Wage Setting and School Enrollment: The Influence of Collective Agreements on Human Capital Accumulation in Italy, 1960s-1980s

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Abstract

Do wage-setting institutions influence human capital accumulation? This paper studies the impact of collective wage bargaining on school enrollment exploiting a quasi-natural historical experiment from Italy around 1969, when labour unions coordinated to bargain steep wage raises. Italy's weakly-selective educational system—whereby high-school students choose between specialist curricula at age fourteen—allows to separately identify the impact on enrollment rates from the substitution effect between alternative school tracks. Absent microdata for the period under study, I present original estimates of education and labour-market variables for ninety-two provinces with annual frequency between 1962 and 1982. Using an instrumental variable approach and flexible Difference-in-Differences with a continuous treatment variable, I find that the wage hike was associated with a temporary increase in early school leaving and a permanent substitution away from vocational schools preparing for manufacturing jobs. The length of the adjustment is found to explain a significant long-term loss in Italy's potential human capital stock.

Keywords: Educational attainment, wage setting, Italy

JEL classification: J24, J31, N34

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1 Introduction

Governments across Western countries are under pressure to address widening income inequalities, either by strengthening redistributive policies, enforcing statutory minimum wages, or promoting collective bargaining (OECD, 2019; European Commission, 2020). But altering relative wages 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 Wascher, 2007; Long, Goldhaber, and Huntington-Klein, 2015; Altonji, Arcidiacono, and Maurel, 2016). The resulting impact on human capital accumulation and skill mismatch could influence individuals' lifetime earnings and the economy's growth potential 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 collective agreements that are bargained at the sector level 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).

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

The paper focuses on the period between the 1960s and the 1970s. This period provides 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 raised the minimum industrial wage, on average, by 14% per year for a decade and reduced the skill premium of blue-collar workers by one third. I hypothesize that this egalitarian wage hike affected teenagers' post-compulsory education through two distinct mechanisms. 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 regarding the *ex-ante* return 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 industries. 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 average treatment effect of increasing the mean contractual minimum wage on school enrollment. I find that enrollment 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 0.3%-0.45%. The drop in enrollment appears particularly concentrated in vocational schools that prepare for skilled blue-collar jobs in the manufacturing sector.

Secondly, I use the spatial equalization of 1969-1972 as a natural experiment to study the dynamic response of school enrollment 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 disappeared. The length of the impact suggests that only cohorts that turned 14 during the wage hike were influenced by it.

Even though the response on in early school leaving was only temporary, it was economically significant. Our estimates can explain four fifths of the reduction in gross enrollment rates that is observed at the national level between the 1970s and the 1980s. The dynamic estimates show that the negative impact on school enrollment was quick but temporary, as enrollment rates reverted to the mean by the early 1980s. These results support the hypothesis that, by setting high entry-level minimum wages with respect to the wage distribution, 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. Enrollment in vocational schools preparing for skilled blue-collar jobs showed a negative response three times larger than the average secondary school. Moreover, enrollment 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. In the long-run, inframarginal students shifted their demand for education towards curricula that were not as affected by the egalitarian wage hike.

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 due to males' sagging enrollment. Counterfactual estimates find that the sag in male enrollment caused a substantial loss for Italy's potential human capital stock (between 1.3 and 2.3 million graduates), 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 enrollment 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; C. H. 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 enroll 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 proposes, for the first time, a detailed breakdown of enrollment 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; Ballarino, Bison, and Schadee, 2011; Ballarino, Panichella, and Triventi, 2014; Panichella, 2014; Ballarino and Panichella, 2016; 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 the employers' association. For each sector, the agreements established minimum wage floors according to the workers' skill level, which depended on the tasks performed on the job (Traversa, 1975).[2] These minimum wage floors represented a major component of the

[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).

workers' take-home pay, and firm-level agreements could only improve on their terms.[3] 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 (see Figure 1). 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

[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 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).

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 2 shows its evolution across twenty-four industrial sectors from 1962 to 1982, at constant prices (sources and harmonization procedures are described in detail in appendix A.2). 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,

[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.

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 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 3 shows that, on average, this skill premium decreased by 56% between 1968 and 1980.[6] The rigid structure of the centralized wage-setting system 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 3 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 4). Moreover, in the industry sector between 1968 and 1984 the skill premium for hig-skill white-collar workers decreased

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

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 enrollment 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.

2.3.1 Implications for post-compulsory school enrollment

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; C. H. Lee, 2020; Smith, 2021). Assuming that marginal students were equally distributed across

[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 4 and are available upon request.

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.) 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 enrollment cannot be anticipated, and we will need to control for youth unemployment throughout the analysis.[9]

[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, 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 enrollment to be proportionally greater in vocational schools.

[9] A formal test of the impact of the minimum wage hike on prime-age and teenage unemployment, by sex, is provided in a companion paper in progress (Ramazzotti, 2022). Preliminary results are available upon request.

2.3.2 Implications for the choice of school field

The second potential influence of egalitarian collective agreements on school enrollment 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.[10] 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. 6. 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.[11] Within each track, schools offered distinct curricula, which students could not mix or modify at will. Professional

[10] 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).

[11] 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.

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).[12] 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*).[13]

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 education to choose vocational schools for industry. *Ceteris paribus*, we would predict a shift in the composition of enrollment 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).[14]

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

[12]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.).

[13]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).

[14]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).

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

Figure 21 shows that enrollment 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.[15] 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. Enrollment 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

[15] Gross enrollment 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).

(A'Hearn and Vecchi, 2017).

What might be the proximate causes of this pause? By distinguishing enrollment rates according to the students' sex, Figure 5 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), enrollment 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 enrollment in post-compulsory secondary schools throughout the 1960s, with enrollment rates reaching 50% for males. However, starting in the early 1970s, the expansion of male enrollment 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 sagged through the following ten years: in 1985, the gross enrollment rate was about the same as in 1976 (56.4% and 56.7%, respectively). Female enrollment 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 enrollment resumed after 1985 at the same rate as the female, with enrollment 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 enrollment 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 enrollment rates for women and men, respectively, distinguishing between the top-six types of schools, by track and curriculum, and a residual 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 enrollment 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 one in four male pupils was enrolled in a technical school preparing for manufacturing jobs.[16] However, the expansion of technical schools with an industrial curriculum slowed down in the first half of the 1970s, and enrollment 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 enrollment 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 enrollment 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

[16] Including professional schools, the share of male students enrolled in upper secondary education preparing for manufacturing jobs reached 40% in 1970. Enrollment 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).

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 enrollment rate in technical schools for industry dropped by 25% between 1973 and 1982 (respectively, its peak and through). Female enrollment 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—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 enrollment levels, the bell-shaped trajectory of enrollment in technical schools for manufacturing was replicated across all of Italy's macroregions. Figure 8 shows that gross enrollment 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.[17] The contrast between the two periods is most accentuated in the North-Western provinces, but the decline in enrollment 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 enrollment expansion for both men and women in the 1970s, the sagging of enrollment 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

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

strategy to credibly identify the effect of the minimum wage hike on post-compulsory school enrollment by sex, track and curriculum.

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 representativity.

To circumvent these limitations, I present a new historical GIS dataset which contains information on enrollment rates, minimum and effective industrial wages and youth unemployment for Italy's ninety-two provinces from 1962 to 1982, with annual frequency. The data has been digitized and harmonized from a vast 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 (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 appendix A.

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 twenty-four 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).

The local industry shares are computed according to two alternative procedures, for greater robustness of the estimates: in one set of reconstructions, the local industry composition is given by 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 9, 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.

However, these local share might be endogenous, for firms might react to the minimum wage hike by laying off workers or freezing new hires. To control for this source of endogeneity bias, I compute alternative industry weights using only the national trends in sectoral employment. This method allows to control for global factors in industrial employment (e.g. trade and technological shocks) while removing the local endogenous response to the minimum wage hike. The sectoral shares are thus computed as:

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

And the resulting mean minimum wage takes the following form:

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

Despite the adjustment, the resulting series does not differ substantially from the one computed with local employment shares.

3.2 School enrollment and control variables

The main dependent variables are gross enrollment rates 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. Appendix A.5 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 enrollment 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 enrollment rates are in line with the aggregate time series published by Istat. Additional details on these procedures and checks are presented in appendix A.

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, which allows to check the bite of the minimum wage. 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 C.3—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 (see appendix A.3 for details and methdology). The local mean effective wage is obtained as the average of mean wages across the ten macro-sectors, using both weighting procedures detailed above (for details on sources and methodologies, see appendix A.1).

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 accounting estimates at the province level produced by Guglielmo Tagliacarne and the namesake Institute (Tagliacarne, 1963; Tagliacarne, 1972; Tagliacarne, 1975; Tagliacarne, 1979; Istituto Guglielmo Tagliacarne, 1986).[18] Youth unemployment is obtained from registrations at job centres at the provincial level (see section A.4 in the Appendix 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.

[18] Data for years 1978-79 are linearly interpolated from 1977 and 1980.

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 variables' trends before 1968 (Joshua David Angrist and J. Pischke, 2009; Caetano et al., 2022). Robustness checks that include the full set of time-varying values does not produce qualitatively different results.

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

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 M 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] The resulting map of wage zones is shown in Figure 10.

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. After a prolonged contrast which was punctuated by a series of strikes and required the mediation of the Ministry of Labour, the employers' associations caved in to the unions' requests. The resulting interconfederal agreement of 18 March 1969 established a convergence process which would partly reduce the wage differentials on 1 April 1969, halve them on 1 October 1970 and remove all remaining

[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).

differentials on 1 July 1972.[20]

The resulting compression of spatial differentials in minimum wage levels can be appreciated from Figure 12, 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 13 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]

[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 unions and the labour relations' representative for state-owned companies (Intersind). The resulting interconfederal agreement signed on 21 December 1968 established that all minimum wages should converge to the respective nominal levels of the Milan province, in three installments: the first installment would remove 40% of the nominal difference on 1 January 1969, the second installment would remove 30% of the difference on 1 April 1970 and the 30% would be removed on 1 July 1971 (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). On 8 March 1969, a similar agreement was reached with the association of small and medium-size industries: see article five of the *Accordo dell'8 marzo 1969 per il conglobamento dell'indennità di contingenza e per l'unificazione dei minimi di paga e di stipendio (aziende associate alla CONFAPI)* available 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

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 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 and the 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:

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

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

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. To ensure that results are robust to these outliers, I either exclude them from some of the regressions or I set their difference with respect to Milan equal to zero, without significant consequences for the estimates.

[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).

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 inverse of the gap between the nominal minimum wage in the province in 1968 and that of Milan in the same year, expressed in log differences. The inversion facilitates the interpretation of the estimated sign, but is not consequential for the results. 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 time-varying controls (X'_{it}) is built using the pre-treatment values and the coefficient for year 1968 is set to zero as a reference point (Borusyak, Jaravel, and Spiess, 2021). I also include time (τ) and province (α) fixed effects. As a robustness check, I alternatively include province-specific time trends, in which case I set to zero both the coefficient for 1968 and that for 1962 to avoid over-parameterization, following Kawaguchi and Mori (2021). 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 15. 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 1% 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]

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 11. 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 new currency for longer (Harris, 1957, pp. 445-449), so much so that these wage zones

[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, 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.

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 either a time trend for Southern provinces or macroregion 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 enrollment 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]

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 enrollment to the minimum wage shock of 1969, over time. Following Kawaguchi and Mori (2021) again, we can specify the reduced-form regression:

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

[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. Adding or excluding the region from the analysis does not significantly alter the results.

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 14 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. While there are no explicit tests for this assumption, as a robustness check we will estimate the average causal response only for marginal increases in the dose at different dose levels, rather than averaging across the whole range. This will allow to individuate possible heterogeneity in the causal response for different levels of the minimum wage hike.

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 the vector of trended pre-1968 controls includes the province's population, GDP per capita, the share of value added produced in the industrial sector, and the number of both young and prime-age individuals (under and over 21) registered as unemployed at local job centres.[28]

The OLS estimate without macroregion trends suggests an elasticity of .5 with respect to the mean minimum wage, which increase up to 0.89 in the fully-saturated specification. 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

[28]For this section only, the observations for Sicily are dropped due to missing data on the outcome variable.

bargaining system, which reduced room for firm-level adjustments, and by the fact that they affected 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 periods

Having established that the minimum wage hike had the potential to bite the wage distribution and 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 enrollment. 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 trended controls Table 3 presents the estimates for β and θ, both with and without macroregion trends.

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 0.45% increase in early school leavers, which is attenuated to 0.33% after we control for trended pre-1968 confounders. The 2SLS estimates present the same sign and are significantly larger. In the specification without controls, the coefficient increases to 0.612 and, after the inclusion of the controls, it reaches 0.631.

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 14% and 20%. This is close to but larger than the 12% difference between the enrollment rate extrapolated from the trend before 1969 and its effective value in 1972.

5.2.2 Robustness check: fractional response model

The OLS and IV estimators applied in the previous section assume linearity, which may cause inaccurate estimates with bounded variables, as is the case with enrollment rates. To avoid such biases, we chose to estimate the effect in levels of the dependent variables of interest, controlling for the size of the relevant demographic group. However, in order to check that the results are robust to alternative estimation strategies, in this section I normalize the dependent variables by the size of the relevant demographic group and I implement a fractional response model that allows for bounded endogenous variables. A presentation of fractional response models with panel data and IV estimators and the code for their implementation is provided by Papke and Wooldridge (2008), which I follow in the rest of the section.

To perform this analysis, I define a new dependent variable at the share s of individuals between the age of 14 and 18 in the province that are not enrolled in upper secondary education. This is essentially the inverse of the gross enrollment rate divided by 100. The variable is bounded by construction between 0 and 1, which respectively represent the extreme cases where either the whole cohort is enrolled in upper secondary school or none is. These are obviously theoretical bounds, for the lowest value registered in the dataset is .17, the greatest is .90 and half of the observations fall between .48 and .66. The linear specification is thus modified accordingly as:

$$s_{it} = \beta \, ln(M_{it}) + X'_{it}\,\gamma + \tau_t + \alpha_i + \xi_{it} \qquad [5]$$

Where X is the usual vector of trended pre-1968 controls, and τ and α are time and province fixed effects. The fractional response model, instead, is modelled as a generalized estimating equation (GEE) and takes the following specification:

$$E(s_{it}|M_{it}, \mathbf{X}_{it}, a_i) = \Phi(\psi + M_{it}\,\beta_1 + \bar{M}_i\,\beta_2 + X'_{it}\,\gamma_1 + \bar{X}'_i\,\gamma_2 + \tau_i + a_i) \qquad [6]$$

In contrast to the linear specification, this model does not use province fixed effects, but it controls for the province-specific time averages of all variables on the right-hand side ($\bar{M}$ and $\bar{X}$), which is akin to demeaning them. It does include, however, the full set of time dummies. $\Phi(.)$ is the standard cumulative distribution (cdf) of the Probit regression. This choice allows the response variable s of being defined as $0 \leq s_{it} \leq 1$. The term a indicates the residual. The discussion of this specification is provided by Papke and Wooldridge (2008).

First, for comparison purposes, I assume that the minimum wage is entirely exogenous. Hence, I estimate Equation 5 by OLS and the Equation 6 by Probit. The results are presented in columns 1 and 3 of Table 4. Both models find a positive and statistically significant association between the level of the minimum wage in the province and the share of people leaving school early. However, the coefficients are not directly comparable due to the different estimation techniques. To perform a more appropriate comparison, I compute the average partial effect (APE) from the Probit estimates, which can be compared with the marginal effect estimated by OLS. The APE for the exogenous specification (0.149) is close but slightly lower than the coefficient estimated by OLS (0.163), suggesting that the linear model overestimates the impact of the minimum wage hike on school enrollment.

I then relax the exogeneity assumption and I instrument the log of the nominal minimum wage with the usual gap in minimum wages interacted with the full set of time dummies (setting 1967 equal to zero to avoid multicollinearity), both in the linear and in the fractional response model. Column 2 of Table 4 shows the results obtained by 2SLS, while columns 5 and 6 the coefficient and the APE from the Probit model, respectively. As was the case in the main analysis, the IV strategy finds a larger coefficient than the OLS model, across both models: the average marginal effect estimated by OLS is 0.20, while the APE estimated by Probit is 0.193. In this case, too, it appears that the linear model slightly overestimate the effect with respect to the fractional probit model.

5.2.3 The causal 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 16 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 5.2%-7.3% (respectively, with and without controls). For reference, one quarter of the provinces had a wage gap larger than 10% with respect to Milan in 1968.

By 1972 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 enrollment 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 enrollment 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 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.

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 enrollment, 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 enrollment 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 enrollment as identified by the aggregate descriptive statistics also lasted about six years before starting to recover (cf. Figure 5), albeit with a small lag with respect to our estimates.

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 enrollment 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 enrollment, 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 enrollment 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 enrollment to decrease in both tracks. Either way, results would inform us about changes in the quantity of education demanded.

