Temporary Jobs, Permanent Consequences*

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Abstract This paper examines the medium-run implications of entering the labor market through a fixed-term rather than an open-ended contract. Using rich Italian administrative matched employer-employee population data, we track workers' earnings and employment trajectories over the first ten years of their careers. We document that fixed-term entrants initially earn less, but catch up rapidly and, by the third year of labor market experience, surpass open-ended entrants in terms of annual earnings. This convergence is primarily driven by a quick reduction in time spent in non-employment, while wages per unit of time worked remain broadly similar over time across the two entry groups. We first provide descriptive empirical evidence of these patterns and then estimate an instrumental variable model, exploiting exogenous variation in the probability of starting a first job under a fixed-term versus an open-ended contract, to address potential endogenous selection bias.

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1 Introduction

Many European labor markets are characterized by a dual contractual structure where stable long-term jobs coexist with short-duration, fixed-term contracts. In these markets, workers can start their career on an open-ended or a fixed-term job. Does the entry contract type have any persistent effect on labor market attainment over time?

There is a large body of evidence on the long-run effects of labor market entry conditions. In most cases, it is focused on business cycle fluctuations and it finds persistent negative effects on income and occupational attainment for those who enter the labor market during downturns (Kahn, 2010; Oreopoulos et al., 2012; Brunner and Kuhn, 2014; Altonji et al., 2016; Fernández-Kranz and Rodríguez-Planas, 2018; Schwandt and Von Wachter, 2019; Bentolila et al., 2021; Bertheau et al., 2022). Another, more recent strand of research documents instead the role of workers' first employer, finding that initial matches with larger firms – or, more generally, with high-growth firms – have positive and substantial effects on long-term individual outcomes (Arellano-Bover, 2024; Gregory, 2020).

We conjecture that the degree of contractual stability at labor market entry could have implications that are as important as those related to initial firms' characteristics or to business cycle fluctuations. First, obtaining a fixed-term job as an initial employment opportunity may negatively impact the rate of human capital growth over time. Employment under a fixed-term contract, as opposed to an open-ended contract, is associated with an increased probability of transitioning to unemployment (Boeri and Garibaldi, 2024) – resulting in direct consequences in terms of diminished on-the-job learning opportunities – as well as a lower availability of workplace training (Cabrales et al., 2017).

Furthermore, for individuals with limited financial resources, the heightened risk of unemployment faced under fixed-term contracts could lead to suboptimal long-term decisions. Specifically, fixed-term workers at risk of becoming unemployed may choose to accept lower wages or worse working conditions, or even take any job offer, including those for which their skills and experiences do not align with the job requirements. In equilibrium, these phenomena could subsequently hinder opportunities for career growth, as young workers may choose to forgo prospects of higher earnings growth in favor of achieving greater job stability – that is, securing

an open-ended contract.

These drawbacks notwithstanding, fixed-term jobs may also offer several advantages for young workers. First, their typically short duration induces frequent transitions between different employers, allowing young individuals to gain experience across diverse sectors and work environments. Second, fixed-term contracts facilitate the swift termination of unproductive matches with minimal frictions – unlike open-ended contracts, which involve severance costs for employers and may align with workers' preferences for job stability, irrespective of the wage or the productivity level of the job match.

In this paper, we study the medium-term effects of the type of initial contract, primarily focusing on earnings as the outcome of interest. To do so, we analyze a sample of young workers entering the labor market for the first time, starting from 2005. The data source is the Italian Social Security Institute (INPS) administrative records, which cover the entire population of Italian employees and employers. We construct a balanced panel of workers and follow them over the first ten years of their careers, which can include non-employment spells.

We begin by examining how yearly earnings evolve over the first ten years of a worker's career, distinguishing between individuals who initially held a fixed-term contract, called *fixed-term entrants*, and those with an open-ended contract, called *open-ended entrants*. A key challenge in this analysis lies in accounting for self-selection of workers into initial contract types. The same underlying factors that influence early-career earnings growth may indeed also drive the choice or assignment of contract type at labor market entry. Failing to account for this selection mechanism could lead to an incorrect assessment of the contribution of contract type to earnings differences.

To address this issue, we start by exploiting the rich information set available in administrative data, which allows us to compare earnings trajectories among individuals with similar demographic characteristics who begin their careers in comparable firms – defined by size, productivity, and age – at the same point in both age and calendar time. This descriptive comparison is conducted in an OLS framework. We then complement this evidence from a residual approach with an instrumental variable analysis, which relies on exogenous variation in the probabilities of starting a first job under a fixed-term or an open-ended contract.

We find that workers who begin their careers with a fixed-term contract initially have lower gross annual earnings than those starting with an open-ended contract. However, already after the first year in the labor market, a convergence process begins, marked by relatively faster earnings growth for fixed-term entrants. By the third year, the gap reverses and fixed-term entrants earn more on average.

These dynamics are not uniform across education levels. In particular, workers with lower levels of education experience more pronounced earnings growth, that even exceeds mere convergence. Specifically, workers with less than a high school diploma experience a substantial positive earnings gap since entry compared to peers who entered directly into an open-ended job. In contrast, highly educated workers are, on average, relatively disadvantaged by starting with a fixed-term contract, although a partial convergence process is still evident.

To better understand the dynamics behind these results, we decompose annual gross earnings into two determining factors: working time and earnings per unit of time. We focus specifically on the number of full-time equivalent weeks worked during the year and full-time equivalent weekly earnings to study which of the two margins mostly explains the different earnings dynamics of workers who started their careers with different contracts. We find that the trend in the earnings gap between the two groups of workers over time is almost entirely explained by the amount of time worked, with only a marginal role of the differences in weekly earnings between the two groups, which remains relatively constant over the observed period. These results indicate that the convergence in annual earnings primarily results from workers entering the labor market with fixed-term contracts initially working for a relatively lower number of weeks during the year. Over time, however, the number of full-time equivalent weeks worked increases, eventually exceeding the amount of time worked by those who entered with open-ended contracts, again except in the case of workers with a university degree, for whom the gap just disappears.

Our results offer new empirical evidence to the academic and public debate on the heterogeneous implications of the widespread use of fixed-term contracts. Specifically, these findings challenge the prevailing view in the academic literature, which typically associates fixed-term contracts with earnings disadvantages in a cross-sectional and static framework. In a more dynamic

setting, some studies have documented lifetime convergence in earnings between fixed-term and open-ended workers. Our analysis extends this line of research by focusing specifically on new entrants to the labor market and by adopting a dynamic medium-run perspective. In this context, by tracking the careers of this group over time, we find not only convergence in earnings but also that, for a subset of workers, fixed-term entrants eventually surpass their open-ended counterparts.

This paper relates to two main strands of the literature. The first one examines the long-term consequences of initial labor market conditions. As already mentioned, entering the labor market during a recession or under unfavorable conditions – such as joining a low-productivity or small firm – can have persistent negative effects on earnings and career progression. Beyond the interest in these so-called *scarring effects* per se, this approach provides a reduction of selection bias in the estimation of causal effects due to labor market entrants having very little or no employment history.

The second strand focuses on the implications of fixed-term employment (Boeri and Garibaldi, 2024). Prior work has documented wage penalties associated with these contracts (Leonardi and Pica, 2013; Daruich et al., 2023), with evidence of persistent earnings losses – for instance, a 7 percent reduction for low-skilled workers who begin their careers on fixed-term contracts (García-Pérez et al., 2019). Several mechanisms have been proposed to explain these penalties. One highlights the greater bargaining power of open-ended workers, which may result in both higher wages and improved future opportunities (Bentolila and Dolado, 1994). Another line of work emphasizes how fixed-term contracts are linked to lower returns to accumulated experience (Garcia-Louzao et al., 2023). Additional studies have underlined that fixed-term workers often receive less training, which may hinder long-run labor market outcomes (Berton and Garibaldi, 2012; Cabrales et al., 2017; Bratti et al., 2021). This literature about the implications of fixed-term employment has also explored whether fixed-term contracts serve as stepping stones to more stable positions or instead act as dead ends (Booth et al., 2002; Bentolila et al. 2020), with recent work emphasizing the role of firms in enabling or obstructing transitions to permanent employment (Güell and Petrongolo, 2007; Filomena and Picchio, 2022).

This paper bridges these two strands of the literature, by investigating how the type of initial

contract affects the performance of early career trajectories. It is organized as follows. section 2 outlines the data used for analysis. section 3 details the estimation strategy. In section 4, we discuss the main empirical findings. section 5 offers concluding remarks.

2 Data

The analysis is based on administrative data provided by the Italian Social Security Institute (INPS, Istituto Nazionale della Previdenza Sociale), which record the employment and pay histories of the entire population of private-sector employees workers outside agriculture. This includes workers in publicly owned companies but excludes public administration employees and the self-employed.[1] The dataset spans the period from January 1974 to December 2023.

It offers rich information on individual worker characteristics and job details. Specifically, it includes the start and end dates of each registered employment spell, contract type (open-ended, fixed-term, or seasonal), occupation, part-time status, job classification (white-collar, blue-collar, manager, or apprentice), and the geographic location of employment. The data also report annual taxable labor earnings, firm identifiers, and records of unemployment benefits and other earnings support programs. On the demographic side, the dataset provides information on gender, education, place of birth, and dates of birth and death.

The dataset also includes detailed information on firms, enabling the matching of the complete population of Italian workers to their respective employers. At the firm level, it includes industry classification, average wages, workforce size at various points in time, sector, and firm age. For a subset of firms, depending on legal disclosure requirements, balance sheet data are also available.

2.1 From administrative data to the working sample

The working sample is constructed by applying several restrictions to the administrative data. First, the analysis focuses on young individuals aged 16 and over who entered the labor market as employees for the first time from 2005 onward. We exclude individuals who have ever held a seasonal contract over their observed career, as well as those whose first job was as an apprentice

[1] The data are sourced from forms that employers are required to submit for the payment of Social Security contributions on behalf of their employees.

or contractor.[2] This allows us to consider only two entry contract types: open-ended and fixed-term.[3] Information available prior to 2005 is used to exclude workers who were already present in the tax records, so as to only include individuals who are first employed from 2005 onward.

Using this sample of workers we construct a balanced yearly panel dataset. First of all, for each individual we select the main employment spell during each year, based on the highest paying contract or the longest one in case of equal earnings across different contracts. Specifically, we retain all the information related to the main employment spell and we sum all the earnings earned across multiple contracts during the year – so that the overall yearly earnings take into account earnings from each job, not only the main one. We then keep workers who can be observed in the dataset for at least ten years after the entry year and we only consider this 11-year period in the analysis. As a result, in our sample we keep workers who enter between 2005 and 2011.[4]

To ensure a minimum level of labor market attachment, workers who, at any point in their observed ten-year career after entry, begin a period of non-employment lasting for strictly more than two years are excluded. This data adjustment is required by the fact that periods of absence in the dataset may represent situations other than non-employment, such as engaging in educational activities or employment under contracts not covered by INPS. For instance, a worker could be absent if she holds a public-sector position, she is self-employed, or she works abroad.

For the same reason, only individuals whose initial employment spell – in terms of a given contract type and firm – lasts at least 26 working days within the first three months are kept. We also require workers to be employed during the entry year – where employment is defined as accumulating at least four full-time equivalent working weeks within the calendar year, whether from a single contract or multiple contracts.

[2] Apprentices have a hybrid status: they formal hold open-ended contracts but experience higher instability than regular workers on such contracts (Fenizia et al., 2024). Contractor jobs refer instead to positions where individuals provide services independently rather than as employees. These contractors are responsible for their own taxes and social contributions, and they typically work on a freelance basis, often under specific contracts that outline the terms of their work.

[3] We assess the robustness of this restriction by including apprentices and contractors in the sample. The results remain essentially unchanged and are available upon request.

[4] The available years after 2021 are used in the sample construction procedure but are not included in the analysis, so that we only consider workers who enter in 2011 or earlier.

Lastly, we impose additional restrictions to improve the identification of the first entry contract as an employee. In particular, we exclude individuals who enter the labor market relatively late compared to similar peers with the same education level, place of birth, and gender. This adjustment is intended to reduce the likelihood of misclassifying as new entrants individuals who may have had prior labor market experience not captured in the Social Security data – such as informal employment, time spent abroad or self-employment.

