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Wage Setting Institutions and Internal Migration: The Effect of Regional Wage Equalization in Italy after 1969

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Abstract
Should minimum wages adjust to local productivity? Italy's sectoral collective agreements make no adjustment, as they establish national wage floors irrespective of regional variation in income or cost of living. While some favour its equalizing action, many have argued that this approach causes inefficiencies that include low migration to more productive areas and high structural unemployment in less productive ones. This paper addresses these concerns by studying the spatial equalization of minimum wages in 1972, when the system was first introduced, using an original dataset of labour market variables covering the period 1962-1981. First, the paper presents an augmented gravity model of internal migration showing that spatial differentials in nominal minimum wages were a strong pull factors for both short- and long-distance migration before the reform, but not afterwards. Then, discussing potential mechanisms, the paper shows that the decrease in internal migration during the 1970s was associated with the inception of the spatial mismatches that characterize Italy's labour markets to this day.

JEL Classification: J31, J61, N34, R23.

Keywords: Wage Differentials, Internal Migration, Labor Economic History.

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1 Introduction

Geographical mobility within countries is a key mechanism for the efficient allocation of labour between local markets. Migration from shrinking regions to booming areas allows to contain aggregate unemployment and its social costs, especially during periods of fast technological change that destroys old jobs and creates new ones, often in different locations (Glaeser, Ponzetto, and Tobio, 2014; Berger and Frey, 2016). Moreover, internal migration underpins agglomeration economies, which drive specialisation and growth (Moretti, 2011). However, internal migration has fallen in the past decades across many developed countries (Molloy, Smith, and Wozniak, 2011; Alvarez, Bernard, and Lieske, 2021), despite the fact that spatial differentials in productivity and income have consistently risen, both in the US and in Europe (Rosés and Wolf, 2019; Gaubert et al., 2021). This contradicting behaviour causes significant misallocation of labour and a net loss for the economy (Diamond, 2016), and it hinders employment opportunities and life-cycle trajectory for the affected individuals and their children (Ludwig et al., 2013; Nakamura, Sigurdsson, and Steinsson, 2022).

Understanding what factors might reduce the propensity to migrate is thus crucial to designing effective policy interventions that can improve the efficient allocation of labour between areas. Growing attention has been given to institutions that alter the costs of, or the returns to, internal migration. For instance, rent controls and land regulations can reduce the elasticity of the housing supply, leading positive productivity shocks to be capitalized into higher rents and prices and thus deterring immigration (Moretti, 2011, pp. 1263-64; Jia et al., 2023, pp. 163-68). This paper explores instead factors that can reduce the return to migration, focusing on wage-setting institutions that aim to increase nominal wages in low-income regions, using the historical case of Italy.

Between the 1950s and the 1960s, Italy's economic miracle was accompanied by large migration flows from low-income areas to the country's industrial core in the North-West (Gomellini and Ó Gráda, 2017). These internal flows sustained economic growth and regional convergence in income per capita and employment levels, reducing spatial divides that dated at least to the unification of the country, in 1861 (Daniele, Malanima, and Ostuni, 2016; Felice, 2019). However, in the early 1970s, internal migration rates suddenly dropped, despite the fact that unemployment in low-income regions was rising faster than in the industrial core. Internal migration rates remained at historically low levels for the next four decades (Bonifazi and Heins, 2000), while unemployment and income per capita have continued to diverge, erasing the improvements of the Golden Age (Felice and Vecchi, 2015).

Despite the wide-ranging debate on the present situation and possible reforms, very limited research has been conducted on the influence of the previous system on local labour markets.[1]

[1] The only empirical research on this institution that I am aware of is Blasio and Poy (2017), but the authors studied the introduction of the system in the 1950s, not its repeal, and focused on the impact on employment, not migration. Using a regression discontinuity design, the authors find higher industrial employment in areas with lower

To the best of my knowledge, there is no direct test of the impact of the 1969 reform on internal migration that spans the period before and after.[2] This paper fills the gap in the literature by testing the hypothesis that the nominal equalization of contractual wages in 1969 contributed to the fall in internal migration rates during the 1970s, using newly-digitized data from a range of printed primary sources.

The dataset combines new estimates of contractual and effective wages and of spatial indices of the cost of living with bilateral migration flows for Italy's provinces from 1961 to 1982, at annual frequency. The wage data is estimated from twenty manufacturing industries covering over 97% of industrial workers and 63% of all dependent employees in 1971. The focus on the industrial sector is justified by the effect of the institutional reform—for the spatial equalization of nominal wages originally affected industrial wages—and by evidence that the search for high-paying manufacturing jobs was a major economic motivation for industrial migration in the period under study. A Laspeyres spatial index of the cost of living is computed from official time series of the cost of living at the province level which I make comparable using a new spatial index of the cost of living for a benchmark year, and I validate the results with reconstructions that use alternative methodologies and sources.

The migration data has been digitized from matrices of residential status changes collected by the National Statistical Institute from municipality registry offices and published in annual statistical demographic reports. By using the complete matrices of migration flows at the province level, the paper can distinguish between long- and short-distance migration, which allows to account for unobservable heterogeneity between the two migration patterns. After harmonization to account for border changes of the geographical units, I have complemented the dataset with a range of control variables including, but not limiting to, resident population, local GDP and unemployment measures. The resulting dataset concerns 8,464 dyads for twenty years, totalling 169,280 observations, which are reduced to 167,440 after we exclude migration within provinces (the diagonal of the matrices).[3]

The analysis is divided in two parts. First, I assess whether nominal differentials in contractually-bargained minimum wages represented a pull factor for internal migration. To this end, I estimate a gravity model that is augmented to take into account differences in contractual and effective wages, unemployment rates, price levels, demographic factors and occupational structure. This

nominal wages, suggesting that the differentials were sizeable enough to have significant economic effects. Mauro and Carmeci (2016) have estimated an endogenous growth model, calibrated using aggregate macroregional data, which finds the repeal of the wage zones one of two permanent institutional shocks that explain the end of regional convergence in the 1970s.

[2] Manacorda and Petrongolo (2006, p. 157) suggested that 33% of the increase in unemployment between 1977 and 1988 could be attributed to an excess growth of the labour supply in the South. Even though their model did not include migration, they presented back-of-the-envelope calculations showing that 60% of the change in the relative labour supply could be explained by falling migration, and mentioned the reduction in nominal wage differentials as one leading factor. Following a similar argument, Caponi (2008, p. 4) claimed that the abolition of the wage zones 'undoubtedly contributed to the end of the internal migration between the [South and the North],' even though he suggested that the main contribution to stopping migration flows came from government transfers, and that this was a conscious aim of the Italian parties.

[3] Moreover, note that this version of the paper excludes from the analysis two provinces, Arezzo and Ancona, due to incomplete coverage of the wage data (which would amount to 7,280 observations, 4.3% of the total). Hence, the total number of observations in the dataset is 160,200.

approach follows several recent examples in the specialist literature which have employed gravity models to study internal migration in Italy in the long run (Etzo, 2011; Piras, 2012; Piras, 2017; Piras, 2021). In contrast to this literature, however, I am able to test separately for the role of contractual nominal wages and average effective wages.

Leveraging the longitudinal dimension of the paper to account for time-invariant omitted variables and common trends, I find that the level of contractual minimum wage in the province of destination was a large and statistically significant pull factor, and that the decline in minimum wage differentials can explain much of the decrease in internal migration flows between 1962-1968 and 1975-1981. Moreover, I find that nominal minimum wages at destination lost their significance as a pull factor of migration in the latter period.

Secondly, I discuss potential mechanism that would explain these dynamics. Using the new spatial series of wages and cost of living, I show that nominal wages adjusted to local productivity before the reform of 1969—especially in the Centre-North—, but not afterwards. This led to a significant increase in real wages in low-income provinces, while unemployment levels increased. This analysis shows that the current situation, described by Boeri, Ichino, et al. (2021), was in fact the result of the spatial equalization of nominal wages after 1969.

The paper contributes to different streams of literature. First, it provides—to the best of my knowledge—the first explicit test of the influence of minimum wage equalization for internal migration in Italy, a thesis that has been proposed since the 1970s and that maintains contemporary policy relevance. Second, the paper contributes to the broader historiography of the Italian economy in the 20th century, providing an additional explanation for the divergence in regional economic performance since the 1970s. Third, the paper connects to the wider literature on wage-setting institutions and local labour markets, showing that letting wages adjust to local productivity can stimulate internal migration and increase the efficiency of local labour markets. This evidence also speaks to contemporary debates on the opportunity of equalizing nominal wages across distinct locations, debates that are often dominated by fairness considerations over economic fundamentals.

The paper is organized in six sections. section 2 provides the literature review and historical background, describing secular trends in internal migration and the wage setting reform of 1969; section 3 presents the new dataset with a brief description of the sources and harmonization procedures; section 4 tests the main hypothesis of the paper; section 5 discusses potential mechanisms; section 6 concludes.

2 Literature review and historical background

Traditional models of inter-regional mobility predict that migrants flow from low- to high-income areas and from high- to low-unemployment areas to equalize geographical differences (Harris and Todaro, 1970; Pissarides and Wadsworth, 1989). Hence, the permanent drop in internal migration observed in Italy at a time of rising regional divides has puzzled researchers, spurring several attempts to identify possible causes. A recent review by Piras (2017, pp. 575–578) lists eighteen empirical papers—published between 1977 and 2014—that address this question with a longitudinal approach, even though only a handful include data from before 1970. This stream of research has focused on the estimation of push and pull factors, paying particular attention to labour market variables.[4]

The first conceptualization of the puzzle represented by low internal migration in the presence of high unemployment differentials was proposed by Attanasio and Padoa Schioppa (1991), who argued that workers' low geographical mobility in the presence of widening regional differentials implied significant mismatch in the labour market. The authors listed a series of possible causes, but focused especially on the 1969 reform of the wage-setting system that equalized nominal minimum wages across Italian regions for all manufacturing sectors. During the 1960s, industrial wages were largely set by sectoral collective bargaining, which established minimum wage floors by skill level for both blue- and white-collar workers.

These nominal wage floors, however, were scaled according to regional coefficients, which meant to adjust for differences in local productivity and cost of living, in order to ensure that workers performing the same job tasks within each sector received similar real wages. Following the shift in labour unions' stance with respect to collective bargaining in 1969, the regional scaling system was eliminated, so that sectoral minimum wages became nominally equal across the whole country. According to Attanasio and Padoa Schioppa (1991), this spatial equalization of nominal contractual wages increased the relative standards of living in low-income areas—where price levels were lower—and reduced incentives to migrate. The authors argued that the rise in unemployment differentials was not enough to contrast the shrinking incentives to migrate because destination regions also saw an increase in unemployment rates, while rising housing prices increased the cost of moving.

Even though Attanasio and Padoa Schioppa (1991) did not formally test their hypothesis—relying instead on descriptive statistics for six macro areas—, their argument inspired many of the following studies (for an early example see Faini et al. (1997, pp. 572-74)). In particular, Manacorda and Petrongolo (2006) argued that the centralization of collective bargaining implied that sectoral wages were set according to the conditions of the labour market in the North of the country, and suggested that up to one third of the high unemployment observed in the South between 1977 and 1998 could be attributed to regional mismatch, represented by an excess of labour supply in the South—see also Pagani and Dell'Aringa (2005) for a comparable argument. Similar results had been obtained by Brunello, Lupi, and Ordine (2000) for a longer time span (1951-1996). However,

[4]A pioneering analysis by Salvatore (1977) used time series econometrics to explain short-term fluctuations in migration from the South to the North of the country, and found that relative unemployment rates and industrial wages had a strong explanatory power. This research, however, predated the structural break in internal migration, which became evident only after the mid-1970s.

both papers stopped short of testing the mechanism suggested by Attanasio and Padoa Schioppa (1991).

A longer-term analysis was instead provided by Brunello, Lupi, and Ordine (2001) who, using data for eight Southern regions for the period 1970-1993, showed that real wages in the South since the 1970s had not been affected by local unemployment, as they were tied to labour market conditions in the North. Consequently, the authors argued that the reduction in real wage differentials between the two areas reduced internal mobility. However, since their data only covered the period after the reform of 1969, their analysis did not provide a direct test for the mechanism proposed by Attanasio and Padoa Schioppa (1991).

One of the most recent and explicit analyses of nominal wage equalization for Italy's low internal mobility and high unemployment differentials has been conducted by Boeri, Ichino, et al. (2021). The article compares Italy's wage-setting system with Germany's, using information on wages, prices, productivity, unemployment and migration for 103 Italian provinces and 96 German 'Spatial Planning Regions' from around 2010. The authors show that, in Germany, firm-level bargaining allows wages to adjust to local productivity, which is not possible for most Italian firms. As a consequence, in Germany real wages are higher in high-productivity regions, and migration flows ensure that unemployment rates are similar across all regions. In Italy, instead, the spatial equalization of nominal wages implies that real wages are the same or higher in low-productivity areas within occupations, causing greater local unemployment. Boeri, Ichino, et al. (2021) suggest that this mismatch is not compensated by internal migration because individuals in low-income areas rationally choose to queue for a high-paying local job rather than sustain a costly move to a low-unemployment, low-real-wage area. The result is a loss of efficiency in labour markets, with extra aggregate unemployment and lower labour income.

This and previous evidence pointing to the nominal wage equalization as a cause for inefficient labour markets spurs frequent debates regarding the opportunity of reintroducing elements of geographical variability in nominal wages, either through firm-level bargaining or a new version of the pre-1969 scaling mechanism (Poy, 2015; Poy, 2017). For the proponents of the reforms, a greater flexibility would allow for lower unemployment in the South thanks to greater job creation in low-productivity regions and internal migration to high-productivity regions. On the other hand, opponents often claim that the proposed reforms would effectively recreate the old system, which they describe as iniquitous—see for instance the discussion in AREL (2019) and Damiani, Pompei, and Ricci (2020). Different views stem also from different readings of the data: for instance, Daniele (2021) shows that *average* wages are in fact correlated with local productivity, so a significant divide remains between the North and South due to their different occupational structure. However, these differences translate in lower prices for services in the South, which means that, within occupations, Southern employees enjoy a greater purchasing power than their counterparts in the North. Hence, it seems unclear that a greater decentralization of wage determination might significantly reduce the mismatches in local labour markets.

Figure 1 shows the evolution of gross internal migration rates in Italy from 1902 to 2012 according to the registry offices of the resident population. The graph clearly identifies two long cycles: 1922-1942 (peaking in 1937) and 1952-1978 (peaking in 1963). The first cycle was comparable in magnitude to the second and ended abruptly, possibly due to the chilling effect of the Second World War and to Fascist policies introduced in 1939 to contrast urbanisation—even though their effectiveness in restraining internal migration flows is contested by the prevalent historiography (Treves, 1976).

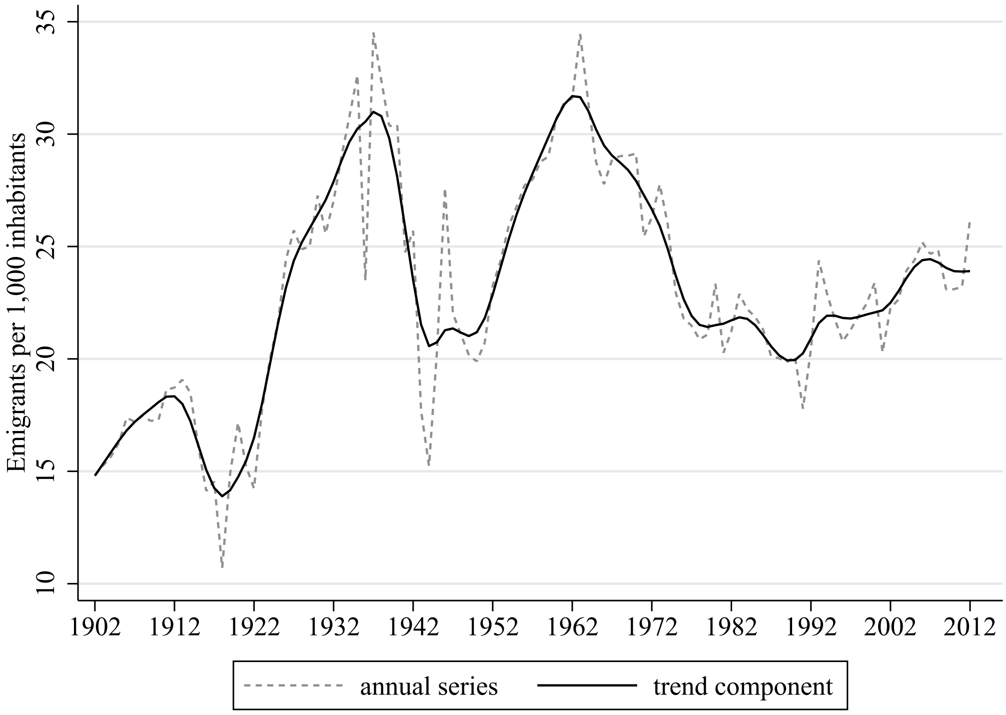


Figure 1: Gross internal migration rates

Total number of internal migrants per 1,000 residents at historical borders. The annual series is the number of individuals registering at the municipalities' registry offices every year and is computed from Istat, *Serie storiche*, Tav. 2.11.1, while the mid-year resident population is computed on data from *Ibid.*, Tav. 2.3, both available for download from `https://seriestoriche.istat.it/` (last retrieved July 2022). The trend component is obtained by applying 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).

The second cycle shows consistently high rates between 1957 and 1971, when at least 28 people per 1,000 inhabitants changed their residence every year. The even higher values recorded in 1961-1963 (when the rate peaked at 34.5‰) are often attributed to the repeal of the Fascist anti-urbanisation laws, which had been *de facto* disapplied since the post-war period but might have prevented many internal migrants to formalize their residence status—this 'clandestine' population was estimated at over one million in 1960 (Gallo, 2012, pp. 156-170). However, the years 1963-64 also coincided with a deceleration of GDP growth and with a contractionary monetary policy that led to higher unemployment—possibly forcing the unemployed to move and adding to the migration flows. Nonetheless, once the extra migration of 1963 is accounted for, the drop after 1971 appears even more stark. In fact, Panichella (2014, ch. 2.1) maintains that, once the spurious registrations of 1962-63 are excluded, the peak of internal migrations happened around 1970, and that the 1970s represented the end of the 'golden age' of internal migrations. Migration rates stabilized at low levels through the 1980s and the 1990s, showing a tendency to grow again only in the 2000s.

The drop in internal migration during the 1970s was also accompanied by changes in their geographical composition. Figure 2 shows the net immigration rates for the five macroregions of Italy: the migration cycle of the 1950s-1960s was characterized by massive long-distance migration from the continental South and the Islands to the North-West, where the industrial core was located. In this graph, the extreme value corresponding to the repeal of the Fascist anti-urbanisation laws can be clearly identified around 1961, as well as the large drop in immigration to the North-West during the 1970s, which is mirrored by the recovery of the South and Islands, two areas that had been characterized by negative net immigration in the previous three decades. Another notable evolution is represented by the North-East—which turned from net negative to net positive around 1965—, while regions in the Centre remained net receivers of internal migrants throughout the century, largely due to the continued migration to the capital city of Rome.