The choice between curricula, instead, reveals 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. Figure 17 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 enrollment 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 enrollment rate by over 2%. This effect is larger, in absolute values, than the general increase in early school leavers, which suggests that teenagers who would enroll in technical schools for manufacturing were disproportionally affected by the minimum wage hike. Moreover, 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 17c shows that there was no strong association between mean minimum contractual wages and male enrollment 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 enrollment in technical schools for manufacturing and the increase in enrollment in technical schools for business suggests that the two mechanisms discussed in the introduction 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 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 enrollment 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 enrollment 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, enrollment 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.

Figure 18 presents the results from the same estimation, but only for male students. The 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 man represented the vast majority of students choosing these school fields.The panels 18c and

18d, 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.

Figure 19 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 enrollment 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).

5.4 Implications for human capital accumulation in the long run

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 enrollment rates in upper secondary education and permanently decreased enrollment 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

[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.

individuals that would have graduated if the dip in enrollment 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 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 20 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 21 and Figure 22). 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

[30]The use of educational attainment rather than enrollment rates, as in the previous section, is justified by the fact that the same dynamics can be observed for both gross enrollment rates and gross graduation rates (see Figure 21), 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`.

the expansion of secondary school enrollment 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 A.5 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 enrollment, under the condition that the year of enrollment is between 1945 and 1969. Hence, I use the fitted values of the regression as a counterfactural 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 23 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 enrollment 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.

These estimates suggest that the loss in terms of human capital stock was sizeable, but was it also historically significant? To answer this question we need to compare Italy's trajectory with similar economies. To provide a comparative assessment of Italy's educational attainment with similar economies since the postwar period, Figure 24 presents a scatterplot of secondary school enrollment and GDP per capita across twenty-three European countries between 1945 and 2010. As we would expect, the fractional-polynomial prediction shows a strong positive association between the two measures, which weakens only after enrollment rates reach 80%. This is an expected observation since higher income levels are associated with a larger share of educated people in the population, but variation in GDP per capita remains relatively large even between countries with comparable levels of education.

The Italian graph lies below the expected value in every year, except possibly for 2005 and 2010, which means that, conditional on the level of GDP per capita, Italy had lower enrollment in secondary education than the European average through whole period. However, performance varied significantly over time: in 1945, the gross enrollment rate was ten percentage points lower than the conditional average, as would be expected given the historically lagging educational performance of the country. However, the Italian peculiarity is that enrollment expanded slower than expected given GDP growth in the post-war decade, so that by 1955 the gap with the conditional average had grown to 30 percentage points. Secondary school enrollment eventually accelerated in the second half of the 1950s and by 1975 it had almost converged to the conditional average. Starting in the second part of the 1970s and through the 1980s, instead, enrollment stagnated while GDP per capita continued to grow—by 1990, Italy's enrollment rate was twenty percentage points lower than expected, making it a negative outlier in comparison to all other comparable European countries. Enrollment rates rose again in the 1990s through the 2000s, which allowed Italy to recover from the pause of the 1970s-1980s and converge to the conditional average.

Thus, the slow-down of educational attainment led Italy along a temporary diverging path from comparable European economies. But by just how much did the this diversion 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 enrollment between the 1970s and the 1980s explains between 24% and 44% of Italy's lag in 2015.

6 Conclusions

Collective agreements are labour market institutions that, like statutory minimum wages, can regulate entry-level wages and address income inequalities. However, their relatively large bite with respect to the wage distribution can significantly raise the opportunity cost of investing in formal education for marginal student, 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 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 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 enrollment 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 enrollment rates continues to affect Italy's ranking in educational attainment to this day. Depending on the counterfactual scenario, the pause in enrollment 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 educational choices and, through them, their opportunities when entering the formal labour market. These results also reinforce our argument that empirical analysis 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.

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Figures

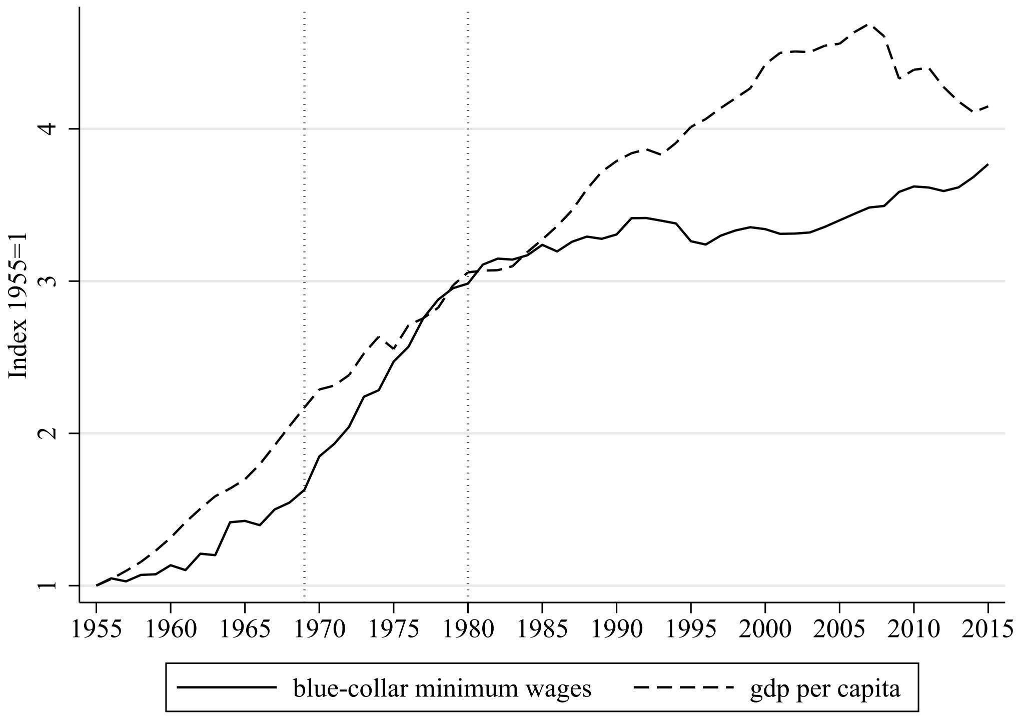


Figure 1: Minimum wages and GDP

Index of contractual hourly wages across for blue-collar workers, 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. Source: own elaborations on data from Istat, *Serie Storiche*, Tav. 10.21, available at https://seriestoriche.istat.it/, and Bank of Italy, *La contabilità nazionale in Italia dall'Unità a oggi, 1861-2017*, available at https://www.bancaditalia.it/statistiche/tematiche/stat-storiche/stat-storiche-economia/index.html—for information see also Baffigi (2015).

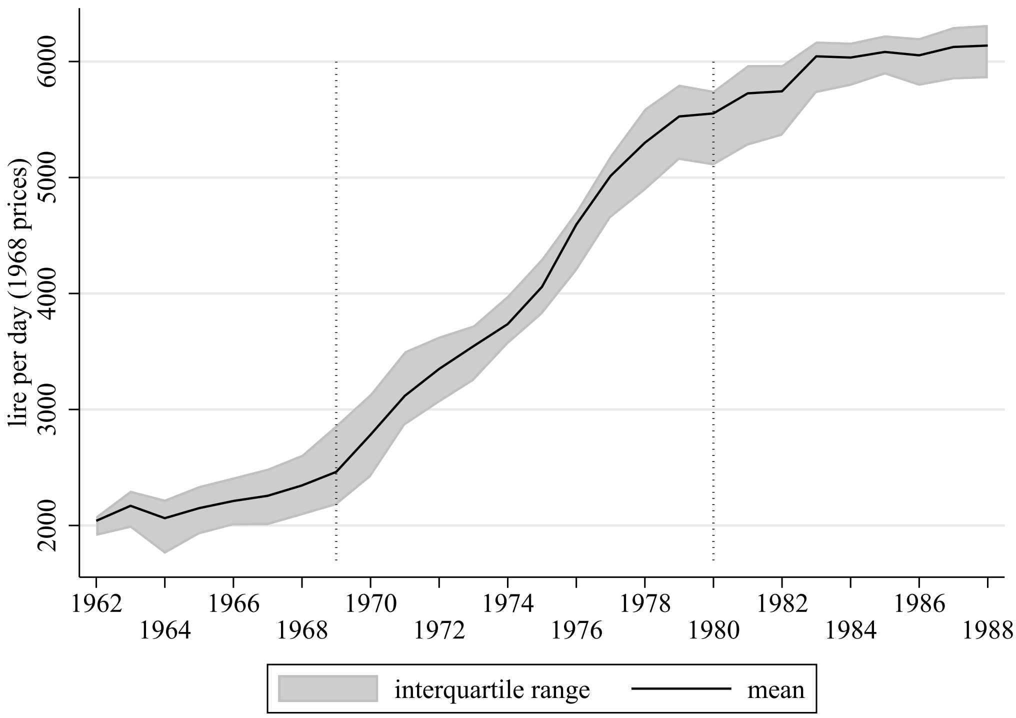


Figure 2: Minimum wage in manufacturing

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 appendices A.2 and A.1. 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.

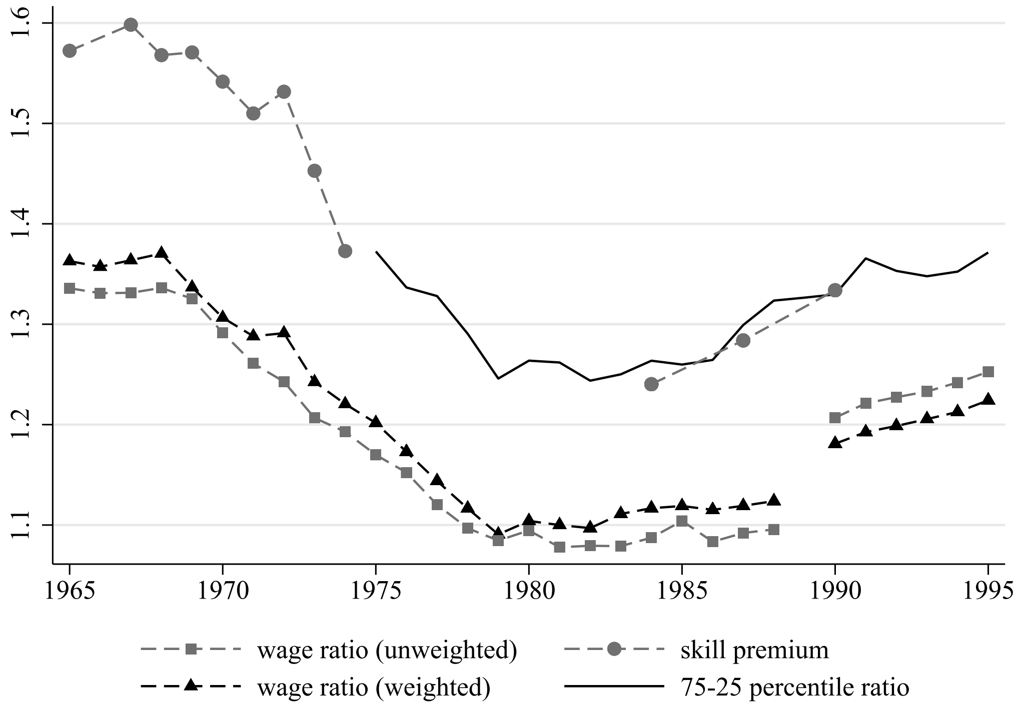


Figure 3: Wage ratio, skill premium and wage dispersion

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 data appendix for harmonization methods.

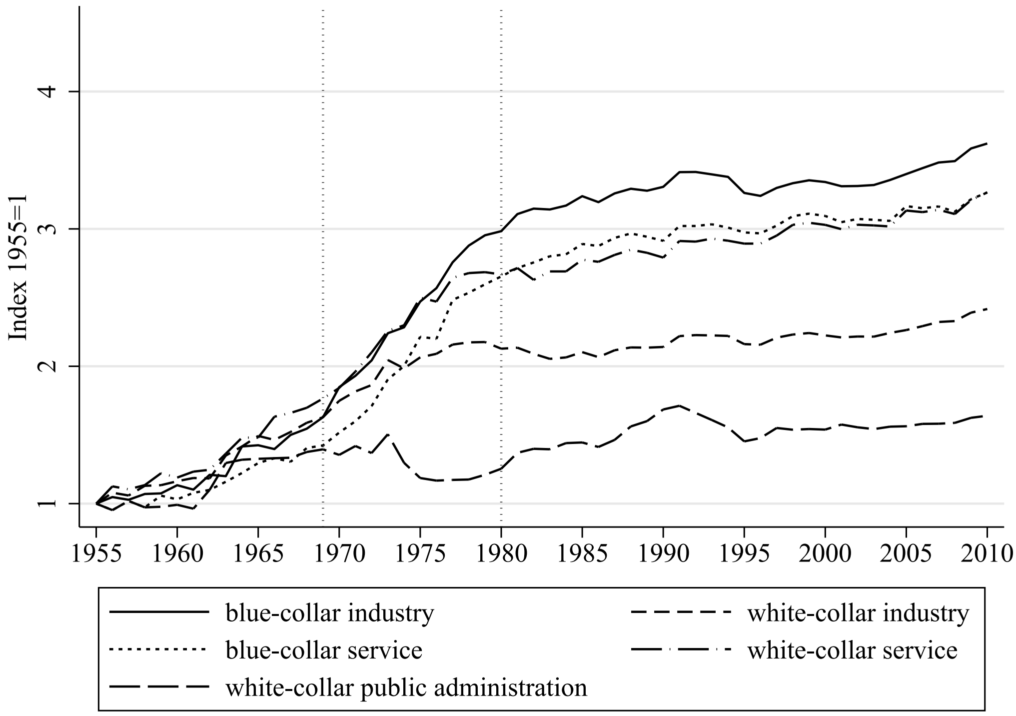


Figure 4: Changes in wage floors by worker category and macro-sector

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.

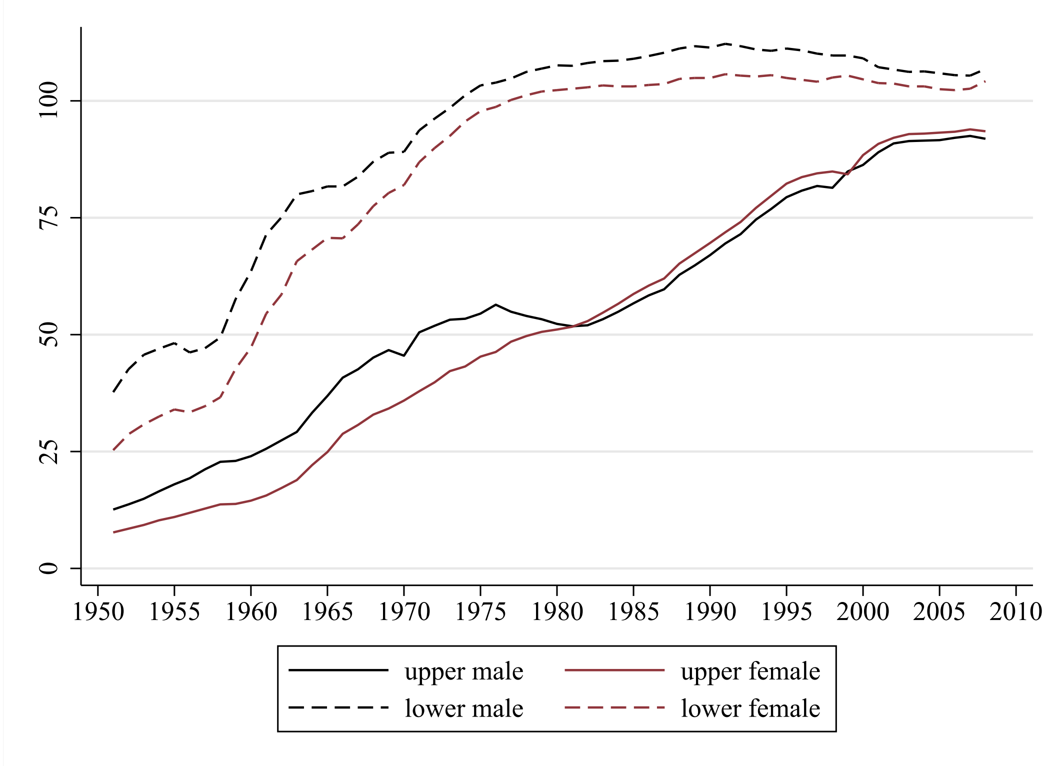


Figure 5: Enrollment in lower and upper secondary education by sex

Gross enrollment 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: Sistema Statistico Nazionale and Istituto Nazionale di Statistica (2011, p. 369).

