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Abstentionism and class divide in Italy: The labor market roots of economic and political inequality

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WORKING PAPER SERIES

Abstentionism and Class Divide in Italy: The Labor Market Roots of Economic and Political Inequality

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Abstentionism and Class Divide in Italy

The Labor Market Roots of Economic and Political Inequality

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Abstract

Abstentionism represents one of the most evident failures of modern democracies. The root causes of the phenomenon are still quite difficult to understand. Why do people abstain? Because of lack of political offer or, rather, because of exclusion and marginalization? This paper addresses a potential root cause, linking political behavior at the ballot box to material conditions in Italy, over the period 1983-2018. Through the lens of a long-run structural analysis, conducting both non parametric and parametric estimates, we connect the emerging stratification of the Italian labor market into social classes and the rise in non-voting. Our empirical results give support to the prediction of *Power Theory*, according to which economic marginalization, here measured through social class divides, fosters political marginalization.

Keywords: Wage inequality, Workforce stratification, Social classes, Power Theory

JEL CODES: D72, D62

1 Introduction

Abstentionism represents one of the most recurrent patterns in modern democracies since the Great Recession. Recent evidence (Schäfer and Schwander, 2019; OECD, 2019) has shown that citizens express dissatisfaction in terms of voting, therefore rendering one of the foundational tools for democratic functioning ineffective. Implications of non voting are manifold: first, non voting implies lack of transmission channels from citizen needs toward the political sphere, undermining the very objective of democratic settings; second, it renders the actual weight of voters more and more influential, amplifying divides between those that count and those that are excluded; third, it favors the formulation of political agendas targeting the interest of voters, feeding a dysfunctional interaction between parties' political offers and citizens' demands.

Why do people in Western democracies abstain from voting, when voting itself has long been the foundation of democratic decision-making and the product of intense political and social struggles? One of the most mobilized candidate explanations is related to the quality of the political offer. The declining quality of the manifesto of parties fosters abstentionism, as people feel themselves as not part of the *élite* of political interest. Whether abstentionism is a passive or rather an active political choice remains an open question. The very same behavior may, in fact, be due to intention to punish the political offer or simply because of detachment from the political matters (Weber, 2011).

Beyond the individual perspective, at the collective level, abstention may signal the erosion of democratic systems. In fact, echoing the debate on the incompatibility between capitalism and democracy (Krahé, 2022), increasing abstentionism can be seen as one of the channels through which the growing intertwining between political and economic power is expressed. This, in turn, puts pressure on the legitimacy of models based on the majority rule, when this majority is not truly representative. And in fact, episodes challenging the legitimacy of elections are becoming recurrent, examples are the attacks to Capitol Hill and to President Lula (Boix, 2020; Mair, 2013; Fumagalli and Ottonelli, 2022; Acemoglu, Naidu, et al., 2015, Iversen and Soskice, 2020).

More recently, however, the rise and consolidation of the phenomenon are leading to the search of explanations that go beyond the mere political sphere, and that include motives connected to the economic sphere. In particular, the rise in economic inequality and the widening of social class divides experienced over the last thirty years represent a good candidate explanation for political detachment from the urns. According to *Power Theory* (Verba et al., 1995; Urbinati, 2025), democratic and political rights are exercised when individuals perceive a sense of security, both social and economic. In contrast, when individuals perceive themselves as weak, they tend to self-marginalize and distrust forms of participation to political decision-making. Inaction, that is non-voting in our case, is the expected outcome of feelings and perceptions of insecurity, and particularly of lack of power. Usually, the phenomenon is studied at the individual-level, and it has been mainly addressed by means of individual survey data looking at political participation. However, during phases of widespread marginalization, as the current one, inaction may spread from individual to collective forms and become a social signifier of progressive marginalization of an increasing number of segments of society.

In the economic literature, the implications of capitalism, and particularly dysfunc-

tional capitalism, on democracy have been mostly addressed looking at voting rather than non-voting behaviors. Research has mainly examined issues of political polarization and increasing consent for populist parties, with motivations often linked to market-based explanations such as technology and globalization (Dorn et al., 2020; Frey et al., 2017).

In recent years, scholars and public opinion have renewed interest in the issue of political abstention. Non-voting and abstentionism have been traced back to multiple causes, namely, disaffection, dismantling of mass parties, decaying participation to democratic processes, forms of opposition. Most causes are traced back to the ideological and political dimensions (Koch et al., 2021; Mair, 2013; Adams et al., 2006; Mannheim and Sani, 2001; Mair, 2013).

Lately, a growing number of studies have instead challenged the material and economic dimensions behind the lack of participation in voting (Pianta, 2020; Bloise et al., 2021; Schäfer and Schwander, 2019; Dalton, 2017). However, they tend to rely on individual-level survey data (Tuorto, 2023; Smets and Van Ham, 2013), which limits the analysis to a point-in-time and individual-specific perspective. This approach has the risk of over-emphasize compositional factors at the expense of contextual or structural dimensions, such as geographical location and historical phases, which the literature has highlighted as particularly relevant (Rodríguez-Pose, 2018; Cagé and Piketty, 2025). What is missing is the analysis of abstentionism as a major collective process, potentially resulting from broader societal factors.

The importance of class identities to understand societies' dynamics was central during the 19th and 20th centuries, but nowadays many argues that class identities are increasingly less relevant. On the one hand, the subjective dimension of social classes has been overcome by the individualism of modern societies, but on the other, classes continued to show signs of social and political relevance. In fact, it is precisely the persistent economic division emerging from growing inequalities that can be recognized as class division (Eidlin, 2014).

Recent literature has increasingly highlighted the importance of class-based analysis in understanding economic inequality, emphasizing power dynamics within the labor process and organizational structures. Cuestas and Philp (2012) argue for a class-based approach centered on income sources—wages, profits, rents—bringing class power explicitly into distributional outcomes. In light of recent development of capitalism and managerial capitalism, this can be translated into differential bargaining power across hierarchical layers within the workforce. Indeed, Braveman (1998) and Majeed (2021) show how technological advancements and managerial control have led to widespread deskilling and weakening workers' autonomy. Building on these ideas, Wright (2000) introduces the concept of “contradictory class locations,” particularly for managers and executives who, though employees, align with capital interests due to their organizational roles (Jung, 2016; Cetrulo, Sbardella, et al., 2021). These perspectives suggest that hierarchies in labor markets affect inequality and societal stratification.

In the following, we examine the extent to which economic inequality, and more specifically social class divides as channels of economic insecurity leading to economic marginalization, drive abstention. We intend to pursue such research objective with reference to Italy, over a medium-term horizon, 1983-2018. Our research design does not investigate individual perceptions or feelings, but rather actual voting behavior, using official electoral

records from the Ministry of the Interior. By doing so, we link the well documented patterns of rising income inequality (Albertini and Ballarino, 2019) and social class divides (Albertini, 2013) with non-voting. In this respect, we delve into the economic and material roots of abstentionism, and we contribute to the literature discussing the implications of social class stratification for modern democracy. In such a respect, we go beyond the aggregate measurement of inequality and we intend to connect social and group stratification, interacting occupational, gender, geography, and age categories as indicators of social classes.

More in detail, we adopt a class-based approach, complementing the theoretical and empirical literature on inequality studies, specifically looking at Italy (Grusky and Weeden, 2001; Albertini, 2013; Cetrulo, Sbardella, et al., 2023; Devicienti et al., 2019; Cetrulo, Guarascio, et al., 2020a) and we build measures of wage inequality by segments of the working population. Once ascertained the emerging class divide in the period under study led by increasing wage disparities, by exploiting the regional and temporal variation in electoral behavior, we assess the association between labor market inequalities and erosion of modern democracies, bridging the literature on economic and political inequality (Blais, 2006; Pianta, 2020; Bloise et al., 2021; Cetrulo, Lanini, et al., 2023).

According to our evidence, class-based wage inequalities, measured over 36 years and 20 regions, exert a statistically significant positive effect on abstention. This relationship emerges over cycles lasting from 5 to 8 years, reflecting a silent cumulative process of growing inequality, visible today, that will manifest in the medium term, shaping outcomes tomorrow. Notably, distinguishing category-specific wage inequality measures by occupation, age, and gender allows us to evaluate different class divides' responses. The results are robust to the inclusion of other forms of inequality, particularly inequality in the distribution of capital income at the regional level.

The paper is organized as follows: Section 2 presents the theoretical framework, connecting electoral participation and social class stratification in modern democracies; Section 3 presents data and methods. In section 4 we present the results of wavelet analysis, assessing the temporal comovements at the aggregate level between inequalities and abstention, while in Section 5 we show the results of panel regressions conducted at the regional level. Section 6 performs a set of robustness estimates, including interactive fixed effects. Section 7 concludes the paper.

2 Abstentionism, Democracies and Class Divides

The link between abstentionism, democracy, and class divides connects with a broader debate on the incompatibility between capitalistic modes of production and the functioning of democratic systems, dating back to the 18th century (Krahé, 2022; Merkel, 2014; Boix, 2020; Mair, 2013; Fumagalli and Ottonelli, 2022; Iversen and Soskice, 2020). Alternate positions were taken during distinct historical periods, also depending on the dominant economic ideology, oscillating between stances supporting compatibility between capitalism and democracy, and others instead constitutively opposing the coexistence of the two processes.

In the economic literature, the debate has been dominated by the emergence of the New

Institutional Political Economy (Acemoglu and Robinson, 2006; Robinson and Acemoglu, 2012; Acemoglu, Johnson, et al., 2008), which emphasizes that democracy and capitalism can mutually reinforce each other through inclusive institutions, secure property rights, and the distribution of economic freedom. According to this perspective, democracy fuels capitalism because inclusive institutions enable entrepreneurial initiatives and innovation, potentially fostering widespread prosperity. Unfortunately, empirical evidence led to a different outcome, particularly in recent years. Indeed, it became evident how even established democratic institutions can become insufficiently inclusive, allowing concentration of economic and political benefits for a few, the *élite*, and the exclusion of large segments of the population, the mass. In this respect, the rise of dysfunctional forms of capitalism in recent decades has eroded democratic processes, limiting the ability of democratic systems to mitigate inequality and ensure broad-based participation.

Why Abstentionism

The exercise of abstentionism, whether active or passive, alters the *de jure* and *de facto* distribution of power in societies. Power is a key dimension in determining policy outcomes and fueling inequality in modern democracies. Let us now analyze why this is so.

First, let us start with the question: why is the universality of the vote important? In principle, universal voting allows us to make decisions on collective outcomes via the expression of all voices and interests that are affected by the process in itself; universal vote also creates a more tight link with parties, as induces politicians to gather information about voters' needs and to articulate their programs accordingly. Moreover, voting is considered an essential tool for democracies, as electoral systems are ways to express preferences for delegates, who will then have the power to govern with consent. Democratic elections grant legitimacy to govern by avoiding the use of violence or repression, allowing freedom of expression, and also broadening access to political power (Fumagalli and Ottonelli, 2022; Acemoglu, Naidu, et al., 2015).

The study of voter turnout began with Powell (1982), which posited electoral participation as one of the three main indicators of democratic performance (Blais, 2006). Perea (2002) asserts that abstentionism can be considered both as a symptom of dissatisfaction of citizens with the political system, and as a symptom of alienation of citizens resulting in lack of interest and distance from political institutions and parties. According to these approaches, what may underlie abstention is the detachment of politics and institutions from collective needs, the questioning of the principle of universality, equality, and influence of the vote, and therefore a more general weakening of the democratic system itself.

In fact, abstention and alienation from the political sphere can result in the exclusion from representative and democratic processes of entire parts of the population, and thus in an implicit rescaling of citizens' weight in decision-making processes. Specifically, a distrusted citizen, who does not believe that it is useful to vote to exercise some decision-making power, decides not to cast a vote. At the same time, as abstentionism increases, the power of voting citizens to influence electoral outcomes grows, thus ex-post distributing decision-making power unequally among citizens. Moreover, abstention can reinforce political demand-supply feedback loops. Abstainers disengage from politics through non-voting, and in response, parties calibrate their interests targeting their policies toward

those who vote, meaning that the political supply adapts to the composition of the voting body. This implies a transformation of the role of political participation as an exclusive right of those who count, and a legitimate discarding of those who do not (Nie, 1972).

A crucial question therefore pertains to the selection into voters and abstainers as linked to society's class structure, the latter exercising differentiated influence over political agendas. From the perspective of Power Theory, political participation is deeply intertwined with citizens' resources and capacities to influence decision-making. Those with fewer resources may perceive their capacity to affect outcomes as limited, reinforcing a sense of political powerlessness. Complementarily, the Withdrawal Hypothesis suggests that this perceived powerlessness leads citizens to disengage from formal political processes, including voting, as a rational response to the feeling that participation will not produce meaningful change (Verba et al., 1995; Urbinati, 2025).

If the divide between voters and abstainers at the individual level mirrors socioeconomic stratification at the collective level, this might imply that exclusion from political processes is neither random nor incidental, but systematically structured along class lines. Lower-income or less-resourced groups may be disproportionately represented among abstainers, while higher-income or better-resourced groups dominate the voting population. Such a dynamic not only undermines the representativeness of the democratic system but also perpetuates existing inequalities, as policy-making becomes increasingly tailored to the interests of those who participate.

Abstentionism, therefore, provides a critical lens to assess whether democratic systems effectively reflect the needs of the entire population or, instead, might exacerbate multidimensional patterns of alienation and marginalization.

