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The Influence of Sectoral Minimum Wages on School Enrollment and Educational Choices: Evidence From Italy in the 1960s-1980s

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The Influence of Sectoral Minimum Wages on School Enrollment and Educational Choices: Evidence From Italy in the 1960s-1980s

Andrea Ramazzotti*

Abstract
Do minimum wages influence post-compulsory school enrollment and educational choices? This paper studies the effect of sectorally-bargained minimum wages using a quasi-natural historical experiment from Italy around 1969, when labour unions obtained steep wage raises for manufacturing workers. Italy's weakly-selective educational system—whereby students choose specialist educational curricula at age fourteen—allows to separately identify the impact on enrollment from that on educational choices. Absent microdata for the period under study, I present original estimates of education and labour-market variables at the province level with annual frequency between 1962 and 1982. Exploiting exogenous spatial variation in the intensity of the minimum wage hike between provinces with an instrumental variable approach and flexible Difference-in-Differences, I find a temporary increase in early school leaving and a permanent substitution away from vocational schools preparing for manufacturing jobs. The length of the adjustment might have caused a significant long-term loss for Italy's human capital stock.

JEL Classification: J24, J31, N34.

Keywords: Wage Differentials, Education, Schooling, Economic History: Europe post-1913.

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1 Introduction

Governments are under pressure to address income inequalities and low-wage jobs through enforcing statutory minimum wages or promoting collective bargaining (OECD, 2019; European Commission, 2020). However, raising the relative wage of low-skill jobs could affect the opportunity cost of—and *ex-ante* return to—formal education, influencing the decision to stay in school and the choice between alternative curricula (Neumark and Nizalova, 2007; Long, Goldhaber, and Huntington-Klein, 2015; Altonji, Arcidiacono, and Maurel, 2016). The resulting impact on human capital accumulation and skill mismatch could reduce individuals' lifetime earnings and the economy's growth in the long run.

While an expanding literature is exploring these implications in the case of statutory minimum wages,[1] relatively little attention has been paid to other wage-setting institutions. This paper focuses on the effect of minimum wages that are agreed at the sector level through collective bargaining and affect both unionized and non-unionized workers. Sector-level collective bargaining remains the prevailing wage-setting institution in most European countries, and extension mechanisms to non-signatory parties ensure that the agreements' coverage is greater than union membership figures would suggest (ILO, 2014, pp. 41-67). In fact, collectively-bargained sectoral minima with high coverage are functionally equivalent to national statutory minimum wages (Garnero, Kampelmann, and Rycx, 2015) and are usually set at higher levels in relation to the wage distribution (Boeri, 2012).

Italy provides a relevant case study because its centralized bargaining system sets wage floors that effectively produce minimum wages at the sectoral level (Pagani and Dell'Aringa, 2005; Boeri et al., 2021). Even though limits to the enforcing mechanisms and the liberalisation of the labour market have eroded their bindingness in recent years (Lucifora, 2017; Garnero, 2018), collectively-bargained minimum wages played a prevalent role in shaping the evolution of the wage distribution through the past decades (Devicienti, Fanfani, and Maida, 2019).

The paper focuses on the period between the 1960s and the 1980s. This period provides

[1] For recent examples see Neumark and Shupe, 2019a; Lee, 2020; Smith, 2021; Alessandrini and Milla, 2021.

a quasi-natural experiment thanks to the labour unions' sudden shift in bargaining strategy, from wage moderation to egalitarianism, which dated to the autumn of 1969 and was largely precipitated by exogenous political pressure (Accornero, 1992; Franzosi, 1995). Presenting new data digitized from historical sources, I show that coordinated sectoral bargaining steeply raised the minimum industrial wage and reduced the skill premium of blue-collar workers. I hypothesize that this egalitarian wage hike affected teenagers' post-compulsory education through two distinct channels. First, by raising the wage rate for entry-level jobs, it increased the opportunity cost of staying in school for the marginal student. Second, by reducing the skill premium for blue-collar workers, it decreased the *ex-ante* returns to vocational education for manufacturing jobs—relative to other types of specialist curricula—for inframarginal students. Italy's weakly-selective educational system—whereby students that intend to stay in school after the compulsory age of fourteen choose a track and specialist curriculum—allows to disentangle these the two effects. The decision to enroll in post-compulsory school and the track chosen reveal preferences on the desired number of years in school, while the choice between curricula reveals students' expectations on *ex-ante* returns to education for each curriculum.

Methodologically, I propose an identification strategy that exploits a contemporaneous but distinct institutional reform: the abolition of nominal wage differentials between geographic areas. Until 1968, minimum wages were bargained at the national level for each industry but their nominal value varied locally according to fixed scaling coefficients. This system was abolished in March 1969 and phased out by the end of 1972, leading to the spatial equalization of nominal minimum wages within each industry. Consequently, provinces that started from lower nominal levels experienced a steeper increase in local minimum wages through the transition period. This differential growth provides a source of exogenous variation in treatment intensity.

Exploiting this reform for identification purposes, I first apply an instrumental variable approach to estimate the effect of increasing the mean contractual minimum wage on school enrolment. I find that enrolment in post-compulsory secondary school is very responsive to minimum wages established by collective agreements. Estimating the marginal effect across the whole 1962-1982 period, a 1% increase in the local mean industrial minimum wage is associated with an increase in early school leavers by 0.4%-0.6%.

Secondly, I use the spatial equalization of 1969-1972 as a natural experiment to study the dynamic response of school enrolment between provinces. I do so by treating the reduced form regression as a generalized Difference-in-Differences estimator, which can recover the average causal response to marginal increases in the mean contractual minimum wage, over time. I find that provinces which experienced a steeper minimum wage hike in 1968-1972 saw a significant increase in the number of early school leavers through 1976. By 1980, however, the effect had substantially attenuated and, in some specifications, disappeared. The length of the impact suggests that mainly cohorts that turned 14 during the wage hike were influenced by it, either because they experienced the treatment first-hand, or because later cohorts were influenced by rising disemployment effects.

Even though the response of early school leaving was largely temporary, it was economically significant. Our estimates can explain four fifths of the reduction in gross enrolment rates that is observed at the national level between the 1970s and the 1980s. The dynamic estimates show that the negative impact on school enrolment was quick but temporary, as enrolment rates reverted to the mean by the early 1980s. These results support the hypothesis that, by setting high entry-level minimum wages, egalitarian collective agreements increased the opportunity cost of schooling and influenced the decision to stay in school for marginal students.

Repeating the analysis for a subset of school tracks and curricula, I find that the egalitarian wage hike provoked a permanent shift in educational choices for inframarginal students. Enrolment in vocational schools preparing for skilled blue-collar jobs showed a negative response three times larger than the average secondary school. Moreover, enrolment remained depressed ten years after the end of the shock and showed no sign of recovery. Vocational schools for white-collar jobs, instead, showed no reaction to the minimum wage hike during the phasing out of the wage zones (1968-1972) nor in the following five years, but possibly a positive effect at the end of the period. These results support the hypothesis that the compression of wage differentials for blue-collar workers permanently reduced the perceived *ex-ante* returns to specialist education.

These effects are equally found for male and female teenagers, but the larger number of male students enrolled in technical schools for manufacturing before the wage hike imply that the aggregate loss of human capital was mostly due to males' sagging enrolment. Counterfactual estimates find that the sag in male enrolment caused a substantial loss for Italy's potential human capital stock, which explains between 25% and 44% of the current lag in educational attainment with respect to the OECD average. This finding suggests that Italians' comparatively low educational attainment is not only a consequence of the late expansion of mass education, nor just a constant feature of the educational system, but also the consequence of a contingent compression in enrolment rates that, whilst temporary, will continue to linger on Italy's growth potential until the affected generations will exit the labour force.

The paper's contribution is threefold. First, it complements the growing literature on the impact of statutory minimum wages on post-compulsory education (Neumark and Shupe, 2019a; Lee, 2020; Smith, 2021; Alessandrini and Milla, 2021) by studying the influence of collective agreements with extra coverage, a hitherto unexplored connection. Moreover, the paper distinguishes between the decision to enrol in post-compulsory education and the choice of alternative tracks and curricula, which is relevant for the many countries that allow students to choose between different school tracks (Ariga et al., 2005; Manning and J.-S. Pischke, 2006; Brunello and Checchi, 2007; Betts, 2011).

Second, the paper contributes to an expanding literature within economic history that studies the institutional determinants of educational development since the 19th century (Mitch, 2013; Mitch and Cappelli, 2019), and in Italy specifically (Cappelli and Ciccarelli, 2020). This recent stream of research has focused on the institutional legacy of former states (A'Hearn and Vecchi, 2017; Ciccarelli and Weisdorf, 2019), the expansion of primary education (see for instance Vasta and Cappelli, 2020) and, more recently, on the reform of lower secondary school in the postwar period (Cappelli, Ridolfi, and Vasta, 2021). The paper extends the analysis to upper secondary education at the time of its mass expansion and presents a breakdown of enrolment by sex and type of school. The paper establishes the relevance of the pause between the 1970s and the 1980s for Italy's human capital stock, and identifying its intermediate and root causes.

Third, the paper connects with research on education inequalities in contemporary Italy. While the current research on the Italian case focuses predominantly on educational inputs and parental background to explain unequal attainment within cohorts (Checchi, 2003; Checchi and Flabbi, 2007; D. Contini and Scagni, 2010; Ballarino, Bison, and Schadee, 2011; Ballarino, Panichella, and Triventi, 2014; Panichella and Triventi, 2014; Ballarino and Panichella, 2016; D. Contini, Di Tommaso, and Mendolia, 2017; D. Contini, Cugnata, and Scagni, 2018; Giancola and Salmieri, 2020; Ballarino, Meraviglia, and Panichella, 2021), this paper studies inequality between cohorts focusing on labour market factors that affect the demand for upper-secondary education.

The rest of the paper is organized as follows: section 2 presents the historical background, formulates the hypotheses and provides descriptive evidence; section 3 discusses the data; section 4 presents the identification strategy; section 5 provides the results of the analysis and the discussion of the counterfactual scenarios; section 6 concludes.

2 Historical background and hypotheses

2.1 The rise of contractual minimum wages after 1969

Since the early 1950s, wages in the manufacturing sector were regulated by collective agreements signed at the industry level between the most representative labour unions and employers' associations. For each sector, the agreements established minimum wage scales according to the workers' skill level, which depended on the tasks performed on the job (Traversa, 1975).[2] These minimum wages represented a major component of the workers' take-home pay, and firm-level agreements could only improve on their terms.[3]

[2]The sectoral agreements provided representative lists of the most common tasks in the industry and objective criteria to classify workers (Giugni, 1963, pp. 327-346).

[3]Other fixed components of the workers' earnings were the inflation 'bonus' (*contingenza*), which was tied to the movements of a nationally-defined price index; seniority and family bonuses which depended, respectively, on experience at the current employer and on the marital and parental status. The only components that could be bargained at the firm level were collective and individual productivity premiums (*superminimo collettivo* and *superminimo individuale*). Additionally, wage earnings could be increased

While in theory the wage floor only applied to unionised workers and/or to employees of firms that were members of signatory employers' associations, it was *de facto* applicable to non-covered workers through the courts—thus providing extra coverage via judiciary extension.[4]

During the 1950s and through most of the 1960s, labour unions followed a strategy of wage moderation (Bedani, 1995). As a consequence, contractual wages lagged behind productivity for two decades. However, union membership dwindled as growing numbers of young workers joining the labour force remained unsatisfied with the unions' strategy (Checchi and Corneo, 2000). Between the autumn of 1968 and the spring of 1969, grassroots movements organized workers outside of traditional labour unions, requesting higher wages and better working conditions. To avoid losing their capacity to represent workers, in the autumn of 1969 union leaders begrudgingly but decisively adopted a new bargaining strategy, requesting higher entry-level wages, more equal pay, and strengthening workers' rights (Lange and Vannicelli, 1982). Consequently, minimum sectoral wages outpaced productivity growth until the end of the 1970s, when macroeconomic conditions and factional differences within the workers' movement led to a new period of wage moderation (Accornero, 1992).

Due to the prevalence of sectoral collective agreements in wage determination, the average wage floor for workers performing low-skill tasks (henceforth, *low-skill workers*) represented the *de facto* minimum wage for legal employment in manufacturing. Figure 1 shows its evolution across industrial sectors from 1962 to 1982, at constant prices (sources and harmonization procedures are described in detail in Ramazzotti (2023), methodological

individually by working overtime or at piecework (Guidi et al., 1971, pp. 36-37).

[4]Article 39 of the Republican Constitution in principle dictate that collective agreements would be effective *erga omnes* provided that the signatory labour unions were organized as registered democratic organizations. However, due to the unions' opposition, the article was never regulated by legislation and, consequently, all collective agreements remain private contracts between the signatories and their members. Attempts to extend their coverage through legislation (Law 751/59 'Vigorelli') failed in 1960, due to the opposition of the Constitutional Court. Nonetheless, judicial coverage has been consistently motivated with reference to article 36 of the Constitution, which recognizes the workers' right to a fair wage—i.e. one that is proportionate to the quantity and quality of the job and in any case sufficient to ensure a free and decent livelihood for them and their households. Judicial practice identifies as 'fair' the wage level bargained between the most representative labour unions and employers' association for each sector and type of job. The courts' power to apply the contractual minimum wage in case of an inferior private agreement between the worker and the employer is justified with reference to article 2099 of the Civil Code (Martone, 2016, pp. 103-158; Lucifora, 2017; Treu, 2019; Ponterio, 2019).

appendix). The series shows that the average minimum wage in industry remained stable at comparatively low levels during the 1960s—when unions followed wage moderation—, but starting in 1969 it experienced rapid growth, which decelerated only in the early 1980s. In real terms, the average minimum wage for low-skill workers in industry grew by an annualized rate of 14.6% per year between 1968 and 1980 (ranging from 8% in constructions to 24.2% in food and beverage).[5]

The minimum wage hike was especially concentrated in two sub-periods: 1969-1972 (+40%) and 1975-1978 (+31%). Growth in the former period was entirely caused by coordinated collective bargaining at the sector level (Dell'Aringa, 1976). Growth in the latter period, instead, was also due to the reform of the wage indexation system in 1975, which provided lump some wage raises in each quarter for every percentage point increase in the reference price index (Spaventa, 1976; Modigliani and Padoa-Schioppa, 1977). To avoid confusion between the two causes and keep the focus of the paper on the role of collective agreements, the identification strategy will only exploit the wage hike of 1969-1972.

2.2 The compression of the skill premium for blue-collar workers

The strategic turn of the labour unions with respect to collective bargaining after 1969 affected not only the growth rate of the contractual wage floors, but also the wage distribution of dependent workers in general and blue-collar workers in particular. In fact, both the rapid growth of entry-level wage floors in manufacturing and the reform of the wage indexation system in 1975 were strongly egalitarian, causing the minimum wages for low-skill blue-collar workers to increase faster than for all other groups. This resulted in a compression of the wage distribution both within blue-collar workers and between them

[5]It should also be noted that, before 1969, those under 21 years of age received lower contractual minimum wages than prime age workers, according to sectoral scales. For example, the 1966 collective agreement for state-owned enterprises in the engineering sector established that the minimum wage floor for low-skill blue-collar workers between the age of 16 and 18 was only 74% of the adults' rate, and for workers under 16 it was just 52% (FIM, FIOM, and UILM, 1966, pp. 236-237). However, post-1969 agreements significantly reduced these age differentials: by 1971 the reduced wage rate for under 16 was scrapped (FIM, FIOM, and UILM, 1970, pp. 255-308) and, in the following years, all age differentials would entirely disappear from most collective agreements. As a consequence, the rise in minimum wages during the 1970s was even steeper for teenagers than for prime age workers.

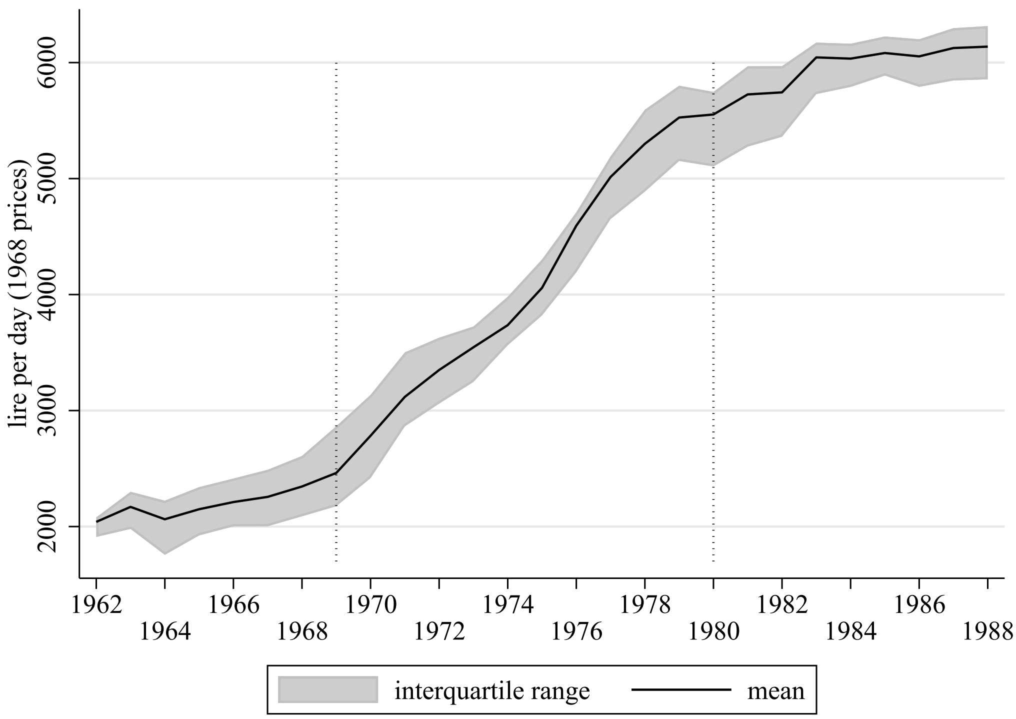


Figure 1: Minimum wage in manufacturing

Minimum wage floor for low-skill blue-collar jobs according to sectoral national collective agreements across 24 industrial sectors (manufacturing proper, mining, energy and construction) in ninety-two provinces. Sectoral minima are weighted using the estimated number of employees in each sector-province cell. The estimated number of employees is obtained as the linear interpolation from decennial industrial censuses. Details on sources and estimation strategy are provided in Ramazzotti (2023), methodological appendix). Conversion at constant 1968 prices performed using official coefficients from Istat, *Il valore della moneta in Italia dal 1861 al 2020*, available for download at `https://www.istat.it/it/archivio/258610` (last retrieved July 2022). Dotted vertical lines indicate 1969 and 1980, respectively the beginning and the end of the contractual wage hike.

and other categories (Erickson and Ichino, 1995; Manacorda, 2004).

To quantify the compression of the wage distribution within blue-collar workers I have reconstructed the skill premium for high-skill blue-collar workers using the tabulations from surveys conducted by the Ministry of Labour on a representative sample of manufacturing establishments between 1965 and 1974, and a similar source for 1984-1988. The skill premium is computed as the ratio between the average hourly effective wage of high-skill and low-skill blue-collar employees across all manufacturing sectors (excluding mining, construction and utilities). Figure 2 shows that, on average, this skill premium decreased by 56% between 1968 and 1980.[6] The rigid structure of the centralized wage-setting system

[6]The gap in the series is filled with the 75th-25th percentile ratio, which I estimate from the population of matched employer-employee administrative microdata for the region of Veneto, which has been shown

implied that changes to the wage distribution were channelled through the collective agreements. In fact, the drop of the skill premium followed closely the evolution of the contractual wage floors: Figure 2 shows that, in 1968, the ratio between the average wage floors for high-skill and low-skill blue-collar workers was 1.35, but it dropped to 1.10 by 1980.

The egalitarian turn in collective bargaining also affected the wage distribution between blue- and white-collar workers, mainly due to a reform of the wage-setting mechanism in 1972 (*inquadramento unico*) which introduced a single wage floor scale for both blue-collar and white-collar workers, effectively equalizing the entry-level wages for low-skill workers in both manual and clerical jobs (Libertini, 1974). However, the extent and significance of the compression varied between sectors and groups of workers. The average contractual wage floor for blue-collar workers outpaced the white-collar workers' minima in industry, but not with respect to the service sector (see Figure A.1). Moreover, in the industry sector between 1968 and 1984 the skill premium for hig-skill white-collar workers decreased by 29%, less than for high-skill blue-collar workers (-40% in the same period), and it remained higher in levels (1.8 in contrast to 1.35).[7] Hence, this evidence suggests that egalitarian collective agreements compressed the wage distribution for all workers, but with heterogeneous distributional effects. While nominal wages increased for all workers, high-skill blue-collar workers were relative losers, as they saw their skill premium rapidly eroding.

