i = \sum_{k=0}^{N} k1(c_k \leq u_i + r < c_{k+1}) \qquad (2.2)$$

where 1(x) is a dummy indicating whether x is true or not. Now assume that the thresholds are exogenously shifted downwards at the end of a specific year T (i.e. 1968). Let us denote $c'_1 < c'_2 < \ldots < c'_N$ as the new thresholds. Also assume $r \geq c_k - c'_{k+1}$, meaning that the effect of grade repetition is at least as large as the advantage provided to students by the downward shift in the thresholds. In this case, students who are of "normal" age in year T (and who still have the possibility to repeat a grade) cannot be affected by the shift. In contrast, the educational career of students who have already been held back a grade may well be altered significantly.

Specifically, consider students who have repeated a grade and are in the last year of secondary school (i.e. *terminal*). In our case, this corresponds to the 1949 cohort. There are students in this cohort such that $c'_1 < u_i + r < c_1$ and hence would not have passed the examinations for the *baccalauréat* if the threshold had remained constant. They obtain the *baccalauréat* only on account of the 1968 events. The proportion of such students can be written as follows:

$$\Delta P = F_u(c_1 - r) - F_u(c'_1 - r)$$

When we compare the 1949 cohort either with a cohort who were far beyond the stage of taking selective examinations in 1968 (such as 1946) or with a cohort who had not yet arrived in *terminal* (such as 1951 or 1952), we find that a proportion

ΔP has attended one supplementary grade within the higher educational system. This generates an outcome $\Delta y=\alpha\Delta P$.

In this very simple setting, a regression of y on n provides an unbiased estimator ($\Delta y/\Delta P$) of the returns to higher education (α) when a dummy variable to indicate whether the individual belongs to the 1949 cohort is used as an instrumental variable. This strategy identifies the impact of the treatment on the treated group.

In the more general case, where the returns to education and the impact of grade repetition vary across individuals and grades (α_{ik} and r_{ik} for each i), this strategy provides us with an estimate of the average causal response (ACR) as defined by Angrist and Imbens (1995). The ACR is a generalization of the local average treatment effect (LATE, Imbens and Angrist, 1994), where the treatment is not binary[19]. In our case, the treatment can take on four values: below the *baccalauréat*; the *baccalauréat* but no higher; a university diploma (*bac+2*); a university degree (>*bac+3*) and there are three possible elementary changes in the intensity of the treatment: an increase from below to above the *baccalauréat* (Change 1); an increase from below to above a university diploma (Change 2); an increase from below to above a university degree (Change 3). In this context, the ACR captures a weighted average of causal responses to these elementary changes for those whose educational attainment has actually been affected by the relaxation of the exams in 1968. Specifically, the ACR can be written $p_1\alpha_1 + p_2\alpha_2 + p_3\alpha_3$ where the weight p_k (k=1,2,3) is proportional to the number of people who were induced by the 1968 examinations to complete Change k and where α_k (k=1,2,3) represents the average impact of

[19] We must also assume that the treatment is a monotone function of the instrument. In our case, this amounts to an assumption that people born in 1949 complete at least as much schooling as they would have completed had they been born in 1946 or 1952. This identifying assumption is not verifiable, but has the testable implication that the empirical CDF of the educational attainment of people born in 1949 does not cross (and actually dominates) the CDF of people born in 1946 or 1952. Table 1 shows that this is true.

Change k for those induced to complete this change by the events. As shown by Angrist and Imbens (1995), the ACR weights p_k can be very simply estimated from the difference between the empirical CDF of the educational attainment of people born in the treated cohort (1949) and the non-treated cohorts (1946 and 1952). This difference is actually provided by Table 1. It shows that Change 3 affected relatively few people in the 1949 cohort (about .6% of the cohort), whereas the events led to Change 1 or Change 2 for a larger proportion of the cohort (about 2.5% of the cohort in both cases). Hence the ACR weights are as follows: p_1=.44; p_2=.46; p_3=.10.

2.2. Estimation strategy

Following on from the simple model described above, we write the wages w_i earned by worker i from cohort c_i at age a_i as follows:

$$w_i = \alpha n_i + \omega a_i + \delta c_i + u_i \qquad (2.3)$$

Holding age constant, the cohort trend is equivalent to a time trend and captures the effect of a general increase in productivity on wages. Building on the previous subsection (eq. 2.2), we assume that years of higher education (n_i) vary in a non-linear way across cohorts:

$$n_i = d_{47}C_{i47} + d_{48}C_{i48} + d_{49}C_{i49} + d_{50}C_{i50} + d_{51}C_{51} + e\tau_i + v_i \qquad (2.4)$$

where the C_{it} variables represent cohort dummies, τ_i a cohort trend and v_i a random variable which captures the influence of unobserved ability on years of education. The d coefficients reflect the effect of the 1968 events on the minimum levels of ability that are required to pass the exams. This effect varies across the different cohorts depending on the stage they were at in the education system in 1968. The potential correlation between v_i and unobserved factors affecting wages (as reflected in u_i) is the reason why estimation of equation (2.3) may be misleading without the benefit of the natural experiment, which is provided by the 1968 events.

In our empirical strategy, we first estimate the reduced form of the earnings equation by substituting equation (2.4) into (2.3). If our understanding of the 1968 events is correct (and if α is positive), dummy variables for the 1948 and 1949 cohorts should have a strong positive effect in addition to the cohort trend. Secondly, we restrict our sample to the most affected cohort (1949) and two cohorts that appear to be unaffected (1946 and 1952) and we identify the return to education using C_{i49} as an instrumental variable.

This approach assumes that C_{i49} does not affect wages other than through an individual's educational attainment. It might be argued that some students benefited directly from the experience of being involved in the student protests. If the benefits were different for students of age 19 (born in 1949) than for older or younger students (i.e. born in 1946 or 1952), the events might have generated differences in labour market outcomes across cohorts. However, as mentioned above, there was no mass student movement pre-May 1968.[20] Apart from those involved in student militant groups[21], the involvement of the more general student body was concentrated within a short time period (May-June 1968).

A further potential problem with our strategy is the implicit assumption that employers have not been able to discriminate against individuals who have benefited from easier exams at a point in their educational career.[22] This might be a reasonable

[20] Touraine (1971) notes weaknesses in the student union bodies in France. He states that it was the repression and especially the foul play at the Sorbonne on 3 May that provoked the student reaction which rapidly became a mass movement. It must be emphasised that the events of 1968 were wholly unexpected. Capdeville and Mouriaux (1986) state 'the events of May/June 1968 were all the harder to grasp because they had not been foreseen, and did not seem predictable'.