Why Social Classes

As inequality and marginalization have grown over time, the relevance of a class-based approach to tackle inequality has been reconsidered on the basis of theory, society's evolution, and empirical analyses. A class-based analysis implies the inclusion of forms of economic and social stratification that encompass not individual characteristics but rather group attributes. In fact, scholars have been questioning the adoption of a single methodology to effectively grasp the key determinants and characteristics of growing inequalities in an expanding set of dimensions. The need of a "multidimensional approach" is evoked for distinguishing several "spaces" of disparities shaped by unbalanced endowment of economic, cultural, social and physical assets to characterize the individual or group condition (D. Grusky, 2018).

Among the many alternative spaces for the generation of inequalities, labor markets represent a privileged arena. This perspective is very relevant in light of the flexibilization wave that hit labor markets at the beginning of the 1990s. Flexibilization has not only worsened the average condition of workers, but, in particular, hit the hardest the weakest segments of the working class (Fana et al., 2016; Weisstanner, 2021; Cetrulo, Sbardella, et al., 2021; Hoffmann et al., 2022). Here, the concept of segments, or groups, opposite to the one of individuals, is fundamental to understand why it is not enough to consider only overall measures of inequality, but it is crucial to complement them with "segmentation measures", in the direction of better understanding the phenomenon of marginalization in

which we are most interested.

The first channel is the gender dimension. The segregation of women in the workplace and, more generally, in the labor market has always been a structural character in many advanced economies. Although women participation in formal labor markets has increased (Thévenon, 2013), the empowerment of this social group does not increase accordingly. The interplay between the gender dimension and other social factors, such as age and social status, has been underlined to explain the widening gap between men and women in terms of wages in the labor market. The channels include maternity leave, increased demand for work amenities and less internal career advancement for women (Goldin et al., 2017).

The second dimension is age, a source of inequality that derives from structural conditions that disadvantage younger age groups by placing them in a restricted choice set (Mulkeen, 2023; Puauschunder, 2017; Wildman et al., 2022). Labor markets create imbalances in terms of opportunities and ex-post rewards for younger workers (Bianchi and Paradisi, 2022; Blanchflower and Freeman, 2000). Wage inequality by age can reflect the worsening of labor market institutions in the last decades (Bukodi et al., 2006).

The third dimension is geography. Geographic inequalities influence incomes, wages and overall welfare between different spatial scales, such as neighborhoods, metropolitan regions, or countries (Bathelt et al., 2024). Interregional inequalities have started to grow mostly in high-income countries since the beginning of the 1980s, and the implications of geographical inequalities are not trivial (Ganong and Shoag, 2017; Rosés and Wolf, 2021). In fact, available opportunities and social mobility are affected by origins and location, beyond personal endowments, therefore differential individual and material well-being are influenced by geographical factors (Chetty et al., 2014; Granström and Engzell, 2025).

Finally, the analysis of social classes framed in terms of occupational categories is crucial to understand the undergoing transformations of society (Grusky and Weeden, 2001). This lens has proven particularly useful in assessing educational opportunities, healthcare access, and intergenerational transmission of status (Albertini, 2013), but also inequality during the COVID-19 pandemic (Cetrulo, Guarascio, et al., 2020b). The use of occupational categories to frame social classes is rooted in both theoretical and empirical underpinnings. Several authors argue for a focus on occupations and job qualifications as meaningful proxies for social classes. Grusky and Weeden (2001), critical of high-level aggregate class analysis, advocate a meso-level Durkheimian perspective that centers on occupational roles as functionally distinct social groups. Albertini (2013) further emphasizes the enduring role of occupational class in inequality, especially in Italy, where low career mobility and early-life inequalities persist throughout working life. Cetrulo, Sbardella, et al. (2021) reinforce this point by showing that while wage disparities across gender, age, and geography are mainly explained by within-group variation, job qualifications exhibit a dominant between-group component, suggesting that occupational class is a key explanatory factor for wage inequality.

Such dimensions are not only central to the wage inequality literature, but can exert a crucial role in affecting electoral participation. First, gender remains crucial for understanding inequalities in electoral participation. Although the traditional participation gap between men and women has narrowed over time, gender still structures the ways in which

individuals engage with politics, even beyond the ballot (Beauregard, 2014; Kittilson, 2018; Skorge, 2023). In contexts where women face persistent wage gaps and occupational segregation, they may perceive weaker connections to political institutions, which in turn can dampen electoral participation.

Territorial cleavages are relevant in explaining both low electoral participation and distrust in traditional political parties, as an increasing number of studies indicate (Rodríguez-Pose, 2018). Regional disparities in economic opportunities, infrastructure, and access to public services can create feelings of marginalization among citizens, which, in turn, reduce engagement in formal political processes. These territorial divides often reinforce existing inequalities, shaping patterns of political behavior and contributing to persistent variations in voter turnout across different geographic areas.

Age is another key dimension in explaining non-voting, as younger cohorts have consistently displayed lower levels of turnout compared to older citizens (File, 2014; Bhatti et al., 2012). Scholars have identified both life-cycle effects—where participation tends to increase with age as individuals accumulate resources, stability, and political experience—and cohort effects, whereby younger generations raised in contexts of political distrust or economic precarity may develop weaker habits of participation (Smets, 2016). These findings suggest that age related wage inequalities and precarious labor market conditions do not only structure material life opportunities but also help explain generational divides in political participation.

The occupational structure provides a comprehensive indicator of social status, it captures both economic resources and opportunities for political engagement. Individuals in higher-status occupations—such as professionals and managers—are consistently more likely to participate politically than those in manual or routine jobs, reflecting differences in income, education, and social networks (Tuorto, 2023). Lower levels of participation among the working class are closely related to more limited resources, including available time, and weaker mobilization opportunities, particularly when classical intermediary organizations, such as parties and unions, lose their capacity to activate and represent these groups. Beyond occupational class itself, the quality of employment also matters: individuals in precarious or insecure jobs often experience detachment from political processes, while stable and secure employment is associated with stronger civic engagement. Moreover, evidence further suggests that unemployment can have a “scarring effect” on political participation, discouraging individuals from re-engaging with democratic institutions even after returning to work (A. E. Rovny and J. Rovny, 2017; Marinova, 2022; Azzollini, 2023).

Finally, disparities in capital ownership and capital income can exert influence on voting patterns. The ownership of financial and real capital and its distribution are highly relevant for understanding stratification within the workforce and, therefore, might influence voting behaviour. Even when restricting the analysis to dependent employees, differences in access and accumulation of financial assets can reinforce socio-economic divides and generate additional forms of inequality not captured by wages alone. Including this measure allows us to account for the larger significance of wealth-based resources in shaping workers’ opportunities and, in turn, their potential participation in political and social processes (Piketty, 2014; Milanovic, 2019).

3 Data and Descriptive Evidence

Integrated Datasets

The database to retrieve information on wage inequality and construct indexes of stratification is the Italian Institute of Social Security Longitudinal Sample – *Rilevazione dei contratti di lavoro*, a micro-aggregated dataset based on administrative records. The dataset has a longitudinal structure and is based on a large representative sample of employees in the private sector – with the exception of agricultural and domestic jobs – from 1983 to 2018. Therefore, it does not include information on public employees or any type of self-employed jobs. The database does not contain individual information but averaged over different segments of the workforce, characterized by gender, age cohort (under 30, 30-50, over 50 years old), region, geographical area (North-West, North-East, South, Center, Islands), and occupational status (trainee, blue-collar, white-collar, middle manager, executive). For each year since 1983 to 2018 it reports information on yearly and weekly salary, number of work contracts, weeks of work, labor market entry age, the percentage of part-time and permanent jobs, and the percentage of jobs in 1-digit ATECO-2007 economic sectors. The dataset does not include wage incomes above 200.000 euro per year, thus it is top censored.

Data on voting behavior are drawn from the website of the Ministry of Interns “Eli-gendo”. The section “Open data” gives access to all results of electoral consultations at the aggregate and disaggregate level. The available data include the elections of the “Assemblea Costituente”, political elections, European elections, regional and administrative elections. Selected data refer to national elections in Italy from 1948 to 2022. Data report for both the election of the chamber deputies and the election of senate deputies. Data provide information on the number of electors and voters for each province, constituency, and municipality, reporting also the number of blank ballots. Finally, information on the number of male and female electors and voters are reported only from 2001 onward. The dataset contains the full historical range of the 19 Italian national elections.¹ The following analyses are performed considering only chamber’s data. The absolute number of abstentions is calculated by including blank ballots.

Data on capital income in Italy are drawn from the Survey on Household Income and Wealth (SHIW) conducted by the Bank of Italy, specifically the 2022 historical database. The survey covered 9,644 households and 23,063 individuals. Of the 9,644 households, 5,184 are panel households, i.e., households that were also interviewed in the previous survey (or in previous surveys), and the remaining 4,460 are households interviewed for the first time in this survey. Panel households have the same questionnaire number used in previous surveys, in order to be identifiable over time. The variable we employ is Capital Income which captures total income derived from both real and financial assets. It includes income from real estate, such as actual and imputed rents, as well as financial capital income, including interest from deposits, government bonds, and other securities, net of interest paid.

Lastly, data on regional characteristics were drawn from the Italian Statistical Office.

¹Election years are the following: 1948, 1953, 1958, 1963, 1968, 1972, 1976, 1979, 1983, 1987, 1992, 1994, 1996, 2001, 2006, 2008, 2013, 2018 and 2022.

Variables

In Italy, the percentage of abstainers increased from less than 10% to more than 35% between 1976 and 2022. As reported by *Openpolis* (2022), after the last parliamentary elections, the issue of abstentionism has emerged even more clearly. More and more voters are, in fact, not going to the polls, and the party of non-voting is now firmly the “first party” in Italy. Figure 1 presents different macro periods in the abstention trend identified by the Bai and Perron test (Bai and Perron, 2003), where abstentionism is computed as $\frac{\text{Abstainers}_t}{\text{Electors}_t} \%$. In particular, five breaks are identified, roughly corresponding to different decades. These breaks correspond to distinct political and social cycles, influenced by changing economic conditions, policy shifts, and socio-cultural transformations (Cetrulo, Lanini, et al., 2023). Periods of economic growth, economic crisis, or significant political reforms coincide with sharp changes in the rate of abstention. Importantly, while cyclical fluctuations are evident, the overall trend remains unmistakably upward, with certain periods showing pronounced acceleration, highlighting the gradual consolidation of non-voting as a structural phenomenon in Italian electoral behavior.

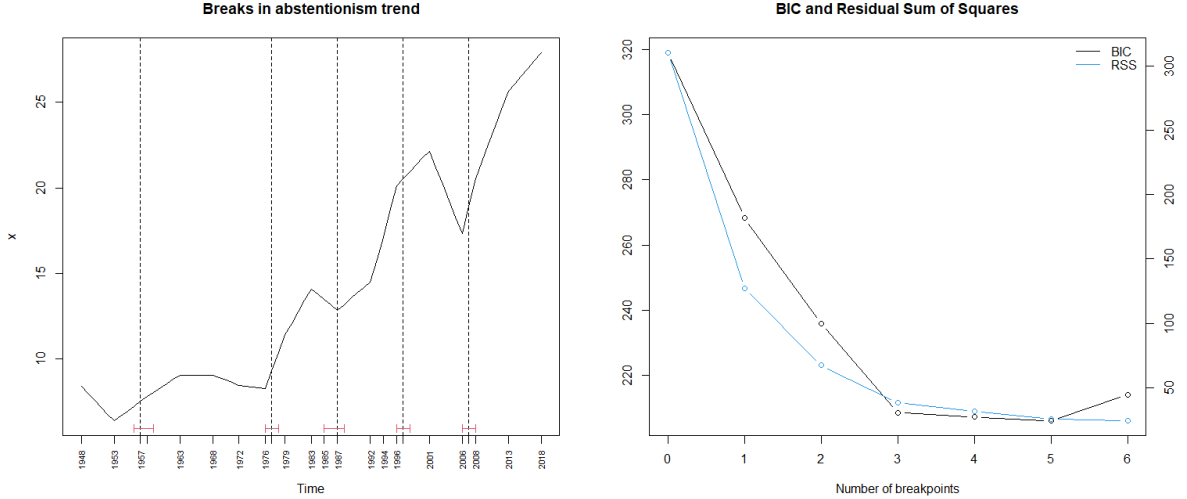


Figure 1: Bai-Perron test for structural breaks in the time series of abstentionism in Italy.

The empirical measures used for class stratification are category-specific inequality indexes that examine social stratification through four dimensions: job qualification, gender, age, and geographical area. These dimensions are analyzed alongside the overall Gini index. In what follows, we are going to display each index with the corresponding formulation, all computations are based on the standard Gini formula.

Between-category indexes are computed by first calculating weighted averages of salaries for each level of the macro-category, where weights are given by the number of work contracts by category. The unweighted Gini formula is then applied to these weighted averages to measure the divergence across levels within each macro-category. The overall Gini index is computed applying to the real wage the standard weighted Gini index formula, using as weights the registered number of work contracts.

Figure 2 and Equation 1 refer to the overall Gini index.

$$Gini_t^{\text{overall}} = \frac{\sum_{i=1}^N \sum_{j=1}^N w_i w_j |x_i - x_j|}{2 \left(\sum_{i=1}^N w_i \right) \left(\sum_{i=1}^N w_i x_i \right)} \quad (1)$$

where x_i is the real salary of segment i , w_i is the number of work contracts of segment i , and N is the total number of segments.