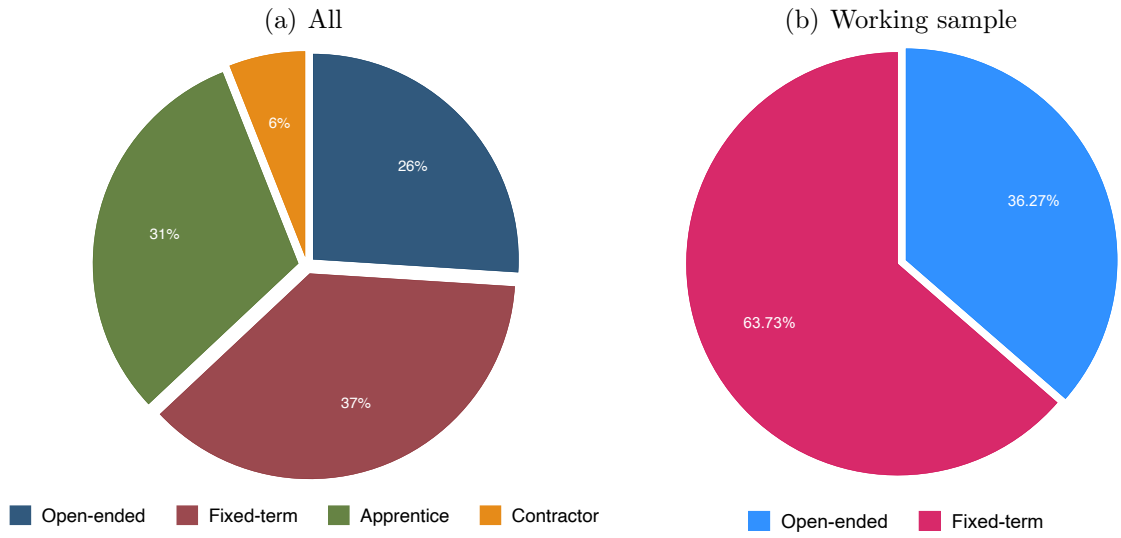
To implement this criterion, we remove workers whose age when they are first observed in the labor market is more than three years above the average entry age of their peer group, defined by education, location of birth, and gender. We do not base this restriction solely on the predicted formal entry age derived from the person's attained education level – i.e. the last year of study – since the average entry age in the general population may differ systematically from that formal age in our sample. Our working sample includes only regular private-sector employees and requires continuous labor market attachment for ten years following entry. In this context, low-educated individuals may systematically become employees later than the year when they complete their education – either because they begin their working career under alternative contractual arrangements, they experience extended periods of non-employment or they return to education.

The main variable in this study is the level of individual earnings. Specifically, the earnings measure used in the analysis is defined as the yearly sum of earnings, which include all pre-tax earnings – both regular and irregular in time – received under registered contracts.

2.2 Working sample characteristics

The working sample consists of 441,262 individuals, each observed for ten years following their first year in the labor market. On average, between 2005 and 2011 – excluding seasonal workers – around 63 percent of individuals tracked by the Italian Social Security Institute entered the labor market as either open-ended or fixed-term employees, the rest starting as either apprentices or contractors. In our final working sample, about 35 percent of workers began with an open-ended contract while the remaining 65 percent entered with a fixed-term contract (Figure 1).

Figure 1: Share of workers by entry labor market status



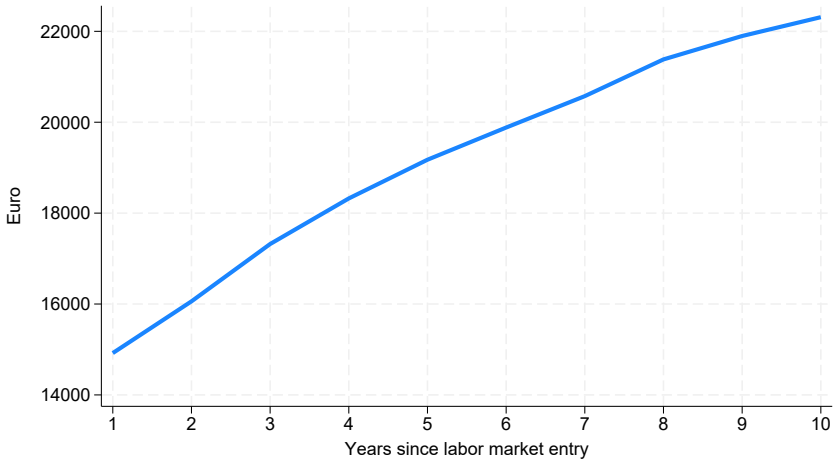
Note: The figures show the distribution of workers by entry labor market status from 2005 to 2011. Workers who entered the labor market with seasonal contracts are excluded.

Going beyond averages, the share of workers entering with a fixed-term contract has increased over time: it was around 59 percent in 2005 and it had risen to 70 percent by 2011. There is also substantial variation across individual characteristics. Perhaps surprisingly, the incidence of fixed-term entry is higher among more educated individuals – around 70 percent among those with a high school diploma or a university degree – compared to 54 percent among those with less than high school education. Geographically, the share is significantly higher in the North (76 percent) than in the South (39 percent), while it remains similar across genders.[5]

Once these and other differences in sample composition are taken into account, the key question becomes whether the type of entry contract has persistent consequences for workers' labor market performance, with a particular focus on their earnings. Figure 2 illustrates that earnings steadily increase with time spent in the labor market, rising from approximately €14,000 in the first year to over €20,000 by the tenth year. Does this growth trajectory vary depending on the type of contract held at entry?

[5]We report in Table 2 and Table 3 in Appendix B the distribution of workers by individual characteristics at labor market entry, as well as the distribution of workers across different entry firms. In both tables, we report the shares separately by entry contract type. Table 1 reports instead the share of fixed-term workers by individual characteristics, considering the cross-sectional distribution of contract types, at a quarterly frequency, over the observed career.

Figure 2: Average earnings by time in the labor market



Note: The figure reports average annual total gross earnings, adjusted for CPI, in 2022 euros. Year one refers to the first year after labor market entry.

3 Empirical strategy

To assess the empirical relationship between the entry contract type and labor market performance over the initial years of a worker's career, a key challenge lies in accounting for self-selection into initial contract types. The same underlying factors that influence early-career earnings growth may indeed also drive the choice or assignment of contract type at labor market entry. In other words, selection into different contract types at entry and subsequent earnings dynamics may be influenced by the same set of worker and firm characteristics, some of which may not be observable. Failing to account for this selection mechanism would lead to an incorrect assessment of the contribution of contract type to differences in labor market performance.

This issue has been addressed in the literature using various approaches. For example, Castellanos et al. (2023) implement an instrumental variable strategy that exploits variation in the aggregate incidence of open-ended contract hiring at the time of potential contract conversion in the worker's region of residence. Other studies have adopted structural approaches based on search and matching models, where contract choice is endogenous. In these models, firms face a trade-off between the speed of filling vacancies and the flexibility of dismissing workers (Berton and Garibaldi, 2012) or they evaluate production opportunities of different expected durations (Cahuc et al., 2016).

In our context, the focus on the initial contract type at the time of labor market entry offers a relatively clean setting for analysis. Workers entering their first job have no prior experience in the formal labor market and latent individual productivity is difficult to assess for both firms and workers. This reduces some of the possible concerns about endogenous selection, based for instance on unobserved past performance.

We adopt a twofold approach. First, we estimate a linear regression model (OLS), leveraging the richness of the administrative data to condition on a wide set of characteristics observed at the time of labor market entry. This helps account for observable heterogeneity that may influence both initial contract assignment and subsequent labor market performance. The linear model provides descriptive evidence on career progression patterns based on entry contract type.

We then turn to the main empirical strategy, implementing an instrumental variable analysis that exploits exogenous variation in the probability of starting a first job under a fixed-term or an open-ended contract. This quasi-experimental approach addresses potential unobserved confounders that may simultaneously influence both entry contract type and subsequent labor market outcomes, enabling a causal interpretation of the impact of the initial contract type.

3.1 Linear regression model

We start by estimating a flexible linear regression model that conditions on a rich set of characteristics observed at the time of entry, to examine whether the type of contract at labor market entry is associated with different earnings trajectories over the first ten years of a worker's career. This approach allows us to account for observable heterogeneity that may influence both the assignment to different contract types and subsequent earnings dynamics. In particular, we control for variation in demographic characteristics, educational background, geographic location, calendar time, and employer features at the time of the first job.

Specifically, we rely on the following regression framework:

$$y_{it} = \gamma_0 + \Sigma_s \gamma_1^s FT_{it_0} \cdot \mathbf{1}_{t=s} + \gamma_2 X_{it_0} + \gamma_3 H_{j_i t_0} + \gamma_4 T_{it} + \varepsilon_{it} \tag{1}$$

where y_{it} denotes total annual earnings for individual i after t years in the labor market, with $t = 1$ corresponding to the year immediately following entry. FT_{it_0} is a dummy variable for

whether the worker entered the labor market, at t_0, with a fixed-term contract. We interact this indicator with a full set of years of experience dummies to allow the influence of the entry contract to vary over time.

The vector X_{it_0} includes individual characteristics fixed at entry: gender, age, calendar month and year of entry, education, qualification at the first job (white-collar, manager, or blue-collar), province of the first job, and province of birth (there being 110 provinces). It also includes indicators for whether the worker migrated across provinces between birthplace and first job location, and whether they were born abroad. The vector H_{j,t_0} captures in turn characteristics of the entry firm j at entry year t_0: age, sector (classified into 21 industries), average wage, and size (both grouped into quintiles). Finally, T_{it} denotes fixed effects for labor market experience – i.e. years since entry.[6]

3.2 Instrumental Variable model

We then turn to an Instrumental Variable (IV) strategy. This quasi-experimental approach addresses concerns that unobserved factors – such as latent worker productivity – may jointly influence both the type of contract at entry and subsequent labor market performance. While the linear model controls for a rich set of observable characteristics, unobservable differences could still bias the estimates.

The instrument To address this issue, we exploit variation in the share of fixed-term hiring across locations and time. This is inspired by the approach in Arellano-Bover (2024), which estimates the impact of firm size at the first job, and Castellanos et al. (2023), which measures how different contract types – fixed-term and open-ended – affect workers' careers. These shares capture local labor market conditions related to the use of different contract types – such as province-specific demand fluctuations or policy changes – that are plausibly exogenous to unobserved individual characteristics, yet still influence the type of contract offered to labor market entrants.

In particular, we construct an instrument based on the share of workers hired with a fixed-

[6] We include additional robustness tests in Appendix A, where we evaluate the sensitivity of our results from the linear residual model to different sets of controls, as well as to the inclusion of entry contractor and entry apprentice workers in the sample.

term contract among all hires – both from non-employment and from other firms – observed in each province and calendar year.[7] We include all hires rather than focusing solely on new entrants, to plausibly reduce endogeneity concerns. To further address potential issues related to endogenous timing of labor market entry and sorting into local labor markets, we assign each worker the fixed-term hiring share corresponding to their province of birth and predicted year of labor market entry, rather than using their actual entry location and year.[8]

The predicted entry year is computed using the average age at labor market entry observed in our sample, stratified by education, province of birth, and gender. This approach avoids relying on formal schooling durations, which may not align systematically with actual entry into employment – particularly in a sample limited to private-sector workers with requirements for continuous labor market attachment over ten years.

Identifying assumption The key identifying assumption for the IV strategy is the exclusion restriction, which requires that the fixed-term hiring share in an individual's province of birth and predicted year of labor market entry affects future labor market outcomes solely through its impact on the type of initial contract. A potential threat to this assumption is that local labor market conditions – such as a province-specific demand shock – could affect both entry contract type and future earnings growth independently of the initial contract assignment.

We take several steps to mitigate this potential threat. First, as mentioned, by using the predicted year of entry rather than the actual year we reduce the risk that individual timing decisions in response to local shocks drive the observed variation. Second, by anchoring the instrument to the province of birth we avoid endogenous sorting based on unobserved labor market potential. Third, the fixed-term hiring share only reflects relative propensities within local hiring flows and not overall labor demand.

It is worth noting that a violation of the exclusion restriction – if the instrument partly reflects overall labor market strength – would likely bias the estimates upward, overstating the earnings growth of entry open-ended workers. A positive labor demand shock could simultaneously increase the likelihood of being hired on an open-ended contract and improve earnings

[7] We only consider hires on fixed-term and open-ended contracts, excluding apprentices and contractors.

[8] Since we compute the share of hires on fixed-term contracts by province of birth, we exclude from the IV analysis workers who were born abroad.

trajectories across the board. In such a case, the instrument would capture a higher share of open-ended hires in expanding labor markets, leading to an overestimation of the earnings growth gap between fixed-term and open-ended workers, favoring the latter.