[21] As a matter of fact, the main leaders of the students' movement (Daniel Cohn-Bendit, Jacques Sauvageot, Alain Geismar) were all born before 1946.

[22] There is also a possibility that an increase in the supply of educated workers in the 1949 cohort depressed skill prices and hence depressed the return to education for this cohort relative to other

assumption in the present context as most people affected by the 1968 events subsequently went on to achieve a higher level of education. If employers place more weight on the highest level of education attained than on earlier qualifications, then one would not expect to observe a penalty for these workers. Furthermore, employers would need to have good information about the educational consequences of 1968.

To account for the possibility that there is a direct effect of birth cohort on wages, we develop a second identification strategy which does not rely on cohort effects for identification. Given that the effects of the 1968 events are expected to vary by family background for the reasons discussed above, our second identification strategy involves using the interaction between birth in the affected cohort and coming from a middle-class family background. Hence, we augment equations (2.3) and (2.4) by including a cohort dummy for year of birth and a dummy for family background. In equation (2.3), we also include an interaction term between the dummy for 'middle class' and whether the worker was born in the affected cohort. This interaction term enables identification while also allowing for cohort effects to exert a direct influence on wages.

2.3. Regression results

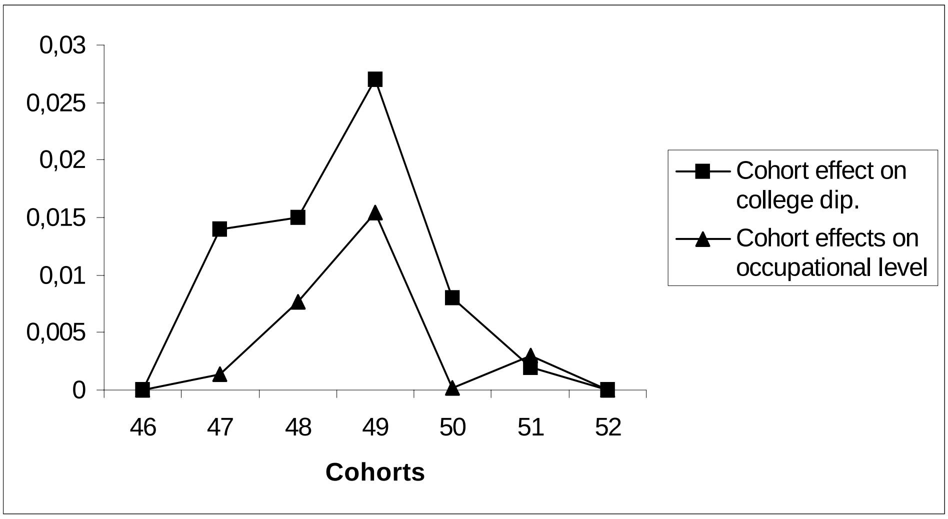
The essence of our results can be simply illustrated. Figure 4 shows the impact of birth cohort on the probability of obtaining a university diploma and the probability of holding an upper-level white collar position (i.e. *cadre*)[23] in OLS regressions where we control for a cohort trend and an age trend (eq. 2.4). The birth cohorts are

cohorts. In this case, estimates will capture the general equilibrium effect to the extent that depends on the elasticity of substitution across cohorts. If workers from different cohorts are highly substitutable (which might be expected, since we are observing workers at an advanced stage of their career), then estimates of the wage return to education will only capture this effect to a small extent (if at all).

[23] About 15% of the population of male workers in our sample hold these positions. Their wages are on average twice as high as the wages of other workers. In the French context, being a *"cadre"* may be interpreted as a proxy for permanent income.

observed in LFS surveys between 1990 and 1999. Hence, the occupation of survey participants reflects their position when in their forties. The pattern in the data shows a marked similarity for the two different effects and provides good *prima facie* evidence of the relationship between educational and labour market outcomes, with a pronounced upward shift for persons born in 1949.

Figure 4: The net effects of birth cohort on the probability of holding at least a university diploma and on the probability of holding an upper-level white-collar position (*cadre*)



A more comprehensive summary of the impact of birth cohort on the education and labour market outcomes of workers is provided in Table 3. Columns (1) to (3) show the probability of obtaining various educational qualifications for each birth cohort (relative to a reference group born in 1946 or 1952), controlling for a cohort trend and the age of the worker. The educational qualifications are: *baccalauréat* only; at least a university diploma (*bac+2* or more); at least a university degree (*bac+3* or more). We have converted qualifications to the equivalent number of years in higher education[24] and this is the dependent variable in column 4. Then in

[24] We have constructed the variable "number of years in higher education" with a value of 3 for workers with *bac* only (the three years of upper-secondary education), 5=3+2 for workers with *bac+2* and 7=5+2 for workers with more than *bac+2*.

columns 5 and 6, we report these regressions where the dependent variables are labour market outcomes – log wages and the probability of holding an upper-level white collar position (i.e. '*cadre*').

Table 3: The impact of birth cohort on the education and labour market outcomes of male workers

	Bac only	At least university diploma (*bac+2* or more)	At least university degree (*bac+3*)	Number of years: higher education*	Wages	*Cadre* (Upper White collar occupations)
1947	-.009	.014	.008	.06	.006	.001
	(.006)	(.008)	(.006)	(.05)	(.010)	(.008)
1948	**.007**	**.015**	**.012**	**.08**	**.031**	**.008**
	(.006)	**(.008)**	**(.006)**	**(.05)**	**(.010)**	**(.008)**
1949	**-.001**	**.027**	**.009**	**.15**	**.021**	**.016**
	(.006)	**(.008)**	**(.006)**	**(.05)**	**(.010)**	**(.008)**
1950	-.001	.008	-.002	.03	.005	.000
	(.006)	(.008)	(.006)	(.05)	(.010)	(.008)
1951	-.005	.002	-.001	.01	.003	.003
	(.006)	(.008)	(.006)	(.05)	(.010)	(.008)
Trend	-.000	.001	-.001	.005	.010	-.005
	(.001)	(.008)	(.001)	(.010)	(.002)	(.001)
Age	-.000	.001	.0002	.004	.023	.003
	(.001)	(.008)	(.0006)	(.005)	(.001)	(.001)
N	26370	26370	26370	26370	26370	26370

Source: LFS, 1990, 1993, 1996 and 1999.
Sample: Male wage earners born between 1946 and 1952

The results show a significant positive impact of belonging to cohorts particularly affected by the 1968 events (1949 and 1948) in terms of the probability of obtaining a university qualification or higher years of education and in terms of labour market outcomes. There is a wage premium of 2-3 percent from having been born in 1948 or 1949. There is also a higher probability of obtaining a high status occupational position (*cadre*) for those born in 1949.