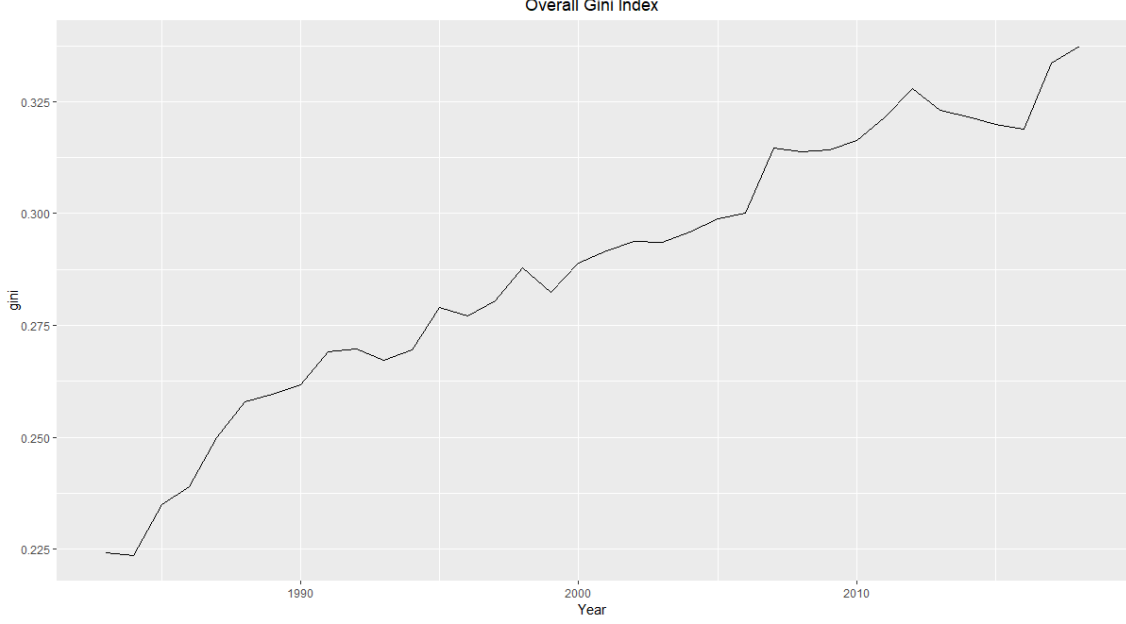


Figure 2: Overall inequality Gini index.

Figure 3 and Equations 2 and 3 refer to the gender dimension, where the levels are Male and Female.

$$\bar{x}_{g,t}^{\text{gender}} = \frac{\sum_{i \in g} x_i w_i}{\sum_{i \in g} w_i}, \quad g = 1, 2 \quad (2)$$

$$Index_t^{\text{gender}} = \frac{\sum_{k=1}^2 \sum_{l=1}^2 |\bar{x}_{k,t} - \bar{x}_{l,t}|}{2 \cdot 2 \sum_{k=1}^2 \bar{x}_{k,t}} \quad (3)$$

where x_i is the real salary of segment i , w_i is the number of work contracts, g identifies a specific category level in the computation of the weighted mean, and k and l are summation indexes running over all category levels when computing the divergence index.

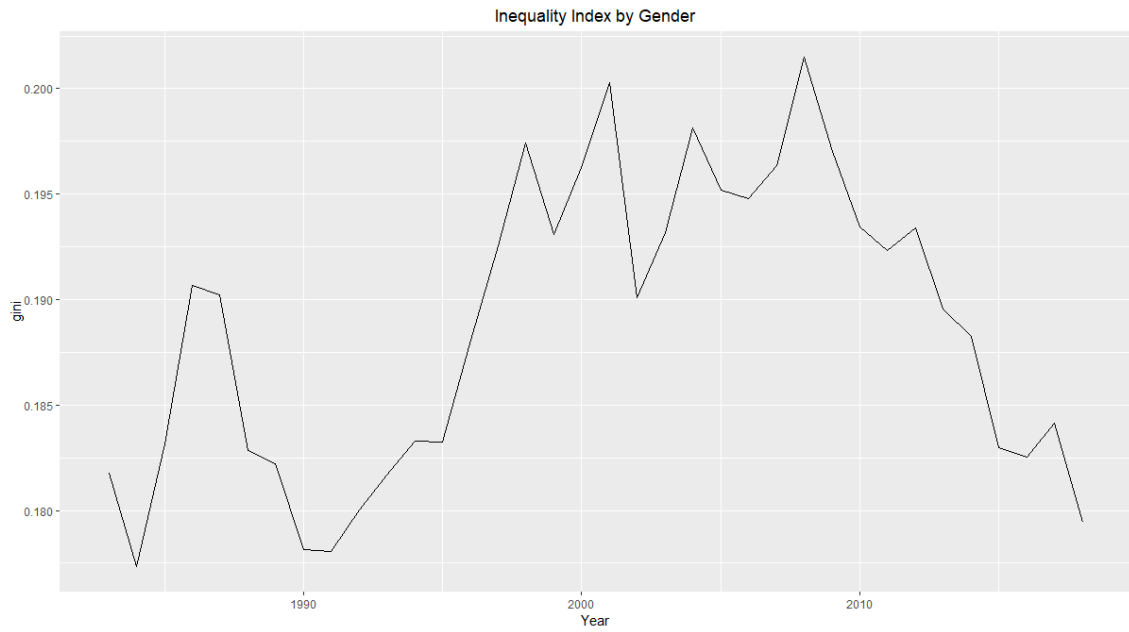


Figure 3: Wage inequality index computed between male and female workers.

Figure 4 and Equations 4 and 5 refer to the geographical dimension, the levels are North-East, North-West, Center, South, and Islands.

$$\bar{x}_{g,t}^{\text{geo}} = \frac{\sum_{i \in g} x_i w_i}{\sum_{i \in g} w_i}, \quad g = 1, \dots, 5 \quad (4)$$

$$Index_t^{\text{geo}} = \frac{\sum_{k=1}^5 \sum_{l=1}^5 |\bar{x}_{k,t} - \bar{x}_{l,t}|}{2 \cdot 5 \sum_{k=1}^5 \bar{x}_{k,t}} \quad (5)$$

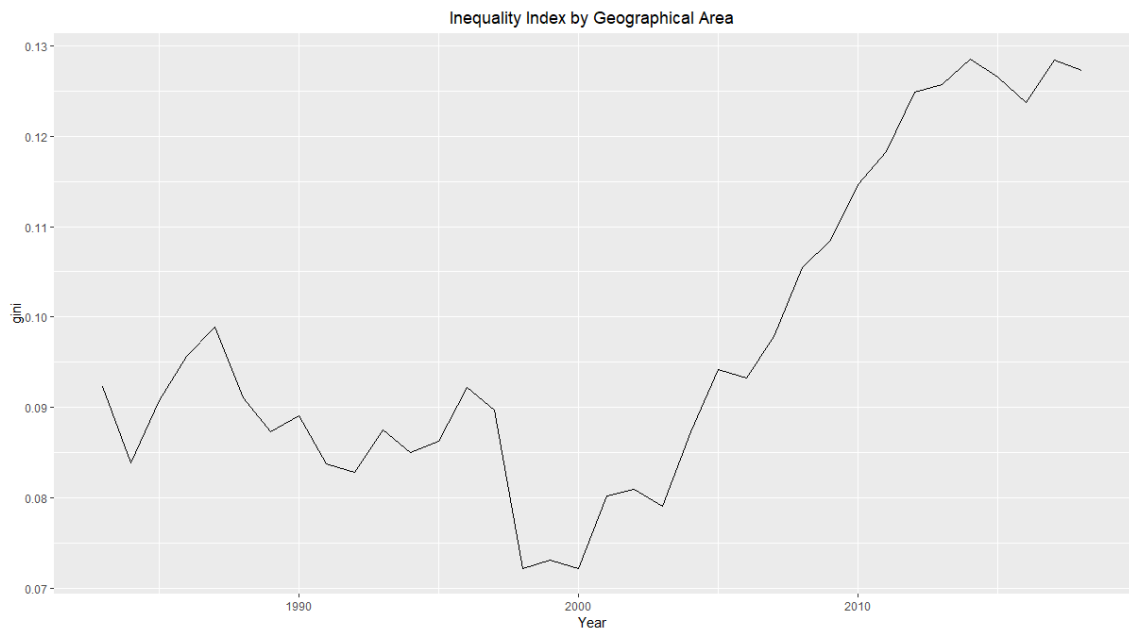


Figure 4: Wage inequality index computed between geographical areas.

Figure 5 and Equations 6 and 7 refer to the age class dimension, the age levels are under 30, 31-50 years old and over 50.

$$\bar{x}_{g,t}^{\text{age}} = \frac{\sum_{i \in g} x_i w_i}{\sum_{i \in g} w_i}, \quad g = 1, 2, 3 \quad (6)$$

$$Index_t^{\text{age}} = \frac{\sum_{k=1}^3 \sum_{l=1}^3 |\bar{x}_{k,t} - \bar{x}_{l,t}|}{2 \cdot 3 \sum_{k=1}^3 \bar{x}_{k,t}} \quad (7)$$

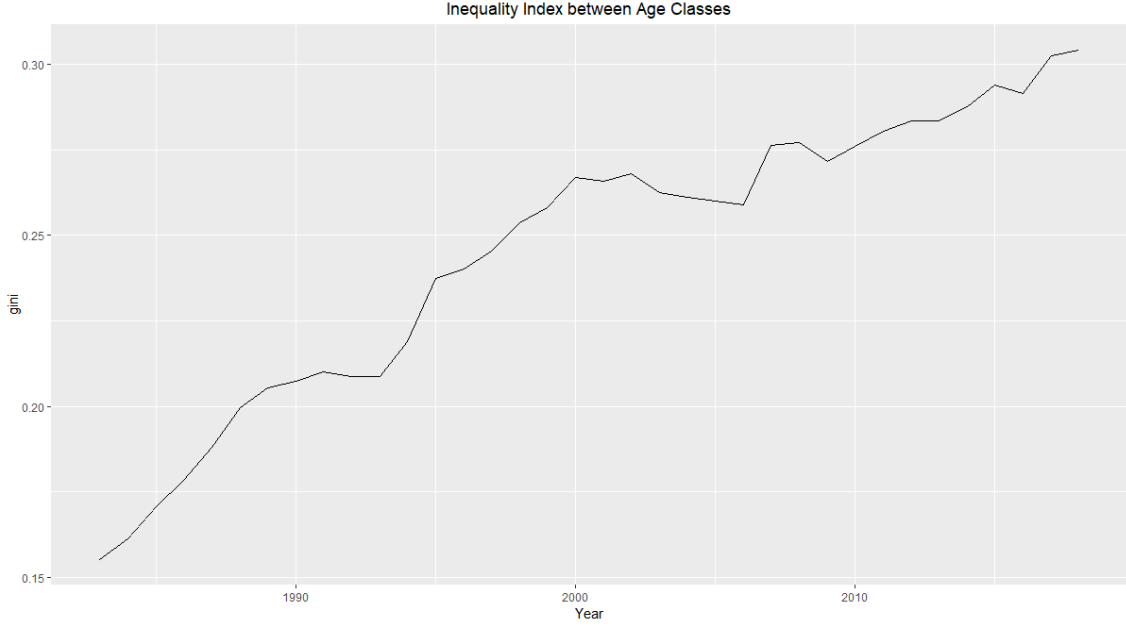


Figure 5: Wage inequality index computed between age classes.

Figure 6 and Equations 8 and 9 refer to the job qualification dimension, the levels are trainee, blue-collar, white-collar, middle managers and managers. Given that the middle manager category is registered in the sample only after 1995, we exclude it from the analysis for now in order to avoid artificial distortions in the inequality measure.

$$\bar{x}_{g,t}^{\text{job}} = \frac{\sum_{i \in g} x_i w_i}{\sum_{i \in g} w_i}, \quad g = 1, \dots, 4 \quad (8)$$

$$Index_t^{\text{job}} = \frac{\sum_{k=1}^4 \sum_{l=1}^4 |\bar{x}_{k,t} - \bar{x}_{l,t}|}{2 \cdot 4 \sum_{k=1}^4 \bar{x}_{k,t}} \quad (9)$$

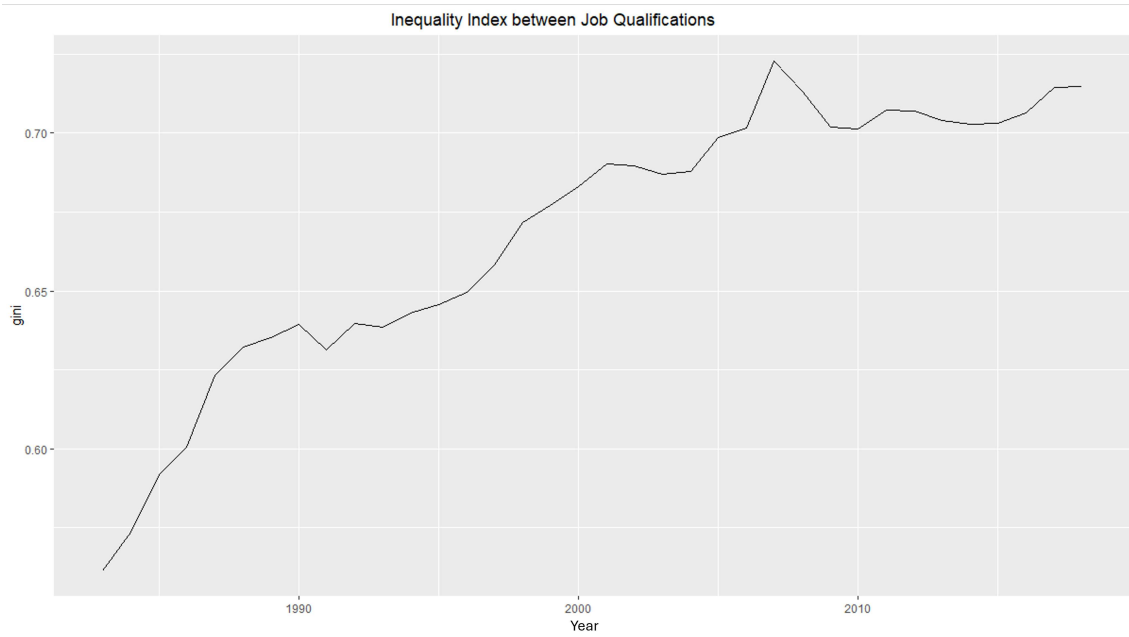


Figure 6: Wage inequality index computed between job qualifications.

