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Employee-Owned Firms and the Careers of Young Workers

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Employee-Owned Firms and the Careers of Young Workers

Abstract

Using detailed administrative data from Spain, we characterize how a first work experience in an employee-owned firm (EOF) versus a conventional firm can affect workers' careers. We find that workers' exposure to EOFs at the time of entry reduces daily wages by 8% over the first 15 years in the labor market. The wage penalty appears to be driven by differences in job mobility and wage returns to experience rather than by non-random selection. We show that workers who had their first job in EOFs have a strong attachment to this organizational model and are less likely to experience both voluntary and involuntary job separations over their careers, with quit and layoff rates 8% and 4% lower, respectively. In addition, we quantify lower wage returns to experience in EOFs, although there are no differences in subsequent career progression in terms of promotions. Taken together, the analysis suggests the existence of other job amenities offered by EOFs that may compensate for flatter wage profiles.

JEL-Codes: J310, J500, J620.

Keywords: employee-owned firms, careers, wages, job mobility.

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1 Introduction

Growing evidence of firms' power to set wages has renewed interest in institutional arrangements aimed at strengthening the position of workers in the labor market (Ashenfelter et al., 2022; Card, 2022). While certain labor institutions, such as unions, collective bargaining, or minority workers' representation on company boards, have been extensively studied, important gaps remain in the understanding of how more extensive forms of worker participation affect individual and firm outcomes.

Employee-owned firms (EOFs) represent a limiting case of such arrangements, as the workforce exerts control over corporate decisions and shares profits. EOFs are often praised for their potential to benefit workers and the economy as a whole (Dow, 2018; Kruse, 2022). Numerous studies have documented how EOFs contribute to protecting workers against adverse shocks (Pencavel et al., 2006; Burdin and Dean, 2009; Kurtulus and Kruse, 2018; Garcia-Louzao, 2021) and, in certain environments, raise productivity (Fakhfakh et al., 2012; Montero, 2022; Young-Hyman et al., 2022) and reduce income inequality (Burdin, 2016; Magne, 2017). However, little is known about the dynamic consequences of EOFs for workers, despite growing evidence that the type of firm at labor market entry matters for future career paths (von Wachter and Bender, 2006; Deutscher et al., 2020; García-Trujillo et al., 2023; Arellano-Bover, 2024).

Drawing on rich Spanish administrative data, this paper sheds light on how the ownership structure of the first employer shapes labor market trajectories. Our study focuses on a particular type of EOF: worker cooperatives, which are enterprises in which worker-members have ultimate control rights (Dow, 2003). In practice, cooperatives combine both employee financial participation and voice, i.e., worker-members own the company, share net profits, and control strategic management decisions on a "one-person, one-vote" basis, regardless of their capital contribution. Hence, in our setting, we compare two types of firm ownership regimes: employee-owned enterprises organized as cooperatives and conventional investor-controlled firms.

The Spanish context offers a unique framework for investigating how alternative ownership and organizational arrangements of firms can shape the careers of young workers. First, the dataset is particularly suitable for our purposes as it allows us to follow workers' trajectories since their entry into the labor market and to identify, along with other demographic and firm characteristics, the legal ownership form of their first employer unambiguously. Importantly, we can control for crucial determi-

nants of wage growth, such as education and the skill content implicit in individuals' professional classifications, information that is usually missing in previous studies on cooperatives using administrative data. Second, Spain is an international case study for the historical development of its cooperative sector, including the famous example of the Mondragon Cooperative group in the Basque Country. Third, the Spanish government has systematically promoted the cooperative sector or, more generally, the Social Economy, as highlighted by the 2011 Social Economy Law and the 2017-2020 Social Economy Strategy. The Social Economy accounts for around 10% of GDP and 12.5% of employment in Spain, and cooperatives are the main business model, representing around 45% of these organizations. Finally, during the period of our analysis, 1985-2018, the country experienced significant economic growth, as well as a large variation in macroeconomic conditions, which allows us to isolate the effect of the first job experience from the national economic developments.

Our empirical analysis consists of comparing the early careers of Spanish workers whose first labor market experience after graduation was in a worker cooperative with those who started in a conventional firm. We find that the cooperative status of an individual's first employer is associated with a wage penalty of 8% over the first 15 years of their career. This negative effect persists even after controlling for province-specific unemployment rates at entry and is not explained by individual characteristics such as gender or education, nor by characteristics of the first job. Importantly, we implement an IV strategy based on the proportion of peers starting their careers in a cooperative and find a penalty of 14%, suggesting that the penalty is not simply due to (negative) selection on unobservables.

To complement the wage effects, we characterize the role of the ownership of the first job in the mobility patterns of workers during their first 15 years in the labor market. Our analysis reveals that workers whose first job after entering the labor market was in a cooperative are significantly less likely to voluntarily leave their first job and their future employer, i.e., a quit probability 8% lower. Similarly, these individuals are almost 4% less likely to be fired from their current firm compared to workers who started their careers in a conventional firm. Interestingly, we document that both wage effects and mobility patterns are entirely driven by workers with an above-median incidence of cooperative employment. This suggests that workers in cooperatives are willing to pay for the non-pecuniary job attributes offered by these organizations.

Finally, since time in the cooperative sector plays a crucial role in the career profiles of workers whose first job was in a cooperative, we quantify the wage returns to employer-specific accumulated experience. Our results indicate that wage returns to cooperative-employer experience are lower compared to the experience accumulated in conventional firms, which could be interpreted as lower human capital accumulation. However, the dynamics of promotions along the professional category ladder reveal that greater cooperative-employer experience does not restrict access to top-level professional positions, which is usually considered an indicator of career success.

The paper contributes to several strands of the literature. We add to an extensive literature that documents the long-lasting effect of initial labor market conditions on workers' careers. Most of these studies have focused on the role of macroeconomic conditions at labor market entry (see von Wachter, 2020, for a literature review). Using similar administrative data on Spanish workers, Arellano-Bover (2024) shows that the type of the first employer may also matter: young workers whose first job was in a large firm have better labor market outcomes than workers who started out in smaller firms. We add to this line of work by investigating the role played by another organizational dimension, namely the ownership form of the first employee. We show that even controlling for differences in firm size, individuals whose first job was in a cooperative have flatter wage profiles but more (voluntary) early career stability.

Our analysis also contributes to the understanding of non-conventional organizational arrangements, such as cooperative firms and employee ownership, a relatively under-researched topic in organization and labor studies (Gibbons and Roberts, 2015). A number of papers examine pay differentials between individuals employed in worker cooperatives and conventional firms (Blasi et al., 1996; Pencavel et al., 2006; Kruse et al., 2010; Clemente et al., 2012; Magne, 2017; Burdin, 2016; Garcia-Louzao, 2021). All these studies focus on contemporaneous worker-firm relationships. By contrast, our paper adds a dynamic perspective to the analysis of wage differentials between cooperatives and conventional firms by investigating the careers of young workers. Our results suggest that young workers initially employed in a cooperative exhibit a strong attachment to the cooperative sector and are less likely to quit their jobs, plausibly revealing their job satisfaction.

The paper also connects with recent work on the career consequences of past job experiences in heterogeneous firms (Gregory, 2020; Sorenson et al., 2021; Di Addario

et al., 2023; García-Trujillo et al., 2023; Arellano-Bover, 2024; Arellano-Bover and Saltiel, 2024). We add to this line of work by analyzing the impact on labor market trajectories of an under-explored dimension: the ownership structure of the firm. Our results indicate that heterogeneous ownership regimes may also have a first-order effect on the careers of young workers.

Finally, our analysis relates to the recent literature on co-determination and worker voice (Kim et al., 2018; Jäger et al., 2021, 2022). The evidence suggests that these institutions have no or small positive wage effects, no effect on voluntary separations, and seem to reduce involuntary separations. These studies focus on arrangements that give workers limited power, such as minority or quasi-parity representation, whereas our paper analyses the limiting case of such mechanisms: labor-managed firms, where worker-members exert control over firm management and ownership.

The remainder of the paper is organized as follows. In Section 2, we describe the institutional set-up governing the operation of worker cooperatives in Spain. In Section 3, we present our main data source, while Section 4 introduces the empirical approach. In Section 5, we discuss our main findings. Section 6 concludes.

2 Institutional background

Employee-owned firms. Two types of employee-owned firms exist in Spain: cooperatives (*sociedades cooperativas*) and labor societies (*sociedades laborales*).¹ Cooperatives are enterprises that meet the international principles of employee-owned enterprises: “an autonomous association of persons united voluntarily to meet their common economic, social, and cultural needs and aspirations through a jointly-owned and democratically-controlled enterprise”. They are organized according to cooperative principles: voluntary and open membership; democratic member control (one member, one vote); member economic participation; autonomy and independence; education, training, and information; cooperation among cooperatives; and concern for the community (Levin, 2002). Labor societies are subject to exactly the same conditions as conventional, capital-owned enterprises in terms of capital requirements, taxes, and the transmission of property rights. The key difference lies in the ownership schemes, as in labor societies, at least 51% of the capital must be held by working partners, and there are limitations on hiring non-partner workers. Given the simi-

¹See Online Appendix A for a more detailed comparison between cooperatives and labor societies.

larities between labor societies and conventional firms, in our analysis, we focus on cooperatives as employee-owned enterprises and leave labor societies as part of traditional capital-owned organizations, as is usually done in official statistics.²

Cooperatives. In order to create a cooperative, two fundamental requirements must be met: (i) the company must have at least three working partners, and (ii) each partner is entitled to only one vote in the governing bodies, regardless of capital contributions.³ It is important to note that cooperative partners can choose their Social Security contribution regime: self-employed or General Regime.⁴ In either of these two regimes, the partners decide on the labor income they declare as Social Security contributions. In practice, the difference between regimes is only related to future social benefits, as partners tend to declare the legal minimum contribution.

Cooperatives are allowed to hire external labor under the same conditions as mainstream capitalist enterprises.⁵ Wage-earners in cooperatives are thus covered by the same Labor Code and collective agreements that apply to workers in conventional enterprises. However, labor legislation restricts the number of hours worked by salaried employees to a maximum of 30% of the total hours worked in a year by partners but does not set any limit on the number of salaried employees.

Finally, cooperatives must allocate at least 20% of their net revenues to a reserve fund and 5% to an education and training fund. Profits paid into these funds are exempt from corporate income tax.⁶ Net revenues in excess of the amount contributed to the funds may be distributed to the partners in proportion to their work, and dividends may be paid on the capital contributed within the limits imposed by the law.

3 Data

Social Security records. The main data source is the Spanish Continuous Sample of Employment Histories (*Muestra Continua de Vidas Laborales* or MCVL), an admin-

²In the rest of the paper, we will use the terms employee-owned firms (EOFs) and cooperatives interchangeably.

³The incorporation of new partners requires the approval of the organization's governing bodies, and new members must subscribe to the minimum capital required and, in some cases, pay an admission fee. If a member decides to leave, the value of the capital contributed is reimbursed.

⁴Recent data from aggregate statistics indicate that about one-third of cooperative partners contribute to the self-employed regime.

⁵Aggregate figures indicate that 20-30% of cooperatives' employment corresponds to wage-earners.

⁶The corporate tax rate for cooperatives is 20%, compared to 25% for conventional corporations.

istrative dataset collected annually by the Spanish Social Security administration and linked to Tax Records from 2005 up to 2018.⁷ The MCVL is a representative 4% random sample of individuals who had any relationship with the Social Security system at any time in the reference year.⁸ The MCVL has a longitudinal design since an individual present in a year who subsequently remains registered with the Social Security administration stays as a sample member.⁹

For each member of the sample, the MCVL includes information on all Social Security relationships from the date of first employment or from 1967 for workers who were employed before then, allowing us to track individuals over time from their first job. For each of these relationships, we observe start and end dates along with part-time status, occupation category, workplace location and sector of activity, type of contract (with reliable information only as of 1997), and labor income.¹⁰ Demographic information such as age, gender, education, and nationality is also observed.

Importantly for the purposes of this analysis, the dataset provides information on the legal setting of the firm, which is determined by firm-specific tax IDs. The first element of these IDs stipulates the legal status of the company in terms of corporate taxation and makes it possible to unequivocally differentiate the employees of cooperatives from those of conventional companies.¹¹ Unfortunately, the data does not allow us to identify self-employed individuals who are partners of a cooperative as compared to other more standard forms of self-employment. Therefore, in our analysis, we focus on individuals whose first job was in the General Regime of Social Security and compare the careers of those individuals who entered the labor market in a cooperative versus a conventional firm. However, we do not limit their career to being exclusively in the General Regime but allow them to become self-employed at any time after their first job.

⁷The first version of the MCVL corresponds to 2004, but the information structure differs from that available for subsequent years.

⁸This includes employed and self-employed workers, recipients of unemployment benefits, and pension earners but excludes individuals registered only as medical care recipients or those with a different social assistance system (civil servants, such as the armed forces or the judicial power).

⁹Individuals who stop working remain in the sample while they receive any benefits from the Social Security, while they exit the sample when they die or leave the country permanently. Moreover, each wave adds individuals who enter the labor market for the first time.

¹⁰Labor income from standard wage-employment relationships is bottom- and top-coded. In our main analysis, we use the censored corrected earnings using a cell-by-cell Tobit model to impute censored observations (see Online Appendix C for more details on the correction method).

¹¹See Online Appendix B for the precise definition of the two employer categories and additional variables.