Model specification To address endogeneity in the entry contract type variable, we follow the approach outlined by Wooldridge (2002), beginning with a logit regression where the binary regressor – the entry contract type dummy FT_{it_0} – is modeled as a function of our instrument and other relevant controls. This first prediction step is necessary because the binary nature of the endogenous regressor requires a non-linear specification to ensure that the predicted probabilities remain within the [0, 1] range. The fitted probabilities from this model are then used as instruments in the first stage of a two-stage least squares (2SLS) model, allowing for consistent estimation of the effect of entry contract type on earnings. Specifically, the prediction step relies on the province-year fixed-term hiring share, as previously discussed, alongside a set of exogenous individual characteristics – gender, education, year of birth, and province of birth:

$$\Pr(FT_{it_0} = 1 \mid FTShare_{p_i,\hat{t}_{i0}}, E_i) = \Lambda(\alpha_0 + \alpha_1 FTShare_{p_i,\hat{t}_{i0}} + \alpha_2 E_i) \tag{2}$$

where $\Lambda(\cdot)$ is the logistic function, $FTShare_{p_i\hat{t}_o}$ denotes the fixed-term hiring share in individual i's province of birth p_i and predicted labor market entry year $\hat{t}_{i0}$, and E_i denotes the set of exogenous individual controls.

We then use the fitted values $\widehat{P}_{it_0}$ from this logistic regression as an instrument in a conventional 2SLS framework. Specifically, we instrument the endogenous observed entry contract-type dummy with the exogenous predicted probability of entering the labor market with a fixed-term contract. To further capture the dynamic impact of the entry contract type over time, in the second stage we interact the fitted entry contract type dummy with a full set of years of experience dummies (multiplying the predicted probability IV as well). The first stage thus estimates:

$$FT_{it_0} = \delta_0 + \delta_1 \widehat{P}_{it_0} + \delta_2 E_i + \delta_3 T_{it} + \delta_4 C_t + \nu_{it} \tag{3}$$

The second stage uses the predicted interactions $\widehat{FT}_{it_0}$ from the linear first stage to estimate:

$$y_{it} = \beta_0 + \Sigma_s \beta_1^s \widehat{FT}_{it_0} \cdot \mathbf{1}_{t=s} + \beta_2 E_i + \beta_3 T_{it} + \beta_4 C_t + \eta_{it} \tag{4}$$

We include the same set of exogenous controls (E_i) as in the first prediction step and control for the business cycle by including a set of calendar year dummies (C_t). Since these dummies, when interacted with labor market experience, could potentially introduce endogenous effects related to the timing of labor market entry, we also produced results excluding the calendar dummies and the findings remained qualitatively unchanged. Note that we omit the firm characteristics included in the OLS model (Equation 1) due to their potential endogeneity.

This setup allows us to estimate how the effect of entering with a fixed-term contract evolves over time, isolating the exogenous component of the initial contract assignment and thereby enabling a causal interpretation of the findings.

4 Results

We now present our empirical results, starting with some descriptive evidence that takes an aggregate perspective and considers the total amount of earnings accumulated over the first ten years of a worker's career. Next, we investigate how earnings performance varies over time depending on the entry contract. To gain deeper insights into the sources of these differences, we then conduct a decomposition analysis that quantifies the contributions of the two main components of annual earnings: working time and wage per unit of time. Finally, we explore additional mechanisms that may explain the observed earnings patterns.

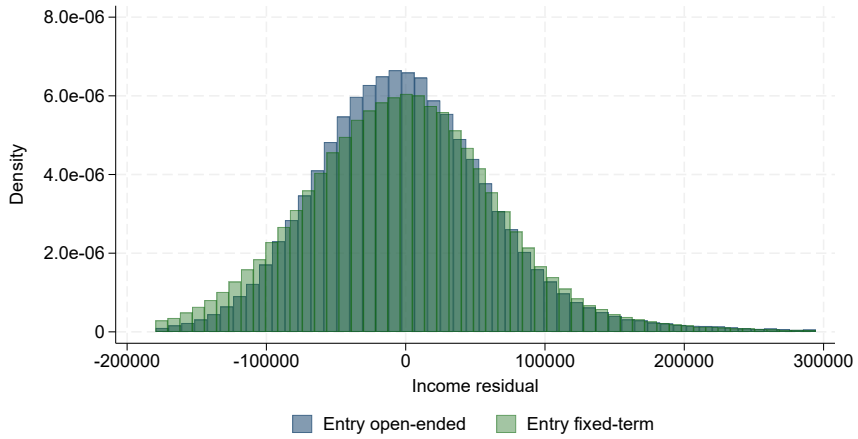
4.1 Introductory evidence

As a first step, we consider the total amount of earnings that individuals earned during their first ten years in the labor market, after entry. This aggregate perspective serves as a starting point to highlight broad disparities between individuals who began with each type of contract.

We regress the logarithm of cumulative earnings on the full set of covariates included in the specification outlined in Equation 1, comparing workers with similar entry conditions in the labor market.[9] Importantly, we first exclude the entry contract type from this regression in order to simply examine earnings differences net of observed entry characteristics. Specifically, we compare the distribution of the resulting residuals between individuals who entered on fixed-

[9] In this case, we use a logarithmic transformation since cumulative earnings are always strictly positive.

Figure 3: Linear residual of cumulative earnings over first ten years of career



Note: The figure reports the linear residual of the log cumulative earnings earned over the first ten years of each worker's career. The residual is computed using a linear OLS regression model including the set of covariates in Equation 1.

term and open-ended contracts. A positive residual indicates that a worker earned more than predicted based on their individual observable characteristics at entry, the year of entry, and the characteristics of their first employer.[10]

We find only minimal differences in total earnings across the two groups. On average, workers who began with a fixed-term contract earned slightly less than those who started with an open-ended one, though the gap is small (Figure 3). When we then include the entry contract dummy directly in the regression specification – rather than computing the residual – we find that its coefficient is slightly negative (-0.002 in log points) but statistically insignificant, with a 95% confidence interval ranging from -.005 to .001.

We then repeat the regression analysis separately by educational attainment – less than high school, high school, and university. This reveals substantial heterogeneity in cumulative earnings performance across groups. Specifically, low-educated workers who start their careers on a fixed-term contract earn 5.8 percent more on average over the first ten years than their counterparts who enter with an open-ended contract. Among high school graduates, the difference turns slightly negative: those who start on a fixed-term contract earn 2.8 percent less. The earnings gap becomes more pronounced instead for university graduates, who accumulate a 6.6 percent lower earnings over the first decade when entering on a fixed-term contract compared

[10]Note that we do not apply any time discount to earnings after the first year. Guvenen et al. (2022) provide several reasons why it is advisable not to discount these real earnings flows.

to those starting on an open-ended position.[11]

Figure 13 to Figure 15 in Appendix C display the cross-sectional distribution of residual earnings – controlling for individual characteristics at entry, year of entry, and characteristics of the first employer – by entry contract type and educational level, where these differences can be observed in terms of probability mass disparities along the relevant support.

To address potential endogeneity concerns, we estimate the relative differences in cumulative earnings using the IV model described in section 3, focusing directly on the different education groups. We find a similar gradient across education levels, although the magnitude of the differences between entry contract types is larger. Low-educated workers who start their careers on a fixed-term contract earn about 39 percent more on average over the first ten years than their counterparts who enter with an open-ended contract. Among high school graduates, the difference is slightly smaller, at 29 percent. By contrast, the earnings gap reverses for university graduates, who accumulate 114 percent lower earnings over the first decade if they start on a fixed-term contract compared to those entering on an open-ended position.

4.2 Earnings trajectories

Behind the performance in total accumulated earnings over the first ten years of a career, there may be systematic differences in how earnings evolve year by year, depending on the entry contract type. We next use the OLS and IV models presented in section 3 to measure observed differences in earnings trajectories over time.

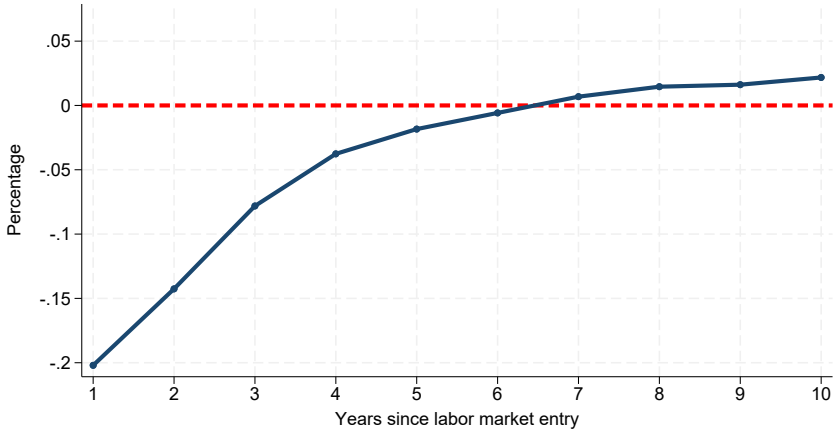
OLS model To investigate this, we start from the OLS specification detailed in Equation 1 and we report the time-varying coefficients on the fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract. This allows the descriptive estimates to be interpreted as the percentage difference in yearly earnings between fixed-term and open-ended entrants. Figure 4 shows the results.[12]

We find that in the first year after entry workers who started on fixed-term contracts are relatively penalized, earning on average 20 percent less than those who began with open-ended

[11] All entry contract dummy coefficients by education are statistically significant at the 0.05 level.

[12] We report in the first two columns of Table 4 in Appendix C the coefficients on the entry contract dummy interacted with years of experience, along with the corresponding robust standard errors.

Figure 4: Yearly total earnings - Relative difference: FT *vs* OE at entry (OLS)



Note: The figure reports the results from the linear OLS model in Equation 1. It displays the coefficients on the fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract.

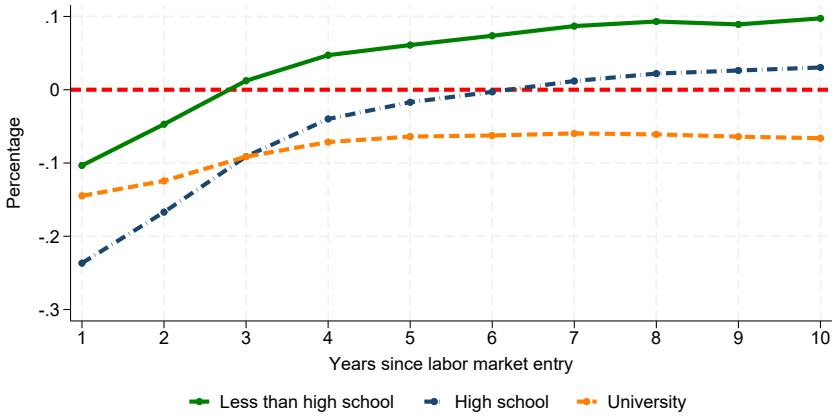
contracts. However, the gap narrows steadily and turns positive after the sixth year in the labor market. By the tenth year, fixed-term entrants earn about 2.2 percent more on average per year than their open-ended counterparts.

Since earnings grow over the life cycle for both groups, the small but positive earnings gap in the later years of the career weighs more heavily in the computation of total accumulated earnings. This helps explain the modest overall difference observed earlier, despite the sizable negative gap during the initial years.

In Figure 16 in Appendix C, we further illustrate that the reversing earnings gap is driven by a relatively steeper earnings growth profile for workers who entered on fixed-term contracts. Earnings increase over time for both entry groups, but at different rates. To highlight this, we re-estimate the model in Equation 1, saturating the regression with both entry contract dummies and removing the constant term. We then plot the coefficients on the fixed-term and open-ended entry dummies, which represent the portion of yearly earnings attributable to the entry contract type after controlling for other entry characteristics. In this setup, the portion of earnings associated with the fixed-term entry dummy grows by about 21.7 percent over time, compared to approximately 12 percent for the open-ended group.

Motivated by the preliminary results on accumulated earnings over the first ten years of workers' careers, we next examine differential earnings trajectories by education level – less than high school, high school, and university. In particular, we split the sample and estimate Equation 1

Figure 5: Yearly total earnings - Relative difference by education: FT *vs* OE at entry (OLS)



Note: The figure reports the results from the baseline model in Equation 1, by education groups. It displays the coefficients on the fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract.