We can observe a general increasing trend in almost all dimensions, with the exception of the gender index, which remains more or less constant. If we compare the different levels of inequality registered among the different categories, the macro-category of job qualification presents the highest level of between-groups wage inequality in terms of magnitude. Overall, these results signal a general wage divergence process entrenched in the Italian labor market, underlying the increasing asymmetries within the workforce.

4 Temporal Analysis - Identification of Shared Cycles

The relationship between electoral abstention and inequality is inherently complex, as it unfolds across multiple temporal dimensions. In particular, abstention trends are influenced by structural medium-term factors such as legislative cycles and the cumulative impact of socioeconomic policies. Electoral cycles in Italy, typically spanning five years (though often interrupted by political instability), offer a natural lens for examining short- and medium-term dynamics. During these periods, governments implement policies that can gradually reshape the distribution of income and influence public trust and political participation, making inequality a plausible driver of abstention on such horizons. The main objective is to observe on a time-frequency domain whether there exist areas of common variability of the series under study and, in particular, to determine whether these areas of high power and volatility of the series coexist at particular frequency domains through time. The analysis will allow us to assess how the intensity of the relationship evolves over time, if it exists, and, second, the lag structure of the temporal relations. The information, non parametric, will support the parametric analysis of the following section.

The results presented are based on the business-cycle (2-8 years) component of the series, isolated through the Christiano-Fitzgerald filter, removing the structural increasing trend present in all series. This approach serves to give relevance to the cyclical dimension, also given the different nature of the events under study in terms of time scales and

frequencies. The time sample goes from 1983 to 2018, and we proceed to analyze, one by one, the relation between the national percentage of abstention and the different between-category inequality indexes presented in Section 3. The abstention series was linearly interpolated throughout the time interval.

In Figure 7, we report the Cross-Wavelet power spectra, representing a measure of signal intensity that can reveal the presence of strong co-shared fluctuations between the series, identified both in the time and in the frequency domain. What emerges from the bi-variate analysis is a positive relationship both at low and high frequencies, which is underlined by the presence of right-pointing arrows in the area of significance. The only exception is the case of the gender dimension which also shows signs of negative comovements (left-pointing arrows).

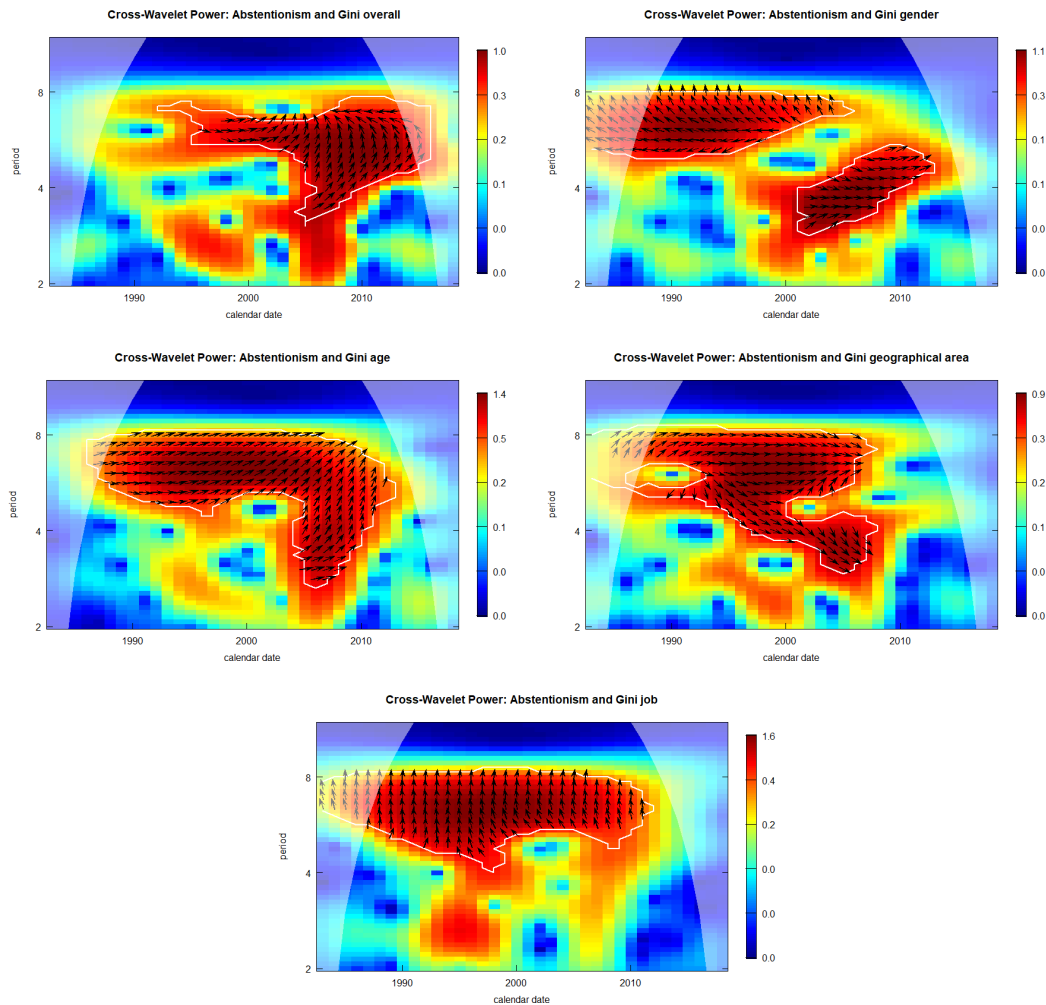


Figure 7: Cross-Wavelet Power between overall and between-groups wage inequality indexes and abstentionism. The red color scale indicates the intensity of the relation and the white line indicates areas of significance at 5% level. Right-pointing arrows indicate positive comovements while left-pointing arrows signal asynchronicity.

To better understand the nature of the different relations observed in Figure 7, it is useful to compare their intensity in Figure 8. Note that the downward trend is expected at the boundaries, as it corresponds to regions outside the cone of influence (Torrence and Compo, 1998). The job qualification index shows the highest intensity along with the age

dimension, followed by the overall Gini index. Moreover, the intensity of the relationship exhibits an overall increasing trend in all categories, except for the geographical dimension. This geographical exception may be attributed to a gradual catch-up, though modest, in abstention rates in the Northern-Central regions relative to the South, which historically recorded the highest levels (Cetrulo, Lanini, et al., 2023). This gradual catch-up likely reflects growing distrust and discontent in the North-Central regions due to impoverishment and deindustrialization processes, progressively reducing the relevance of the geographical component over time. Returning to the overall pattern, the non-linear increasing trend signals a certain degree of inertia—or sluggishness—in the manifestation of the underlying relationships. This pattern may be consistent with the cumulative effects of long-term socioeconomic transformations, which tend to persistently impact the most vulnerable categories of workers: women, individuals in lower hierarchical positions, younger cohorts, and those residing in Southern Italy. These dynamics may gradually crystallize disparities over time, laying the groundwork for self-reinforcing mechanisms and relational patterns that only emerge more clearly in later stages.

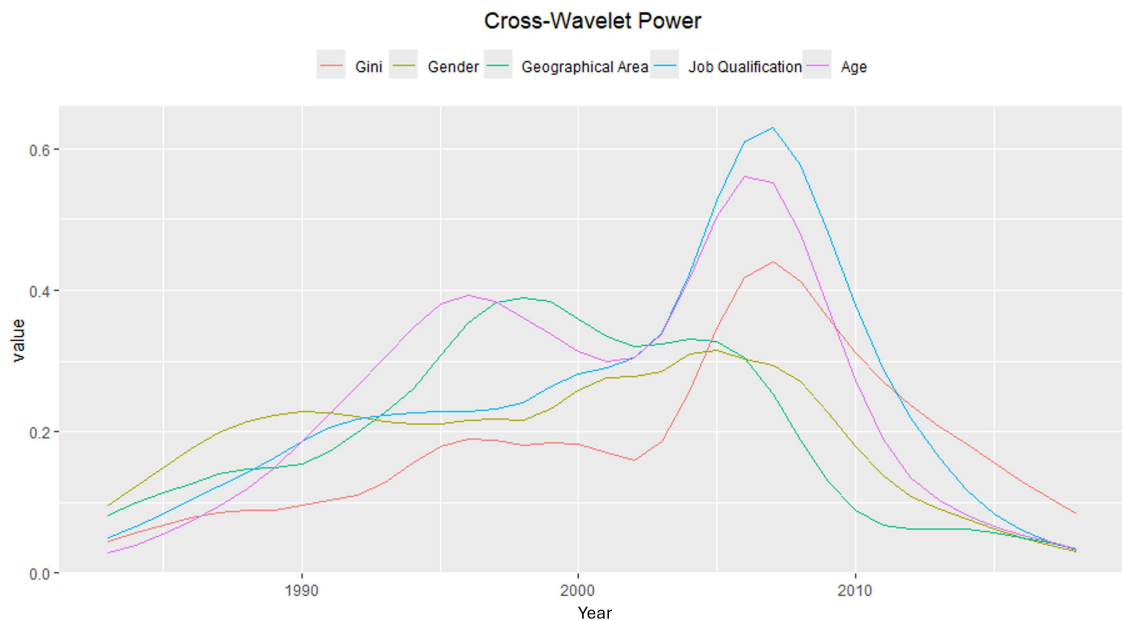
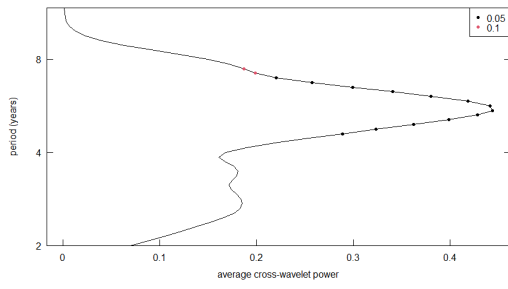


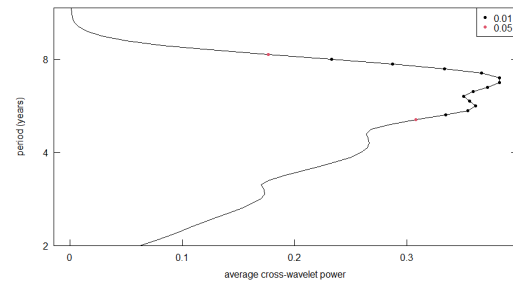
Figure 8: Cross-power intensity.

Moreover, Figures 9 and 10 report the comparison between the intensity of the relation for medium frequencies and for higher frequencies. In particular, the high-frequency intensity (short period) is computed as the average intensity between frequencies in the interval of 2 and 4 years, while the medium-frequency is computed averaging frequencies between 5 and 8 years. What emerges from the analysis is that, in the medium term, the relation is stronger for all measures of inequality. This evidence confirms, therefore, the presence of a significant medium-run relationship between overall and between-group wage inequality indicators and abstention dynamics. In fact, the highest power of the comovement is registered at medium frequencies (identified between 5 and 8 years). Of course, there is variability in time with respect to the short vs medium intensity, in particular there is a general spike in the second part of the time span in the short-run intensity (Figure 10), but overall on average the medium frequency is higher (Figure 9). This information will

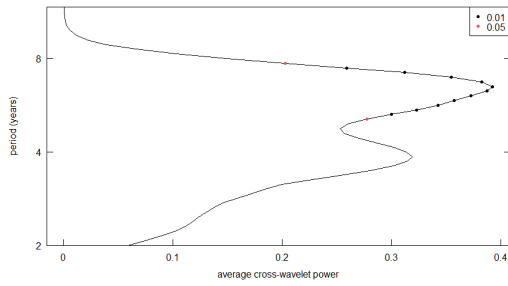
be used to better model the lag structure in the parametric analysis.



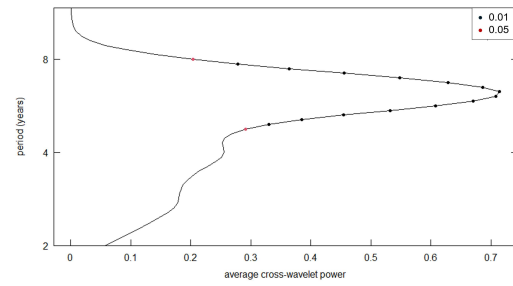
(a) Overall Inequality



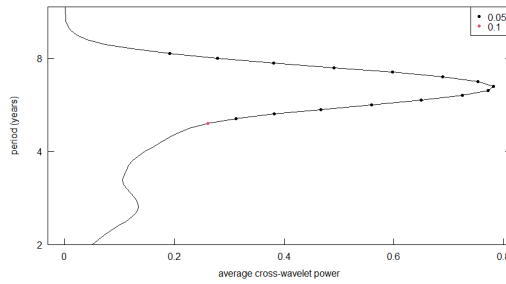
(b) Geographical Inequality



(c) Gender Inequality



(d) Age Inequality



(e) Job Inequality

Figure 9: Average cross wavelet power by category. For each category is reported the average intensity across time by frequency. On the y axis is reported the period in years (2-8 years) and on the x axis the average cross-wavelet power.