Analysis sample. We exploit all the 2005-2018 MCVL files to select individuals born in Spain and reconstruct their labor market histories, as we cannot track the complete history of those born abroad. For this group of people, we follow Arellano-Bover (2024) and rely on information on educational attainment to calculate each individual's expected graduation year in order to predict when a worker *should* start her career, thus addressing selective graduation decisions.¹² We define education-specific graduation years in terms of the usual graduation age: when high school dropouts turn 16, when high school graduates turn 18, and when college graduates turn 23. Using the predicted graduation year, we define the first job as the first six months after the year of graduation when individuals worked for more than 100 days. We then classify workers according to the ownership structure of the first employer to differentiate between those whose first job was in a cooperative versus a conventional firm.¹³ Finally, to avoid the inclusion of individuals whose first employment is likely to have taken place outside Spain, in the informal sector, or occurred abnormally late, we eliminate those workers whose first employment is observed more than five years after their predicted year of graduation.

From this sample, we select cohorts of graduates between 1984 and 2003 to ensure that we follow each cohort for 15 years and have reliable information on earnings from the first year after graduation.¹⁴ Next, we construct a monthly panel of individuals to study the career effects of having a first job in a cooperative compared to a conventional business form. Our final sample consists of 214,024 individuals observed in a total of 24,659,247 employment (worker-month-year) observations between January 1985 and December 2018. Table A1 in the Appendix reports summary statistics.¹⁵

4 Econometric model

To investigate the impact of entering the labor market in a worker cooperative relative to a conventional firm, we closely follow the literature on the effects of aggregate conditions at the time of labor market entry on workers' careers (e.g., Kahn, 2010; Ore-

¹²We use this approach to define an exogenous moment for labor market entry since we only observe workers from the first relationship with Social Security, which is ultimately an endogenous decision.

¹³If during the first six months, a worker has more than one job, we consider as the main employer the one in which she worked most of the time.

¹⁴Information on labor income prior to 1985 exhibits several missing values and inconsistencies and, hence, is not reliable.

¹⁵Online Appendix D offers further descriptive evidence.

opoulos et al., 2012). Specifically, we estimate models of the following form

$$y_{icpt} = \alpha + \beta Coop_{it_0} + \theta_p + \lambda_c + \delta_e + \gamma_t + X_{it_0}\Omega + \epsilon_{icpt} \quad (1)$$

where y_{icpt} refers to the (log) real daily wage in month-year t of individual i categorized by graduation cohort c and province of birth p . $Coop_{it_0}$ is the main independent variable that identifies individuals whose first job after labor market entry, t_0 , was in a cooperative. Thus, β quantifies the effect of the first labor market experience in a cooperative on the average daily wage over the first 15 years on the labor market.

To purge the coefficient of interest from confounding factors, we include unrestricted fixed effects for graduation cohort (λ_c) and the province of birth (θ_p), as well as years of potential experience (δ_e) and current calendar time (γ_t).¹⁶ This set of fixed effects is intended to account for unobserved heterogeneity related to non-linear experience profiles, different economic conditions at the current time as well as the time of graduation and among cohorts, respectively. To the extent that province-cohort specific variations in the likelihood of having a first work experience in a worker cooperative are uncorrelated with entrants' traits that are not loaded into our set of fixed effects, β would produce the average wage effect of having a first job in a cooperative relative to a conventional firm.

Despite the fact that we take into account several potential sources of heterogeneity via the broad set of fixed effects, there could still be other determinants of wage profiles that are arguably correlated with the ownership of the first employer (e.g., differences in risk aversion). To mitigate these concerns, we include three different sets of entry-level controls, which are summarized in X_{it_0} .¹⁷ To account for predetermined differences in earnings potential, we add indicators for gender and educational attainment. In addition, we include a cubic polynomial of the provincial unemployment rate at the time of entry to take into account local labor market conditions. Similarly, we control for the characteristics of the first job (skill level, full-time status, sector, and

¹⁶Recall that cohort, time, and experience effects cannot be identified separately in our framework without any additional assumptions. Given our interest in the effect of having a first job in a cooperative, we adopt a modeling strategy similar to that of Oreopoulos et al. (2012) and simply estimate $C - 2$ cohort effects.

¹⁷We do not include contemporaneous controls to avoid a *bad* controls problem, i.e., conditioning on post-treatment outcomes that ultimately affect current earnings (Angrist and Pischke, 2009; Cinelli et al., 2022). Therefore, we allow the indicator variable of having a first work experience in a cooperative to capture all the (potentially) different career paths of workers whose first employer was a cooperative relative to a conventional firm.

firm age) aimed at isolating the effect of ownership from other factors that are plausibly accounted for by the observed characteristics of the first labor relationship.¹⁸

Finally, ϵ_{icpt} represents the error term, and we allow for group-specific error components at the graduation cohort \times province birth level by means of clustered standard errors (Abadie et al., 2022).

5 Career effects of first employment in a cooperative

In this section, we characterize the first 15 years of the careers of workers who started in a cooperative relative to those whose first job was in a conventional enterprise.

Wage trajectories. In Table 1, we present the benchmark results with respect to daily wages. Column (1) reports estimates from a regression that only includes the basic set of controls (fixed effects for time, potential experience, graduation cohort, and province of birth). In Columns (2) to (4), we sequentially add controls to assess changes in the coefficient and the role of the observed factors on it. Column (2) adds gender and education controls, whereas Column (3) accounts for the initial local labor market conditions on workers' outcomes and the counter-cyclical role of workers' cooperatives in terms of job creation and employment stabilization by means of a cubic polynomial of the province-specific unemployment rate at the time of labor market entry. In our preferred specification, reported in Column (4), we add controls for first-employer and job attributes, such as skill level, full-time status, sector, and firm age.

[Table 1 around here]

Our results in Column (4) indicate that young workers' exposure to worker cooperatives at the time of labor market entry reduces earnings by about 8% during the first 15 years in the labor market. Given the higher incidence of workers with college education and in high-skilled jobs in cooperatives (see Appendix Table A1), the pay gap increases when controlling for demographics and characteristics of the first job with respect to the basic set of controls in Column (1). Interestingly, the penalty appears to be higher for men, as well as for those with a college degree (see Table A2 in the

¹⁸For example, workers may have a preference for greater job stability and therefore search for jobs in less volatile sectors.

Appendix). A possible explanation for a higher wage penalty for this group of workers, who tend to be those with wage premiums, could be related to concerns about inequality within cooperatives.¹⁹ In addition, in Figure 1, we document the dynamics of the wage gap by plotting the coefficients associated with the interaction of the cooperative dummy variable with indicators for each year of labor market experience from an extended version of Equation (1). The results show that wages are not significantly different at the beginning of a worker’s career but increase over time, becoming noticeable after two years in the labor market, peaking after eight years of labor market experience, and remaining stable or declining slightly thereafter.

[Figure 1 around here]

In the Online Appendix E, we perform several sensitivity tests that validate our results. We show that our findings hold under alternative definitions of the first job, the definition of the first job, the type of employment relationship considered, the size of the first employer, and our definition of earnings or EOFs. Importantly, for a subset of workers for whom we observe the size of the first employer, we find that while the size of the first employer matters for the dynamics of wages after entry, in line with the findings of [Arellano-Bover \(2024\)](#), it does not explain the differences between workers who start their careers in cooperatives compared to conventional firms. Moreover, using a set of cohorts that we can follow for 30 years, we show that the estimated gap hardly changes when we extend the time horizon over which we follow workers, suggesting that the 15 years is the critical period for life-cycle wage growth.

A fundamental concern with our OLS estimates is that the first employer type may be correlated with unobserved worker characteristics (motivation, ability), which could explain the wage profiles due to the selection of certain types of individuals into cooperatives at the moment of labor market entry. To assess the role of non-random selection, we implement an IV strategy that adapts the leave-one-out approach developed by [Arellano-Bover \(2024\)](#) in the context of the career premium associated with first-employer size. In his context, hiring shocks driven by the decisions of large firms generate variation in the regional composition of labor demand, which affects the allocation of workers to small and large firms. This is closely related to the power law

¹⁹For example, to the extent that individuals with a first job in a worker cooperative develop a career in the cooperative sector, this heterogeneous pattern is consistent with existing evidence on wage compression in EOFs ([Kremer, 1997](#); [Abramitzky, 2008](#); [Burdin, 2016](#)).

distribution of firm size, and thus this type of instrument may not work for other firm attributes (Arellano-Bover, 2024, p.564). With this caveat in mind, in our setting, we instrument the first job in a cooperative for each worker i with the share of her peers who got their first job in a cooperative, i.e., individuals who have the same education level as i , started their first job in i 's province of birth, and did so in i 's predicted year of graduation.

Column (5) of Table 1 shows that the penalty still emerges under the IV strategy, but the magnitude becomes almost 2 times larger: matching with a cooperative at labor market entry leads to 14% lower wages over the first 15 years of their careers. We complement the IV results with the strategy proposed by Oster (2019) to characterize the degree of selection (see Online Appendix F). The approach is based on the omitted variable bias formula and allows to obtain bounds on the degree of selection bias under the assumption that selection on unobservables is proportional to selection on observed characteristics. Note that this assumption seems to hold in our context since accounting for selection on unobservables using the IV moves the coefficient of interest in the same direction as when controlling for workers' observed characteristics, as suggested by the comparison between Columns (1) and (4) of Table 1. This exercise points to a maximum wage penalty of 17%. In addition, the relative magnitude of selection on unobservables would have to be 11 times larger than that on observables and of the opposite sign for the wage penalty to disappear.

Taken together, the IV strategy and Oster's bounds suggest that the wage penalty is downward biased, which is consistent with other studies that find larger IV estimates when looking at the consequences of labor market entry i.e., initial employer size (Arellano-Bover, 2024) or economic conditions at labor market entry (Kahn, 2010). This suggests that negative selection is unlikely to explain the wage penalty. In our context, the IV-OLS gap could be driven by individuals with greater earnings potential joining a cooperative firm when a large fraction of their peers do so.²⁰ Thus, given the egalitarian pay policies that characterize cooperatives (Kremer, 1997; Abramitzky, 2008; Burdin, 2016), these individuals are relatively more penalized by working in these firms. This is consistent with the wage penalty being zero at entry and building

²⁰For example, the fact that other individuals in the reference group are joining a cooperative may lend greater legitimacy to this occupational choice and facilitate the diffusion of information and knowledge, thereby reducing the potential costs associated with joining a relatively rare organizational form (Hannan, 2005).

up only over time, as shown in Figure 1, and with the heterogeneous effects reported in Table A2 in the Appendix, which points to larger penalties for men or highly educated individuals.

Job mobility. A large literature identifies (voluntary) job mobility as a critical driver of young workers' wage progression due to the search for a better job (e.g., Topel and Ward, 1992; Light and McGarry, 1998; Keith and McWilliams, 1999; Davia, 2010).²¹ We then turn to investigate whether an initial work experience in a cooperative affects the average probability of separating from a given employer during the first 15 years in the labor market.

In Table 2, Columns (1) and (2), we report estimates from a linear probability model and a discrete duration model specification for the likelihood of separation, using the same right-hand side variables as in Equation (1). The results suggest that workers whose first work experience is in a cooperative are less likely to separate from their employer over the course of their career. In Columns (3)-(5), we investigate these differences further by estimating a competing risk duration model where we consider different types of separations (voluntary, involuntary, and other types). We find that starting a career in a cooperative reduces the probability of both employee-initiated and employer-initiated separations, but the effect is larger for the former. This suggests that the higher stability during the first 15 years in the labor market for these workers is driven by both labor supply and labor demand dimensions. On the one hand, the reduction in voluntary turnover could be interpreted as a preference-revealed indicator of job satisfaction with workplace amenities provided by cooperatives (Böckerman and Ilmakunnas, 2009; Cottini et al., 2011; Harju et al., 2021).²² On the other hand, the lower probability of involuntary separations may reflect the well-documented propensity of cooperative enterprises to avoid layoffs and to provide employment insurance for their employees (Pencavel et al., 2006; Burdin and Dean, 2009; Garcia-Louzao, 2021).²³

²¹Involuntary mobility (job displacement) tend to be followed by wage losses (Bertheau et al., 2022).

²²Common ownership of assets in cooperatives increases individuals' exit costs, serving as a lock-in device (Abramitzky, 2008).

²³In Table A3 in the Appendix, we show that having a first job experience in a cooperative significantly reduces both the probability of making job-to-job transitions (unemployment duration of less than a month) and experiencing spells of mid-length unemployment (6-11 months) over the first 15 years in the career. However, we find no differences for short unemployment spells (2-5 months) or the incidence of long-term unemployment (12 months or more).

[Table 2 around here]

The lower frequency of job separations among individuals who started their career in a cooperative compared to those who started in a conventional firm may explain the flatter wage profiles for the former group of workers, as they may be partly a consequence of foregone wage gains associated with (voluntary) job mobility.²⁴ However, less intense job mobility may reflect the transformative experience of working in a cooperative. In other words, individuals may develop preferences for certain (cooperative-specific) non-pecuniary job attributes that compensate for lower wages and lead them to stay in this sector (Rosen, 1986). This would be consistent with Figure A1 in the Appendix that shows that even after 15 years of labor market experience, individuals who started in a worker cooperative are still 20% more likely to remain employed in the cooperative sector.²⁵

Thus, if flatter wage profiles reflect certain job amenities specific to cooperatives, our results should be driven by individuals with a stronger attachment to the cooperative sector and not by those who, despite having a first job in a cooperative, spent most of their careers in conventional firms. Table 3 divides the indicator of workers initially matched with a cooperative by the below- and above-average incidence of cooperative employment during the first 15 years of their careers, based on the time spent in a cooperative relative to total working time. The results of this exercise indicate that the flatter wage profiles and lower mobility are entirely determined by workers who have spent a larger part of their careers in the cooperative sector.