In Table 4, we show results from estimating the earnings equation and the probability of being in a high level occupation (*cadre*) using OLS (columns 1 and 3) and using birth in 1949 as an instrumental variable for years of higher education (columns 2 and 4). In this case, only cohorts born in 1949 and the two control years (1946 and 1952) are included in the analysis.[25] The results indicate that each additional year spent in higher education increases wages by about 14 percentage points and the probability of holding a high-status position by 10 percentage points.[26] In Appendix 5, we show estimates obtained using different possible control groups. The results are very similar.

In the earnings equation, the point estimate for years of higher education is higher in the IV regression than in the OLS regression, indicating a possible downward bias in the former case (although they are not statistically different). A higher estimate in the IV specification is a fairly typical finding in the literature estimating the returns to education (Card, 2001). For example, it may be due to errors in the measurement of education or reflect the fact that the return to education is higher for the 'marginal' student affected by 1968 than for the average student. When looking at the probability of holding a high status position (*cadre*), point estimates from the two approaches are very close**.** The estimate is almost exactly the same (and estimated with greater precision) if the same approach is applied to Census data (see Appendix 2).

[25] Results are similar when the sample is restricted to cohorts born in 1946, 1948 and 1952, where birth in 1948 is used as the instrumental variable. Estimates are larger, though less well determined.

[26] The results are qualitatively similar when we measure education by whether the worker holds at least a university diploma.

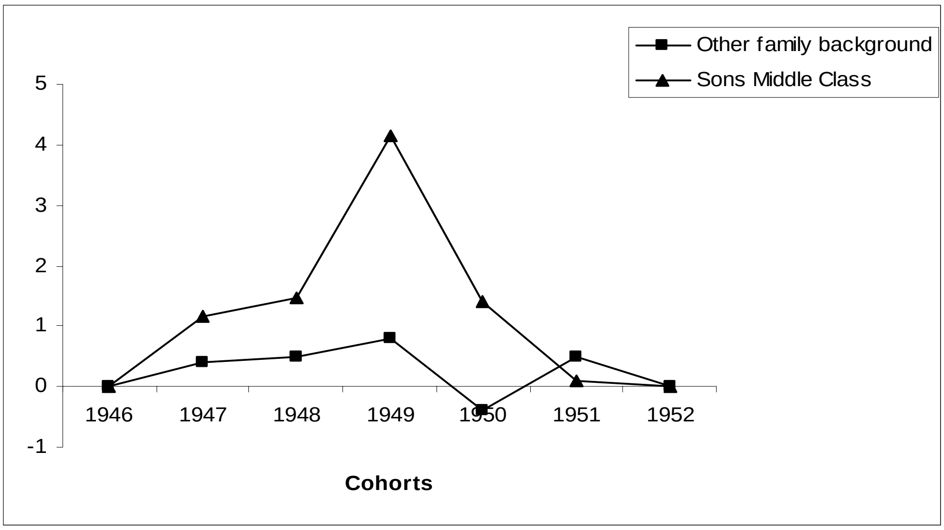
Table 4: An evaluation of the return to education using 1949 as an instrumental variable

	Wages		*Probability of upper white collar occupation (cadre)*	
	OLS	IV	OLS	IV
Years of higher education	.09 (.002)	.14 (.06)	.10 (.001)	.10 (.04)
Cohort trend	.010 (.002)	.010 (.02)	-.006 (.001)	-.006 (.001)
Age	.023 (.001)	.023 (.002)	.003 (.001)	.003 (.001)
N	11171	11171	11171	11171
R-squared	.25	.04	.36	.01

Source: LFS, 1990, 1993, 1996 and 1999.
Sample: Male wage earners born in 1946, 1949 or 1952.

As discussed above, there may be a concern that cohort dummies should enter directly into the wage equation, if the 1968 events had a direct effect on wages (as well as through years of higher education) that varies across the treatment and control cohorts. As shown above, students born in 1949 from a middle-class family background were considerably more affected by the consequences of the 1968 events on examinations than students born in the same year but from another social background. Given this fact, a simple test for our basic identifying assumption is to examine whether this is reflected in labour market outcomes. This is very simply illustrated in Figure 5 (OLS regressions where we control for birth cohort, a cohort trend and an age trend - eq. 2.4). This shows that the peak for the 1949 cohort in terms of occupational status is much stronger for workers coming from a middle-class family background than for workers from other social groups.

Figure 5: The net effects of birth cohort on the probability of holding an upper-level white-collar position (*cadre*) by family background



Hence, we implement our second instrumental variable strategy, which uses the interaction between birth in 1949 and being from a middle-class family background to identify the return to an additional year of higher education, and thus enables the birth cohort dummy to enter directly into the wage equation. Regression results are shown in table 5. The first column shows the first stage regression, where years of higher education are regressed against age, birth cohort, social background and interactions between cohort and coming from a middle-class family background. The interaction term between birth in 1949 and coming from a middle-class family background is shown to have a positive and significant effect. Columns (2) and (3) show reduced form regressions, where the dependent variables are, respectively, log wages and holding an upper-level white-collar occupation (*cadre*). Again, positive and significant effects are shown for having been born in 1949 and coming from a middle-class family background. In columns (4) and (5), we present IV estimates for the two labour market outcomes using the interaction between 1949 and middle-class family background as an instrumental variable. The results are remarkably similar to those using the first identification strategy.

In this approach, an additional year of education raises wages by 17 percentage points and the probability of holding an upper-level white-collar position (*cadre*) by 10 percentage points. Our estimates of the return to an additional year of education are close to the upper bound reported in studies which use reforms to compulsory schooling as the source of identification. This suggests that the private return to an additional year of higher education may be at least as large as the return to an additional year of compulsory schooling.

Table 5 : The effect of years of higher education on wages and the probability of holding an upper level white collar position (*cadre*).