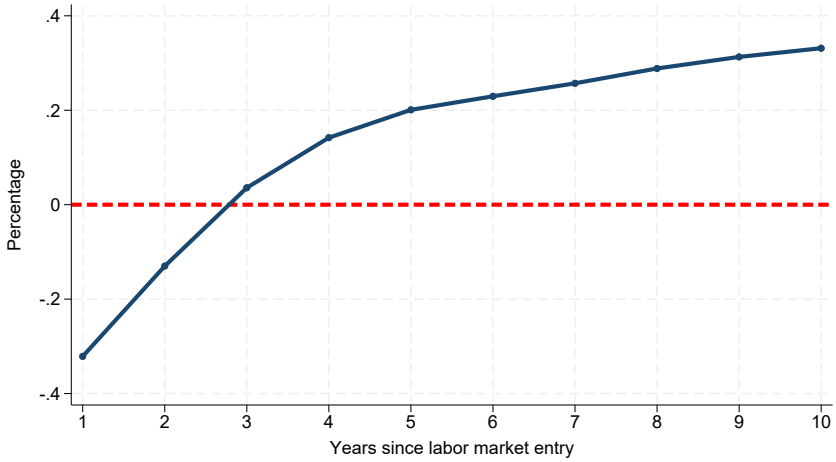
separately for each education group.

The results reveal a clear gradient pattern. The advantage associated with entering the labor market on a fixed-term contract declines as educational attainment rises (Figure 5). Workers with less than a high school diploma who start on a fixed-term contract obtain a positive earnings gap compared to their peers who began with an open-ended contract as early as the third year in the labor market. By the eighth year, they already earn about 10 percent more on average. At the other end of the spectrum, university graduates who enter on fixed-term contracts face a persistent negative earnings gap – of around 13 percent at entry, stabilizing at around 7 percent by the sixth year. Workers with a high school diploma fall in between, with a more balanced and converging trajectory: their gap is negative in the first year (22 percent) but turns slightly positive after the sixth year in the labor market, reaching about 3 percent in the tenth year.[13]

Quasi-experimental evidence We next presents the main results, based on the IV model outlined in section 3, which addresses potential endogeneity concerns and enables a causal interpretation of the impact of entry contract type on subsequent labor market outcomes. Specifically, we instrument the type of entry contract with the share of fixed-term hiring among all hires in an individual's province of birth and predicted year of labor market entry.

[13]We report in the last six columns of Table 4 in Appendix C the coefficients on the entry contract dummy interacted with years of experience, along with the corresponding robust standard errors.

Figure 6: Yearly total earnings - Relative difference: FT *vs* OE at entry (IV)



Note: The figure reports the results from the IV model in Equation 4. It displays the coefficients on the instrumented fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract.

While our linear regression approach accounts for a broad set of individual and firm-level characteristics at entry, unobserved factors may still influence both the selection into contract types and earnings trajectories over time. This being said, for such factors to substantially bias our findings they would have to imply that workers with stronger medium-term earnings prospects are systematically more likely to begin their careers on fixed-term contracts – a hypothesis that runs counter to the prevailing evidence of wage discounts in the empirical literature.

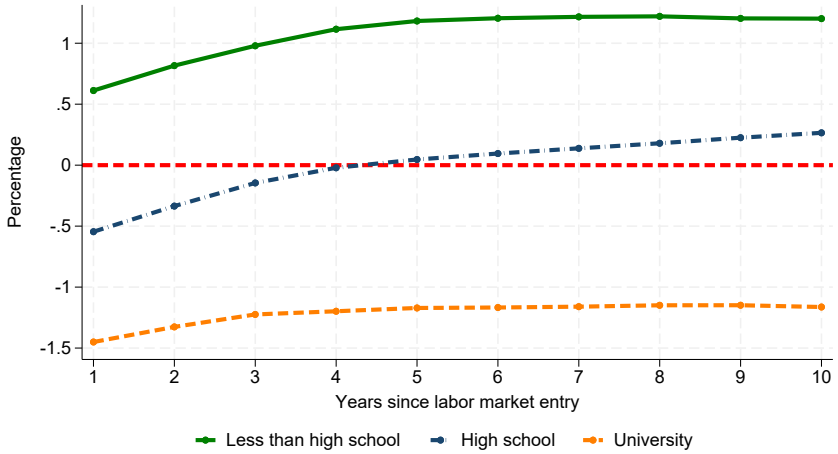
As a first step, Figure 17 in Appendix C reports the correlation between the observed and predicted share of entry fixed-term workers. Specifically, the predicted share is computed following Equation 2, using the share of fixed-term hiring among all workers in a given local labor market at a specific point in time. The figure shows an almost perfect correlation between the predicted and observed measures.[14]

As a further check on instrument strength, we report a Kleibergen-Paap rk Wald F-statistic of 270.2, well above the conventional threshold of 10, indicating a strong first stage. The underidentification test (Kleibergen-Paap rk LM statistic) also strongly rejects the null, with a value of 2694.4 ($p < 0.001$).[15]

[14]Pseudo $R^2 = 0.0928$. Sensitivity test = 70.45 percent of observations are correctly classified. We also estimate a Probit model (instead of the Logit model), which yields very similar results. Including calendar year dummies does not improve the model's performance.

[15]When we use the share of fixed-term hiring directly as the instrument, the model performs worse.

Figure 7: Yearly total earnings - Relative difference by education: FT *vs* OE at entry (IV)



Note: The figure reports the results from the IV model in Equation 4. It displays the coefficients on the instrumented fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract. The sample is split by educational attainment.

Figure 6 presents the IV estimates. We report the coefficient on the instrumented fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered on an open-ended contract. This scaling allows the results to be interpreted as percentage differences between the two groups.

The IV results confirm the pattern of earnings differences identified with the descriptive linear residual approach. In both cases, workers who entered on fixed-term contracts earn less initially but gradually experience convergence, eventually surpassing their open-ended counterparts. Under the IV model, however, accounting for endogenous selection, the initial earnings shortfall is more pronounced (about 30 percent at entry) and the later advantage is stronger (about 35 percent by the tenth year in the labor market), indicating a steeper path of catch-up and overtaking. Moreover, the reversal in the earnings gap occurs slightly earlier – around the start of the third year, compared to the sixth year in the baseline analysis. Overall, addressing endogeneity strengthens the core findings.

The linear residual results on differential earnings trajectories are also confirmed when splitting the sample by educational attainment. Specifically, the IV approach similarly shows that those who benefit most from entering the labor market with a fixed-term contract are workers with lower education levels. As in the OLS framework, the results display a clear gradient: the advantage of entering on a fixed-term contract declines as educational attainment rises

Kleibergen-Paap F statistic = 120.3; Kleibergen-Paap LM statistic = 1201.4.

(Figure 7). Workers with less than a high school diploma experience a substantial positive earnings gap since entry (around 50 percent) compared to peers who entered directly into an open-ended job. This finding is even stronger than under the OLS model, where the initial gap was negative (about -10 percent) in the first year after entry.[16]

In contrast, workers with a university degree who enter with a fixed-term contract experience a negative earnings gap, which also remains almost constant over the first ten years. Workers with a high school diploma experience instead a converging earnings gap, which is negative in the first year and turns positive after the fifth year in the labor market. Both of these dynamics are broadly consistent with the OLS results.

Overall, the results from the IV approach reinforce our descriptive findings, showing that earnings convergence between fixed-term and open-ended entrants is not driven by confounding factors related to entry contract type. The IV estimates point to even larger earnings gaps between the two groups.

Our findings show that although workers who begin with fixed-term contracts initially earn less, they experience faster earnings growth and, within a few years of labor market experience, surpass those who entered with open-ended contracts. This pattern of convergence and overtaking is more pronounced among less-educated workers, while those with a university degree who start on a fixed-term contract are unable to fully close the earnings gap over the first ten years of their careers. These results – which focus on new entrants and take a medium-run perspective – provide new empirical evidence that challenges the prevailing view in the academic literature, which typically associates fixed-term contracts with earnings disadvantages.

4.3 Decomposition analysis

We next investigate the mechanisms underlying this empirical pattern, starting with a decomposition of yearly earnings. Specifically, total earnings during the year can be broken down into two main components: the amount of time worked and the earnings per unit of time. Which of these two margins drives the steeper earnings profile observed for fixed-term entrants?

[16] We report in Table 5 in Appendix C the coefficients on the entry contract dummy interacted with years of experience, on aggregate and estimated separately by education group, along with the corresponding robust standard errors.

To explore this issue, we again rely on the two models specified in section 3 using as outcomes the number of full-time equivalent working weeks and the corresponding average weekly earnings during the year. In this context, the extensive employment margin is captured by the number of working weeks, which is equal to zero in years of full non-employment. In contrast, the earnings per full-time equivalent working week is conditioned on being positive, so that we exclude zero-earnings observations from this regression.

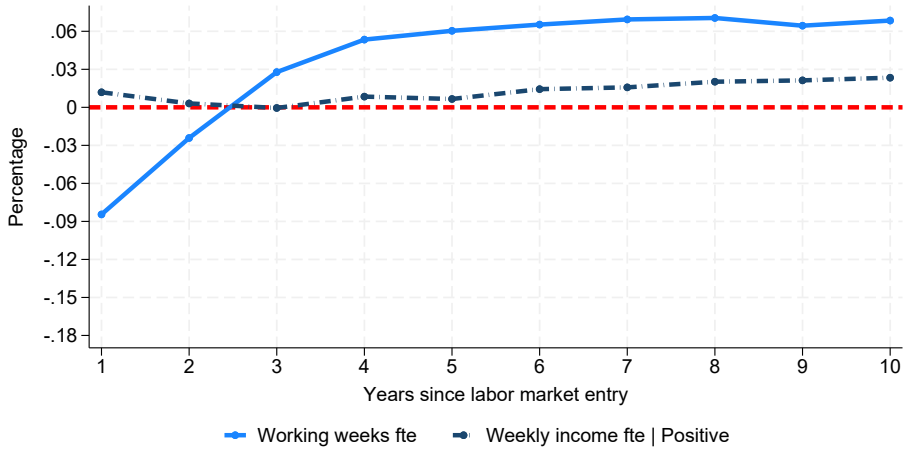
In this case, we again report the coefficient on the fixed-term entry dummy, scaled by the time-varying average number of full-time working weeks or the average full-time equivalent weekly earnings observed among workers who entered with an open-ended contract. This allows the results to be interpreted as percentage differences between the two entry groups. Since our results on earnings dynamics differ substantially by educational attainment, an aggregate decomposition analysis would be of limited value. For this reason, we directly proceed by performing the decomposition analysis of yearly earnings separately by education group.

OLS model We start by presenting the OLS descriptive evidence in Figure 8.[17] We find that differences in working time are the key factor driving the observed earnings trajectories. Specifically, in the early years of the labor market, workers entering on fixed-term contracts work fewer full-time equivalent weeks per year, with the gap ranging from -9 percent to -18 percent, depending on the education group. However, by the fourth year – even earlier for less educated workers – this gap reverses, and fixed-term entrants end up working more weeks than those who entered with open-ended contracts. In contrast, earnings per unit of time worked do not exhibit significant differences between the two groups.

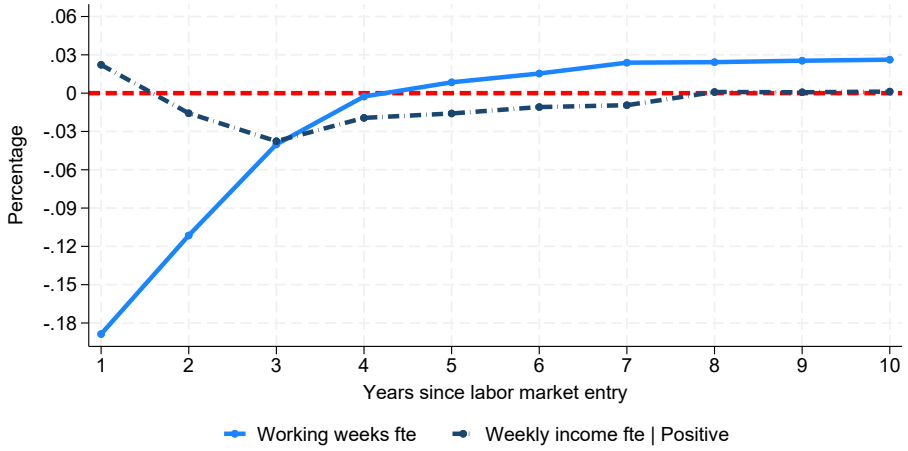
[17] We report in Table 6 and Table 7 in Appendix C the coefficients on the entry contract dummy interacted with years of experience, estimated separately by education group, along with the corresponding robust standard errors.

Figure 8: Decomposition of yearly earnings - Relative difference (OLS)

(a) Less than high school



(b) High school



(c) University

Note: The figures report a decomposition of yearly earnings using the OLS model in Equation 1, with the outcomes being full-time equivalent working weeks and weekly earnings full-time equivalent. Coefficients are percentage differences. They focus on workers with different educational attainment.