[Table 3 around here]

Returns to experience. Flatter wage profiles and lower separation probabilities are driven by individuals who, having started their careers in a cooperative, remain in this type of organization most of the time. To gain a better understanding of this

²⁴In the Online Appendix D, Table D.1 shows the estimates of a similar model for the probability of separation from the first job, and only voluntary separations differ by type of first employer. However, this is unlikely to explain the flatter wage profiles, as we find that the wage penalty is around 6% even for individuals who spent less than a year with their first employer (Figure D.7).

²⁵This does not necessarily mean that all remaining workers leave the cooperative sector, as cooperative partners can register as self-employed even when they belong to a cooperative group. In Panel B of Figure A1, we document that the probability of becoming self-employed is increasing over the career path for individuals who had their first job at a worker cooperative. Although we cannot directly determine whether these individuals remain self-employed in the cooperative sector, the parallel evolution of both figures may be suggestive of such a pattern.

dimension behind the flatter wage profiles, we estimate a Mincer-type wage equation using as the main explanatory variables the *actual* experience accumulated in each type of firm since the first job up to 15 years after labor market entry.²⁶ In these regressions, we include as controls a wide set of contemporaneous characteristics, including the cooperative status of the current employer as well as worker fixed effects to account for time-invariant characteristics that may affect both wages and experience with different types of firms.

[Table 4 around here]

Our estimates in Table 4 indicate that cooperative-employer experience has lower returns than experience at conventional employers and that the differential decreases over time. Differential returns to experience across jobs or firms are typically used as a proxy to measure differences in human capital (e.g., Garcia-Louzao et al., 2023; Arellano-Bover and Saltiel, 2024; Arellano-Bover, 2024). According to this literature, the lower return to cooperative experience may suggest that workers in these organizations have lower human capital accumulation. This would be consistent with, for example, EOFs' pay policies that favor equality within firms may affect their ability to recruit and retain highly skilled workers (Kremer, 1997; Abramitzky, 2008; Burdin, 2016), which may reduce the opportunities for young workers to learn from talented managers and peers. In addition, young workers' expectations of more job security in a cooperative may hold back the acquisition of alternative skills, i.e., skills that are not useful in the current job but would be useful elsewhere (Kuhn and Sweetman, 1999). Another potential channel relates to skill specificity: skills developed in an EOF may be too specific, having limited value once workers switch to a conventional business, but also the routines and habits in those organizations may not fit into the conventional sector (Dokko et al., 2009; Wu and Paluck, 2020).²⁷

Alternatively, lower returns to experience can be consistent with statistical discrimination if conventional employers use past experience in a cooperative as a noisy signal of low productivity.²⁸ According to this interpretation, negative returns should vanish over time as employers learn the true productivity of workers (Altonji and Pierret,

²⁶In this exercise, we rely exclusively on wage-employment observations.

²⁷For example, individuals' past experience in a participatory workplace may contribute to developing critical attitudes towards authority, signaling to conventional employers a lack of discipline, thus undermining their chances of employment in the conventional sector.

²⁸In Online Appendix G, we compare the distribution of worker fixed effects obtained from augmented Mincer equations of the type in Column (2) of Table 4. The comparison shows that workers

2001). However, the slow catch-up process that we observe in the data (see Figure A2 in the Appendix) is difficult to reconcile with any plausible specification of the employer's learning process (Lange, 2007). To be precise, we estimate that it would take 10 years of tenure for the gap in the returns to experience to disappear (while the median tenure in the sample is 4 years).

[Table 5 around here]

Given that the dynamic of promotions is usually considered the result of a skill acquisition process and a strong indicator of career success (Gibbons and Waldman, 2006; Waldman, 2012; Frederiksen and Kato, 2018), in Table 5, we re-estimate the Mincer-equation using promotions as a dependent variable. Promotions are identified by analyzing within-individual changes in professional categories defined by both the level of education required for the specific job and the complexity of the task.²⁹ The results indicate that accumulated experience in cooperatives is not associated with slower career progression, as measured by promotions to more skilled job categories. Thus, the fact that we do not observe significant differences along this dimension suggests that workers in cooperatives might not necessarily acquire less valuable skills.³⁰

The lower return to experience in cooperatives compared to conventional firms may instead reflect differences in pay policies between the two types of organizations. Consistently with this interpretation, Columns (2) and (4) of Table 4 indicate that the gap in returns to experience is smaller when we control for the cooperative status of the current employer. Therefore, attributing heterogeneous returns to experience to differences in human capital may not be entirely appropriate when firms have very different objective functions.³¹

with the lowest (highest) ability, as measured by their fixed effects, are only slightly overrepresented (underrepresented) among those who started their careers in cooperatives.

²⁹See Online Appendix B for detailed definitions of the professional categories. Online Appendix D Figure D.9 shows raw correlations between the professional category in the first job and the one in the last observed job after 15 years in the labor market.

³⁰Promotions to high-level positions within the cooperative sector may be driven by political considerations rather than by meritocratic factors. However, as shown in Column (6) of Table 5, results remain unchanged when we control for the cooperative status of the current employer.

³¹Existing evidence suggests that most conventional, capital-owned firms are primarily concerned with profit maximization, while worker-managed firms are concerned with both income and employment or, more generally, the welfare of their members (e.g., Burdin and Dean, 2012).

6 Conclusions

This paper examines how the ownership structure of the first job can shape the careers of young workers. Our analysis shows that individuals who had their first job in an EOF have 8-14% lower wages over their first 15 years in the labor market relative to similar individuals who were initially employed in conventional, capital-owned firms. We also document negative wage returns to accumulated experience in EOFs relative to experience in conventional firms, but no significant differences in the dynamics of promotions. Moreover, we find that initial experience in EOFs reduces both voluntary and involuntary mobility. In other words, workers who started their careers in a cooperative enjoy greater job stability during their first 15 years in the labor market. Thus, in contrast to studies that find small effects of institutions that give limited power to workers (Jäger et al., 2021), our results suggest that organizations with broader forms of worker participation in corporate governance can have large and persistent effects.

Altogether, our evidence is indicative of the existence of non-pecuniary job attributes offered by EOFs that could compensate for lower earnings. In this regard, recent estimates of willingness to pay for workplace amenities (e.g., job security or work environment) indicate that moving from a job with the lowest to a job with the highest amenities is equivalent to a 20-50% wage increase (Maestas et al., 2022; Dube et al., 2022; Sockin, 2022). Therefore, our results are consistent with these estimates, provided that the average job in a cooperative or the average job in a conventional firm does not correspond to either the lowest or the highest amenity in the economy.³²

Our findings have potential implications for policy initiatives to promote the role of EOFs in facilitating school-to-work transitions and improving employment conditions for young people. These initiatives typically focus on addressing informational barriers and cultural factors that are thought to prevent young workers from joining these organizations. Our analysis suggests that it is equally important to understand the long-term costs and benefits of working in such firms and to assess whether they provide sufficient incentives for individuals to join and remain in them.

Importantly, we do not have information on individuals' compensation packages, and in particular the incidence of profit sharing. If profit-sharing is more common in

³²Experimental evidence from Wiswall and Zafar (2017) on the willingness to pay for a specific job attribute, i.e., job security, indicates that the average subject is willing to give up 2.8% of annual earnings for a job with a one percentage point lower probability of dismissal.

EOFs, we may overestimate the earnings penalty. It would also be interesting to analyze differences in the provision of (on-the-job) training and other workplace amenities (e.g., child care, health, or pension plans). Thus, further research using firm-level data could provide a more in-depth analysis of the mechanisms behind our results.

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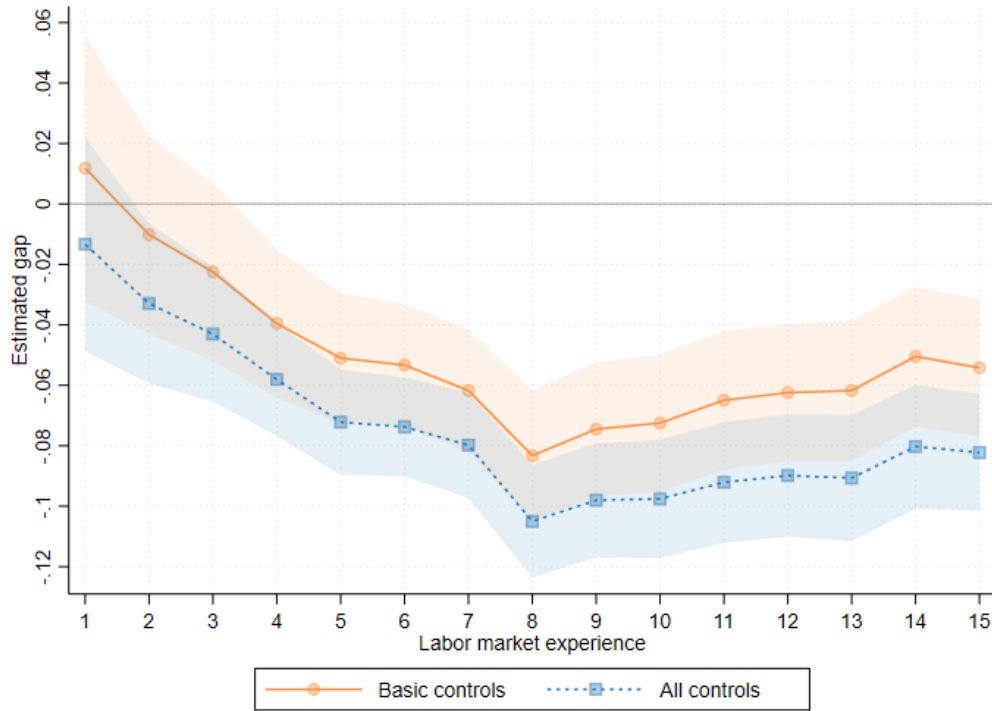
Figures and tables

Table 1: Impact of first job in a cooperative

	(1)	(2)	(3)	(4)	(5)
First job at cooperative	-0.056*** (0.009)	-0.071*** (0.007)	-0.071*** (0.007)	-0.080*** (0.007)	-0.1425*** (0.047)
Observations	24,659,247	24,659,247	24,659,247	24,659,247	24,659,247
F-stat					3969.73
Basic controls	Yes	Yes	Yes	Yes	Yes
Gender/Education	No	Yes	Yes	Yes	Yes
Unemployment rate at entry	No	No	Yes	Yes	Yes
Job-firm characteristics of first employer	No	No	No	Yes	Yes

Notes: The dependent variable is the log of daily labor income (traditional wage employment as well as income from self-employment and other forms of dependent employment) after the first job. All regressions include basic controls (calendar month-year, potential experience, graduation cohort, and province of birth). In column (2), we add controls for gender and education. In column (3), we control for the province-specific unemployment rate at the time of labor market entry (cubic polynomial). In column (4), we include additional controls for first-job-employer characteristics (skill level, full-time status, sector, firm age). Column (5) instruments the first job in a cooperative for each worker i with the share of her peers who got their first job in a cooperative, i.e., individuals who have the same education level as i , started their first job in i 's province of birth and did so in i 's predicted year of graduation. Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure 1: Experience-wage gap of first job in a cooperative



Notes: The figure shows the experience-specific wage differential for first job experience in a cooperative relative to a conventional firm, estimated from an augmented version of Equation 1 in which the cooperative dummy variable interacts with the indicators for each year of experience. Basic controls refer to calendar month-year, potential experience, graduation cohort, and province of birth. All controls include basic controls plus province-specific unemployment rate at the time of labor market entry (cubic polynomial) along with variables to account for gender, education, entry-level skill, full-time status, sector, and firm age differences. 95% confidence bands based on standard errors clustered at the level of the province of birth \times graduation year.

Table 2: Job mobility over the career

	OLS	Logit	Multinomial Logit		
			Voluntary	Involuntary	Other
First job at cooperative	-0.002*** (0.001)	-0.050*** (0.015)	-0.085*** (0.021)	-0.038** (0.017)	-0.073* (0.042)
Observations	24,659,247	24,659,247	24,659,247	24,659,247	24,659,247

Notes: The dependent variable is an indicator for the probability that a worker separates from her current employer. Column (1) estimates a linear probability of the model for the likelihood of separating. Column (2) estimates the separation probability by maximum likelihood using a Logit link for the hazard function. Columns (3) to (5) extend Column (2) to a competing risk setting (multinomial logit) using three types of separations: voluntary, involuntary, and other. Voluntary and involuntary movers stand for workers whose separation from their first job was initiated by the employee or the employer, respectively. Other types of separation is a residual category including, among others, sickness or parental leave, but mostly refers to undefined/administrative causes. All specifications include the same set of controls as Column (4) in Table 1. Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 3: Wages and mobility by incidence of cooperative employment

	(1) Wage effect	Mobility effect		
		(2) Voluntary	(3) Involuntary	(4) Other
First job at cooperative, low incidence	-0.009 (0.008)	0.161*** (0.026)	0.162*** (0.021)	0.087* (0.052)
First job at cooperative, high incidence	-0.152*** (0.011)	-0.310*** (0.037)	-0.365*** (0.029)	-0.361*** (0.066)
Observations	24,659,247	24,659,247	24,659,247	24,659,247

Notes: Column (1) reproduces Column (4) of the Table 1, while Columns (2) to (4) reproduce Columns (3) to (5) of the Table 2, but we divide workers whose first job was in a cooperative into low and high incidence of cooperative employment during the first 15 years in the labor market. Incidence of cooperative employers is defined as the ratio of time spent in a cooperative to total working time. Workers with low (high) incidence are those with time spent in a cooperative below the median (above the median) relative to total working time. The median incidence is 0.31. All specifications include the same set of controls as Column (4) in Table 1. Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 4: Returns to experience at cooperatives