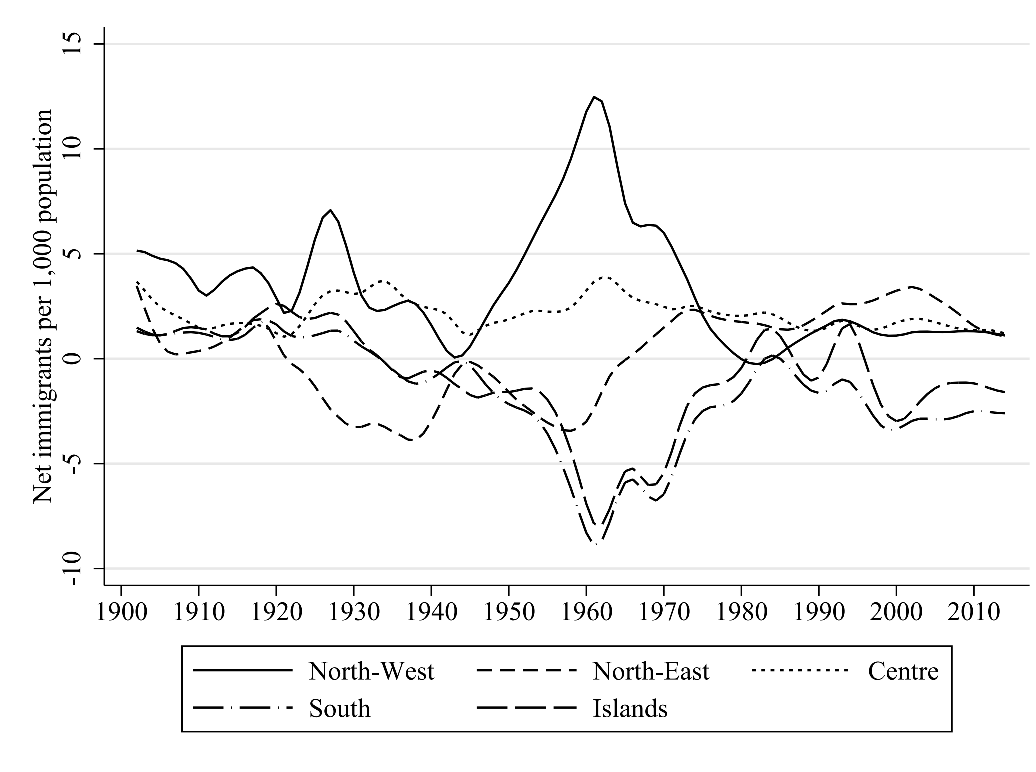


Figure 2: Migration rates by macroarea

Total number of internal migrants per 1,000 residents at historical borders, trend component from a Hodrick-Prescott filter with a smoothing parameter of 6.25. For source and methodology see note at Figure 1.

This sharp drop in long-distance migration from the South to the North motivated researchers to focus on comparisons at the aggregate level. However, more complete long-term reconstructions show that an even larger drop can be identified for migration rates between provinces in the same regions (-47.1% between 1955 and 1995) and between regions in the same macroarea (-50%). In fact, in the same period migration rates between macroareas declined by 32.1%, while rates within the same province dropped only by 12.8% (Bonifazi and Heins, 2000, p. 114). This suggests that, during the 1970s, Italy underwent a generalized migratory decline between provincial borders at all distances, while migration within provincial borders was not affected to the same extreme.

It is possible that the different evolution of migration within and between provinces is due to distinct push and pull factors. Using data from the 2000s, Biagi, Faggian, and McCann (2011) suggest that long- and short-distance migration in Italy respond to different sets of factors: the former appear to be driven by economic fundamentals (income differentials, relative unemployment), while the latter would depend on the quality of life, proxied by the availability of local amenities. Hence, the authors caution against explaining the two types of movements with the same model.

However, it is also possible that the different evolution of migration inside and outside provincial borders is further evidence for the role of the spatial equalization of contractual wage floors. By construction, the equalization of nominal wages could affect only migration between provinces, but not within.[5] Hence, we can hypothesize that some structural factors might have caused a reduction in the Italians' general propensity to leave their place of origin, but the spatial equalization of minimum wages removed a potent economic motivation for longer distance migration.

This hypothesis would also be compatible and separately testable with respect to alternative causes that were originally suggested by Attanasio and Padoa Schioppa (1991) and later assessed empirically by other researchers. This list included a decreasing matching efficiency of the labour market, changing differentials in the cost of living, and shifts in labour demand due to structural transformations. Faini et al. (1997) presented a cross-sectional study of mobility choices in 1995 which pointed to two distinct causes: inefficiencies in regional job-matching and high mobility costs, which they attributed to rent controls and house price differentials. Focusing on the role of the housing market, Cannari, Nucci, and Sestito (2000) found that differential regional dynamics in housing prices were strongly associated with a reduction in internal migration flows both between the South and the North and within the two macro areas, in 1967-1992, but they also highlighted that a large share of mobility remained unexplained.

Focusing on demand factors, Murat and Paba (2002) posited that the structural and technological transformations underwent by the manufacturing sector between the 1970s and the 1980s shifted labour demand in favour of workers already embedded in the local economy—for they would have acquired tacit human capital and more specialist knowledge. However, they also find that wage differentials between locations has a strong predictive power between the 1950s and the 1960s, and much less in the following decades, which would also lend support to our hypothesis. However, their analysis pools data for the two macro-periods and does not account for bilateral migration flows, limiting the possibility to identify causal effects.

[5] Even though we cannot exclude the possibility of spillover effects, for less out-migration from the province can also decrease opportunities for migration within the province.

Hence, despite taking a centre stage in early interpretations of the migration puzzle, as far as I am aware the role of spatial equalization of contractual wages has seldom been directly tested with historical data. To test the hypothesis that the equalization in spatial minimum wages provoked the structural decrease in internal migration observed in the 1970s, I have assembled a new dataset of bilateral migration flows, contractual and effective wages, and local price indexes, which I will describe in the next section.

3 Data and descriptive evidence

3.1 Bilateral migration flows

To observe changes in internal migration before and after the reform of 1969, I have digitized and harmonized annual matrices reporting the number of emigrants and immigrants between any couple of Italian provinces from 1961 to 1981. The data originate from a regular publication of the National Statistical Institute (Istituto Centrale di Statistica, 1964-Istituto Centrale di Statistica, 1985) and from the supplement to the Monthly Statistical Bulletin for the year 1970 (Istituto Centrale di Statistica, 1972). The resulting dataset contains information on migration flows for 8,464 dyads for twenty years, totalling 169,280 observations, including 1,840 observations regarding intraprovincial migration (i.e. when the province of origin and destination coincide).

The matrices were computed by the National Statistical Institute according to regular communication by all municipalities regarding the number of incoming and leaving residents in the previous year. Changes in residential status were recorded by the municipality's registry office every time an individual declared their residence status in the municipality, and they were transmitted to the previous municipality of residence for confirmation of cancellation.[6] Hence, the data originates from administrative sources, and as such it carries both pros and cons that affect the analysis. According to Bell et al. (2015, p. 10), administrative records have the benefit of registering all migration events over time—in contrast to surveys, which have non-complete coverage, and censuses, which typically take snapshots of migrant stocks—, but their reliability is affected by the laws governing resident status.

In the case of Italy, this observation is particularly relevant because of the anti-urbanisation laws limiting changes of residence that had been introduced during the fascist period. As section 2 mentioned, the preservation of the regulation during the 1950s did not prevent migration from happening, but it limited its transparency, causing under-reporting for a sizeable share of the migrant population. To avoid introducing biases in the estimations, I refrain from including in the dataset all years previous to the repeal of the anti-urbanisation laws, even though this implies that we cannot check for the potential effect of the first reform of the wage zone system, which happened

[6] Details on the procedure are described by Istituto Centrale di Statistica (1957).

in the same year.

Caution should also be used with respect to the data recorded after the repeal of the urbanisation laws, for it is possible that they are skewed by the regularisation of the residence status for people that had migrated in previous years. Unfortunately, a distinction between new registration and regularisation is only available for the year 1962, when 27% of all changes of residence was due to regularisation (Istituto Centrale di Statistica, 1965, pp. 282-285). This information, moreover, is only available for the total number of immigrants in each province, without information on the province of origin, making it impossible to correct our data. To account for this eventuality, I run robustness check excluding the year 1962, without obtaining qualitatively different (not reported).

The analyses will be performed using both migration flows and migration rates. To compute the latter, I divide the number of emigrants by the mid-year population in the province, as recorded by the residence statistics from the registry offices. This choice ensures that the source of the population and migration data is the same, reducing the risk of introducing external biases. The mid-year population for year t is the arithmetic average of the population registered at time t and at time $t-1$.

Figure 3 plots the evolution of gross migration rates for all ninety-two provinces, excluding internal migration, from 1965 to 1981. The graph shows that migration rates did not appear to react immediately to the repeal of the wage zone system, but following the reform's completion in 1972 they quickly dropped by one third, and remained at the new low levels through the decade. This drop is slightly larger than that suggested by aggregate data reported in the previous section, and shows a steeper decrease after 1972.

However, section 2 argued that internal migration in this period was a combination of short- and long-distance flows, which were possibly influenced by distinct factors. Hence, it is possible that Figure 3 masks some relevant heterogeneity. In order to check for this, we can first disaggregate the migration flows by type of destination. Figure 4 shows that the year 1972 marked a structural break in the series for all destinations, with migration dropping by circa 30% in the next three years. However, migration within provinces stabilized and quickly recovered towards the end of the decade. Migration to all other destinations, instead, continued falling throughout the 1970s, with limited sign of stabilization at the end of the period.

It is also noticeable that migration peaked across all destinations in 1972, right around the completion of the spatial equalization of contractual wages. We cannot exclude that this temporary spike was caused by individuals' reacting to the wage equalization by returning to their place of origin or moving to a province which offered higher living standards under the new system. It should also be highlighted that migration within regions had started decreasing immediately after the 1968/69 agreements repealing the wage zone system, while all other flows remained stable throughout the transition period. We could speculate that within-region migration is less costly (both in monetary and non-monetary terms), hence it reacted more quickly to the early phases of

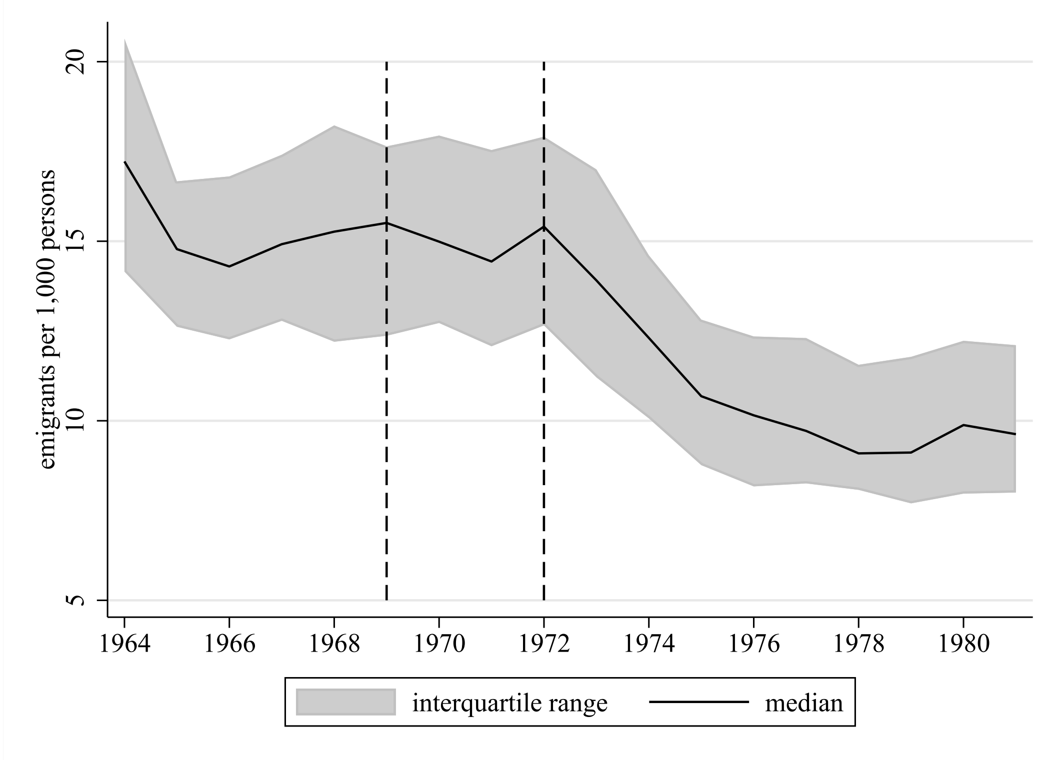


Figure 3: Gross migration rates between provinces

The graph reports the median value and the interquartile range of the gross migration rate (number of emigrants per 1,000 persons). Values are computed as the total number of emigrants from each of the ninety-two provinces, divided by the mid-year population of the province. The graph excludes migration within the province. For sources and methodology, see section 3.

the transition. To conclude, we notice that migration between macroareas experienced a temporary decreased between 1964 and 1966 that is unexplained by our hypothesis. Hence, for robustness, the analysis will also separate migration flows within and between macroareas.

Type of destination, however, is not the sole possible source of heterogeneity. Another relevant factor concerns how migration figures could be affected by outliers. The underlying data shows that the distribution of bilateral migration flows was right-tailed throughout the period, suggesting that a few provinces showed very high migration rates—about two-times larger than the median value (see Figure A.21). Hence, it is important to distinguish whether the reduction in migration observed during the 1970s was due to a generalized decrease in migration from all provinces, or to the decline of mass migration from these few outliers. To begin examining this problem, Figure 5 plots the kernel density distributions of gross migration rates by period. The figure shows that the distribution of migration rates shifted to the left after the repeal of the wage zone system, which can be attributed to a general decrease in migration rates across all provinces. In addition, however, the right tail of the distribution is less fat after 1972, suggesting that mass migration from the outliers also decreased. In fact, the whole distribution is more concentrated in the period after the repeal of the wage zones. This observation suggests that both phenomena happened at the same time, which sustains our hypothesis that the spatial equalization of nominal wages might have had a broad impact on migration flows.

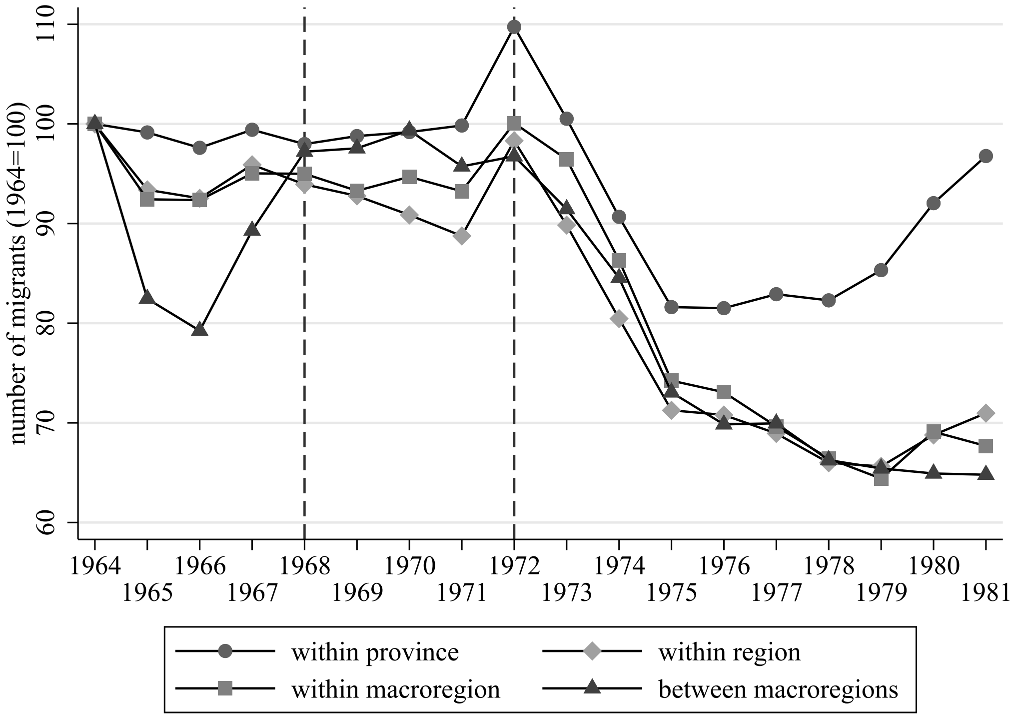


Figure 4: Evolution of gross migration by destination

The graph reports the change in the total number of emigrants for all 92 provinces by type of destination, including migration within the province. To ease the comparison, the series are set equal to 100 in 1964. For sources and methodology, see text and section 3.

This evolution also had a spatial connotation. Figure 6 shows that 64% of provinces were net senders between 1965 and 1968, but this share dropped to 39% in 1973-1981. This change can be attributed to the decline of emigration from provinces in the Centre and in the North-East: 28 provinces in this areas showed negative net migration rates in the first period, but only 8 in the second. Southern provinces, instead, continued to show negative net migration in the second period, but the rates decreased substantially, with only eleven provinces showing a rate larger than five—in absolute level—, down from 21 in the first period. To account for this heterogeneity, the analysis will control for macroarea trends across specifications.

3.2 Estimates of minimum industrial wages

There have been few attempts to quantify the impact of the nominal equalization on spatial wage differentials, possibly due to limited data availability and the almost contemporaneous occurrence of the wage push that started in the autumn of 1969. An early reconstruction by Dell'Aringa (1976, pp. 91–94) found that the coefficient of variation in average effective wages between regions dropped by 30% between 1966 and 1974 (or by 46%, using different wage series), but not much in the following years. The author interpreted these results as evidence that the repeal of the 'wage zones' was a distinct shock to the wage distribution that predated the wage push of 1969. However, it should also be noted that the regional data used by the author would mask part of the

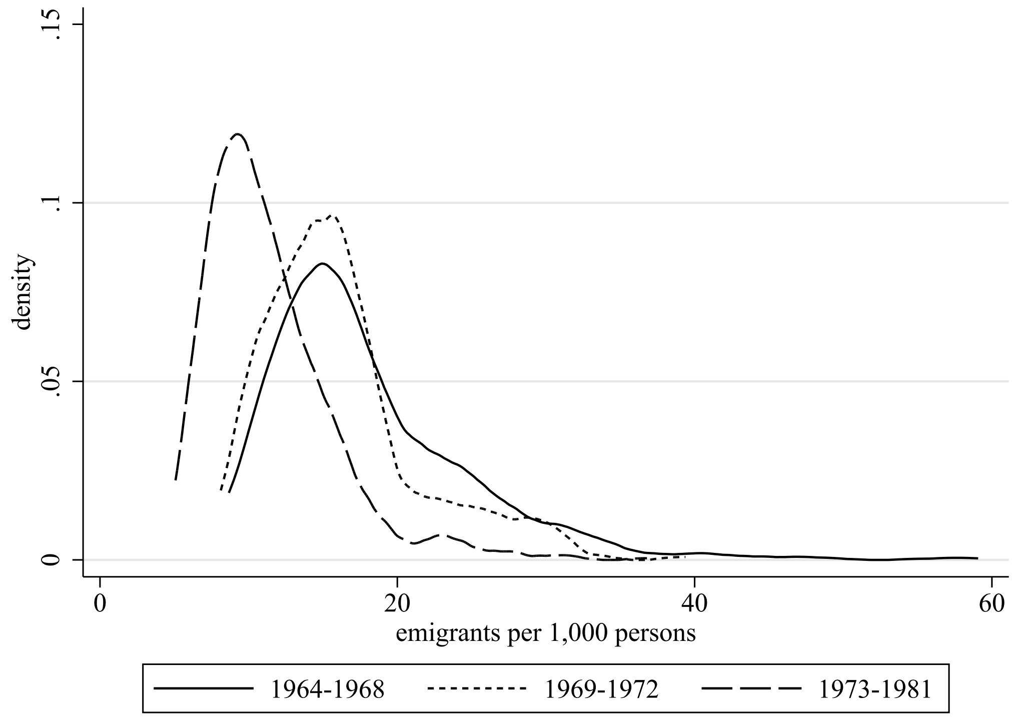


Figure 5: Distribution of internal migration rates by period

The graph reports the kernel density distribution of the gross migration rate (number of emigrants per 1,000 persons) for 92 provinces by period. Values are computed as the total number of emigrants from each of the ninety-two provinces, divided by the mid-year population of the province. The graph excludes migration within the province. The periods are defined as before the repeal of the wage zone system (1965-1968), during the phasing out of the system (1969-1972), and after its complete repeal (1973-1981). For sources and methodology, see text and section 3.

original spatial variation, for most regions were divided into different wage zones, while provinces belonging to two different regions could be assigned to the same wage zone. Moreover, the use of average effective wages rather than contractual minimum wages could bias the computation of the original variation, for this would also depend on spatial differences in workers' skill levels, and on the amount of wage drift accumulated over time in each province.