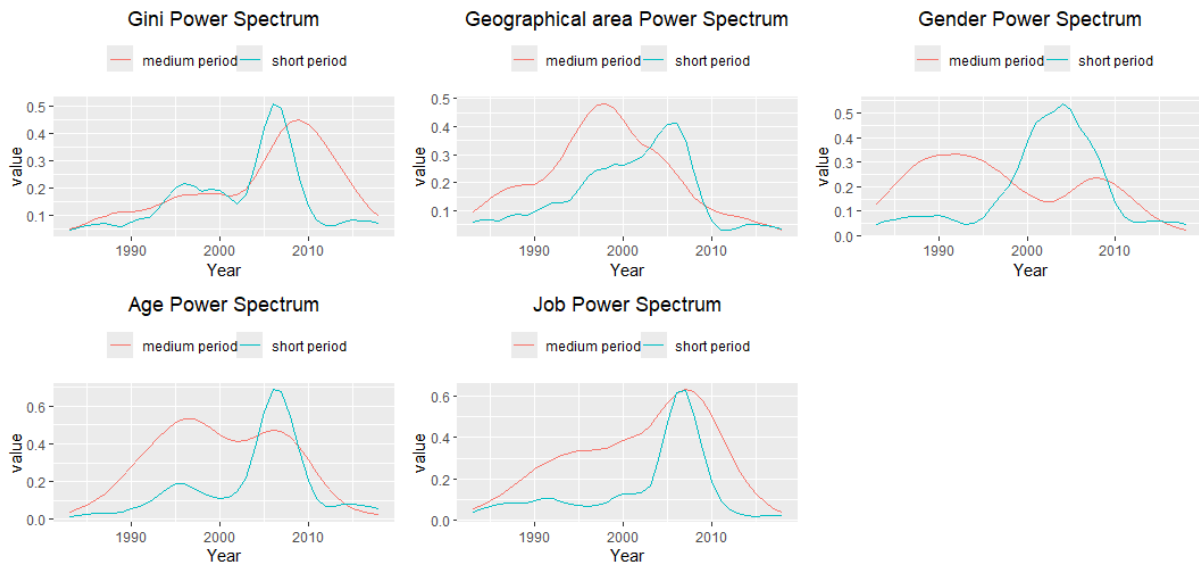


Figure 10: Wavelet Power by frequency range. Medium (5-8 years) and short period (2-4 years) intensity.

5 Parametric Analysis

Building on the preliminary findings, the next step involves a parametric analysis aimed at evaluating the effects of changes in between-category inequality indexes on abstention rates. We construct a panel dataset, whose units of observation are regions through time. Abstentionism and inequality indexes are therefore recomputed at the regional level, exploiting both time and cross-sectional variability.

The main specification used in the analysis is reported in Equation 10. The panel structure is characterized by 20 Italian regions observed over 5–7 non consecutive years between 1983 and 2018, considering election years only.² We estimate a two-way fixed effects panel model including fixed effects for regions (r) and years (t): region fixed effects control for all time-invariant regional characteristics (e.g., geography, institutional features, demographic composition), while year fixed effects absorb common shocks and nationwide trends affecting abstention. The regression is estimated separately for each of the four wage inequality indexes.³ In the corresponding equation, the inequality variable is summarized as X^G , where G indexes the group-specific measure: Overall Gini, Job inequality, Age inequality, and Gender inequality. The key coefficient of interest is computed as the temporal average over the previous five years ($\hat{T} = 5$). This estimator is intended to capture the medium-term time structure highlighted by the wavelet analysis. Given that we observed a predominant pattern of medium-run cyclical relationships, this specification aims to reflect how the effects of inequality on abstention unfold over medium-term horizons, shaped by political cycles and the cumulative impact of socioeconomic policies undertaken in the previous period of interest. Since electoral cycles occur discretely over

²The panel goes from 1983 to 2018 thus counting 10 election years, then the years numerosity in the regression analysis diminishes to 5-7 because of the construction of the main regressors - the lagged 5-year moving average of the inequality indexes.

³We consider only the overall Gini, the gender index, the age index and the job qualification index. The geographical dimension is excluded in this section since we recomputed all the indicators at the regional level, so it was not possible to disaggregate also the geographical indicator.

time, it is reasonable — and has also been empirically supported in the previous section — to assume that the effect is neither immediate nor limited to a one-year lag. Rather, it tends to emerge through medium-term dynamics, either in the form of consolidated trends or more transitory effects within regional or national economies. In addition to regional and year fixed effects, the model includes two controls. These comprise the average age of the population in each region, denoted as \overline{Age} , — also the share of the population over 60 years old was considered and the results were consistent — and an indicator of the dispersion of capital income between dependent workers, denoted as K . This variable is included to account for disparities in material conditions beyond wage income alone (Milanovic, 2019), while data on average age were included coherently with the literature on electoral behavior (File, 2014). Other variables such as employment level and GDP could not be included due to multicollinearity.

$$Abst_{r,t}^G = \beta_0 + \beta_1 \frac{\sum_{j=1}^{\hat{T}} X_{r,t-j}^G}{\hat{T}} + \beta_2 Year_t + \beta_3 Region_r + \beta_4 K_{r,t} + \beta_5 \overline{Age}_{r,t} + \varepsilon_{r,t} \quad (10)$$

The estimation procedure follows a within model with two-way fixed effects and computes Driscoll-Kraay standard errors (Driscoll and Kraay, 1998), robust to general forms of heteroskedasticity, serial correlation, and cross-sectional dependence. These properties are particularly relevant in a context where regions are exposed to common national-level dynamics that may generate correlated errors across units and over time. Unlike classic cluster-robust standard errors, D-K corrections account for both temporal and spatial dependencies simultaneously.⁴

Finally, as the focus of the analysis is on statistical associations rather than causal effects, we do not require strict assumptions on parallel trends or treatment homogeneity. To mitigate potential confounding from time-region specific changes in abstention that could be correlated with wage inequality (i.e., region \times year unobserved shocks), we implement as additional robustness check regressions including region-specific linear trends, reported in Appendix C.2 (Wooldridge, 2010).

5.1 Results

The general findings, presented in Table 1, highlight the heterogeneity of the estimated effects. Specifically, we observe a positive and statistically significant relationship between abstention levels and both the overall inequality index and the inequality index based on job qualifications. In particular, a 1% increase in the five-year average of overall inequality or wage divergence between job qualifications is associated, on average, with a 0.4% increase in abstention. In contrast, the inequality index based on age classes shows a positive but non-significant association, while the gender-based inequality index displays a negative and significant effect. Interestingly, we observe that an increase (or decrease) in the remu-

⁴While Driscoll-Kraay estimators are typically recommended for panels with a longer time dimension ($T \geq 20$), recent Monte Carlo evidence (e.g., Hoechle, 2007) suggests that they perform reasonably well even in panels with small T . Although T is relatively small (5–7), the years included span a long historical period (1983–2018), covering multiple electoral cycles and diverse political and socioeconomic environments. This increases the potential for correlated errors across units and over time, reinforcing the rationale for using Driscoll-Kraay standard errors to obtain robust inference.

neration gap between male and female workers is associated with a decrease (or increase) in the abstention rate of approximately 0.37%. This pattern can be traced back to the underlying conditions and structural characteristics driving labor market trends in relation to the gender inequality index, essential to understanding the nature of this relationship (Section 5.1.2).

The positive coefficients suggest the presence of a reverberation channel from economic inequality and class stratification to voting behavior and political inequality at the regional level and over time. The sign of the coefficients supports the prediction of the Power Theory. In fact, unfavorable material and labor market conditions positively influence non voting, showing the presence of a transmission channel that departs from structural economic marginalization and lands to political marginalization.

Table 1: The table reports the results of the estimation associated with Equation 10. A two-way fixed-effects within model is used, and Driscoll–Kraay standard errors are applied to account for cross-sectional dependence and serial correlation.

<i>Dependent variable: Percentage Abstention</i>				
	(1)	(2)	(3)	(4)
Gini overall	0.466*** (0.166)			
Ineq job		0.362*** (0.124)		
Ineq age			0.161 (0.233)	
Ineq gender				−0.371** (0.158)
K	0.258*** (0.063)	0.217*** (0.067)	0.271*** (0.070)	0.247*** (0.058)
Age	−0.781*** (0.241)	−1.024*** (0.220)	−0.956*** (0.226)	−0.803*** (0.278)
<i>Two-way fixed effects</i>	Yes	Yes	Yes	Yes
<i>Unbalanced Panel:</i>				
n = 20, T = 5-7, N = 138	Yes	Yes	Yes	Yes

Note: Driscoll-Kraay standard errors in parentheses

Note: *p<0.1; **p<0.05; ***p<0.01

In Table 2, we present additional results including an interaction term between our main explanatory variables and a structural break variable (Equation 11). The Bai–Perron analysis in Section 3 identifies five breaks over the long-run series of abstention.

For the regression analysis, we adopt a more parsimonious specification with a single structural break in 2007 — the variable *Break* has two levels, pre-2007 and post-2007. This

choice is motivated by the very limited number of election years in the regression sample, which makes estimating multiple period-specific effects statistically fragile. Moreover, the year 2007 represents a substantive turning point as it anticipates the global financial crisis and the subsequent economic and political transformations. Focusing on the 2007 break therefore balances statistical reliability and theoretical relevance. Moreover, 2007 is also the most appropriate point to introduce the structural shift, which is most central in the distribution of the time sample we employ in the regression.

$$Abst_{r,t}^G = \beta_0 + \beta_1 \frac{\sum_{j=1}^{\hat{T}} X_{r,t-j}^G}{\hat{T}} * Break_t + \beta_2 Year_t + \beta_3 Region_r + \beta_4 K_{r,t} + \beta_5 \overline{Age}_{r,t} + \varepsilon_{r,t} \quad (11)$$

Notably, the association coefficient increases over time, in line with the increasing intensity of the wavelet power shown in Figure 8, with the exception of the gender dimension.

Table 2: The table reports the results of the estimation correspondent to Equation 11. The estimation was performed through a two-way fixed effect within model, computing Driscoll-Kraay standard errors. Here are reported only the interaction coefficients of interest, the full table is available in the Appendix.

<i>Dependent variable: Percentage Abstention</i>				
	(1)	(2)	(3)	(4)
Gini overall * post 2007	0.564*** (0.207)			
Ineq job * post 2007		0.205*** (0.067)		
Ineq age * post 2007			0.259* (0.137)	
Ineq gender * post 2007				-0.008 (0.249)
<i>Two-way fixed effects</i>	Yes	Yes	Yes	Yes
<i>Unbalanced Panel:</i>				
n = 20, T = 5-7, N = 138	Yes	Yes	Yes	Yes

Note: Driscoll-Kraay standard errors in parentheses

Note: *p<0.1; **p<0.05; ***p<0.01

Moving to the additional covariates in Table 1, we observe quite a stable homogeneity in the results across the four specifications. In fact, we observe a general negative relation between mean age and percentage of abstention, which is in line with the literature connecting age patterns and voting (File, 2014), while, in line with the expectation of Power Theory, we find a positive relation between inequality in capital income and percentage of abstention. This evidence is noteworthy because it highlights how different forms of

material conditions influence voting behavior. In addition, the coefficient is smaller than that of the wage inequality, consistent with the fact that, first, capital income is only a component of inequality, and it is also important to account for wage inequality, as the latter exerts a notable role; second, the capital income dimension — particularly the financial sector — is a variable that suffers from under-reporting and might be underestimated. The negative contribution of capital income concentration to voting behavior also reflects that concentration of economic power — what the literature has called the power of the *élite* (Franko et al., 2016; Gilens and Page, 2014) — might discourage electoral activation.

5.1.1 Age

Although the baseline results show no significant association between intergenerational wage divergence and abstention, this may obscure important territorial differences. In particular, the Italian labor market is characterized by strong regional disparities: many younger individuals from the South migrate to the North and Center in search of better opportunities, suggesting that wage divergence between age groups may have different political implications depending on the region.

To investigate the interaction between macro-areas and age, we estimate the model in Equation 10 separately for Northern-Central and Southern regions. Although this reduces the number of cross-sectional units in each subsample, the results are informative. In Table 3, the wage divergence variable is statistically significant in the North-Center, where active young workers are more integrated in formal employment, which could imply greater relevance of the inequality variable in this framework. In contrast, no significant association is found in the South, possibly due to weaker ties to the labor markets among younger cohorts or greater structural exclusion (Odoardi and Liberatore, 2021). Notably, the sign of the coefficient for the North-Center is positive, in line with the overall results, confirming the Power Theory hypothesis. These results should be interpreted with caution due to the reduced sample size, but still highlight how territorial heterogeneity can condition the relationship between wage-age inequality and political behavior.

Table 3: The table reports the results of Equation 10 estimated separately for North-Central and Southern regions to investigate how the relationship between abstention and the age inequality index differs across macro-areas. A two-way fixed-effects within model with Driscoll–Kraay standard errors is estimated.

<i>Dependent variable: Percentage Abstention</i>		
	North-Center	South
Ineq age	0.738*** (0.213)	-0.351 (0.256)
K	-0.022 (0.079)	0.232** (0.091)
Age	0.251 (0.156)	2.988*** (0.663)
<i>Two-way fixed effects</i>	Yes	Yes
<i>Unbalanced Panel:</i> n = 12, T = 5-7, N = 82	Yes	No
<i>Balanced Panel:</i> n = 8, T = 7, N = 56	No	Yes

Note: Driscoll-Kraay standard errors in parentheses

Note: *p<0.1; **p<0.05; ***p<0.01

To rationalize our finding, we checked the number of employment contracts by age group in Figure 11. The data show that disparities between the North-Center and the South are particularly marked for the youngest cohort (under 30), with substantial differences in both medians and means of employment contracts. This highlights that young people in the North-Center are significantly more active in the labor market compared to their Southern peers, likely due to greater job opportunities and territorial mobility, for example for study or work reasons. For other age cohorts, particularly in the 31-50 year class, the difference in the median employment contracts is lower, while the difference in mean level is higher, suggesting greater variability and the presence of outliers in the distribution. For workers over 50, territorial differences are more limited (Table 4). The distribution by age of the number of contracts differentiated across geographical areas corroborates our expectation.