	First stage	Reduced form		IV	IV
	Years of higher education	Log wage	Upper-level white-collar occupation *cadre*	Log wage	Upper-level white-collar occupation *cadre*
Years of higher education	**-**	**-**	**-**	**.17**	**.10**
				(.07)	**(.05)**
Cohorts					
Middle*1949	**.34**	**.060**	**.034**	**--**	**--**
	(.12)	**(.024)**	**(.020)**		
1949	.12	.045	-.005	.02	-.017
	(.06)	(.013)	(.010)	(.02)	(.013
Middle*1952	.27	.004	.007	-.04	-.019
	(.12)	(.025)	(.020)	(.02)	(.015
1952	.04	.074	-.025	.07	-.029
	(.07)	(.015)	(.012)	(.02)	(.010
Age	.013	.024	.004	.022	.003
	(.006)	(.001)	(.001)	(.002)	.001
Social background					
High (upper level)	3.42	.53	.46	-.07	.12
	(.13)	(.03)	(.02)	(.26)	(.17
High (lower level)	1.68	.35	.21	.06	.04
	(.12)	(.03)	(.02)	(.13)	(.09
Middle (upper)	.72	.22	.11	.09	.04
	(.15)	(.03)	(.02)	(.07)	(.05)
Middle (lower)	.36	.19	.05	.12	.01
	(.14)	(.03)	.02	(.05)	(.03)
Farmer	-.03	.03	-.04	.04	-.03
	(.12)	(.02)	(.02)	(.03	(.02)
Manual worker	-.31	.05	-.05	.10	-.02
	(.11)	(.02)	(.02)	(.03	(.02
N	11171	11171	11171	11171	11171

Source: LFS, 1990, 1993, 1996 and 1999.
Sample: Male wage earners born in 1946, 1949 or 1952.
Note: the excluded category for socio-economic background is ‘unknown’

The fact that our estimates do not differ radically across specifications gives us some confidence that estimates are robust and credible. The cohorts affected by 1968 did indeed benefit from the events of that year. As with other papers in this literature, we cannot say whether employers are rewarding the human capital gains from additional years of higher education or the signalling value of these years[27] (and the educational credential). However, we can say that enabling the 'marginal' university entrant to gain years of higher education led to high private gains at this time[28] – as reflected in their wages and occupational standing. On this basis, one might hypothesise that a similar group could be affected in this way by policies designed to widen access to higher education (to the extent that this involves some lowering of thresholds and hence making the system less selective). The next question is whether this gain in education and earnings for the affected cohorts had any impact on the outcomes of their children.

[27] However, Chevalier *et al.* (2004) show evidence for the UK that strongly supports the human capital interpretation of returns to education rather than the signalling hypothesis.

[28] Assuming a discount rate of 5% and a return of 3% for each year of labour market experience, a 15% effect on wages corresponds to a gain in permanent income which exceeds private costs (i.e. foregone earnings) by a factor of about 7 (i.e., .15/(.05-.03)) and which exceeds the social costs of one year of education by an even larger factor (in France, the annual per student cost of higher education is about $8,000. This is lower than annual earnings for a worker earning the minimum wage).

3: Intergenerational transmission of human capital

An important policy issue is whether providing additional education to parents (and resources more generally) transmits to the next generation, thus breaking the cycle of intergenerational immobility that occurs in many countries (e.g. see for example, Solon, 2002; Blanden *et al.* 2002). An important problem is the difficulty in distinguishing causal effects of parental education on the outcomes of children from a correlation which arises from unobserved factors that transmit across generations (e.g environmental or genetic factors). There have been a few recent papers which consider an exogenous policy change to the education system in an attempt to identify the causal impact of parental years of education on children's outcomes (Black *et al*, 2003; Chevalier, 2004; and Oreopoulos *et al*., 2003). All these studies use an extension to compulsory schooling as the basis for identification. They compare outcomes for the children of parents who were born just after the extension with those who were born just before. Black *et al*. (2003) do not find evidence of strong causal effects[29] but suggest that a policy which increased enrolment in higher education might be transmitted more successfully. We have an opportunity to consider this issue by comparing outcomes for children of cohorts affected by the impact of 1968 on the examinations (1948 or 1949) with outcomes for children of the adjacent unaffected cohorts (1946 or 1952).

We consider the sample of 15 year-old adolescents who are observed in the Labour Force Surveys between 1990 and 2001. We have information on their date of birth, their grade and the date of birth, education and occupation of their father. Our primary measure of school performance corresponds to children's 'educational

[29] However, in a UK context, Chevalier (2004) finds evidence for a causal effect of maternal education, though not for paternal education. Oreopoulous *et al.* (2003) find significant positive effects for both parents, which double when instrumented.

advancement' at the age of 15. It describes the difference between the actual grade of the adolescent and the normal grade for his/her age group (i.e. the 9th grade). This takes a value of 1 when the adolescent is one year ahead (10th grade); 0 when he/she is of 'normal' age; -1 when he/she is one year behind (8th grade); and –2 when he/she is two years behind (7th grade). The distribution for children in this sample is as follows: 54.7% are at 'normal' age; 3.1% are one year ahead; 29% are one year behind; and 13.2% are two years behind. We will also present results when using a dummy variable to indicate whether the 15 year-old has been held back for at least two grades at school.[30]

Similarly to Oreopoulos *et al.* (2003), the educational outcomes we consider are related to grade repetition. Grade repetition is a widespread phenomenon in many countries and is correlated with other measures of educational achievement. The Program for International Student Assessment (PISA) conducted by the OECD shows that 15 year-old French adolescents who have repeated a grade are likely to obtain much lower scores in maths, reading or science than those who progress to this point without grade repetition. In terms of standard deviations, the difference is equivalent to about 1.14, 1.26 and 1.17 in maths, reading and science respectively (Murat and Rocher, 2003). In the French context, this outcome can be considered as an appropriate measure of school performance.

Using these data and measures, a simple model of the relationship between parental and child outcomes can be described as follows:

$$s_i = \alpha f_i + \lambda c_i^p + \delta c_i + \beta g_i + u_i \qquad (3.1)$$

where the school performance of child i (denoted s_i) is potentially affected by his/her parental resources (or education in particular) f_i, the birth cohort of his father c_i^p,

[30] The results are qualitatively similar when we use a dummy indicating that the child is one year ahead or a dummy indicating that he/she is at least one year behind.

his/her gender g_i and birth cohort, c_i. Holding the birth cohort of the child constant, the birth cohort of the parent may be correlated with outcomes because those who have children at a later stage may have a greater capacity to invest in the human capital of their children. The cohort trend captures a possible linear trend in school performance.