To evaluate more precisely the potential effect of the spatial equalization on nominal wage differentials across the industrial sector, it is first necessary to observe the variation in contractual wage floors. To this end, I have digitized annual publications from the National Statistical Institute reporting information on the wage floors established by collective agreements for twenty-three industrial sectors (manufacturing proper, mining and utilities, but not construction), employing 97% of industrial workers and 66% of dependent employees.[7] Until 1972, the publication reported the wage floors separately for each wage zone, while afterwards the only the national value for each sector is reported, as by then the repeal of the wage zone system had been completed.

The gradual but quick phasing out of the system can be appreciated from Figure 7, which plots both the median value of the wage floor for low-skill workers (i.e. the entry-level minimum

[7]Values estimated on data from Istat (2014) and Istituto Centrale di Statistica (1973, p. 51). For sources and harmonization methodologies on the minimum wage data see Ramazzotti (2023, especially the methodological appendix).

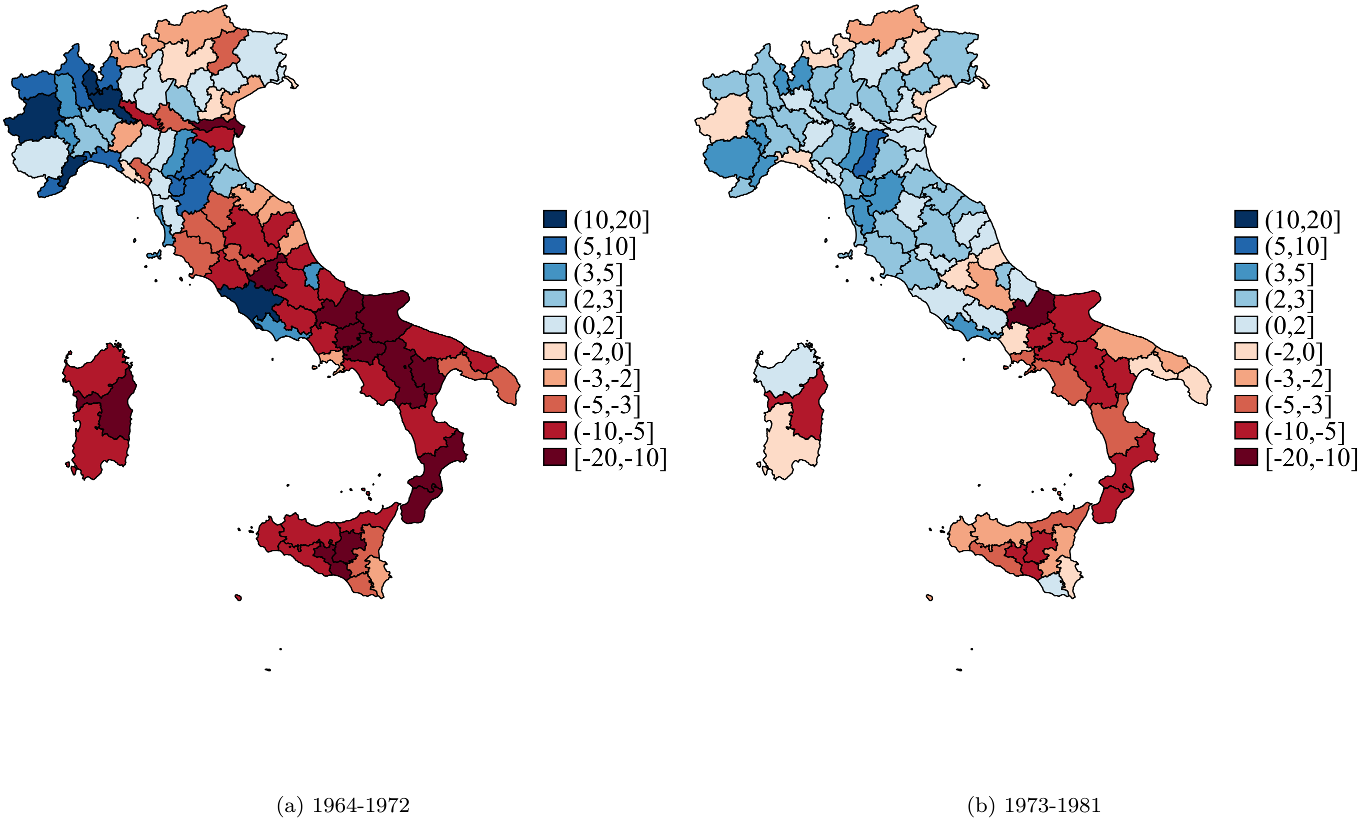


Figure 6: Average net migration rates by period

The maps show the average net migration rates (number of immigrants minus emigrants per 1,000 persons) for each province by period. Computations exclude migration within the province. The periods are defined as before the repeal of the wage zone system (1965-1968) and after its complete repeal (1973-1981). For sources and methodology, see text and section 3.

wage in industry) and the coefficient of variation between provinces. Looking at the latter, we notice that the coefficient of variation dropped by almost 40% between 1968 and 1969, by 35% the next year, and by over 50% per year until 1972, following closely the stipulations of the 1968/69 agreements. By 1973, all sectors had repealed the wage zone coefficients, leading to effectively no spatial variation in nominal minimum wages across the national territory. Meanwhile, the wage push that started during the Hot Autumn of 1969 provoked a substantial increase of real minimum wages, which continued in the next decades. The comparison between the two series shows that the drop in spatial variation started one year in advance to the acceleration of minimum wage growth, supporting the argument that the reduction of spatial wage differentials predated the Hot Autumn.

However, spatial differentials in wage floors did not depend solely on the variation within sector, but also on the local industrial structure. A potential migrant, in fact, would compare the average wage floor available in the province of origin with that offered in the province of destination, which depends on which sectors are present in the two places and the possibility of employment in any of them. Assuming that the probability of employment in any given sector is proportional to the

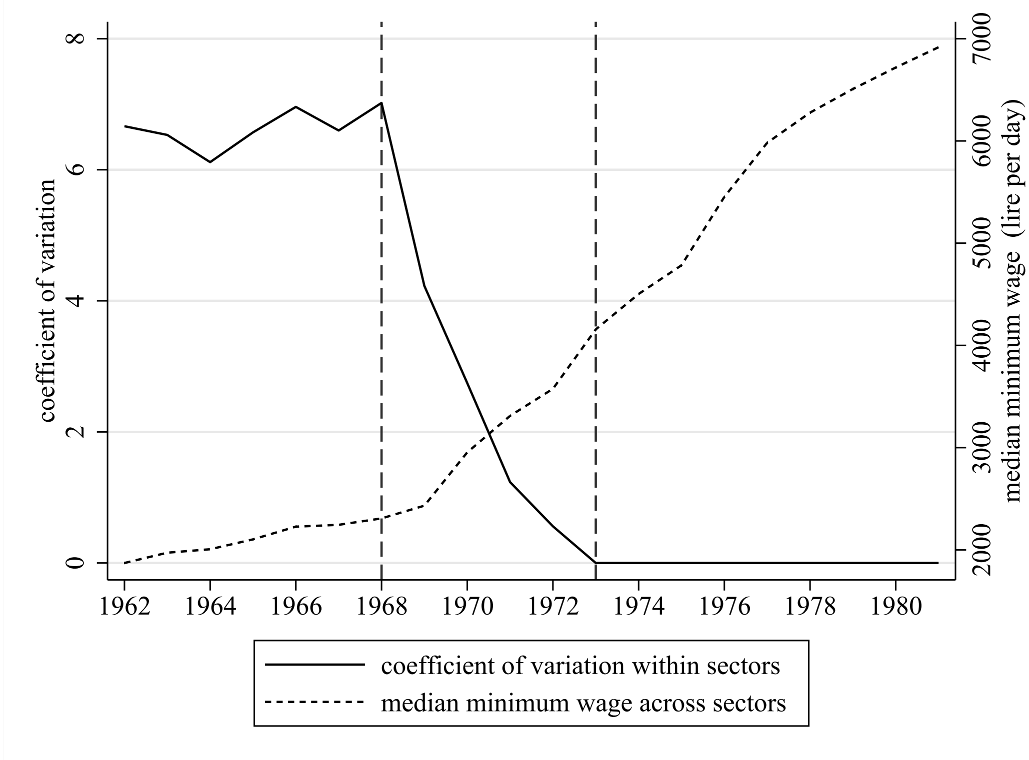


Figure 7: Median minimum wage and spatial variation

Median minimum wage for low-skill blue-collar workers in sixteen industries and ninety-two provinces, and median coefficient of variation between provinces, within each sector. The nominal value of the minimum wage is expressed at constant 1968 prices. The coefficient of variation is the median of the coefficients for each sector. Each sector's coefficient of variation is computed as the standard deviation of the minimum wages for low-skill blue-collar workers across the provinces, divided by their mean value and expressed in percentages. For sources and methodology, see section 3.

number of employees, we can compute the weighted average of the entry-level wage floor $\overline{M}$ in province j at time t as:

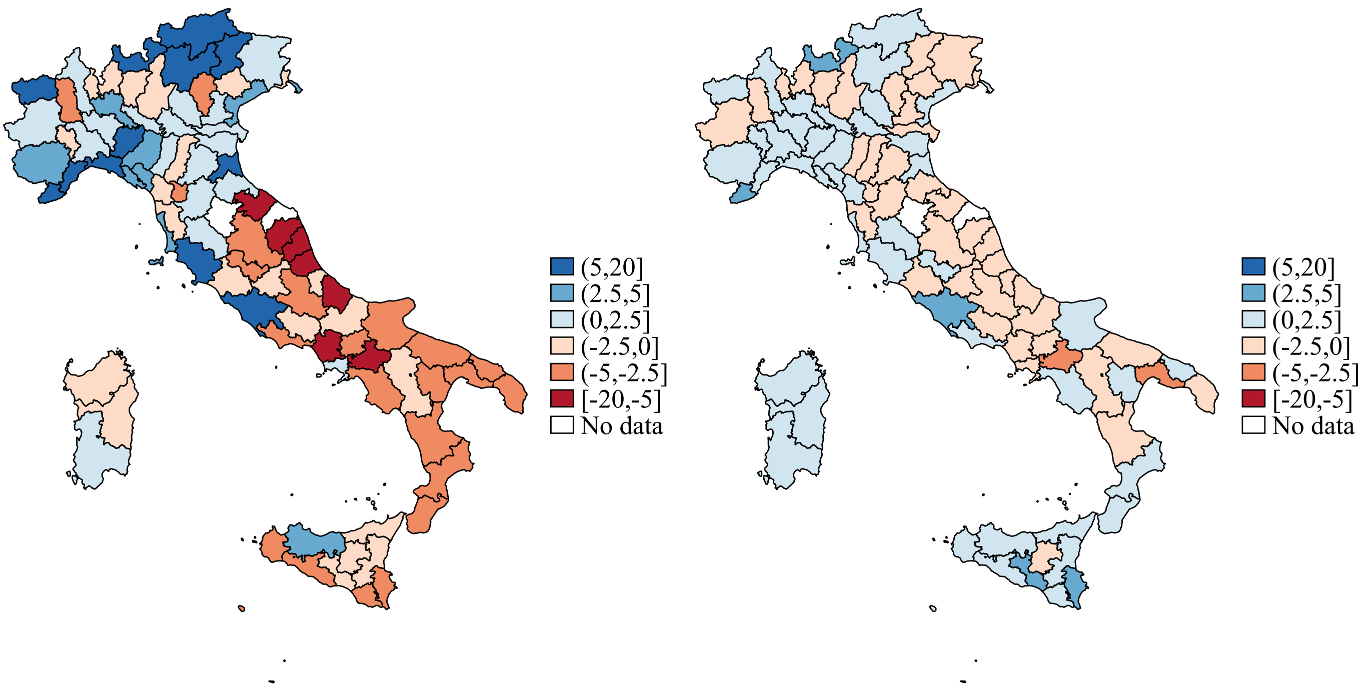
$$\overline{M_{jt}} = \frac{\sum_{i=1}^{24} M_{ijt} \cdot \overline{S_{ijt}}}{\sum_{i=1}^{24} \overline{S_{ijt}}} \quad [1]$$

Where $\overline{S}$ is the share of employees in province j and sector i at time t. Due to the absence of detailed data on sectoral employment by province with annual frequency, $\overline{S}$ is obtained from the linear interpolation of decennial census according to the formula:

$$\overline{S_{ijt}} = S_{ijT} \cdot \frac{(S_{ijT+10} - S_{ijT})/S_{ijT}}{10} \quad [2]$$

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. It is worth noticing that the average minimum wage would be particularly relevant for low-skill migrants, who would have limited specialist knowledge and thus would be indifferent to being employed in any of the industrial sectors considered. Skilled workers would instead give greater weight to the sector for which they are trained and close substitutes. However, since our data does not allow to differentiate between low-skill and high-skill migrants, we focus on the weighted mean industrial minimum wage through the rest of the analysis.

Figure A.22 shows that the coefficient of variation of the weighted mean minimum wage between the Italian provinces also dropped after the repeal of the wage zone system, but some spatial differentials remained during the 1970s, due to the different sector shares in the provinces and the variation in level of minimum wages between sectors. The effect on spatial differentials can be observed from Figure 8, which maps the provinces' mean minimum industrial wage with respect to the national average before and after the repeal of the wage zone system. The figure shows that, in 1962-1968, about 20% of the provinces had a minimum wage either 5% higher or lower than the national average. In 1973-1981, no province deviated as much from the national average. In fact, in the latter period, only fifteen provinces deviated more than 2.5% from the national average, down from 57 in the former period. The analysis will test whether this variation was sufficient to pull internal migration in either periods.



(a) 1962-1968

(b) 1973-1981

Figure 8: Deviation of minimum wages from national average

The maps show the percentage deviation from the national average minimum industrial wage. Provinces' mean minimum industrial wages are the average of the nominal wage floors for low-skill workers in sixteen industries, weighted by the estimated number of employees in each industry. Data are averaged by period. For sources and methodology, see text and section 3.

3.3 Effective wages and the bite of sectoral minima

The hypothesis that the equalization of nominal minimum wages could affect internal migration hinges on their bindingness. If minimum wages were disapplied—or if they were set too low with respect to the wage distribution—, then their influence on the propensity to migrate would be questionable. Researchers agree that wage floors established by contractual agreements have strongly influenced the wage distribution through the decades, even though the partial liberalization of the labour market since the 1990s has decreased their bite in recent years.

Detailed historical assessments, however, are lacking, due to the absence of useful data to estimate the wage distribution. Typical administrative sources (such as matched employer-employee records from social security data) are only available since the mid-1980s. The same limitation applies to common household surveys. The microdata available for the period under study usually cover only a subset of the population or the national territory. To fill this gap, I have digitized aggregate statistics that were originally compiled by the National Institute for Insurance Against Accidents at Work (INAIL), which collected mandatory insurance from all dependent workers in the private sector. INAIL gathered information on the workers insured, including their sector of occupation and wage. Annual publications reported the average wage of blue-collar workers in ten industrial macro-sectors, including industries in manufacturing proper, construction, utilities and trucking.[8]

I used this data to compute the mean industrial wage in the province using the estimates of local employees in each sector as weights, using the formula:

$$\overline{W_{jt}} = \frac{\sum_{i=1}^{10} W_{ljt} \cdot \overline{S_{ljt}}}{\sum_{i=1}^{24} \overline{S_{ljt}}} \quad [3]$$

Where W is the mean wage in macro-sector l and province j, and $\overline{S}$ is defined as in Equation 2, but with employment data aggregated at the macro-sector level.

As before, I assess the evolution of spatial differentials by looking at the coefficient of variation between provinces. Figure 9 shows that spatial differentials were stable before the agreements of 1968/69, decreased sharply immediately thereafter (1969-1970) but rebounded in the following two years. Following the complete repeal of the wage zone system in 1972, however, the coefficient of variation started falling continuously. In 1968, the coefficient of variation was 35% smaller than in 1968. In the same period, the average blue-collar hourly wage doubled, in real terms—a significant growth, but not as steep as that of the mean minimum wage, which almost tripled. This would suggest that the weight of the contractual wage floors in determining the wage level increased.

To establish whether the minimum wages were set high enough to influence the wage distribution, I compute its bite as the ratio between the mean minimum wage and the average industrial wage,

[8]For sources and harmonization methodologies on the effective wage data see Ramazzotti (2023, especially the methodological appendix).

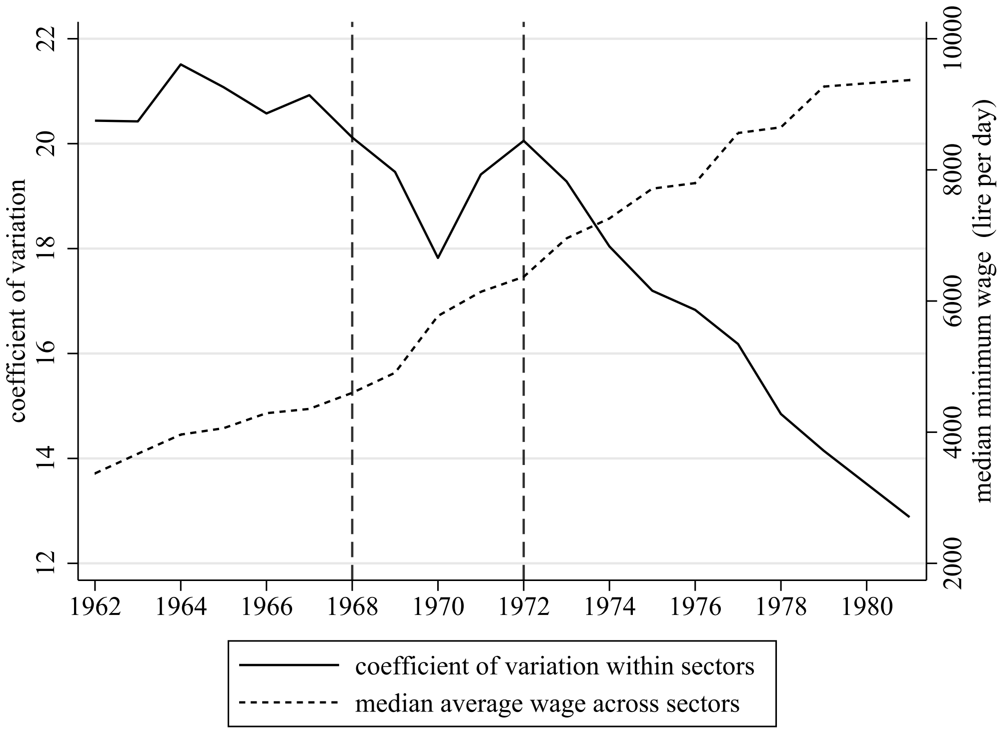


Figure 9: Median average effective wage and spatial variation

Median average effective wage for low-skill blue-collar workers in ten macro-sectors and ninety-two provinces, and median coefficient of variation between provinces, within each sector. The nominal value of the minimum wage is expressed at constant 1968 prices. The coefficient of variation is the median of the coefficients for each sector. Each sector's coefficient of variation is computed as the standard deviation of the average wages for blue-collar workers across the provinces, divided by their mean value and expressed in percentages. For sources and methodology, see section 3.

in each province, expressed in percentage. The mean bite increased from 58% in 1962-1968 to 70% in 1972-1981 (standard devation 6.4 and 7.8, respectively). This increase can be appreciated by the rightward shift of the distribution of the bite values between the two periods, which is represented in Figure A.23. It is also noticeable the wide dispersion of the distributions, with a range of circa 40 percentage points.