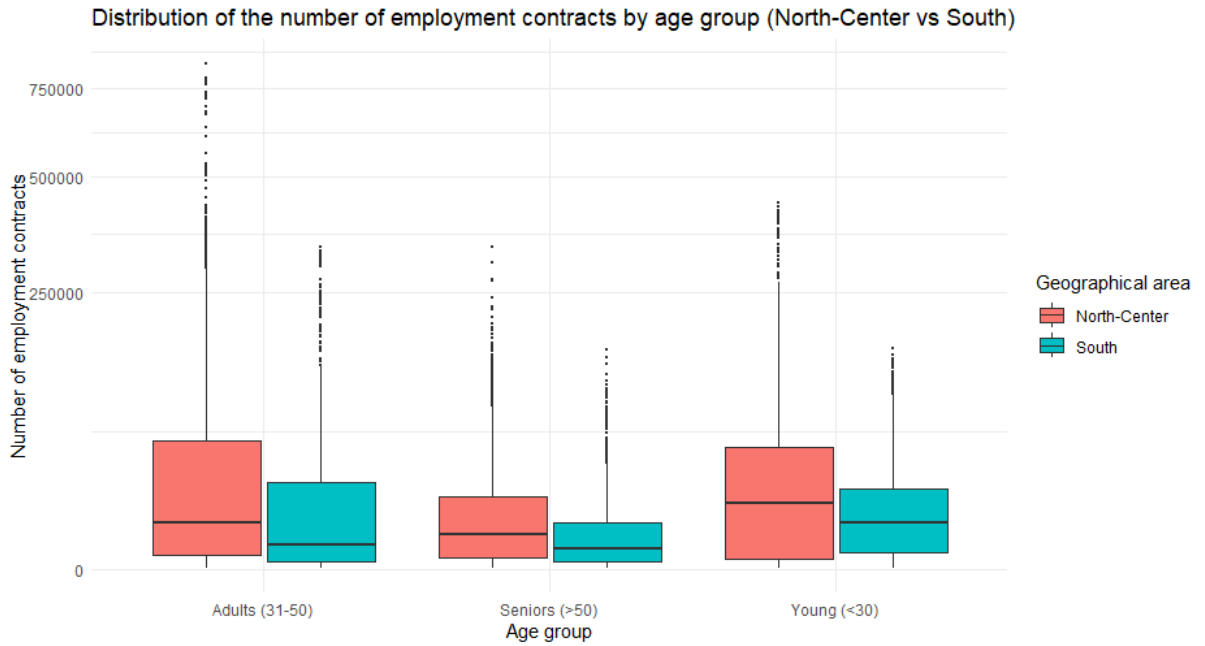


Figure 11: Employment contracts by age group and by area.

Table 4: Summary statistics of employment contracts by age group and territorial area

Age Group	Diff. Median	Diff. Mean
Young (≤ 30)	7,404	19,853
Adults (31-50)	5,092	29,157
Seniors (>50)	2,507	8,124

5.1.2 Gender

To address the negative association between political abstention and the gender wage inequality index, we assess the underlying dynamics driving the evolution of the indicator. The index has remained relatively stable over time in most regions, oscillating between 15 and 25%, with minor downward trends in a few cases. At the national level (Figure 3), the index shows some fluctuations — first an increasing, then a decreasing phase starting from the 2000s — but overall displays long-term stability over the past 36 years.

This stability can be understood by considering two opposing channels in the labor market. On the one hand, there has been a substantial increase in female labor force participation in Italy, especially since the late 1990s, as shown in Figure 12. This trend tends to reduce the gender wage inequality, as a broader inclusion of women in the labor market mitigates the wage gap. However, the increase in participation has been accompanied by a growing share of women employed in low-paid and temporary positions (Figure 12), particularly in the tertiary sector. These positions are often involuntary, precarious and lacking career progression, which exerts upward pressure on wage inequality.

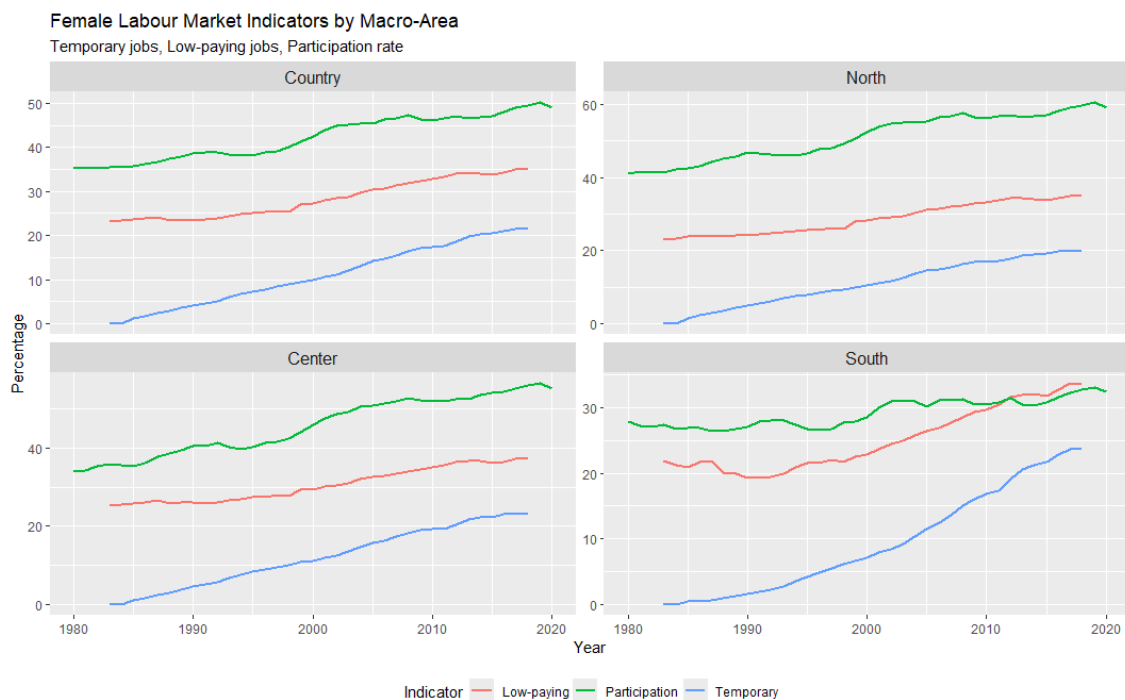


Figure 12: Female labor market indicators.

The interplay of these two channels — a participation channel pushing inequality down and a segregation/low job quality channel pushing it up — produces an overall stability in the gender inequality index. Importantly, the observed negative coefficient of gender inequality in its association with abstentionism suggests that the participation channel has prevailed in shaping the inequality trend, keeping it relatively flat. Thus, under a Power Theory interpretation of political and economic exclusion, despite stable or slightly declining gender wage inequality, political engagement has not improved. The concurrent increase in female labor force participation and the concentration in low-wage, precarious jobs have stabilized a structural pattern: women are more present in the workforce, yet remain largely segregated into low-quality positions, a dual dynamic that contributes to hinder political activation and entrenches existing inequalities.

5.1.3 Comparison and Relative Contribution of the Different Indicators

To compare the relative importance of the different inequality variables, we estimate a standardized OLS model in which both the dependent variable and the main regressors are expressed in standard deviations. The results are shown in the first column of Table 5. This approach allows for a direct comparison of the magnitude of the coefficients, irrespective of the original measurement units. The results indicate that income inequality between job types and overall Gini is associated with larger effects on abstention compared to gender- or age-based wage gaps. This suggests that structural and occupation-related inequality dimensions may play a more specific role in shaping political disengagement.

Finally, to quantify the contribution of each inequality variable to the explanatory power of the model, we also compute adjusted partial R^2 values, reported in the second column of Table 5. The explained variance is limited but not negligible, and overall inequality and job-wage inequality exhibit higher explanatory power (4 - 8.5%). The limited

net explanatory power is consistent with the idea that abstention is a complex phenomenon whose roots go beyond our included explanatory variables.

Table 5: The table presents in column (1) the standardized coefficients corresponding to Table 1 standardized results, while in column (2) reports the correspondent adjusted R^2 for the main explanatory inequality variables.

<i>Dependent variable: Percentage Abstention</i>		
	Scaled Coefficient	Partial R^2
Gini overall	0.199	4.1%
Ineq job	0.171	8.5%
Ineq age	0.069	0.1%
Ineq gender	-0.102	2.1%

6 Robustness: Interactive Fixed Effects and Latent Factors

We conduct a robustness test that estimates the model with interactive fixed effects, as in Bai (2009), to overcome the restrictive additivity assumption of region and year fixed effects by allowing for unobserved common factors that affect units heterogeneously over time, thus capturing and exploring unobserved yet persistent dynamics, potentially linked to economic, cultural, institutional, or historical factors that are difficult to quantify.

In Table 6 we observe the estimated results of Equation 12. When accounting for unobserved components, only the job qualification and age inequality measures remain statistically significant, whereas the others lose explanatory power. This extension serves to unpack the large share of variance absorbed by the region- and year-fixed effects in the baseline models. Given the relatively low within-year variability of our inequality measures and the fact that latent factors can themselves capture macroeconomic dynamics — such as structural shifts in the labor market or national economic trends — it is not surprising that some of our indicators lose statistical significance in this specification. Rather than weakening the argument, this result highlights the complexity of political abstention as a phenomenon shaped by overlapping economic, institutional, and cultural dimensions, some of which may only be indirectly proxied by observable economic variables.

$$Abst_{r,t}^G = \beta_0 + \beta_1 \frac{\sum_{j=1}^{\hat{T}} X_{r,t-j}^G}{\hat{T}} + \lambda_r * f_t + \beta_2 Year_t + \beta_3 Region_r + \beta_4 K_{r,t} + \beta_5 \overline{Age}_{r,t} + \varepsilon_{r,t} \quad (12)$$

Table 6: Interactive Fixed Effects Models corresponding to Equation 12.

	<i>Dependent variable: Percentage Abstention</i>			
	(1)	(2)	(3)	(4)
Ineq job	0.383*** (0.040)			
Gini overall		0.390 (0.405)		
Ineq age			0.449*** (0.077)	
Ineq gender				0.269 (0.182)
K	-0.004 (0.012)	0.012 (0.012)	0.005 (0.017)	-0.022*** (0.005)
Age	0.421 (0.149)	0.920 (0.178)	1.270*** (0.146)	0.811*** (0.082)
<i>Interactive Fixed Effects</i>	Yes	Yes	Yes	Yes

Note: Standard errors in parentheses.

Note: * $p < 0.1$; ** $p < 0.05$; *** $p < 0.01$

We briefly discuss the nature of the two factors emerging from the analysis which capture distinct but complementary dynamics behind regional abstention levels. Figure 13 displays the trends, the first factor in black, while the second in red. Moreover, regional factor loadings are reported in Figures 14 and 15 respectively. The first aspect to note concerns loadings. In fact, for both factors, we can observe a divide between the central-northern and southern regions. In particular, the two macroregions show opposite signs, signaling a structural division in terms of latent responses of the two factors isolated in the analysis.

The first factor shows a decreasing trend until 2005, followed by a period of relative stability and another of slight decline after 2014. Its positive loadings are concentrated in Southern regions (e.g., Calabria, Campania, Sicily), while Northern regions have negative loadings. This factor could reflect the persistent structural economic and institutional gap between the North and South of Italy, which in turn has determined differentiated levels of institutional trust. Over time, this gap has appeared to be shrinking, possibly due to a relative convergence in low levels of institutional trust, as signaled by the progressive acceleration of abstention also in Northern regions. In particular, between 2008 and 2022, abstention more than doubled in the Center (+103%) and in the North (+99%), compared to a +93% increase in the South. Thus, while Southern regions continue to display structurally higher abstention in levels, the acceleration observed in the Center and North suggests a convergence driven by a deterioration of institutional trust also in these areas.

The second factor exhibits a clear cyclical pattern: it declines before 2005, increases

until around 2008, drops sharply during the financial and PIIGS crises (2008–2014) and rises again in the post-crisis years. Its positive loadings are mostly in wealthier Northern and Central regions (e.g., Lombardy, Emilia-Romagna, Veneto), while Southern regions have negative loadings. This factor could capture short- to medium-term fluctuations in abstention linked to distrust or disaffection arising after economic shocks and, therefore, more genuinely linked to the political offer. In particular, the 2006 election represents a turning point in lowering abstention rates. The 2000s in general were characterized by high volatility in electoral cycles, with governments lasting less than 5 years.

Together, the two factors explain both structural long-run lasting divides in electoral behavior, distinctly characterizing the northern-central toward the southern area (the first factor), and cyclical changes in electoral phases (second factor). Evidence suggests that abstention is shaped both by long-term structural divides between North and South and by regionally differentiated responses to economic and institutional crises, highlighting the multidimensional nature of electoral participation in Italy.