	(1)	(2)	(3)	(4)
CoopExp	-57.333*** (8.317)	-34.264*** (8.615)	-55.386*** (8.393)	-33.759*** (8.642)
CoopExp \times Exp	0.011*** (0.002)	0.006*** (0.002)	0.007** (0.003)	0.003 (0.003)
CoopExp \times Tenure			0.005*** (0.002)	0.004** (0.002)
Exp	236.670*** (2.009)	236.437*** (2.009)	236.643*** (2.009)	236.426*** (2.009)
Exp ²	-0.024*** (0.000)	-0.024*** (0.000)	-0.024*** (0.000)	-0.024*** (0.000)
Tenure	38.131*** (1.100)	37.952*** (1.100)	38.154*** (1.100)	37.980*** (1.100)
Tenure ²	-0.007*** (0.000)	-0.007*** (0.000)	-0.007*** (0.000)	-0.007*** (0.000)
Observations	23,086,331	23,086,331	23,086,331	23,086,331
Current coop status	No	Yes	No	Yes

Notes: Dependent variable is log daily labor earnings from wage-employment. Experience and tenure are measured in days. Exp is overall *actual* experience. CoopExp is experience accumulated at cooperative firms. Tenure equals days worked for the current employer. Estimated coefficients and standard errors multiplied by 10^6 for readability. All specifications include fixed effects for workers, skill level, full-time, firm age, sector, location, and time. Columns (2) and (4) include a dummy variable equal to one if the current employer is a cooperative firm and zero otherwise. Standard errors clustered at the worker level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 5: Returns to experience at cooperatives: Professional mobility

	Promotions		Promotion, w/o demotions		High-level promotion	
	(1)	(2)	(3)	(4)	(5)	(6)
CoopExp	-0.983** (0.497)	-0.562 (0.533)	-0.722 (0.484)	-0.544 (0.503)	-0.233 (0.216)	-0.348 (0.235)
CoopExp × Exp	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)
CoopExp × Tenure	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	0.000 (0.000)	-0.000 (0.000)	-0.000 (0.000)
Observations	23,086,331	23,086,331	14,534,828	14,534,828	14,534,828	14,534,828
Current coop status	No	Yes	No	Yes	No	Yes

Notes: Dependent variable is an indicator variable equal to one if the individual moves up in the professional category ladder in $t + 1$ and zero otherwise. Columns (1) and (2) include all upward movements on the occupation ladder. Columns (3) and (5) restrict the sample to individuals who have never experienced a downward career move. Columns (4) and (6) define only movements to the highest level of occupational categories as promotions. Experience and tenure are measured in days. Exp is overall *actual* experience. CoopExp is experience accumulated at cooperative firms. Tenure equals days worked in the current employer. Estimated coefficients and standard errors multiplied by 10^6 for readability. All specifications include the same set of controls as Table 4 Column (4), referring to period t . Standard errors clustered at the worker level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Appendix

Table A1: Descriptive statistics by firm ownership of first job

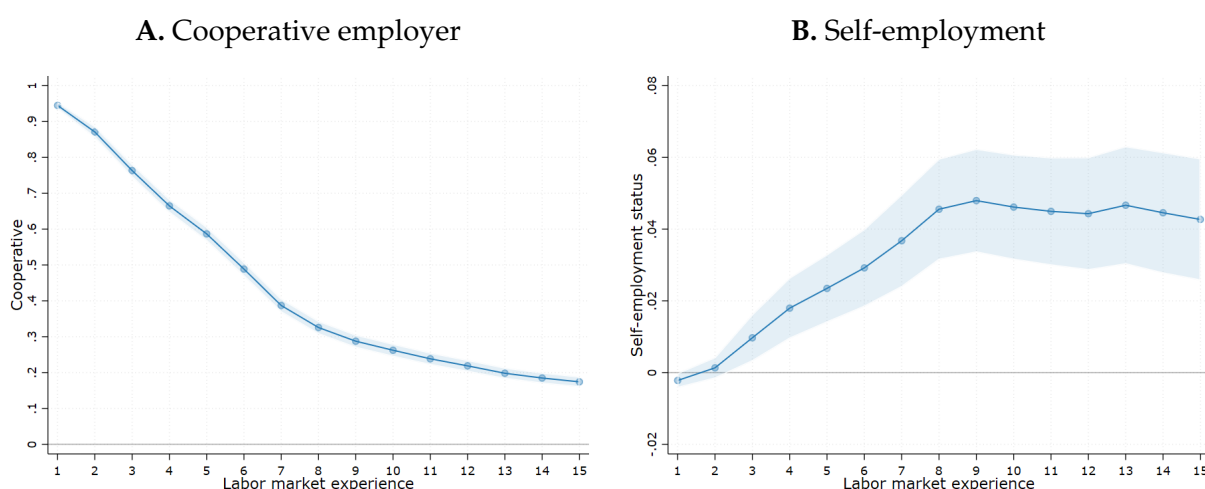
	Conventional firms	Cooperatives
Demographics		
Female	0.43	0.47
Age	20.83	21.03
High-school	0.30	0.27
College	0.21	0.25
First labor market experience		
Time to first job (yr)	1.74	1.72
First job in province of birth	0.83	0.85
Unemployment rate	0.18	0.19
Earnings first 6 months	5,557.70	5,732.15
Days worked first 6 months	163.82	163.80
No. employers first 6 months	1.21	1.19
Total earnings in first job	43,018.94	44,703.85
Total days worked in first job	843.65	909.51
First job in a new firm	0.18	0.15
Entry partner	-	0.09
Mid-skill occupation	0.30	0.27
High-skill occupation	0.08	0.12
Full-time job	0.77	0.74
Firm age	9.06	10.86
Manufacturing	0.21	0.31
Construction	0.13	0.10
Big city	0.43	0.29
Outcomes during first 15 years		
Wage-employment income	175,593.41	157,182.23
Self-employment income	6,133.74	10,650.69
Unemployment benefits	9,868.97	8,687.70
Wage-employment days	3,118.58	2,921.34
Self-employment days	191.06	314.76
Days registered as unemployed	218.22	196.69
No. employers	4.10	3.72
No. cooperatives	0.04	1.20
Always same firm type	0.40	0.16
Workers	209,627	4,397

Notes: Time to first job refers to the years between graduation year and the year of the first employment. The first job in a new firm stands for individuals who were hired within the year the firm was founded. Big city refers to metropolitan areas with over 1 million inhabitants (Madrid, Barcelona, Sevilla, and Valencia). Income is expressed in 2018 euros deflated using the Spanish consumer price index.

Table A2: Impact of first job in a cooperative: Heterogeneity by worker demographics

	(1) Men	(2) Women	(3) Non-College	(4) College
First job at cooperative	-0.097*** (0.009)	-0.058*** (0.009)	-0.073*** (0.008)	-0.103*** (0.014)
Observations	14,356,008	10,303,239	18,811,330	5,847,917

Notes: All specifications include the same set of controls as Column (4) in Table 1 when appropriate. Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure A1: Differential probability of being in a cooperative or self-employment by labor market experience and initial employer type

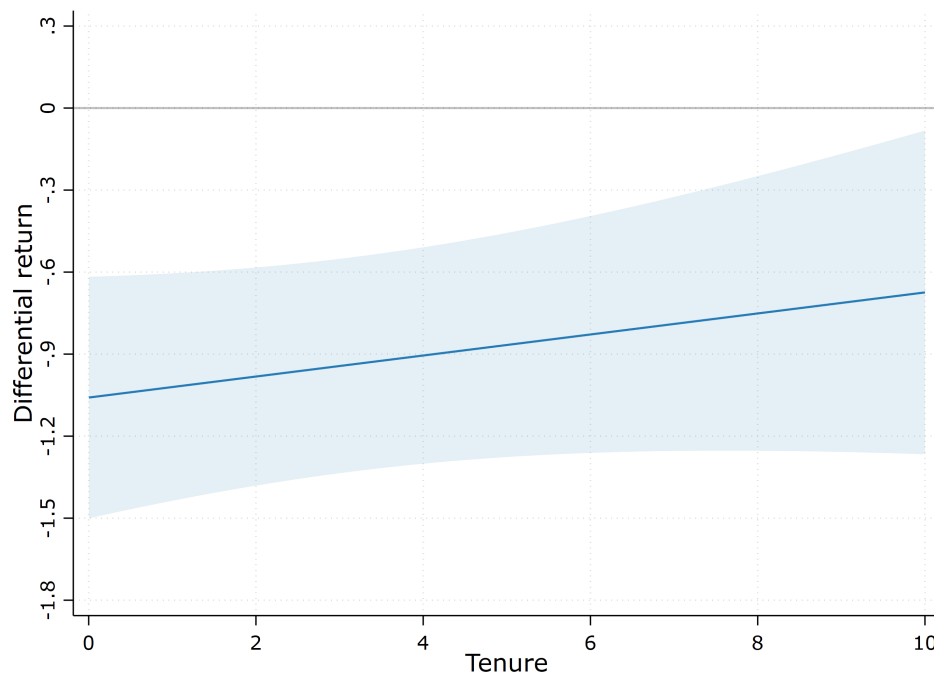
Notes: The figure shows the impact of having a first job in a cooperative relative to a conventional enterprise on the probability of working for a cooperative (Panel A) or being self-employed (Panel B) by labor market experience. Results are obtained from separated linear probability models where the first job in a cooperative dummy is interacted with the indicators for each year of experience. Both linear probability models include the same set of controls as Column (4) in Table 1. 95% confidence bands based on standard errors clustered at the level of the province of birth \times graduation year.

Table A3: Unemployment incidence over the career

	≤ 1 month	2 to 5 months	6 to 11 months	≥ 12 months
First job at cooperative	-0.093*** (0.018)	-0.015 (0.024)	-0.087*** (0.029)	0.018 (0.014)
Observations	24,659,247	24,659,247	24,659,247	24,659,247

Notes: All panels estimate competing risk (multinomial logit) models using four separation types. Panel A uses separation types based on the length of the unemployment spell, i.e., the time between leaving the current job and starting a new job. Panel B divides the risks into job-to-job (up to a month to change job) transitions and employment-to-unemployment (more than a month to change job) transitions, with the latter further disaggregated to reflect whether the separation was voluntary, involuntary, or some other type as in Table 2. Panel C classifies separations into short unemployment spells (less than 6 months) and long unemployment spells (6 months or more), and the latter into voluntary, involuntary, and other types of separations. Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure A2: Catch-up rate: Wage return to one year of experience in a cooperative relative to a conventional firm, by current employer tenure



Notes: The figure shows the differential return (in percent) to one additional year of experience acquired in a cooperative (CoopExp = 360 days) relative to a conventional enterprise by current employer tenure. The differential return is calculated using point estimates, CoopExp, Coop×Exp, and CoopExp×Tenure, from Table 4 Column (4). Experience is fixed at its median value of the actual experience distribution in the baseline sample, 1465 days (~4 years). Tenure is measured in days but expressed in years for readability. 95% confidence bands are computed using the delta method.

Online Appendix

The online appendix is organized as follows. First, in Appendix A, we discuss differences in the legal framework governing the two types of employee-owned firms (cooperatives and labor societies) in the Spanish context. Second, in Appendix B we provide a definition of our main variables. Third, in Appendix C we explain the correction made to censored earnings. Fourth, Appendix D reports supplementary tables and figures that are referred to in the main text. Fifth, in Appendix E, we report results from a series of robustness tests to validate our main results. Sixth, Appendix F discussed the calculation of the Oster's bounds. Finally, in Appendix G, we present our analysis of the distribution of worker unobserved heterogeneity by firm ownership of the first job.

A Cooperatives versus labor societies

The Spanish law distinguishes two types of employee-owned firms: cooperatives and labor societies. These business forms share the common characteristic of being majority-owned by their workers, i.e., they control more than 51% of the capital. However, there are key elements that differentiate them:

- i. *Restrictions on hiring non-partner workers.* In cooperatives, wage-earners cannot work more than 30% of the total hours worked by cooperative members during a year, while for labor societies, this limit is set at 49%.
- ii. *Minimum capital requirement.* labor societies are regulated by the same legislation as conventional corporations: firms must deposit a minimum level of capital to set up the organization, 3,000 euros in case of limited liability labor societies, and 60,000 euros for limited liability companies that are publicly traded. For cooperative businesses, there is no legal capital minimum requirement, but partners must stipulate the amount of the initial capital in the company bylaws.
- iii. *Capital ownership.* In cooperatives, the owners of the firm are the working partners, but the firm bylaws may establish the possibility of accepting collaborative partners whose total contributions cannot exceed 45% of the social capital. In the case of labor societies, there are two types of partners: working partners who own at least 51% of the capital and capitalist partners who can have a share in the firm that cannot exceed 49% of the capital.
- iv. *Voting power of partners.* The voting power in labor societies depends on the amount of capital provided by each member, while in cooperatives, each member has the same voting power. If there are collaborative partners in the cooperative, they cannot represent more than 30% of the votes in the company governing bodies.
- v. *Corporate taxation.* labor societies are taxed as any other type of for-profit business in Spain (25% of the profits). In the case of cooperatives, the corporate tax rate is lower (20%).³³

³³Prior to 2016, the corporate income tax rate for conventional companies was 30%. In the case of cooperatives, the tax rate was not changed, but a distinction was made between the tax rate for cooperative and non-cooperative profits. Cooperative profits were taxed at 20%, while non-cooperative profits—those obtained from business activities that do not correspond to the purpose of the organization—were taxed at the general rate (30%).

B Variables definition

Gender. Obtained from the Spanish Residence registry. We select this information from the most recent wave, and if there is any inconsistency, we choose the mode over the waves in which it is available.

Birth date. Obtained from personal files coming from the Spanish Residents registry. We select this information from the most recent wave, and if there is any inconsistency, we choose the most common value over the waves for which it is available.