Interestingly, the gap in both components of yearly earnings displays a clear gradient by education group. Among workers with less than a high school diploma, the overtaking in full-time equivalent working weeks by fixed-term entrants is more pronounced, shifting from -9 percent in the first year after entry to +6 percent by year 10. Additionally, the gap in full-time equivalent weekly earnings is slightly positive in favor of fixed-term entrants, though the magnitude is relatively small.

For workers with a high school diploma, the gap in full-time equivalent working weeks for fixed-term entrants ranges from -18 percent in the first year after entry to +3 percent by year 10, indicating that while the gap closes, the overtaking is less pronounced. For these workers, differences in weekly earnings are instead negligible across the two entry contract types.

Lastly, for workers with a university degree, the gap in full-time equivalent working weeks closes by the fourth year in the labor market but does not show any overtaking pattern in the subsequent six years of observed careers. For these workers, the weekly wage gap remains negative for fixed-term entrants, holding steady at around -6 percent throughout the observed period.

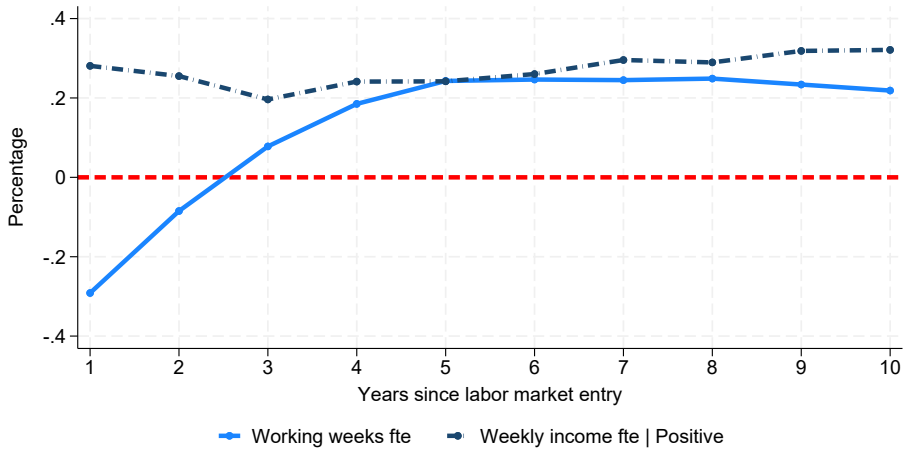
Quasi-experimental evidence We next present the evidence from the IV quasi-experimental setting. Overall, also in this case the results corroborate the OLS descriptive findings, with the main driver of the convergence in yearly earnings being the change in full-time equivalent working weeks over time. In contrast, differences in full-time equivalent weekly earnings between the two entry contract-type groups remain relatively stable over time (Figure 9).[18]

In the early years of the labor market, workers entering on fixed-term contracts work fewer full-time equivalent weeks per year, with the initial gap being in the range of -30 to -40 percent, across all education groups. However, by the third year, this gap reverses, and fixed-term entrants work more weeks than those who entered with open-ended contracts – except for the highly educated group, where the gap remains negative. In contrast, earnings per week worked differ between the two groups but remain relatively stable over time.

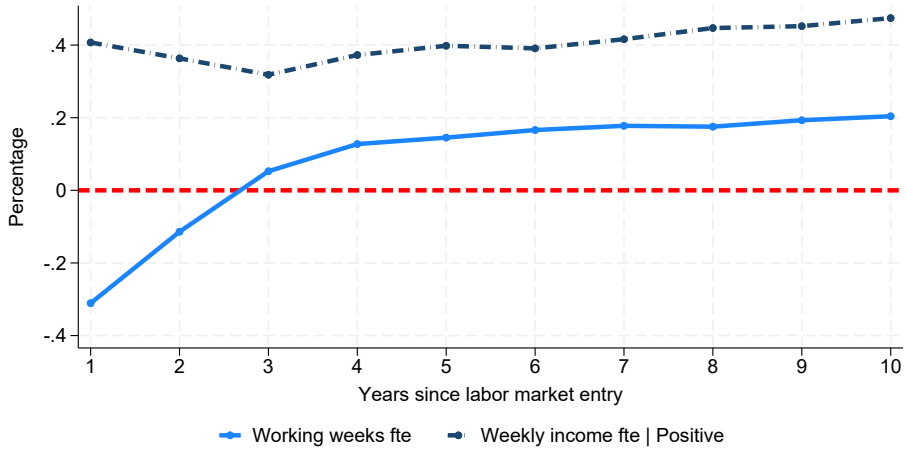
[18] We report in Table 8 and Table 9 in Appendix C the coefficients on the entry contract dummy interacted with years of experience, estimated separately by education group, along with the corresponding robust standard errors.

Figure 9: Decomposition of yearly earnings - Relative difference (IV)

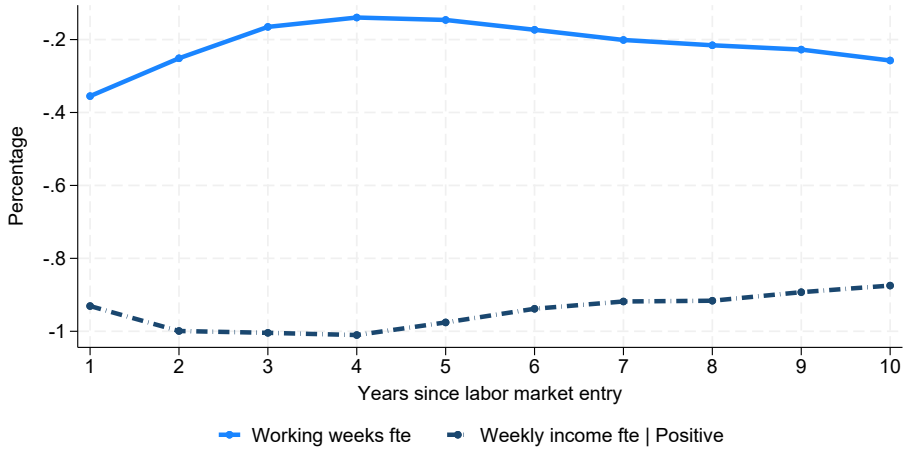
(a) Less than high school



(b) High school



(c) University



Note: The figures report a decomposition of yearly earnings using the IV model in Equation 4, with the outcomes being full-time equivalent working weeks and weekly earnings full-time equivalent. Coefficients are percentage differences. They focus on workers with different educational attainment.

Once again, the gap in both components of yearly earnings shows a clear gradient by education group. Among workers with less than a high school diploma, the overtaking in full-time equivalent working weeks by fixed-term entrants is slightly more pronounced than in the OLS specification, shifting from -40 percent in the first year after entry to +20 percent by year 10. The gap in full-time equivalent weekly earnings remains positive in favor of fixed-term entrants, with a magnitude of around +30 percent. For workers with a high school diploma, the gap in full-time equivalent working weeks for fixed-term versus open-ended entrants shows a slightly less steep trajectory, while differences in weekly earnings are positive and slightly larger compared to workers with less than a high school diploma. Lastly, for workers with a university degree, both the gap in full-time equivalent working weeks and the gap in full-time equivalent weekly earnings remain negative over the observed period and do not display any convergence.

Interestingly, once we account for endogenous selection into entry contract types, we find that among low- and medium-educated workers, weekly earnings per unit of time are actually higher for those who enter on a fixed-term rather than an open-ended contract. This pattern aligns with a compensating differential mechanism found for Italy (Albanese and Gallo, 2020).

In sum, both the OLS and IV decomposition analyses reveal a clear pattern: the gap in annual earnings between fixed-term and open-ended entrants is primarily driven by differences in full-time equivalent working weeks, while differences in earnings per unit of time worked remain relatively stable over time. The IV results confirm these dynamics but show somewhat larger gaps in both working weeks and weekly earnings.

4.4 Discussion and mechanisms

Our results reveal that workers who enter the labor market with fixed-term contracts initially experience a significant earnings disadvantage compared to those entering with open-ended contracts. However, we find that this earnings gap reverses over time, with fixed-term entrants outperforming open-ended entrants by the tenth year. This reversal is particularly pronounced for less-educated workers, with those holding less than a high school diploma experiencing a positive earnings gap, while university-educated workers see a persistent negative gap.

These results challenge the prevailing view in most of the academic literature, which typi-

cally associates fixed-term contracts with earnings disadvantages in a cross-sectional and static framework. We should note that there is mixed evidence on fixed-term contracts being stepping stones as opposed to dead ends for countries such as Austria, Denmark, Germany, Sweden, the Netherlands, the UK or the US, though not for Italy (Bentolila et al., 2020). Also, in a more dynamic setting, some studies have documented lifetime convergence in earnings between fixed-term and open-ended workers (Faccini, 2014). Our analysis extends this line of research by focusing specifically on new entrants to the labor market and by adopting a dynamic medium-run perspective. In this context, by tracking the careers of this group over time, we find not only convergence in earnings but also that fixed-term entrants eventually surpass their open-ended counterparts.

This earnings convergence and eventual overcoming process observed for fixed-term entrants is primarily driven by a reduction in non-employment time over the years. In contrast, the earnings per unit of time worked does not exhibit significant heterogeneity between the two groups over the course of their observed career trajectory. This sharp asymmetry is likely to stem in part from the prevalence of sectoral collective bargaining agreements in earnings setting in Italy. These agreements set earnings floors for all workers within a given industry and professional category, and they limit the ability of firms to adjust nominal earnings over time (Boeri et al. 2021). The gap between the actual earnings and the collective agreement earnings floor, i.e. the so-called *earnings cushion*, is typically small for young workers with little experience. Card and Cardoso (2022) find, for Portugal, average earnings cushions around zero for workers aged 18-20 and below 5% for workers younger than 25 years old. Adamopoulou and Villanueva (2022) study the metalworking industry in Italy and Spain, also finding lower earnings cushions for younger workers (though young is defined as being below 40 years old). These small earnings cushions severely limit the scope for earnings rate gains for young workers, either within a firm or by changing employers, and even if these workers are progressively increasing their productivity.

Mechanisms The empirical patterns observed so far can be understood in the context of the substantial uncertainty surrounding workers' productivity at the time of labor market entry. In such an environment, matches where the worker's ex-post productivity is low should dissolve quickly, while high-productivity matches should persist. Yet frictions often prevent this efficient

sorting. Open-ended contracts create inertia on both sides of the employment relationship: firms face firing costs that induce lower turnover and workers may prefer the stability of the job even when the match is suboptimal. As a result, early assignments to open-ended contracts may lock workers into low-quality matches that persist longer than they should.

Over time these matches may eventually dissolve, but by then the workers involved may have accumulated relatively little valuable experience. This increases their risk of non-employment later in their careers, as their job search is constrained by a narrow set of skills and a limited employment history.

In contrast, fixed-term contracts allow poor matches to be broken early, at lower cost, facilitating screening of workers by firms and learning on their own skills by young workers. This flexibility enables a faster reallocation process, giving workers more opportunities to explore different roles and accumulate a more diverse and relevant set of experiences. Over time, this reallocation mechanism leads to more productive and stable employment relationships. These dynamics might help to explain why fixed-term entrants, despite facing an initial disadvantage, eventually catch up with – and even surpass – open-ended entrants in terms of yearly earnings.

The gradient observed in the empirical findings by education level supports this interpretation. Low-educated workers typically face fewer outside options. As a result, low- and medium-educated workers who enter the labor market with an open-ended contract may choose to preserve suboptimal matches even when the position offers limited learning or mobility opportunities. This stickiness can lead to adverse consequences in the medium run, as remaining in low-quality matches hinders skill and experience accumulation, and limits upward mobility.

To provide further empirical support for the view that entering the labor market on a fixed-term contract can expose workers to a broader range of skills and experiences – with positive effects already emerging a few years into their careers – we measure the cumulative number of distinct firms, sectors, and provinces in which individuals have been employed during the first ten years of their working life. Specifically, we rely on the IV model outlined in section 3 to study the differential trajectories in the number of distinct firms, sectors, and provinces in which the workers have been employed over time.

Figure 18 to Figure 20 in Appendix C report the results. In the early years of their careers, entry fixed-term workers are employed in about 35 percent more distinct firms than those who started with an open-ended contract. This difference remains positive over the entire period, gradually declining to about 20 percent by year ten. A similar pattern is observed for the number of sectors – where a peak of 66 percent is reached – indicating that fixed-term entrants are not only more likely to change employers, but also more likely to switch working sectors, thereby diversifying their set of experiences and skills. Finally, we find evidence of greater geographic mobility: fixed-term entrants are around 25 percent more likely to change their working province during the first year in the labor market. The difference steadily declines thereafter, but it is still large even up to their tenth year of experience.