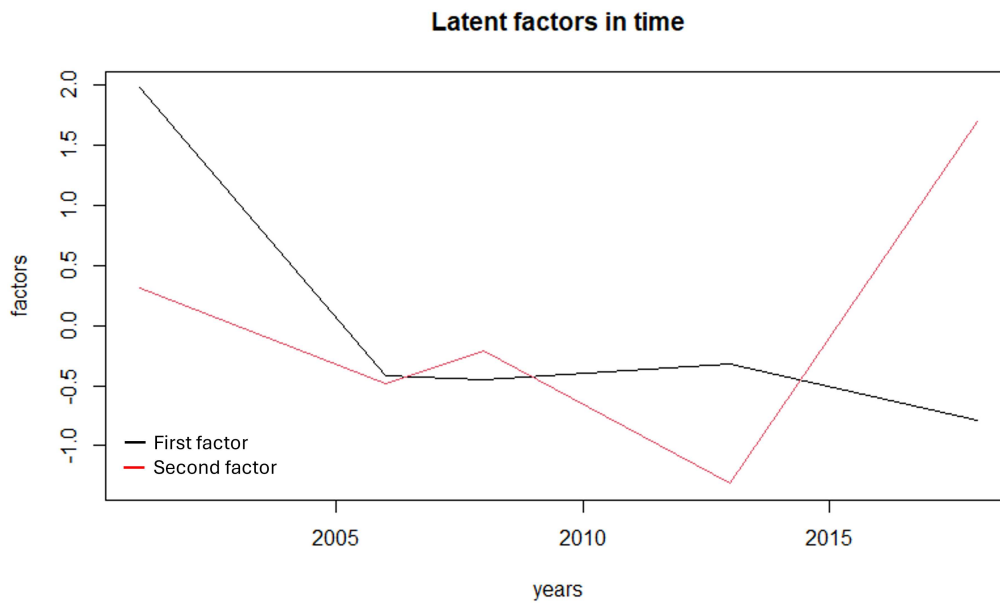


Figure 13: Latent factors resulting from the interactive fixed effect extension. In particular, these factors correspond to the Job qualification model.

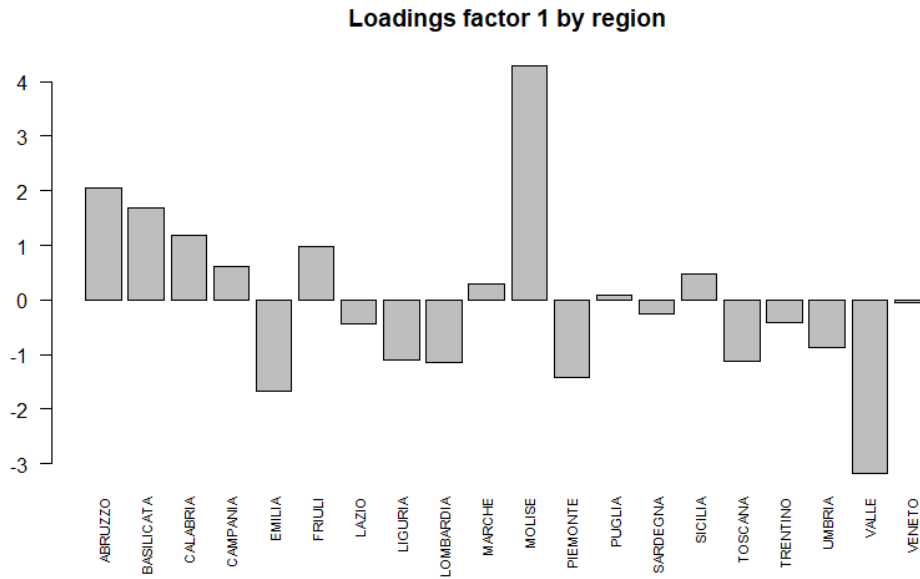


Figure 14: Regional loadings Factor 1.

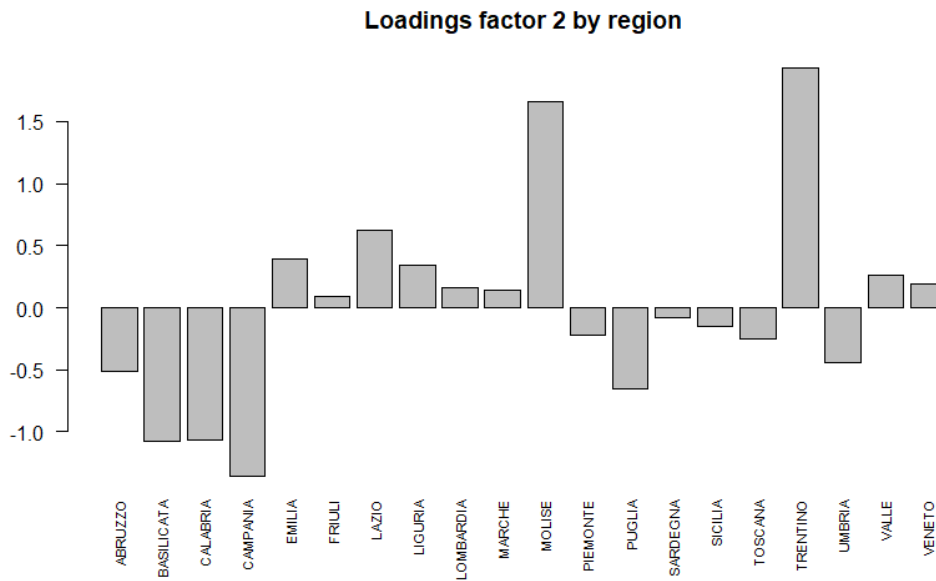


Figure 15: Regional loadings Factor 2.

7 Concluding Remarks

In recent years, the link between economic and political inequality has emerged as one of the most pressing questions. This renewed urgency is rooted in the broader debate on the compatibility between capitalism and democracy, which has gained attention through the rise of antagonistic narratives questioning their coexistence. At the core of the problem lies a fundamental puzzle: how can rising trends in income inequality be reconciled with the legitimacy of democratic processes if political engagement and influence remain deeply shaped by unequal distributions of economic resources?

This study contributes to the theoretical understanding of the relationship between

economic inequality and political behavior in several ways. First, it provides support for Power Theory by linking unequal access to economic resources with political disengagement. Second, it moves beyond the traditional focus on political supply side explanations of political participation, emphasizing the importance of structural economic factors that have become entrenched over time. Third, the findings highlight the emerging role of capital income distribution in modulating the interaction between modern capitalism and democracy, showing how the concentration of economic power can drive processes of marginalization.

Overall, the work underscores the non-linear, medium-term dynamics through which economic inequalities influence political outcomes, offering a framework for analyzing these mechanisms within broader economic and institutional structures. On a societal level, the results raise concerns about the political marginalization of specific social groups. As economic inequalities deepen and labor market segmentization hardens, particularly affecting the most vulnerable workers, the risk of cumulative disenfranchisement increases. Processes of stratification and the reproduction of economic inequalities can undermine democratic legitimacy, highlighting the need for policies aimed at mitigating these structural disadvantages. Understanding these patterns is therefore crucial not only for scholars of inequality and political behavior but also for policymakers seeking to preserve the inclusiveness of decision-making processes. Future research could expand the temporal and spatial coverage, examine additional dimensions of inequality and material conditions, and further explore the mechanisms linking labor market structures, elite power, and political disengagement.

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Appendix

A Methodology

This appendix section is meant to illustrate in greater detail the main techniques utilized in the analysis. The implemented non-parametric method is the Wavelet analysis along with the Christiano-Fitzgerald filter, which are mainly meant for the study of time series. In the last part of the paper, standard parametric regression analyses are performed. Let us now better characterize the wavelet methodology given that is less widespread in the economic literature with respect to parametric regression, details on the Christiano-Fitzgerald filter can be also found below. Finally, the last section of the methodology appendix is dedicated to the interactive fixed effects regression models, implemented in the robustness section of the paper.

A.1 The Wavelet Analysis

A Wavelet is a wave-like oscillation that is localized in time. Wavelets have two basic properties: scale and location. Scale defines how “stretched” or “squished” a wavelet is. This property is related to the frequency as defined for waves. The location defines where the wavelet is positioned in time (or space).

Wavelet analysis is a method to study frequency changes across time, trying to find a better balance in the trade off between the time and the frequency domain, known as the time–frequency dilemma. The frequency structure of the time series is analyzed in this work using the Morlet wavelet, a complex-valued and continuous wavelet which leads to a complex wavelet transform of the time series. These properties allow for information preservation. The transform can be separated into its real and imaginary parts, thus providing information on local amplitude and instantaneous phase of any periodic process. The method performs well both with stationary and non stationary time series, as well as when breaks are present. This is because the method allows for a decomposition of all the periodic characteristics of the time series to be a function of time, which means that unit root analysis is not a prerequisite in these studies.

From the wavelet transform of a time series $x(t)$ it is possible to construct the wavelet power spectrum, which is the square of the amplitude and has an interpretation as time–frequency wavelet energy density. It gives the indication of to what extent a time series is volatile across different time scales.⁵ The power spectrum analysis corresponds to the local variance of a time series across time and across frequencies. The wavelet power spectrum is an energy density in the time–frequency plane. The horizontal axis denotes the time and the vertical axis the period measured in the time unit considered in the analysis (years in the present study). In this work, the power spectrum coefficients are represented by color intensity: warmer colors stand for higher power. The bold contours enclose regions where the wavelet power is significant at the 5% level, as determined by Monte Carlo simulations under the assumption that the time series is generated by a white noise process. This

⁵The Morlet wavelet is defined as $\psi(t) = \pi^{-1/4} e^{-i\omega t} e^{-t^2/4}$ and the wavelet transform of a time series using the Morlet wavelet can be written as $W_x(\tau, s) = \int_{-\infty}^{+\infty} x(t) \frac{1}{\sqrt{|s|}} \psi * \left(\frac{t-\tau}{s}\right) dt$, from this is possible to construct the wavelet power spectrum $WPS = |W_x(\tau, s)|^2$.

suggests the presence of time-localized periodic structure beyond what would be expected from random fluctuations. The cone of influence identifies the regions affected and not affected by the edge effect.

Considering the relation between two time series it is possible to extend the analysis to the bi-variate level using wavelet cross spectrum and wavelet coherence. The wavelet cross spectrum provides appropriate tools for comparing the frequency contents of two time series and to draw conclusions about the series synchronicity at certain periods and across certain ranges of time. The cross-spectrum will have positive values for a phase shift between $[0, 90^\circ)$, negative values between $(90^\circ, 270^\circ)$ and again positive values between $(270^\circ, 360^\circ]$. A complementary measure is the Wavelet Coherence, that is the "normalized" cross-spectrum and is near to the concept of correlation⁶. It analyzes how the time series are correlated across different timescales. The coherence is computed by taking the square value of the cross-spectrum. Moreover, thanks to the smoothing parameter coherence analysis "smooths away" not stable or not synchronized relations — for instance, cases where the phase may vary significantly or where there may be no consistent correlation within the smoothing range. Consequently, the wavelet coherence, by applying smoothing and normalization, attenuates unstable relationships (Rösch and Schmidbauer, 2016; Kumar, 2018; Menegaki, 2020).

Taken together, wavelet coherence is a stringent measure for defining a relationship between two time series because it requires not only simultaneous energy but also a stable relationship (e.g., a consistent phase difference) within a smoothed window. In contrast, the cross-wavelet power spectrum effectively identify time-localized common oscillatory behavior in the time series, regardless of the stability or noisiness of their relationship.

A.2 The Christiano-Fitzgerald Filter

In general, the Christiano-Fitzgerald filter separates a time series into a trend component and cyclical component. Given a time series $\{x_t\}_{t=1}^T$ we are interested in isolating the component of x_t denoted y_t with period of oscillations between p_l and p_u , where $2 \leq p_l < p_u < \infty$. The parameters p_u and p_l are the cut-off cycle length. Cycles longer than p_l and shorter than p_u are preserved in the cyclical term. Consider the following decomposition of the time series:

$$x_t = y_t + \bar{x}_t, \quad (13)$$

the component y_t is assumed to have power only in the frequencies in the interval $\{(a, b) \cup (-a, -b)\} \in (-\pi, \pi)$, and \bar{x}_t has power only in the complementary interval in $(\pi, -\pi)$. a and b are related to p_l and p_u by $a = \frac{2\pi}{p_u}$ and $b = \frac{2\pi}{p_l}$.

If infinite amount of data is available, then we can use the ideal bandpass filter:

$$y_t = B(L)x_t, \quad (14)$$

⁶The cross-spectrum is defined as $W_{(x_j y_j)}(\tau, s) = W_x^*(\tau, s)W_y(\tau, s)$, its modulus can be interpreted as cross-wavelet power. In this formulation "*" stands for the complex conjugate of the wavelet transform (Intuitively, complex conjugation of a wave function describes the time-reversed wave function traveling backward in time). The wavelet coherence between two time series $x(t)$ and $y(t)$ is defined as $R_{(x,y)}^2 = \frac{|S(s^{(-1)}W_{(x_j y_j)}(\tau, s))|^2}{(S(s^{(-1)}|W_{(x_j)}(\tau, s)|)S(s^{(-1)}|W_{(y_j)}(\tau, s)|))}$.

where the filter, $B(L)$, is given in terms of the lag operator L and defined as $B(L) = \sum_{j=-\infty}^{\infty} B_j L^j$ and $L^k x_t = x_{t-k}$. The ideal bandpass filter weights are given by $B_j = \frac{\sin(jb) - \sin(ja)}{\pi_j}$ and $B_0 = \frac{b-a}{\pi}$.

The finite sample approximation to the ideal bandpass filter uses the alternative filter:

$$\hat{y}_t = \hat{B}(L)x_t = \sum_{j=-f}^p \hat{B}_j^{p,f} x_{t-j}, \quad (15)$$

where the weights, $\hat{B}_j^{p,f}$, of the approximation are a solution to $\hat{B}_j^{p,f} = \arg \min E\{(y_t - \hat{y}_t)^2\}$, $p = t - 1$ and $f = T - t$.

The Christiano-Fitzgerald filter is a finite data approximation to the ideal bandpass filter and minimizes the mean squared error defined in the above equation (Christiano and Fitzgerald, 2003).