2.3 Implications for the demand of education

In summary, egalitarian collective bargaining after 1969 caused a steep increase in the entry-level wage for manufacturing jobs and a strong compression of the skill premium for blue-collar workers. Could this influence enrolment in post-compulsory education? To formulate some testable hypotheses, it is necessary to separately address the implications for the opportunity cost of schooling and the return to education.

to be largely representative of the national distribution for the wage earnings of blue-collar workers (Devicienti, Fanfani, and Maida, 2019).

[7] A similar evolution is described by the P75-P25 ratio. These computations are performed on the same sources detailed in the note of Figure 2 and are available upon request.

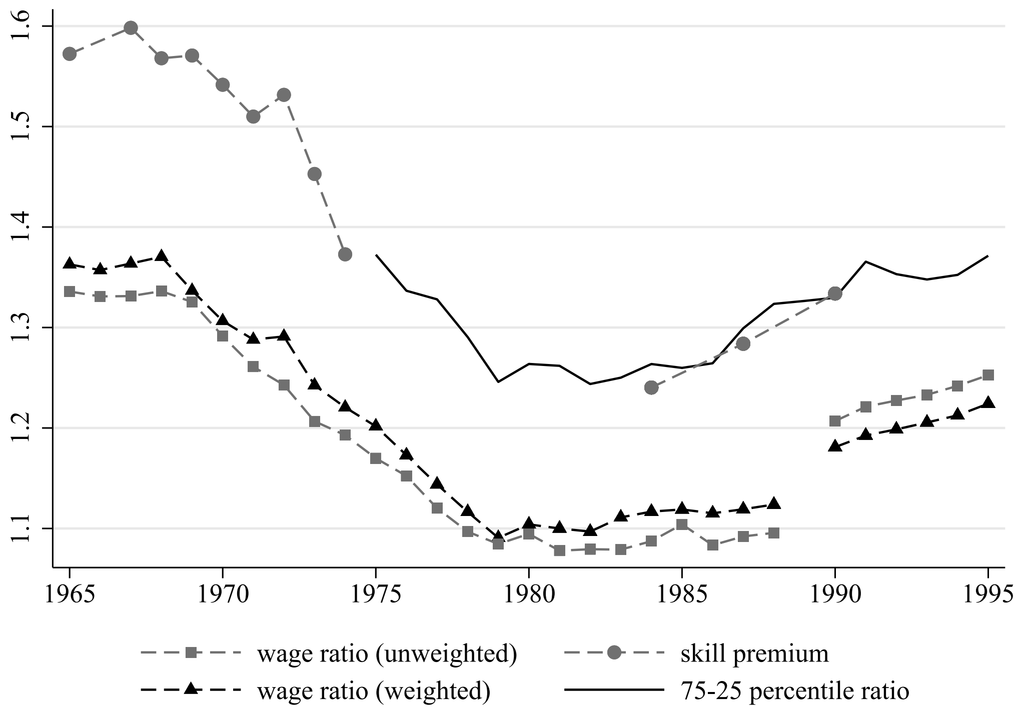


Figure 2: Wage ratio, skill premium and wage dispersion

Skill premium: ratio of the average hourly effective wage of blue-collar workers classified in the high-skill category over that of workers classified in the low-skill category in the manufacturing sector. Own elaborations on aggregate survey data from Ministero del Lavoro, *Statistiche del lavoro*, 1966-1975 and from *Rassegna di statistiche del lavoro*, several years, for 1984-1990. *Wage ratio*: minimum wage by collective agreement for most skill-intensive job class to the least for blue-collar workers in 19 industries (18 for 1990-1995). Own elaborations on contractual wage data from Istat *Statistiche industriali*, Roma, 1955-1990 and *Id.*, *Indagine sulle retribuzioni contrattuali*, Roma, 1998. Wages for 1965-1988 are weighted by the number of employees in the industries and 94 provinces, interpolated from the industrial censuses of 1961, 1971, 1981 and 1991. Wages for 1990-1995 are weighted by the number of employees in 18 sectors interpolated from the industrial censuses of 1981, 1991 and 2001. Armonized census data is extracted from Istat (2014). *Wage dispersion*: ratio of the 75th percentile to the 25th percentile from the distribution of weekly wages of blue-collar workers in the Veneto region. Weekly wages computed from Veneto Worker Histories data for thirteen industries. Weekly wages computed dividing total gross wage per employment spell by the number of weeks worked or, if unavailable, the number of days worked divided by 5.5. In case of multiple employment spells for the same worker and year, only the longest spell was used. Employment spells shorter than 16 weeks have been excluded. The dataset has been trimmed to exclude observations in the 1st and 99th percentile. The resulting sample size is 7,896,796 employment spells for 1,060,713 distinct workers from 1975 to 2000. See Ramazzotti (2023, methodological appendix) for harmonization methods.

2.3.1 Implications for post-compulsory school enrolment

In a standard framework of human capital investment (Becker, 1993), an individual's demand for education is a positive function of the expected return, which is derived by comparing the discounted value of future earnings—net of any financial and psychological costs of education—with the earnings that can be obtained on the labour market at the individual's current endowment of human capital (Checchi, 2006, pp. 18-35). Hence, *ceteris paribus*, the minimum wage hike could indirectly reduce the demand for education by raising the opportunity cost of staying in school. The immediate consequence would be an increase in the risk of dropout for the marginal student, that is an individual who—given their ability, preferences and intertemporal discount rate—was indifferent between school and work before the wage hike (Neumark and Wascher, 1995; Mohanty and Finney, 1997; Neumark and Wascher, 2003; Neumark and Shupe, 2019b; Lee, 2020; Smith, 2021). Assuming that marginal students were equally distributed across schools, we would predict a general decrease in post-compulsory educational attainment.[8]

This second-order effect, however, would be contrasted by the potential first-order effect of the minimum wage hike on teenagers' unemployment. Theoretically, a high enough minimum wage is expected to increase unemployment, even though how high the minimum wage should be set to cause disemployment effects is contingent on the degree of monopsonistic power in local labour markets and on the bite of the minimum wage (Neumark, Salas, and Wascher, 2014; Manning, 2021). If the minimum wage has strong enough disemployment effects among young people—who are typically less skilled than the average worker and thus more probable to receive earnings close to the minimum wage rate—we would expect students to stay in school rather than remain idle. In fact, some research argues that statutory minimum wages increase schooling by incentivizing teenagers to acquire more skills which would make them productive enough to be employed by firms at the higher wage rate (Mattila, 1981; Belman and Wolfson, 2014, pp. 209-217). Contrarian research notices that the effect can be heterogeneous between groups of teenagers (depending on age, sex, parental and socio-economic background, etc.)

[8] In fact, the choice of school track in the Italian educational system was strongly influenced by the student's socio-economic background, most importantly parental occupation (Panichella and Triventi, 2014). Vocational education for manufacturing jobs was the favourite option among the male children of blue-collar workers. A survey conducted by Istat on 88% of all upper secondary school graduates in the academic year 1966-1967 found that over 44% of male children of blue-collar workers graduated from this path, followed by 18% choosing technical schools for white-collar jobs, and 12% opting for technical schools for jobs in the construction sector; only 28% of male children of blue-collar workers graduated from non-vocational tracks (Istituto Centrale di Statistica, 1971, pp. 397-399). Since the probability of being a marginal student was plausibly higher for this group—surveys found a a greater sensitivity to the opportunity cost of schooling and liquidity constraints than the average secondary school student (Padoa Schioppa, 1974)—, we would expect the reduction in enrolment to be proportionally greater in vocational schools.

leading to polarisation in educational attainment and to lower completion rates of further education (Ehrenberg and Marcus, 1982; Landon, 1997; Crofton, Anderson, and Rawe, 2009). For these reasons, the net effect of the minimum wage hike on school enrolment cannot be anticipated, and we will need to explore the impact on youth unemployment during the analysis.

Historically, it is important to test for the effect on unemployment because the 1970s are considered a decade of relatively high youth unemployment in Italy (Pugliese and Rebeggiani, 2004, pp. 78-87). While unemployment rates remained at low absolute levels with respect to the following decades, in comparison to prime-age rates they surged at historical highs: the unemployment rate of 15-19 years old rose from five times the prime age rate in 1965 to twelve times in 1975, while that of 20-24 years old increased from three to seven times (see Figure 3). This period of high relative youth unemployment continued into the 1980s, at which point prime age unemployment reduced the distance to youth rates. Nonetheless, high relative youth unemployment has remained a feature of the Italian labour market with respect to comparable European economies. The formation of this dualistic labour market has been attributed, among other causes, to high employment protection for dependent workers and high entry-level wages, which would cut young people out of the formal labour market and direct them into the informal economy, which swelled in the same years (B. Contini, 1979, pp. 15-57). Support for this thesis is provided by the composition of unemployment, for the growth of youth unemployment was mainly concentrated among first job seekers (Reyneri, 1996, pp. 189-230).

Even though we do not aim to establish the root causes of the Italian dual labour market, these observations are relevant for assessing the plausibility of our interpretation with respect to school enrolment. If we find an immediate large disemployment effect of the minimum wage hike on first job seekers, it would be necessary to justify why marginal students would leave school early only to face a greater risk of unemployment and/or a longer waiting time to enter the formal labour market. If, instead, we do not find such an effect—or we find it with a delay—we could argue more convincingly in favour of the negative impact on schooling choices.

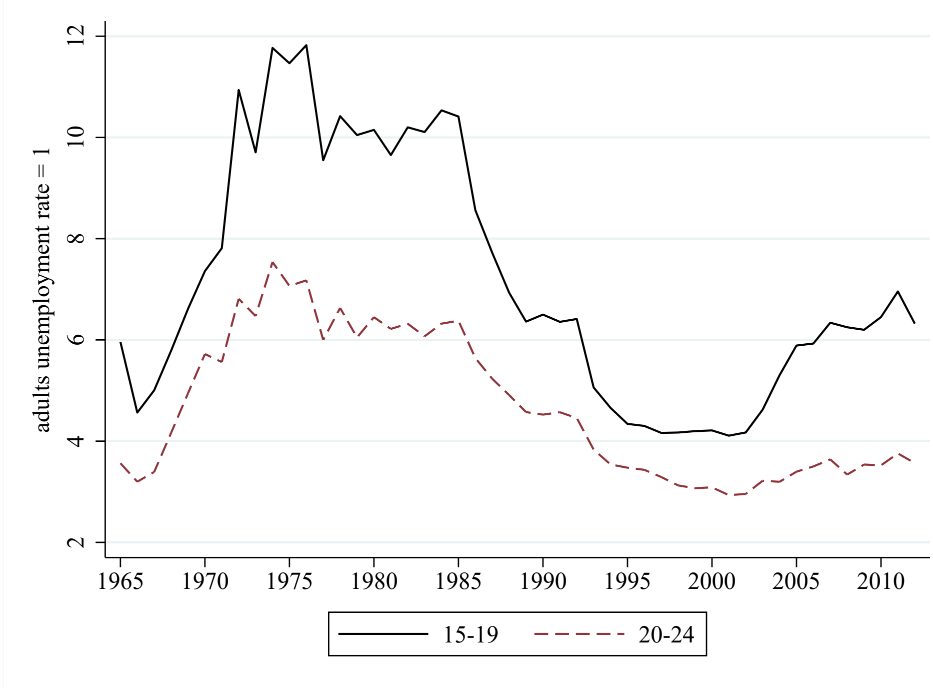


Figure 3: Relative youth unemployment rates

Relative unemployment obtained by dividing the unemployment rate for each group divided by the unemployment rate for over 25-years-old, including men and women. Unemployed include first job-seekers and people previously employed. For 1965-1969, own elaborations on Istituto Centrale di Statistica (1966, p. 52), Istituto Centrale di Statistica (1967, p. 70), Istituto Centrale di Statistica (1968, p. 72), Istituto Centrale di Statistica (1969, p. 72), and Istituto Centrale di Statistica (1970, p. 108). For 1970-2012, own elaborations on data from U.S. Bureau of labour Statistics, *International comparisons of Annual labour Force Statistics, 1970-2012*, June 7, 2013, tables by country, available for download at `https://www.bls.gov/fls/flscomparelf.htm`. From 1965 to 1968, the 15-19 group includes 14-year-olds.

2.3.2 Implications for the choice of school field

The second potential influence of egalitarian collective agreements on school enrolment acts through the compression of the skill premium in manufacturing jobs, which could affect the choice between alternative school tracks and curricula (for short, school field). Basic models of human capital accumulation treat schooling as homogeneous in providing general knowledge to students—that is, skills that can be used in any occupation. In contrast, Altonji, Blom, and Meghir (2012) present a model where schools are heterogeneous in the provision of specialist knowledge—that is, knowledge that can be directly applied only to a limited range of occupations. In this model, 'the field of education conditions occupational path' (Altonji, Blom, and Meghir, 2012, p. 186), so the choice of a school field over another depends in part on the predicted (*ex ante*) relative return to the specialist knowledge that it offers.[9] This model has been shown to explain the choice of college field, for the return to education can vary significantly between different specialisations (Long, Goldhaber, and Huntington-Klein, 2015; Altonji, Arcidiacono, and Maurel, 2016; see also Berger, 1988 for an early formulation).

The same intuition can be applied to post-compulsory education in Italy. At the time, Italy was characterized by a weak tracking system at the upper secondary level, which is schematically represented in Fig. 4. After passing a leaving exam from lower secondary school, at age 14 students could leave school or enroll in upper secondary courses. In the latter case, they had to choose between three main tracks and a range of curricula. With respect to the track, the choice was between professional schools, technical schools and general academic schools. Both professional and technical schools were vocationally-oriented, but they were differentiated because the former were focused less on theoretical contents and more on practical applications.[10] Within each track, schools offered distinct curricula, which students could not mix or modify at will. Professional and technical schools offered four main curricula, each geared towards a different economic sector: agricultural, construction, industrial, and 'business' (i.e. imparting specialist knowledge for clerical jobs).[11] Academic schools (*licei*) offered traditionally a humanistic curriculum (*classici*) but there was a significant share of students enrolled in schools offering a scientific curriculum (*licei scientifici*).[12]

Assuming that vocational schools offering an industrial curriculum were the only ones to provide specialist knowledge for manual jobs in manufacturing, the compression of the skill premium for blue-collar workers would decrease their relative expected returns. This would reduce the incentive for inframarginal students enrolling in upper secondary

[9] In the model, the individual maximizes her expected utility which depend on current consumption and on the expected value in the labour market for graduating in the chosen field, conditional on occupational random shocks and a set of factors—including beliefs about personal ability and preferences—that are influenced by previous experience and parents' genetic, cultural, and financial influence (Altonji, Blom, and Meghir, 2012, p. 187-197; see also Altonji, Arcidiacono, and Maurel, 2016, pp. 333-342).

[10] In addition, professional schools gave students the option to leave school with a professional qualification after three years (age 16) rather than staying until the completion of the usual 5-year grades. Professional schools did not give access to university courses, unlike technical schools.

[11] Other technical and professional schools offered curricula preparing for careers in the merchant navy, in hospitality, and for artistic professions (figurative arts, music, design, etc.).

[12] Other schools qualified to become teachers or offered curricula specifically geared towards female students (technical and vocational schools 'for girls', which taught skills for artisanal and secretarial jobs).

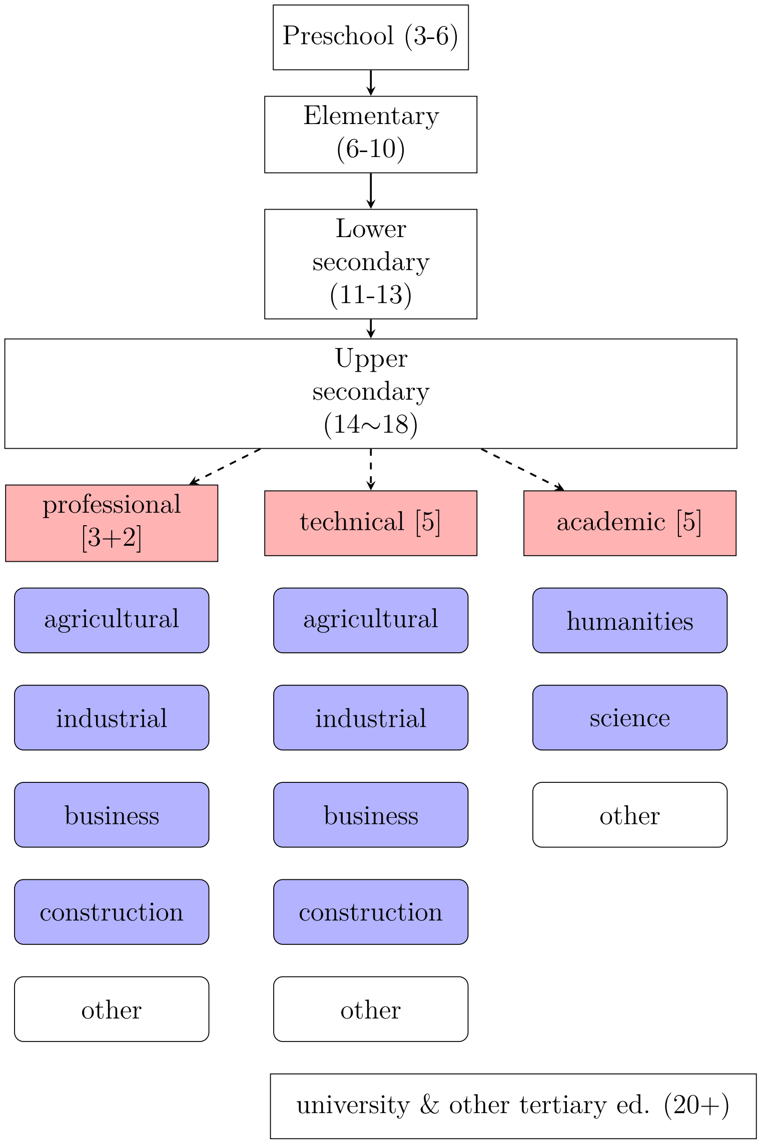


Figure 4: Tracking and curricula in the educational system, 1960s-1980s

This schematic representation highlights the three tracks facing pupils at age 14 (in red) and the main curricula within each track (in blue rectangles with rounded corners). Numbers in parentheses indicate the representative age for each level of school. Numbers in brackets indicate the minimum number of years required to obtain the leaving qualification within each track. ‘Other’ curricula include professional and technical schools for the merchant navy (*istituti professionali marinari* and *istituti tecnici nautici*), female professional and technical schools (*istituti professionali femminili* and *istituti tecnici femminili*), schools preparing teachers at different educational levels (*scuole magistrali*, *istituto magistrale*, etc.), and schools for artistic jobs (*scuole d'arte*, *istituti d'arte*, *licei artistici*; *conservatori di musica*, etc.). New curricula were added within each track over time, including professional and technical schools for hospitality (*istituti professionali alberghieri* and *istituti tecnici per il turismo*) and technical schools for low-management jobs (*istituti tecnici per periti industriali*) in the 1960s, and an academic track with a modern languages curriculum (*licei linguistici*).

education to choose vocational schools for industry. *Ceteris paribus*, we would predict a shift in the composition of enrolment in favour of other curricula and/or tracks.

The assumption that vocational schools with an industrial curriculum prepared specifically for skilled blue-collar jobs in manufacturing is supported by historical evidence. An official survey on the hiring practices of over six thousand large firms in 1960 found that 60% of employers required an upper-secondary qualification from technical or vocational schools for supervisors on production lines, and 40% of the surveyed firms applied the same requirement for high-skill manual jobs. In contrast, about 50% of firms maintained that for simpler industrial jobs a lower-secondary school diploma was sufficient, and only 32% were satisfied with a primary school qualification (Istituto Centrale di Statistica, 1964, p. 25).[13]

However, we cannot exclude that the minimum wage hike of 1969 disproportionately reduced the creation of blue-collar jobs in manufacturing (for instance, because firms could respond with labour-saving technical change). In this case, the minimum wage hike would decrease the expected returns to specialist education not through the compression of the skill premium but rather through the decrease in employment opportunities. Our estimation strategy accounts for this possibility by controlling for the share of local GDP produced in the industry sector, which proxies for the composition of the sectoral structure.

2.4 Descriptive evidence on post-compulsory education

To summarize, we have hypothesized that the egalitarian wage push could affect schooling choices through two distinct channels. First, the steep rise in the average minimum wage increased the opportunity cost of staying in school for the marginal student, raising the risk of dropping out. Second, the egalitarian compression of the skill premium for blue-collar workers reduced incentives to invest in specialist education for manual manufacturing jobs. These implications would predict post-compulsory school enrolment to decrease in the aftermath of the minimum wage hike. This decrease would be especially evident in vocational schools preparing for manufacturing jobs. In the medium term, we would

[13] The surveyed firms (covering 26% of dependent workers in manufacturing) also expected to increase by 1.5 times the number of employees with an upper-secondary school diploma from vocational manufacturing schools, at a time when overall labour demand was expected to grow only by 15%. Demand for employees with lower education was predicted to increase by just 7% (Istituto Centrale di Statistica, 1964, p. 33). It also appeared that demand outstripped supply for the more skilled roles: technical graduates with an industrial curriculum alone made over 34% of all unfulfilled vacancies in manufacturing (Istituto Centrale di Statistica, 1964, p. 48).

expect a shift in the composition of school enrolment between tracks and curricula. This section presents some descriptive evidence in favour of the hypotheses.