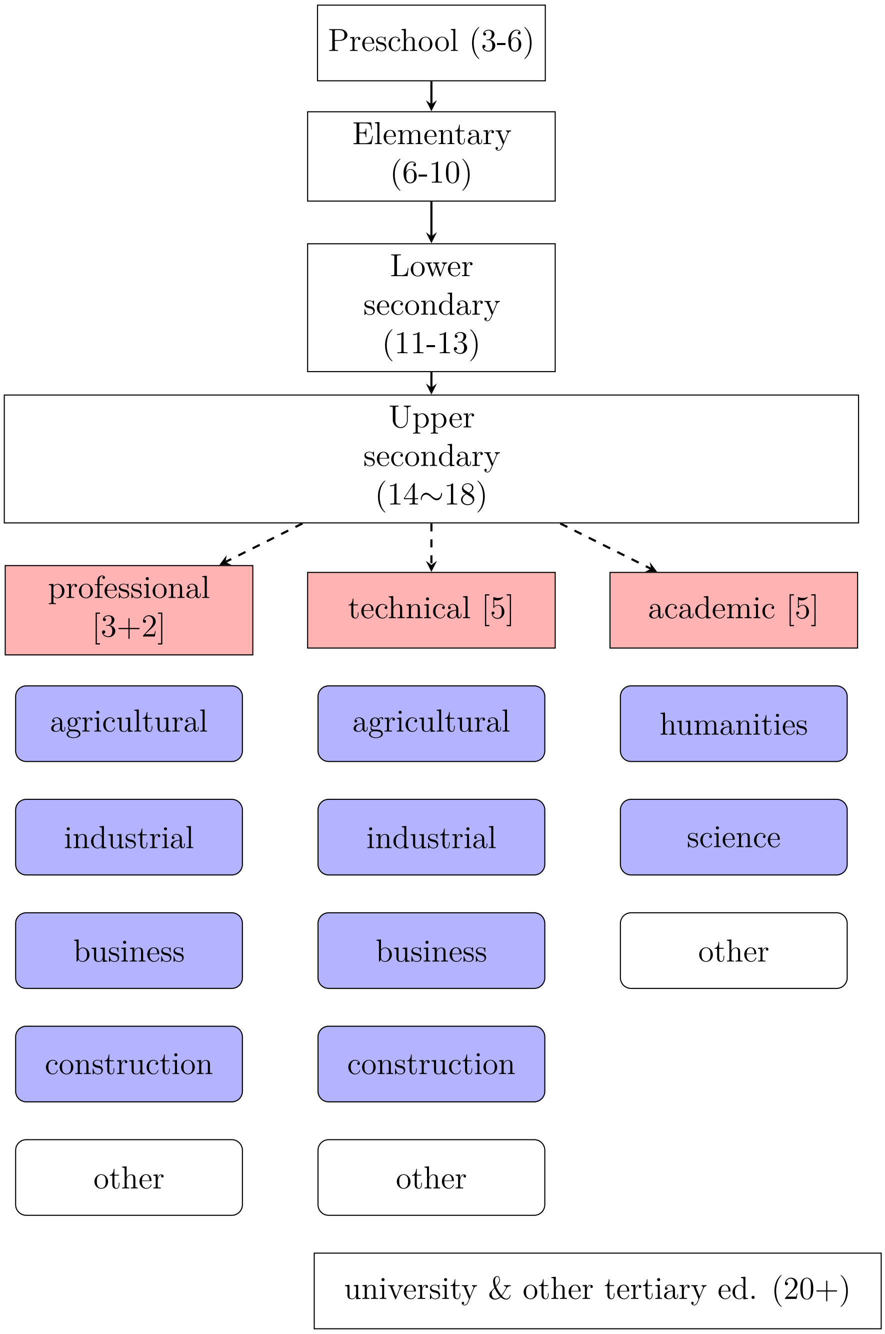
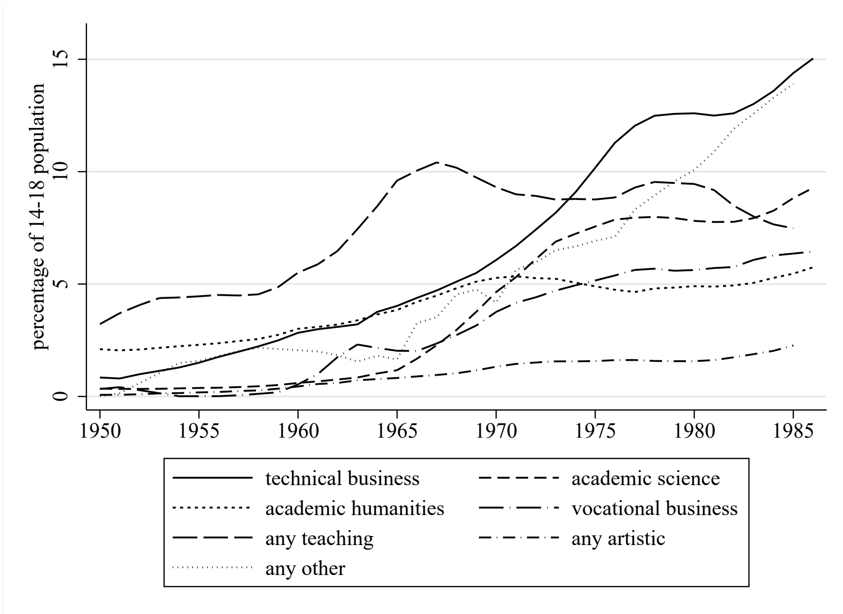
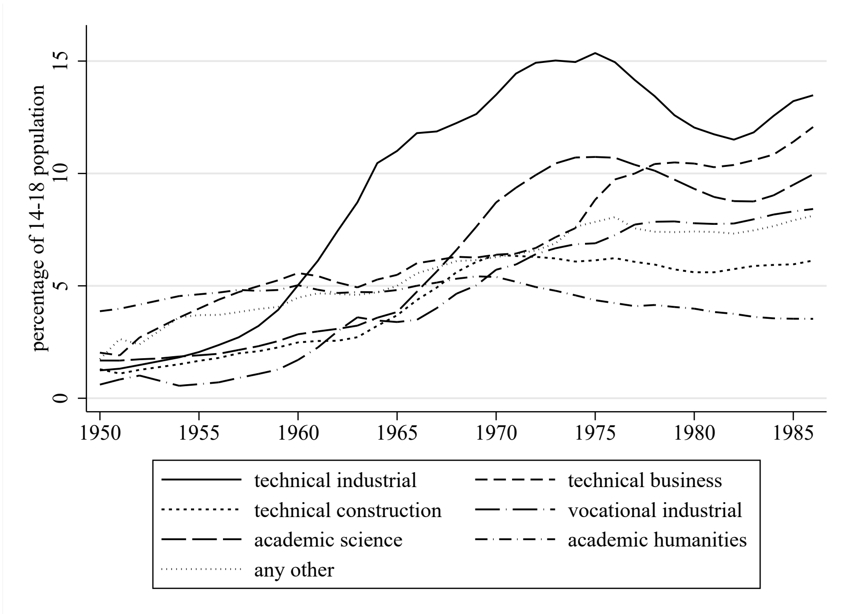


Figure 6: TRACKING AND CURRICULA IN THE EDUCATIONAL SYSTEM, 1960S-1980S

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*).



(a) Female



(b) Male

Figure 7: Enrollment rates by sex, track and curriculum, 1950-1986

Gross enrollment 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). For tracks and curricula translation in Italian see Appendix A. '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.

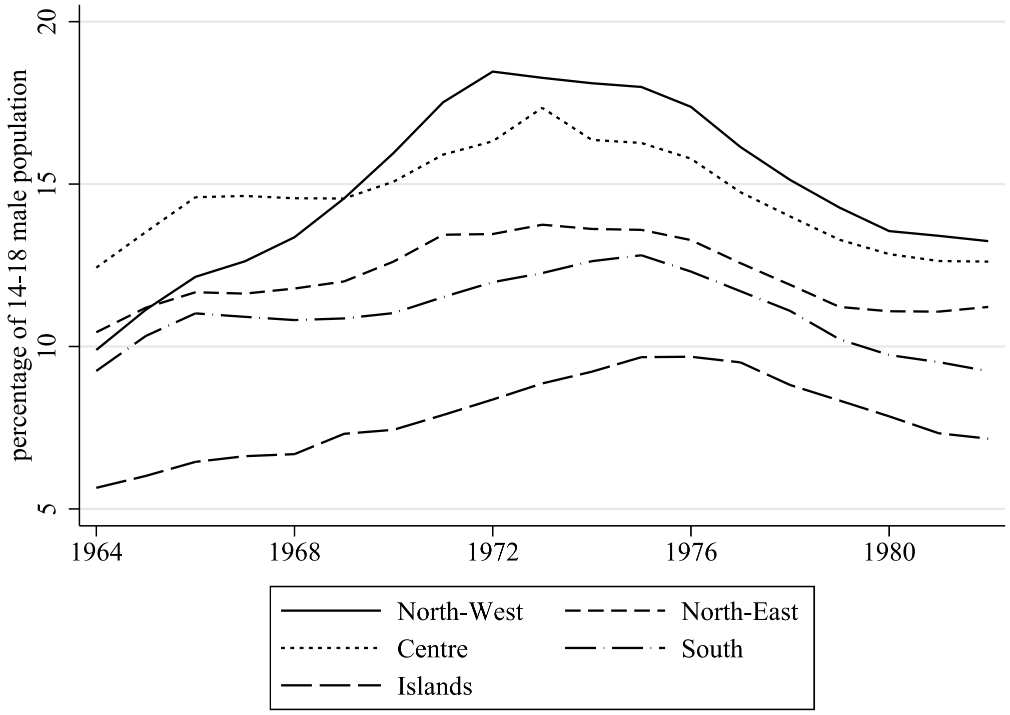


Figure 8: Male enrollment in technical schools for manufacturing by macroregion

Gross enrollment 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 Appendix A.

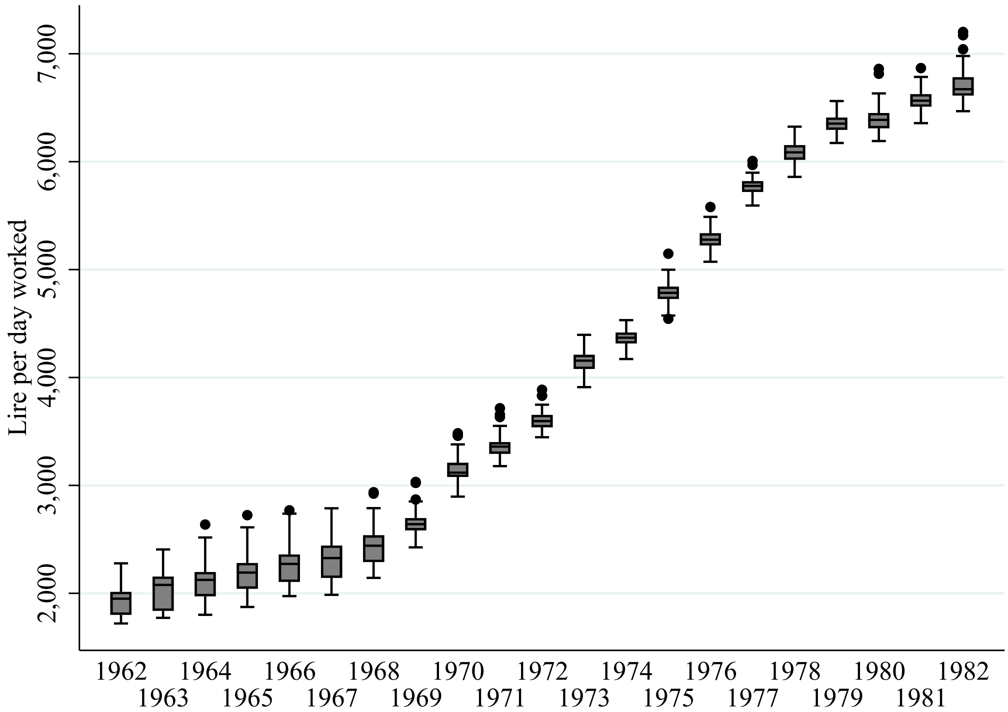


Figure 9: Box plot of mean minimum wages at 1968 constant prices

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 appendices A.2 and A.1. 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).

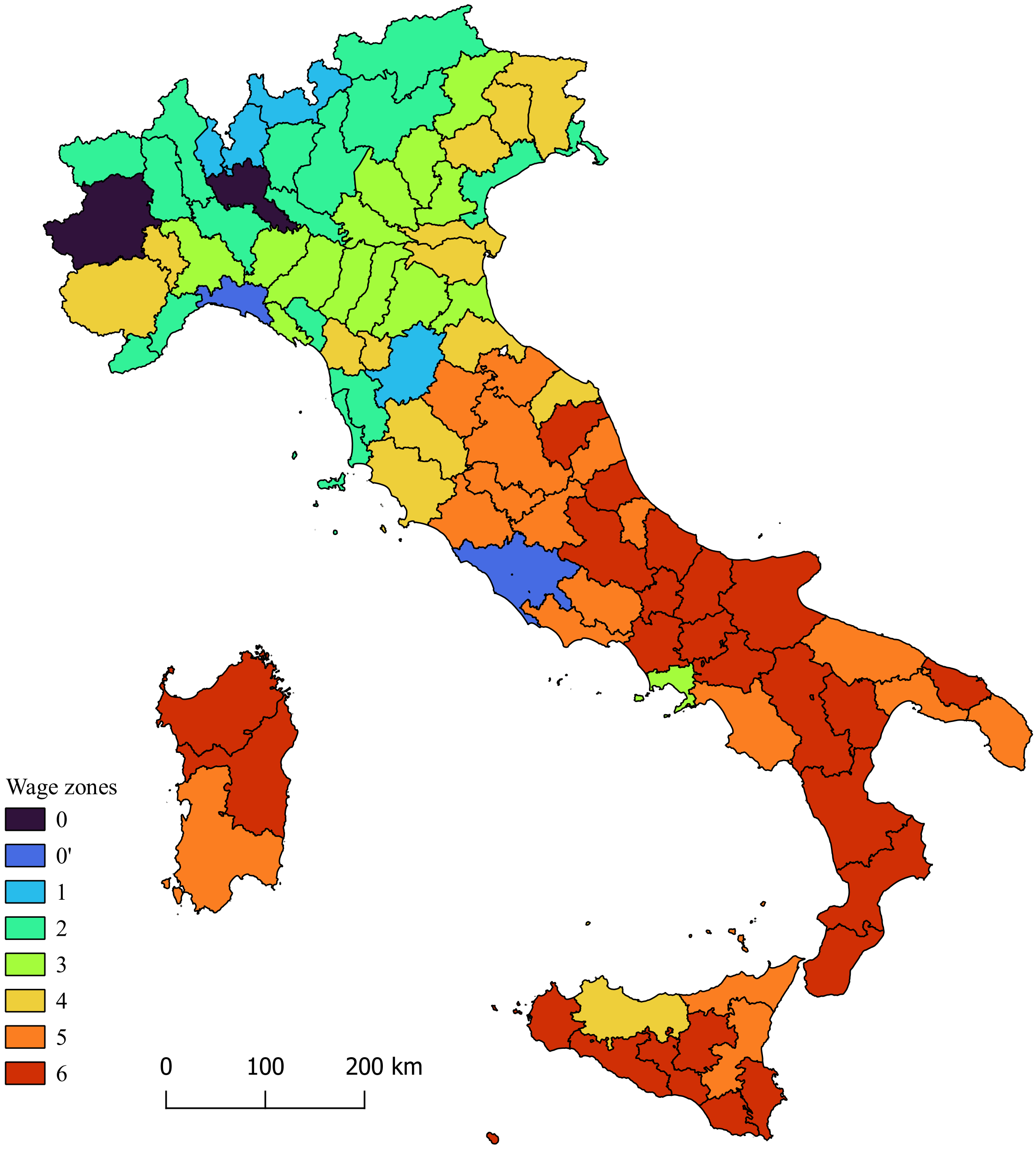


Figure 10: Wage zones before 1969

Classification of provinces according to the assigned wage zone during the 1960s. These wage zones were defined with the interconfederal agreement of 2 August 1961, which established seven wage zones, from zero to six. Wage zone zero included Milan, Turin, Rome and Genoa. However, nominal wage levels differed between the former couple of provinces and the latter. To signal this, the map indicates Rome and Genoa as wage zone 0'. The wage zones were abolished with the interconfederal agreement of 18 March 1969. Source: Istituto Centrale di Statistica (1969b, p. 150, footnote a). The shapefile of the provinces at 1971 historical borders is available at https://www.istat.it/it/archivio/231601 (last retrieved July 2022).