These patterns are suggestive on the types of skills being acquired at the beginning of workers' careers. Young labor market entrants do not know their own skills or their comparative advantage well, nor do they have any experience of how firms operate. As already mentioned, by forcing them to move across different firms and industries, fixed-term contracts are likely to reduce workers' benefits from investments in firm-specific human capital (Cabrales et al., 2017). On the other hand, such turnover can help young workers learn both about the best matches for their skills and about certain noncognitive or *soft* skills that matter for labor market performance.

Soft skills include items like teamwork, communication, and problem-solving, which are highly valued by employers. Deming (2017) finds for the US that both standard noncognitive skills and social skills have positive returns in terms of hourly wages.[19] While key noncognitive skills are innate or shaped in infancy and early youth (Cunha et al., 2010), some social skills may be learned on the job and acquired with cumulative labor market experience. Being exposed to different firms and industries can provide young workers entering on a fixed-term contract with a more rapidly growing stock, and a more varied portfolio, of social skills.

Further analysis of the initial labor paths of young labor market entrants, that focused on their early accumulation of skills measuring both cognitive and noncognitive skills could potentially

[19]Cognitive skills are proxied by standardized scores on the Armed Forces Qualifying Test (AFQT). Standard noncognitive skills are measured via surveys on how much control individuals feel they have over their lives and on self-esteem, as well as tests on the so-called *Big 5* personality factor called conscientiousness. Lastly, social capital is measured by surveys on sociability and participation in clubs and sports.

help account for the disparity between our results and those in the extant literature.

5 Conclusions

This paper examines the medium-run implications of entering the labor market through a fixed-term contract rather than an open-ended one, focusing on earnings and employment trajectories over the first ten years of workers' careers. We find that although fixed-term entrants initially earn lower annual earnings, they quickly catch up and eventually surpass their open-ended counterparts. This convergence is primarily driven by a gradual increase in the number of full-time working weeks during the year, while earnings differences per unit of time worked remain negligible. These patterns are particularly pronounced among workers with lower education levels, who appear to benefit more from entering the labor market on a fixed-term contract.

Further work is needed to ascertain the robustness of our results. If confirmed, our findings could have potentially important policy implications. Contrary to the widespread perception that fixed-term contracts trap workers in unstable positions, our evidence suggests that they can serve as an effective sorting mechanism, providing early exposure to a broader range of firms, sectors, and geographic areas. This variety of experiences may accelerate learning and lead to better job matching in the medium run. Policies aimed at improving the labor market performance of young workers should not focus solely on reducing segmentation. Instead, efforts could be directed toward ensuring that fixed-term positions function as genuine stepping stones – by strengthening job search support, expanding access to training, and promoting transparent pathways toward stable employment – rather than becoming endpoints in precarious careers.

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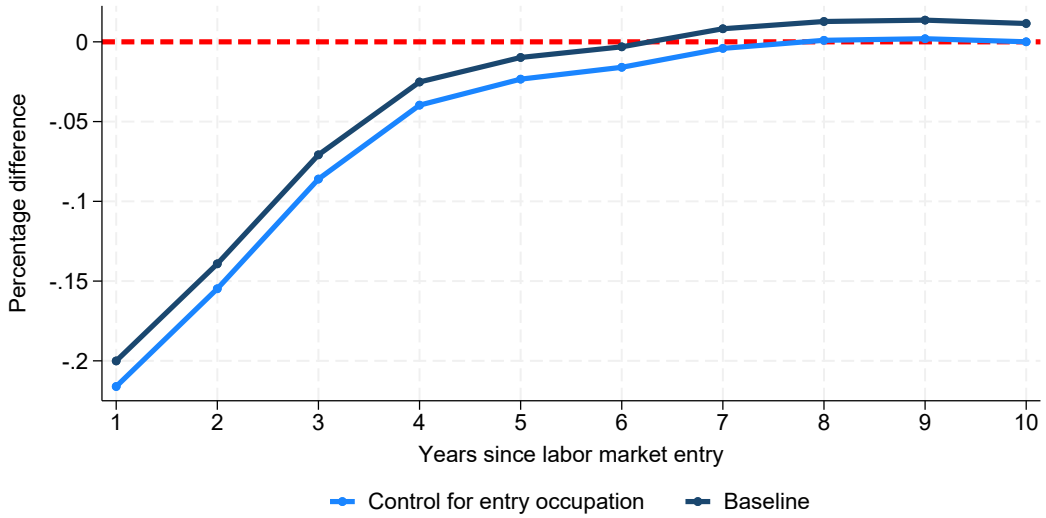
Figure 10: Yearly total earnings - Relative difference: FT *vs* OE at entry (OLS)



Note: The figure reports the results from the linear OLS model in Equation 1, as well as from the specification expanded to include an indicator for occupation at entry. It displays the coefficients on the fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract. The sample is restricted to new labor market entrants between 2009 and 2011.

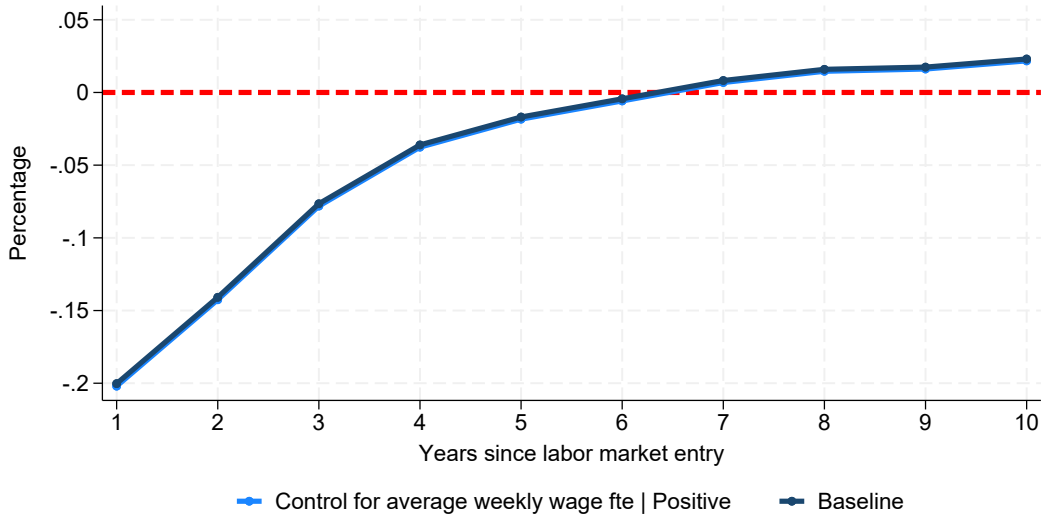
Appendix

A Alternative specifications of the linear OLS model

In this section we assess the sensitivity of our results to different sets of controls. We examine in particular the role of additional controls omitted from the main specification of the linear OLS model (Equation 1). We first test whether our results are sensitive to including the worker's occupation at entry. This variable is excluded from the main analysis because it is only available starting in 2009. To assess the robustness of our findings, we therefore restrict the sample to new labor market entrants between 2009 and 2011.

Within this limited sample, we estimate the baseline regression model and compare it to a specification that additionally controls for occupation in the first job. As in the main analysis, we report the coefficient on the fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract – so that the results can be interpreted as a percentage difference between the two groups. Figure 10 presents the results: the two lines are nearly indistinguishable, indicating that controlling for occupation at entry has little impact on the estimates. That these lines differ from the main results is expected, as they are based on a different and more restrictive sample of workers.

Figure 11: Yearly total earnings - Relative difference: FT *vs* OE at entry (OLS)



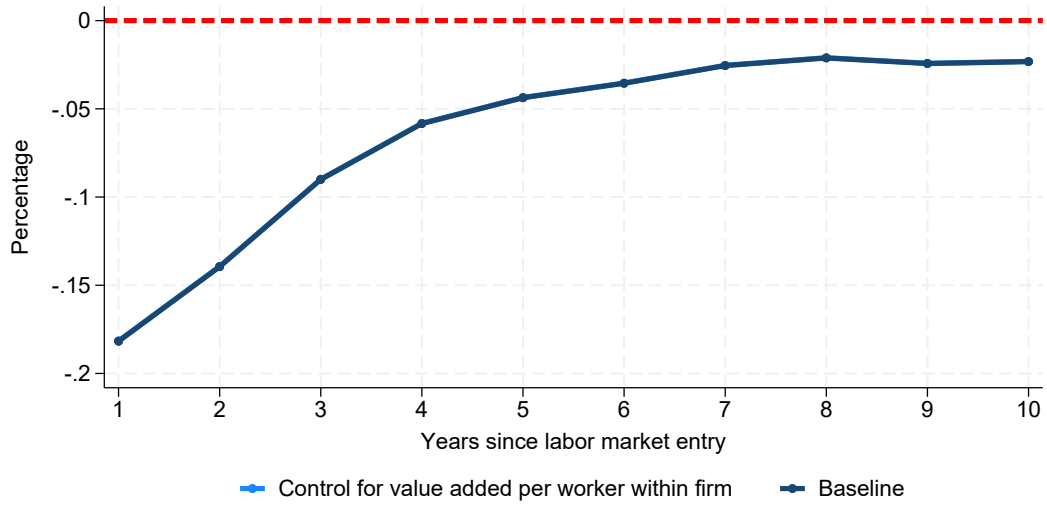
Note: The figure reports the results from the linear OLS model in Equation 1, as well as from the specification expanded to include an indicator for 10-year average weekly earnings full-time equivalent, in the form of quintile groups. It displays the coefficients on the fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract.

We next test whether including in the specification a broad measure of individual productivity affects the results. As a proxy for latent productivity, we use the average full-time equivalent weekly wage earned over the first ten years of a worker's career. This wage measure is not influenced by periods of non-employment and is intended to reflect a (noisy) indicator of the worker's underlying productivity.

We then estimate the baseline regression and compare it to a specification that includes controls for this productivity measure, grouped into quintiles. As in the main analysis, we report the coefficient on the fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract. Figure 11 presents the results. Once again, the two lines are nearly indistinguishable, suggesting that controlling for a proxy of latent worker productivity does not meaningfully alter the findings.

Lastly, we turn to the set of controls related to the entry firm. In the main specification, we proxy firm productivity using the average wage paid within the company. However, for a subset of firms – depending on legal disclosure requirements – we also have access to balance sheet data. In this robustness test, we replace the wage-based productivity measure with value added per worker, calculated using a full-time equivalent measure of the workforce and grouped into quintiles.

Figure 12: Yearly total earnings - Relative difference: FT *vs* OE at entry (OLS)



Note: The figure reports the results from the linear OLS model in Equation 1, as well as from the specification where we replace the earnings-based firm productivity measure with value added per worker. It displays the coefficients on the fixed-term entry dummy, scaled by the time-varying average earnings of workers who entered with an open-ended contract. The sample is restricted to firms with available balance sheet data.

As in the previous cases, we compare the results from this specification to those from the baseline framework. Figure 12 presents the results. In this case, the two lines are perfectly overlapping, indicating that controlling for average wage within the firm performs just as well as using value added per worker. As before, differences relative to the main results are due to a restricted sample, since we only consider firms with available balance sheet data.

Overall, the results of these robustness tests reinforce the validity of our baseline findings. Controlling for additional information on entry occupation, latent worker productivity, or alternative measures of firm performance does not substantially alter the estimated differences in earnings dynamics between workers who begin their careers on fixed-term versus open-ended contracts. This consistency across specifications lends further credibility to the interpretation that the observed patterns are not driven by omitted variables or sample selection.

B Sample characteristics

In this section, we describe the main characteristics of our sample of workers. To start with, Table 1 reports the distribution of fixed-term workers by individual characteristics, considering the cross-sectional distribution of contract types across all observed quarters. The overall share of fixed-term workers is relatively similar by gender, with 22.8 percent among women

Table 1: Share of fixed-term workers by individual characteristics

	%
Gender	
Female	22.8
Male	22.4
Education	
Less than high-school	23.1
High-school	24.2
University	18.8
Macro-region (IT)	
North-west	23.8
North-east	24.6
Centre	23.0
South	19.0
Country of birth	
Italy	22.8
Abroad	22.0

Note: The table reports the share of fixed-term workers by individual characteristics. It considers the cross-sectional distribution of contract types, at a quarterly frequency, over the observed career.

and 22.4 percent among men. Educational attainment shows more variation, with fixed-term contracts being more frequent among individuals with high-school education (24.2 percent) and less prevalent among university graduates (18.8 percent). Geographical differences are also notable: the share of fixed-term workers is highest in the North-East (24.6 percent) and lowest in the South (19.0 percent). Finally, the distribution by country of birth shows a slightly higher incidence of fixed-term contracts among Italian-born individuals (22.8 percent) compared to those born abroad (22.0 percent). Overall, the table highlights moderate heterogeneity in fixed-term contract incidence across observable characteristics, with education and geographic region exhibiting the largest differences.