These values are high but not impossible. A reference point is the Kaitz index observed across European countries in recent years, even though two distinctions should be highlighted. First, our bite measure is the ratio between mean minimum wages and mean effective wages of industrial workers, while the Kaitz index is the ratio between the statutory minimum wage and the median wage of the whole distribution. Second, the Kaitz index establishes the bite of statutory minimum wages, which represent a flat wage floor for all sectors, whilst our measure is the average bite of the wage floors established by national collective agreements in each (industrial) sector. Both features would tend to bias upward our estimate, because 1) our mean wages originate from effectively right-censored distributions (because they are limited to blue-collar workers) and 2) sectoral wage floors are typically set higher than statutory minimum wages with respect to the wage distribution (Boeri, 2012). Keeping these distinctions in mind, we notice that the increase observed is equivalent to a shift from a 'medium' bite—such as the statutory minimum wage in the UK today—to 'very high'—such as the statutory minimum in France or Portugal (Grimshaw, Dingeldey, and Schulten, 2021, pp. 269-270).

In fact, using data from 2008-2015, Garnero (2018) finds that the bite of sectoral minimum wage floors in Italy today ranges between 74% and 80% of the median wage in each sector, and the average minimum wage floor in manufacturing is equal to 73% of the national median wage (considering also workers in agriculture, mining and services). Our results are very close to these estimates, suggesting that our results are plausible and that the liberalization of the labour market initiated in the 1990s has not attacked the bite but rather the coverage of the contractual wage floors. In other words, the wages of incumbent (covered) workers are as bound by the sectoral wage floors, but those of outsiders (non-covered) workers are not anymore. This interpretation is coherent with the dualistic character of the Italian labour market, which is commonly attributed to the liberalization attempts of the 1990s (Boeri and Garibaldi, 2007).

3.4 Control variables

The analysis of the spatial determinants of internal migration will also take into account differentials in productivity, cost of living and unemployment levels, for they all contribute to the decision to migrate. With low labour mobility, higher local productivity increases wages, which in turn pulls migration from low-productivity regions. With high labour mobility, instead, local productivity is capitalized into rents and, more generally, can increase the cost of living. Finally, the effective wage at destination depends on the level of local average wages and the probability of finding a job. To approximate the inverse of this probability, I control for local unemployment. Since rational migration decisions are based on spatial differentials in real effective wages and the opportunity to get them, all these factors need to be included in the analysis.

3.4.1 Productivity

Productivity differentials are typically attributed to agglomeration economies, which in the case of Italy present a strong historical persistence and are typically cast in terms of a North-South divide. The period under consideration, however, coincides with significant transformations in regional divides. The 1960s continued a process of historical convergence between the South and the rest of the country that had started in the postwar period, but this success was only temporary and the South drifted away in the following decades. The 1970s, instead, saw the acceleration of economic growth in the North-East, traditionally a low-income area, and some central regions. The expansion of small manufacturing led to a stable convergence with the North-West, the traditional industrial core of the country. Several hypotheses have been proposed for this evolution, including the effect of labour market reforms on small businesses, the emergence of industrial districts, and the crisis of large companies.

To account for these dynamics, I have produced original estimates of the value added per worker in the industrial sector and GDP per capita, with annual frequency. Data on value added by macrosector and GDP per capita has been digitized from annual publications by Guglielmo Tagliacarne and, later, the namesake institute.[9] Tagliacarne's estimates are the only reconstructions of national accounting data available at the provincial level in the long run. They used province-level data originating from a range of statistical sources, some unpublished. Even though some reservations have been expressed with respect to their accuracy, Tagliacarne's estimates have been shown to be compatible with recent regional-level estimates that have been produced with current standard methodology for a few benchmark years.

Tagliacarne's data only provides information on the total value added in industry. To compute the value added per worker I have estimated the number of industrial employees in each province-year cell. The estimate has been performed by allocating the number of industrial employees in each region (NUTS 2) as reported in the labour force surveys to the constituent provinces (NUTS 3) using census data as weights interpolated between benchmark years (1961-1971-1981).[10]

3.4.2 Unemployment

Annual estimates of unemployment are not available at the province level for the period under study. To overcome this limitation, I have combined two proxy measurements. The first measurement, which is available at the province level with annual frequency, is the number of individuals registered as unemployed at local job centres, that I have digitized from annual publications of the Ministry of Labour and the National Statistical Institute.[11] Throughout the period under study, the Italian government maintained a centralized system of public placement whereby all individuals seeking work were required to register with local job centres, providing information on their age, education, experience and sector of previous or preferred employment. Employers would hire from the local lists of registered unemployed, filtered by selection criteria. Exceptions were allowed for high skill jobs and for small firms (Musso, 2004, pp. 300-309).

The effectiveness of the centralized placement system was highly questioned, as a large share of hires continued happening through private transactions (Musso, 2004, pp. 332-344). However, the economic and legal incentives to register means that the source can be used as a reasonable proxy for unemployment over time, especially after we control for province fixed effects which would

[9] The publications digitized are Tagliacarne (1963), Tagliacarne (1972), Tagliacarne (1979) and Istituto Guglielmo Tagliacarne (1986). In this version of the paper, the data for 1981 and 1982 are extrapolated from the provinces' trends, and 1978-1979 are interpolated. Note also that the data for the period 1962-1969 has been harmonized with the data for 1970-1980 to avoid a jump in the source around 1970. This smoothing has been obtained by adjusting the data for 1962-1969 by multiplying the series for a province-specific adjustment coefficient. This coefficient was obtained dividing the value of the net domestic product in 1970 as reported in Tagliacarne (1979, pp. 84-91) by the value reported in Tagliacarne (1972, pp. 38-45), for each province.

[10] For the labour force surveys see next section. The industrial census data is from the Datawarehouse CIS, an electronic source originally presented by Istat (1998) and until 2020 hosted at `http://dwcis.istat.it/index.html` (last retrieved 26/11/2019). For a detailed description of this source see Ramazzotti (2023, ch. 5 and methodological appendix)

[11] The dataset combines information from Istat's 'Annuario di statistiche provinciali' and 'Annuario di statistiche del lavoro' and from Ministero del Lavoro e della Previdenza Sociale, 'Statistiche del Lavoro' and its 'Supplementi al bollettino di statistiche del lavoro,' several years.

capture time-invariant heterogeneity in the propensity to register.[12]

The number of individuals registered as unemployed, however, does not allow to compute unemployment rates—not only because of the source's limitations, but also because we lack a suitable estimate of the active population. Hence, I produce an original estimate of local unemployment rates by combining the data from the registrations with data from labour force surveys.[13] The surveys are available at the regional level, one degree of statistical aggregation higher than the province level. The regional data has been digitized from annual publications of the National Statistical Institute, which averaged across quarterly data.[14]

I use this source for two purposes. First, I compute the active population in the province by assuming that the participation rate differed between provinces belonging to different regions but was the same within regions. Hence, I estimate the active population L in province i of region I as:

$$L_{it} = \frac{L_{It}}{P_{It}} \times P_{it} \qquad [4]$$

Where the population P of region I is obtained as $P_{It} = \sum_{i=1}^{n} P_{it}$ with $n \in I$. The second purpose for using the labour force surveys consists in adjusting the number of unemployed people registered at local job centres to filter out potential heterogeneity in the propensity to register within each region. To do so, I assume that the undercounting or overcounting of unemployment in the job centres' lists differs between regions but is constant for provinces in the same region. Hence, the adjusted number of unemployed people $\widehat{U}$ in province i of region I is:

$$\widehat{U}_{it} = U_{it} \times \frac{U_{It}}{S_{It}} \qquad [5]$$

Where S is the number of individuals classified as currently jobless and seeking work in region I according to the labour force surveys. It is important to notice that labour force surveys started in 1959 and changed the definition of unemployed over time, most significantly in 1977, 1984 and 1992. The first definition counted as unemployed all jobless individuals over fourteen who, in the week of the survey, were actively seeking work and willing work. The survey distinguished between properly unemployed (i.e., individuals who had lost their previous job and were looking for a new one, including previously self-employed) and first job seekers. The 1977 reform counted also jobless individuals who were already offered a job which would start at a later date, individuals who

[12] A discussion of the sources on labour market statistics and their limitations for the period under study is given by Ioly (1978).

[13] For the region of Valle d'Aosta, I always use the number of people registered at local job centres rather than estimating unemployment because the publications of the labour force surveys reported the number of unemployed as zero when it was below 1,000.

[14] The publications digitized for this purpose are Istituto Centrale di Statistica (1967, pp. 36-39) for the years 1962-1966, Istituto Centrale di Statistica (1977, pp. 13-16) for 1967-1976, and Istituto Centrale di Statistica (1983, pp. 14-17) for 1977-1981.

would start self-employment at a later date, and individuals who declared to be inactive (students, 'housewives,' retirees) but were in fact seeking work.

To ensure that the data is comparable over time, I have systematically excluded all individuals who declared to be inactive. The resulting series thus include properly unemployed individuals and, for the perios 1977-1982, also currently unemployed individuals who would start employment at a later date. The adjustments performed make the resulting series not directly comparable with official time series of unemployment at the national and (since 1977) macroregional level, which are computed by Istat using the current definition of unemployed. Nonetheless, the trends observed match between my estimates and the official series. I finally use the estimate of the active and unemployed population to compute the unemployment rate u in the province as:

$$u_{it} = \widehat{U}_{it}/L_{it}$$

3.4.3 Cost of living

Despite evidence that prices vary significantly between regions, estimates of the local cost of living are very difficult to obtain for Italy due to data limitations (Vecchi and Amendola, 2017). Since the interwar period, Istat has tracked inflation at the province level using indexes of the local cost of living which were standardized to a common benchmark year for each province. Because of this standardization, the available series allow to compare prices' rates of change between provinces but not their levels. Official estimates of local cost of living have been published only since the late 2000s, and their representativity is limited (Istat, 2009).

Few methods have been proposed to circumvent this data limitation. An early attempt was provided by Campiglio (1986), who obtained access to the unpublished prices that Istat collected to estimate the provincial indexes. The analysis focused only on the prices of foodstuff and was limited to 21 provinces. The author found significant spatial variation in the cost of living, reaching almost 30 percentage points between Milan (in the North West) and Bari (in the South). The estimates, however, were only limited to the year 1986.

An explicit analysis of cost of living differentials over time was instead provided by Caruso, Sabbatini, and Sestito (1993), who used time-series analysis to decompose the Istat indexes between short-term deviations in the inflation rate and long-term changes in the cost of living relative to the national average, aggregating the indexes for eight groups of regions. They found evidence of strong convergence between Northern and Southern regions during the 1950s, stability in the 1960s, and divergence through the 1970s and the 1980s. However, their methodology could only identify the long-term dynamics, but could not quantify the differentials. Alesina, Danninger, and Rostagno (2001), instead, assumed that cost of living differentials were small at the beginning of the series (1947) and computed the 'cumulative price divergence' between the North and the South using the indexes of the cost of living from six and seven provinces, respectively. A limitation of this method is that the underlying assumption is untested and the result—a maximum range of 14 percentage points in 1993—appears smalller than contemporary computations.

The most precise estimate of cost of living differentials with a cross-sectional approach was provided Cannari and Iuzzolino (2009) for the year 2006. Extending Istat's official estimates using information on a wider range of goods and services, the authors found a sizeable differential in the cost of living between the South and the Centre-North, averaging between 15% and 17% but reaching as much as 30% for some regions. These more precise estimates were then used by Amendola, Vecchi, and Al Kiswani (2009) to reconstruct the cost of living differentials between Italy's twenty regions and five macroregions from 1947 to 2011. The methodology proposed by the authors uses the official series of provincial deflators to project back in time the spatial differentials estimated by Istat (2009) and Cannari and Iuzzolino (2009) for 2006.

The authors found that the differential in the cost of living between the Centre-North and the South was already large in the 1940s (circa 10%), decreased briefly in the 1950s, stabilized in the 1960s, grew even larger through the 1970s and the 1980s, peaked in the 1990s at just below 20%, and finally decreased somewhat in the late 2000s. These trends, however, masked significant regional heterogeneity: between the 1970s and the 1980s the relative cost of living increased in the North-West and especially in the North-East, while it decreased in the Centre and especially in the Islands, which converged to the low and stable levels of the continental South.

To my knowledge, the estimates by Amendola, Vecchi, and Al Kiswani (2009) represent the best attempt to reconstruct cost of living differentials in the long term with the data available for Italy. However, two limitations prevent us from using directly their estimates for our purposes. First, the estimates are aggregated at the regional, not the provincial level, which is too coarse for our purposes. Second, by projecting back in time the spatial differentials for 2006, the authors explicitly assume that the structure of relative prices and that of national consumption remained stable in the long run. While this choice is motivated by considerations of data availability and by the aim of linking the historical evolution to the current differentials, for our purposes it would appear more appropriate to use a historical consumption structure.

To this end, I have computed a new spatial index of the cost of living for Italy's historical provinces in the year 1966. The price data has been digitized from Istituto Centrale di Statistica (1968) and covers twenty-one consumer goods and services, which account for circa 40% of the household budget estimated by Istat to build the original indexes of the cost of living. The resulting basket is largely determined by the price of foodstuffs (16 products, 34.3 % of the budget), hence it is most representative of the consumption of low-income households. While a greater coverage would be best, this data limitation should not excessively bias our estimates. First, our research question concerns the propensity to migrate of workers earning close to the minimum wage, so the focus on the cost of living for low-income households is not inappropriate. Second, it aids comparisons with estimates by other authors, who have faced similar data constraints—such as Campiglio (1986), Istat (2009) and one of the indexes by Amendola, Vecchi, and Al Kiswani (2009).

Following Amendola, Vecchi, and Al Kiswani (2009), the prices for each product j of the sixteen in the basket are used to estimate the spatial index L for each province i in the year 1966, using the formula:

$$L_{i,1966} = \sum_{j}^{16} w_{j,1966} * \frac{p_{i,j,1966}}{\hat{p}_{j,1966}} \quad \forall\, i \in \, [1, 92] \tag{6}$$

Where w is the weight of the product j in the national representative household budget, and $\hat{p}$ is the average price of product j across all ninety-two provinces. I then use the annual indexes of the cost of living at the province level to project this spatial index backwards to 1961 and forward to 1981, using the formula:

$$L_{i,t} = L_{i,1966} * \frac{P_{i,t}}{P_{i,1966}} * \frac{\widehat{P}_{1966}}{\widehat{P}_t} \quad \forall\, i \in \, [1, 92] \; \wedge \; \forall\, t \in \, [1961, 1981] \tag{7}$$

Where t is the year, P is the index of the cost of living in province i and $\widehat{P}$ is the average price index across all provinces. Further details on the construction of the indexes of the provincial cost of living are provided in the methodological appendix (available upon request).

4 Analysis and results

4.1 Were nominal minimum wages pull factors of migration?

The first test for the hypothesis that the repeal of the wage zone system affected internal migration consists in estimating the economic and statistical significance of nominal wage differentials as a pull factor for migration flows, using a gravity model approach. The gravity model is a workhorse of applied economic research for the analysis of dyadic flows—most notably bilateral trade (Head and Mayer, 2014). Its application to migration data is almost as old as its general formulation (Hua and Porell, 1979), but in recent years it has received growing attention thanks to the development of stronger theoretical foundations, improvements in computational capability, and greater data availability (Anderson, 2011; Ramos, 2016).

Traditional applications augment the gravity model originally introduced by Lowry (1966) to account for push factors of migration in sending areas and pull factors in receiving areas. A common specification takes the following form (Etzo, 2011, p. 954; Poot et al., 2016, p. 64):

$$m_{jkt} = G^{\alpha_0} \frac{P_{jt}^{\alpha_1} P_{kt}^{\alpha_2}}{D_{jk}^{\alpha_3}} \prod_{h=1}^{n} \frac{X_{hkt}^{\beta_h}}{X_{hjt}^{\gamma_h}} \tag{8}$$

Where m is the gross (or net) migration flow from origin j to destination k at time t, G is a context-specific constant, P is the size of the population at origin j and destination k, D is the distance between the two places, X_{hk} indicate any of n factors that pull migration to destination k, X_{hj} any factor pushing migration out of origin j. Setting $ln(G)$ equal to one, we can linearize Equation 8 as:

$$ln(m_{jkt}) = \alpha_0 + \alpha_1\, ln(P_{kt}) + \alpha_2\, ln(P_{jt}) - \alpha_3\, ln(D_{jk}) + \sum_{h=1}^{n} \beta_h\, X_{h,k} - \sum_{h=1}^{n} \gamma_h\, X_{h,j} \qquad [9]$$

This basic equation can be estimated with OLS. We start the analysis by regressing the log of gross (alternatively net) migrants on the traditional gravity variables (population size and distance), and then we progressively augment the model with the economic and labour market variables shown in Table B.3. In our baseline specification, population at the origin is included as an independent variable, so the dependent variable is the total number of emigrants from province j going to province k at time t. The distance between provinces is computed as the crow flies between centroids (in kilometres).

The economic and labour market variables include the level of the mean nominal minimum wage M in the province of origin and destination, the number of unemployed people U in the province of origin and destination, the level of effective average industrial wages W in the province of origin and destination, and the level of GDP per capita, all expressed in logarithms, and the log difference in the cost of living C between destination and origin. As is common in applications, the economic and labour market variables enter the regressions separately for the origin and the destination, which allows for heterogeneous effects of the same variable as either a push or a pull factor. In addition, I control for origin fixed effects ϕ, destination fixed effects ψ and time fixed effects τ. Standard errors are clustered both at the origin and destination, to account for serial autocorrelation. Hence, the augmented gravity model to be estimated takes the following form:

$$\begin{aligned} ln(m_{jkt}) = \alpha_0 + \alpha_1\, ln(P_{kt}) + \alpha_2\, ln(P_{jt}) - \alpha_3\, ln(D_{jk}) + \beta_1 ln(M_{kt}) + \gamma_1 ln(M_{jt}) + \\ \beta_2 ln(U_{kt}) + \gamma_2 ln(U_{jt}) + \beta_3 ln(W_{kt}) + \gamma_3 ln(W_{jt}) + \beta_4 ln(GDP_{kt}) + \gamma_4 ln(GDP_{jt}) + \\ \delta[ln(C_{kt}) - ln(C_{jt})] + \tau_t + \psi_k + \phi_j + \epsilon_{jkt} \end{aligned} \qquad [10]$$

We hypothesize that the nominal minimum wage at the destination is a pull factor for internal migration, after controlling for the minimum wage at the origin, the level of effective wages and the cost of living differentials. Hence, our null hypothesis is that the coefficient β_1 is not statistically significant from zero. We are instead agnostic about the association between nominal minimum wage at the origin and migration, even though we might expect a negative sign for γ_1.

Columns 1-6 in Table 1 show the estimates for the baseline specification when we additively include the control variables. Column 1 estimates the basic gravity model and finds the expected signs: the coefficient for distance is negative and statistically significant with an elasticity close to one, which is common for migration studies; the coefficient for the population of origin is positive and statistically significant, as we expect since we do not normalize the migration flow by the size of the population; population at destination, instead, is not statistically significant, which is not uncommon in similar studies.

Column 2 augments the basic model with the nominal mean minimum wage at origin and destination. Our estimate for the coefficient of the minimum wage at the destination ($\hat{\beta_1}$) is positive and economically significant: a 1% increase in the nominal minimum wage at the destination is associated with an increase in migration by circa 1.6%, keeping minimum wage at the origin constant.

Column 3 adds unemployment to the gravity model. This is a necessary control because it can be a source of omitted variable bias, for instance because unemployment can be caused by setting the nominal minimum wage too high with respect to the market clearing rate ($Corr(M_{jt}, U_{jt}) > 0$), which in turn increases the pressure to emigrate ($Corr(m_{jt}, U_{jt}) > 0$). The coefficients are found to be statistically significant and with the expected sign both at the origin and at the destination, even though the size of the effect is not very large: a 1% increase in the level of unemployment in the province of origin is associated with a 0.18% increase in emigration; the same increase in unemployment at the destination, instead, decreases immigration by 0.14%. Notice that controlling for unemployment at destination increases by 22% the estimate for the minimum wage ($\hat{\beta_1}$).