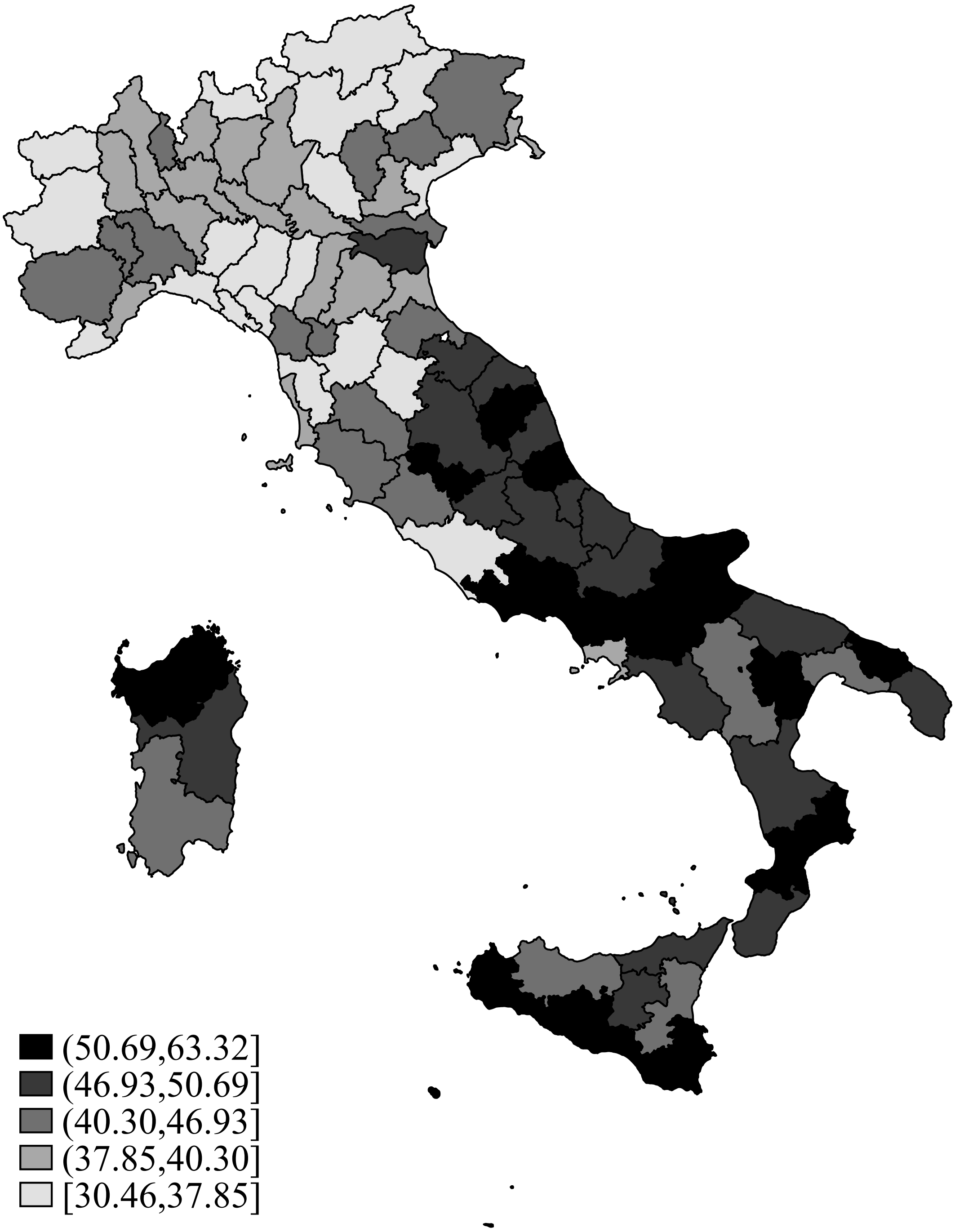


Figure 11: Change in mean industrial minimum wage 1968-1972

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 on the methodology see section 2.1. For the sources of the minimum wage data see section A.2. The shapefile of the provinces at 1961 historical borders is available at `https://www.istat.it/it/archivio/231601` (last retrieved July 2022).

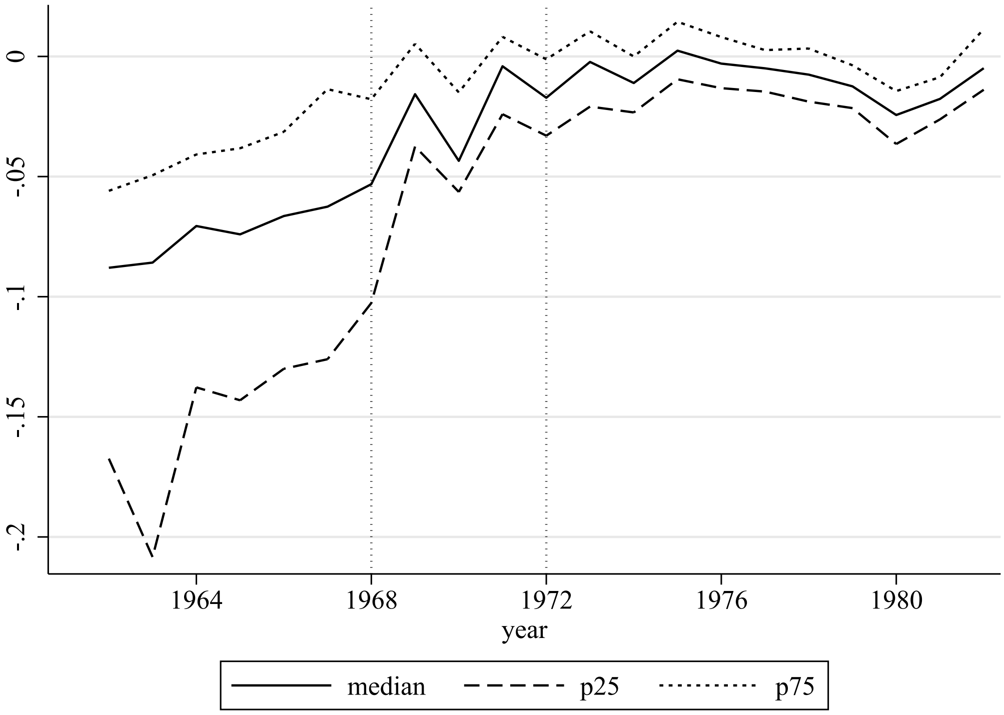


Figure 12: Minimum wage differentials with respect to Milan

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 Appendix A.2.

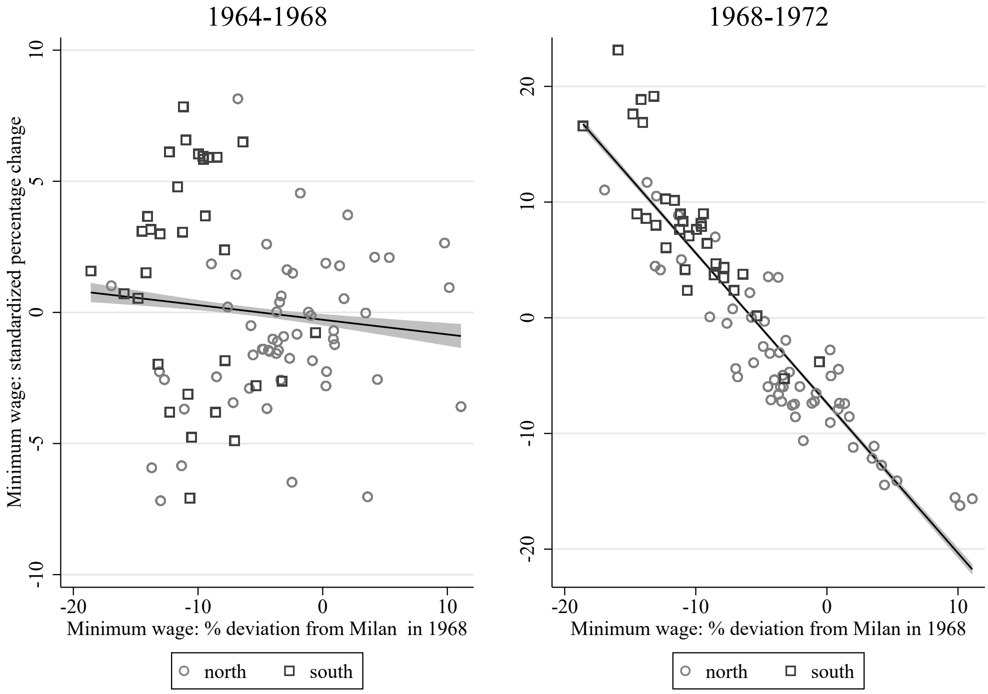


Figure 13: Nominal wage equalization and minimum wage growth

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 section A.2.

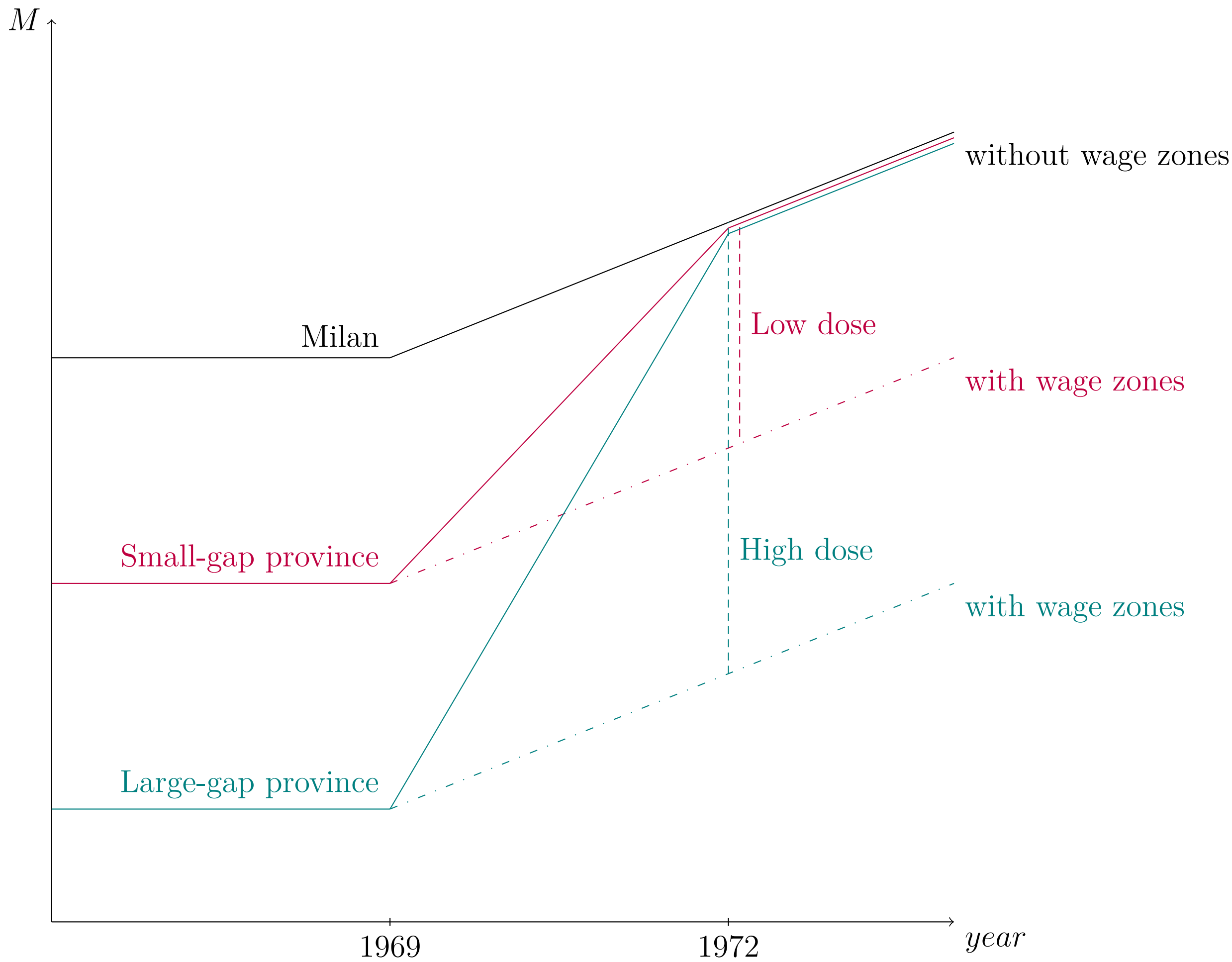


Figure 14: Representation of the identification strategy for DiD

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.

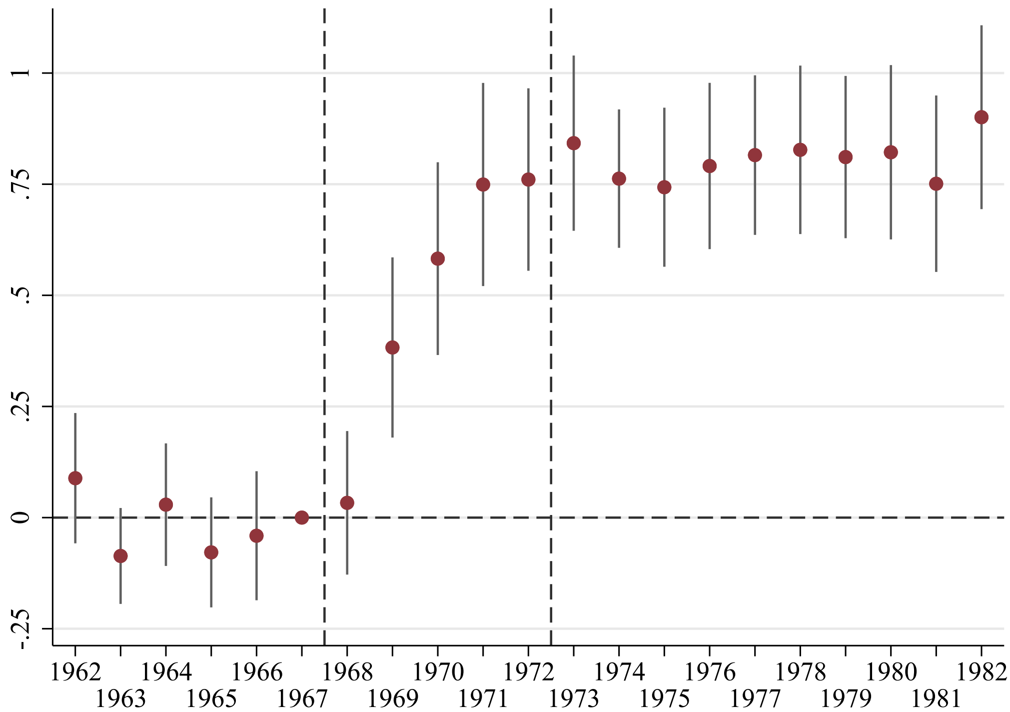


Figure 15: Gap with respect to Milan in 1968 and minimum wage levels

Coefficients for the interaction term between the inverse of the log-point difference with respect to the mean minimum wage of Milan in 1968 and year dummies, with 1968 set to zero. OLS estimates controlling for time and province fixed effects, and time-varying controls. 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.

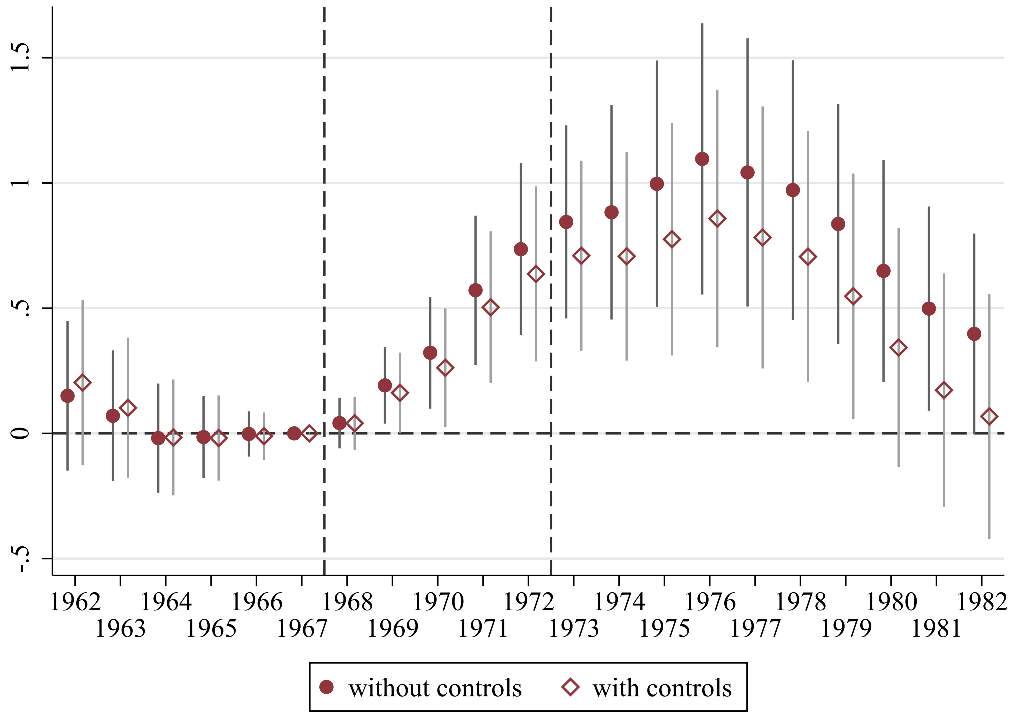
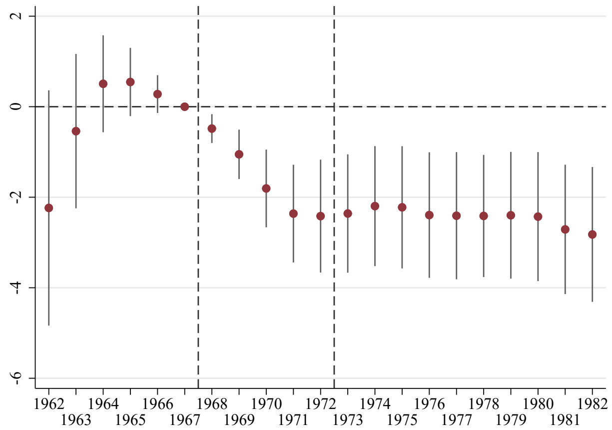
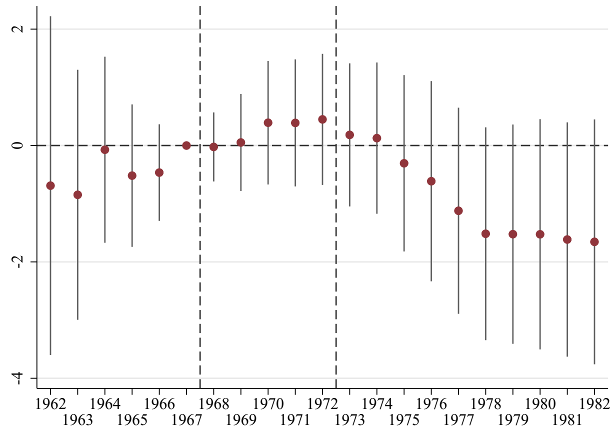