As described in section 2.2, parental resources vary in a non-linear way across cohorts on account of the 1968 events:

$$f_i = d_{47}C^p_{i47} + d_{48}C^p_{i48} + d_{49}C^p_{i49} + d_{50}C^p_{i50} + d_{51}C^p_{i51} + \theta c^p_i + v_i \qquad (3.2)$$

where the C^p variables represent cohort dummies and v_i is a random variable which represents the influence of unobserved factors. The *d* coefficients reflect the impact of the 1968 events on parental resources. In our empirical strategy, we first substitute equation (3.2) into (3.1) to estimate the reduced form. If α is positive, we expect the dummy variables for the 1948 and 1949 cohorts to have significant positive effects on children's outcomes. Secondly, we implement an IV approach where parental resources in (3.1) are instrumented by cohort dummies, measuring whether the father was born in 1948 or 1949 respectively. In this case, our sample is restricted to those born in an affected year (1948 or 1949) and the two 'control' years (1946 and 1952).

Table 6 shows the (first stage) equation (3.2) where the measures of parental resources are the father's years of higher education (column 1) and whether the father is in an upper-level white-collar occupation, "*cadre*" (column 2). In columns 3 and 4, we estimate the reduced form version of equation (3.1), where the cohort of the father is used to explain children's educational advancement.

Table 6: The relationship between father's cohort of birth, his resources and child performance at school

	Father's resources		Children's performance	
	Father's years of higher education	Father=upper level white collar '*cadre*'	(Actual grade-normal grade)	Being two grades behind
Dummies for father's cohort				
1947	.17	.00	-.006	-.006
	(.10)	(.02)	(.030)	(.013)
1948	**.29**	**.043**	**.097**	**-.043**
	(.09)	**(.014)**	**(.027)**	**(.011)**
1949	**.22**	**.022**	**.070**	**-.031**
	(.08)	**(.013)**	**(.025)**	**(.011)**
1950	.12	.016	.034	-.01
	(.08)	(.013)	(.024)	(.01)
1951	.09	.010	.052	-.01
	(.08)	(.013)	(.025)	(.01)
Father's cohort trend	-.06	-.013	-.003	.000
	(.02)	(.002)	(.005)	(.002)
Child's date of birth	.07	.007	.026	-.011
	(.01)	(.001)	(.002)	(.001)
Male	.01	.01	-.17	.056
	(.05)	(.01)	(.01)	(.006)

Source: LFS, 1990-2001; Adolescents of age 15, fathers born between 1946 and 1952.

The results show a strong positive relationship between birth in an affected cohort and father's resources – whether this is measured by his years of higher education or his occupational status.[31] As before, the strongest effects are found for the cohorts born in 1948 or 1949. In columns 3 and 4, we show that this translates into better school performance for their children. The continuous measure of educational advancement (in column 3) suggests that children of the affected cohorts are more likely to be on target or ahead in terms of where they were in the school system at age 15 in comparison with the two cohorts which have been unaffected by the 1968 events (1946 and 1952). The effects are strong for the two most affected cohorts, 1948 and 1949. In column 4, results show that children of fathers born in these years are less likely to repeat two grades by 4.3 percentage points and 3.1 percentage points

[31] When we use the IV method applied in Section 1 to estimate the causal impact of years of higher education on occupational outcomes for the sample of fathers, we obtain similar results.

respectively.[32] Hence, it would appear that the benefits of the 1968 events were transmitted to the children of the affected students.

In table 7, we report regressions (estimating equation 3.1) where cohort dummies denoting whether the father was born in 1948 or 1949 are used as instrumental variables for the father's years of higher education. The dependent variable is educational advancement, as measured by the difference between the child's actual grade at age 15 with the 'normal' grade for this age group. We report the results from two IV strategies, where birth in 1948 is used as the instrument in the first approach (column 4) and birth in 1949 is used in the second approach (column 5). The IV regressions consist of those cohorts born in 1946, 1952 and either 1948 or 1949, depending on the instrument used.

In column 1, we report results from the first stage regression, where the father's years of higher education are regressed against birth cohort dummies, a trend for his own birth cohort and that of his child and the gender of his child. As before, there is a strong positive relationship between birth in a cohort affected by the 1968 events (1948 and 1949) relative to birth in the comparison cohorts (1946 and 1952). In column 2, we show the reduced form regression where the child's educational advancement at age 15 is the dependent variable (similarly to table 6). Then in column 3, we report the OLS regression, before showing regressions using the two IV strategies in columns 4 and 5 respectively.

[32] The average probability of being at least 2 grades behind at age 15 is about 11 per cent for the children of fathers from cohorts 1946-52.

Table 7: An evaluation of the impact of father's education on children's performance at school.

	(1) First Stage Father's education	(2) Reduced -Form Actual –norm. grade	(3) OLS Actual – norm. grade	(4) IV (Z=1948) Actual –norm. grade	(5) IV (Z=1949) Actual –norm. grade
Father's education	-	-	.076 (.004)	.33 (.12)	.32 (.15)
Father's cohort=1948	.29 (.09)	.097 (.027)	-	-	-
Father's cohort=1949	.22 (.09)	.071 (.025)	-	-	-
Father's cohort trend	-.06 (.02)	-.004 (.005)	.004 (.005)	.01 (.01)	.01 (.01)
Child's cohort Trend	.004 (.005)	.030 (.003)	-.023 (.003)	.006 (.012)	-.001 (.014)
Gender of child: Male	.004 (.005)	-.16 (.02)	-.16 (.02)	-.16 (.03)	-.20 (.03)
N	5,087	5,087	5,087	3,710	3,804

Source: LFS 1990-2001
Note: Regressions include all children whose fathers were born in 1946, 1948, 1949 and 1952 (first stage, reduced form and OLS regressions in columns 1 to 3); 1946, 1948 and 1952 (column 4); and 1946, 1949 and 1952 (column 5).