Column 4 adds the level of the effective average wage for blue-collar workers in manufacturing. This variable controls that $\hat{\beta_1}$ captures only the specific pull effect of minimum wages on migration, and not firms' greater ability to pay in the province of destination. As can be expected, a higher average wage at destination acts as a significant pulling factor which increases migration, while at the origin it decreases it (even though the estimate is not statistically significant in this case). Its inclusion also partially reduces the size of $\hat{\beta_1}$, but not its statistical significance. Moreover, the response is smaller for the average effective wage than the minimum wage, which suggests that migrants were more sensitive to the tail of the wage distribution than to its average. This reinforces our argument that the minimum wage was a crucial influence for internal migration flows in the period.

So far, all variables have been expressed in nominal terms. Column 5 indirectly controls for real values by augmenting the gravity model with the cost of living differential between destination and origin. The estimated coefficient is positive but it does not significantly alter the rest of the estimated coefficients, and it loses statistical significance as we saturate the model. Column 6, instead, has a more consequential impact by including the level of GDP per capita, which proxies for income in the province. As expected, GDP per capita is a significant pull factor, but it has limited effect as a push factor. Its inclusion, nonetheless, partly modifies the estimates, reducing the size of $\hat{\beta_1}$ and of most other coefficients, albeit not in an extreme way.

The GDP per capita, however, could be affected by the large income differentials between macroareas. To account for this eventuality, column 7 includes macroregion time trends, which increase slightly the size of $\hat{\beta}_1$. Finally, column 8 substitutes the origin and destination fixed effects with dyadic fixed effects. This is the most conservative of the specifications because it de-means the data cross-sectionally for every combination of origin and destination. In conjunction with the time fixed effects, this procedure performs the estimation only on dyadic-specific time variation.[15] Despite the high saturation of the model, the coefficients are almost unaffected.

Table 1: Augmented gravity model with OLS (1962-1981)

	ln(emigrants)							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln(Distance km)	-0.837***	-0.837***	-0.837***	-0.837***	-0.837***	-0.837***	-0.837***	
	(0.0576)	(0.0576)	(0.0576)	(0.0576)	(0.0175)	(0.0576)	(0.0576)	
ln(Population)$_D$	-0.294	0.184	0.345	0.346	0.362***	0.459	0.464	0.464
	(0.362)	(0.317)	(0.317)	(0.298)	(0.0738)	(0.294)	(0.294)	(0.294)
ln(Population)$_O$	0.548**	0.602**	0.396	0.396	0.380***	0.425*	0.440*	0.440*
	(0.267)	(0.284)	(0.253)	(0.253)	(0.0728)	(0.244)	(0.253)	(0.253)
ln(M)$_D$		1.594***	1.931***	1.284***	1.324***	1.292***	1.308***	1.308***
		(0.290)	(0.297)	(0.278)	(0.0768)	(0.278)	(0.277)	(0.277)
ln(M)$_O$		0.178	-0.252	-0.0725	-0.112	-0.122	-0.472**	-0.472**
		(0.302)	(0.281)	(0.274)	(0.0748)	(0.273)	(0.214)	(0.214)
ln(Unemployment)$_D$			-0.147***	-0.134***	-0.133***	-0.130***	-0.129***	-0.129***
			(0.0421)	(0.0400)	(0.00926)	(0.0392)	(0.0392)	(0.0392)
ln(Unemployment)$_O$			0.187***	0.184***	0.183***	0.184***	0.138***	0.138***
			(0.0323)	(0.0322)	(0.00887)	(0.0319)	(0.0251)	(0.0251)
ln(W)$_D$				0.724***	0.724***	0.687***	0.687***	0.687***
				(0.142)	(0.0401)	(0.137)	(0.137)	(0.137)
ln(W)$_O$				-0.201*	-0.201***	-0.217*	-0.192*	-0.192*
				(0.113)	(0.0386)	(0.119)	(0.0985)	(0.0985)
log difference					0.145***	0.136	0.194	0.194
cost of living					(0.0464)	(0.130)	(0.122)	(0.122)
ln(GDP pc)$_D$						0.200**	0.199**	0.199**
						(0.0917)	(0.0920)	(0.0920)
ln(GDP pc)$_O$						0.0887	0.0890	0.0890
						(0.116)	(0.0888)	(0.0888)
Constant	4.732	-17.40**	-16.37**	-17.19**	-17.19***	-20.30***	-16.69	-21.60**
	(5.672)	(7.349)	(7.123)	(6.986)	(1.795)	(6.986)	(10.05)	(9.993)
Origin FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No
Destination FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	No
Dyad FE	No	No	No	No	No	No	No	Yes
Time FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Macroregion trends	No	No	No	No	No	No	Yes	Yes
Clustered SE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R2	0.661	0.662	0.663	0.664	0.664	0.664	0.665	0.869
N	160200	160200	160200	160200	160200	160200	160200	160200

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

[15] Notice that time-invariant, dyadic-specific variables, such as the distance, cannot be estimated for this specification.

4.2 Did the repeal of the wage zones reduce internal migration?

The previous section has shown that, throughout our panel, higher nominal wages at destination pulled greater migration, after controlling for the level of the minimum wage in the province of origin. This implies that a reduction in the nominal *differential* between the minimum wage at the destination and the minimum wage at the origin could reduce migration on the intensive margin.

In fact, minimum wage differentials between provinces dropped, on average, by 75% after the repeal of the wage zones. The reason for such a large drop is clearly depicted by Figure A.24: in the period prior to the abolition of the wage zones, the average differential in nominal wages between dyads was 8% (median 7%), and for one quarter of the dyads the difference was greater than 12%. The repeal of the wage zone system led to a significant compression of the wage differentials: in 1975-1982, the mean differential had dropped to 1.7% (median 1.4%), and for 99% of the dyads it stood at less than 5%.

Could this strong compression of nominal minimum wage differentials between provinces explain the drop in internal migration flows? To quantify this effect, we regress the log difference of the emigrants on the log difference of the minimum wages within each dyad, controlling for all the other variables included in Equation 10, and using dyadic fixed effects and macroregion trends. The specification thus takes the following form:

$$\begin{aligned} ln(m_{jkt}) - ln(m_{kjt}) = \alpha_0 + \alpha_1 \, ln(P_{kt}) + \alpha_2 \, ln(P_{jt}) - \alpha_3 \, ln(D_{jk}) + \\ \rho \, [ln(M_{kt}) - ln(M_{jt})] + \sum_{h=1}^{n} \beta_h \, X_{h,k} - \sum_{h=1}^{n} \gamma_h \, X_{h,j} + \tau_t + \chi_{kj} + \eta_{jkt} \end{aligned} \quad [11]$$

Where X include all controls in Equation 10, except for the level of the minimum wage at destination and at the origin, which are absorbed by the new regressor of interest. The rationale for this specification maintains that migration flows between two comparable provinces (i.e. similar under all characteristics other than minimum wages) depend only on the individuals' idiosyncratic preferences, hence in the aggregate they would net to zero. The existence of a positive wage differential between the two provinces would thus create an extra migration flow from the low-wage to the high-wage province. The coefficient $\hat{\rho}$ recovers the marginal increase in the net migration flow for a one percent change in the difference between the minimum wages.

The results, reported in column 1 of Table 2, imply that a 1% decrease in the difference between the minimum wage at destination and the minimum wage at the origin was associated with a reduction of the net migration flow by 2.3%.

Such a large predicted drop in net migration might be driven by the fact that wage differentials became altogether irrelevant for migration decisions after the repeal of the wage zones. After all, most of the remaining spatial variation in nominal minimum wages after 1972 was caused by differences in the industrial composition of the provinces, so a potential migrant would need not only to change residence, but also sector of occupation if she wanted to receive the higher minimum wage rate. This could reduce the salience of nominal wage differentials to the mind of the migrants.

To test the hypothesis that nominal minimum wages became ininfluential for migration decisions after the repeal of the wage zones, we can interact our regressor of interest with an indicator *Post* that takes value zero for the period before the repeal of the wage zones and one thereafter. We adapt Equation 11, to test whether the differential in nominal minimum wages lost significance in explaining extra migration between provinces after 1972. This specification takes the following form:

$$\begin{aligned} ln(m_{jkt}) - ln(m_{kjt}) = \alpha_0 + \alpha_1\, ln(P_{kt}) + \alpha_2\, ln(P_{jt}) - \alpha_3\, ln(D_{jk}) + \rho_1\, [ln(M_{kt}) - ln(M_{jt})] + \\ \rho_2\, [ln(M_{kt}) - ln(M_{jt})] \times Post_{1972,t} + \sum_{h=1}^{n} \beta_h\, X_{h,k} - \sum_{h=1}^{n} \gamma_h\, X_{h,j} + \chi_{kj} s + \eta_{jkt} \end{aligned} \tag{12}$$

The results are reported in column 2 of Table 2. The estimate for ρ_1 is 2.314, close to the result obtained in the previous specification, and is statistically significant at the 99%, which supports the hypothesis that nominal wage differentials pulled migration before the repeal of the wage zones. The estimate for ρ_2 is, instead, closer to zero and statistically insignificant even at the 90% level. Hence, we cannot reject the null hypothesis that nominal minimum wage differentials were irrelevant for migration flows after the repeal of the wage zones.

Table 2: Minimum wage differentials and migration flows

	ln(emigrants)-ln(immigrants)	
	(1)	(2)
$\ln(\mathrm{M})_D - \ln(\mathrm{M})_O$	2.313^{***}	2.314^{***}
	(0.722)	(0.763)
$\ln(\mathrm{M})_D - \ln(\mathrm{M})_O$ $\times$ Post1972 = 1		0.109
		(1.105)
Post1972 = 1		-0.0572
		(0.0351)
$\ln(\text{Population})_D$	1.009	1.003
	(1.103)	(1.055)
$\ln(\text{Population})_O$	-0.319	-0.296
	(0.293)	(0.330)
$\ln(\text{Unemployment})_D$	-0.530^{***}	-0.486^{***}
	(0.102)	(0.0854)
$\ln(\text{Unemployment})_O$	0.279^{***}	0.332^{***}
	(0.0347)	(0.0412)
$\ln(\mathrm{W})_D$	1.141^{**}	0.906^{***}
	(0.546)	(0.325)
$\ln(\mathrm{W})_O$	-0.434^{**}	-0.705^{***}
	(0.183)	(0.204)
log difference cost of living	0.992^{***}	1.002^{***}
	(0.366)	(0.369)
$\ln(\text{GDP pc})_D$	-0.365	-0.227
	(0.337)	(0.218)
$\ln(\text{GDP pc})_O$	-0.0350	0.108
	(0.120)	(0.158)
Constant	-21.90	-3.244
	(18.04)	(50.01)
Dyad FE	Yes	Yes
Time FE	Yes	No
Macroregion trends	Yes	Yes
Clustered SE	Yes	Yes
Adjusted R2	0.305	0.304
N	160,200	160,200

Standard errors in parentheses

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

4.3 Falsification test: the influence of scaling coefficients before and after the repeal of the wage zones

To further corroborate our hypothesis that the repeal of the wage zone system modified the flow of internal migrants, we can regress the net migration flow on the difference between the scaling coefficients of the wage zone at destination and at origin. The scaling coefficients determined the spatial differentials for minimum wage floors within each sector, and they were largely equal across all sectors, for they had been established by an intersectoral agreement, last updated in 1961.[16] Hence, we would expect that, before the repeal of the wage zones, migration flows were significantly larger between provinces with a greater difference in scaling coefficients, but not thereafter. If we fail to reject this hypothesis, we would falsify our previous argument that the drop in minimum wage significance for internal migration was caused by the repeal of the wage zone system.

To perform this test, I regress the net emigration flow on the difference between the scaling coefficient C at destination and origin, both computed as logpoint differences. To allow the estimate for the scaling coefficient to vary between periods, I interact the variable with an indicator *Post* that takes three values representing respectively the period before (1962-1967), during (1968-1972) and after (1973-1981) the repeal of the wage zones. Due to the fact that the scaling coefficient were time-invariant, we cannot control for both origin and destination fixed effects. Since what matters most is the coefficient at destination, relative to that of the origin, I control for origin fixed effects only. Not controlling for destination fixed effects, however, might lead to biased estimates because our results would not account for differences between long and short distance migration: given the spatial distribution of wage zones, most migration between macroregions was be from a low-scaling coefficient province to a high-scaling coefficient one. If long-distance migration decreased for reasons other than the repeal of the wage zones, we would find a spurious correlation with the scaling coefficients. To address this threat, I include a triple interaction between the log-difference of the scaling coefficients, the period indicator I, and a variable R taking value one if the migration happened within the same macroregion or between macroregions.

$$\begin{aligned} ln(m_{jkt}) - ln(m_{kjt}) = \alpha_0 + \alpha_1\, ln(P_{kt}) + \alpha_2\, ln(P_{jt}) - \alpha_3\, ln(D_{jk}) + \rho_1\,[ln(C_k) - ln(C_j)] + \\ \rho_2\,[ln(C_{kt}) - ln(C_{jt})] \times I_t + \rho_3\,[ln(C_{kt}) - ln(C_{jt})] \times I_t \times R_{jk} + \sum_{h=1}^{n} \beta_h\, X_{h,k} - \sum_{h=1}^{n} \gamma_h\, X_{h,j} + \\ \phi_j + \eta_{jkt} \end{aligned} \quad [13]$$

The results for the coefficients of interest are reported in Figure 10. They show that, before the repeal of the wage zones, migration sorted to provinces with greater scaling coefficients (i.e. where

[16] I follow the list presented in the 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 on the website of CNEL (National Council for Economics and Labour) at `https://www.cnel.it/Archivio-Contratti` (last retrieved July 2021)

nominal minimum wages where higher) than at the origin, irrespective of whether the destination was inside the same macroregion or outside. These results are also economically significant: comparing migration from the same origin to two otherwise similar destinations, if one destination had a scaling coefficient 10% greater than the other, the net migration flow to that destination was between 10% and 11% larger. The average percentage difference in scaling coefficients between dyads was 6% (median 7.13%), and one quarter of the dyads had a percentage difference greater than 11.5%.

During the repeal of the wage zones (1968-1972), the estimate for migration within wage zones halved, while that for migration between wage zones remained constant. This suggests that short-distance migration was very reactive to the repeal of the wage zones, while long-distance migration continued to sort into provinces that originally had higher scaling coefficients. However, by the end of the period both types of migration appeared to have lost any association with the original scaling coefficients: for both cases, we cannot reject the null hypothesis that the difference in pre-1968 scaling coefficients had no association with net migration flows in 1973-1981. This result allows us to pass the falsification test, which corroborates our argument that the wage zones were a significant pull factors of internal migration, sorting net migration flows both in the short and in the long distance.

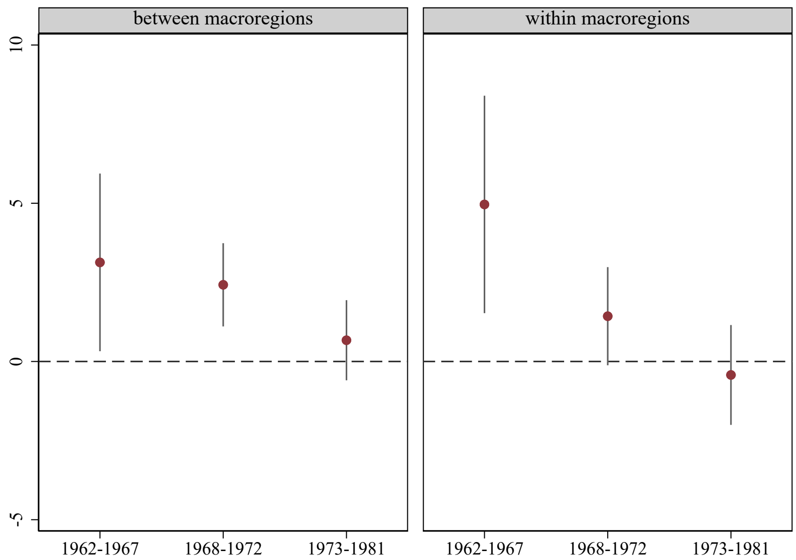


Figure 10: Wage zone coefficients and migration sorting

OLS estimates regressing net emigrants (log-point difference of emigrants and immigrants) on the triple interaction between the log-point difference of the wage zone coefficient between destination and origin, an indicator for time period, and an indicator distinguishing migration within and between macroregions. The regression adjusts for origin fixed effects and a vector of time-varying controls including distance between the provinces and population, unemployment, average industrial wages and GDP per capita both at province and destination, and the difference in the cost of living. The solid vertical lines indicate 95% confidence intervals from standard errors clustered at the origin and destination.

5 Mechanisms and discussion

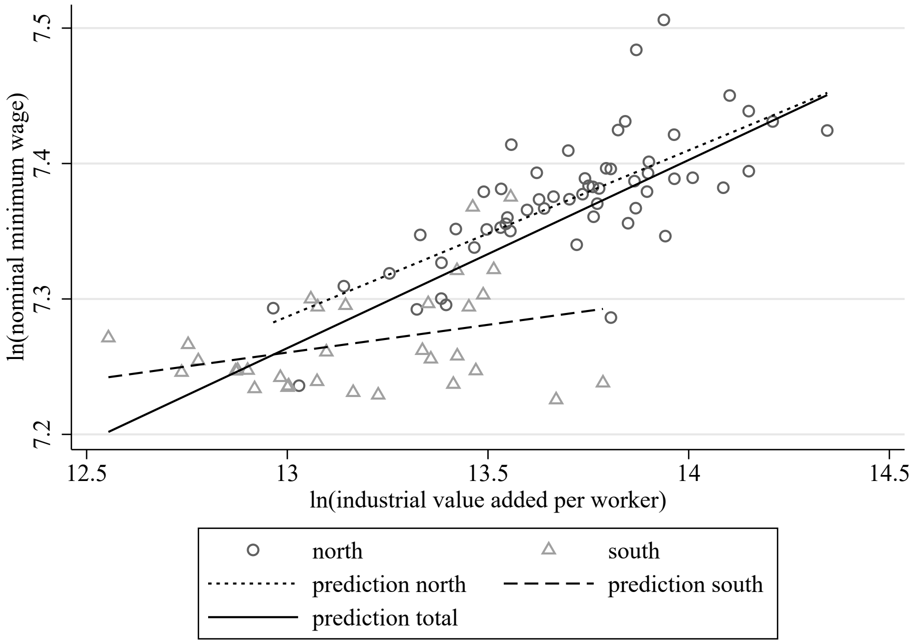
The previous section has shown that an augmented gravity model of internal migration is able to closely predict the drop in internal migration caused by the decrease in minimum wage differentials between Italy's provinces between 1962-1981. Moreover, interacting the minimum wage differentials with an indicator for before and after 1972 has shown that nominal minimum wages lost their role of pull factors for internal migration after the repeal of the wage zones, and migration stopped sorting into high-minimum wage provinces shortly thereafter. This section discusses the potential mechanisms that explain this behaviour, and tests whether the repeal of the wage zones can be considered a root cause for Italy's characteristic spatial mismatches in the labour market.