Nationality. Obtained from Spanish Residents registry. The variable reports the link between the individual and Spain in terms of legal rights and duties. This variable allows for distinguishing between individuals with Spanish nationality (N00 code) and other worldwide nationalities.

Education. Retrieved from the Spanish Residents registry up to 2009, and from 2009 thereafter, the Ministry of Education directly reports individuals' educational attainment to the National Statistical Office, and this information is used to update the corresponding records in the Residence registry. Therefore, educational attainment is imputed backward whenever it is possible, i.e., when a worker is observed in the MCVL post-2009. In the imputation, we assigned 25 years as the minimum age to recover values related to university education.³⁴

Labor income. Refers to Social Security contribution bases adding up both for traditional wage-employment as well as income from self-employment activities and other forms of dependent employment. Wage-employment income captures gross monthly labor earnings plus one-twelfth of year bonuses and is bottom and top-coded. The minimum and maximum caps vary according to the Social Security regime and contribution group, and they are adjusted each year according to the evolution of the minimum wage and inflation rate. In our main analysis, we use censored earnings due to the low incidence in our sample, as it is mainly composed of young workers. However, we test the sensitivity of our results using censored corrected earnings, where we correct the upper tail of the wage distribution by fitting cell-by-cell Tobit models to

³⁴The age threshold is the average graduation age for a Bachelor's degree in Spain: <https://www.oecd.org/education/education-at-a-glance-19991487.htm>

(log) daily wages.³⁵ Self-employment income corresponds to the Social Security contribution declared by individuals carrying out their own activity and is based on their expected stream of labor income for a given period, typically quarters. We express labor income in 2018 euros, deflated using the Spanish consumer price index.

Skill category. Refers to Social Security contribution groups. These groups indicate a level in a ranking determined by the worker's contribution to the Social Security system, which is determined by both the level of education required for the specific job and the complexity of the task. The MCVL contains 10 different contribution groups that are aggregated based on similarities in skill requirements. High-Skill: Group 1 (engineers, college, senior managers—in Spanish *ingenieros, licenciados y alta direccion*), Group 2 (technicians—in *ingenieros tecnicos, peritos y ayudantes*), and Group 3 (administrative managers—in *jefes administrativos y de taller*). Medium-Skill: Group 4 (assistants—in *ayudantes no titulados*) and Group 5-7 (administrative workers—in *oficiales administrativos* (5), *subalternos* (6) and *auxiliares administrativos* (7)). Low-Skill: Group 8-10: (manual workers—in *oficiales de primera y segunda* (8), *oficiales de tercera y especialistas* (9) y *mayores de 18 años no cualificados* (10)).

Reason for termination. Declared by the employer to the Social Security Administration. This variable is relevant for determining entitlement to severance pay and unemployment benefits. Using this information, we create three broad categories based on the following codes: code 51 refers to voluntary resignations or separations, 52, 54, 69, 77, 91, 92, 93, and 94 to dismissals or involuntary separations; and the remaining codes are considered other reasons for dismissal, including among others sickness or parental leave, but mostly referring to undefined/administrative causes.

Plant. A plant is defined by its Social Security contribution account (*codigo de cuenta de cotizacion*). Each firm is mandated to have as many accounts as regimes, provinces, and relation types with which it operates. According to the Social Security Administration, around 85% of the firms are single-unit organizations, i.e. they have just one contribution account per firm. Each firm has one account for each treble province-Social Security regime-type of the employment relation. Thus, the Social Security Ad-

³⁵See Appendix C for a detailed discussion of the correction method and a comparison between the original and corrected wage distributions.

ministration identifies different groups of employees of a given firm within a province.

Conventional firms. We rely on the information on the legal structure coming from tax IDs to select limited liability, joint-stock companies, and foreign entities (first digit of the tax ID: A, B, N, and W) as well as sole proprietor enterprises. Conventional enterprises and labor societies are under the same regulation for corporate taxation, implying that the tax IDs do not allow for the distinction between them. Then, we use the type of the plant to separate capitalist firms (code 9999) from labor societies (5180).

Cooperatives. We use tax IDs to select cooperatives (first digit F), which are the type of organization fulfilling all the international standards to qualify under the cooperatives' class of enterprises. Additionally, we hinge on the type of labor relationship between worker and employer to isolate partners (code 930) from wage-earner employees in cooperatives.

Plant creation date. Date when the first employee was registered in the contribution account.

Plant size. Number of employees in the contribution account at the data extraction moment. In the case of inactive plants, this variable takes the value zero. This variable is available from 2005.

Industry. The MCVL provides information on the main sector of activity at a three-digit level (*actividad economica de la cuenta de cotizacion, CNAE*). Due to a change in the classification in 2009, the MCVL contains CNAE93 and CNAE09 for all plants observed in business from 2009 onwards, but only CNAE93 for those which stop their activity before. We rely on the CNAE09 classification when available and CNAE93 otherwise, exploiting the correspondence table provided by the Spanish National Statistical Office.³⁶ Then, we aggregate the three-digit industry information in 14 categories: primary sector (1 to 99), manufacturing and utilities (100 to 399); construction (411 to 439); wholesale and retail trade (451 to 479); transportation and storage (491 to 532); accommodation and food services (551 to 563); information and communication technologies (581 to 639); financial, insurance and real estate activities (641 to

³⁶http://www.ine.es/daco/daco42/clasificaciones/rev.1/cnae2009_cnae93rev1.pdf

683); professional, scientific and technical activities (691 to 750); administrative, support and other services (771 to 829 and 950 to 970); education, health and social work (851 to 889); entertainment (900 to 949); public administration and international organizations (840 to 849 and 990 to 999).

Unemployment rate. Refers to the provincial annual unemployment rate downloaded from the National Statistical Office: <http://ine.es/>

C Censoring correction

In the MCVL, the labor income concept referring to wage-employment relationships is either bottom or top-coded. In our estimation sample, 5.5% of the real daily wages of worker-month wage-employment observations are coded at the maximum. Following other studies that face censored earnings in administrative data (Dustmann et al., 2009; Card et al., 2013; Bonhomme and Hospido, 2017), we fit cell-by-cell Tobit models to log real daily wages to correct the upper tail of the wage distribution.³⁷

Each cell, c , is defined based on gender, occupational groups (3 categories), 7-year length age groups (3), and 5-year interval time groups (7) for a total of 126 cells. Within each cell, we assume log daily wages follow a Gaussian distribution, as it is common in the literature, with cell-specific mean and variance, i.e., $\log w \sim N(X\beta_c, \sigma_c^2)$.³⁸ Denoting Φ the standard normal cdf, the cell-specific maximum likelihood is (up to an additive constant)

$$\sum_{cens_{it}=0} \left[-\frac{1}{2} \ln \sigma_c^2 - \frac{1}{2\sigma_c^2} (\ln(w_{ijt}) - X_{it}\beta_c)^2 \right] + \sum_{cens_{ijt}=1} \ln \left(1 - \Phi \left(\frac{\ln(\bar{w}) - X_{ijt}\beta_c}{\sigma_c} \right) \right)$$

where w_{it} refers to log real daily wages of individual i in plant j in moment t (a worker-month pair), \bar{w} is the maximum cap, $cens_{ijt} = 1$ if the observation is top-coded. X_{ijt} is a set of controls, including age and categorical variables, for full-time jobs, sector of activity, workplace location, and time dummies. We follow Card et al. (2013) to also include individual-specific components of the wages using the mean log daily wages in other months and the fraction of censored wages in other months.

After the estimation, we replace each censored observation with the sum of the predicted wages and a random component drawn from a normal distribution with mean zero and cell-specific variance. The imputation rule is

$$\ln w_{ijt} = X_{ijt}\hat{\beta}_c + \hat{\sigma}_c \Phi^{-1} \left[\Phi \left(\frac{\ln \bar{w} - X_{ijt}\hat{\beta}_c}{\hat{\sigma}_c} \right) + u_{ijt} \times \left(1 - \Phi \left(\frac{\ln \bar{w} - X_{ijt}\hat{\beta}_c}{\hat{\sigma}_c} \right) \right) \right]$$

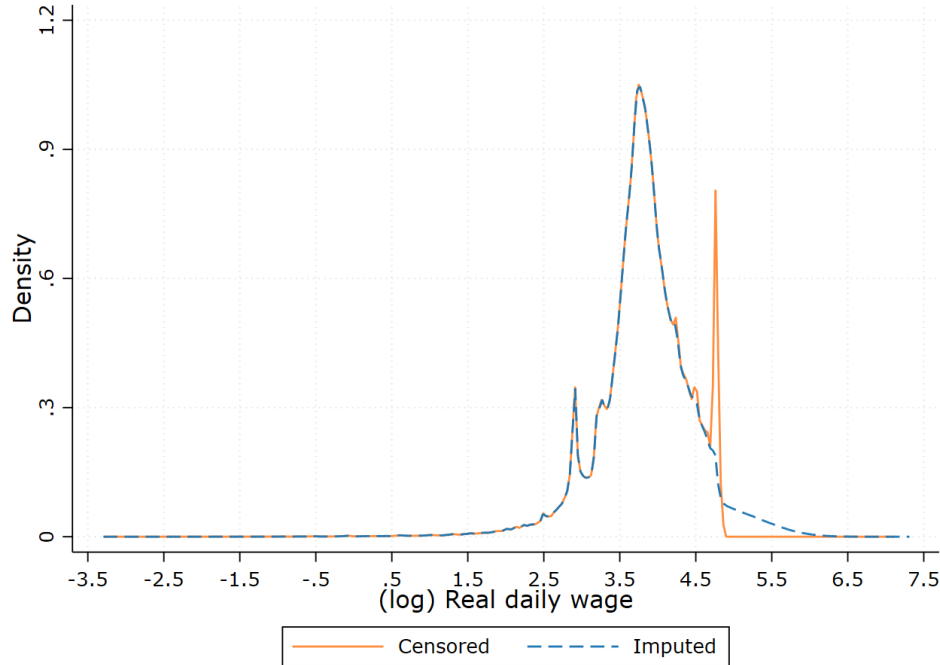
where $(\hat{\beta}_c, \hat{\sigma}_c)$ are the maximum likelihood estimates of each cell, Φ denotes the standard normal cdf, and u represents a random draw from the uniform distribution.

³⁷We do not correct the lower tail due to the existence of a national minimum wage.

³⁸Dustmann et al. (2009) provide a systematic comparison across four different distributional assumptions, and conclude that the results are similar across different specifications. Bonhomme and Hospido (2017) use the same data set as we do to evaluate the performance of the cell-by-cell Tobit model compared to a quantile censoring correction method with respect to uncensored income from tax records and find that the fit is superior with the Tobit model.

Figure C.1 and Table C.1 compare the original (censored) distribution with the imputed distribution resulting from the correction method.

Figure C.1: Censored and imputed wage distributions



Notes: Figure shows the distributions of censored and imputed distributions of (log) real daily wages based on month-worker-firm observations (23,086,331). Censored refers to the top-coded original wage distribution. Imputed stands for the imputed distribution based on the cell-by-cell Tobit model. Wages are expressed in 2018:12 euros deflated using the Spanish monthly consumer price index.

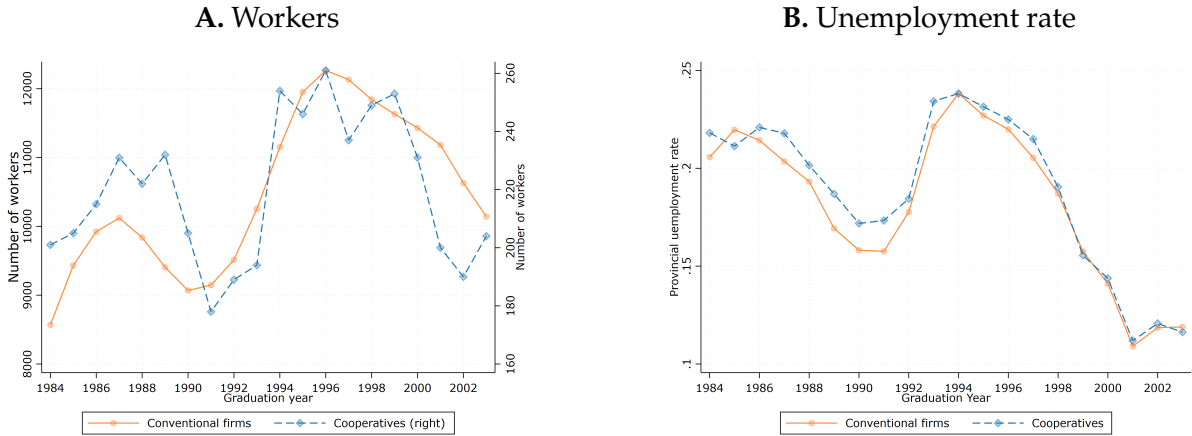
Table C.1: Moments of censored and imputed wage distributions

Percentiles	Censored	Imputed
5th	2.85	2.85
10th	3.12	3.12
25th	3.55	3.55
50th	3.82	3.82
75th	4.16	4.16
90th	4.54	4.55
95th	4.74	4.78
99th	4.79	5.47

Notes: Censored refers to the top-coded original wage distribution. Imputed stands for the imputed distribution based on the cell-by-cell Tobit model. Moments of the log daily wage distribution are computed over month-worker-firm observations (23,086,331). Wages are expressed in 2018:12 euros deflated using the Spanish monthly consumer price index.