Figure 5 shows that enrolment and graduation rates in post-compulsory upper secondary education (age 14-18) followed a path of sustained expansion through the 1950s and the 1960s, starting from low absolute levels in the postwar period.[14] This expansion, however, decelerated in the first half of the 1970s and halted entirely in the second half of the decade, when only one in two young Italians enrolled in secondary school. enrolment and graduation rates would return to their pre-trends only in the late 1980s, finally leading to levels of educational attainment over 90% in the first half of the 2000s. The temporary pause of post-compulsory education in the 1970s is a characterizing feature of the slow expansion of secondary education in Italy with respect to other Western countries (A'Hearn and Vecchi, 2017).

What might be the proximate causes of this pause? By distinguishing enrolment rates according to the students' sex, Figure 6 shows that the pause of the 1970s was largely explained by male teenagers failing to transition from lower secondary education (age 11-13) to upper secondary education. The former had considerably expanded in the postwar period: thanks to the high demand of semi-skilled workers in the fast-growing economy and a 1962 reform that instituted a comprehensive educational system (Brunello and Checchi, 2005; Cappelli, Ridolfi, and Vasta, 2021), enrolment in lower secondary school doubled in less than a decade, reaching 76% of all children by 1965 and virtually 100% ten years later (Checchi, 1996). The expansion of lower secondary education pulled enrolment in post-compulsory secondary schools throughout the 1960s, with enrolment rates reaching 50% for males. However, starting in the early 1970s, the expansion of male enrolment in upper secondary education slowed significantly—the annual growth fell from a five-year moving average of 8% in 1951-1969 to 3% between 1970 and 1975—, and

[14]Gross enrolment rates are computed as the ratio of the total number of students enrolled in the academic year over the total population in the theoretical age group of school attendance (11-13 for lower secondary school, 14-18 for upper secondary). Academic years in Italy started at the beginning of October and finished in June in the period under consideration. For short, the academic year is defined by the year of the first term, so 1961 stands for 1961-1962. This is preferred because it gives greater relevance to the year when students chose between schools (usually, between January and September prior to starting school).

Figure 5: ENROLMENT AND GRADUATION RATES IN UPPER SECONDARY SCHOOL

Gross enrolment rate computed as the ratio between number of students and population between 14 and 18 years old. Gross graduation rate computed as the ratio between the number of high school graduates and the 18-year old population. Computations on data from Checchi, 1996.

sagged through the following ten years: in 1985, the gross enrolment rate was about the same as in 1976 (56.4% and 56.7%, respectively). Female enrolment also slowed down after the growth spurt of the early 1960s, but it did not come to a halt, which allowed it to catch up with the men's rate in 1985. The expansion of male enrolment resumed after 1985 at the same rate as the female, with enrolment rates for both sexes overcoming 80% around 1995 and 90% in the early 2000s: a thirty-year delay with respect to lower secondary education.

Why did post-compulsory school enrolment evolve differently by sex after 1969? One possible explanation has to do with sorting between school tracks and curricula. Figures 7a and 7b show the evolution of enrolment rates for women and men, respectively, distinguishing between the top-six types of schools, by track and curriculum, and a residual

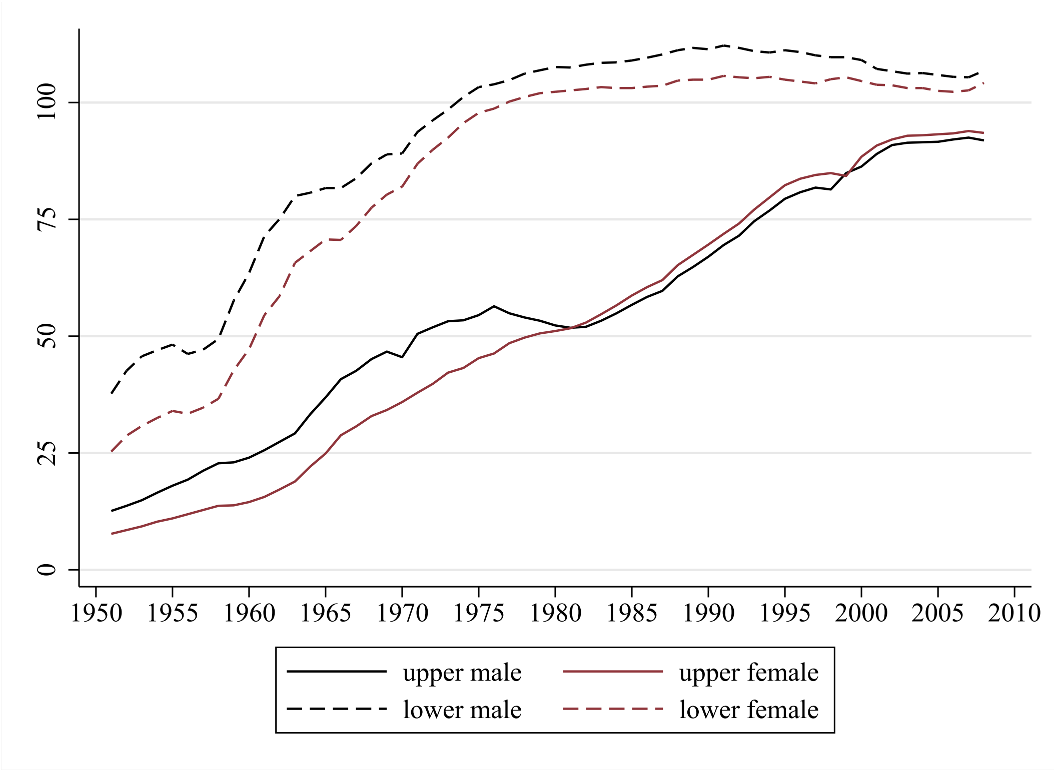
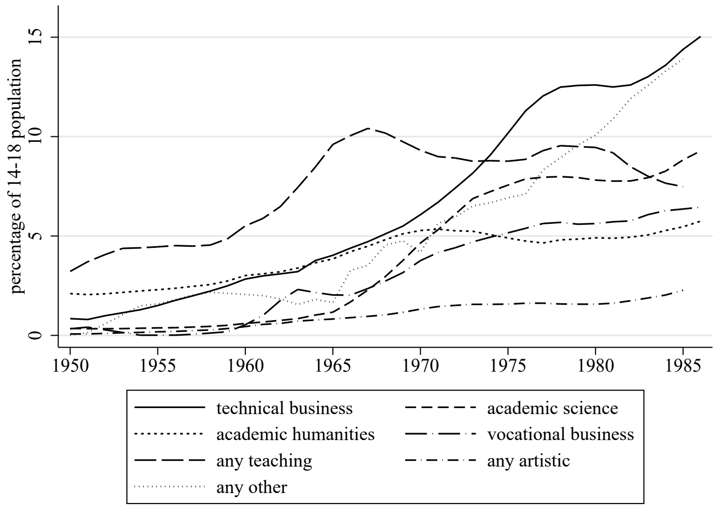


Figure 6: ENROLMENT IN LOWER AND UPPER SECONDARY EDUCATION BY SEX

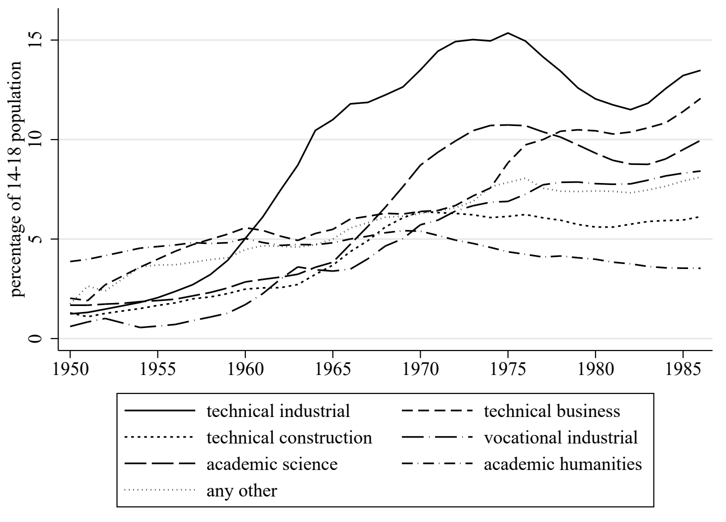
Gross enrolment rates in lower secondary education (age 11-13) and upper secondary education (age 14-18) by sex. Rates can be greater than 100 due to students repeating grades, students enrolling before the standard age, and students not officially residing in Italy. Years are defined as the calendar year at the start of the academic year (i.e. 1951 stands for academic year 1951/52). Data from 1998 to 2000 are estimated in the source due to gaps in coverage. Source: Istat (2011, p. 369).

category. For female students, technical schools with a business curriculum were the fastest-growing choice through most of the period, experiencing slow but continuous increases from the 1950s through the mid-1970s. By 1975 they overtook women's traditional first choice—schools preparing for teaching jobs. A similar evolution was followed by professional schools with business curricula, although with a delay and slower rate of growth. All the traditional choices stagnated in the second half of the 1970s, but overall enrolment was pulled by residual options, foremost academic schools with a foreign languages curriculum. Thanks to the expansion of these alternative options, total enrolment rates among females continued expanding throughout the period.

The fastest growing curriculum among male students, instead, was the industrial. Its expansion accelerated dramatically in the second half of the 1950s through 1965, and continued to grow at high rates in the following years. By the end of the decade, over



(a) Female



(b) Male

Figure 7: ENROLMENT RATES BY SEX, TRACK AND CURRICULUM, 1950-1986

Gross enrolment rates in upper secondary school, by track and curriculum (share of students enrolled on the 14-18 population, by sex). The year refers to the beginning of the academic year (i.e. October). 'Any other' includes all other choices. Source: own computations on education data from Istituto Centrale di Statistica, *Annuario statistico dell'istruzione italiana*, Roma, years 1953-1972 and *Id*, *Annuario statistico dell'istruzione*, Roma, years 1973-1990, and from Istat, *Serie storiche*, *Tavola 7.8* available for download from https://seriestoriche.istat.it (last retrieved June 2022). Population age 14-18 estimated for both sexes from the official intercensual reconstruction by summing the total population of age 14 to 18 in each year, dividing by two and multiplying by .96 for male and 1.04 for female, to account for the average sex ratio in Italy in the period considered. Official intercensual reconstruction available from Istat's *I.Stat* datawarehouse at http://dati.istat.it/Index.aspx?DataSetCode=DCIS_RICPOPRES1971 (for 1952-1972), http://dati.istat.it/Index.aspx?DataSetCode=DCIS_RICPOPRES1981 (for 1972-1981) and, http://dati.istat.it/Index.aspx?DataSetCode=DCIS_RICPOPRES1991 (for 1982-1991), last retrieved October 2021. For 1951, the population considered is that in the age range 15-19 in 1952; for 1950, the population is that in the age range 16-20 in 1952.

one in four male pupils was enrolled in a technical school preparing for manufacturing jobs.[15] However, the expansion of technical schools with an industrial curriculum slowed down in the first half of the 1970s, and enrolment sagged through the following decade. The only type of school to show a similar sag for male students was the academic track with a scientific curriculum, which had experienced comparatively fast growth in the previous decade. Male enrolment in most other school types stagnated through the 1970s: professional schools for manufacturing had expanded significantly between 1965 and 1975, but they made little progress in the following fifteen years, which suggests that they did not compensate for the missing students in the technical schools with a similarly-oriented curriculum.

A compensatory effect can instead be detected with respect to technical schools with a business curriculum: a growth spurt around 1975 made this type of school the second most favourite choices among male teenagers, even though it never reached the enrolment rates of the technical schools with an industrial curriculum. No compensatory dynamics can be detected within the academic track, instead, for the traditional humanities curriculum continued to follow a downward trend that had initiated in the late 1960s, and the new curricula (e.g. foreign languages) were not as popular among male teenagers as their female peers—in 1983, just over one male enrolled for every ten females.

This evidence suggest that the pause in the expansion of secondary school observed between the 1970s and the 1980s can be largely ascribed to male teenagers, whose enrolment rate in technical schools for industry dropped by 25% between 1973 and 1982 (respectively, its peak and through). Female enrolment in this type of schools also stagnated in the second half of the 1970s, following a decade of continuous growth; however, females amounted to only 3.3% of the total students enrolled in this type of schools in 1973—

[15] Including professional schools, the share of male students enrolled in upper secondary education preparing for manufacturing jobs reached 40% in 1970. enrolment in technical schools was buttressed by the extension of university access in 1965, which removed existing caps to the number of technical school graduates that could enroll in higher education. Using data for students in Milan, Bianchi and Giorcelli (2020, pp. 2617-2619) find that the reform increased graduates in STEM degrees that originated from technical schools, although the effect of this positive shock waned in the early 1970s. Transition rates from upper secondary school to university, in fact, increased almost continuously from 1954 to 1970, rising from a historical low of 40% to a peak of 67%. However, university students from technical schools remained a minority, and their family background often implied a greater necessity to work part-time, which increased the time to graduation and the risk of dropping out (Martinotti, 1969, pp. 89-204).

in fact, only 1.2% of female secondary school students opted for technical schools for manufacturing, compared to over 20% choosing technical school for business in the same year.

It is worth noting that, despite significant geographical differentials in enrolment levels, the bell-shaped trajectory of enrolment in technical schools for manufacturing was replicated across all of Italy's macroregions. Figure 8 shows that gross enrolment rates were predictably higher in the industrial core than in the South and especially in the Islands, but all areas exhibit rapid growth in the 1960s followed by stagnation in the first half of the 1970s and decrease in the second half of the decade.[16] The contrast between the two periods is most accentuated in the North-Western provinces, but the decline in enrolment rates can be identified in all areas. The downturn appears to start around the same time for most areas, possibly with a delay in the Islands.

These reconstructions suggest that, while there are common elements that could explain the slow-down in enrolment expansion for both men and women in the 1970s, the sagging of enrolment rates between the 1970s and the 1980s is largely explained by the contraction of technical schools for manufacturing jobs, which were largely attended by males. Is it possible that this evolution was caused by the egalitarian turn in collective bargaining after 1969? The evidence presented in this section is coherent with our hypotheses but cannot substantiate a causal claim. The next sections introduce the data and an identification strategy to credibly identify the effect of the minimum wage hike on post-compulsory school enrolment by sex, track and curriculum.

[16]Historical path dependency caused Southern Italy (including Sicily and Sardinia) to start from lower enrolment rates in secondary schooling than the rest of the country in the postwar period (A'Hearn and Vecchi, 2017).

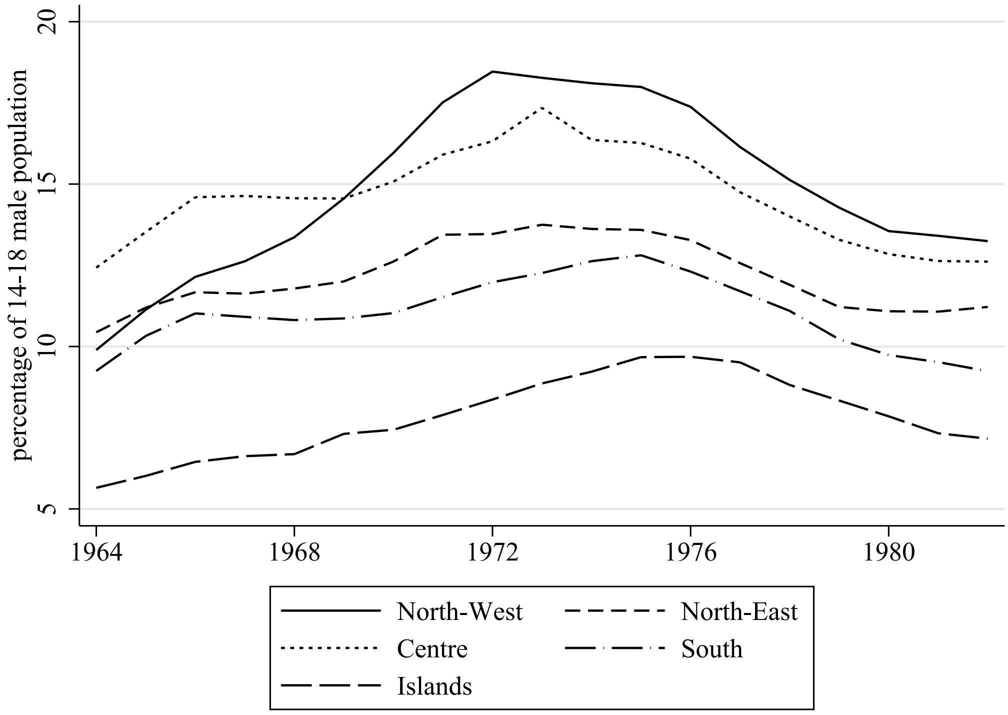


Figure 8: Male enrolment in technical schools for manufacturing by macroregion

Gross enrolment rates of male students in technical schools for manufacturing jobs (*istituti tecnici industriali*) by macroregion. The GER is computed as the ratio between the number of male students enrolled and the male population between the age of 14 and 18, in the relevant macroregion. North-West includes provinces in Valle d'Aosta, Piedmont, Lombardy and Liguria, North-East includes provinces in Trentino-Alto Adige, Emilia Romagna, Veneto, Friuli-Venezia Giulia, Centre includes provinces in Tuscany, Marche, Latium and Umbria, South includes provinces in Abruzzi, Campania, Molise, Apulia, Basilicata and Calabria, Islands include provinces in Sicily and Sardinia. For data sources and estimations see text and Ramazzotti (2023, methodological appendix)).

3 Data

The ideal setting to study the effect of the egalitarian wage bargaining on educational choices would provide information on career progression, wage earnings and detailed educational attainment for a representative sample of dependent workers across multiple birth cohorts before and after 1969. Unfortunately, such microlevel data does not exist in usable format for Italy during the period under study: official matched employer-employee datasets are only available since the 1980s, as are commonly-used household and labour force surveys—moreover, neither source provides information about the school track and curriculum attended. Italian census microdata are not available before 1971. Smaller surveys that collect such information are usually more limited in time range, geographical scope or statistical representativeness.

To circumvent these limitations, I present a new historical dataset which contains information on upper-secondary school enrolment, minimum and effective industrial wages and youth unemployment for Italy's provinces from 1962 to 1982, with annual frequency. The data has been digitized and harmonized from a range of printed primary sources, and adjusted to constant historical borders to ensure comparability. The province is the second-smallest administrative division for which data is consistently available in Italy, they are functionally comparable to the American counties, and the spatial level of detail is NUTS-3 in the EU framework for statistical units. During the whole 1962-1982 period, provinces had an average population of 596,746 (standard deviation 608) and mean area of 3,276 km^2 (standard deviation 1,849). The section provides a brief description of the computations behind each variable, while more detailed information on sources, comparability and harmonization procedures is provided in Ramazzotti (2023), methodological appendix).

3.1 Local minimum wages

To measure the opportunity cost of staying in school for the marginal student, I compute the mean local minimum wage in the province as the weighted average of the collectively-bargained wage floors for low-skill blue-collar workers, across industrial sectors (manufacturing proper plus construction, mining and utilities). This range of sectors covers 97% of manufacturing establishments and 98% of industrial workers according to the 1971 census. The weights are given by the number of employees in each sector and province, for I assume that a young individual with a lower-secondary school endowment of general human capital faces an expected entry-level wage which depends on the minimum wages payed in the industries that are present in the province, and on the probability of being employed in any such industries, which is assumed to be proportional to the number of employees in each industry. This procedure is also similar to adjusting the statutory minimum wage for coverage, as it is conventional in empirical applications where the minimum wage does not cover all workers in the area (Neumark and Wascher, 1992a). The minimum wage data has been digitized and harmonized from printed primary sources that consistently report the wage floors bargained in each sector, with annual frequency, at the province level until 1972 and at the national level thereafter (the motivation for this distinction is given in the next section). Note that due to missing data in some primary sources, the provinces of Arezzo and Ancona are excluded throughout the analysis.

The local industry shares are computed as the linear interpolation of industrial employees in each sector at the province level, from the industrial censuses of 1961, 1971, 1981 and 1991. The interpolation is necessitated by the lack of disaggregated annual data on industrial employment at the province level, but it also allows to avoid that the minimum wage series is affected by short-term shocks to local employment. Consequently, the mean minimum industrial wage $\overline{M}$ in province j at time t is obtained as:

$$\overline{M_{jt}} = \frac{\sum_{i=1}^{24} M_{ijt} \cdot \overline{S_{ijt}}}{\sum_{i=1}^{24} \overline{S_{ijt}}}$$

Where $\overline{S}$ is the share of employees in province j and sector i at time t, computed as the intercensual interpolation according to the formula:

$$\overline{S_{ijt}} = S_{ijT} \cdot \frac{(S_{ijT+10} - S_{ijT})/S_{ijT}}{10}$$

Where T is the earliest census year in any two consecutive, starting with 1961. This weighting procedure ensures to capture local long-term trends in sectoral composition, that affect a teenager ex-ante employment opportunities. The resulting series is represented by Figure A.2, with annual box plots showing the interquartile range and the median, adjacent and outside values. The graph clearly shows the acceleration of the contractual minima after 1969, but also the compression minimum wage differentials between provinces, which will be described in the next section.