Figure 16: Dynamic response of early school leavers

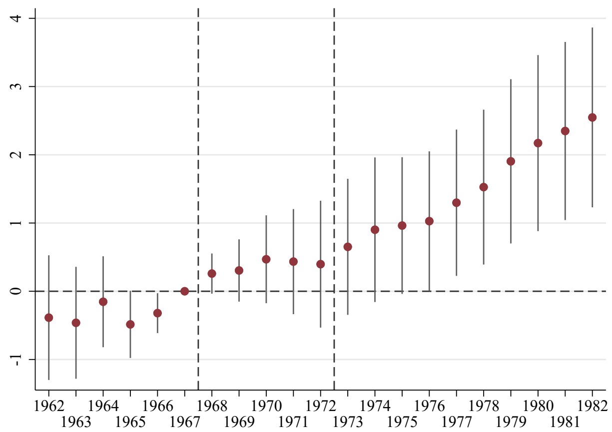
OLS estimates if the coefficients for the interaction term between the inverse of the minimum wage gap with respect to Milan in 1968 and year dummies, with 1967 set to zero to avoid multicollinearity. The dependent variable is the log of the number of individuals between the age of 14 and 18 not enrolled in post-compulsory upper secondary education. The regression controls for the size of the age cohort in every province-year cell. Both specifications control for time and province fixed effects. The second specification also includes trended pre-treatment controls. 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: 1344. The estimation excludes provinces in special statute regions and provinces that had a higher mean minimum wage than Milan in 1968.



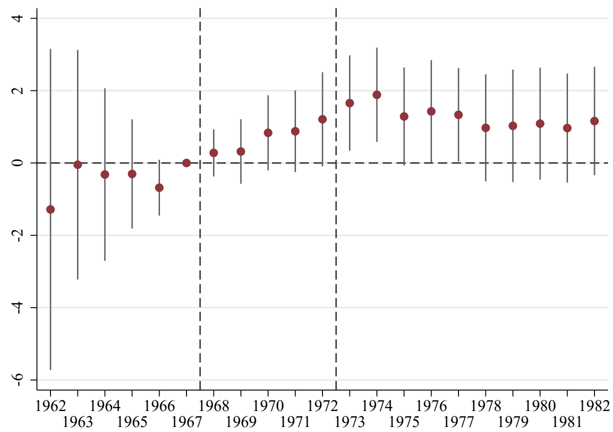
(a) technical manufacturing



(b) professional manufacturing



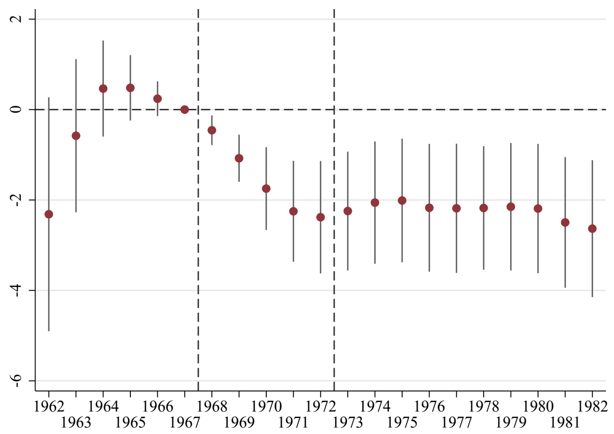
(c) technical business



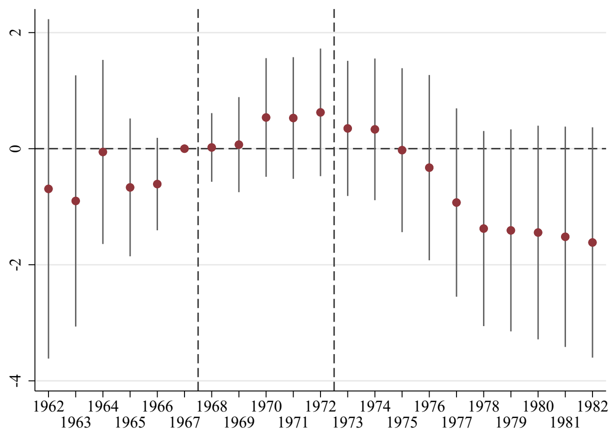
(d) professional business

Figure 17: Dynamic response by school type (male and female)

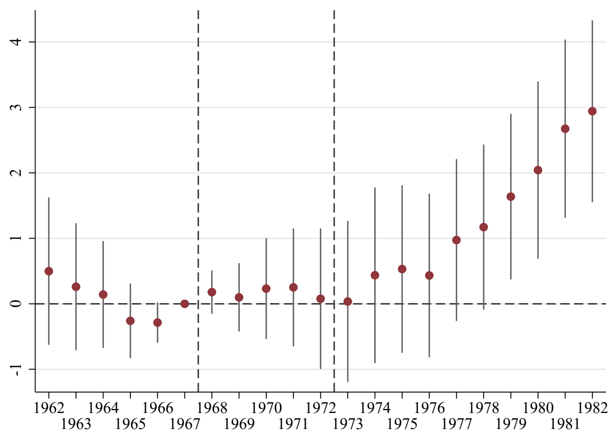
Coefficients for the interaction term between the inverse of the log-point difference with respect to the mean minimum wage of Milan in 1968 and year dummies, with 1968 set to zero. OLS estimates controlling for time and province fixed effects. The vertical lines indicate the 95% confidence interval 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.



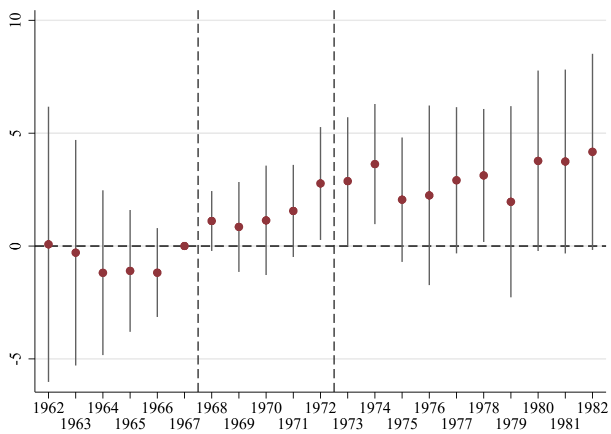
(a) technical manufacturing



(b) professional manufacturing



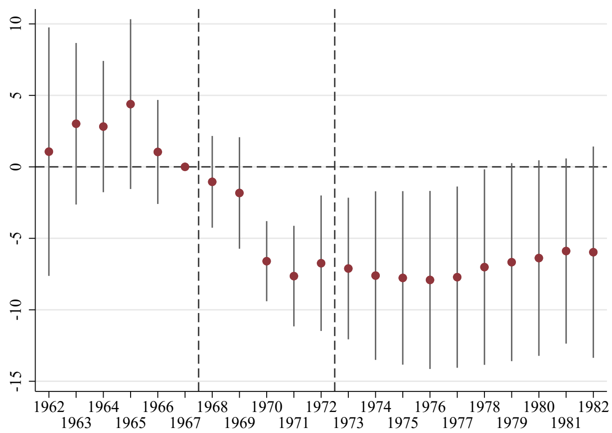
(c) technical business



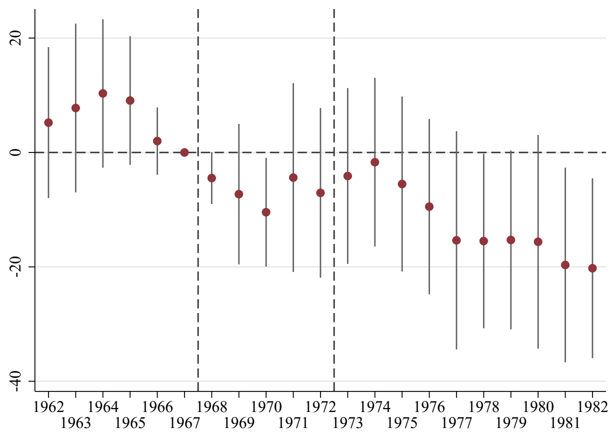
(d) professional business

Figure 18: Dynamic response by school type (male only)

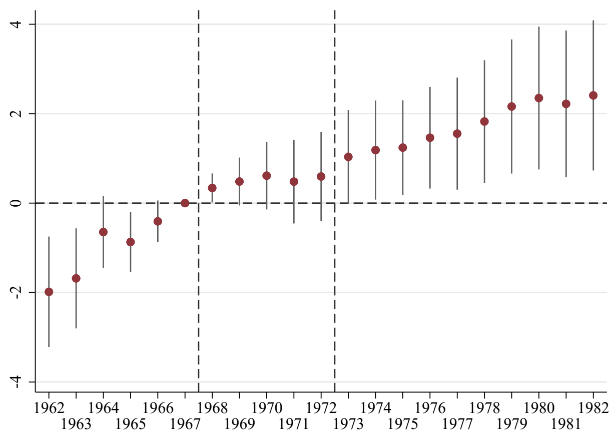
Coefficients for the interaction term between the inverse of the log-point difference with respect to the mean minimum wage of Milan in 1968 and year dummies, with 1968 set to zero. OLS estimates controlling for time and province fixed effects. The vertical lines indicate the 95% confidence interval 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.



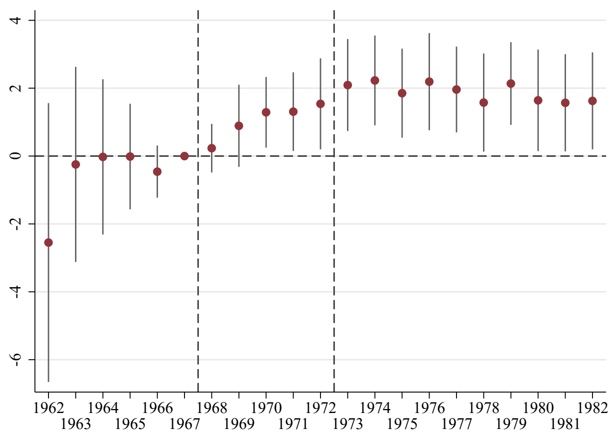
(a) technical manufacturing



(b) professional manufacturing



(c) technical business



(d) professional business

Figure 19: Dynamic response by school type (female only)

Coefficients for the interaction term between the inverse of the log-point difference with respect to the mean minimum wage of Milan in 1968 and year dummies, with 1968 set to zero. OLS estimates controlling for 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.

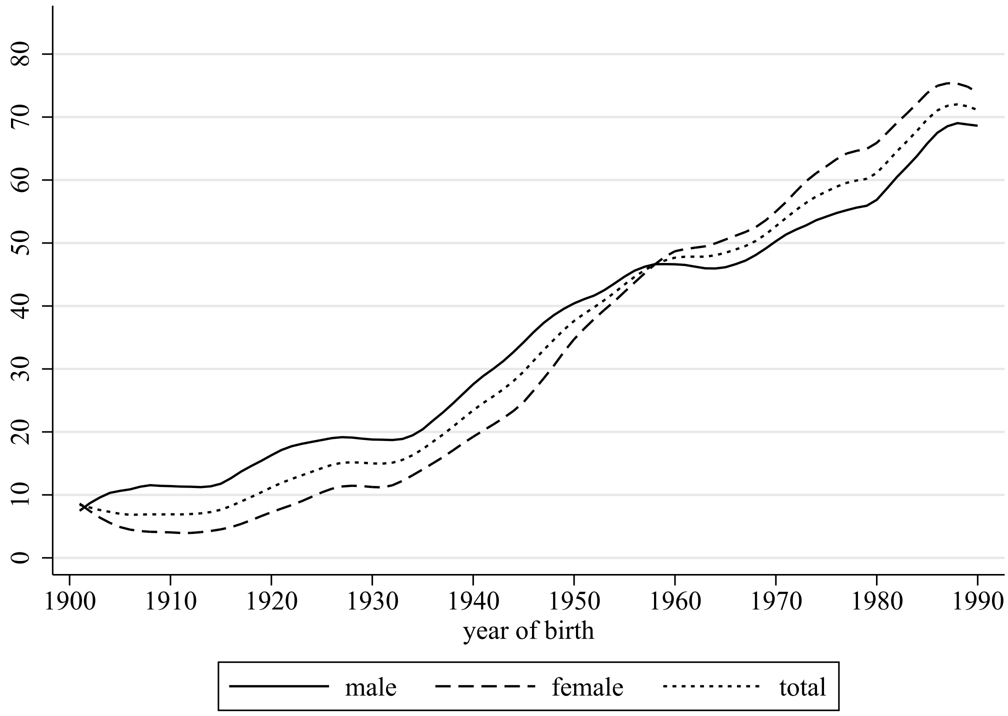


Figure 20: Upper secondary school attainment by birth cohort

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 21: Enrollment and graduation rates in upper secondary school

Gross enrollment 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.

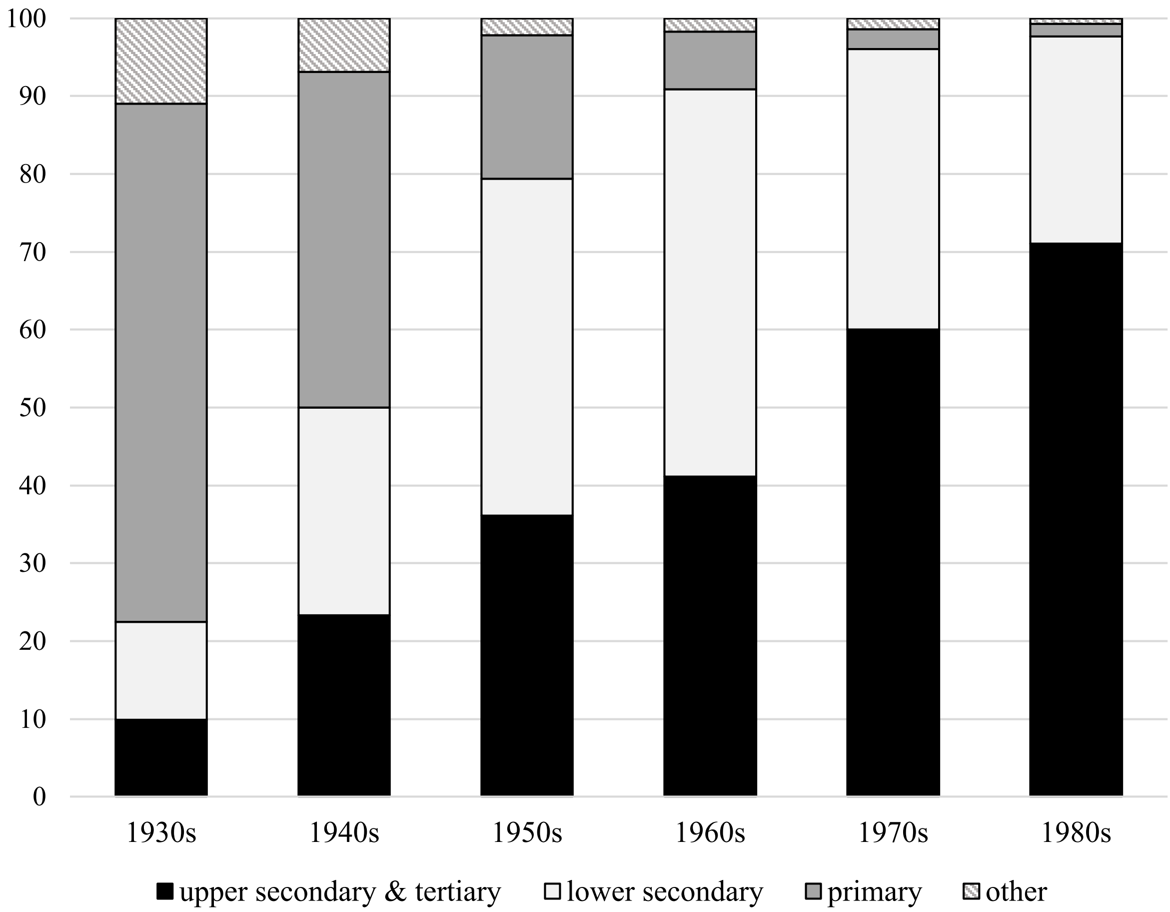


Figure 22: Educational attainment by decade of birth

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 underestimatation 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).