We then focus on differences in sample characteristics based on the entry contract type. The sample displays statistically significant differences in entry characteristics between the two groups, as confirmed by chi-squared tests for categorical variables (Table 2). Fixed-term entrants are more likely to be female (42.6 vs. 37.7 percent), to hold a university degree (24.5 vs. 18.6 percent), and to begin their careers in white-collar positions (45.6 vs. 34.4 percent).

Table 2: Sample characteristics by entry contract type (%)

	Entry Open-ended	Entry Fixed-term
Gender		
Female	37.7	42.6
Male	62.3	57.4
Education		
Less than high-school	49.6	33.6
High-school	31.8	41.9
University	18.6	24.5
Entry qualification		
Blue-collar	65.5	54.3
White collar	34.4	45.6
Manager	0.1	0.1
Entry macro-region (IT)		
North-west	26.0	48.6
North-east	13.7	18.8
Centre	17.7	17.2
South	42.6	15.4
Country of birth		
Italy	67.6	80.6
Abroad	32.4	19.4

Note: The table reports the characteristics of workers, by entry contract type.

They are also more geographically concentrated in the North-west of the country (48.6 vs. 26.0 percent) and less likely to enter the labor market in the South (15.4 vs. 42.6 percent). Finally, fixed-term entrants are more likely to be Italian-born (80.6 vs. 67.6 percent).

The average entry age for workers on an open-ended contract is 23.2, compared to 22.4 for those who entered with a fixed-term contract.

Lastly, in terms of sector at entry, the largest share of fixed-term entrants (29.5 percent) start in the renting and service sector, while the largest share of open-ended entrants (19.9 percent) are employed in the commerce sector.

We further examine the characteristics of the firms where workers enter the labor market, conditional on entry contract type. As shown in Table 3, fixed-term entrants are more likely to start in firms with higher average wages, with 25.4 percent in the top wage quintile compared to

Table 3: Share of workers in different entry firms, by entry contract type (%)

	Entry Open-Ended	**Entry Fixed-Term**
Average wage within firm		
First Quintile	30.9	10.7
Second Quintile	21.1	18.6
Third Quintile	16.8	22.5
Fourth Quintile	16.2	22.8
Fifth Quintile	15.0	25.4
Firm size		
First Quintile	35.3	10.2
Second Quintile	27.9	14.8
Third Quintile	18.9	20.3
Fourth Quintile	13.9	24.4
Fifth Quintile	4.0	30.3
Firm age		
First Quintile	17.7	8.9
Second Quintile	32.5	23.3
Third Quintile	13.8	16.8
Fourth Quintile	18.8	26.6
Fifth Quintile	17.2	22.4

Note: The table reports the share of workers in different entry firms, by entry contract type.

15.0 percent among open-ended entrants. Similarly, they are more likely to enter larger firms: 30.3 percent of fixed-term entrants work in the largest size quintile, versus only 4.0 percent of open-ended entrants. In terms of firm age, fixed-term entrants are slightly more concentrated in older firms, with 22.4 percent in the oldest quintile compared to 17.2 percent among open-ended entrants. Chi-squared tests confirm that these differences in firm characteristics are statistically significant.

C Additional figures and tables

Table 4: Yearly earnings (OLS)

	All sample		**< High school**		**High school**		**University**	
	Coefficient	SE	Coefficient	SE	Coefficient	SE	Coefficient	SE
Period 1	-2948.9	25.5	-1215.1	33.3	-3246.4	41.2	-3431.9	69.0
Period 2	-2173.1	27.1	-563.4	35.4	-2394.7	44.0	-3176.5	74.5
Period 3	-1243.4	27.9	152.5	36.4	-1364.4	44.8	-2463.5	77.8
Period 4	-621.4	28.8	605.9	37.2	-617.0	45.9	-2009.7	81.6
Period 5	-315.3	29.9	811.1	38.1	-275.2	47.2	-1875.4	87.1
Period 6	-103.1	31.6	1009.1	39.4	-47.4	49.2	-1903.8	93.9
Period 7	125.2	33.5	1224.1	40.6	202.2	51.4	-1884.8	102.4
Period 8	277.3	35.7	1361.8	42.0	388.9	53.8	-2012.1	111.8
Period 9	314.8	38.1	1330.4	43.3	470.9	56.4	-2201.3	122.5
Period 10	431.5	40.8	1465.1	44.7	551.7	59.2	-2349.1	134.7
Observations	4,359,100		1,722,200		1,663,010		973,890	

Note: The table reports the coefficient of the entry fixed-term dummy — and the corresponding robust standard error — in the OLS model in Equation 1.

Table 5: Yearly earnings (IV)

	All sample		**< High school**		**High school**		**University**	
	Coefficient	SE	Coefficient	SE	Coefficient	SE	Coefficient	SE
Period 1	-5336.1	479.4	7791.4	1829.5	-7741.5	954.4	-35634.4	5986.3
Period 2	-2280.9	478.5	10616.4	1826.7	-4993.3	952.7	-35240.3	5986.2
Period 3	655.6	478.9	12984.6	1826.8	-2259.8	952.6	-34385.9	5998.7
Period 4	2683.5	480.0	15091.9	1830.7	-340.2	953.5	-35080.6	6017.3
Period 5	3933.5	480.6	16389.9	1833.2	771.1	953.8	-35684.6	6023.3
Period 6	4638.9	481.0	17120.7	1833.9	1619.7	953.5	-36942.3	6019.4
Period 7	5350.9	481.8	17697.1	1834.2	2406.4	954.1	-38097.4	6010.1
Period 8	6232.5	483.0	18310.8	1835.7	3237.9	954.7	-39385.1	6021.4
Period 9	6975.3	484.9	18556.8	1837.9	4155.9	955.8	-40981.5	6054.7
Period 10	7550.9	487.3	18841.1	1840.8	4944.2	957.3	-42938.8	6089.1
Observations	2316930		548,500		1,076,810		691,620	

Note: The table reports the coefficient of the entry fixed-term dummy – and the corresponding robust standard error – in the IV model in Equation 4.

Table 6: Working weeks full-time equivalent (OLS)

	< **High school**		**High school**		**University**	
	Coefficient	SE	Coefficient	SE	Coefficient	SE
Period 1	-2.7	0.1	-6.5	0.9	-3.9	1.0
Period 2	-0.8	0.1	-3.9	0.9	-2.7	1.0
Period 3	0-9	0.1	-1.4	0.9	-1.1	1.0
Period 4	1.7	0.1	-0.1	0.9	-0.9	1.0
Period 5	1.9	0.1	0.3	0.9	0.1	1.0
Period 6	2.2	0.1	0.6	0.9	0.1	0.9
Period 7	2.4	0.1	0.9	0.9	0.2	0.9
Period 8	2.4	0.1	0.9	0.8	0.1	0.9
Period 9	2.2	0.1	0.9	0.8	-0.1	0.9
Period 10	2.4	0.1	1.0	0.8	-0.1	0.9
Observations	1,722,200		1,663,010		973,890	

Note: The table reports the coefficient of the entry fixed-term dummy – and the corresponding robust standard error – in the OLS model in Equation 1.

Table 7: Weekly earnings full-time equivalent (OLS)

	< **High school**		**High school**		**University**	
	Coefficient	SE	Coefficient	SE	Coefficient	SE
Period 1	4.9	1.6	9.6	1.7	-20.9	1.9
Period 2	1.3	1.5	-6.9	1.6	-33.2	1.8
Period 3	-0.2	1.4	-17.1	1.6	-37.2	1.8
Period 4	3.6	1.3	-8.9	1.5	-40.8	1.9
Period 5	2.8	1.3	-7.4	1.5	-39.8	1.9
Period 6	6.2	1.3	-5.1	1.5	-37.6	2.0
Period 7	6.9	1.2	-4.5	1.5	-38.4	2.1
Period 8	8.9	1.2	0.4	1.5	-40.1	2.3
Period 9	9.5	1.3	0.3	1.5	-42.7	2.5
Period 10	10.5	1.3	0.6	1.5	-43.6	2.7
Observations	1,632,845		1,588,705		948,374	

Note: The table reports the coefficient of the entry fixed-term dummy – and the corresponding robust standard error – in the OLS model in Equation 1.

Table 8: Working weeks full-time equivalent (IV)

	< **High school**		**High school**		**University**	
	Coefficient	SE	Coefficient	SE	Coefficient	SE
Period 1	-10.0	3.1	-10.9	1.7	-15.8	3.7
Period 2	-2.9	3.0	-4.1	1.6	-11.3	3.8
Period 3	2.6	3.0	1.9	1.7	-7.5	3.8
Period 4	6.3	3.0	4.6	1.7	-6.3	3.8
Period 5	8.3	3.0	5.4	1.7	-6.6	3.8
Period 6	8.5	3.0	6.2	1.6	-7.9	3.8
Period 7	8.5	3.0	6.6	1.7	-9.1	3.8
Period 8	8.7	3.0	6.6	1.7	-9.8	3.8
Period 9	8.3	3.1	7.3	1.7	-10.3	3.7
Period 10	7.8	3.0	7.6	1.7	-11.6	3.7
Observations	548,500		1,076,810		691,620	

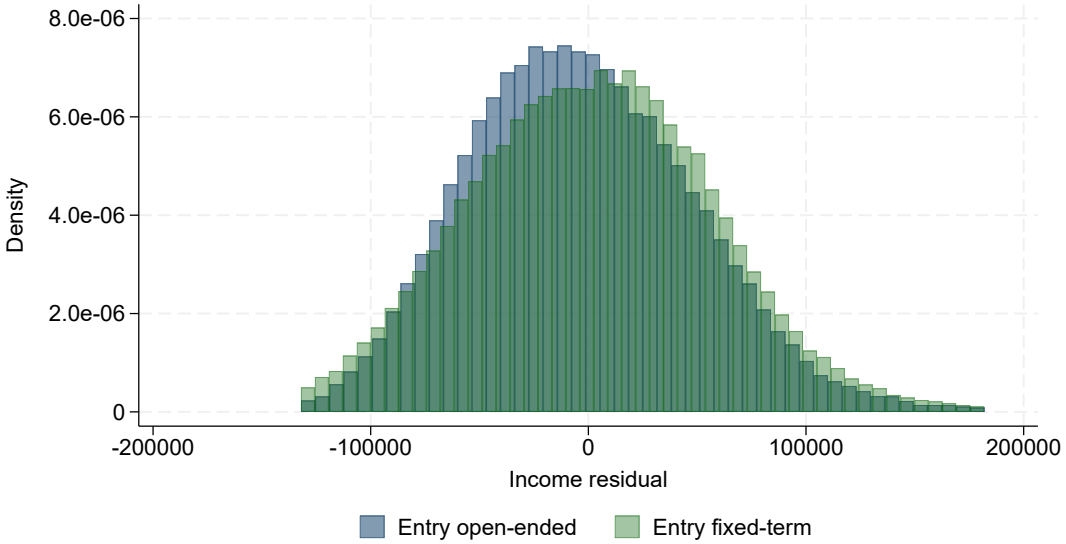
Note: The table reports the coefficient of the entry fixed-term dummy – and the corresponding robust standard error – in the IV model in Equation 4.