3.2 School enrolment and control variables

The main outcome variable is enrolment in upper secondary school, by track and curriculum. To obtain these rates at the province level with annual frequency, I have first collected and digitized annual official statistics on the number of students enrolled, distinguishing by sex and province, and I have harmonized the resulting series to constant historical 1961 borders. Secondly, I have reconstructed the size of the relevant age group (14-18 years old) at the province level from census statistics.

The use of population censuses is necessitated by the lack of official reconstructions of intercensual population at the province level before the year 1982. To perform this reconstruction, I have digitized tables reporting the age distribution of the resident population in each province in 1961 and in 1971, I have linked them to the official intercensual reconstructions for 1982, and I have harmonized the data to constant historical borders. To obtain the intercensual estimates I first identified the year of birth for each age group, and then I ran a linear interpolation between the benchmark years for each birth cohort, by sex and province. Finally, the size of the 14-18 age group was computed by summing the number of individuals in the respective age range, for each year-province cell. Ramazzotti (2023, methodological appendix)) provides additional details on the interpolation method, corrections to the data, and limitations of this methodology and sources.

Since these are entirely new reconstruction obtained from a range of different sources, I have checked their compatibility with aggregate statistics that are available at the national level. The tracks and curricula included in the dataset cover about 90% of all secondary school students in any given year, and an even larger percentage for male students. The total number of students, including those enrolled in the residual curricula, is virtually identical to the numbers presented by Checchi (1996). The gross enrolment rates, instead, are lower than Checchi's estimates due to differences in the size of the age group which are attributable to errors in the population census of 1971. Following the correction proposed by Caselli, Golini, and Capocaccia (1989), I obtain an age-group profile that is compatible with official intercensual reconstructions at the national level. Consequently, my computations of gross enrolment rates are in line with the aggregate time series published by Istat. Additional details on these procedures and checks are presented in Ramazzotti (2023, methodology appendix)).

To test that the minimum wage had a sizeable effect on the wage distribution, one would need to know the complete earnings distribution for blue-collar workers at the province level in the period under consideration. While this is not possible due to the mentioned data limitations, I have collected, digitized and harmonized aggregate statistics on the average wage of blue-collar workers by industrial macro-sector. The data were published with annual frequency by INAIL, the National Institute for Insurance against Workplace Accidents. The publications reported the mean daily earnings of blue-collar workers that suffered a temporary incapacitating accident on the workplace in the solar year. The earnings were reported separately for each province in nine macro-sectors. To harmonize the series with the minimum wage and industrial census data, I have devised a conversion system—reported in table B.1—and I have rescaled all wages by a common coefficient to correct for an underestimate of the wage level in the source—the correction does not alter relative wages between provinces and sectors. The local mean effective wage is obtained as the weighted average of mean wages across the ten macro-sectors.[17]

The control variables include the provinces' population size, income per capita, value added in industry, and prime age and youth unemployment. The economic variables have been digitized and harmonized to constant historical borders from the income

[17] For details on sources and methodologies, see the appendix in Ramazzotti (2023).

accounting estimates at the province level produced by Guglielmo Tagliacarne and the namesake Institute (Tagliacarne, 1972; Tagliacarne, 1975; Tagliacarne, 1979).[18] Youth unemployment is obtained from registrations at job centres at the provincial level (see the appendix in Ramazzotti (2023) for a discussion of the sources). Table 1 shows the descriptive statistics from the dataset, distinguishing by time period, that is before 1969, during the convergence period, and after 1972.

Table 1: Descriptive statistics by period

	(1) 1962-1968		(2) 1969-1972		(3) 1973-1982	
	mean	sd	mean	sd	mean	sd
ln(minimum wage)	7.62	0.17	8.16	0.17	9.62	0.61
ln(average wage)	8.15	0.25	8.72	0.21	9.99	0.57
% early school leavers	69.53	9.32	56.10	9.29	48.31	8.68
% GER tech manufacturing male	9.75	5.61	12.10	5.44	11.54	4.86
% GER tech manufacturing female	0.25	0.37	0.43	0.49	0.66	0.66
% GER tech business male	5.20	1.90	6.29	2.05	9.37	2.99
% GER tech business female	4.54	2.15	7.40	2.84	13.08	4.54
% GER voc manufacturing male	3.82	2.19	5.96	2.78	7.64	3.30
% GER voc manufacturing female	0.13	0.34	0.27	0.51	0.65	1.07
% GER voc business male	0.64	0.51	0.84	0.73	0.89	0.87
% GER voc business female	2.33	1.67	4.08	2.35	5.76	3.02
% GER academic science male	3.95	2.20	8.13	2.92	8.69	2.72
% GER academic science female	1.70	1.26	5.39	2.30	7.36	2.57
ln(prime-age male unemployed)	8.47	0.79	8.24	0.85	8.16	0.97
ln(prime-age female unemployed)	7.17	1.11	7.25	0.97	7.79	0.87
ln(under-21 male unemployed with previous job)	6.11	0.89	5.79	0.88	6.20	0.94
ln(under-21 female unemployed with previous job)	5.35	1.04	5.28	0.88	6.13	0.89
ln(under-21 male first job seekers)	6.56	1.03	6.53	1.04	7.01	1.22
ln(under-21 female first job seekers)	5.83	0.90	6.02	0.88	7.16	1.14
% 14-18 pop	7.65	1.11	7.13	1.20	7.65	1.14
% 15-21 pop	10.59	1.22	9.97	1.54	10.47	1.46
ln(population)	13.02	0.65	13.03	0.66	13.05	0.68
ln(industrial value added)	25.16	0.96	25.72	0.85	27.03	1.07
ln(gdp per capita)	13.19	0.41	13.74	0.29	15.01	0.70
Observations	629		360		900	

[18] Data for years 1978-79 and 1981-82 are linearly interpolated and extrapolated from 1977 and 1980, respectively, due to missing sources at the time of writing. Data for 1962 are not included in this version of the paper.

4 Identification strategy

4.1 Baseline specification and endogeneity concerns

We argued that the egalitarian collective agreements of the 1970s were functionally equivalent to statutory minimum wages, hence we can take references for the identification strategy from the vast empirical literature on the latter. A common requirement for the identification of the causal effects of minimum wages is, in fact, the availability of credible research designs. Since the path-breaking contributions of Katz and Krueger (1992) and Card and Krueger (1994), quasi-natural experiments that exploit exogenous spatial variation in treatment are considered among the most appropriate strategies for minimum wage studies. In the standard approach, Difference-in-Differences estimates are employed to compare labour markets that receive a minimum wage hike with similar localities that remain untreated in the period under study. To strengthen the external validity of the results, two-way fixed-effect estimation with panel-data have been applied, for they generalize the Difference-in-Differences approach by exploiting identifying variation from multiple localized differences in minimum wages over time (Neumark and Wascher, 1992b; Neumark, Salas, and Wascher, 2014, Wolfson and Belman, 2019), even though debates continue regarding the most appropriate strategies to select comparable groups (Allegretto et al., 2017; Neumark and Shupe, 2019b; Manning, 2021).

In our case, the spatial variation in minimum wage levels originates from differences in the industrial structure of the provinces and from the different evolution of minimum wages established by collective agreement in each sector. In the baseline approach, we would estimate the following structural equation as two-way fixed-effects model where the dependent variable Y (for instance, early school leavers) in province i at time t is regressed on the level of the mean minimum wage M, controlling for a vector of time-varying covariates X, and including province and time fixed effects (respectively, α and τ). Given the possibility that our time-varying covariats could be influenced by the minimum wage hike, the vector of control includes only the trended variables' mean value before 1968 (Joshua David Angrist and J. Pischke, 2009; Caetano et al., 2022).

$$ln(Y)_{it} = \beta\, ln(M)_{it} + X'_{it}\,\gamma + \tau_t + \alpha_i + \varsigma_{it} \qquad [1]$$

The coefficient β in Equation 1 would thus provide the marginal effect of increasing the mean minimum wage floor in the province on the dependent variable. However, the endogeneity of minimum wage determination is a major threat to causal identification in this type of designs. In the literature on statutory minimum wages this possibility mainly arises from self-selection by local authorities (Card and Krueger, 1995, pp. 183-186; Baskaya and Rubinstein, 2015). Statutory minimum wages that are set at the state or city level, in particular, are the product of political processes that incorporate information on the local economy. The case of Italy differs because minimum wages were bargained between labour unions and employers' associations at the national level, for each sector. Our estimate of β would be biased if the bargaining process within each sector incorporated unobservable information on local labour markets, and/or if the change in sectoral minimum wages modified the industrial structure of the province, thus biasing the computation of the average minimum wage.

Of the two potential threats, the first appears less plausible. The centralization of collective bargaining at the sector level implies that local labour market conditions were less relevant than sectoral trends. However, it is possible that employment and wage levels in core industrial areas were considered in the bargaining process, especially for regionally concentrated sectors (Manacorda and Petrongolo, 2006). The second threat appears, instead, more plausible, for we cannot exclude that the local minimum wage affected the composition of the industrial structure within each province, creating a feedback loop with respect to our independent variable. This effect appears particularly plausible for the period after 1969, following the steep increases in minimum wages across sectors. To address these potential threats of endogeneity, we need to isolate variation in the minimum wage hike that is uncorrelated with the error term in the post-1968 period.

4.2 Exogenous variation in treatment intensity

One source of exogenous variation in the intensity of the minimum wage is provided by a contemporaneous institutional change in the wage-setting system, which offers a credible research design. Before 1969, sectoral collective agreements established a single nominal wage floor for each skill category, but its level applied only to the provinces of Milan and Turin. In all other provinces, its effective value was automatically rescaled according to local coefficients which were based on differences in the cost of living, in order to equalize real wages over the national territory. This practice had been established in the postwar period to contrast high inflation that showed significant spatial variation depending on the impact of the war. Originally, thirteen local coefficients were established. Each of Italy's ninety-two provinces was assigned to an index, according to the similarity of local price levels in 1946, thus creating thirteen 'wage zones'. For each sector, minimum wage nominal levels were equal within, and differed between, wage zones (Poy, 2015). The system was partially reformed in 1961, when the number of wage zones was reduced to seven, and the maximum difference between the local minimum and Milan's nominal level was set to 20%.[19]

This system was maintained through the 1960s, but it met growing aversion from the labour unions. The unions' proposal to repeal the wage zones hinged on equity grounds: reformist union leaders argued that workers performing the same tasks should be paid the same irrespective of the location of the factory (Poy, 2017). The contrast grew in 1968, when calls for equal nominal pay for the same jobs found the support of more radical groups inside and outside unions. An agreement was eventually reached in March of 1969, establishing a convergence process which would partly reduce the wage differentials on 1 April 1969, halve them on 1 October 1970 and remove all remaining differentials on 1 July 1972.[20]

[19] Attachment to the Interconfederal agreement of 2 August 1961 (*Accordo interconfederale per la revisione dell'assetto zonale delle retribuzioni e il conglobamento della contingenza 2 Agosto 1961*), available for download from the website of CNEL (National Council for Economics and Labour) at https://www.cnel.it/Archivio-Contratti (last retrieved July 2021).

[20] See article one of the *Accordo Interconfederale 18 marzo 1969 per il conglobamento della contingenza e per la revisione dell'assetto zonale delle retribuzioni* available at https://www.cnel.it/Archivio-Contratti. Notice that a similar agreement had already been reached between the labour

The resulting compression of spatial differentials in minimum wage levels can be appreciated from Figure A.4, which plots the log difference between the average minimum wage in Milan and the rest of Italy's provinces. The graph shows that the median difference decreased from about 5 percentage points in 1968 to effectively zero in 1972. Most importantly, however, the extent of the reduction shows significant variation, ranging from less than one percentage point for the 25th percentile to over ten percentage points for the 75th percentile. While a mild reduction appeared underway in the early period (possibly driven by the Intersind agreement of 1968), it is clear that most of the convergence took place in 1969-1972 and was completed by 1976.

Thus, the compression in minimum wages between provinces after the repeal of the wage zones represents an exogenous source of spatial variation in treatment intensity because, during the adjustment period, provinces that started at lower nominal levels relative to Milan experienced a steeper minimum wage hike than provinces whose mean minimum wage in 1968 was closer in levels to that of Milan, irrespective of their industrial composition and local labour markets conditions. Figure 9 shows that there was a strong association between the deviation of mean minimum wages from Milan's level in 1968 and minimum wage growth during the adjustment period (1968-1972), while no significant association is found between 1964 and 1968, which strengthens our argument that the repeal of the wage zones was an exogenous shock to the determination of the local minimum wages and did not correlate with pre-trends.[21]

A similar source of spatial variation in the intensity of a minimum wage hike has been used by Kawaguchi and Mori (2021) for identifying the impact on unemployment after 2007 in Japan. In that case, the variation originated from the introduction of the

unions and the labour relations' representative for state-owned companies (Intersind). See article two of the *Accordo Interconfederale 21 dicembre 1968 per il conglobamento dell'indennità di contingenza e per il graduale superamento delle differenze zonali delle retribuzioni* available on the digital Historical Archive of the collective labour agreements maintained by CNEL (Italy's National Council of the Economy and Labour) at https://www.cnel.it/Archivio-Contratti).

[21]Notice that there are few observations with a higher mean minimum wage than Milan in 1968, which is due to a greater concentration of high-wage industries in these provinces and their location in high minimum wage zones. As expected, these provinces experience a lower increase in minimum wage between 1968 and 1972 than the average, but higher than the one predicted by the linear regression. In the main analysis I include these provinces in the estimates, but to ensure that results are robust to these outliers, I have also run the regressions including only provinces that had a mean minimum wage smaller than Milan in 1968 (available upon request).

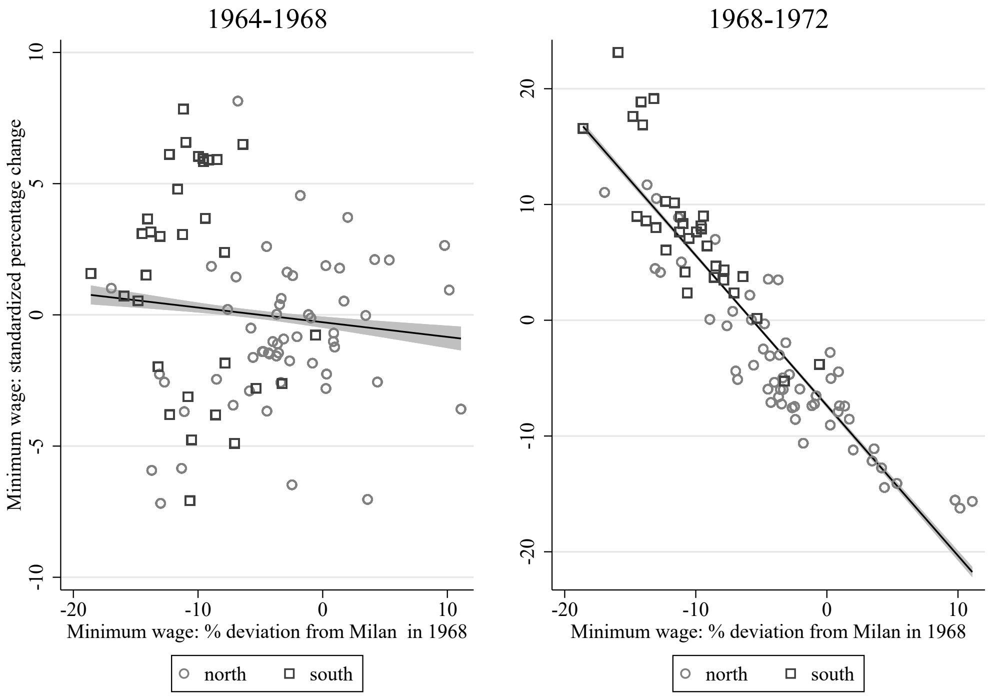


Figure 9: Nominal wage equalization and minimum wage growth

The figure shows the relationship between the four-year change (expressed in log-point differences) of the minimum wage between 1964 and 1968 (left panel) or between 1968 and 1972 (right panel), and the log-point difference with respect to the minimum wage level in Milan in 1968. Each circle represents one of 91 provinces. The size of the circle is proportional to the size of the province population. The solid line represents the linear prediction from the scatterplot, while the shaded area represent the 95% confidence interval. For the sources of the minimum wage data see Ramazzotti (2023, appendix).

indexation of province-specific minimum wages to the local cost of living, rather than its repeal, but the effect on the wage hike was comparable to our case.[22] Given the similarity between the two natural experiments, we can adapt the estimation strategy of Kawaguchi and Mori (2021) to our case. The analysis is realized in two steps. First, we estimate the average treatment effect of the minimum wage hike on schooling using an instrumental variable approach.[23] Second, we explore the dynamic response over time (both before, during and after the transition period) by estimating the reduced-form regression. This second step is akin to a natural experiment where treatment assignment

[22]In our case, minimum wages rose more steeply in provinces with lower cost of living after the repeal of the indexation; in the case of Japan, minimum wages rose more steeply in places with higher cost of living after the introduction of the indexation (Kawaguchi and Mori, 2021, pp. 390-391). To the best of my knowledge, this is the only published research exploiting a similar natural experiment for identification purposes.

[23]This approach builds on Joshua D. Angrist and Imbens (1995).

depends on the province's wage zone. In this case estimation is equivalent to a generalized Difference-in-Difference design with a continuous variable.

4.3 Instrumental variable approach

In the first step, the instrumental variable approach allows to identify the average marginal effect of the minimum wage hike on the outcome variables by exploiting, for the period after 1968, only the variation that is predicted by the repeal of the wage zones. To retrieve this effect, I estimate the following baseline model by two-stage least squares:

$$ln(Y)_{it} = \pi\, ln(\widehat{M})_{it} + \psi X'_{it} + \tau_t + \alpha_i + \eta_{it} \qquad [2]$$

$$ln(\widehat{M})_{it} = \sum_{y=1962}^{1982} \theta_y \left[ln(M)^{1968}_{Milan} - ln(M)^{1968}_{i}\right] * 1(Year = y) + \phi X'_{it} + \tau_t + \alpha_j + \epsilon_{jt} \qquad [3]$$

In Equation 2, Y is the dependent variable in province i at time t and M is the instrumented nominal minimum wage (the endogenous regressor). The instrument in Equation 3 is the the gap between the nominal minimum wage in the province of Milan in 1968 and that of province i in the same year, expressed in log differences. To allow the coefficients of the instrument to vary over time, I interact the deviation from Milan with time dummies. This instrument predicts the spatial variation in the minimum wage hike which is only caused by the repeal of the wage indexation system in 1969—hence, only the variation that is exogenous with respect to local labour market and industrial characteristics.

Like before, the vector of time-varying controls (X'_{it}) is built using the pre-treatment values (averaged over the years 1962-1967), and the coefficient for the year 1967 is set to zero as a reference point (Borusyak, Jaravel, and Spiess, 2021). I also include time (τ) and province (α) fixed effects. Standard errors are clustered at the province level to control for possible serial autocorrelation. The second stage in Equation 2 regresses the dependent variable on the instrumented minimum wage, and controls for the same variables as in Equation 3, including province and time fixed effects. Standard errors are also clustered at the province level. Thus, assuming that the parallel trends assumptions are met, the coefficient π recovers the ATE for the minimum wage on the dependent variable. It is important to notice that the estimated ATE concerns only the transition period 1968-1972, because it exploits the variation caused by the spatial equalization of nominal wages, and not the total increase in minimum wage levels.

4.4 Assessing the instrument

To establish the relevance of the instrument, I run the first stage regression from Equation 3. The estimate of the θ coefficient of the instrument is plotted for each year in Figure 10. The figures shows that the gap with respect to Milan in 1968 had no significant association with the minimum wage levels in the pre-treatment period. However, after the repeal of the wage zones in 1968, the coefficient quickly turns positive and statistically significant, stabilizing at around 1 in the post-transition years. These results imply that a 1% difference in the mean minimum wage with respect to Milan in 1968 predicts a 0.8% higher minimum wage level after 1972, which is additional proof that nominal wages became substantially equalized between provinces. This confirms the argument that the repeal of the wage zones introduced a source of exogenous variation in the steepness of the minimum wage hike that is uncorrelated with previous levels, which we will use for identification in the second stage.[24]

The random assignment of the instrument cannot be formally tested, but the historical context can provide a motivation. The spatial variation in treatment intensity appears to be independent from local labour market conditions by construction: the classification into the different wage zones was based on post-war inflationary pressure—over twenty-five year prior to the minimum wage hike—and partly on the simplification of 1961, eight years earlier.[25]

[24] Alternatively, I run the regression using the mean minimum wages computed at constant industry share, without obtaining significantly different results (not reported, results are available upon request).