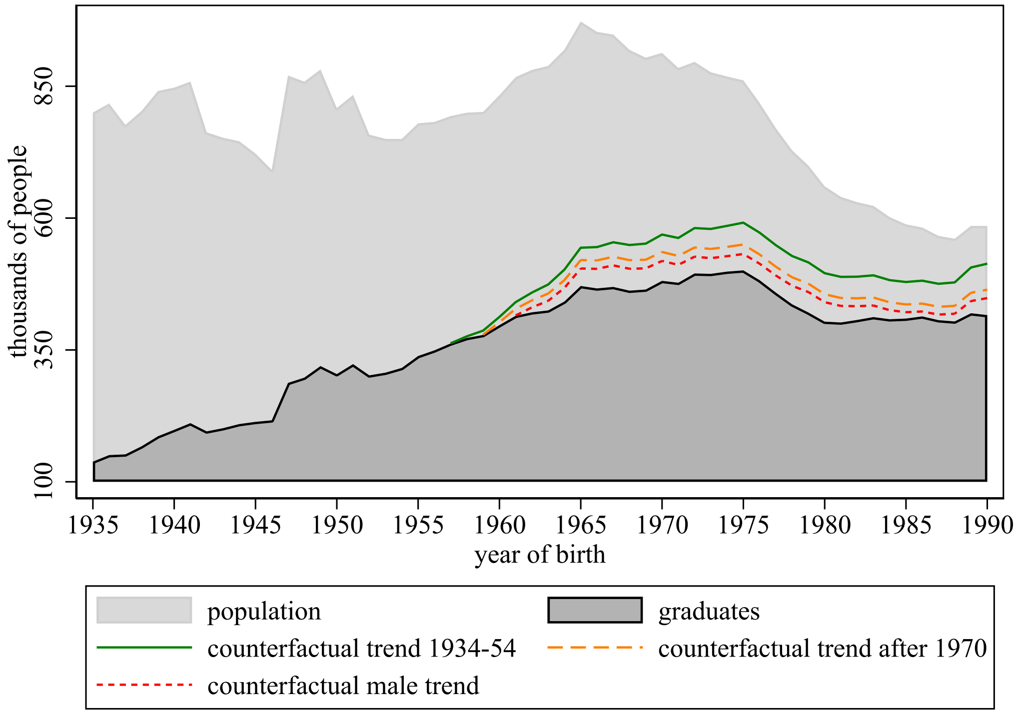


Figure 23: Upper secondary school graduates and counterfactual estimates

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 to account for the annual frequency of the data (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 birth cohorts born in 1934-1954. The counterfactual trend after 1970 shows the total number of graduates if attainment for the 1954-1969 birth cohorts had expanded following the same trend as the cohorts born since 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 obtained after linearly regressing the smoothed number of graduates from the HP filter on a time trend, restricted to the relevant birth cohorts (female only for the third counterfactual), and multiplying the resulting predicted enrollment 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. The year of birth for the population of age 18 in each year is computed by subtracting 18 to each year. Hence, the first cohort available for analysis was born in 1934 (1952-18). The graph only shows cohorts 1935-1990.

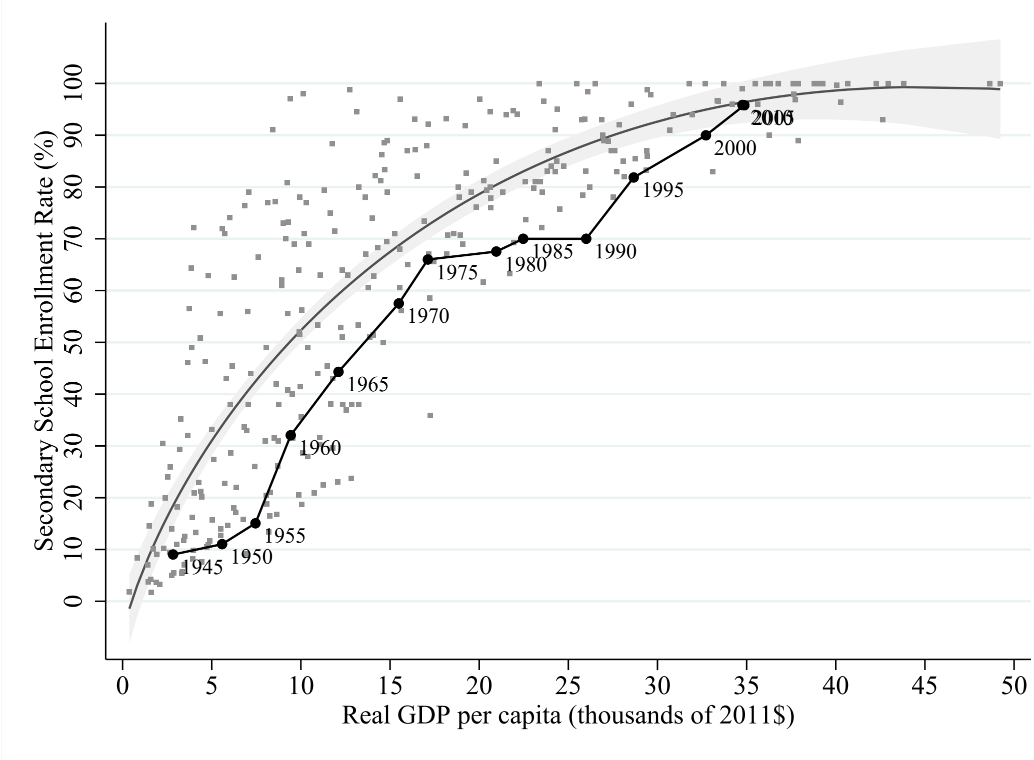


Figure 24: Secondary education and GDP across Europe

The black connected markers represent Italy in the labelled year (years 2005 and 2010 overlap). Each square marker in the scatterplot represents a European country in a given year between 1945 and 2010. Countries included are Albania, Austria, Belgium, Bulgaria, Czech Republic, Denmark, Finland, France, Greece, Hungary, Iceland, Ireland, Malta, Netherlands, Norway, Poland, Portugal, Romania, Spain, Sweden, Turkey, and the United Kingdom. The solid line represents a twoway fractional-polynomial prediction plot estimated on all countries excluding Italy, the shaded area represents 95% confidence intervals. Quinquennial data between 1945 and 2010 or latest available. Adjusted secondary school enrollment rate are obtained from J. W. Lee and H. Lee (2016), GDP per capita from Maddison Project Database, version 2020 (Bolt and Zanden, 2020).

Tables

Table 1: Descriptive statistics by period

	(1) 1962-1968 mean	(1) 1962-1968 sd	(2) 1969-1972 mean	(2) 1969-1972 sd	(3) 1973-1982 mean	(3) 1973-1982 sd
ln(minimum wage)	7.62	0.17	8.16	0.17	9.62	0.61
ln(minimum wage const)	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
ln(average wage const)	8.14	0.25	8.70	0.21	9.98	0.58
% 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	

Table 2: Log average wage elasticity to mean minimum wage

	ln(average effective wage)			
	(1) OLS	(2) OLS	(3) OLS	(4) 2SLS
ln(minimum wage)	0.533***	0.622***	0.745***	0.887***
	(0.108)	(0.104)	(0.187)	(0.166)
Province FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Pre-treat controls	Yes	Yes	Yes	Yes
Macroregion FE	No	Yes	No	Yes
Clustered SE	Yes	Yes	Yes	Yes
Adj R2	0.998	0.998		
Adj within R2	0.323	0.34		
Kleibergen-Paap F statistic			156.69	147.66
N	1701	1701	1701	1701

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 trends in total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector. All specifications include time and province fixed effects. Standard errors are clustered at the province level. The observations exclude provinces with an average minimum wage higher than Milan in 1968, due to the instrument's restriction.

Table 3: MINIMUM WAGE AND EARLY SCHOOL LEAVERS

	ln(early school leavers)			
	(1) OLS	(2) OLS	(3) 2SLS	(4) 2SLS
ln(minimum wage)	0.446***	0.325***	0.612***	0.631***
	(0.133)	(0.119)	(0.197)	(0.233)
Province FE	Yes	Yes	Yes	Yes
Time FE	Yes	Yes	Yes	Yes
Pre-treat controls	No	Yes	No	Yes
Time-variant cohort size	Yes	Yes	Yes	Yes
Clustered SE	Yes	Yes	Yes	Yes
Adj R2	0.996	0.952		
Adj within R2	0.696	0.744		
Kleibergen-Paap F statistic			164.747	133.593
N	1344	1344	1344	1344

Cluster-robust 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 trends in total population, prime age and youth unemployment, GDP per capita, and share of GDP produced in the industrial sector. All specifications include time and province fixed effects. Standard errors are clustered at the province level. The observations exclude provinces in special statute regions and those with an average minimum wage higher than Milan in 1968, due to the instrument's restriction.

Table 4: ROBUSTNESS TEST: FRACTIONAL RESPONSE MODELS

	share not enrolled					
	Linear		Fractional probit			
	OLS	2SLS	GEE			
	Exogenous	Endogenous	Exogenous		Endogenous	
	Coefficient (1)	Coefficient (2)	Coefficient (3)	APE (4)	Coefficient (5)	APE (6)
ln(minimum wage)	0.163***	0.202***	0.393***	0.149***	0.510***	0.193***
	(0.046)	(0.0623)	(0.120)	(0.0454)	(0.160)	(0.0607)
scale factor				0.379		0.378
Pre-treat controls	YES	YES	YES		YES	
Time FE	YES	YES	YES		YES	
Province FE	YES	YES				
Clustered SE	YES	YES	YES		YES	
N	1512	1512	1512	1512	1512	1512

Standard errors in parentheses

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

Estimates obtained from regressing the share of individuals between the age of 14 and 18 not enrolled in upper secondary education on the natural logarithm of the minimum wage and the usual pre-treatment trended controls. Column 1 reports the coefficient estimated by OLS in the baselines specification. Column 2 reports the coefficient estimates by 2SLS, after instrumenting the independent variables with the minimum wage gap with respect to Milan in 1968, interacted with the time dummies. Both specifications include time and province fixed effects. Column 3 and 4 report the coefficient obtained with the generalized estimating equation (GEE) approach described by (Papke and Wooldridge, 2008). Column 3 assumes strict exogeneity and regresses the outcome on the dependent variable, while column 5 is the second stage from the predicted values obtained by regressing the log of the minimum wage on the wage gap interacted with time dummies. In both cases, the model controls for time averages of the dependent variables. Standard errors are clustered at the province level across all models. For the fractional probit model, standard errors are obtained by bootstrapping the 1512 provinces using 1500 bootstrap replications. Columns 4 and 5 report the Average Partial Effects estimated by multiplying the coefficient for the reported scale factor.

Appendix

A Primary sources and harmonization procedures

The dataset used for the main analysis has been assembled from a range of primary sources, the majority of which have been specifically digitized for this project. Since the primary sources were produced by a multitude of institutions for different purposes and using different classification criteria, the design and execution of harmonization procedures have represented critical preliminary steps for the research, and an additional contribution to the literature. This appendix provides a description of the sources used and the harmonization procedures undertaken, with as detailed information as possible to the location of the data and the methodologies, to ensure replicability of the data construction process.

Both average minimum and effective earnings at the province level have been computed by taking into account the local industrial composition. Hence, the first step in the analysis consisted in the identification of the covered sectors and the computation of local industry weights. The next section describes the computation of local industry weights from the industrial censuses, while A.2 and A.3 detail the sources and methodologies to compute the average minimum and mean effective wages. Section A.4 discusses the sources of unemployment data. Section A.5 details the reconstruction of the intercensual age structure of the provincial population, which was instrumental to the computation of enrollment rates and the normalization of unemployment data. Special attention is given to resolving some errors in the sources.

A.1 Provinces' industrial composition

Local industry weights are computed from an electronic database reporting the number of establishments and employees at the municipality level (LAU 2) for 1961, 1971, 1981 and 1991. The data, originally from the industrial censuses carried out in the respective years, was harmonized by Istat to allow intertemporal comparisons between industrial sectors. Access to the electronic data was possible at the following website: http:

//dwcis.istat.it/index.html (last retrieved 26/11/2019). However, since September 2020, the website has been disabled. Nonetheless, information on the dataset, including sources and methodology, can be obtained from the archived snapshots at the Internet Archive through the WayBack Machine (see for instance https://web.archive.org/web/20161110223356/http://dwcis.istat.it/cis/index.htm. The archived version does not allow to download the data at the municipality level due to login procedure failing, but the raw files can be provided upon request.

Data were originally output into a distinct csv file for each census year of 1951, 1961, 1971, 1981, 1991, 1996 and 2001. Four fields were reported in each file: code and name of the municipality; code and description of the subsector (*categoria economica*) at four-digit level; number of establishments and number of employees in the sector. I have cleaned the dataset dividing codes and descriptions into separate fields, and I appended the cross-sectional csv files to obtain a panel dataset. An important notice is that municipality-sector cells were reported only if they had at least one establishment in one census year. Consequently, the resulting panel was unbalanced, as it missed all municipality-sector cells that were equal to zero in all census years. To produce a complete balanced dataset, I obtained the full list of municipalities in each census year from Istat's *Atlante Statistico dei Comuni*, 2014 edition (available at https://www.istat.it/it/archivio/113712), and I added all missing municipality-sector cells, setting the value of establishments and employees equal to zero. The resulting dataset contains information on the number of establishments and employees in 8,141 municipalities for 57 sectors in each census year, for a grand total of 1,852,272 observations. Sectors include extractive industries, manufacturing, construction, energy and services, but not agriculture. For this paper, I only consider 21 sectors in manufacturing proper, in addition to extraction of metallic minerals, construction, and energy. Tab. C.1 lists the sectors considered and the number of establishments and employees in 1971. The 21 sectors in manufacturing proper cover over 97% of all manufacturing establishments and almost 98% of all employees in 1971. Other manufacturing sectors not considered are tobacco, production of cinematographic, photographic and phonographic materials (which is originally combined with services such as movie production), and 'other manufacturing', a residual category from the census classification. Among non-manufacturing sectors, I exclude extraction of non-metallic minerals, oil and gas refinery, and the water industry. I exclude from the manufacturing classification all repair shops (which were considered so until the 19 census) due to their later and more appropriate classification into the service sector. For a discussion, see Istat (1998).

The data was not originally harmonized for geographical comparisons. To measure the industrial composition at the provincial level I have first harmonized the data to historical provincial boundaries (choosing 1961 as the benchmark year) by aggregating the municipality data into contemporary provinces first and then aggregating provinces to 1961 boundaries. This implied adding the values for the new provinces of Pordenone (established in 1968) to Udine, that of Isernia (established in 1970) to Campobasso and that of Oristano (established in 1974) to Cagliari, when present. Secondly, I have harmonized the data for comparison with the wage series, by using the sector definitions of the minimum wage series. The harmonization consisted in summing the number of establishments and employees from the dataset's subsectors according to the minimum wage sectors, in each year-province cell. Intercensual values have been estimated with a linear interpolation of the census data.

A.2 Minimum wages

National collective agreements were at the sectoral level established minimum wage scales according to the complexity of tasks performed on the job and the worker's skill level. Istat, the National statistical institute, published a harmonized wage scales for unmarried adults (i.e. 21 years old) blue-collar workers according to the most recent national collective agreement for a selection of industries, with annual frequency. The wages reported by Istat only included the basic pay and the inflation benefit; additional pay components (e.g. productivity premiums) were included only if they were paid to all workers in the sector.

Before 1968 and for most sectors, wages were expressed as daily values. From 1968, wages were expressed as hourly values. To ensure comparability with the previous series, hourly values were converted to daily by multiplying them by eight. This simplifying assumption is coherent with available statistics on hours worked per day: according to a survey conducted by the Ministry of Labour, blue-collar employees worked an average of seven hours and forty-one minutes per day in 1973, with a standard deviation of 14 minutes between the provinces, across all industrial sectors excluding energy, gas and water (Ministero del Lavoro e della Previdenza Sociale, 1974b, pp. 256-259). Monthly wages, instead, have been rescaled to daily by dividing the monthly value by twenty, that is assuming an average number of five days worked in a four-week month. Other methods, including rescaling all wages to monthly or to hourly values, do not alter qualitatively any result in the analysis. These robustness checks are available upon request.

Skill categories were traditionally four for male workers. However, since the 1950s, the collective agreements in several industries introduced additional classes and modified their names to better suit the characteristics of the sector. The source adjusts the data to take into account these differences, so that wages are directly comparable both between industries and across time. The minimum wage considered for each sector is that for the lowest class at any given time.