The regression results suggest a strong causal relationship between fathers' educational resources and the school performance of their children. Most interestingly, the evaluation of this effect is almost exactly the same whether we use 1949 or 1948 as an instrumental variable.[33] Results are qualitatively similar when we use other measures of educational advancement, such as the probability of being two years behind at school. Also, results are qualitatively similar to Oreopoulos *et al.* (2003) in that larger effects are estimated in the IV approach.

It is likely that father's education is not the only family resource whose distribution across children has been affected by the 1968 events. Since similarly educated men and women are more likely to marry each other, the distribution of mother's education has plausibly been affected, and our IV estimate should be

[33] The results are similar when we use the two instruments together. We have also checked that results are robust to the inclusion of a set of dummies indicating the age of the father.

interpreted in this light. Table 8 (column 1) confirms that children of fathers born in 1948 or 1949 are more likely to have highly educated mothers than children born in the control cohorts of 1946 and 1952. Given this fact, the estimated impact of the father's years of higher education reflects both the direct effect of his resources and the indirect effect of his wife's resources on the educational performance of children.[34]

When years of higher education of mothers and fathers are entered into the regression separately, OLS estimates suggest a similar order of magnitude for the two effects (.042 and .06 for fathers and mothers respectively – column 3). These estimates will have a similar endogeneity bias if the unobserved attributes of children are similarly correlated with the education levels of both parents. Under this hypothesis, the true causal effect of paternal and maternal education is close in value and it can be identified using the average level of education of mothers and fathers as a measure of family resources.[35] Columns 2, 4 and 5 re-estimate the effect of family resources using this new measure of parental education. The IV estimate is close in value to that estimated when father's education is used (Table 7). However, in this case the gap between the OLS and IV estimates is reduced, although the latter is still significantly higher. The IV estimate indicates that an additional year of higher education by both parents leads to an increase of about .38 standard deviations in the

[34] If α_f denotes the true effect of fathers' education and α_m the true effect of mothers' education, our IV estimates actually correspond to $\alpha_f + \alpha_m\gamma/\pi$, where γ and π are the effects of the instrument on mothers' and fathers' education respectively.

[35] If $\alpha_f = \alpha_m = \alpha$, then α can be defined as the IV regression coefficient of children's outcomes on the sum of the fathers' and mothers' education. One might like to relax the $\alpha_f = \alpha_m$ hypothesis and use the birth cohort of the mother as an instrument for identifying the specific effect of maternal education. Given the father's cohort, one actually observes an increase in education for mothers born between 1948 and 1950 relative to the control cohorts. However, the mother's cohort simultaneously affects her own education and that of her husband (exactly the same way as father's birth cohort simultaneously affects his own education and that of his wife). Hence, this instrument can not be used to separate the two effects.

child's educational advancement at age 15, whereas the OLS estimate indicates an increase of only .13 standard deviations.

Table 8: The relationship between father's birth cohort, parents' average education and child performance at school.

	Parental resources		Children's performance		
	Mother's years of higher education	Parents' average years of education	(Actual grade-normal grade) OLS	(Actual grade-normal grade) OLS	(Actual grade-normal grade) IV
Parents' average years of higher education	(-)	(-)	(-)	.10 (.01)	.29 (.08)
Father's years of higher education			.042 (.005)		
Mother's years of higher education			.060 (.005)		
Dummies for father's cohort					
1948	.35 (.08)	.30 (.07)	(-)	(-)	(-)
1949	.27 (.07)	.26 (.07)	(-)	(-)	(-)
N	5,087	5,087	5,087	5,087	5,087

Source: LFS, 1990-2001; Adolescents of age 15.
Sample: Children whose fathers were born in 1946, 1948-1949 and 1952.
Other controls included in the regressions: father's cohort trend, time trend, 6 social background dummies and a gender dummy

4: Conclusion

The famous short-lived 'revolution' of 1968 had consequences for those undertaking important examinations in that year. The protracted delays, modifications to exam procedures and deliberate intention that any exams taking place should be 'light touch', led to higher pass rates than would otherwise have been the case. While for some students, this would have made no difference or only modified the timing of their progression through the education system, there is a sizeable group for which these events made a real difference to the years of higher education obtained. This group consists of birth cohorts where many would have been in very selective stages of the higher education system at the time of the 1968 events and corresponds to 1948 and 1949 in particular. The importance of 1968 for the latter cohort is mainly due to the effect of the 1968 events on the examination for the *baccalauréat*. We also show that the 'marginal' student is more likely to be from a middle-class family background, which may be explained by the fact that students from lower socio-economic groups are much less likely to progress to this level of education and students from higher socio-economic groups are more likely to pass the examinations in any given year. The consequence of 1968 for the affected group was that it became easier to progress to a further stage of higher education, and thus obtain more years of higher education than would otherwise have been the case.

The regression results show a very strong relationship between birth in an affected cohort (especially 1948 and 1949) and labour market returns, whether these are measured by wages or the probability of holding a high-status occupation. For fathers, these effects are transmitted to the next generation since children are shown to be much less likely to repeat a grade at school.

We use these exogenous cohort-specific shifts in years of higher education to implement an instrumental variable strategy in which we estimate the labour market return to years of higher education and analyse whether these additional years have any causal impact on the educational outcomes of the next generation. The regression results show large private returns to a year of higher education. The returns are similar for both outcome measures (earnings; probability of holding an upper-level white collar occupation) and for the OLS and IV approaches. The estimated return is higher than in many other studies in this literature and suggests that the return to an additional year of higher education may be at least as large as an additional year of compulsory schooling (around which most 'experimental' studies in this literature are based). This large return is transmitted to the next generation. The 'natural experiment' of 1968 suggests a strong causal relationship between the education of fathers and that of their children.

The treatment group in this study is of much policy interest since it consists of people on the margin of the education system, who would also have been affected by any deliberate attempt to widen access to higher education at this time. Such people are shown to benefit greatly from the reduction of selectivity in the university system. This is manifest in their labour market returns. Furthermore, the strong positive effects of higher education are transmitted to the next generation.

We contribute to two important debates in the economics of education: What is the causal impact of an additional year of higher education? What is the true relationship between the education of parents and that of their children? We show the importance of an additional year of higher education for those on the margins of the higher education system. Thus, our results are of contemporary relevance. We also show the success of the revolutionaries of 1968, in ways that they were unlikely to have foreseen.