[25] Wage zone 0 included provinces from three regions in the North-West (Turin, Milan and Genoa) and one from the Centre (Rome), each from a different region; zone 1 included three provinces in Lombardy plus Florence (in Tuscany); zone 2 included provinces from eight different regions (Valle d'Aosta, Piedmont, Liguria, Lombardy, Trentino-Alto Adige, Veneto, Friuli-Venezia Giulia, and Tuscany); zone 3 covered provinces from six regions (Piedmont, Lombardy, Liguria, Emilia-Romagna, Veneto and the province of Naples in Campania); zone 4 included provinces from Piedmont, Tuscany, Emilia-Romagna,

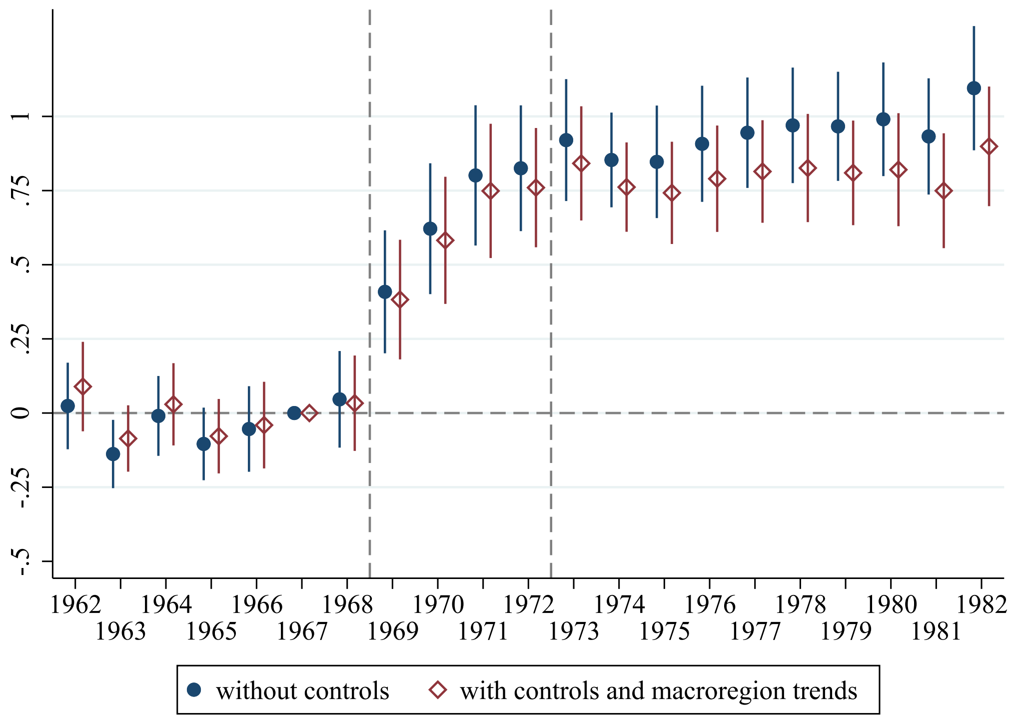


Figure 10: Gap with respect to Milan in 1968 and minimum wage levels

OLS estimates of the coefficients for the interaction term between the log difference between the mean minimum wage in Milan and in the province in 1968 and year indicators, with 1967 set to zero. Both specifications include both province and time fixed effects. In the second specification for both models, the regression controls for pre-1968 averages of total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector are interacted with a linear trend, as well as macroregion dummies interacted with a linear trend. The vertical solid lines indicate the 95% confidence interval with standard errors that are clustered at the province level. The vertical dashed line indicate the start and the end of the convergence period, following the repeal of the wage zones in 1968. The number of provinces is 90 (minimum wage data for the provinces of Arezzo and Ancona are missing from the sources).

However, since low-wage zones were predominantly located in the continental South, the minimum wage growth during the transition period shows a strong North-South gradient, as represented in Figure A.3. This is due to two distinct causes: first, the Allied liberation of Italy moved from the South to the North of the country, leaving the provinces south of the Gothic Line exposed to the inflationary pressure of the military government's

Veneto, Friuli-Venezia Giulia, the province of Ancona in the Marche region and that of Palermo in Sicily (seven regions in total); zone 5 covered provinces across the Centre, the South and the Islands (nine regions, including Tuscany, Umbria, Marche, Latium, Abruzzi, Campania, Apulia, Sicily and Sardinia); zone 6 covered most Southern regions (all of Calabria, Basilicata and Molise, and parts of Abruzzi, Campania, Apulia, Sicily and Sardinia) but it also included one province in the Marche region (Macerata). Consequently, there was significant geographical variation in wage zone assignment: the North-West included six wage zones (considering Genoa separately from Milan and Turin), the North-East three, the Centre six and the South and islands four.

new currency for longer (Harris, 1957, pp. 445-449), so much so that these wage zones were assigned a different indexation mechanism once the country was entirely liberated;[26] second, the reform of the wage zone system in 1961 decreased the number of wage zones especially in the South, thus reducing the spatial variation of the instrument within this macroarea. Given the historical differences in economic development, social structure and culture between the North and the South, these otherwise unrelated causes could correlate with unobservable variables that can also affect schooling decisions. To control for this, robustness checks include macroregion indicators interacted with linear time trends.

As usual, the exclusion restriction that the instrument affects the outcome variables only through the endogenous regressor cannot be formally tested, but the contextual evidence suggests that a province's wage zone had little effect on enrolment rates except through the minimum wage level. Wage zones did not overlap with administrative divisions, which reduces the risk that the spatial variation correlates with local policies. One case where this case might not apply are special statute regions, who had legislative autonomy on several areas, including education. For this reason, I exclude from the main analysis regions that had special statute in 1962.[27] Adding or excluding the regions from the analysis does not significantly alter the results, except for the case of Sicily, which introduces a contrarian response in the baseline analysis with respect to all other regions. Future research should address this specificity in light of the special characteristics of this region (few wage zones, greater legislative autonomy, prevalence of organized crime).

4.5 Generalized Difference-in-Differences

The second step of the analysis consists in using the repeal of the wage zones as a natural experiment, which allows to estimate the dynamic average causal response of school enrolment to the minimum wage shock of 1969, over time. Following Kawaguchi and Mori (2021) again, we can specify the reduced-form regression:

[26]See 'Concordato del 23 Maggio 1946 per la perequazione del trattamento economico dei lavoratori dell'industria nelle provincie dell'Italia centro-meridionale', reprinted in *Gli accordi interconfederali di lavoro dal 1944 al 1954* (1955, pp. 44-62)

[27]Special statute regions in 1962 were Sicily, Sardinia, Valle d'Aosta and Trentino-Alto Adige, totalling fifteen provinces out of ninety-two. Friuli-Venezia Giulia was made special statute region in 1963.

$$ln(Y)_{it} = \sum_{y=1962}^{1982} \delta_y \left[ln(M)^{1968}_{Milan} - ln(M)^{1968}_{i}\right] * 1(Year = y) + \rho X'_{it} + \tau_t + \alpha_j + \zeta_{jt} \quad [4]$$

Where the dependent variable Y is regressed directly on the gap with respect to Milan in 1968. This regression is equivalent to a generalized Difference-in-Differences approach where the continuous treatment variable (the log difference in nominal minimum wages in 1968 between each province i and Milan) predicts the extra increase of the province's minimum wage that was caused only by the repeal of the wage zones.

Figure 11 provides a simplified diagram of this variation for two representative provinces, one with a small gap with respect to Milan before 1969, and one with a larger gap. The diagram assumes that there is only one industry, so the starting difference between provinces in the nominal level of the minimum wage is entirely attributable to their being assigned into different wage zones. Following the egalitarian turn of 1969, the minimum wage set by collective agreement for Milan starts rising. Meanwhile, however, the repeal of the wage zones requires that the minimum wage in the other province converge to Milan's level. Consequently, the wage hike for these provinces is steeper than in the counterfactual scenario where the wage zones are not repealed. This extra wage raise represents the treatment in our design. All provinces but Milan are treated at the same time, but the intensity of treatment differs on a continuous scale. Following the terminology in Callaway, Goodman-Bacon, and Sant'Anna (2021), the larger the gap in 1968 with respect to Milan, the greater the 'dose' of the minimum wage hike that the province receives in 1969-1972. In the example, the province starting with the smaller gap receives a low does, while the province starting with a large gap receives a high dose. The coefficient δ would thus recover the average causal response of the treated provinces (i.e., provinces where the 1968 gap was larger than zero, in absolute value) in each year, provided that the stricter parallel trends assumptions identified by Callaway, Goodman-Bacon, and Sant'Anna (2021) hold. These require the usual conditions of DiD designs as well as that 'for all doses, the average change in outcomes over time across all units if they had been assigned that amount of dose is the same as the average change in outcomes over time for all units that experienced that dose' (Callaway, Goodman-Bacon, and Sant'Anna, 2021, p. 11). This assumption requires that provinces receiving a smaller dose of treatment are a good counterfactual for provinces receiving a larger dose. This assumption does not hold if observations self-select into the dose levels, but this does not seem to apply to our case: the dose depended only on which wage zone the province had been assigned to in 1953 (as reformed in 1961), which increases our confidence that the extra minimum wage hike was as good as randomly assigned between the provinces.

5 Results and discussion

5.1 Impact on the opportunity cost of schooling

To verify that the minimum wage hike effectively increased the opportunity cost of schooling, I first test the impact of contractual minima on the effective wages of blue-collar workers in manufacturing. If we cannot reject the null hypothesis that the minimum wage hike had no effect on blue-collar wages, it would be difficult to argue that it was influential enough to affect teenagers' schooling decisions.

Table 2 reports the OLS estimates for the structural Equation 1, and the 2SLS results from Equation 2, separately with and without macroregion trends. The dependent variable is the average wage of blue-collar workers in the industrial sector, while we control for a vector of pre-1968 averages of total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector, interacted with a linear trend.

The OLS estimate without controls and macroregion trends suggests an elasticity of .76 with respect to the mean minimum wage, which decreases to 0.46 in the fully-saturated specification, and the results are essentially confirmed by the IV estimates. The implied elasticities are particularly high, especially compared to the wage elasticities commonly found in studies on statutory minimum wages. Our results can be explained by the greater bite of wage floors established by collective agreements and the centralized bargaining system, which reduced room for firm-level adjustments, and by the fact that they affected

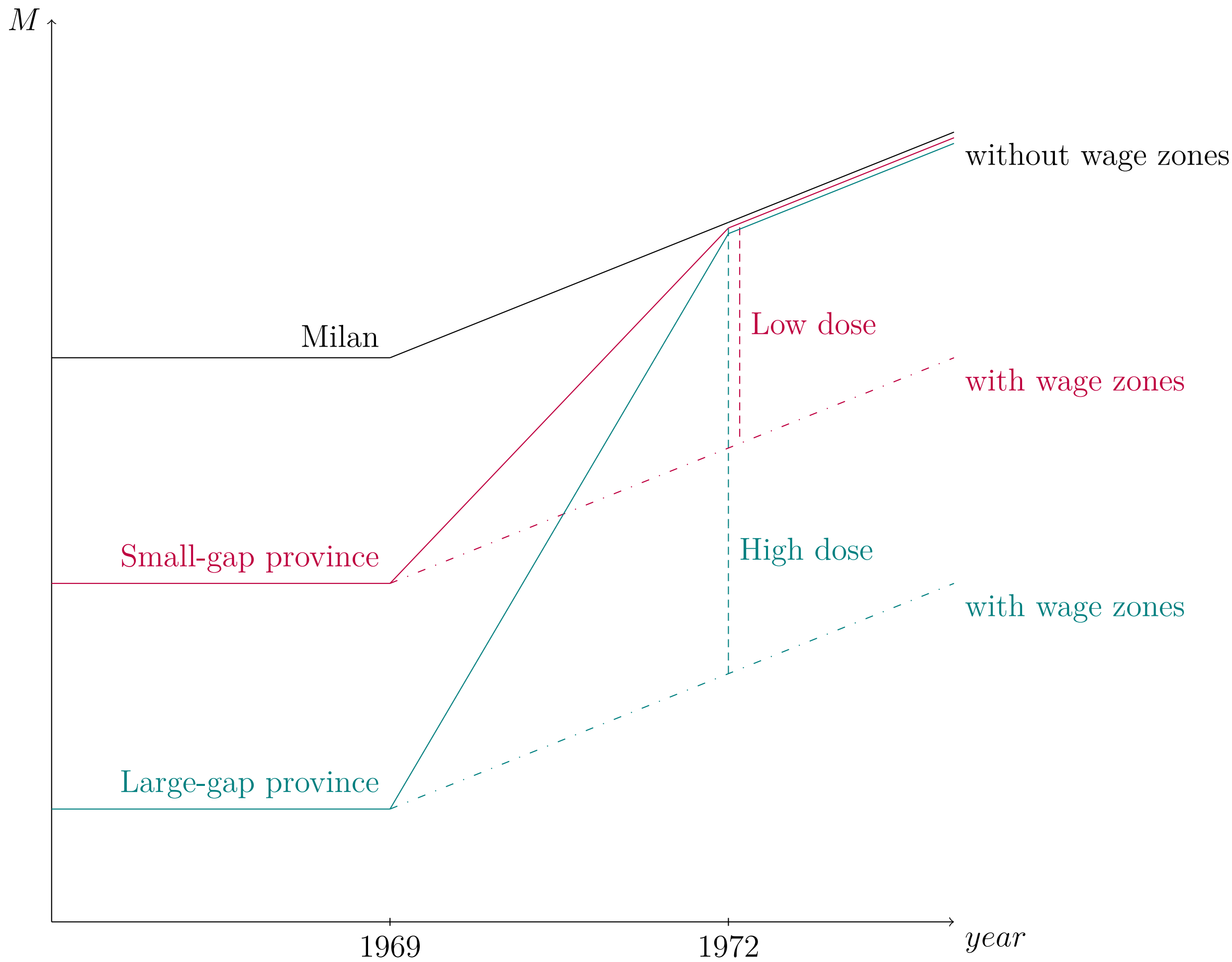


Figure 11: Representation of the identification strategy for DiD

This diagram shows a schematic representation of the identification strategy for the generalized Difference-in-Differences. The solid lines represent the level of the mean minimum wage (M) in the province of Milan and in two representative provinces—one with a small gap with respect to Milan in 1968, and another with a large gap. The size of the gap at the start of the period depends on the wage zone to which the province is assigned before 1969. After the repeal of the wage zones in 1969, the minimum wage level in all provinces must converge to that of Milan by 1972. The dotted lines represent the counterfactual minimum wage if the wage zones had not been abolished. The dashed vertical line represent the variation in minimum wage caused only by the repeal of the wage zones. Both provinces are treated with this extra wage hike, but the province with a small starting gap receives a lower 'dose' of treatment than the province with a larger starting gap. Milan represents the control group.

Table 2: Wage elasticity to average minimum wage

	ln(average effective wage)			
	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
ln(minimum wage)	0.865***	0.493***	0.986***	0.536***
	(0.0778)	(0.0929)	(0.107)	(0.148)
Province FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
Macroregion trends	No	Yes	No	Yes
Clustered SE	Yes	Yes	Yes	Yes
Within R2	0.998	0.998		
Between R2	0.478	0.612		
Overall R2	0.984	0.971		
Kleibergen-Paap rk Wald F statistic			1249.1	360.6
N	1890	1890	1890	1890

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

OLS estimates for the baseline model (columns 1-2) and 2SLS estimates for the IV model (columns 3-4). The dependent variable is the natural logarithm of the average effective wages of blue-collar workers employed in manufacturing, construction and utilities. In the second specification for both models, the regression controls for pre-1968 averages of total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector, each interacted with a linear trend, as well as macroregion dummies interacted with a linear trend. All specifications include time and province fixed effects. Standard errors are clustered at the province level.

the wage distribution across all sectors. Even though the absence of detailed data on the wage distribution does not allow to test the for compliance in levels, these results suggest that collective agreements influenced growth rates, for changes to the sectoral minima were almost entirely incorporated into the growth rate of effective wages.

5.2 The response of early school leavers

5.2.1 The marginal effect across across the whole period

Having established that minimum wage increases impacted effective wages and thus could influence the opportunity cost of schooling, we turn our attention to testing the main hypotheses of the paper, i.e. that high contractual minimum wages can discourage post-compulsory school enrolment. I start the analysis by estimating the baseline specification in Equation 1 by OLS and the IV model in Equation 2 by 2SLS, using as dependent variable the log of the number of early school leavers, which is defined as all individuals between the age of 14 and 18 that are not enrolled in upper secondary education, controlling for the size of the same age cohort in the province, besides all other pre-1968 averaged controls interacted with a linear trend. Table 3 presents the estimates for β and θ, both with and without macroregion trends.

Table 3: Minimum wage and early school leavers

	ln(average effective wage)			
	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
ln(minimum wage)	0.644***	0.387***	0.868***	0.894***
	(0.120)	(0.126)	(0.201)	(0.316)
Province FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Controls	No	Yes	No	Yes
Macroregion trends	No	Yes	No	Yes
Clustered SE	Yes	Yes	Yes	Yes
Within R2	0.930	0.946		
Between R2	0.958	0.967		
Overall R2	0.953	0.964		
Kleibergen-Paap rk Wald F statistic			353.8	308.5
N	1575	1575	1575	1575

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

OLS estimates for the baseline model (columns 1-2) and 2SLS estimates for the IV model (columns 3-4). The dependent variable is the natural logarithm of the number of individuals enrolled in upper secondary education in the academic year running from October to June. The models control for the size of the cohort between the age of 14 and 18 in the province. In the second specification for both models, the regression controls also for pre-1968 averages of total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector, each interacted with a linear trend, as well as macroregion dummies interacted with a linear trend. All specifications include time and province fixed effects. Standard errors are clustered at the province level. The observations exclude provinces in special statute regions, so the total number of provinces is 75 (minimum wage data for the provinces of Arezzo and Ancona are missing from the sources).

A higher mean minimum wage is associated with a larger number of teenagers not enrolled in upper secondary education, although the size of the effect varies between specifications. In the structural specification a 1% increase in the mean minimum wage is associated with 0.64% increase in early school leavers, which is attenuated to 0.39% after we control for pre-1968 averages of total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector, interacted with a linear trend. The 2SLS estimates present the same sign and are significantly larger. In the specification without controls, the coefficient increases to 0.87 and, after the inclusion of the controls, it reaches 0.89. The difference in the estimates between the structural equation and the IV approach corroborate our argument for instrumenting the nominal minimum wage. Moreover, the Kleibergen-Paap F Statistic for the first stage is confidently larger than the critical value, reassuring us about its strength. All results are statistically significant at the 99% level.

These results appear also economically significant: with the log of the minimum wage increasing by circa 34% between 1968 and 1972, our estimates would predict an increase in early-school leaving between 13% and 27%. This is close to but larger than the 12% difference between the enrolment rate extrapolated from the trend before 1969 and its effective value in 1972.

5.2.2 The dynamic response to the 1969 minimum wage hike

The previous estimations have established that, across our panel of provinces between 1962 and 1982, exogenous minimum wage hikes were associated with significant increases in early school leavers. However, our research question focused the attention on the steep rise of minimum wages after 1969, for this is the natural experiment that can provide a clean identification of the causal effect. To perform this analysis, I estimate the reduced-form regression from Equation 4, which allows for a dynamic response to the shock. As previously discussed, this approach is equivalent to a generalized DiD that exploits the exogenous variation in the intensity of the minimum wage hike to recover the causal response.

The dependent variable is defined as the log of the number of individuals between 14 and 18 not enrolled in post-compulsory secondary school, and the vector of controls includes the size of the cohort (logged). This definition allows us to immediately interpret the coefficient of interest as the marginal increase in the number of young people that leave school early for a 1% increase in treatment. Treatment is continuous and is defined as the gap between the minimum wage in Milan and the minimum wage in the province in 1968. Recalling our previous discussion, we know that provinces with a larger starting gap with respect to Milan experienced a steeper increase in minimum wages after 1968. This differential increase provides the source of variation in treatment intensity to estimate the causal response over time. To recover the dynamic causal response, I interact this variable with a full set of time dummies, setting the coefficient for 1967 equal to zero.

Figure 12 reports the coefficient from the interaction term, both including and excluding trended pre-treatment controls. The inclusion of the controls attenuates slightly the causal response but does not modify the interpretation. The figure shows that there was no association between early school leaving and the 1968 minimum wage gap before the wage shock, providing indirect evidence in support of the parallel trend assumption. The coefficient turns positive and statistically significant during the transition period (1968-1972), when provinces with a larger wage gap experienced a proportionally steeper increase in the minimum wage level. By 1972, a province that in 1968 had a gap of 10% with respect to the minimum wage in Milan see early school leaving in excess of 5%. For reference, one quarter of the provinces had a wage gap larger than 10% with respect to Milan in 1968.