Until 1972, wages are reported separately for wage zones. From 1962 to 1972, collective agreements identified ten wage zones (classified from one to six, plus two additional special zones for Turin and Milan and Genoa and Rome, and two separate zones for the provices of Arezzo and Ancona). Wage zones comprised one or more provinces, and were further divided in two sets (A and B) that computed the inflation bonus on different price indexes. Table C.2 provides a list of the wage zones with the respective provinces and inflation indexation sets. For the period 1972-1988, national-level contractual wages have been imputed to all provinces.

The main source provided a wide but incomplete coverage of wage zones for most industries. To obtain a balanced panel, missing industry-province cells were filled digitizing another primary source, Istat's *Statistiche provinciali*, an annual publication containing a wide range of statistics at the province level, including a more complete coverage of the contractual wages by wage zone. In a limited number of instances, the combined use of the two sources left some cells empty, in which case the contractual wages were derived directly from the original text of the industry's most recent collective agreement (available in digital format for the majority of industries from the historical archive of CNEL - Consiglio nazionale dell'economia e del lavoro, https://www.cnel.it/Archivio-Contratti). When the most recent collective agreement was not available, the missing industry-province cells were filled through linear interpolation from lead and lag values or adjacent cells. Interpolation was applied to compute only 35 industry-province cells out of 47,196 (i.e. 0.07 percent of the observations).

A.3 Average effective wages

Average effective wages in the manufacturing sector are digitized from an annual publication by INAIL, the national institute for insurance against workplace accidents, titled *Notiziario statistico*. Among other statistics, the publication 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 and ten macro-sectors. To harmonize the series with the minimum wage and industrial census data, I have devised a conversion system that is reported in table C.3. Each year-province mean effective wage is obtained as the average of mean wages across the ten macro-sectors, weighted by the employment shares according to the local industrial composition, in each province-year cell.

This weighting procedure allows to control for changes in composition between macro-sectors, while changes within macro-sectors would be accounted for originally by the source, as long as the frequency of accidents is a function of the number of employees in each sub-sector. The underlying assumption requires that the probability of a temporary incapacitating accident is equally distributed among sub-sectors in each macro-sectors, and that the probability distribution does not change over time. While it is not possible to directly verify this assumption, it appears plausible given that macro-sectors are relatively narrowly defined and they share similarities in the production processes. The main exception is macro-sector 2, which aggregates chemicals, rubber and plastics with paper

& printing and with leather & hide. To ensure comparability with the minimum wage series, I excluded sector 9—which is more appropriately classified in the service sector.

The INAIL series is the only source of effective wage data for blue-collar workers at the province level with annual frequency through the period under consideration. Hence, the series has been routinely used for research that requires spatial disaggregation with relatively high frequency (Salvatore, 1977; Padoa Schioppa, 1991). Nonetheless, the source presents some idiosyncratic characteristics that need to be addressed. First, the series only covers individuals that were temporarily incapacitated due to an accident on the job in the solar year. This selection criteria can introduce several distortions, both in the cross-section and over time. A possible source of distortion arises from the correlation between the probability of an accident and unsobservable workers' characteristics. For instance, if the probability of an accident decreases with experience, young workers will be over-represented in the sample.

To check that this idiosyncratic weighting does not alter the evolution of the wage series, I have compared the mean blue-collar wage (obtained as a weighted average from the province-level data, using population as weights) with other series of industrial earnings that are available in the same time period at the national level. The series have been digitized from quarterly surveys conducted by the Ministry of Labour on a large sample of firms employing ten or more workers. The sources present two series for the period 1962-1974: one only considers 'direct' wages, i.e. average wage rates per hour effectively worked; the second series also includes common additional pay components, such as paid holidays and family bonuses. Figure B.1 plots all wage series for the national average. INAIL wages appear consistently lower than the comprehensive Ministry of Labour series untiul 1974, but it matches the direct series—except possibly for the early years. Most importantly, all series show a strong comovement, even though the gap between the INAIL wages and the comprehensive series tends to decrease, especially since 1969. It is possible that this convergence is explained by the compression of the wage distribution that followed the mininimum wage hike of 1969: if the INAIL data is negatively selected, we would expect it to rise faster than the average industrial wage after 1969.

The Ministry of Labour series between 1975 and 1982 cannot be directly compared with the previous years, due to a switch in sources and computation methods (see figure's note). Hence, it is not possible to know whether the convergence between the two series is entirely spurious or it is partly justified by structural changes in the wage distribution. It is nonetheless possible that the convergence is due to the reform of the wage indexation system in 1975, which produced larger percentage increases for low wages every new quarter. It is in fact well established that the reform of the wage indexation system caused a strong compression of the wage distribution (Manacorda, 2004) and that, by the late 1970s, the 'direct' components of industrial wages (payscale minima and inflation bonus) accounted for over 80% of average blue-collar earnings (Brunetta, Cucchiarelli, and Tronti, 1994, pp. 160-161). Taking into consideration that the 1970s also saw a reduction of pay scale stratification (Regini, 1974, p. 77), it seems plausible that the gap between the two series dropped significantly in the later period.

An additional cross-sectional source of concern about the INAIL series arises from the possibility that the idiosyncratic weights differ between provinces. This is the most relevant threat of distortion for the econometric analysis because it might influence the identifying variation and bias the estimates in unpredictable ways. However, comparisons with the limited available data at the province level dispels such concerns. Starting in 1972, the Ministry of Labour temporarily and occasionally published its comprehensive wage series at a more disaggregated provincial level. The two series show a strong correlation (r>.8) in all available years, suggesting that the idiosyncratic weighting of the INAIL series does not alter significantly the provinces' relative positions. Figure C.4 shows the scatterplots of the INAIL and Ministry of Labour series at the province level for the available years, which confirm the positive correlation between the two series.

Moreover, occasional discordances can also be attributed to the the Ministry of Labour series: the Ministry surveyed only firms employing over ten employees (five in the construction sector), which excluded a significant number of low-wage firms in several manufacturing sectors. Given the variation in the local industrial composition and firm size, it is possible that the Ministry of Labour series overestimates wages in provinces that had a hollowed-out firm-size distribution. This was particularly the case in the South, where native firms were mostly of small size and remained concentrated in traditional sectors, meanwhile state subsidies promoted large industrial plants in heavy sectors. Thus, the INAIL series—which did not select on firm size—could arguably be more representative of the local industry composition and firm size distribution. Hence, it is unclear whether province-level adjustments of the INAIL series are justified.

Nonetheless, the underestimation of average earnings due to the idiosyncratic weighting of the INAIL series can lead to overestimating the minimum wage bite, especially for the early period. To provide a more plausible estimate of the bite, I have adjusted the INAIL series by multiplying the wages by the ratio between the Ministry of Labour series and the INAIL series, for every year. Alternatively, to avoid potential distortions in the period after 1975, I have multiplied the INAIL wages by the national average of the ratio between the province-level Ministry of Labour wages and the province-level INAIL data in 1973. It is worth noticing that such adjustments do not affect the econometric analysis, because they only apply a common scalar to all provinces in each year, which is absorbed by time-fixed effects. They are simply performed to provide a more plausible estimate of the mean minimum wage bite in the descriptive section.

A.4 Youth unemployment

Estimates of youth unemployment from labour force surveys are not available at the provincial level for the period under study. To circumvent this data limitation, I opt to use the number of young people registered at local job centres. Statistics from the local job centres have well-known limitations, due to the administrative nature of the data—in contrast to statistical sources, such as the labour force estimates—as highlighted, for instance, by warnings contained in Istituto Centrale di Statistica (1970, pp. 18-19). On the one hand, not all unemployed would register, especially individuals that searched for jobs that did not require compulsory registration—however, this typically did not apply to blue-collar jobs—, and those that were searching for local jobs through informal networks. On the other hand, not all people registered at local job centres were effectively looking for a job, for registration alone was necessary to obtain unemployment benefits and other allowances. For instance, in 1968 the labour force survey reported that 331,000 individuals under 21 were searching for their first job, but only 150,502 were registered as such in the job centres' lists, that is about 45% of the total.[32] Similarily, according to the responses given to the labour force survey in 1977, fewer than 53% of unemployed individuals under 21 searching for a job (including both first job seekers and other) had registered at a local public job centre.[33] However, in the same year, the number of people under 21 that were registered at job centres as looking for their first job was higher than that estimated by Istat's survey, suggesting that a significant number was not in active search—although it should be also noted that comparisons over time are also affected by changes in the definition of unemployed according to Istat's own labour force survey.[34]

Because of these limitations, registrations at job centres do not allow to precisely estimate unemployment rates. Hence, I will focus on the number of individuals registered, distinguishing between prime age workers and under 21 and, among the latter, between those that had a previous employment and first job seekers, by sex. Since the reconstruction required to combine a range of different sources, I have checked the plausibility of the series by comparing them with aggregate figures at the national level reported in Istat's annual labour statistics publication. The two series exhibit a strong co-movement and are virtually identical until the latest period (after 1975), when my series tends to slightly under report the number of registrations. This can be attributed to the limited selection of monthly registrations that is used for the later period, due to the absence of sources reporting the annual averages at the province level. Nonetheless, to the extent that the under-reporting is common to all provinces, this deficiency should not systematically bias the estimates.

Figure B.2 shows the total number of individuals registered as unemployed, by sex and

[32]Own computations on data from Istituto Centrale di Statistica (1969a, pp. 71, 100).

[33]Own computations on public microdata from the historical quarterly labour force surveys, available from 1977 to 1992, at https://www4.istat.it/it/archivio/206993, last retrieved July 2022. Estimates obtained by computing the share of individuals actively looking for a job that registered at public job centres, between the age of 15 and 21, in the year 1977. Survey respondents are weighted according to the sample probability coefficients reported in the dataset.

[34]Own computations on national-level data from Istituto Centrale di Statistica (1978, pp. 5-7, 94).

category, across all provinces over time. Male prime-age workers were consistently the largest group, but their number remained relatively stable in the long run, oscillating 650 and 450 thousand; female prime-age unemployed remained stable at a considerably lower level until circa 1975, when they started converging to the men's. The number of male first job seekers, instead, grew at fast rates after 1975, converging towards the prime age level. An even steeper increase is shown by women seeking their first job, which overtook the number of prime-age unemployed females. This convergence is particularly remarkable considering that the cohort between the age of 15 and 21 accounted for only 10% of the population. Young people with previous employment experiences also show a tendency to increase towards the end of the period, but much more slowly than first job seekers. These tendencies closely matched national series of unemployment rates by group and sex—see Reyneri (1996, p. 66) and Pugliese and Rebeggiani (2004, p. 80)—, reinforcing our opinion that job-centre registrations provide a reliable proxy for the dynamics of local unemployment.

A.5 Estimates of age groups between census years

School data need to be compared with the population in the relevant age groups to compute enrollment rates, and job centre registrations should be normalized by the number of young individuals in the province. However, annual time series reporting the size of age groups at the province level are not available before 1982. To estimate the number of individuals (male and female) that were aged 14-18 (to compute enrollment rates in upper secondary education) and 15-21 (to normalize job registrations) in each year-province cell between 1962 and 1982, I have first digitized tables from the population censuses of 1961 and 1971, which report the resident population in each province at the time of each census, by age and sex. The census of 1961 defined age as number of years since birth (that is, counting the first year of life as number one), while the census of 1971 computed age as number of birthday anniversaries (thus counting the first year of life as number zero). In order to harmonize the definitions, I have subtracted one to the age reported in the census of 1961. The resulting data has been further harmonized to historical provincial boundaries by aggregating the province of Pordenone (established in 1968) to Udine and the province of Isernia (established in 1970) to Campobasso, in 1971. The same procedure has been applied to Istat's official reconstructions of the intercensual population in 1982: in this case, in addition to the harmonization of Pordenone and Isernia, the province of Oristano (established in 1974) was aggregated to the province of Cagliari. Istat's reconstructions from 1982 appear to show a lag with respect to

These procedures provided an intermediate dataset containing the size of age groups in each province in 1961, 1971 and 1982. To obtain estimates of the size of each age group in the missing years (1962-1970 and 1972-1981) I first identified the year of birth for each age group. Thus, I obtained the number of individuals born in each year from 1898 to 1982 (i.e. the birth cohort). Restricting the analysis to individuals that were no older than 21 in 1962 (born 1941) and no younger than 14 in 1982 (born 1968), I performed a linear interpolation between the benchmark years for each birth cohort, by sex and province. Finally, the size of the 14-18 and 15-21 age groups was computed by summing the number of individuals in the same age range, for each year-province cell.

This methodology has allowed to present new time series of the 14-18 and 15-21 age groups in each province and year, at 1961 historical boundaries. However, some limitations to this reconstruction must be acknowledged. First, the linear interpolation imposes the assumption that any changes in the size of the birth cohorts accrue evenly over time, which could be too restrictive. Changes to the size of a birth cohort in each province can be attributed to deaths and net migrations. Thus, the assumption implies that age-specific death rates are similar and that the probability of migration does not vary with age between birth and age 21. It is obvious that age-specific risk factors can negate this assumption. A more precise procedure would consist in deflating the number of individuals using age-specific death rates and accounting for emigration and immigration rates by age. However, the data requirements for such an analysis cannot be satisfied for the earlier period at the relevant frequency and level of disaggregation, due to the lack of accesible sources. Nonetheless, the focus on short age ranges (lasting five or six years) suggests that age-specific risk factors do not vary significantly: mortality at 14-21 for the age groups considered was extremely low from 1962 through 1982. The probability of emigration did in fact increase in the age range, but as long as this did not vary extremely *between* consecutive birth cohorts the potential bias should not affect the empirical analysis.

It is also necessary to highlight that the census data and the official reconstruction of the 1982 population considers only *resident* individuals in the province. An individual was considered resident only if she was registered at the local population office (*anagrafe*) at the time of the census, and she would be counted by the census enumerators even if she were not present at the stated address at the time of the enumerator's visit. An alternative measure would be the size of the present population, that is the number of people that were present at the address at the time of the enumerator's visit. To obtain the present population from the resident population one needs to add individuals that were present at the time of the visit but were registered in a different province and subtract the number of individuals that were officially resident but absent at the time of the visit. Hence, the resulting totals would inflate the size of the age groups in provinces with negative net migration (in the age group) and reduce it in provinces with positive net migration. While the present population can provice a more accurate description of recent migrations, the choice of the resident population stems from the consideration that the resident data is more conservative, thus appearing more appropriate for the purpose of computing enrollment rates in school, for it is assumed that students mantain a stable residence in the province. Moreover, Ista's official 1982 reconstructions report only resident population, so this choice is forced to ensure correct historical comparability.

Several studies highlighted possible errors in the 1971 census, for the size of a limited number of birth cohorts appears too large when compared with the previous and the following censuses. This is in fact reflected in our data: the size of certain birth cohorts spikes in 1971 with respect to consecutive cohorts, while this does not happen for the census of 1961 and Istat's reconstructions for 1982. To eliminate any potential biases that this might cause, I follow the methodology proposed by (Caselli, Golini, and Capocaccia, 1989) only for the birth cohorts that are relevant for the 14-18 and 15-21 age groups in the period under consideration (1962-1982). These are the cohorts born in 1960, 1950 and

1948. The formula suggested by the authors to correct the error in each affected cohort g is the following:

$$P_g^{71^*} = P_g^{81} \sqrt{\frac{P_{g-1}^{71} \, P_{g+1}^{71}}{P_{g-1}^{81} \, P_{g+1}^{81}} K}$$

Where P is the size of cohort g, the apex indicates the year of the census that reports the information (1971 or 1981), and K indicates a coefficient which accounts for mortality and net migration between the two censuses. The corrections are performed separately for male and female. Using the values reported by Caselli, Golini, and Capocaccia (1989, pp. 9-3), the K coefficient for males is 0.999995 for the 1960 cohort, 1.000034 for the 1950 cohort and 0.999865 for the 1948 cohort. For females, the coefficient is 0.999995, 1.000025 and 0.999950, respectively. The correction does appear to solve the unexpected spikes in the data from the 1971 census: Figure B.3 plots the ratios between the size of each cohort in 1971 with respect to 1961 and with respect to 1982, both with and without adjustments, aggregated at the national level. The spikes that are present in the unadjusted data are not present in the adjusted data, for the relevant birth cohorts. The same check is performed for each province separately, reaching the same conclusion (graphs available upon request). Nonetheless, to check that the correction does not alter the results, regressions are run both both with and without adjustments and results remain unchanged (available upon request).