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Appendix 1

The Labour Force Survey and Educational Qualifications

The annual Labour Force Survey (*l'Enquête Emploi*) is the largest household survey conducted by the French Statistical Office, *INSEE* (with a rate of coverage of 1/300; 150,000 respondents). It is the survey used to measure the rate of unemployment. Also, it provides a measure of the distribution of educational attainment across the French population which is used as a reference by *INSEE* when conducting smaller surveys. These measures are the subject of particular care. The surveyor poses a series of three questions to determine: (a) the highest qualification obtained in '*l'enseignement secondaire général*' (general secondary education); (b) the highest qualification obtained in '*l'enseignement secondaire technique*' (vocational secondary); (c) the highest qualification obtained in '*l'enseignement supérieur general ou technique*' (tertiary education). There are additional questions about the speciality of a person's education.

With regard to the *baccalauréat*, there are different routes that can be followed (i.e. in terms of subjects chosen). The most common route is '*baccalauréat général*' (i.e academic). However, there are also vocational options – '*technical baccalauréat*' and equivalent technical diplomas. Table A1-1 shows the percentage of male workers in cohorts 1946-1952 who obtain a *baccalauréat général* and the percentage who obtain one of the technical diplomas which make it possible to pursue higher education.[36]

Table A1-1 The percentage of *bacheliers* among male workers by cohort and type of *baccalauréat*

	Bac. Général	Technical baccalauréat and technician diploma			
		All	*Bac. de techniciens et bac. technologiques*	*Diplômes equivalent to bac. techno.and bac. techniciens*	*Brevets de l'enseignement industrie ou commercial*
1946	**17.2**	**6.9**	***2.5***	***2 .4***	***2.0***
1947	18.2	6.5	*2.8*	*1.9*	*1.8*
1948	18.2	6.7	*3.3*	*2.4*	*1.0*
1949	**20.2**	**7.0**	***3.6***	***2.7***	***0.7***
1950	18.2	7.0	*4.6*	*1.8*	*0.6*
1951	17.8	7.0	*4.7*	*2.2*	*0.2*
1952	**18.4**	**7.1**	***4.9***	***1.7***	***0.3***
All	**18.3**	**6.9**	***3.8***	***2.2***	***0.9***
Nb Obs.	26,371	26,371	*26,371*	*26,371*	*26,371*

Sample: Male wage-earners born between 1946 and 1952.

[36] The rise in *baccalauréat technologique* (taken for the first time in 1969) was accompanied by a decline in *brevets* of the same magnitude. Hence, the percentage of workers with a technical diploma is fairly constant over these years. The introduction of *baccalauréat technologique* was not followed by an increase in the probability of obtaining a university diploma for persons with a non-general *baccalauréat*.

The percentage of male workers with a *baccalauréat général* is about 18.3% for cohorts 1946-1952, with a peak of 20.2% for the 1949 cohort. It is only slightly larger than that suggested by administrative data when we divide the number of persons who obtain the *baccalauréat general* at date *t* by the size of the cohort born at date *t-19*.

The percentage of male workers with some form of technical diploma is about 6.9% and very stable across cohorts. This suggests that it has not been affected by the 1968 events. Thus the shift in the distribution of university credentials observed in the LFS is driven by a shift in the percentage of those with the *baccalauréat général*. Further analysis (not reported) confirms that the shift in the percentage of workers with the *baccalauréat général* is very strong for those from a middle class family background (+4 points) and only very modest for those from other social classes (+1 point).

It is possible to identify the type of institution from which university qualifications were obtained: (a) a qualification from technological institutes and other specialised institutes (*insitut de technologie, ecole d'infirmiers, d'instituteurs, ecole d'architecture, etc…*) (b) general university (either from *premier, deuxième* or *troisième cycle*), (c) *grandes écoles* (the elite institutes of higher education). Table A1-2 shows that the increase in the percentage of workers obtaining the *baccalauréat général* for the 1949 cohort has been followed by an almost equivalent increase in the percentage of workers graduating from vocationally-oriented and other specialist institutes.

In summary, the events have shifted a fraction of the 1949 cohort from below the *baccalauréat* to above the *baccalauréat* and the data suggests that the vast majority of these students have used the opportunity to obtain a upper-level 'technicien' diploma from one of the many specialist institutes in France.

Table A1-2 : Percentage of workers with a university diploma or above, by type of higher education institute.

	Institutes of Technology and other specialised Institutes	General University (first, second & third cycles)	Grandes Écoles
1946	8.4	7.3	.8
1947	8.7	8.5	1.0
1948	8.7	8.5	1.0
1949	**11.1**	**7.9**	**.8**
1950	9.4	7.4	1.0
1951	9.2	7.2	.7
1952	8.8	7.8	.7
All	9.2	7.8	.8
N	26371	26371	26371

Sample: Male wage earners born between 1946 and 1952

Appendix 2

A Re-evaluation using the Population Census

The Population Census is much less suitable for analysis of the 1968 events compared to the LFS as it does not have information on wages or social background. Furthermore, it is much less reliable for measuring individual characteristics (notably, education and date of birth). However, it is of interest to see if similar effects can be found for the probability of obtaining a higher-level qualification and an upper-level white collar position (*cadre*). Furthermore, given the large sample sizes, effects can be estimated with greater precision than when using the LFS.

In the most recent Census (conducted in 1999), 19.6% of men born in 1949 state that they have a higher-level qualification than the *baccalauréat*. This compares to 18.6% of those born in 1952 and 18.2% of those born in 1946. The proportion of men with a higher-level qualification has a similar profile in the Census as in the LFS for cohorts born between 1946 and 1952. However, when one compares 1949 to 1946 or 1952, the increase in education is only half the size in the Census (1.2% on average) compared to the LFS (+2.5%). The lower peak in the Census may reflect the lower quality of information collected in the Census and hence the effect of measurement error.

We have repeated this exercise for earlier Censuses, conducted in 1982 and 1990 respectively. The proportion of persons who say that they have a higher-level qualification is lower compared to the most recent Census. However, there is a similar pattern for the relevant cohorts to that observed in the 1999 Census and the LFS. For example, in the 1982 Census, the percentage of men with a higher-level qualification is 15.1% for those born in 1949, as compared to 14.3% in 1946 or 1952. Furthermore, whichever Census is used, the proportion of persons with the *baccalauréat* but no higher is fairly stable over these cohorts. In 1999, such persons constitute about 10.5% of the male population, which is close to that observed in the LFS.