By 1972 the spatial convergence in nominal minimum wages was achieved, so our treatment switches off. Nonetheless, the estimated coefficient continues to grow until 1976. Why would the response continue to increase after the treatment switched off? A plausible explanation attributes this evolution to the fact that our measure of enrolment averages across five birth cohorts every year (the expected duration of most secondary school courses). This means that the cumulative opportunity cost of enrolling in secondary school varies between cohorts within each year. It is possible that the minimum wage hike that started in 1969 was steep enough to immediately discourage enrolment among the younger cohorts (those turning 14 in 1969), but not enough to cause similarly large dropouts in the older cohorts (those approaching 18 in 1969). Hence, the coefficients for the early years would underestimate the true effect. In fact, the first year for which all

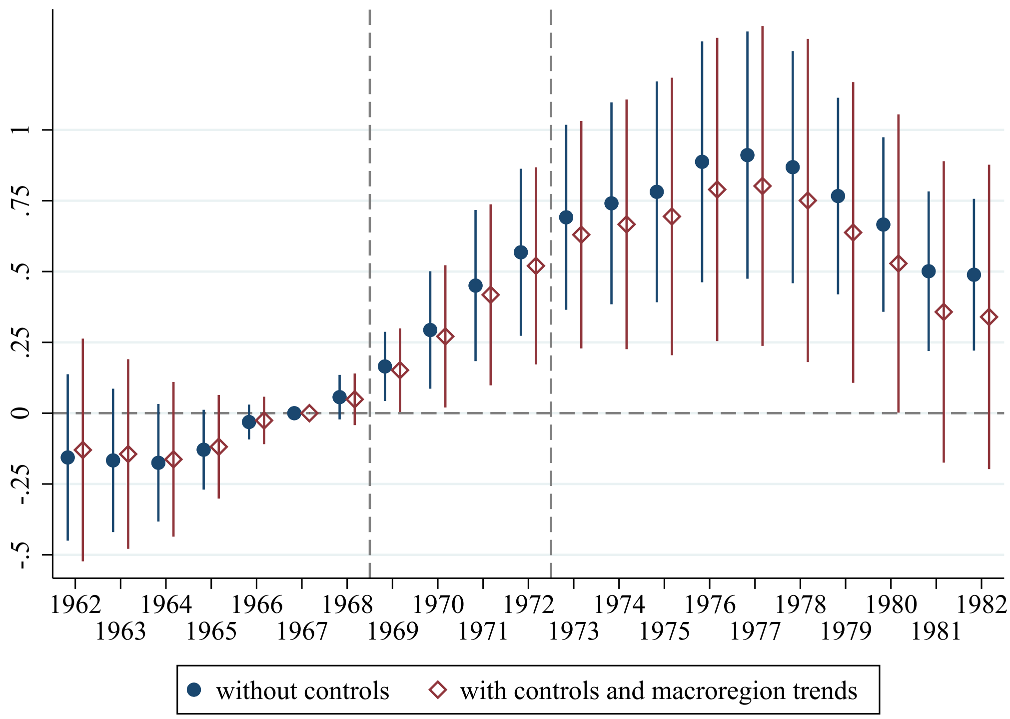


Figure 12: Dynamic response of early school leavers

OLS estimates of the coefficients for the interaction term between the log difference between the mean minimum wage in Milan and in the province in 1968 and year indicators, with 1967 set to zero. The dependent variable is the log of the number of individuals between the age of 14 and 18 not enrolled in post-compulsory upper secondary education. The regression controls for the size of the 14-18 age cohort in every province-year cell (in log). Both specifications control for time and province fixed effects. The second specification controls also for pre-1968 averages of total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector, each interacted with a linear trend, as well as macroregion dummies interacted with a linear trend. The vertical lines indicate the 95% confidence interval with standard errors that are clustered at the province level. The vertical dashed line indicate the start (1968) and the end (1972) of the convergence period, following the repeal of the wage zones in 1968. Number of observations: 1575. The estimation excludes provinces in special statute regions, so the total number of provinces is 75 (minimum wage data for the provinces of Arezzo and Ancona are missing from some sources).

individuals had turned 14 after the start of the wage hike is 1974. Moreover, the peak year of 1976 is the final year in which all individuals had turned 14 before the switching off of the treatment in 1972. Starting in 1977, we begin observing the response of the cohorts that turned 14 after 1972, i.e. that made their educational choices after the end of the differential wage shock.

Assuming that the coefficient in 1976 recovers the 'true' average causal response across the treated cohorts, our estimate implies that a 1% larger minimum wage gap with respect to Milan in 1968 is associated with a 1.1% increase in early school living (0.86% with controls). To understand what this means in terms of the impact of the minimum wage on enrolment, we recall from the first stage of the IV approach that the coefficient obtained from regressing the minimum wage on the wage gap was circa 0.85 in the same period. This yields indirect least squares of about 1.29 (1.01 with controls), which suggests that enrolment responded fairly elastically to the minimum wage shock of 1969-1972.

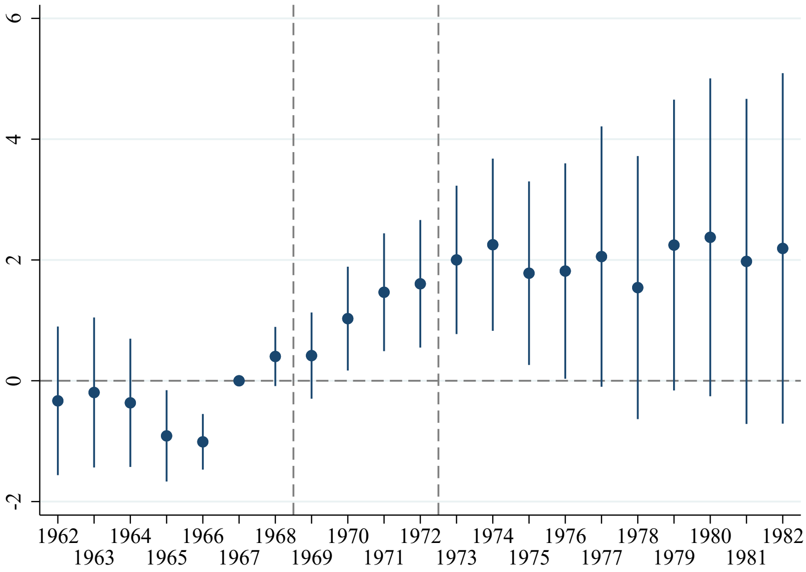
This large effect, however, did not carry over to the cohorts that turned 14 after the switch-off of the treatment, as shown by our coefficients' tendency to revert to the mean after 1976. By 1982, provinces that had experienced a steeper increase in the minimum wage due to their greater starting gap with respect to Milan showed no sign of extra early school leavers. It should be noted that the compression of male enrolment as identified by the aggregate descriptive statistics also lasted about six years before starting to recover (cf. Figure 6), albeit with a short lag with respect to our estimates.

5.2.3 Youth unemployment and the recovery of school enrolment

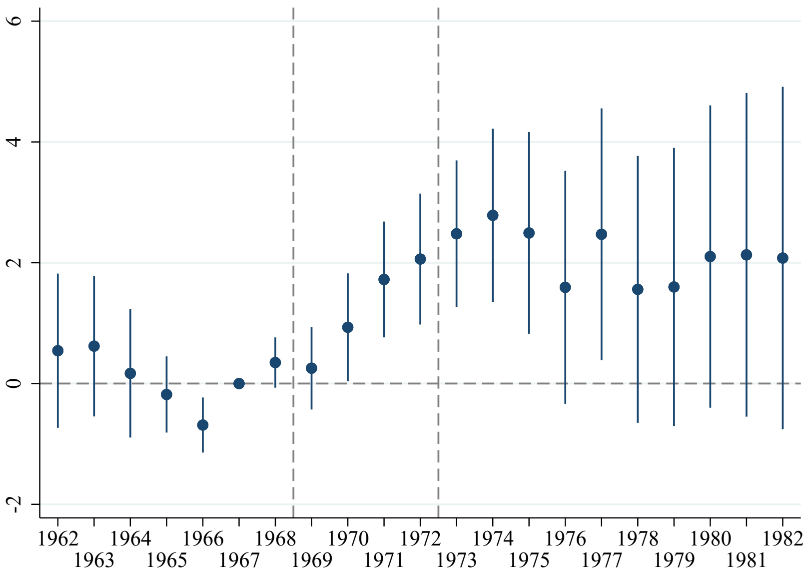
The previous analysis has shown that the egalitarian wage hike provoked a temporary increase in early school leavers, and we suggested that the disappearance of the effect after 1976 is due to the substitution of the treated cohorts (teenagers who decided to enroll during the 1968-1972 period) with untreated cohorts (those that turned 14 after 1972). However, an alternative interpretation might point to the disemployment effect of the minimum wage hike. This is a plausible explanation because, for such a steep wage hike, we would expect to find some impact on youth employment—especially if the labour demand did not adjust to the extra supply of early school leavers. In order to test this alternative hypothesis, I estimate the causal response of youth unemployment using the preferred generalized DiD model. The estimate is performed separately by sex and, for each sex, by whether the unemployed individual was a first job seeker or had been previously employed.

Figure 13 presents the results for men under 21. Both groups show no pre-trends and a significant increase during the transition period. The increase is comparable between the two groups, although the estimates lose statistical significance after 1976. While this result appears counter-intuitive given the aggregate trends in relative unemployment, it would explain why we observe an increase in early school leaving during the transition period: while youth unemployment was rising across all provinces, those that experienced a steeper increase in the minimum wage hike did not disproportionally penalize first-job seekers. Thus, it is possible that, in the short run, opportunity cost considerations prevailed over the rising risk of unemployment. This coincides with that of early school leavers, and it is possible that the two dynamics are connected, for a larger number of people in school would reduce the number of individuals searching for their first job. Hence, it is possible that, in the long run, the risk of unemployment prevailed over opportunity cost considerations for marginal students, who opted to stay in school. On the other hand, the permanent increase in unemployed young people with previous work experience could be explained by the cumulative stock of early school leavers that had accumulated over the transition period, and by general equilibrium effects that we cannot account for, such as lower job creation in the provinces which experienced the steeper increase in minimum wages.

Figure 14 shows the same analysis for women under 21. Both groups show a similar reaction to the male counterparts during the transition period. However, the coefficients quickly revert back to the mean after 1974 for young women previously employed, so we cannot reject the null hypothesis that there was no association between the deviation from Milan's nominal wage in 1968 and the number of young women registered as unemployed after this period. This is less the case for first job seekers, although the estimates lose statistical significance towards the end of the period. These results raise two interesting points. First, young women were equally responsive to the wage hike than men, which is coherent with the rising female labour force participation among young cohorts in this period (Reyneri, 1996, p. 115). Second, for previously employed women we do not find the same permanent effect that we found for men. Two interpretations seem most plausible: either the labour market was strongly segmented by gender—for instance, because women would more frequently find occupation in the service sector (Fullin and Reyneri, 2015)—, or women were more easily discourage when facing the prospect of



(a) previously employed



(b) first job seekers

Figure 13: The response of male youth unemployment

OLS estimates of the coefficients for the interaction term between the log difference between the mean minimum wage in Milan and in the province in 1968 and year indicators, with 1967 set to zero. The regression controls for the size of the 15-21 year cohort (males), for pre-1968 averages of total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector, each interacted with a linear trend, as well as macroregion dummies interacted with a linear trend. The vertical solid lines indicate the 95% confidence interval with standard errors that are clustered at the province level. The vertical dashed line indicate the start (1968) and the end (1972) of the convergence period, following the repeal of the wage zones in 1968. The number of observations is 1,890 and the number of provinces 90 (minimum wage data for the provinces of Arezzo and Ancona are missing from some sources).

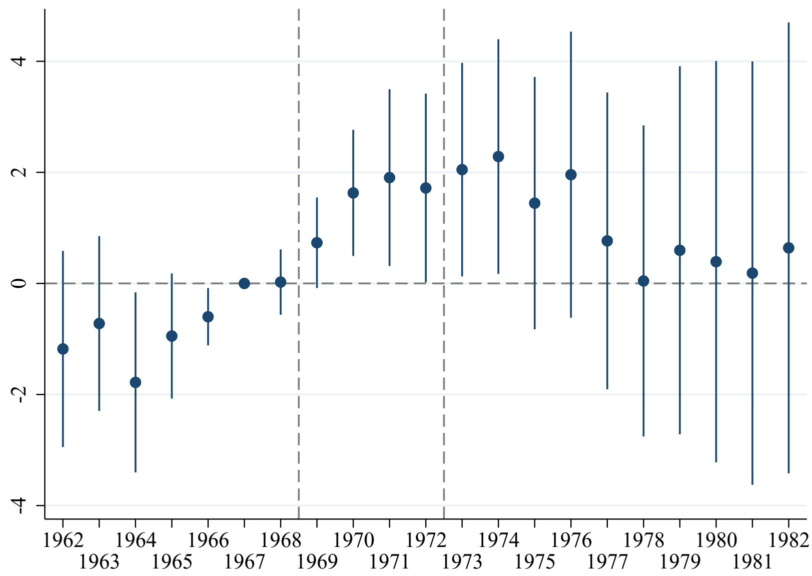
unemployment. Distinguishing between these two interpretations is not possible with our data, but the results suggest an avenue for future research.

5.3 The effect on the choice of school field

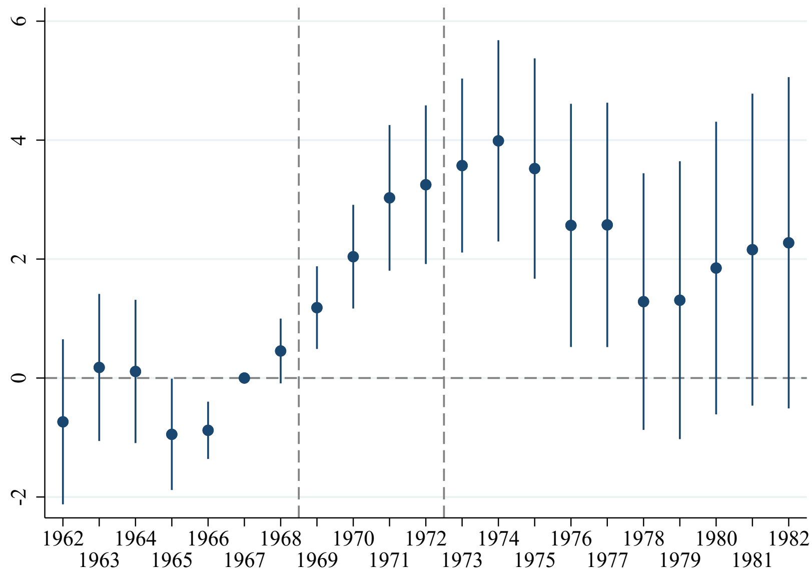
This analysis has shown that the minimum wage hike caused by the egalitarian collective agreements affected educational investment on the extensive margin, as higher mean minimum wages were associated with a significant increase in early school leaving. This, however, does not exclude that the egalitarian wage hike also affected investment on the intensive margin by provoking a shift in the composition of school enrolment through its effect on the relative returns to specialist education. This section tests this second hypothesis by looking at alternative school choices.

To test whether the egalitarian wage hike modified the composition of school enrolment, I distinguish between the choice of track and the choice of curriculum. I focus specifically on two alternative tracks—technical and professional schools—and two alternative curricula—schools for manufacturing and schools for business—which were chosen by teenagers that sought a practical education with immediate use in the private sector, for this group would be the most sensitive to changes in the *ex-ante* return to education.

Both the professional and the technical track were vocationally oriented, but the former was focused on application and gave students the option of leaving with an intermediate diploma after three years (age 16) instead of the customary five (age 18), while the latter had more theoretical components and only offered five-year courses. Hence, the choice between tracks reveals students' preferences with respect to the intensive margin of educational investment: choosing the technical track implied a longer time to completion but potentially access to further education. If the egalitarian wage hike was strong enough to reduce the ex-ante return to education for all students, we would expect students to shift their demand from technical to professional schools. Hence, we would observe enrolment decreasing in the former and increasing in the latter. Alternatively, in the extreme case that the ex-ante (discounted) return to education dropped below the reservation wage for the median student, we would expect enrolment to decrease in both tracks. Either way,



(a) previously employed



(b) first job seekers

Figure 14: The response of female youth unemployment

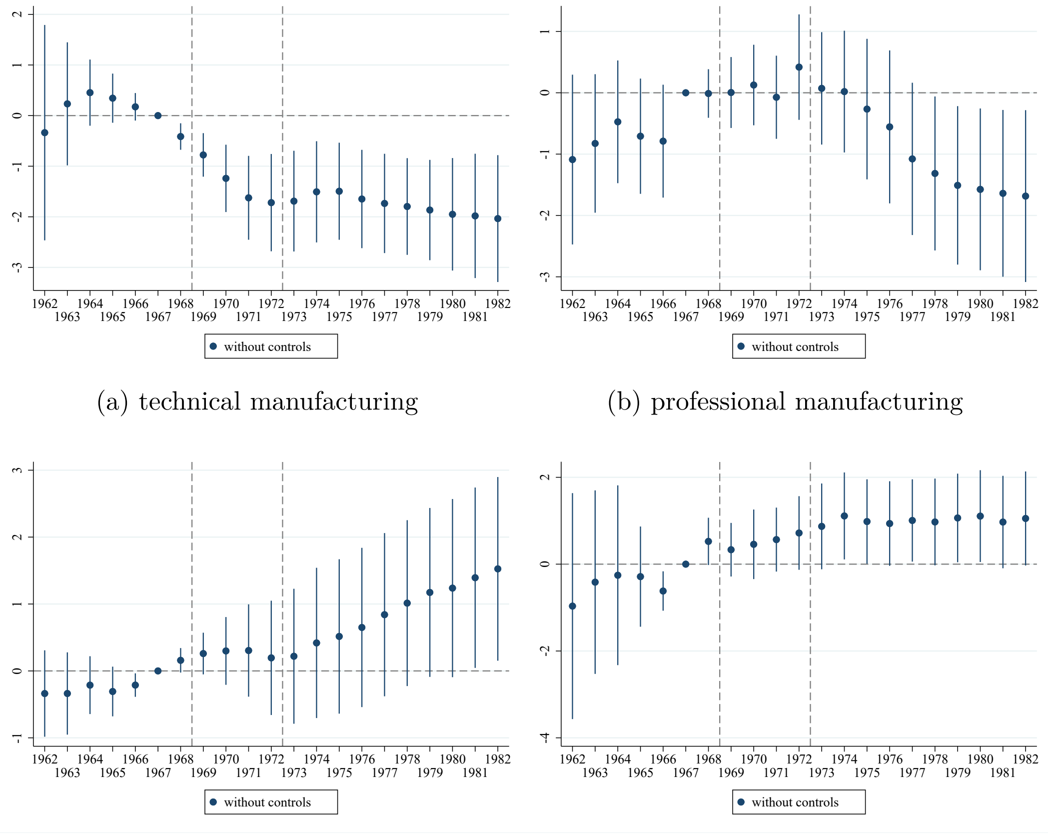
OLS estimates of the coefficients for the interaction term between the log difference between the mean minimum wage in Milan and in the province in 1968 and year indicators, with 1967 set to zero. The regression controls for the size of the 15-21 year cohort (females), for pre-1968 averages of total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector, each interacted with a linear trend, as well as macroregion dummies interacted with a linear trend. The vertical solid lines indicate the 95% confidence interval with standard errors that are clustered at the province level. The vertical dashed line indicate the start (1968) and the end (1972) of the convergence period, following the repeal of the wage zones in 1968. The number of observations is 1,890 and the number of provinces 90 (minimum wage data for the provinces of Arezzo and Ancona are missing from some sources).

results would inform us about changes in the quantity of education demanded.

The choice between curricula, instead, should reveal students' preferences with respect to the content of the specialist education offered by the schools. Schools for manufacturing provided specialist knowledge for high-skill blue-collar jobs across a range of industries, qualifying graduates to be employed as machine operators, maintenance workers, technicians and floor managers. Schools for 'business' (*commerciali*), instead, prepared for clerical jobs or white-collar professions that did not require a tertiary degree. Hence, the choice between the two curricula would be influenced by the relative return to specialist education for blue-collar workers. If the egalitarian wage hike decreased the ex-ante return to specialist education for blue-collar jobs more than for white-collar jobs, we would expect to see a shift from schools offering manufacturing curricula and to schools offering business curricula.

To test these hypotheses, I estimate the reduced form regression separately for each type of school and sex, using as dependent variable the log difference between the number of individuals enrolled in the relevant school field and the size of the 14-18 age cohort. Using equivalent definitions such as the log of the share enrolled or the log of the individuals enrolled after controlling for the size of the cohort does not produce significantly different results. The main regressor of interest is the interaction between the minimum wage gap with respect to Milan in 1968 and the time dummies. The coefficient of the interaction would recover the causal response an increase in the minimum wage gap, which proxies for an increase in the minimum wage after 1968, but not before.