B Additional figures

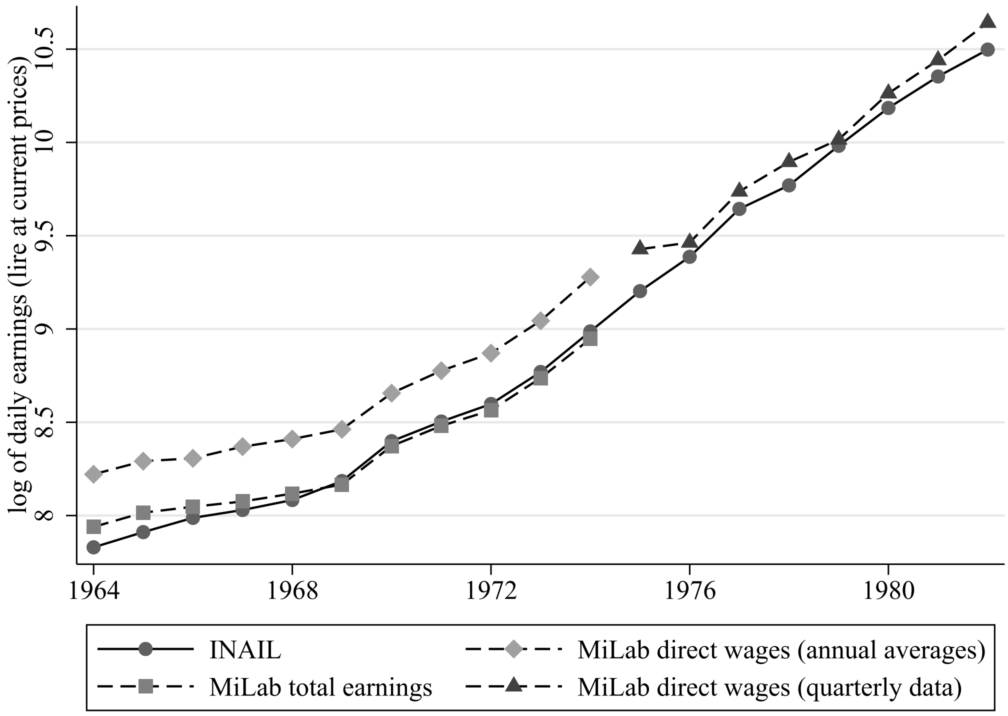
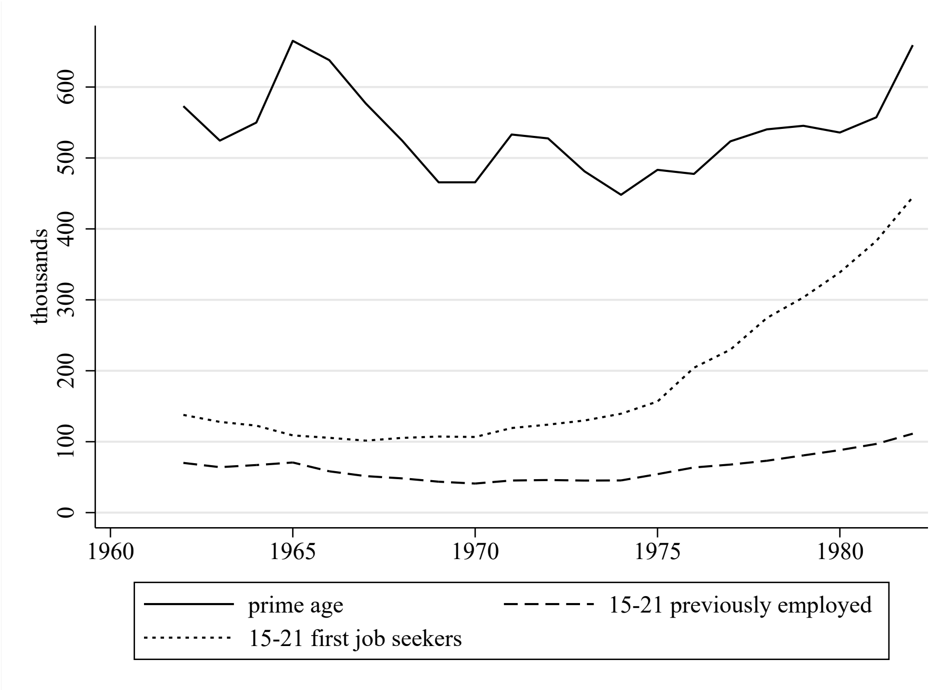
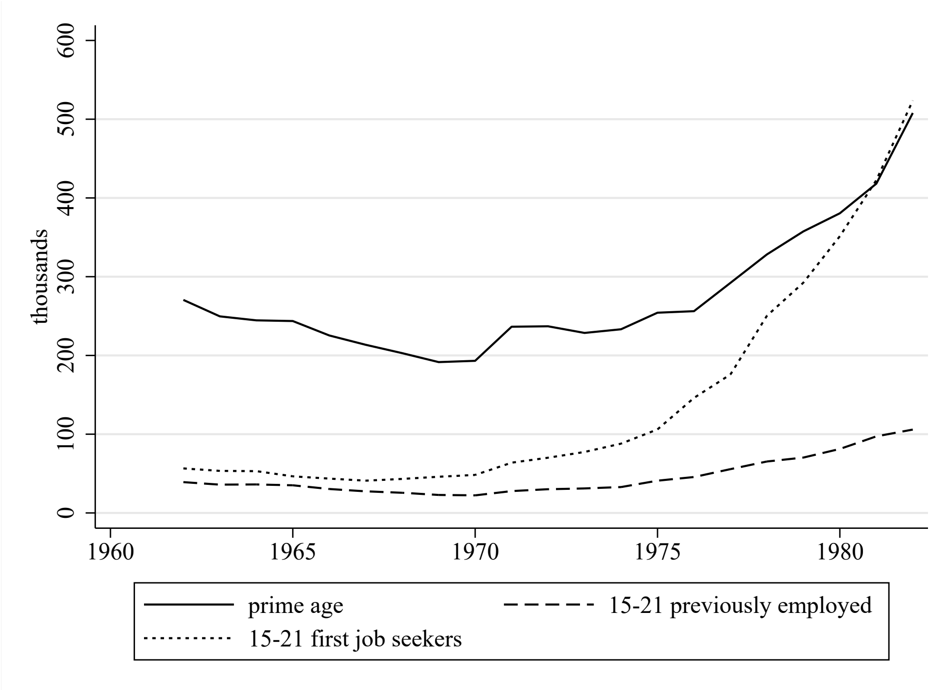


Figure B.1: Comparison of blue-collar mean daily wage series

Log of daily earnings for blue-collar workers at nominal value from different sources. INAIL data include daily wages for blue-collar workers injured on the job, excluding any additional component. The Ministry of Labour series of direct wages include average wages for one hour of effective work. The 'total earnings' series includes also other indirect earnings, including paid holidays, personal and family bonuses, and additional components. Sources and methodology: the INAIL series has been digitized from Istituto Nazionale per l'Assicurazione contro gli Infortuni sul Lavoro (1964)-Istituto Nazionale per l'Assicurazione contro gli Infortuni sul Lavoro (1975) for 1964-1974 and from "Statistiche di base e retrospettive" (1978)-"Statistiche di base e retrospettive" (1986) for 1975 to 1982. Data for 1982 have been linearly interpolated by province between 1981 and 1984. The Ministry of Labour series between 1962 and 1974 have been digitized from Ministero del Lavoro e della Previdenza Sociale (1968, p. 177-179), Ministero del Lavoro e della Previdenza Sociale (1971, p. 262), Ministero del Lavoro e della Previdenza Sociale (1974b, p. 262). Data for 1975 to 1982 have been digitized from a quarterly publication which originally supplemented and later substituted the previous source (Ministero del Lavoro e della Previdenza Sociale, 1975-Ministero del Lavoro e della Previdenza Sociale, 1982). For data availability, the series only refer to the second quarter of each year, except for 1982, which reports annual averages. All Ministry of Labour series presented only hourly wages. Daily wages have been computed by multiplying the hourly wage by the average number of hours worker per day, whenever the information was available (until 1966 included). Afterwards, the number of hours worked per day was obtained by dividing the average number of hours worked per month by 21, that is the average number of hours worked in 1962-1966, which is equivalent to four five-day weeks plus one day of overtime.



(a) male



(b) female

Figure B.2: Unemployed registered at job centres

Total number of individuals registered at job centres (thousands), by sex and category. Prime age workers are between the age of 22 and 60. Sources: see section A.4

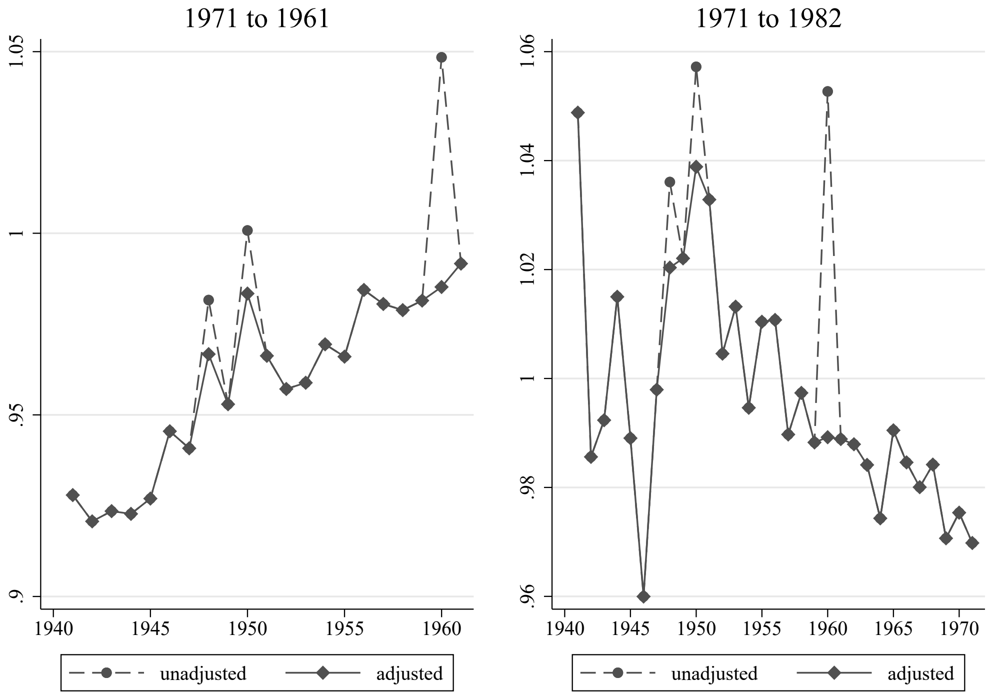


Figure B.3: Comparison between raw and adjusted population data

Ratio of the cohort sizes (1971 to 1961 and 1971 to 1982) with and without adjustments. The ratio is equal to one if the cohort size reported in the 1971 census is the same in the 1961 census (left panel) or Istat's 1982 official reconstructions (right panel). The ratio can differ from one due to deaths and net migration. The adjusted series is the same as the unadjusted except for three birth cohorts: 1948, 1950 and 1960. Adjustment method and sources described in the text.

C Additional tables

Table C.3: 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

Table C.1: Sectors in the census dataset: codes, establishments and employees

census sectors	codes	establishments	employees
Food & beverage	3010	49,272	381,215
Leather & hide	3030	6,680	56,811
Textiles	3040	49,280	541,030
Clothing	3051	97,041	416,447
Footwear	3052	36,390	172,052
Wood & wooden products	3061	68,597	221,062
Furniture	3062	31,072	175,532
Paper & packaging	3070	3,491	94,256
Printing & publishing	3080	13,603	141,020
Iron & steel	3101	2,641	221,354
Forging, pressing, stamping and roll forming of metal	3102	911	24,294
Non-electric engineering, metallic carpentry, secondary smelting	3111	35,026	703,473
Electrical & telecom engineering	3112	5,370	318,125
Precision engineering, goldsmithing & silversmithing	3113	9,210	125,630
Transport vehicles	3115	2,498	335,844
Non-metallic minerals	3120	23,985	330,487
Chemicals	3131	6,230	252,280
Oil derivatives	3132	266	22,579
Artificial textiles fibres	3133	71	47,332
Rubber	3140	5,629	84,568
Plastics	3151	6,619	101,485
Total manufacturing proper		*453,882*	*4,766,876*
Extraction of metallic minerals	2010	140	9,521
Construction	4010	158,553	997,534
Electricity & gas	5010	5,366	134,037
Total industry		*631,226*	*6,011,586*

Four-digit sectors in the census dataset (manufacturing proper and other industries). Establishments and employees refer to the 1971 census, national total from 8,141 municipalities. Tobacco, cinematographic, photographic and phonographic materials, and 'other manufacturing' (totalling 13,305 establishments and 103,618 employees) are excluded from the analysis. Sources: see text.

Table C.2: Provinces by wage zone, 1962-1972

wage zone	inflation set	provinces
0	A	Milano, Torino
0*	A	Genova, Roma
1	A	Como, Firenze, Sondrio, Varese
2	A	Aosta, Bergamo, Bolzano-Bozen, Brescia, Cremona, Gorizia, Imperia, Livorno, Massa-Carrara, Novara, Pavia, Pisa, Savona, Trento, Trieste, Venezia, Vercelli
3	A	Alessandria, Belluno, Bologna, La Spezia, Mantova, Modena, Napoli, Padova, Parma, Piacenza, Ravenna, Reggio nell'Emilia, Verona, Vicenza
4	A	Asti, Cuneo, Ferrara, Forlì, Grosseto, Lucca, Palermo, Pistoia, Rovigo, Siena, Treviso, Udine
4	B	Ancona
5	A	Arezzo
5	B	Ascoli Piceno, Bari, Cagliari, Catania, Frosinone, Latina, Lecce, Messina, Perugia, Pesaro, Pescara, Rieti, Salerno, Taranto, Terni, Viterbo
6	B	Agrigento, Avellino, Benevento, Brindisi, Caltanissetta, Campobasso, Caserta, Catanzaro, Chieti, Cosenza, Enna, Foggia, L'Aquila, Macerata, Matera, Nuoro, Potenza, Ragusa, Reggio di Calabria, Sassari, Siracusa, Teramo, Trapani

Wage zone 0 is separately defined for Milan and Turin and for Rome and Genoa. An agreement to phase out wage zones was reached in 1968 between the confederal labour unions and Confindustria, the employers' association. Wage zones were eliminated by each industry individually through the renewal of collective agreements in the following years. By 1972, all industries had abolished wage zones, introducing same nominal level contractual wages by skill category across whole of Italy, except for the construction sector, where province-level differences in contractual wages remained through the period under study. Source: Istat, *Statistiche industriali*, 1962, tav. 97, p. 151, footnote a.

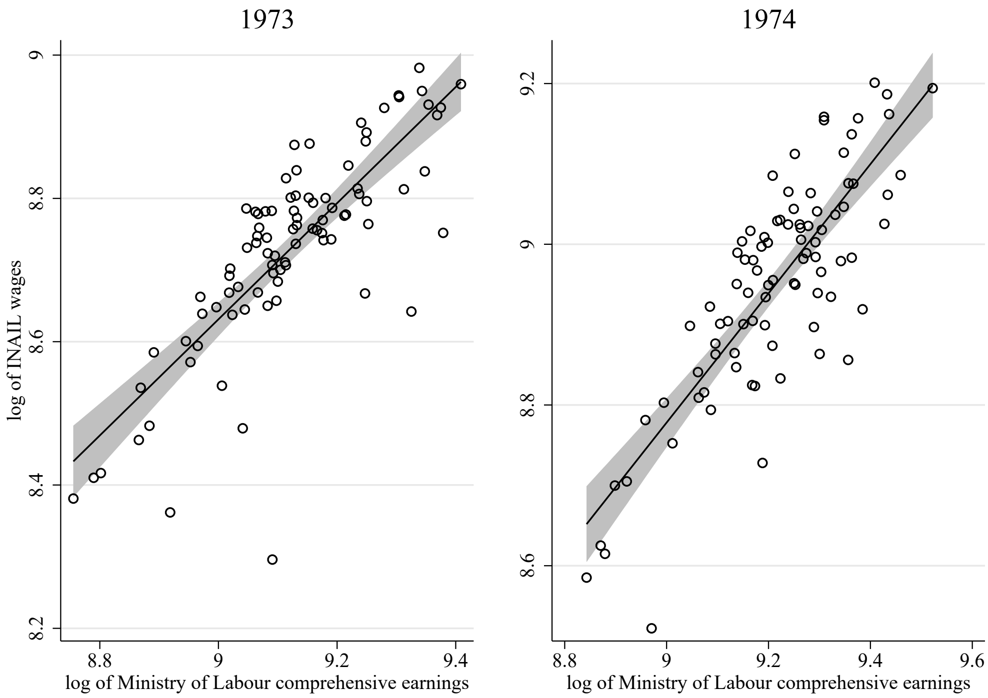


Figure C.4: Blue-collar daily wage series at the province level

Log of daily earnings for blue-collar workers at current nominal value from different sources. INAIL data include daily wages for blue-collar workers injured on the job, excluding any additional component. The Ministry of Labour series of comprehensive earnings include paid holidays, personal and family bonuses, and additional components. Sources and methodology: for the INAIL series see footnote Figure B.1. The Ministry of Labour data are from Ministero del Lavoro e della Previdenza Sociale, 1974a and Ministero del Lavoro e della Previdenza Sociale, 1974b, Tab. OP/2. The Ministry of Labour earnings are for all industry, including construction but excluding energy and water. The solid line represents the linear interpolation of each scatterplot. The shaded area represents the 95% confidence interval.