A.3 Interactive Fixed Effects

The interactive fixed effects (IFE) model (Bai, 2009) generalizes the standard two-way fixed effects model by incorporating unobserved common factors that vary over time and affect cross-sectional units heterogeneously. The model is specified as:

$$Y_{it} = X'_{it}\beta + \lambda'_i F_t + \alpha_i + \gamma_t + \nu_{it}, \quad (16)$$

where F_t is a vector of unobserved common factors, λ_i are unit-specific factor loadings, α_i and γ_t denote individual and time fixed effects, and ν_{it} is an idiosyncratic error term.

In contrast, the standard two-way fixed effects model only includes α_i and γ_t , implicitly assuming that residuals are cross-sectionally uncorrelated. In the presence of common shocks or omitted time-varying variables, this assumption is violated. The IFE model addresses this by decomposing the residuals of the conventional model as:

$$\varepsilon_{it} = \lambda'_i F_t + \nu_{it}, \quad (17)$$

where the term $\lambda'_i F_t$ captures the unobserved interactive effects. The factors F_t and loadings λ_i can be consistently estimated via principal component analysis (PCA) on the residuals from an initial fixed effects regression. This allows for flexible cross-sectional dependence and reduces bias in the estimation of β caused by unobserved common factors.

B Statistics

B.1 Gender section

This section clarifies the computation of the share of low paying jobs and part time jobs represented in Figure 12. First of all, the average wage by ATECO sector was computed as a weighted average of wages, using the number of jobs in each sector as weights, as shown in Figure 16. In this figure, the wage trends are reported separately for each ATECO sector, along with the overall average. Figure 17, instead, shows only the aggregated average wages, grouping together sectors with wages above the mean and those with wages below the mean. Then, the following step was to sum the total number of jobs in sectors paying

a salary below the average and then computing the share of "low paying jobs" employing women.

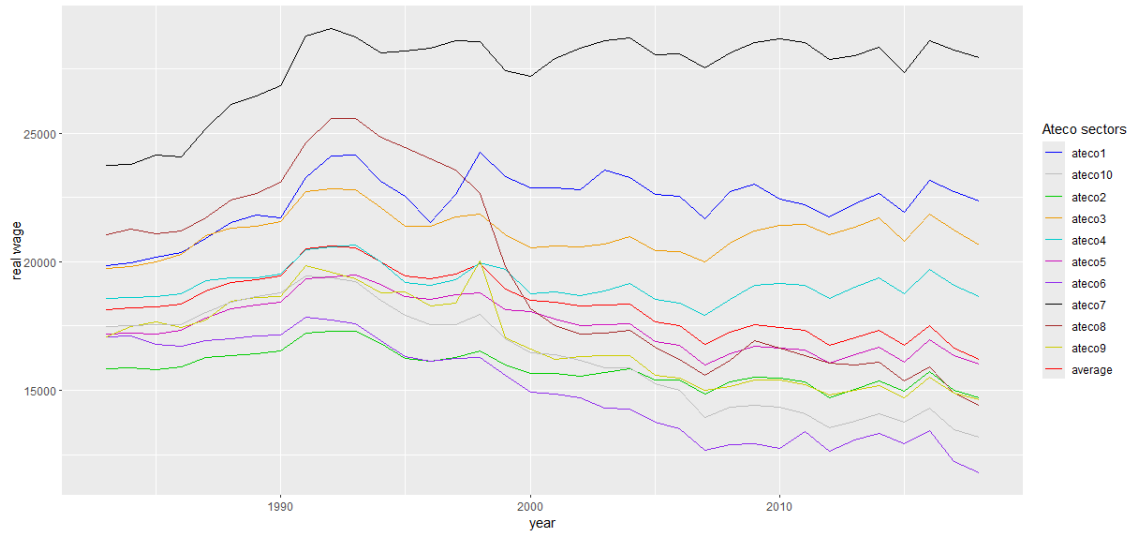


Figure 16: Average wage level by ateco sector aggregates.

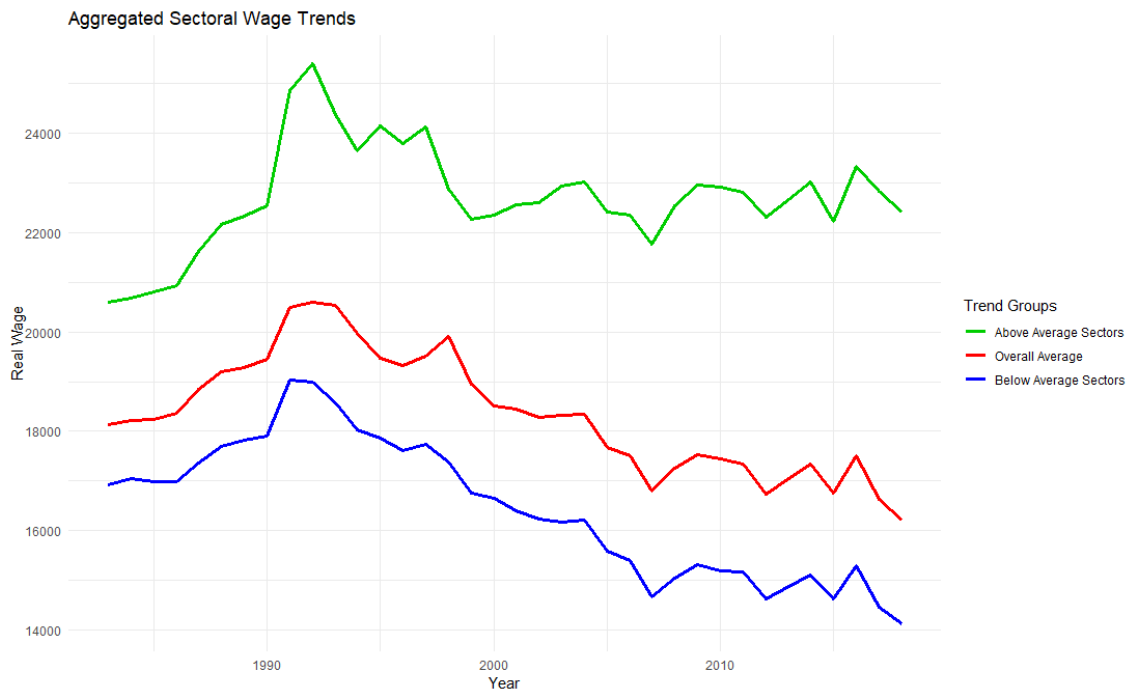


Figure 17: Aggregated average wage of sectors above and below the mean.

Similarly, also the evolution of the share of part-time jobs was computed as a sum of the shares of part-time jobs employing women in the economy.

C Parametric analysis

C.1 Baseline regression

Here we report few complementary tables to the parametric analysis and results sections. In Table 18, correlation levels between covariates are reported.



Figure 18: Correlation Plot.

In Table 7 we report the full Table 2 presented in Section 5.1.

Table 7: The table reports the complete estimation results related to Table 2.

	<i>Dependent variable: Percentage Abstention</i>			
	(1)	(2)	(3)	(4)
Gini overall	0.472*** (0.143)			
Ineq job		0.379*** (0.096)		
Ineq age			0.144 (0.237)	
Ineq gender				-0.333* (0.198)
post 2007	-9.267 (5.923)	-5.548 (3.646)	-4.059 (3.855)	4.556 (4.555)
K	0.222*** (0.059)	0.188*** (0.064)	0.248*** (0.052)	0.233*** (0.070)
Age	-0.833*** (0.214)	-1.299*** (0.154)		-0.978*** (0.332)
Gini overall * post 2007	0.564*** (0.207)			
Ineq job * post 2007		0.205*** (0.067)		
Ineq age * post 2007			0.259* (0.137)	
Ineq gender * post 2007				-0.008 (0.249)
Constant				46.005*** (9.830)
<i>Twoways Fixed Effect Model</i>	Yes	Yes	Yes	Yes

Note: Driscoll-Kraay standard errors in parentheses

Note: *p<0.1; **p<0.05; ***p<0.01

C.2 Additional Robustness Tests

Table 8 reports estimated results following Equation 10, using not only election years but rather the linearly interpolated time series of abstention at the regional level to increase the number of observations and exploit fully available information of the independent

regressors. To account for the interpolation, a linear time trend is included as control. Results are robust to this specification and show similar coefficients and significance levels as in the baseline analysis (Table 1).

Table 8: The table reports the results of Equation 10 estimated on linearly interpolated regional abstention series. A two-way fixed-effects within model is used, with Driscoll–Kraay standard errors. A linear time trend is included to account for the interpolation.

<i>Dependent variable: Percentage Abstention</i>				
	(1)	(2)	(3)	(4)
Gini overall	0.563*** (0.210)			
Ineq job		0.347* (0.209)		
Ineq age			−0.281 (0.241)	
Ineq gender				−0.590*** (0.134)
Time trend	0.320*** (0.105)	0.481*** (0.058)	0.581*** (0.126)	0.477*** (0.038)
Age	−0.622*** (0.221)	−0.788*** (0.230)	−0.945*** (0.240)	−0.577*** (0.196)
K	0.157** (0.068)	0.154** (0.065)	0.156** (0.069)	0.143* (0.074)
<i>Twoways Fixed Effect Model</i>	Yes	Yes	Yes	Yes
<i>Unbalanced Panel:</i>				
<i>n = 20, T = 21-26, N = 515</i>	Yes	Yes	Yes	Yes

Note: Driscoll-Kraay standard errors in parentheses

Note: *p<0.1; **p<0.05; ***p<0.01

Table 9 reports an additional robustness test, which entails to control for regional specific time trends to account for year x region specific shocks. Results are consistent with previous estimations, indeed coefficients and p-values are similar. While the coefficient of *Ineq job* loses statistical significance in this specification, its p-value remains around 0.13, not far from the level observed in Table 8, around 0.1. In addition, we observe that the significance of region-specific time trends is particularly pronounced in the northern-central regions, which are generally wealthier, exhibit greater inequality due to more complex labor markets, and have a higher share of managerial positions. The estimation highlights that the upward trend in job inequality (and its relation to abstention) is significantly driven by these northern regions, suggesting that the increase is absorbed to a larger extent in

areas with a higher incidence of managerial roles and more stratified labor structures. This pattern reinforces the interpretation that regional economic and labor market characteristics play a key role in shaping the dynamics of job-related inequality captured by our regressions.

Table 9: The table reports an extension of the results in Table 8 including region-specific linear time trends to account for region–year specific shocks. A two-way fixed-effects within model with Driscoll–Kraay standard errors is employed.

	<i>Dependent variable: Percentage Abstention</i>			
	(1)	(2)	(3)	(4)
Overall gini	0.875*** (0.292)			
Ineq job		0.350 (0.242)		
Ineq age			−0.208 (0.211)	
Ineq gender				−0.550*** (0.176)
time trend	−0.123 (0.480)	−0.496*** (0.162)	0.169 (0.517)	0.357 (0.560)
Age	−0.564 (2.264)	2.313*** (0.573)	−0.408 (1.978)	−0.860 (2.370)
K	−0.016 (0.086)	0.022 (0.084)	0.010 (0.089)	0.027 (0.090)
time trend*BASILICATA	0.262 (0.237)	−0.010 (0.089)	0.189 (0.178)	0.282 (0.235)
time trend*CALABRIA	0.395 (0.241)	0.003 (0.071)	0.156 (0.151)	0.220 (0.207)
time trend*CAMPANIA	0.429** (0.179)	0.058 (0.100)	0.299** (0.118)	0.147 (0.163)
time trend*EMILIA	0.480 (0.332)	0.837*** (0.123)	0.457 (0.302)	0.242 (0.365)
time trend*FRIULI	0.335** (0.151)	0.456*** (0.080)	0.217 (0.180)	0.112 (0.195)
time trend*LAZIO	0.371*** (0.084)	0.485*** (0.046)	0.435*** (0.072)	0.188* (0.100)
time trend*LIGURIA	0.533*** (0.196)	0.676*** (0.092)	0.463** (0.198)	0.254 (0.236)
time trend*LOMBARDIA	0.426*** (0.150)	0.621*** (0.079)	0.487*** (0.140)	0.319** (0.157)
time trend*MARCHE	0.262* (0.155)	0.406*** (0.066)	0.250* (0.145)	0.144 (0.169)

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	<i>Dependent variable: Percentage Abstention</i>			
	(1)	(2)	(3)	(4)
time trend*MOLISE	-0.287*** (0.092)	-0.405*** (0.107)	-0.403*** (0.088)	-0.514*** (0.100)
time trend*PIEMONTE	0.488*** (0.142)	0.617*** (0.080)	0.406** (0.162)	0.263 (0.177)
time trend*PUGLIA	0.520** (0.253)	0.034 (0.093)	0.343* (0.177)	0.278 (0.244)
time trend*SARDEGNA	0.794** (0.372)	0.104 (0.120)	0.546** (0.253)	0.510 (0.340)
time trend*SICILIA	0.493*** (0.148)	0.122** (0.060)	0.205*** (0.073)	0.231** (0.114)
time trend*TOSCANA	0.442** (0.223)	0.641*** (0.095)	0.420** (0.211)	0.206 (0.257)
time trend*TRENTINO	0.367*** (0.115)	0.356*** (0.095)	0.148 (0.185)	0.123 (0.154)
time trend*UMBRIA	0.496*** (0.182)	0.540*** (0.101)	0.314 (0.204)	0.110 (0.231)
time trend*VALLE D'AOSTA	0.805*** (0.125)	0.809*** (0.128)	0.695*** (0.182)	0.577*** (0.196)
time trend*VENETO	0.395*** (0.089)	0.431*** (0.069)	0.307*** (0.110)	0.212* (0.110)

Note: *p<0.1; **p<0.05; ***p<0.01