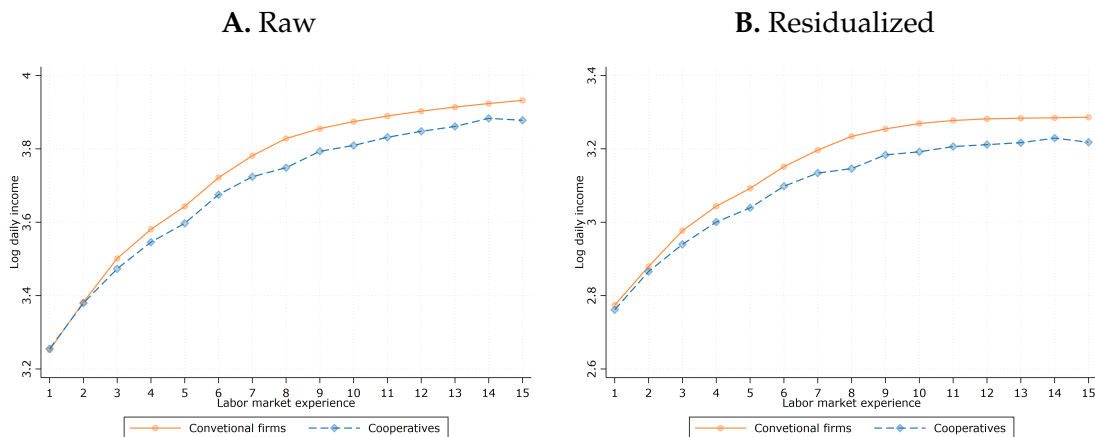
D Supplementary tables and figures

Figure D.1: Graduation cohorts by firm ownership of first job



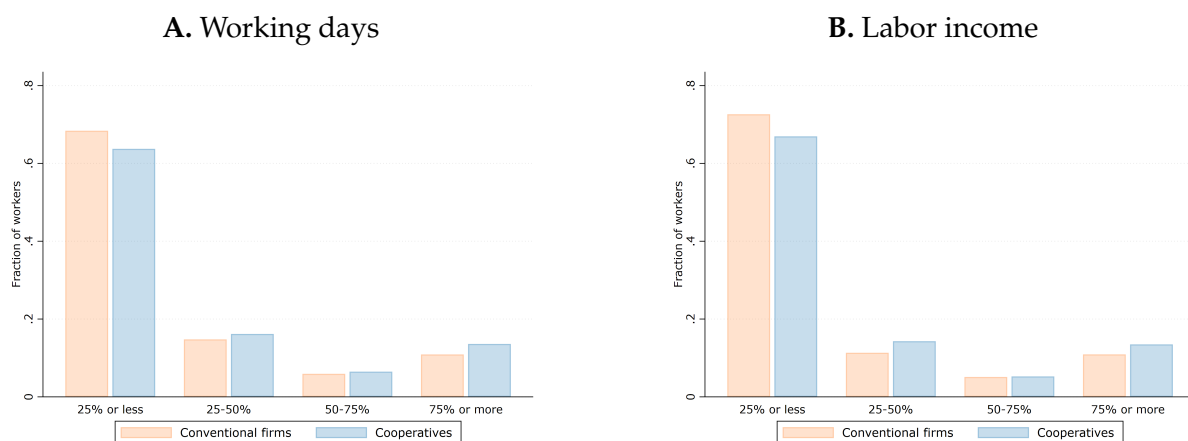
Notes: Panel A shows the number of workers by graduation cohort and ownership of the firm of the first job. Panel B displays the prevalent unemployment rate in the year of graduation in the province of workers' birth.

Figure D.2: Experience-wage profile by firm ownership of first job



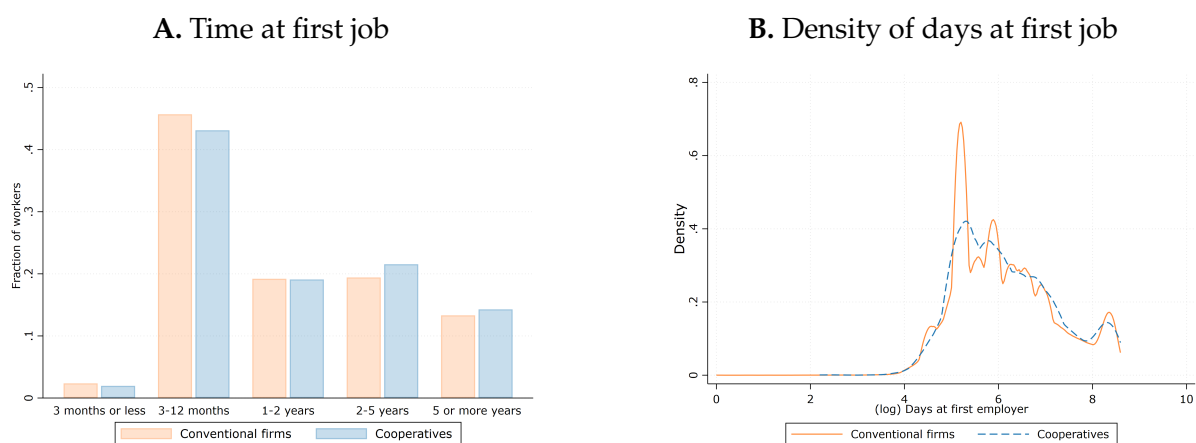
Notes: The figure shows raw (Panel A) and residualized (Panel B) experience-wage profiles by ownership of the firm of first employment. Residualized stands for daily wages after removing year and demographic (education and gender) effects. Wage refers to total labor income (traditional wage employment as well as income from self-employment and other forms of dependent employment) divided by days worked each year. Wages are expressed in 2018 euros deflated using the Spanish consumer price index.

Figure D.3: First job outcomes relative to career outcomes



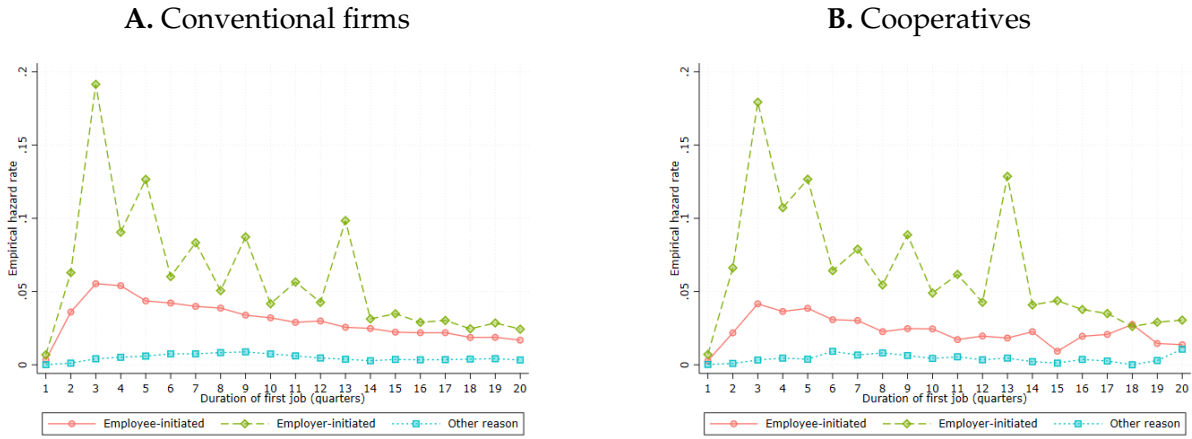
Notes: Figures show the distribution of total days worked (Panel A) and income earned (Panel B) in the first job relative to overall days worked and income earned during the first 15 years in the labor market by firm ownership of the first job.

Figure D.4: Duration of first job



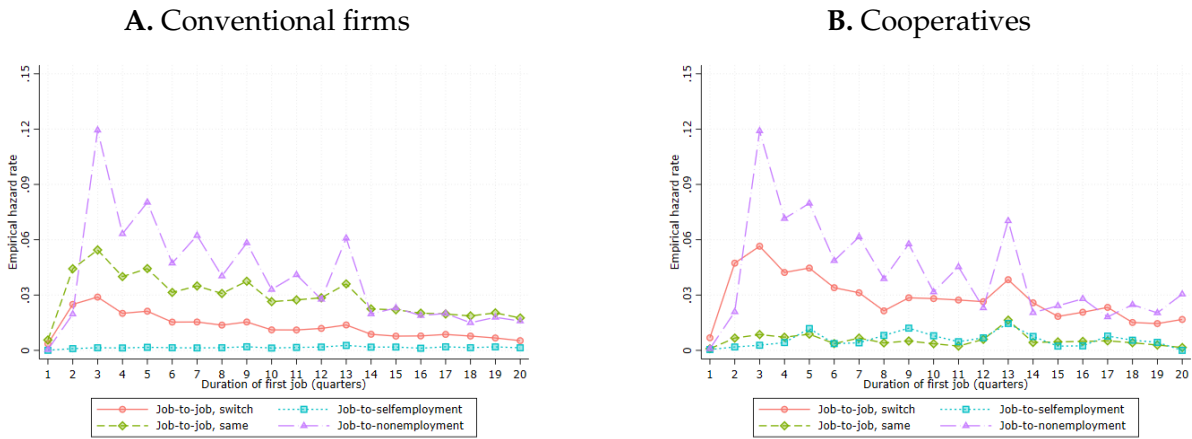
Notes: Panel A shows the share of workers by time spent in the first job. Panel B reports the distribution of (log) days worked in the first job.

Figure D.5: Exit from first job by type of separation



Notes: The figure shows the empirical rates of exit from the first job by the type of separation and ownership of the first employer, i.e., conventional firms (Panel A) and cooperatives (Panel B). "Other reason" is a residual category including, among others, sickness or parental leave, but mostly refers to undefined/administrative causes.

Figure D.6: Exit from first job by destination state

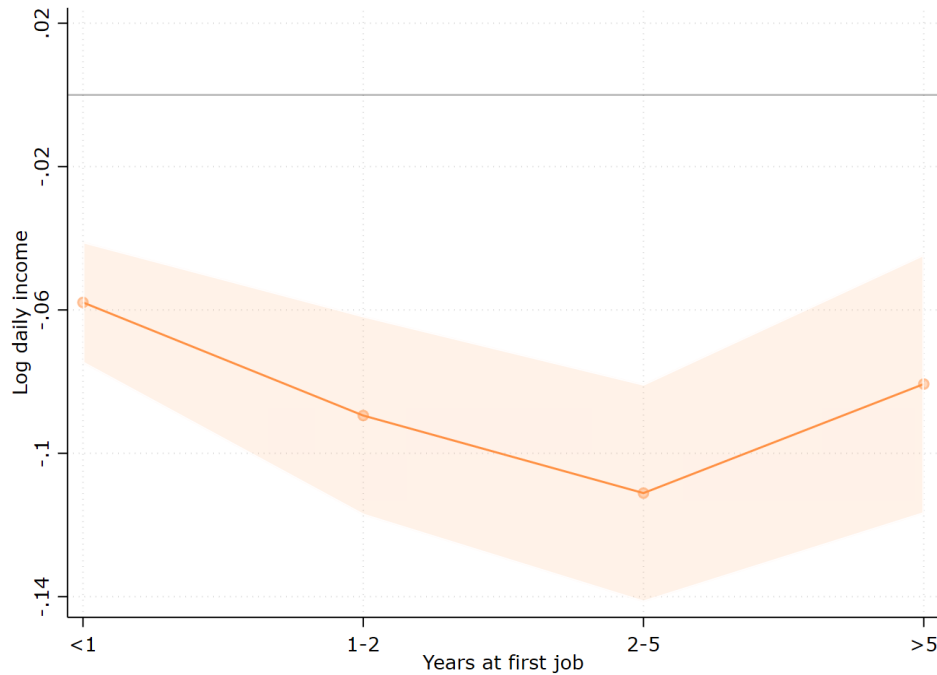


Notes: The figure shows the empirical rates of exit from the first job by destination state and ownership of the first employer, i.e., conventional firms (Panel A) and cooperatives (Panel B). "Switch" refers to workers who changed employer type, while "same" refers to those who remained with the same employer type as in the first job. "Non-employment" means transitions with 6 or more months between the end of the first job and the beginning of the next job.