5.1 The decoupling of minimum wages and productivity

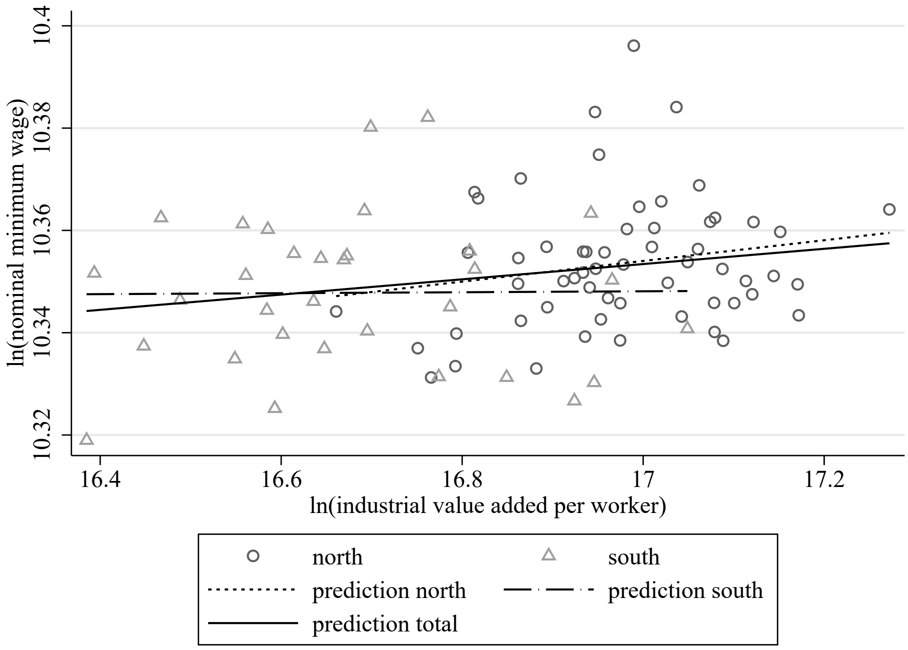
The hypothesis that the spatial equalization of nominal minimum wages causes spatial mismatches hinges on the argument that it decouples wages from productivity. An explicit test of this hypothesis has been provided by Boeri, Ichino, et al. (2021), using a comparison with Germany today. However, no study has ever tested whether this decoupling was originally caused by the repeal of the wage zones. Our dataset allows to perform such a test.

First, I provide some descriptive evidence on the relationship between nominal wages and productivity at the beginning and at the end of our period. Throughout this section, productivity is computed as the value of industrial value added per worker according to the procedure described in section 3.4.1. Figure 11 describes the relationship between the nominal minimum wages and productivity in 1962 and in 1981 by separately plotting the scatterplots at the province level and fitting both the whole graph and the two groups of provinces in the Centre-North and in the South with linear predictions.

The graph for 1962 shows a strong positive relationship between the two variables at the national level (the solid line), suggesting that the wage zones were effective in adapting nominal wages to productivity differentials on a national scale. In 1981, after the repeal of the wage zones, we find instead a very weak correlation, illustrating how the spatial equalization of nominal minimum wages created a disconnect with respect to local productivity differentials. In fact, by construction, all remaining correlation between minimum wages and productivity in 1981 is due to differences in the industrial structure between the provinces. Looking at the sub-national level, we notice that the disconnect between nominal wages and productivity holds particularly for the provinces in the Centre-North. Provinces in the South, instead, already showed a lower correlation between nominal minimum wages and the local industrial productivity. Whilst the dataset does not include information on the nominal minimum wages *before* 1962, we can speculate that the 1961 reform of the wage zones—which reduced the number of wage zones in the South—already produced an



(a) 1962



(b) 1981

Figure 11: Nominal minimum wage and value added

Scatterplot of nominal minimum industrial wages and value added per employee in industry, in ninety-two provinces. For sources and methodology, see text and section 3.

excessive compression of wage differentials within this area prior to their repeal in 1968.

These graphs, however, simply show that the repeal of the wage zones was effective in eliminating the majority of spatial variation in nominal contractual wages between Italian provinces. What was the effect on their *real* value? Figure 12 presents the correlation between the real value of the local minimum wages and local productivity. The real value is obtained by deflating the nominal minimum wages using the spatial price indexed described in section 3.4.3. The data for 1962 show a positive correlation at the national level, even though this is less steep than in the case of nominal differentials. However, this is to be expected, since higher productivity is commonly not entirely passed through to real wages but rather partly captured by greater rents. The data from 1981, instead, show an overall negative correlation between real wages and local productivity, similarly to the situation described by research on present time data. The inversion of this relationship implies a crucial modification in the spatial equilibrium of the Italian labour market, hinting to the 1970s as the period when the current disconnect between real wages and productivity first appeared.

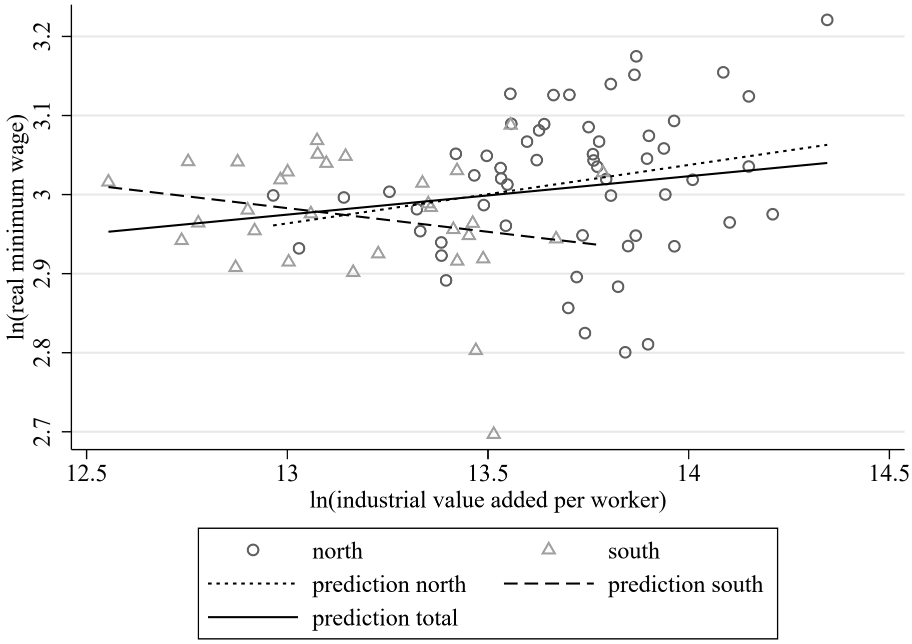
Distinguishing between provinces in the Centre-North and in the South, nevertheless, suggests a more complex interpretation: the correlation between real minimum wages and local industrial productivity was positive in the Centre-North in 1962, but negative in the South, reinforcing our argument that the wage zone system had already provoked an inversion of the relationship within Southern provinces. This distinction would carry over to the end of the period, when the correlation between real minimum wages and productivity remained strongly negative within Southern provinces and became mildly negative within provinces in the Centre-North.

The simple correlations depicted in the graphs for the beginning and the end of the period, however, can be influenced both by changes in the spatial variation of minimum wages within sectors, and by changes in the sectoral composition of industrial employment within provinces. Moreover, it is unclear whether these two years are representative of the period before and after the repeal of the wage zones, respectively. Hence, to formally test the the hypothesis, I regress the nominal minimum wage M on the value added V in the province i, interacted with an indicator variable for the year. Moreover, I control for the local cost of living and for the number of workers employed in each of the sectors that I used to compute the mean minimum wage in the province.[17] Thus, we estimate by OLS the following two specifications, which include also province and time fixed effects.

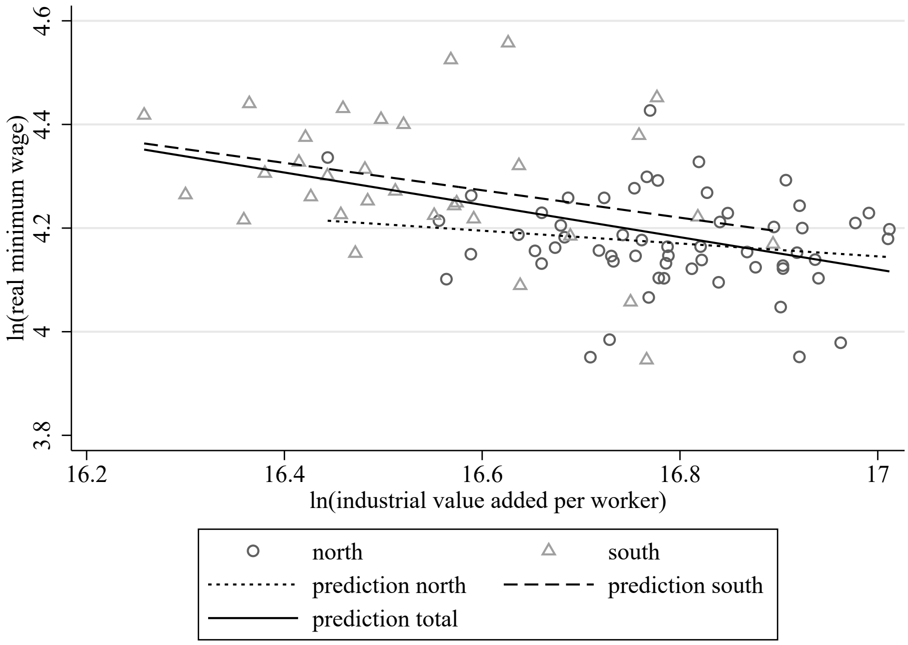
$$ln(M_i) = \alpha + \sum_{y=1962}^{1981} \beta_y \, ln(V_i) \times 1(Year = y) + X'_{it}\gamma + \tau_t + \phi_j + \sigma_{it} \qquad [14]$$

Figure 13 shows the coefficients β_y estimated for each year. It shows that, during the wage zone system, there was a positive and statistically significant association between the minimum wage and industrial productivity in the province, which disappeared during the transition period and turned negative after 1972. This dynamic evolution clearly shows the inversion of the relationship between minimum wages and productivity that we had discussed previously by comparing 1962 and 1981. Moreover, controlling for the sectoral composition of industrial employment ensures that

[17]The number of workers employed in each industrial sector is obtained from the census of 1961 at the province level and extrapolated using the annualized rate of growth of each sector at the national level, obtained from the decadal rate of growth with respect to the census in 1971 and in 1981. This procedure, in contrast to a simple linear interpolation between census years, allows to filter out the local endogenous response in employment to minimum wage changes.



(a) 1962



(b) 1981

Figure 12: Real minimum wage and value added

Scatterplot of real minimum industrial wages and value added per employee in industry, in ninety-two provinces. For sources and methodology, see text and section 3.

the effect is driven by the compression of spatial differentials within sectors, rather than by changes in the industrial composition. Controlling for the local cost of living ensures that our estimates are not biased by the different evolution of the cost of living between provinces.

However, our descriptive analysis suggested that results at the national levels could obfuscate different dynamics between macroregions. To check whether this is the case, we run a similar

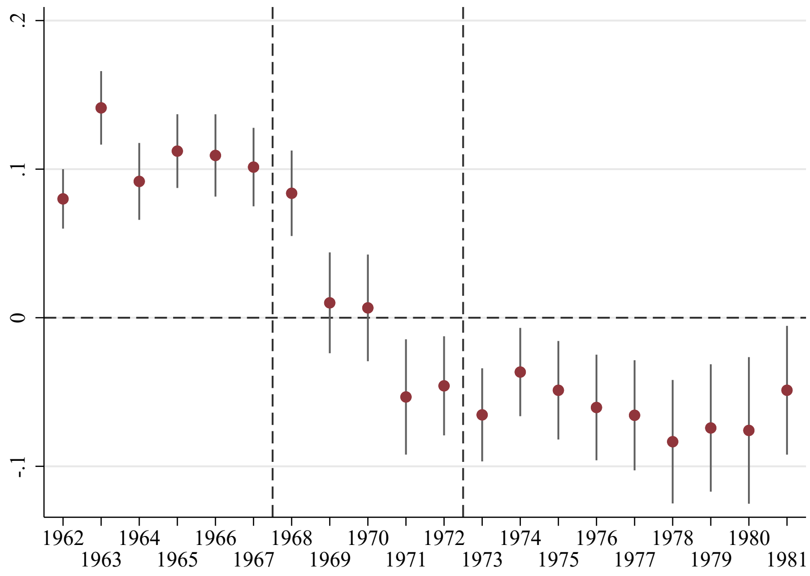


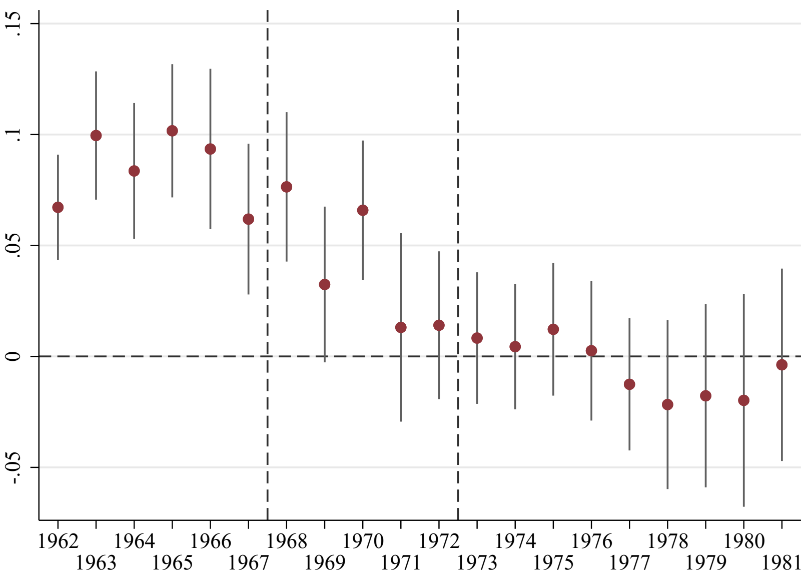
Figure 13: Minimum wages and productivity

Association between the nominal minimum wage and value added per employee in manufacturing, controlling for province and time fixed effects, composition of industrial employment in the province by subsector and local cost of living. Vertical solid lines represent the 95% confidence intervals obtained from standard errors clustered at the province level.

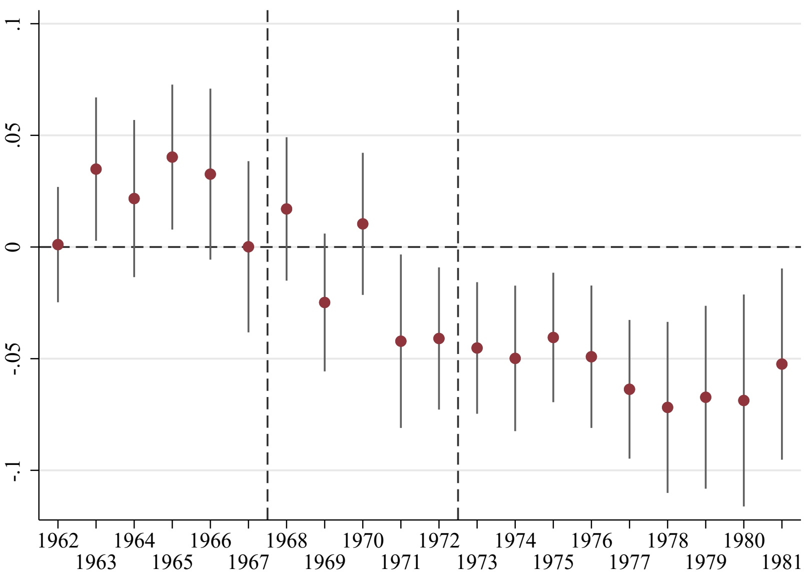
estimation, this time including an indicator S for whether the province was located in the Centre-North or in the South.

$$ln(M_i) = \kappa + \sum_{y=1962}^{1981} \lambda_y \, ln(V_i) \times 1(Year = y) \times S_i + X'_{it}\mu + \tau_t + \phi_j + \varsigma_{it} \qquad [15]$$

Figure 14 plots the coefficients λ_y separately for provinces in the Centre-North and in the South. These estimates largely support our interpretation: with the wage zones, there was a stable positive association between minimum wages and productivity in the Centre-North, which quickly disappeared after their repeal; in the South, instead, there was a much weaker association in the first period, and the spatial equalization after 1968 established a negative association, whereby provinces with greater value added per employee offered lower entry-level minimum wages.



(a) Centre-North



(b) South

Figure 14: Minimum wages and productivity, by macroregion

Association between the nominal minimum wage and value added per employee in manufacturing, interacted with an indicator for whether the province is located in the Centre-North or in the South, controlling for province and time fixed effects, composition of industrial employment in the province by subsector and local cost of living. Vertical solid lines represent the 95% confidence intervals obtained from standard errors clustered at the province level.

5.2 Did wage equalization create spatial mismatches in effective wages?

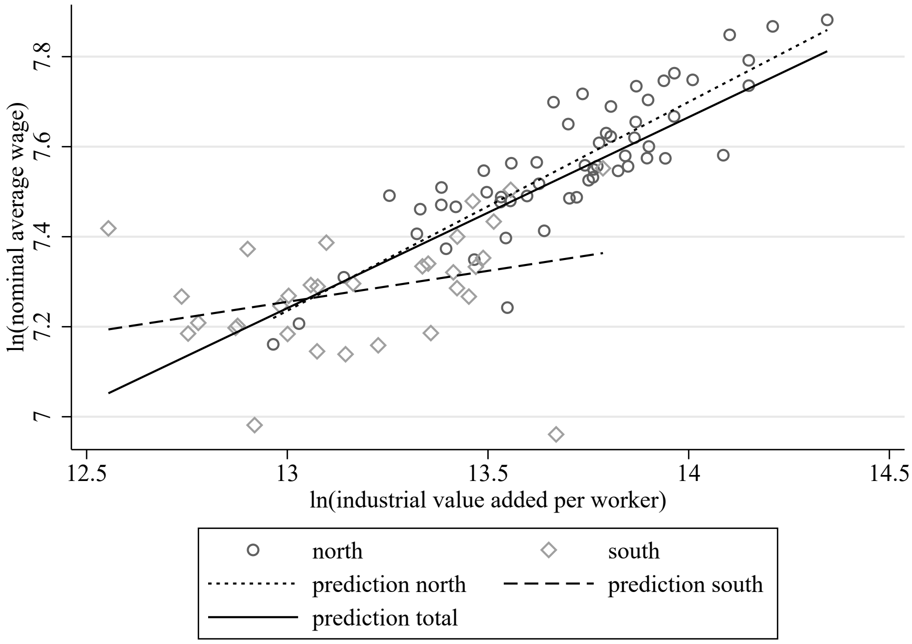
The previous analysis has shown that the wage zone system already influenced wage rates both within and between macroregions, and that its repeal caused a complete disconnect between productivity and minimum wages. This evolution would explain the drop in mobility for the marginal low-skill worker. However, were contractual minimum wages set high enough with respect to the wage distribution to force the same dynamics on average wages—potentially modifying the mobility of a larger share of workers?

To answer this question, I first describe the correlation between average effective wages for blue-collar workers and value added per-employee in manufacturing at the beginning and at the end of the period. Figure 15 shows that, in 1962, there was a strong positive association at the national level between the nominal average industrial wage and productivity. At the end of the period, in 1981, the association was still positive but apparently weaker. Distinguishing between macroregions, it appears that the South underwent the opposite trajectory, as it shows a stronger correlation in 1981 than in 1962. At a superficial level, these correlations might seem to disprove our argument: the relationship between nominal effective wages and productivity does not seem to be disrupted by the repeal of the wage zones.

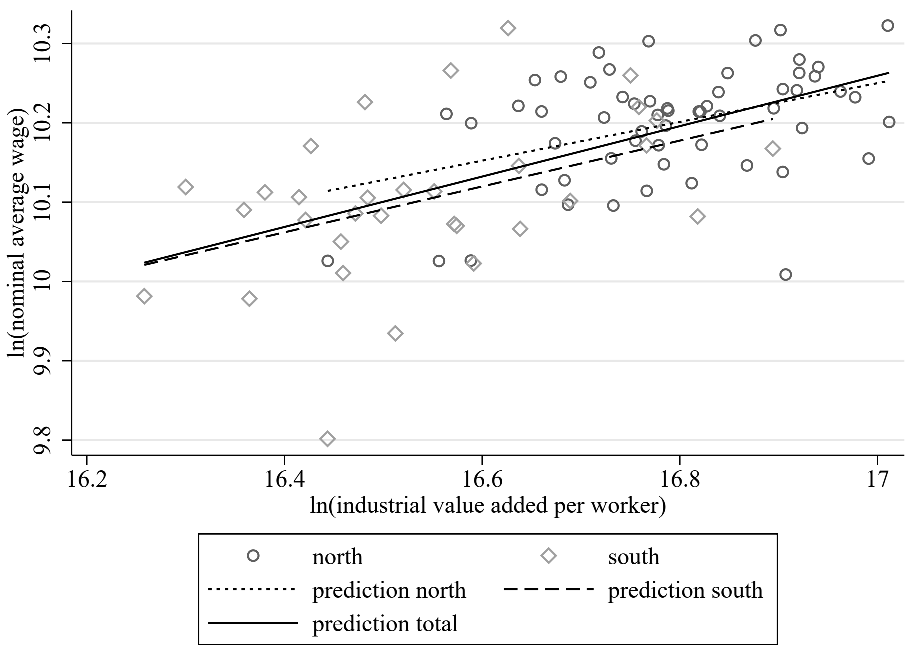
However, plotting the same graphs with real wages hints to a different conclusion: Figure 16 shows that a strong positive relationship between effective wages and productivity was present in 1962, but it disappeared entirely in 1981. Moreover, the figure shows that the disconnect between effective wages and productivity was already present within Southern regions in 1962 (represented by the flat dashed line in panel 16a). Hence, the repeal of the wage zones impacted most strongly the wage rates within the Centre-North and between the two macro-areas. By the end of the period, it appears that an average blue-collar workers would obtain no systematic gains from moving between a low-productivity and high-productivity province. This observation would support our argument that the repeal of the wage zones removed a large monetary incentive for internal mobility, especially for long-distance migration.