Figure 15 shows the estimated interacted coefficients between the instrument and the time dummies for both male and students, separately by type of school. The first panel shows that enrolment in technical schools offering a manufacturing curriculum dropped after the repeal of the wage zones: starting around 1970, a 1% deviation from Milan's mean minimum wage in 1968 is associated with a decrease in the gross enrolment rate by about 2%. This effect is larger, in absolute values, than the general increase in early school leavers, which suggests that teenagers who would enrol in technical schools for manufacturing were disproportionally affected by the minimum wage hike. Moreover,



(a) technical manufacturing (b) professional manufacturing

(c) technical business (d) professional business

Figure 15: Dynamic response by school type (male and female)

Coefficients for the interaction term between the log difference between the mean minimum wage in Milan and in the province in 1968 and year indicators, with 1967 set to zero. OLS estimates controlling for the size of the 14-18 cohort and the number of people of the same age not enrolled in upper secondary education. The regression includes time and province fixed effects. The vertical lines indicate the 95% confidence interval with standard errors that are clustered at the province level. The vertical dashed line indicate the start (1968) and the end (1972) of the convergence period, following the repeal of the wage zones in 1968. The analysis exclude provinces in special statute regions, the provinces of Arezzo and Ancona (due to missing minimum wage data) and all provinces where the number of students was zero before 1968, for each type of school (i.e. the regressions excludes provinces without the school type before treatment).

unlike the estimates for early school leavers, the coefficient does not show a tendency to revert to the mean by the end of the period, hinting to a permanent effect.

So, did students shift their demand in favour of other types of specialist education in the long run? Panel 15c shows that there was no strong association between mean minimum contractual wages and male enrolment in technical schools for business either before or immediately after the repeal of the wage zones. However, the coefficients turn positive, large and statistically significant by the end of the period.This result would suggest that the loss of students in schools preparing for high-skill blue-collar jobs in manufacturing was partly compensated, in the long run, by an increase in schools preparing for white-collar jobs.

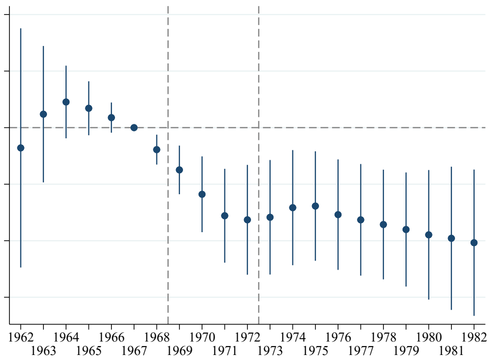
The delay between the reduction in enrolment in technical schools for manufacturing and the increase in enrolment in technical schools for business suggests that the two mechanisms discussed in the introduction might have acted in sequence: first, the raise in minimum wages increased the opportunity cost of schooling, pushing marginal students out of post-compulsory education. In the long-run, however, the compression of the skill premium for blue-collar workers might have prevailed, provoking a shift in the composition of educational demand from curricula preparing for manufacturing jobs and those preparing for blue-collar jobs.

This interpretation seems to be supported by the evolution of enrolment in professional schools. As mentioned above, these schools allowed to graduate in a shorter time, so the a relative preference for these schools implies a reduction in the investment in formal education on the intensive margin. Even though the estimates are imprecise, it appears in fact that enrolment in these schools did not decrease in the aftermath of the minimum wage hike and, possibly, increased somewhat during the repeal of the wage zones. Over the longer run, however, enrolment in professional schools for manufacturing jobs decisively decreased, while that in professional schools for business increased. This diverging evolution supports the hypothesis that concerns with respect to the relative return to education prevailed over the influence of the opportunity cost of staying in school for educational decisions.[28]

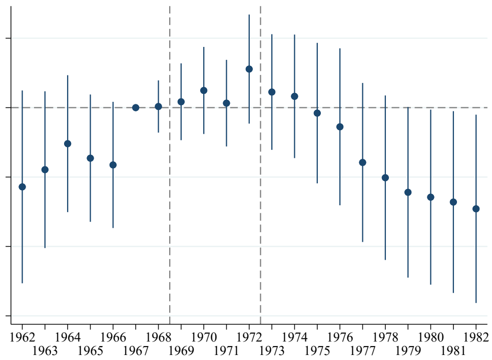
Figure 16 presents the results from the same estimation, but only for male students. The

[28] It is important to note, however, that the coefficients estimated for the vocational schools with a business curriculum is highly sensitive to the inclusion of pre-treatment trended controls and macroregion trends (not reported, available upon request). The inclusion of the trended controls significantly reduces the point estimates and does not allow us to reject the null hypothesis that the wage shock had no effect on enrollment in vocational schools for business (either technical or professional). At further inspection, it appears that this sensitivity is mostly due to the inclusion of the provinces with a mean minimum wage larger than Milan's in 1968. Excluding these provinces or setting the difference with Milan's minimum wage equal to zero yields positive coefficients in the post-treatment period, similarly to the uncontrolled regression, albeit with smaller magnitude and larger confidence intervals. Further research will focus on understanding the root causes of this issue. This note applies also to the following regressions, which estimate male and female enrollment separately by type of school. Note that the inclusion of the controls does not affect the point estimates and interpretation of enrollment in vocational schools with a manufacturing curriculum (either technical or professional).

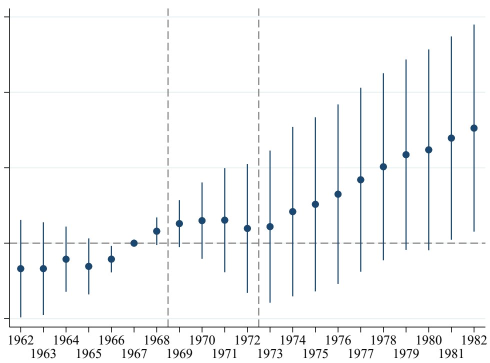
dynamic response for technical and professional schools for manufacturing is essentially identical to that for the full sample—which is to be expected, considering that men represented the vast majority of students choosing these school fields. The panels 16c and 16d, instead, show an even larger shift in favour of schools preparing for white-collar jobs than the previous estimates, suggesting that male students were more responsive to changes in the relative return to specialist education than aggregate statistics might lead to believe.



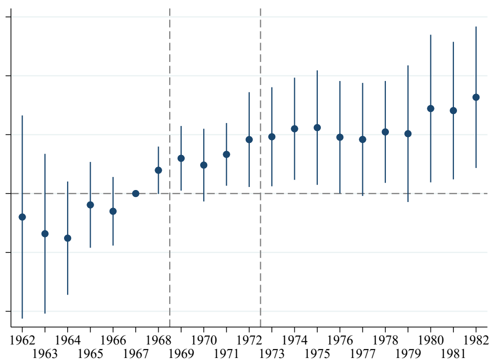
(a) technical manufacturing



(b) professional manufacturing



(c) technical business



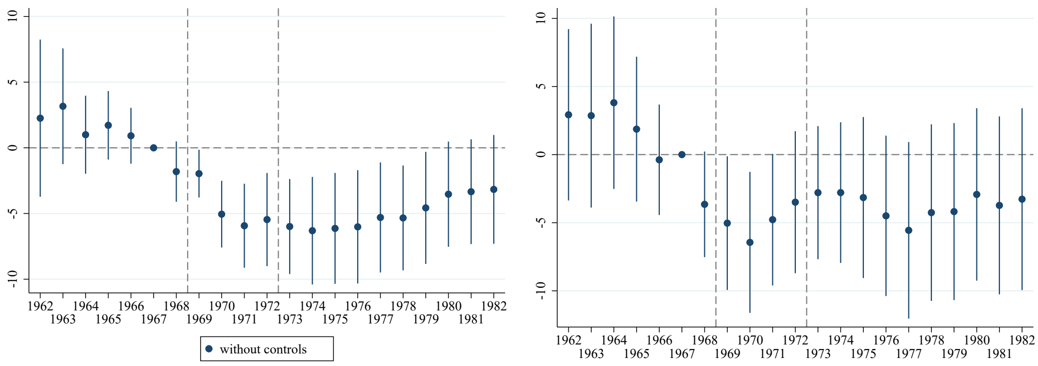
(d) professional business

Figure 16: Dynamic response by school type (male only)

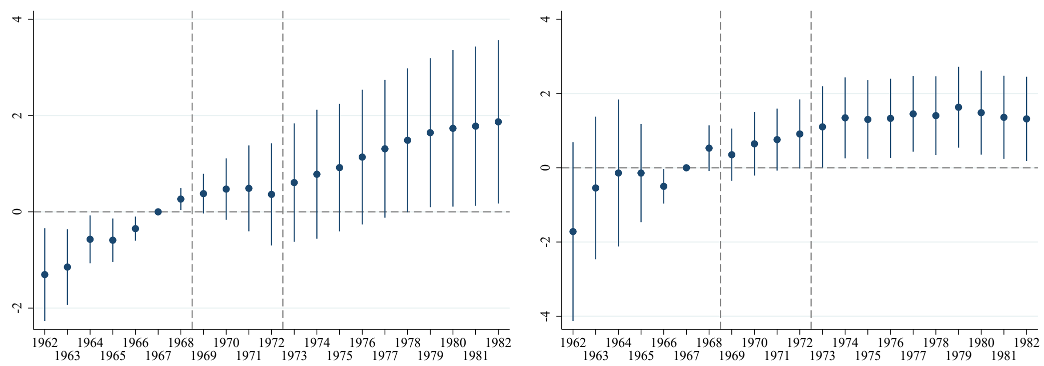
Coefficients for the interaction term between the log difference between the mean minimum wage in Milan and in the province in 1968 and year indicators, with 1967 set to zero. OLS estimates controlling for the size of the 14-18 cohort (male only) and the number of people of the same age not enrolled in upper secondary education (total). The regression includes time and province fixed effects. The vertical lines indicate the 95% confidence interval with standard errors that are clustered at the province level. The vertical dashed line indicate the start (1968) and the end (1972) of the convergence period, following the repeal of the wage zones in 1968. The analysis exclude provinces in special statute regions, the provinces of Arezzo and Ancona (due to missing minimum wage data) and all provinces where the number of students was zero before 1968, for each type of school (i.e. the regressions excludes provinces without the school type before treatment).

Figure 17 repeats the analysis for female students. It should be noted that very few female teenagers enrolled in schools with a manufacturing curriculum in the first place, so the precision of the estimates is affected by the small cross-sectional and longitudinal variation of the dependent variables.[29] Nonetheless, the exercise is useful because, if the results point to the same direction as for their male classmates, it would add evidence to the argument that the effect identified can be causally attributed to the egalitarian wage hike and not to unobservable factors that only affected male enrolment in secondary school. The figure shows that the response of female students was indeed similar to that seen before for men: technical and professional schools for manufacturing show a negative response after the abolition of the wage zones (with a tendency to mean reversion for the former); the coefficients for technical and professional schools for business turn positive in the long run (even though we detect some pre-trends in this specific case).

[29] Across all provinces and years, the average number of female students choosing manufacturing curricula is 78 for professional schools and 98 for technical schools, against 2,747 and 1,288 for male, respectively.



(a) technical manufacturing (b) professional manufacturing



(c) technical business (d) professional business

Figure 17: Dynamic response by school type (female only)

Coefficients for the interaction term between the log difference between the mean minimum wage in Milan and in the province in 1968 and year indicators, with 1967 set to zero. OLS estimates controlling for the size of the 14-18 cohort (female) and the number of people of the same age not enrolled in upper secondary education (total). The regression includes time and province fixed effects. The vertical lines indicate the 95% confidence interval with standard errors that are clustered at the province level. The vertical dashed line indicate the start (1968) and the end (1972) of the convergence period, following the repeal of the wage zones in 1968. The analysis exclude provinces in special statute regions, the provinces of Arezzo and Ancona (due to missing minimum wage data) and all provinces where the number of students was zero before 1968, for each type of school (i.e. the regressions excludes provinces without the school type before treatment).

5.4 Implications for the human capital stock

The empirical analysis has shown that the steep rise of contractual wage floors caused by egalitarian collective bargaining affected teenagers' educational decisions. In particular, the minimum wage hike temporarily reduced enrolment rates in upper secondary education and permanently decreased enrolment in technical and vocational schools that prepared for blue-collar jobs in the manufacturing sector. But just how large was the impact on the accumulation of human capital in the long run? To quantify the impact on the stock of secondary school graduates we can compare the number of individuals holding at least a secondary education diploma with a counterfactual estimate, that is the number of individuals that would have graduated if the dip in enrolment rates had not materialized.[30]

For this analysis I use the 'Historical Database' (version 10.1) of the Bank of Italy's Survey on Household Income and Wealth (SHIW), which contains cross-sectional microdata from all the surveys that were conducted annually from 1977 to 1987 (excluding 1985) and every other year from 1989 to 2016.[31] I exclude the waves from 1977-1983 because the individuals' age is estimated and only takes four possible values. Among other variables, the database provides information on the highest level of education attained by the surveyed individuals. The SHIW defines six levels of education: none, primary school, lower secondary school, upper secondary school, university degree and post-graduate degree. An individual is classified in the uppermost possible category, conditional on having completed the relevant educational cycle. Hence, the SHIW tends to underestimate years of education because it does not account for the number of years spent in school before dropping out. Information on the level of education is available for 374,755 individuals across all survey waves, that is 85% of all observations. Hence, I exclude observations with missing values from the analysis. Additionally, I drop all individuals younger than 26, to account for late graduates and cohorts that are still in education at the time of the

[30]The use of educational attainment rather than enrolment rates, as in the previous section, is justified by the fact that the same dynamics can be observed for both gross enrolment rates and gross graduation rates (see Figure 5), and that educational attainment can be observed also at a later stage in life, allowing us to perform estimates for each birth cohort based on individual-level data.

[31]The survey has been conducted since the 1960s on a representative sample of Italian households and is available for download from the Bank of Italy's website: https://www.bancaditalia.it/statistiche/tematiche/indagini-famiglie-imprese/bilanci-famiglie/index.html.

survey.

To estimate the educational attainment for each cohort, I compute the share of individuals with an upper secondary school diploma or higher in each cohort, accounting for the relevant population sampling weights. Figure A.5 shows the resulting graduate-cohort profile, distinguishing also between men and women. These estimates confirm the dynamics depicted by the educational statistics and comparable with alternative estimates that can be computed on data from population censuses (see Figure 5 and Figure A.6). All statistics confirm that the generation born after 1955 and before 1970 (theoretically to be enrolled in upper secondary school between 1970 and 1985) contributed the least to the expansion of secondary school enrolment throughout the period 1930-1990.

To compute a plausible counterfactual without the pause of the 1970s, I first estimate the absolute number of secondary school graduates in each cohort, multiplying the share of graduates by the size of the cohort at age 18, which I obtain from my intercensual reconstructions (see appendix in Ramazzotti (2023) for sources and methodology). Then, I estimate the number of additional individuals that would have graduated if growth had continued following the trend of the cohorts 1930 to 1954. For this purpose, I regress the graduation rate on the year of enrolment, under the condition that the year of enrolment is between 1945 and 1969. Hence, I use the fitted values of the regression as a counterfactual estimate of the gross graduation rate for subsequent age groups. To compute the number of graduates in the counterfactual scenario, I multiply the predicted gross graduation rate by the size of the age group at 18 for the cohorts born after 1954.

Figure 18 shows the number of graduates estimates for each cohort and the counterfactual scenario, together with the size of each the cohorts. The cumulative net loss until the 1990 cohort is over 2.3 million graduates. A back-of-the-envelope calculation finds that, had secondary education continued to expand at the same rate, the share of individuals aged 25 to 64 with at least a diploma of upper secondary education in 2015 would have increased from 60% to 67%. This is possibly an upper bound estimate, because it assumes that, without the dip in enrolment of the 1970s-1980s, graduation rates would have grown at the same trend as in the 1950s and in the 1960s, the time of fastest expansion of secondary school. To account for a plausible natural decrease in the rate of growth, we can repeat the same computations using the trend after the pause (the cohorts born after 1970). In this case, the cumulative net loss is estimated at 1.3 million graduates. The additional graduates would increase the share of people between 25 and 64 with an upper secondary school diploma or higher in 2015 to 64%. Moreover, it appears that the missing graduates can be mostly attributed to the specific dynamics of male educational attainment, rather than factors that were common to both sexes: 97% of the missing graduates from the more conservative counterfactual can be entirely attributed to men's sagging attainment.

But by just how much did this diversion from the trend matter for Italy's current lag in its human capital stock? Considering that, in 2015, the average share of upper secondary graduates in OECD countries was 76% of the population aged between 25 and 64 *vis-à-vis* 60% in Italy, our counterfactuals would see the gap decrease by between 9 and 16 percentage points. In other words, the pause in the expansion of secondary school enrolment between the 1970s and the 1980s would explain between 24% and 44% of Italy's lag in 2015.

6 Conclusions

Collective agreements are labour market institutions that, like statutory minimum wages, regulate entry-level salaries and influence the wage distribution. However, their relatively large bite with respect to the median wage can significantly raise the opportunity cost of investing in formal education for marginal students, and their egalitarian influence can alter the ex-ante return to education for inframarginal students. Thus, collective agreements that set high contractual minimum wage floors can reduce the accumulation of formal human capital and the relative supply of specialist knowledge, causing skill mismatch and lower growth potential in the long run.

The paper has explored these implications studying a historical natural experiment from Italy between the 1960s and the 1980s. Reconstructing new series of contractual and

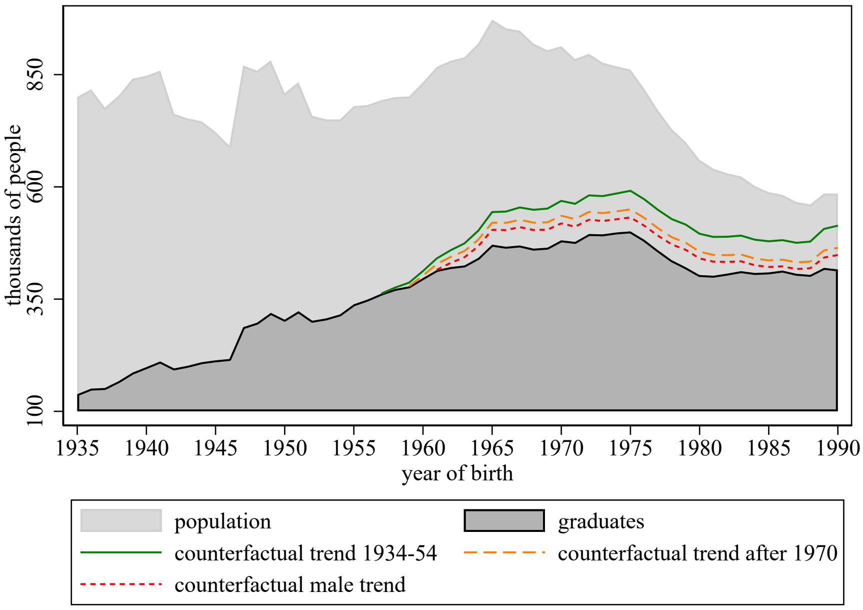


Figure 18: Upper secondary school graduates and counterfactual estimates

Population size at age 18 and number of upper secondary school graduates by birth cohort, with counterfactual estimates. Number of graduates computed by multiplying the share the smoothed graduation rate in each birth cohort by the size of the cohort at age 18. The smoothed graduation rate is computed as the trend component from a Hodrick-Prescott filter applied to the share of individuals with an upper secondary diploma or higher in the SHIW surveyed sample, for each birth cohort. The HP filter is applied with a smoothing parameter of 6.25 (cf. Ravn and Uhlig, 2002). The counterfactual trend 1934-1954 shows the total number of graduates in case educational attainment of the birth cohorts born after 1954 had expanded at the same rate as for the cohorts born in 1934-54. The counterfactual trend after 1970 shows the total number of graduates if attainment for the 1954-69 cohorts had expanded following the same trend as the cohorts born after 1970. The counterfactual male trend show the total number of graduates had the attainment of the male cohorts born after 1954 followed the contemporary female trend. Counterfactual trends are the predicted values from linearly regressing the smoothed number of graduates on a time trend, restricted to the relevant birth cohorts (female only for the third counterfactual), and multiplying the predicted enrolment rates by the size of each birth cohort at age 18. Source: own estimates on microdata from the Bank of Italy's Survey on Household Income and Wealth (SHIW), Historical Database, version 10.1, waves 1984-2016 pooled together. Year of birth computed subtracting the individual's age from the year of the survey. Individuals born before 1900 are excluded from all waves, as well as individuals younger than 26 in each wave. Before the 1989 wave, only the educational level of income earners was recorded. Upper secondary school is considered attained if the educational qualification is upper secondary school (*medie superiori*), graduate degree (*laurea*) or post-graduate degree (*specializzazione post-laurea*). Total sample size: 245,116 observations. Data available for download at `https://www.bancaditalia.it/statistiche/tematiche/indagini-famiglie-imprese/bilanci-famiglie/distribuzione-microdati/index.html` (last retrieved October 2021). Size of cohorts at age 18 is obtained from the official reconstruction of the national population on January 1st of each year by age group since 1952, available from Istat's *I.Stat* datawarehouse at `http://dati.istat.it/Index.aspx?DataSetCode=DCIS_RICPOPRES1971` (for 1952-1972), `http://dati.istat.it/Index.aspx?DataSetCode=DCIS_RICPOPRES1981` (for 1972-1981) and, `http://dati.istat.it/Index.aspx?DataSetCode=DCIS_RICPOPRES1991` (for 1982-1991), last retrieved October 2021.

effective wages across the manufacturing sector, the paper has shown that labour unions' shift in favour of egalitarian bargaining in 1969 provoked a steep increase in entry-level minimum wages and a compression of the skill premium for blue-collar workers. New estimates on educational data have also shown that the wage hike was accompanied by a dip in male enrolment in upper secondary school and by a shift in the composition of curricula chosen by those who stayed in education.