Table D.1: Reason for separation from the first job

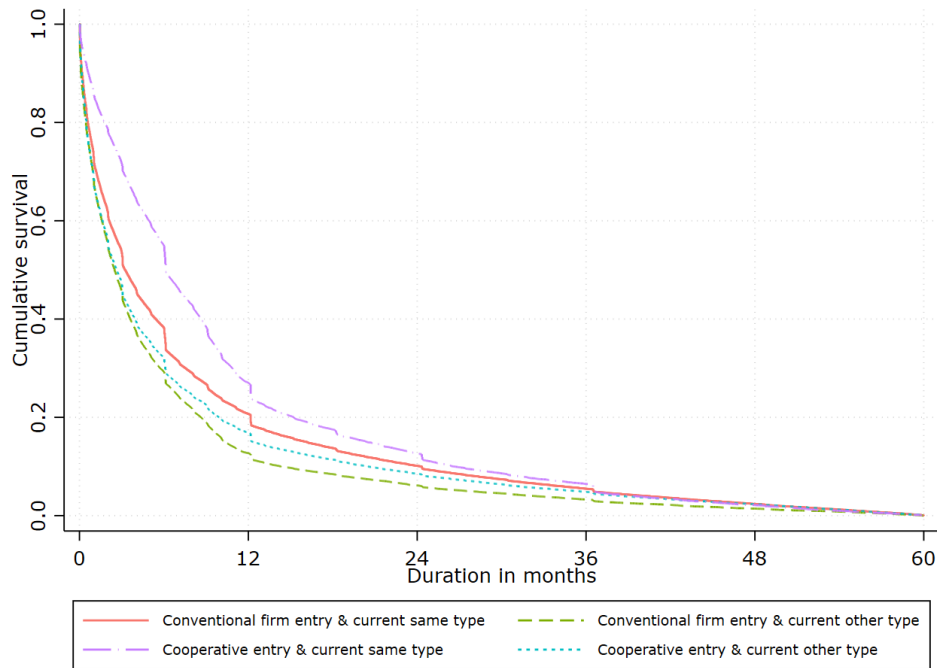
	Competing-risk		
	(1) Voluntary	(2) Involuntary	(3) Other
First job at cooperative	-0.141** (0.059)	0.017 (0.032)	-0.194 (0.128)
Observations	2,198,595	2,198,595	2,198,595

Notes: The table reports the results from a discrete-time duration model with competing risks (multinomial logit) for the likelihood of separating from the first job using three types of separations: voluntary, involuntary, and other. Voluntary and involuntary movers stand for workers whose separation from their first job was initiated by the employee or the employer, respectively. Other types of separation is a residual category including, among others, sickness or parental leave, but mostly refers to undefined/administrative causes. Controls include a firm-type specific quadratic polynomial on job duration, province of birth and graduation cohort fixed effects, province-specific unemployment rate at the time of labor market entry (cubic polynomial), seasonal effects, gender, education level, and first-job-employer variables such as skill category, full-time status, sector, and firm age. Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Figure D.7: Impact of first job in a cooperative: Heterogeneity by the duration of the first job

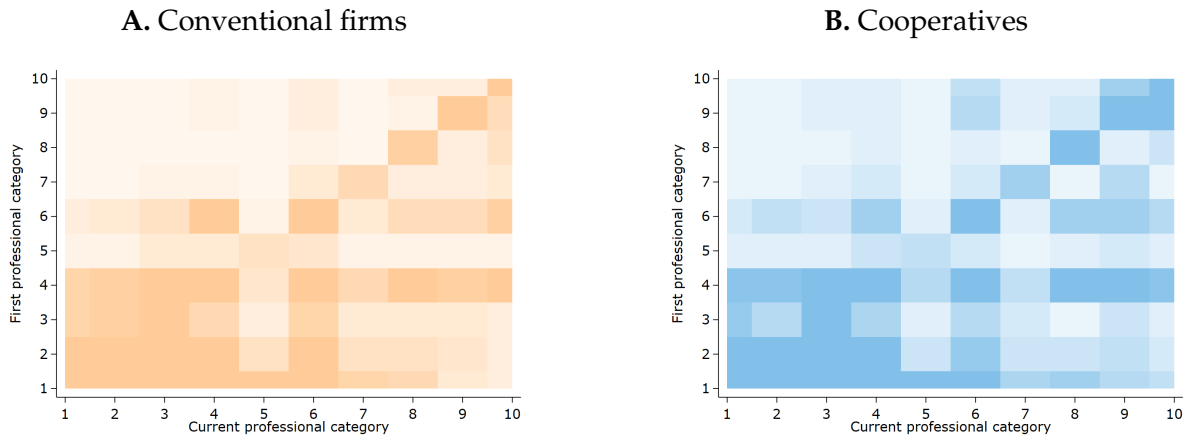
Notes: The figure shows the wage differential for first job experience in a cooperative relative to a conventional firm, estimated from an augmented version of Equation 1 in which the cooperative dummy variable interacts with the indicators for each first job length category: less than a year, between 1 and 2 years, 2 to 5 years, and 5 years or more. The regression includes the same set of controls as Column (4) in Table 1. 95% confidence bands based on standard errors clustered at the level of the province of birth \times graduation year.

Figure D.8: Kaplan-Meier survival estimates of job duration



Notes: The figure shows the survival functions of the duration of the labor relationship separating spells by workers' firm ownership of their first job and the type of current labor relationship. "Current other type" refers to wage-employment spells in an organization with an ownership structure different from that of the first job, as well as periods of self-employment.

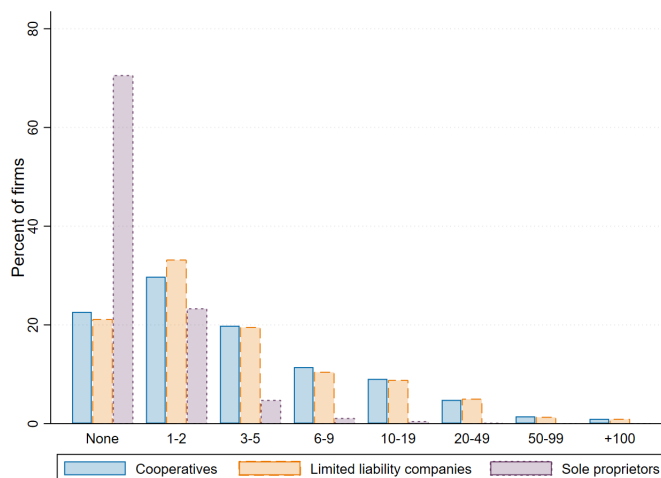
Figure D.9: Professional mobility: Transitions between first and last observed job



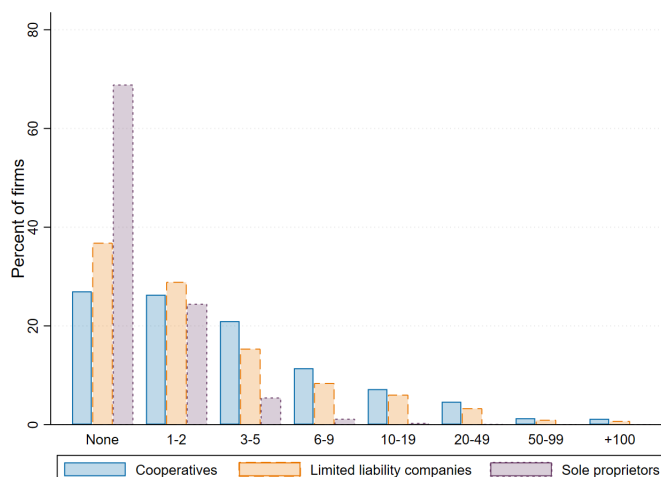
Notes: The figure shows a heat plot of transition matrices of professional categories between the first and last observed job for workers who started in conventional firms (Panel A) and cooperatives (Panel B). Professional categories are constructed using Social Security contribution groups. The MCVL contains 10 different contribution groups that are aggregated based on similarities in skill requirements. Group 10 (engineers, college, senior managers—in Spanish *ingenieros, licenciados y alta direccion*), Group 9 (technicians—in *ingenieros tecnicos, peritos y ayudantes*), Group 8 (administrative managers—in *jefes administrativos y de taller*), Group 7 (assistants—in *ayudantes no titulados*), Group 6-4 (administrative workers—in *oficiales administrativos* (6), *subalternos* (5) and *auxiliares administrativos* (4)), Group 3-1: (manual workers—in *oficiales de primera y segunda* (3), *oficiales de tercera y especialistas* (2) and *mayores de 18 años no cualificados* (1))

Figure D.10: Firm size distribution by firm ownership in selected years

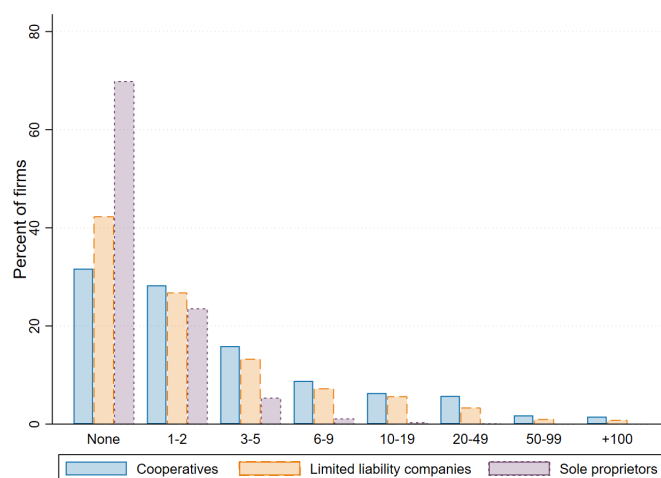
A. 2000



B. 2010



C. 2020



Notes: The figure shows the firm size (average number of employees in a year) distribution for selected years when there is available data from the *DIRCE* database of the Spanish Statistical Office. Firms with no employees (the *none* category) correspond to organizations in which the owner is the only worker.

E Robustness checks

We perform several sensitivity tests to validate our benchmark results. We start assessing the robustness of our estimates to the definition of the first job after graduation. The estimates in Table E.1 indicate that alternative definitions produce virtually the same results. Similarly, our results are not affected by the definition of earnings, as the results remain qualitatively unchanged when using censored income or uncorrected income or eliminating the observation (see Table E.2).³⁹

Table E.1: Impact of first job in a cooperative: Robustness to the definition of first employment after graduation

	(1) Benchmark	(2) Job spell of 100 days	(3) 180 days worked in 12 months	(4) 1st job within 3 years
First job at cooperative	-0.080*** (0.007)	-0.078*** (0.007)	-0.086*** (0.008)	-0.072*** (0.008)
Observations	24,659,247	25,700,745	22,442,022	17,656,608

Notes: Column (1) is our benchmark definition of the first job, referring to the first 6 consecutive months after the year of graduation when individuals work at least 100 days. Column (2) defines first employment as the first job spell after the year of graduation, lasting at least 100 days. Column (3) specifies first employment the first 12 consecutive months after the year of graduation when individuals work at least 180 days. Column (4) considers only workers whose first job, defined as the first 6 months after graduation when individuals worked more than 100 days, occurs no later than 3 years after the year of graduation. All specifications include the same set of controls as Table 1 Column (4). Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table E.2: Impact of first job in a cooperative: Robustness to top-coded income

	(1) Corrected	(2) Censored	(3) No censored
First job at cooperative	-0.080*** (0.007)	-0.067*** (0.006)	-0.054*** (0.006)
Observations	24,659,247	24,659,247	23,395,870

Notes: Column (1) uses censoring-corrected observations following the imputation method described in Appendix C. Column (2) uses uncorrected earnings. Column (3) removes censored observations from the sample. All specifications include the same set of controls as Table 1 Column (4). Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

We also assess the sensitivity of our results to the impact of the incidence of self-employment on workers' careers. This is particularly relevant in the Spanish context, as cooperative partners can decide to contribute either to the General Regime of the Social Security for salaried workers or to the Special Regime for self-employed workers. In the latter case, individuals' Social Security contributions are based on a notional

³⁹The smaller penalty observed when censored observations are not corrected, or not used at all, is due to the maximum earnings gap and a higher incidence of censored observations at the top among workers who started their careers in conventional firms.

income statement rather than on actual earnings. This suggests that the previously documented wage penalty could partly be an artifact of income reporting issues.⁴⁰ To rule out this possibility, in Table E.3, we implement different sample constraints to mitigate the incidence of self-employment. For instance, in Column (1), we restrict the sample to individuals who contributed to the General Regime for wage-employment relationships at least 60% of the time over their first 15 years in the labor market. The wage penalty experienced by individuals who kicked off their careers at a worker cooperative remains negative and statistically significant, albeit smaller in comparison to our baseline estimate. Similar results are obtained when alternative sampling restrictions are introduced, suggesting that the penalty is not due to workers becoming self-employed and under-reporting their earnings.

Table E.3: Impact of first job in a cooperative: Robustness to the type of labor relationships

	(1) Mainly GR	(2) LM Attachment	(3) LLC	(4) WE Income
First job at cooperative	-0.064*** (0.007)	-0.081*** (0.007)	-0.083*** (0.008)	-0.060*** (0.007)
Observations	22,929,433	22,806,452	18,280,594	23,086,331

Notes: Column (1) specification includes only workers who spend at least 60% of the first 15 years in the labor market employed in standard wage-employment relationships (General Regime of the Social Security). Column (2) specification considers only workers who during the first 15 years in the labor market (LM) were employed more than 50% of the time. Column (3) specification uses only employment relationships with limited liability companies (LLC), i.e., excludes employment relationships with individual enterprises as well as other special forms of businesses. Column (4) considers only income coming from wage-employment relationships. All specifications include the same set of controls as Table 1 Column (4). Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Our benchmark analysis excludes Labor Societies, another form of employee-owned enterprise in Spain, and identifies worker cooperatives simply by looking at the legal cooperative status of individuals' first employer. Column (1) of Table E.4 shows that including Labor Societies as part of cooperatives does not affect the estimated penalty. Similarly, by defining cooperatives based exclusively on their legal status as determined by their relationship with the Tax Authority, our cooperative dummy may be picking up the effect of other types of cooperative firms that do not involve substantial levels of employee ownership (e.g., credit cooperatives, consumer cooperatives).

⁴⁰Unfortunately, we cannot track the link between individuals and cooperative firms if individuals contribute to the self-employed regime. In other words, we cannot distinguish between genuine self-employed individuals and worker-partners employed in a worker cooperative who contribute to the regime for self-employed workers.

In Column (2), we restrict the definition of cooperatives to firms specifically coded as worker cooperatives by the Spanish Security Administration. In Column (3), we restrict the analysis to individuals who had an initial job in Manufacturing, a sector in which the cooperative firm legal status unambiguously identifies worker cooperatives. These alternative definitions yield a larger wage penalty, but the interpretation of the results holds. Finally, in Column (4), we distinguish salaried employees and partners among individuals who had an initial job experience in a cooperative. The point estimates in this specification suggest that partners have a larger wage penalty. However, this may just be a reflection of the fact that wages are *not* the only source of income for partners or that these workers are more likely to become self-employed.