To test whether this hypothesis holds throughout the time period, I estimate Equation 15 substituting the level of the minimum wage in the province with that of the average effective wage. In addition to the previous variables, I also control for the *skill* composition of the industrial workers in 1966, before the repeal of the wage zones, interacted with a linear trend. This control is necessary because, contrary to our measure of minimum wages (which was computed as the mean minimum wage for low-skill workers only), effective wages average different skill levels. Hence, provinces where the share of high-skill blue-collar workers is higher would show a greater average effective wage.

Figure 17 plots the estimated coefficients separately for the Centre-North and the South. As was the case with the minimum wages, we find a positive and statistically significant relationship in



(a) 1962

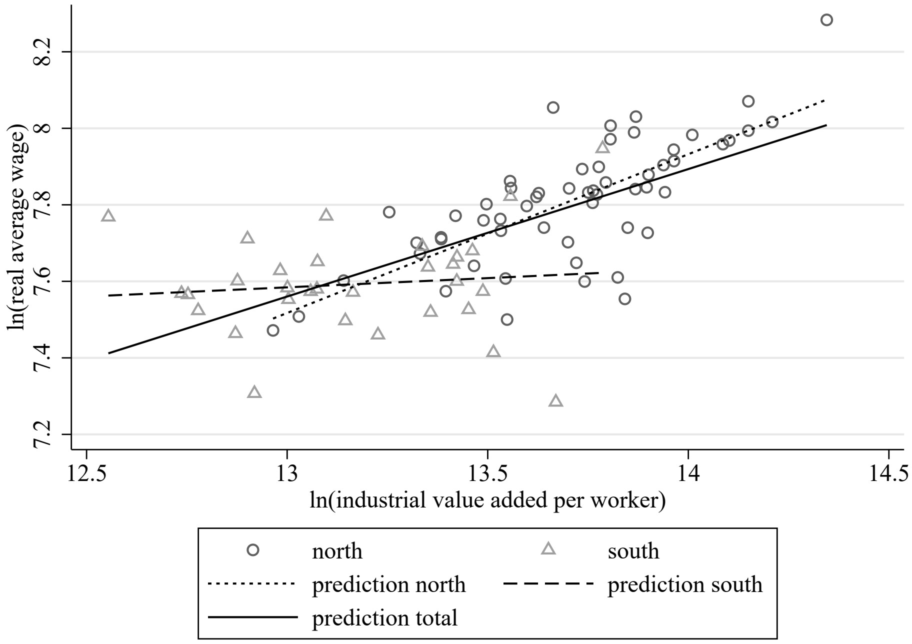


(b) 1981

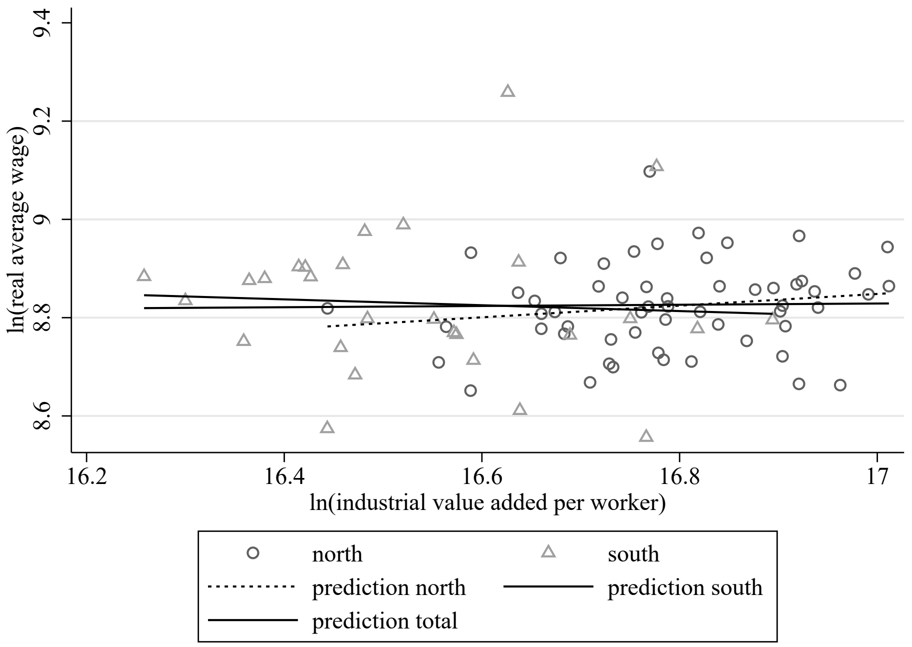
Figure 15: Average nominal wage and value added

Scatterplot of nominal average industrial wages and value added per employee in industry, in ninety-two provinces. For sources and methodology, see text and section 3.

the Centre-North before the repeal of the wage zones, and a weaker one in the South. Following the completion of the repeal, in 1972, we observe the disappearance of the relationship in both macroareas, with the South potentially drifting towards negative coefficients, even though the confidence interval at the 95% level includes the zero. Thus, the results largely confirm the descriptive analysis: the wage zone system allowed real effective wages in the Centre-North to vary



(a) 1962

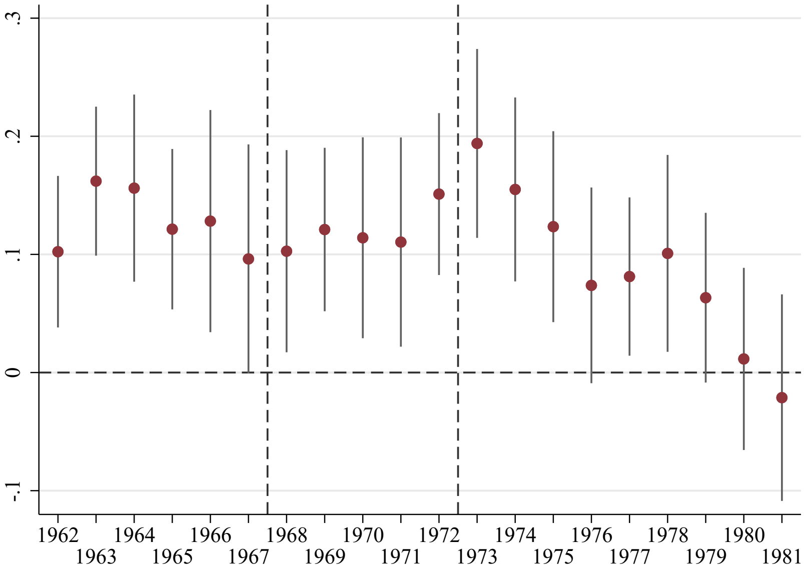


(b) 1981

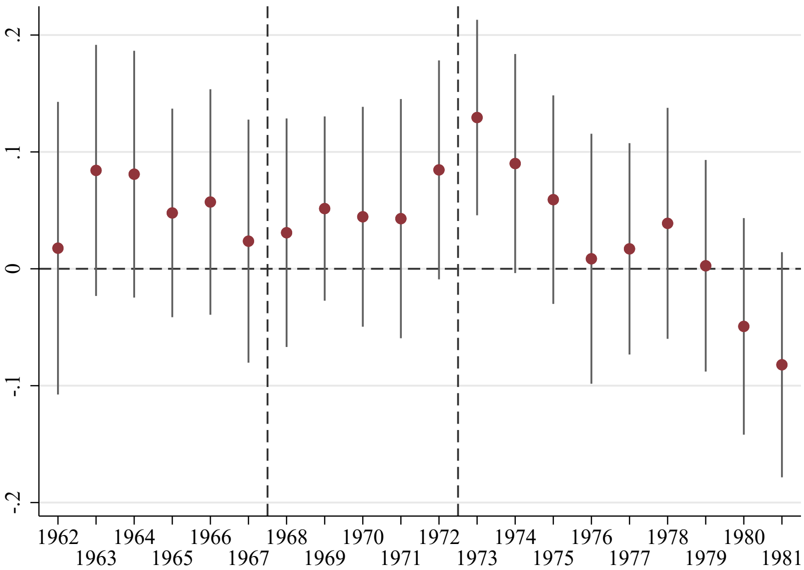
Figure 16: Average real wage and value added

Scatterplot of real average industrial wages and value added per employee in industry, in ninety-two provinces. For sources and methodology, see text and section 3.

with productivity, but less so in the South, possibly due to the already significant compression in minimum wage differentials.The repeal of the wage zones exacerbated the situation in the South and decoupled average wages from productivity in the North, too. This conclusion appears to suggest that the labour unions were successful in their bid to make wages an ‘independent variable’ with respect to productivity, at leas at the aggregate level. The next section discusses the consequences for unemployment and spatial mobility.



(a) Centre-North



(b) South

Figure 17: Real wages and productivity, by macroregion

Association between the real minimum wage and value added per employee in manufacturing, interacted with an indicator for whether the province is located in the Centre-North or in the South, controlling for province and time fixed effects, composition of industrial employment in the province by subsector and local cost of living. Vertical solid lines represent the 95% confidence intervals obtained from standard errors clustered at the province level.

5.3 The polarization of unemployment and the slow-down of macroregional mobility

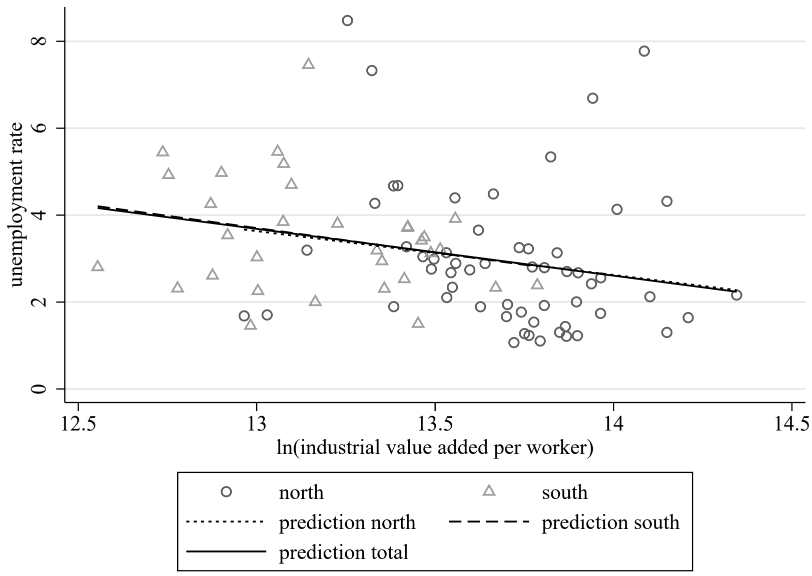
A common argument in the literature maintains that the spatial equalization of minimum wages causes excessive unemployment in low-income regions as inviduals prefer to queue for local jobs rather than move to high-income areas that offer the same, if not lower, real wages. To test whether this mismatch was instigated by the repeal of the wage zones, Figure 18 plots the scatterplot between the unemployment rate in the province and the industrial value added per employee, separately for the beginning and the end of the period. It shows that, in 1962, there was a negative correlation, meaning that unemployment was highest in low-productivity provinces. However, the correlation was relatively weak, as it is supposed to be in a competitive labour markets where unemployed individuals are allowed to move freely and the supply of housing in high-income provinces is allowed to adjust to positive shifts in local demand.

The data from 1981 shows, instead, a more complex picture (panel 18b). The negative correlation at the national level became much stronger, while the correlation within macroregions disappeared entirely. These dynamics are characteristic of a polarized and disconnected labour market: unemployment is systematically higher in Southern regions than in the Centre-North—irrespective of local productivity. The local labour markets within the two macroareas, instead, appear in equilibrium—even more so than at the beginning of the period.

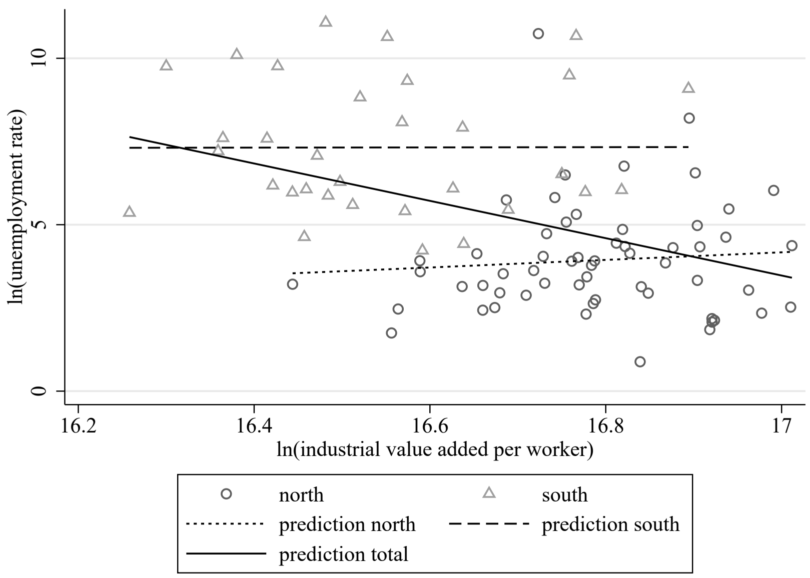
Could this polarization be connected to changes in migration patterns? A comparison with migration rates appears to corroborate this hypothesis: Figure 19 shows that net emigration rates were negatively correlated with productivity in 1962, both at the national level and within macroregions. Most noticeably, a sizeable number of provinces in the Centre-North with lower-than-average productivity were net senders of migrants, while high-productivity regions in the South had comparatively lower emigration rates. The situation was much more polarized at the end of the period, even though mobility in general was significantly reduced. In 1981, almost all provinces in the South were net senders, irrespective of productivity. In the Centre-North, instead, medium- and low-productivity provinces were just as likely to receive immigrants than high-productivity provinces. This situation is compatible with present-day studies which find that local amenities, rather than labour market factors, explain migration within macroregions.

The timing of the inversion in the relationship between emigration and unemployment can be tested by regressing the net emigration rate e (computed as emigrants minus immigrants over mid-year population) on the unemployment rate u interacted with an indicator for the year, according to the following specification:

$$ln(e_i) = \pi + \sum_{y=1962}^{1981} \omega_y \, ln(u_i) \times 1(Year = y) + X_{it}'\xi + \tau_t + \phi_j + \zeta_{it} \qquad [16]$$



(a) 1962



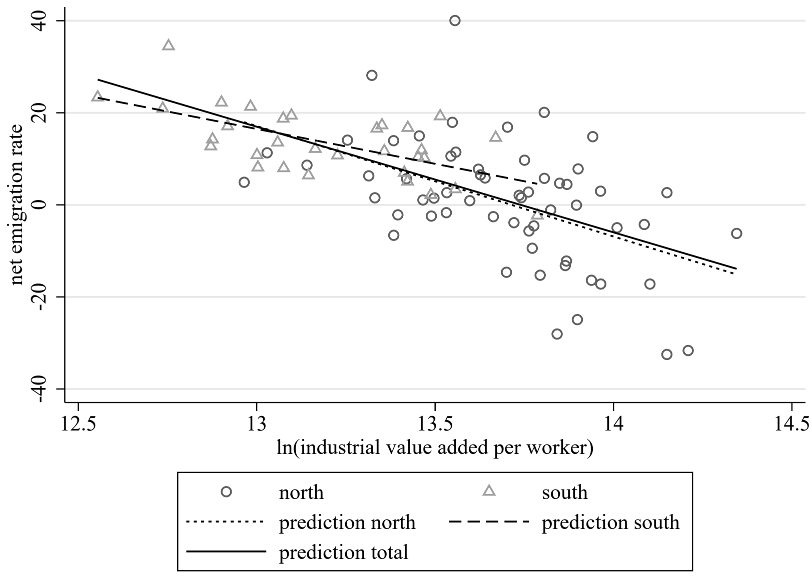
(b) 1981

Figure 18: Unemployment rate and value added

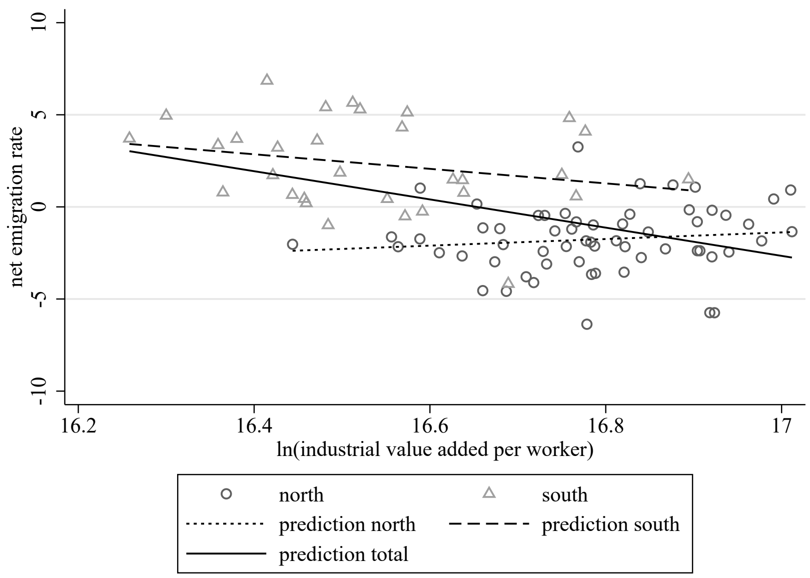
Scatterplot of unemployment rate and value added per employee in industry, in ninety-two provinces. For sources and methodology, see text and section 3.

The specification uses province and time fixed effects and a vector of controls including the province's population, the economic structure (annual value added in agriculture, industry, commerce, services), the sectoral composition within the manufacturing sector (interpolation of employees in 18 subsectors between census years) and skill distribution in manufacturing (shares of workers classified as high skill, medium-high, medium-low, low skill, apprentices and other).

Figure 20 shows the annual estimates for the interacted coefficient. The graph clearly identifies a positive association between unemployment and net emigration before the repeal of the wage zones, which however turned negative at the end of the period. The transition between the two regimes



(a) 1962



(b) 1981

Figure 19: Net migration rates and value added

Scatterplot of net migration rates and value added per employee in industry, in ninety-two provinces. For sources and methodology, see text and section 3.

happened exactly during the gradual introduction of spatial equalization in nominal minimum wages. This evidence reinforces the argument that the current spatial mismatches in the Italian labour market—i.e. the low propensity to migrate from low-income, high-unemployment areas—are not a traditional historical feature but rather a recent acquisition, which can be linked to the repeal of the wage zone system in 1968-1972. The same pattern can be identified both within the North and within the South (Figure A.25).