The paper has hypothesized that the two phenomena are linked. In particular, two mechanisms have been proposed: first, that the increase in minimum contractual wages motivated marginal students to leave post-compulsory school early, either by dropping out entirely or by choosing tracks that offered shorter courses. Second, the compression of wage differentials for blue-collar workers incentivized inframarginal students to shift away from specialist curricula providing skills for manufacturing jobs. While the first effect was only temporary—possibly due to the rising risk of not finding a job at the higher minimum wage—the second effect was permanent.

The paper has also argued that, despite its temporary nature, the negative impact on enrolment rates continues to affect Italy's ranking in educational attainment to this day. Depending on the counterfactual scenario, the pause in enrolment cost between one and two million missing graduates, which would have reduced the distance in educational attainment between Italy and the OECD average by at least 25%. These empirical results support our hypothesis that the steep egalitarian rise of contractual minimum wages modified the incentive structure for young Italians, affecting their post-compulsory educational choices. These results also reinforce our argument that empirical analyses of the impact of minimum wages—either statutory or bargained—on schooling should take into account not only the effect on drop out rates, but also its potential impact on the fields of education chosen by the students and its eventual implications for labour market outcomes and an economy's growth potential in the long run.

Additional materials

A Additional figures

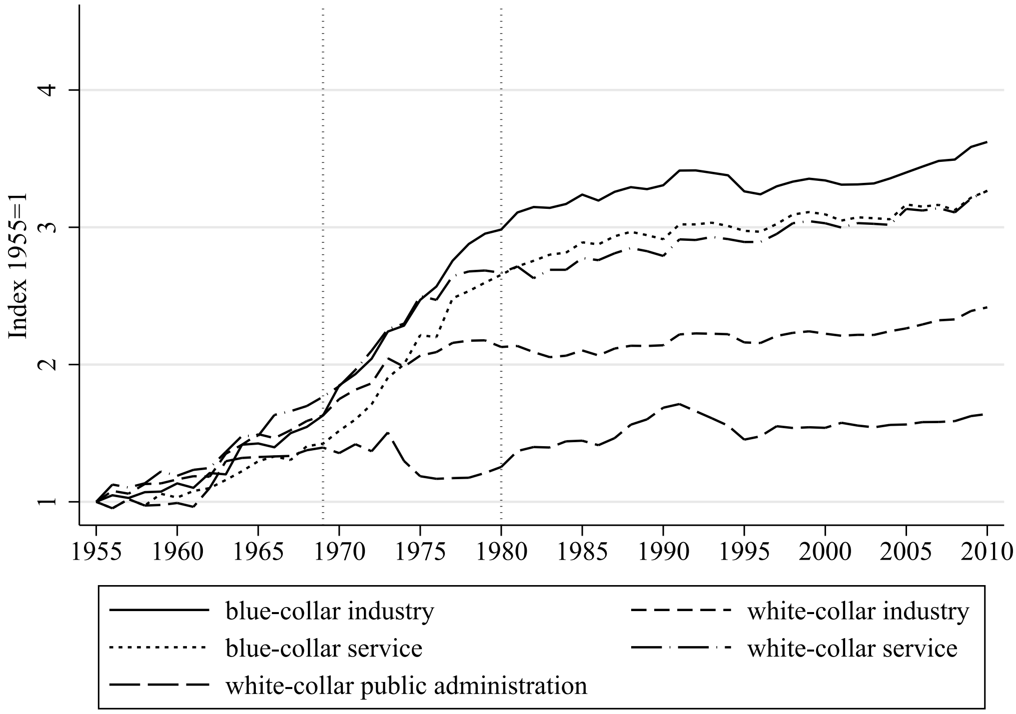


Figure A.1: Changes in wage floors by worker category and macro-sector

Index of contractual hourly wages by economic macro-sector and worker category, rescaled with 1955 as base year. Vertical dotted lines indicate the years 1969 (the Hot Autumn) and 1980 (year of the 'march of the forty thousand'), which identify the period of the egalitarian wage push. The 'industry' series includes mining, food, textile, metal and engineering, chemical, construction and electricity. These official series differ from the authors' reconstructions as they average across skill levels within worker categories. Source: own elaborations on data from Istat, *Serie Storiche*, Tav. 10.21, available at https://seriestoriche.istat.it/, last retrieved June 2022.

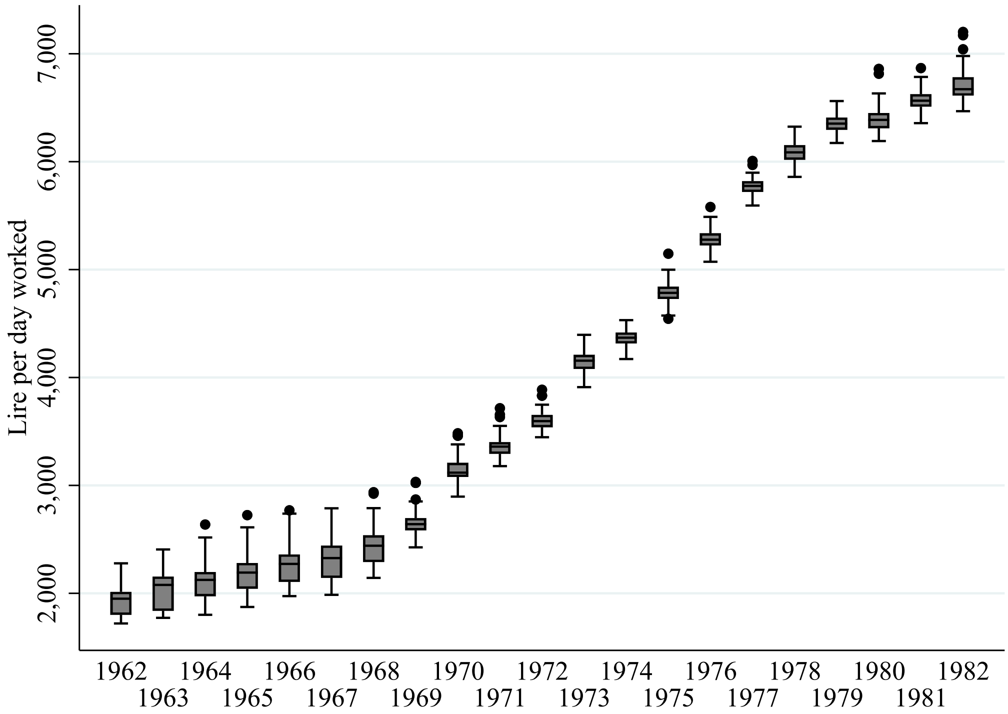


Figure A.2: Box plot of mean minimum wages at 1968 constant prices

Box plot of mean minimum wages for low-skill blue-collar workers in 24 industrial sectors across ninety-two provinces. The box indicates the interquartile range, the line indicates the median, the whiskers connect to the adjacent values and the markers indicate outside values. Sectoral minima are weighted using the estimated number of employees in each sector-province cell, obtained as the linear interpolation from decennial industrial censuses. Details on sources and estimation strategy are provided in the methodology appendix (available upon request). The nominal value of the minimum wage is originally expressed in current Italian lire per hour worked, converted to daily wages multiplying by eight and converted to 1968 prices using official coefficients from Istat, *Il valore della moneta in Italia dal 1861 al 2020*, available for download at https://www.istat.it/it/archivio/258610 (last retrieved July 2022).

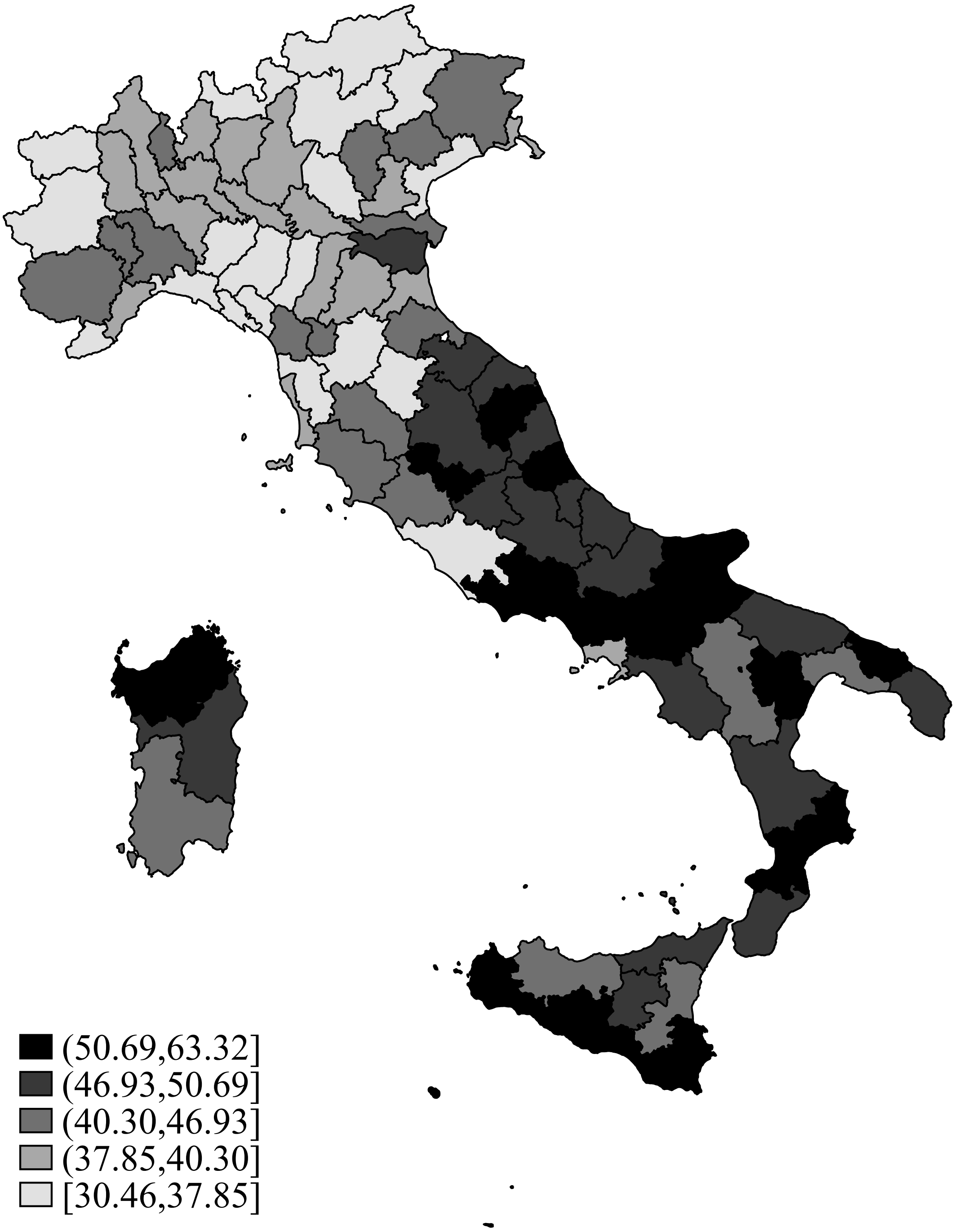


Figure A.3: CHANGE IN MEAN INDUSTRIAL MINIMUM WAGE 1968-1972

The map shows the percentage change in the mean nominal minimum wage across nineteen industrial sectors between 1968 and 1972. The change is computed at constant 1968 prices. Sectoral wages are weighted according to local industry shares, in each province. For additional details see methodology appendix (available upon request). The shapefile of the provinces at 1961 historical borders is available at https://www.istat.it/it/archivio/231601 (last retrieved July 2022).

Figure A.4: Minimum wage differentials with respect to Milan

Log difference of the provincial mean minimum wage with respect to Milan. The mean minimum wage is computed as the weighted average of the minimum wage for the lowest category of blue-collar worker across twenty industrial sectors. Weights are obtained from the industry shares of employees in the province. The median, 25th percentile and 75th percentile are the respective values of the difference with respect to Milan, for all remaining 91 provinces. For data sources see text and methodology appendix (available upon request).

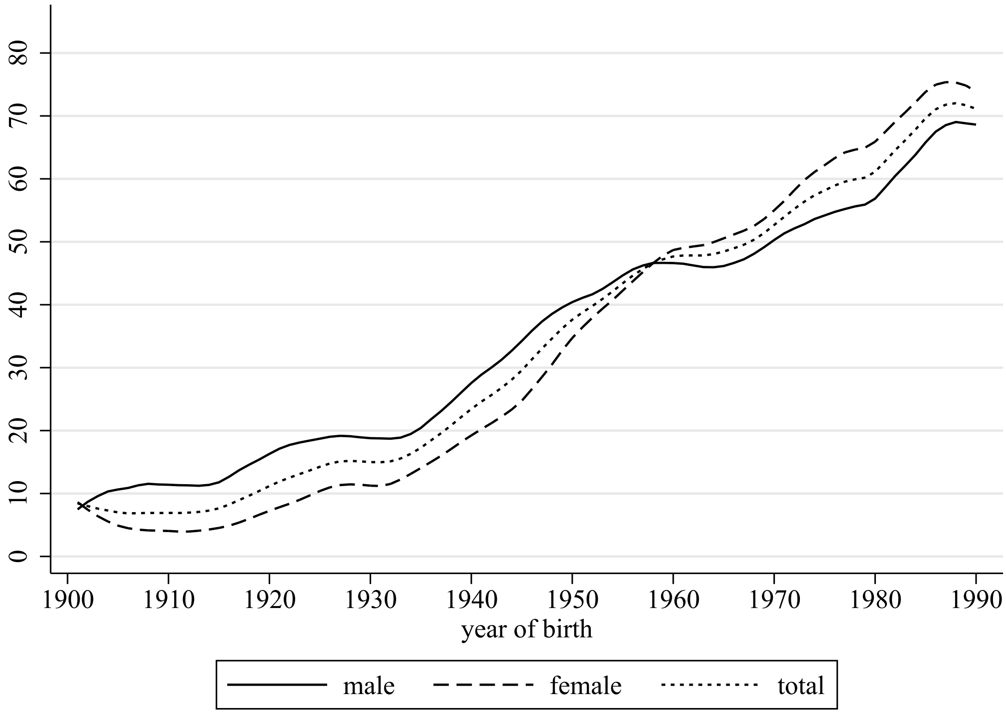


Figure A.5: Upper secondary school attainment by birth cohort

Percentage of individuals with a secondary school diploma or higher by birth cohort. Trend component from a Hodrick-Prescott filter with a smoothing parameter of 6.25 to account for the annual frequency of the data (cf. Ravn and Uhlig, 2002). Source: own estimates on microdata from the Bank of Italy's Survey on Household Income and Wealth (SHIW), Historical Database, version 10.1, waves 1984-2016 pooled together. Year of birth computed subtracting the individual's age from the year of the survey. Individuals born before 1900 are excluded from all waves, as well as individuals younger than 26 in each wave. Before the 1989, only the educational level of income earners is recorded. Upper secondary school is considered attained if the educational qualification is upper secondary school (*medie superiori*), graduate degree (*laurea*) or postgraduate degree (*specializzazione post-laurea*). Total sample size: 245,116 observations. Data available for download at `https://www.bancaditalia.it/statistiche/tematiche/indagini-famiglie-imprese/bilanci-famiglie/distribuzione-microdati/index.html`

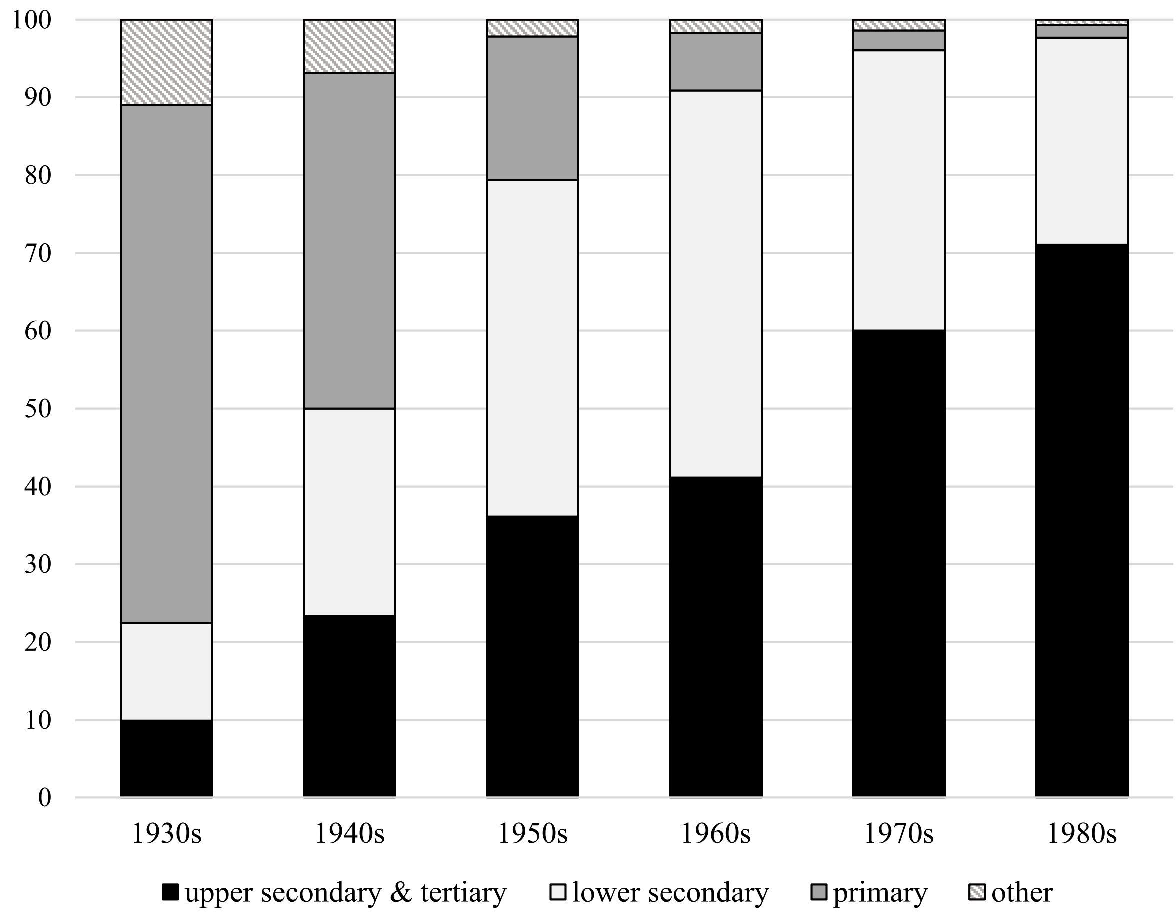


Figure A.6: Educational attainment by decade of birth

Percentage of population by highest level of education attained and decade of birth. Education attained is identified by the highest leaving school qualification declared at age 19-29. Tertiary education is included with upper secondary education to avoid underestimation for individuals born in the second half of each decade, who would still be enrolled in university at the time of the census. The 'other' category includes illiterate individuals and literate individuals without school leaving qualifications. Source: own computations on population censuses of 1961, 1971, 1981, 1991, 2001 and 2011—respectively, Istituto Centrale di Statistica (1975, pp. 32-125), Istituto Centrale di Statistica (1984), Istituto Nazionale di Statistica (1994), Istat, *DaWinci.istat.it* available at http://dawinci.istat.it/ (last retrieved June 2022), and Istat, *Censimento Popolazione Abitazioni*, available at http://dati-censimentopopolazione.istat.it/Index.aspx?lang=it (last retrieved June 2022).

B Additional tables

Table B.1: Conversion table between census, minimum wage and INAIL codes

N	INAIL definition	min wage	census
1	Industrial food processing	5	3010
2	Chemicals, rubber, plastics	2, 3,4	3070, 3131, 3080
	Paper & packaging, printing & publishing	10, 12	3140, 3151
	Leather & hide	18	3030
3	Construction	9	4010
4	Electricity & gas	6	5010
5	Wood & similar	16	3061, 3062
6	Metallurgy, metal carpentry,	15	3102, 3101
	machinery, transport vehicles, instruments	17	3111, 3112, 3113, 3115
7	Mining, mineralogy and complementary	11	2010, 2020
8	Textiles and clothing	1, 7, 19	3052, 3040, 3051
9	Trucking and warehousing	—	—
10	Other	8, 13, 14	3132, 3120, 3133

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