Table E.4: Impact of first job in a cooperative: Robustness to the definition of cooperatives and inclusion of labor societies

	(1) Including LSs	(2) Only WCs	(3) Only Manufacturing	(4) Partners/employees
First job at cooperative	-0.079*** (0.006)	-0.137*** (0.013)	-0.113*** (0.011)	
First job at cooperative - Employee				-0.071*** (0.007)
First job at cooperative - Partner				-0.176*** (0.030)
Observations	24,659,247	24,659,247	5,279,850	24,659,247

Notes: In Column (1) specification, the first job at a cooperative is a dummy variable, equal to one if the individual had an initial job experience either in a cooperative or a labor society, and zero otherwise. Column (2) specification restricts the definition of cooperative to firms coded as worker cooperatives by the Spanish Social Security Agency. Column (3) specification restricts the analysis to individuals who had an initial job experience in Manufacturing, where the cooperative status unambiguously identifies employee-owned firms (worker cooperatives). In Column (4), we report estimates distinguishing salaried employees and partners among individuals who had an initial job experience in a cooperative as employees or partners (the omitted category refers to individuals who held their first job in a conventional business). All specifications include the same set of controls as Table 1 Column (4). Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Recent work by [Arellano-Bover \(2024\)](#) shows that firm size can be considered a sufficient statistic for several earnings-enhancing firm characteristics that are difficult to observe (e.g., management quality). Indeed, using similar data from Spain, he shows that firm size is associated with higher earnings. Therefore, one could argue that the documented wage penalty may simply reflect that cooperative firms are smaller than their conventional counterparts. Unfortunately, in our dataset, information on firm size is only available between 2005 and 2018. This implies that we have to restrict the analysis to workers whose first job is observed when firm size is available, but also to a subset of years in order to be able to observe wage developments over time. In Table E.5, we report estimates of this model using graduation cohorts whose first job occurred between 2005 and 2009 and including first-employer size as an additional

control. In line with [Arellano-Bover \(2024\)](#), firm size is positively correlated with individual earnings. However, the wage penalty associated with having an initial cooperative job remains virtually unchanged, suggesting that firm size is not an important factor underlying our results. The fact that firm size plays little role in explaining the cooperative wage penalty is not surprising. Using aggregate data from the Spanish Statistical Office, [Figure D.10](#) shows no clear differences in the size distribution of the two types of firms.

Table E.5: Impact of first job in a cooperative: Robustness to the size of the first employer

	Graduation year: 1999-2003		
	(1) All workers	(2) First job after 2004	(3) (2) + Firm size
First job at cooperative	-0.064*** (0.011)	-0.112*** (0.027)	-0.114*** (0.026)
(log) Firm Size			0.035*** (0.001)
Observations	6,570,628	1,847,141	1,847,141

Notes: Column (1) includes all workers whose year of graduation is between 1999 and 2003. Column (2) considers only workers from Column (1) whose first job occurs between 2005 and 2009. Column (3) adds to Column (2) (log) the firm size of the first employer as an additional control. All specifications include the same set of controls as Table 1 Column (4). Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

While focusing on the first 15 years of a worker's career excludes several years of working life, existing evidence suggests that the bulk of lifetime wage growth occurs in the first 10 years of labor market entry and stabilizes in the late thirties and early forties see, for example, [Topel and Ward \(1992\)](#) for the US, and [Davia \(2010\)](#) or [Arellano-Bover \(2024\)](#) for the case of Spain. Thus, by looking at the first 15 years, we aim to capture the most relevant part of the labor market career while following the largest number of cohorts in our sample. However, for a subset of cohorts, those who graduated between 1984 and 1989, we can follow them over a maximum horizon of 30 years and thus assess the sensitivity of the wage gap when more years of a worker's career are included. [Table E.6](#) shows that the estimated wage penalty of starting a career in a cooperative compared to a conventional firm for these cohorts barely changed when the career horizon was extended by 15 additional years.

Table E.6: Impact of first job in a cooperative: Robustness to career horizon considered

	All workers	Graduation year: 1984-89	
	(1)	(2)	(3)
	First 15 years	First 15 years	First 30 years
First job at cooperative	-0.080*** (0.007)	-0.047*** (0.012)	-0.049*** (0.013)
Observations	24,659,247	6,195,143	13,692,157

Notes: Column (1) replicates our benchmark results in Table 1 Column (4). Column (2) estimates the wage gap for cohorts 1984-1989 over the first 15 years in their career, whereas Column (3) considers the first 30 years. All specifications include the same set of controls as Table 1 Column (4). Standard errors clustered at the level of the province of birth \times graduation year. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

F Omitted variable bias: Oster test

To quantify the role of unobservable factors in our estimates, we follow the strategy proposed by Oster (2019). The approach is based on the omitted variable bias formula and allows bounding the degree of selection bias under the assumption that the selection on unobservables is proportional to the selection on observed characteristics. Under this assumption, we can (i) estimate the bias-adjusted wage effect, $\tilde{\beta}$, assuming a certain degree of selection on unobservables relative to observed factors, and (ii) quantify the degree of selection, δ , that would be required to our benchmark wage effect, β , becomes zero. In this context, $|\delta| > 1$ would imply that a very large degree of selection on unobservable variables would be necessary to cancel out the uncovered wage penalty, suggesting that omitted variable bias is unlikely to fully account for the wage penalty documented in our baseline estimates.

In Table F.1, we report the results from this bounding exercise.⁴¹ Column (1) reproduces our benchmark wage effect, which would equate to the “true” effect of having an initial work experience in a cooperative under the assumption that there is no omitted variable bias due to selection on unobservables ($\delta=0$). The bias-adjusted wage effect, $\tilde{\beta}$, in Column (2) indicates that if selection on unobservables were equally important than that on observables ($\delta=1$), the identified wage effect would be -0.088, or 10% higher than our benchmark estimate. In Column (3), we present the degree of selection on unobserved attributes relative to observed factors that would be needed for the wage effect to disappear. The results imply that the relative magnitude of selection on unobservables, δ , would have to be 11 times larger and have the opposite sign of selection on observed characteristics for the identified wage effect to disappear. Given the large set of controls we include in our benchmark specification, this degree of selection seems implausibly high to suggest that our identified wage effect is attributable to the non-random selection of workers across firm types at the time of labor market entry.

⁴¹To implement the approach, we assume, as in Oster (2019) that the value of the R-squared in a hypothetical regression including unobserved confounding factors (R_{max}) would be $1.3\tilde{R}$, where \tilde{R} is the R-squared from the controlled regression reported in Column (4) of Table 1.

Table F.1: Oster's bounds estimates

	(1) Baseline	(2) Bias-adjusted effect	(3) Degree of selection (δ)	(4) Max-negative effect
First job at cooperative	-0.080	-0.088	-11.302	-0.170

Notes: The table displays results from the bounding exercise proposed by Oster (2019). Column (1) reproduces our baseline estimate under the assumption of no selection on unobservables, as originally reported in Column (4) of Table 1. Column (2) shows the bias-adjusted β estimate under the assumption that selection on unobservables is equal to selection on observed factors ($\delta = 1$). Column (3) reports the relative amount (and sign) of selection on unobservables needed to drive the coefficient of interest to zero, i.e., δ . The value of the R-squared of the hypothetical regression including unobserved confounding factors (R_{max}) is assumed to be $1.3\bar{R}$, where \bar{R} is the R-squared from the controlled regression in Column (1). Column (4) shows the maximum wage penalty assuming that the selection on unobservables has the same sign as the selection on observables and imposing the highest degree of selection ($\delta = 11.302$).

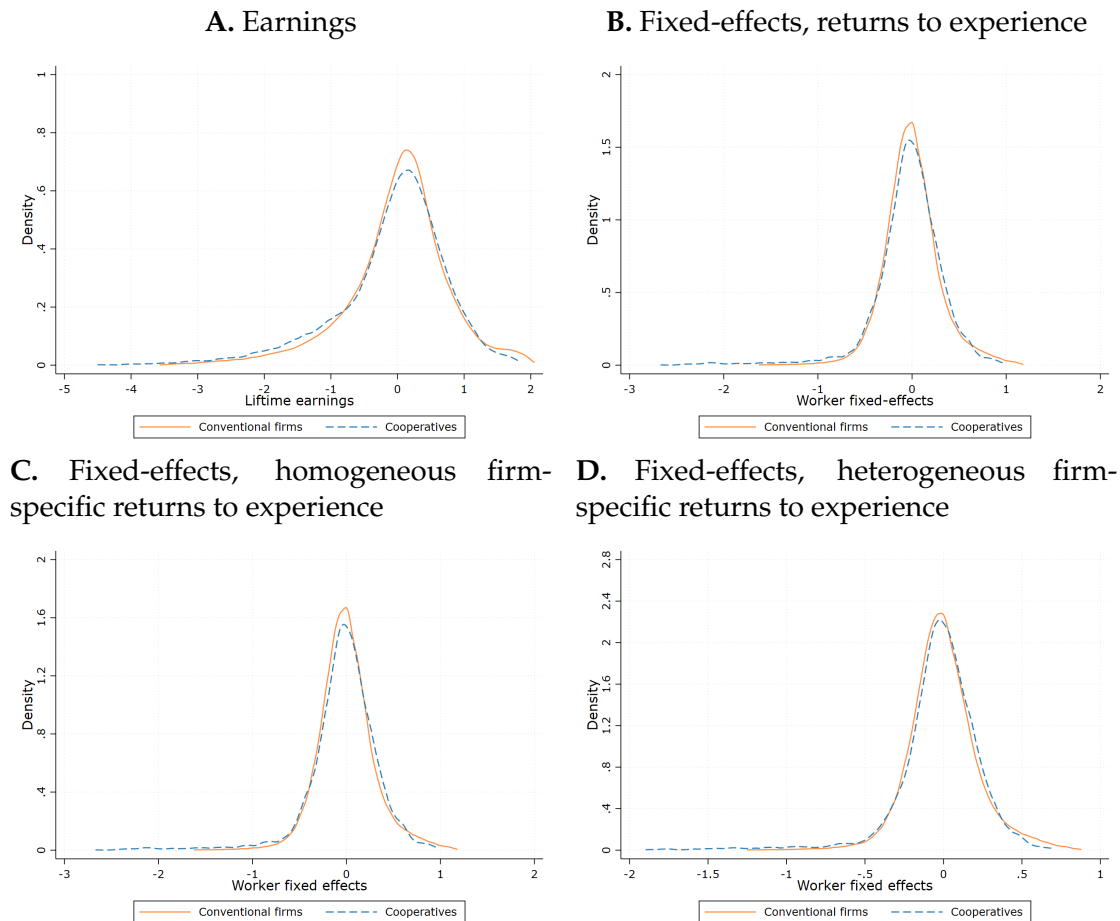
G Distribution of worker fixed effects

The comparison between the observed ability of workers in cooperatives relative to conventional enterprises suggests that workers are more educated and occupy more skilled positions, but there remains the possibility that they may not be the “best” individuals within these categories. In other words, sorting into unobservables might be more salient. To delve into this issue, we investigate the distribution of unobserved ability, as measured by different estimates of worker fixed effects, across workers classified by the type of their first employer. We closely adopt the strategy of [de la Roca and Puga \(2017\)](#) and estimate different types of Mincer equations that allow the wage returns to experience to be (i) common across firms, (ii) heterogeneous across firms, and (iii) heterogeneous across firms and workers, by interacting firm-specific returns to experience with the worker fixed effect. Note that if having a first job in a regular firm makes workers relatively more productive, the worker fixed effect would partially capture this effect. In other words, these worker FEs could be an indirect function of the first job and would not accurately capture the extent to which selection into cooperatives is driven by individuals’ predetermined characteristics. However, allowing for heterogeneous returns to experience by *both* employer type and worker fixed effects largely mitigate this concern.

In Figure G.1, we plot estimates of unobserved ability based on these alternative regressions and also show the distribution of total earnings for comparison. The total earnings distribution (Panel A) of workers whose first job was in a cooperative compared to those who started their career in conventional firms, the distribution of workers in cooperatives shows a lower mean. In addition, it also shows greater dispersion and a higher degree of negative skewness. However, when comparing the unobserved ability of workers across firm types, the two distributions become more similar. In particular, we show that the distributions of worker fixed effects are also similar when we allow for firm-specific returns to experience to be heterogeneous across firms and workers (Panel D). These differences across ability distributions arise because the worker-fixed effects in Panel B and Panel C not only capture the time-invariant heterogeneity across individuals but also embed the time-varying effect of working for a given type of firm (returns to firm-specific experience), as well as the interaction of that effect with workers’ innate abilities. While accounting for heterogeneous career paths that are plausibly not linked to unobserved ability suggests that permanent

worker heterogeneity among labor market entrants in cooperatives and conventional firms is very similar (Panel D), the two distributions still show statistically significant differences in the tails: there is a larger (smaller) mass of low (high) ability workers in cooperatives.⁴²

Figure G.1: Worker unobserved heterogeneity by firm ownership of first job



Notes: Panel A shows the distribution of total labor income earned during the first 15 years in the labor market by firm ownership of the first job. Panel B shows the distribution of worker-fixed effects from a wage equation controlling for actual experience, tenure, skill level, full-time status, current cooperative employer, workplace location, sector of activity, and time-fixed effects. Panel C plots the distribution of worker fixed-effects, extending the wage equation of Panel B to allow returns to experience to vary by firm-ownership, whereas Panel D further extends such wage equation to allow the returns to experience to vary by firm-ownership *and* worker fixed-effects using the algorithm proposed by de la Roca and Puga (2017). Distributions are centered at the average value of workers who had their first job in a conventional enterprise.

⁴²Using the Kolmogorov-Smirnov test, we can reject the null of equality of the distributions (p -value=0.006). We also estimate quantile regressions for each decile of the fixed effects distribution on the cooperative entry dummy and find statistical differences (at the 5 percent level) only below the 10th percentile and above the 90th percentile.