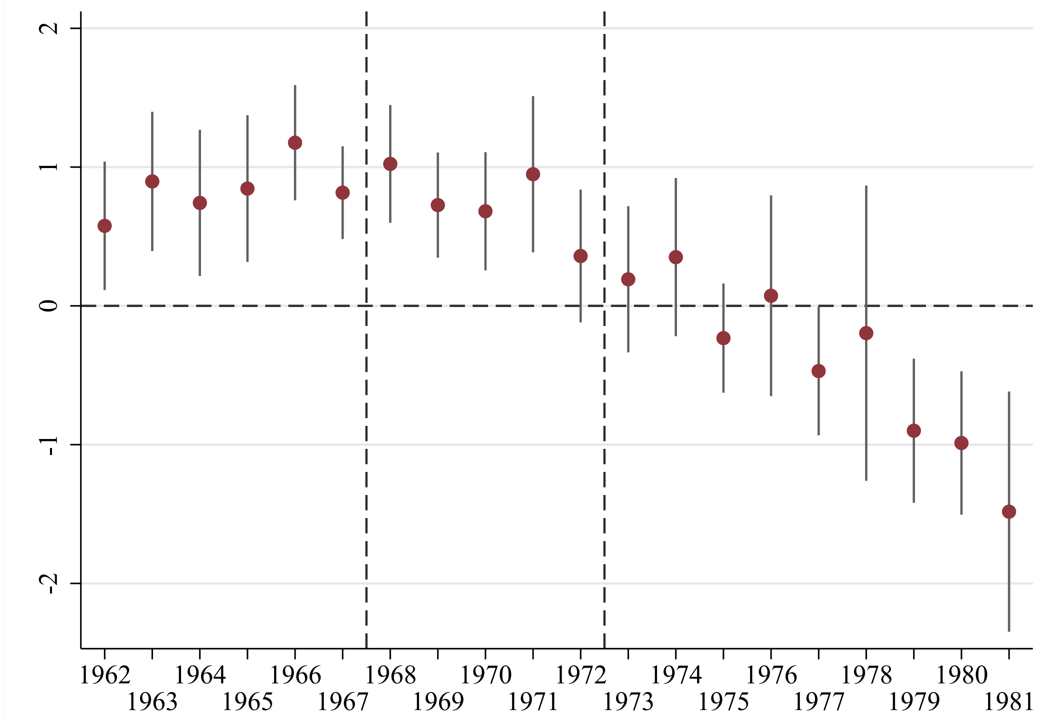


Figure 20: Net emigration and unemployment

Association between the net emigration rate and the unemployment rate, estimating by OLS including province and time fixed effects and a vector of time-varying controls. Vertical solid lines represent the 95% confidence intervals obtained from standard errors clustered at the province level.

6 Conclusions

Italians historically showed great propensity to internal mobility, which was particularly high between the 1950s and the 1960s. In the early 1970s, however, internal migration rates fell by one third, and remained at low levels for the following decades, despite widening spatial differentials in unemployment and income levels. Searching for the possible causes of these spatial imbalances, economists have often stressed the role of labour market institutions—most importantly, the spatial equalization of nominal minimum wages established by sectoral collective agreements. In the most common argument, these agreements incentivize individuals to stay in low-income areas—where real wages are high due to the low cost of living—and queue for local jobs, rather than move to high-income areas which offer greater chances of employment but lower real wages.

This paper has provided the first historical test for this hypothesis by comparing the current wage-setting regime with the previous system, which allowed nominal wages to vary between provinces according to fixed coefficients that proxied for differences in local productivity and price differentials. The paper has done so by reconstructing a range of labour market statistics at the province level from 1962 to 1981, with annual frequency. In particular, the paper has presented new of minimum and average wages of blue-collar workers in the manufacturing sector, and original estimates of local industrial productivity, cost of living and unemployment. All sources have been specifically digitized from printed primary sources and harmonized to allow intertemporal and spatial comparisons. These statistics have been merged with a new dataset of bilateral internal migration flows between Italian provinces, for the same time span.

The analysis has first presented an augmented gravity model of internal migration which tested for the role of nominal minimum wages as a pull factor. After having established a strong pull effect of nominal minimum wages on migration flows, the analysis has found that the model provides good predictions of the fall in internal migration after the spatial equalization of nominal wages. The analysis also found that nominal minimum wages lost their influence as pull factors of migration around the same time.

Discussing potential mechanisms, the paper has shown that the transition from the old to the new wage-setting regime inverted the relationship between real wages and productivity, from positive to negative. This shift also affected average industrial wages: while originally real wages were higher in high-productivity areas, by the end of the period they lost all association with local productivity. This evolution appears to have caused a polarization of the labour market the Centre-North and the South. In particular, unemployment rates appear systematically higher than in the Centre-North, irrespective of productivity.

The paper has wider relevance and policy implications. With respect to the international literature, the paper connects to recent research exploring the causes for falling internal mobility and rising spatial misallocation between local labour markets across developed countries. While research on the United States has stressed the role of constraints to housing supply in high-income areas, this paper highlights the role of wage-setting institutions in altering the monetary incentives to migrate from low-income areas. These results have important policy implications: with growing political pressure to ensure fair wages and higher living standards for disadvantaged groups—including the push for higher statutory minimum wages and a more central role of collective bargaining—, it is necessary to consider the second-order effects of these interventions on local labour markets and on their aggregate performance.

Appendices

A Additional figures

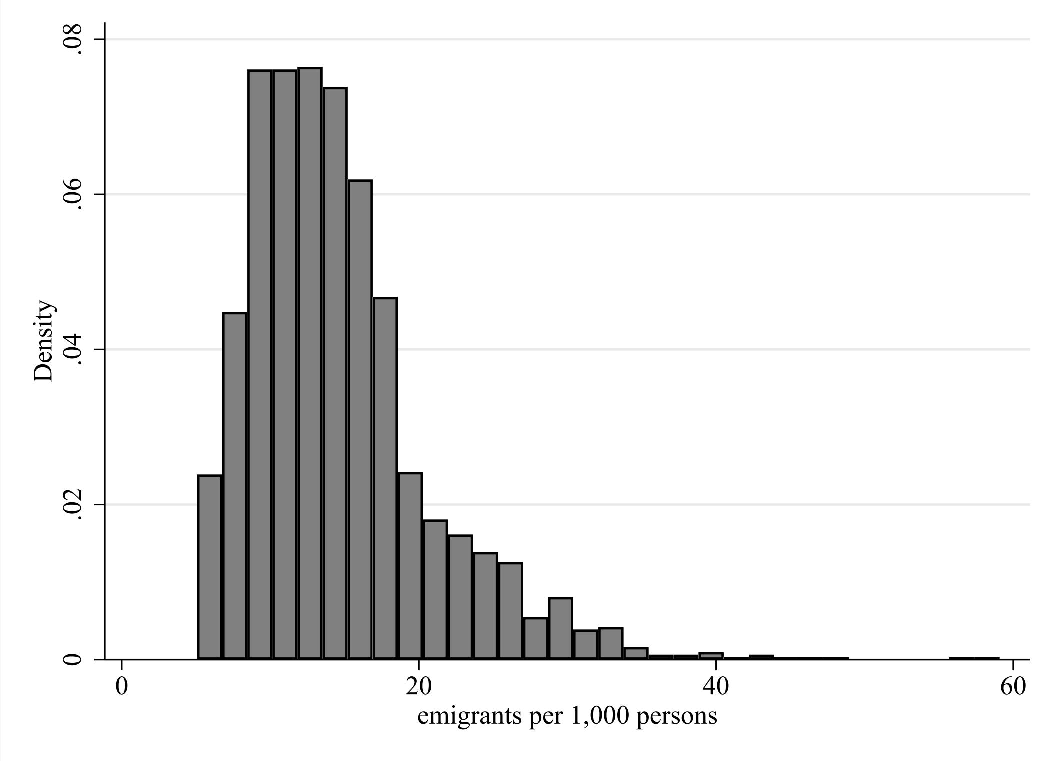


Figure A.21: Distribution of gross migration rates (1964-1981)

The graph reports the distribution of gross migration rates from 92 provinces (excluding internal migration) for the whole period 1964-1981. For sources and methodology, see text and section 3.

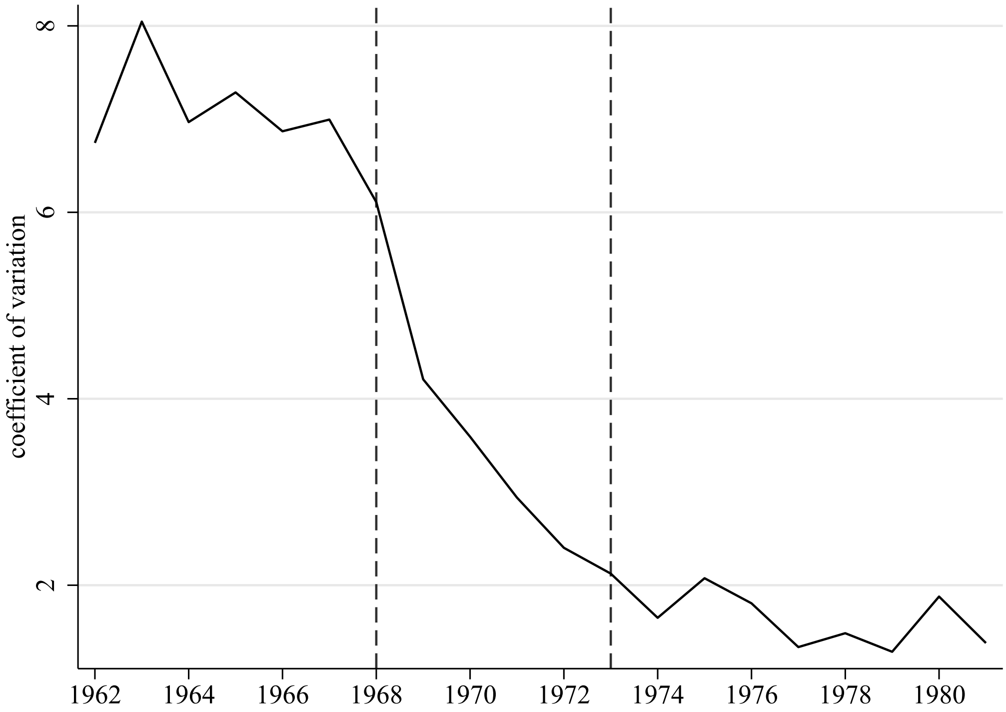


Figure A.22: Spatial variation of mean minimum wages

Coefficient of variation for the mean minimum industrial wage between provinces. The coefficient is computed as the standard deviation of the mean minimum wage for low-skill blue-collar workers across sectors (weighted by the local industry shares), divided by the mean value and expressed in percentages. For sources and methodology, see section 3.

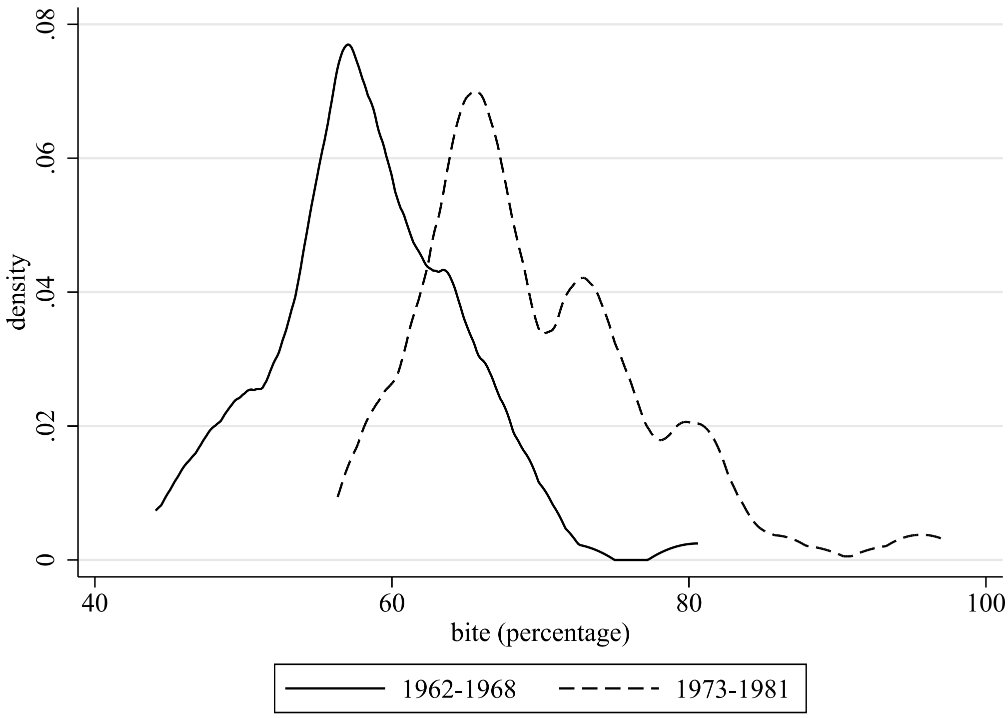


Figure A.23: Distribution of minimum wage bite across provinces

Kernel density distribution of minimum wage bite across ninety provinces, by period. The bite is computed as the percentage ratio between the local mean minimum wage and the mean average effective wage, obtained according to Equation 1 and 3, with annual frequency. Results are then averaged by period. For sources and methodology, see text and section 3.

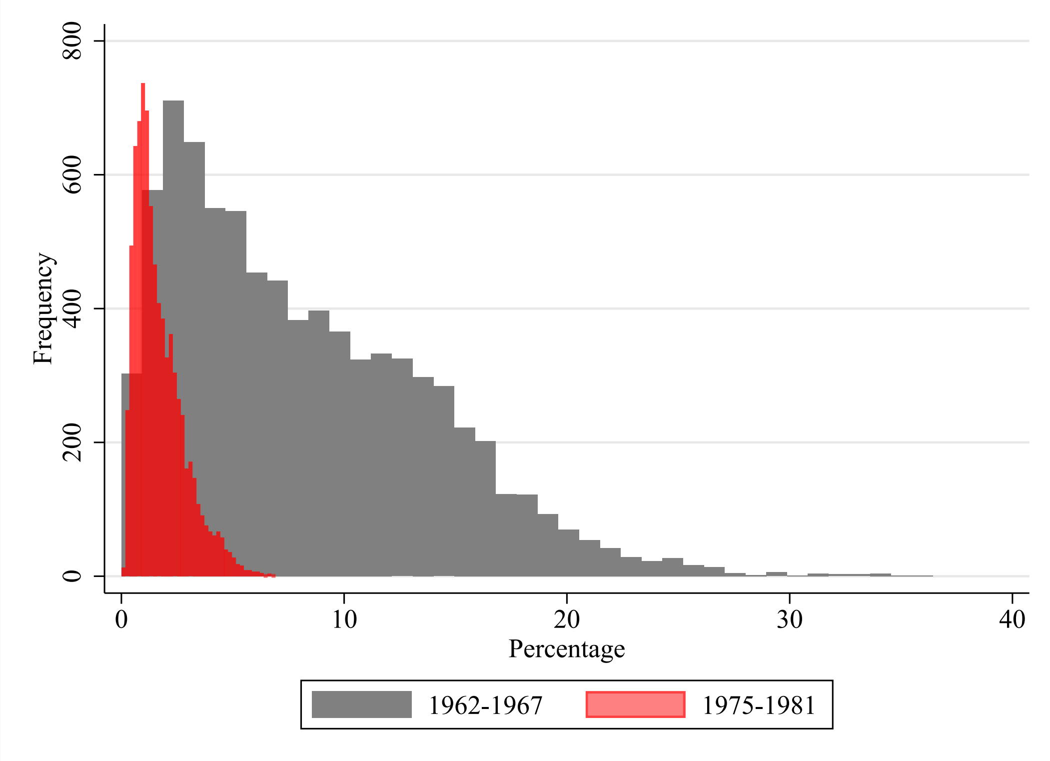
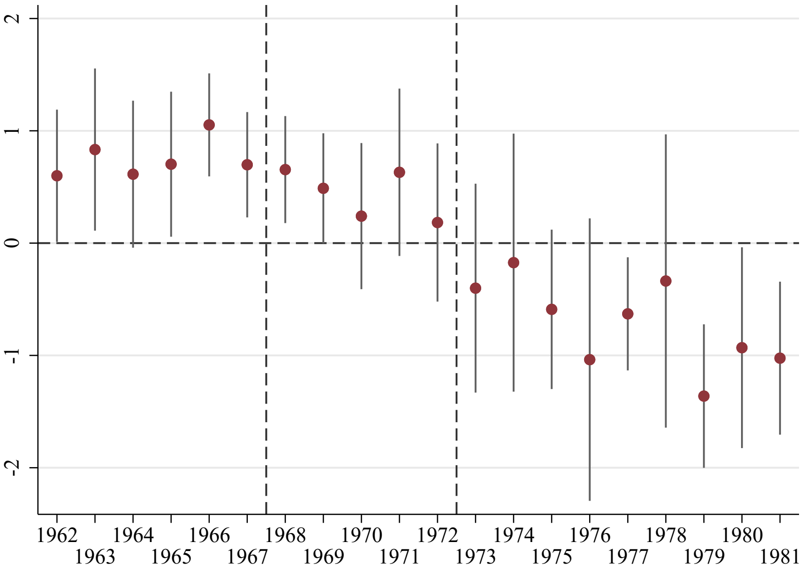
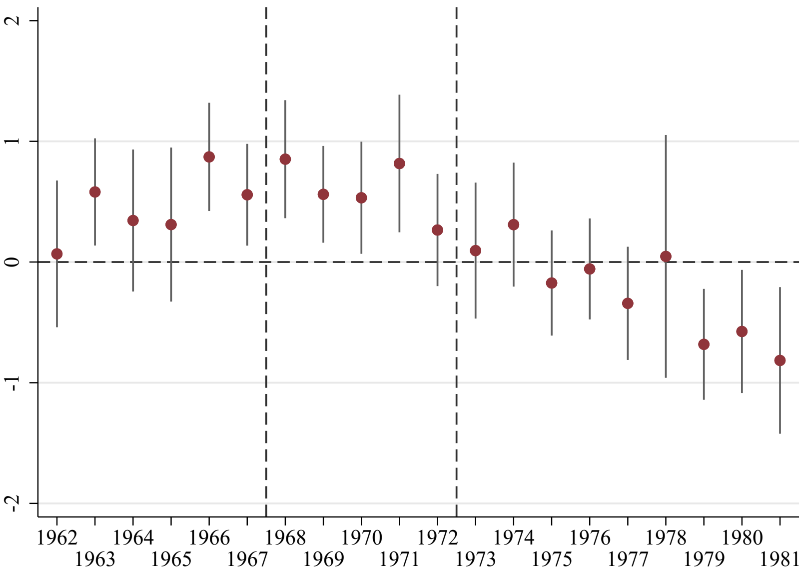


Figure A.24: Percentage differential in nominal minimum wages within dyads

Distribution of percentage differences between the nominal minimum wage at destination and at origin, for all 8,100 province dyads, averaged across the period 1962-1967 and 1975-1981.



(a) Centre-North



(b) South

Figure A.25: Net emigration and unemployment, by macroregion

Association between the net emigration rate and the unemployment rate, estimating by OLS including province and time fixed effects and a vector of time-varying controls. Vertical solid lines represent the 95% confidence intervals obtained from standard errors clustered at the province level.

B Additional tables

Table B.3: Descriptive statistics

	(1) 1962-1971		(2) 1972-1981	
	mean	sd	mean	sd
ln(Emigrants)	3.26	1.53	3.02	1.54
ln(Distance km)	5.87	0.74	5.87	0.74
ln(Population)	13.03	0.65	13.06	0.67
ln(Minimum ind. wage)	7.76	0.27	9.41	0.64
ln(Unemployment)	8.57	0.73	8.80	0.81
ln(Average ind. wage)	8.30	0.33	9.79	0.59
log difference cost of living	0.00	0.13	0.00	0.15
ln(GDP pc)	6.44	0.44	7.90	0.69
% North-West	0.22	0.00	0.22	0.00
% North-East	0.22	0.00	0.22	0.00
% Centre	0.20	0.00	0.20	0.00
% South	0.22	0.00	0.22	0.00
% Islands	0.13	0.00	0.13	0.00
Observations	80,100		80,100	

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