While the French Census does not provide information on wages, some information is collected on occupation. Therefore, we have been able to test whether the increase in qualifications observed for the 1949 cohort translates to an increase in the proportion of male workers with an upper-level white-collar occupation (*cadre*). Using the 1999 Census, one observes an increase of .7% for men born in 1949 who have this occupational status, in comparison to the control cohorts of 1946 and 1952. Similarly to that observed for educational qualifications, this increase is less strong than in the LFS.[37] The same type of pattern is observed in the 1990 and 1982 Census. Overall, in qualitative terms, both outcome measures show a similar peak in the Censuses and the LFS. Only the magnitude of the 1949 effects differs between these data sources. It is interesting to observe that the difference between the magnitude of the peaks in the Census and the LFS is almost identical with respect to both occupational status (*cadre*) and educational attainment. As a result, the return to education – that which can be identified using the relationship between the increase in occupational status (*cadre*) and the increase in qualifications for the 1949 cohort – is very similar using both data sources. This is confirmed by the example in Table A2-1, which shows a re-evaluation of the IV impact of higher education on the probability

[37] The same exercise using the 1982 Census (i.e. earlier in the career of the relevant population) reveals an increase of about 1.5 percentage points in the number of those with a higher-status occupation (*cadre* or *techniciens*).

of holding an upper-level white collar occupation using the 1982 Census. Similarly to the LFS, men born in 1949 are more likely to have obtained a higher-level qualification or a higher occupational status (‘*cadre*’ or ‘*techniciens*’) that those born in 1946 or 1952. Using the 1949 cohort as an instrumental variable, the impact of education on the probability of obtaining a higher occupational status is 9.9%. This is almost exactly the same as that found when using the LFS (see Table 4 of the text). As with the analysis using the LFS, the result is very close to that obtained using OLS (where the estimate is 8.2%). Hence, this analysis confirms the robustness of our results.

Table A2-1 : The effect of years of higher education on the probability of being cadre : a re-evaluation using 1982 census data.

	Dependent variables						
	Education level			Occupational Status			
	Bac only	At least university	Years Higher educ.	*Cadre*	*Cadre or technician*	*Cadre*	*Cadre*
						OLS	IV
Years higher education	-	-	-	-	-	.082 (.0002)	.099 (.023)
[1949=1]	-.0007 (.0012)	.0072 (.0013)	.042 (.008)	.0042 (.0012)	.0106 (.0017)	-	-
Cohort trend	.0006 (.0002)	.00005 (.00025)	-.0013 (.0016)	-.0086 (.0002)	-.0130 (.0003)	-.0085 (.0002)	-.0085 (.0002)
N	328,916	328,916	328,916	328,916	328,916	328,916	328,916

Source: Census 1982 (1/4)
Sample: Men born in 1946, 1949 or 1952

Appendix 3

The survey *Formation et Qualification Professionelle* (FQP), conducted in 1985 by the French Statistical Office, contains information on age at entry into the higher education system as well as on the educational credentials of the respondents. It shows that about 8% of the male respondents born between 1946 and 1952 entered into the system of higher education at age 18 or younger and that about 91% of these 'normal age' male entrants obtain at least a university diploma (see columns 3 and 4). The FQP survey conducted in 1977 provides very similar results. Another FQP survey has been conducted in 1993 on a smaller sample, but with additional (retrospective) information on the educational career of respondents. It shows that the median age of students in *terminal* was 19 and that at age 19, about 10% of a cohort was in *terminal* (see columns 1 and 2). The size of the sample is about 500 male respondents per cohort for FQP 1985 and about 500 respondents per cohort (male and female) for FQP 1993.

Cohort	Median age of students in *Terminal* (year of taking exam for *baccalauréat*)	Proportion in *Terminal* at age 19	Proportion entering higher education system at age 18 or younger	Proportion obtaining university diploma if entering at age 18 or younger
1946	19	.080	.062	.945
1947	19	.111	.063	.894
1948	19	.054	.054	.972
1949	19	.105	.076	.932
1950	19	.112	.086	.863
1951	19	.094	.116	.895
1952	19	.120	.099	.918
Total	**19**	**.096**	**.081**	**.912**

Source: All respondents FQP 1993 (columns 1 and 2) and male respondents FQP 1985 (columns 3 and 4)

Appendix 4

The impact of social background on obtaining qualifications

In the below table, we report results from OLS regressions. In column 1, the dependent variable is whether a person leaves school with a junior high school diploma (*BEPC*); in column 2, the dependent variable is whether he leaves school with this diploma but without a university diploma. The base category is workers from a high socio-economic family background. The first regression shows that the probability of leaving school with the junior high school diploma increases with respect to socio-economic background. The second regression shows that the probability of leaving school with the junior high school diploma but without subsequently obtaining a university diploma is higher if the person comes from a middle-class background and lower if he comes from a lower-class background relative to the base category.

	Junior high school diploma= 1	Junior high school diploma= 1 and university diploma=0
Family background		
Middle	.002	.023
	(.007)	(.007)
Low	-.77	-.041
	(.06)	(.06)
High	Ref.	Ref.
Cohort Trend	.006	.006
	(.001)	(.001)
N	26370	26370

Sample: Male wage earners; Cohorts 1946-1952.

Appendix 5

A robustness check using different possible control groups

In the below table, we show IV estimates of the effect of years of higher education on log wages, using alternative specification for the control cohorts.

The treated cohort is 1949. Similar results are obtained to those reported in the text (Section 2.3) when we use control cohorts that are closer to the treated cohort (i.e. 1947 and 1951) or further from the treated cohort (i.e. 1945 and 1953). Also, we obtain similar results when we rely only on the difference between the treated cohort and cohorts born later (i.e. 1950-1953) or cohorts born earlier (i.e. 1944-1947). Finally, results are similar when we use more control cohorts (i.e. 1944-1947 and 1950-1953) and a quadratic specification for the cohort effects.

The IV effect of years of education: a re-evaluation using alternative specifications.

	47/49/51	45/49/53	44-47/49/50-53	44-47/49	49/50-53
Years of higher education	.14 (.06)	.16 (.06)	.13 (.06)	.16 (.11)	.18 (.08)
N	11,427	10,292	31,520	16,145	19,262

Sample: Male wage earners.
Note: The table shows the results of regressions of (log) wages on years of higher education controlling for age and a cohort trend. Results are reported for five different set of control cohorts, using the cohort dummy [1949=1] as an instrumental variable.