Table 9: Weekly earnings full-time equivalent (IV)

	< **High school**		**High school**		**University**	
	Coefficient	SE	Coefficient	SE	Coefficient	SE
Period 1	115.5	51.4	176.8	29.3	-521.0	106.4
Period 2	104.5	51.5	161.4	29.4	-594.6	106.1
Period 3	82.7	52.3	145.6	29.6	-624.7	107.8
Period 4	102.2	52.7	171.7	29.9	-652.4	108.3
Period 5	104.1	52.4	185.9	29.9	-648.9	108.4
Period 6	111.8	52.3	185.4	29.8	-646.9	108.5
Period 7	129.5	52.1	200.5	29.8	-654.5	107.9
Period 8	128.6	52.3	218.4	29.8	-681.1	108.5
Period 9	143.2	52.2	224.6	29.7	-688.8	108.5
Period 10	146.4	52.2	239.3	29.7	-703.4	109.4
Observations	421,077		1,031,984		676,273	

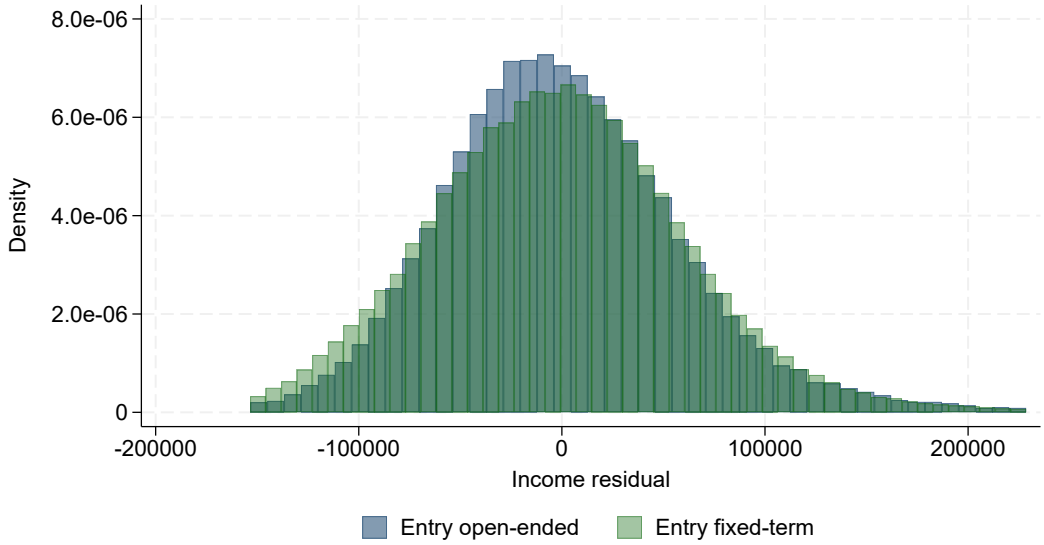
Note: The table reports the coefficient of the entry fixed-term dummy – and the corresponding robust standard error – in the IV model in Equation 4.

Figure 13: Linear residual of cumulative earnings over first 10 years of career - Less than high school



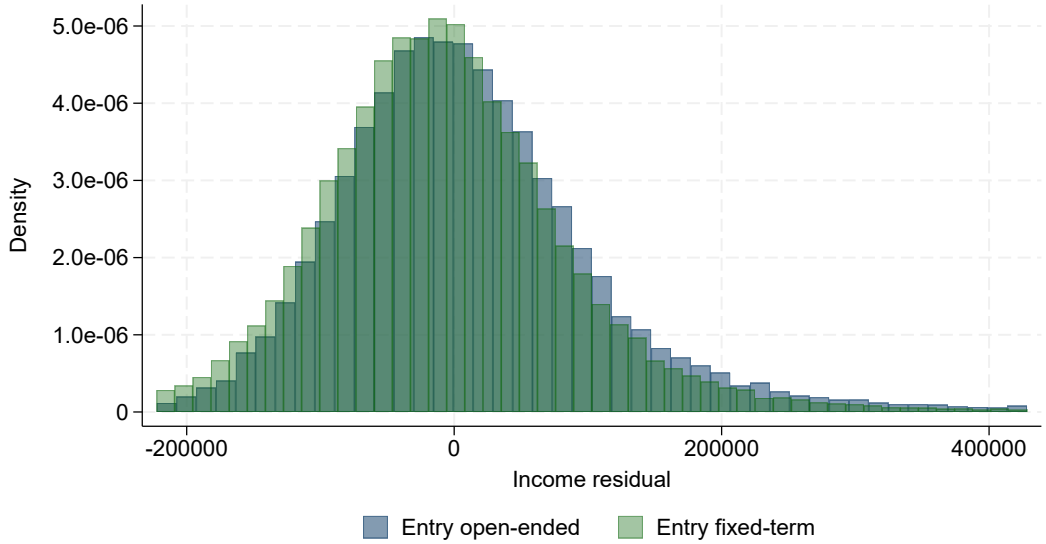
Note: The figure reports the linear residual of the log cumulative earnings earned over the first ten years of each worker's career. The residual is computed using a linear OLS regression model including the set of covariates in Equation 1. The sample is restricted to low-educated individuals.

Figure 14: Linear residual of cumulative over first 10 years of career - High school



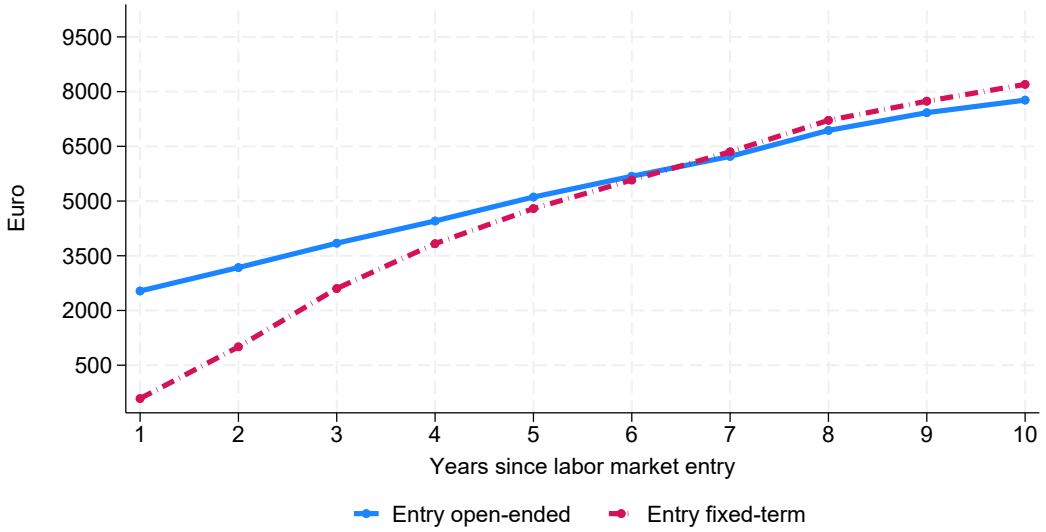
Note: The figure reports the linear residual of the log cumulative earnings earned over the first ten years of each worker's career. The residual is computed using a linear OLS regression model including the set of covariates in Equation 1. The sample is restricted to individuals with a high school diploma.

Figure 15: Linear residual of cumulative over first 10 years of career - University



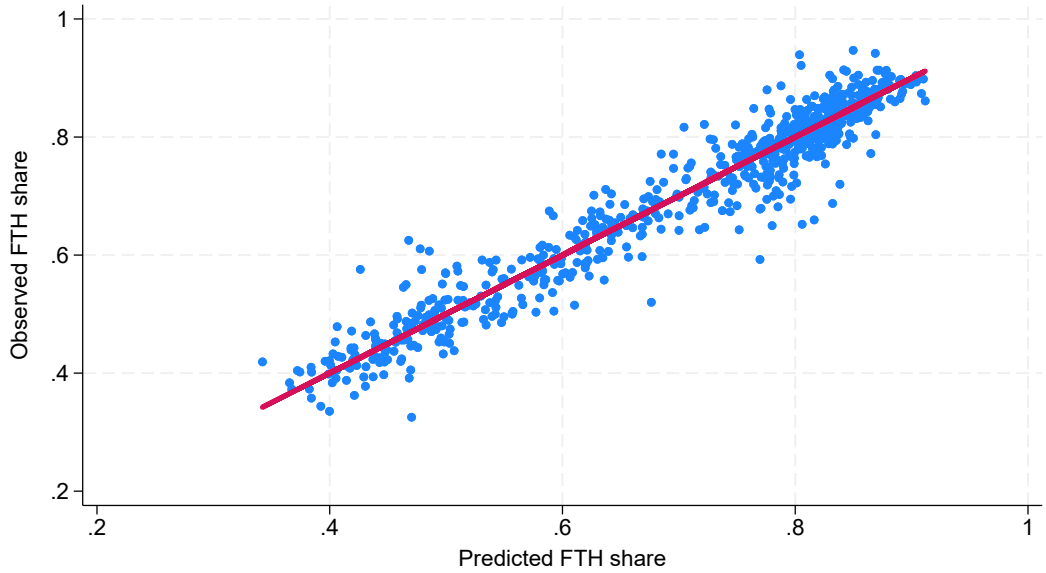
Note: The figure reports the linear residual of the log cumulative earnings earned over the first ten years of each worker's career. The residual is computed using a linear OLS regression model including the set of covariates in Equation 1. The sample is restricted to individuals with a university degree.

Figure 16: Yearly total earnings - Relative difference: FT *vs* OE at entry (OLS)



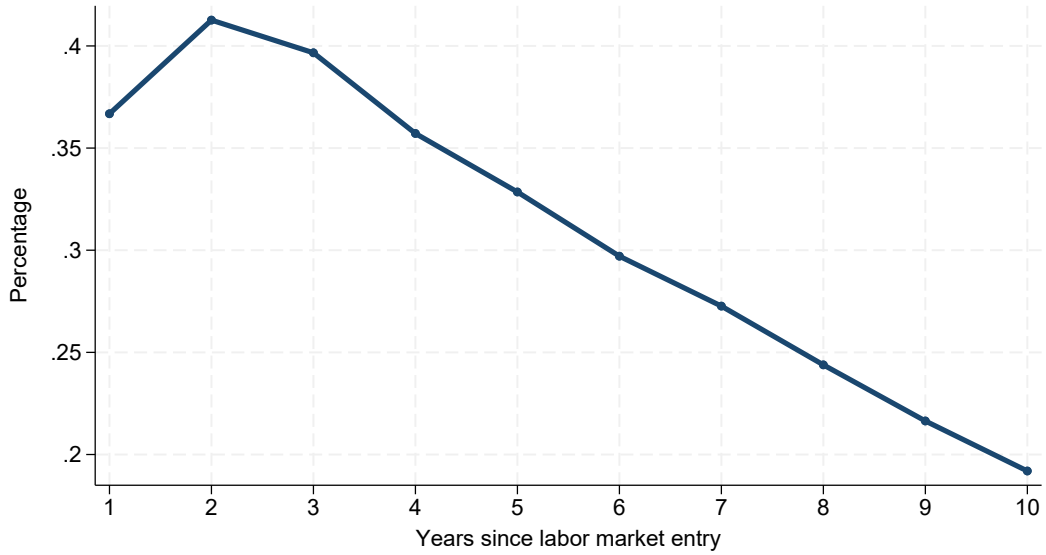
Note: The figure reports results from the linear OLS model in Equation 1, estimated with a saturated specification excluding the constant term. It displays the coefficients on the fixed-term and open-ended entry dummies, capturing the portion of yearly earnings associated with each entry contract type.

Figure 17: Predicted and actual share of open-ended entrants by province and calendar year



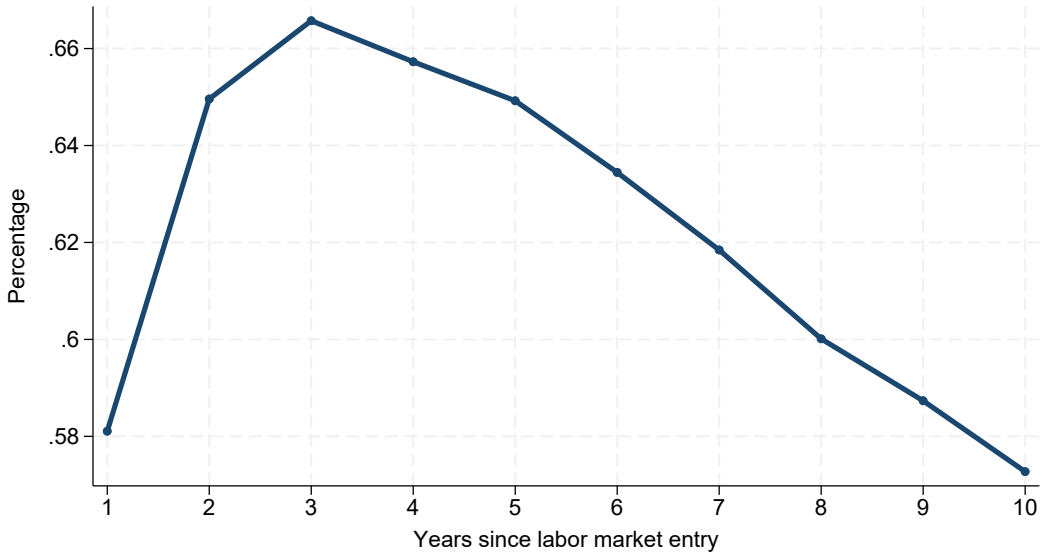
Note: The figure reports the correlation between the observed and predicted share of entry open-ended workers, based on our instrument. The predicted share is computed according to Equation 2.

Figure 18: Cumulative number of firms - Relative difference: FT *vs* OE at entry (IV)



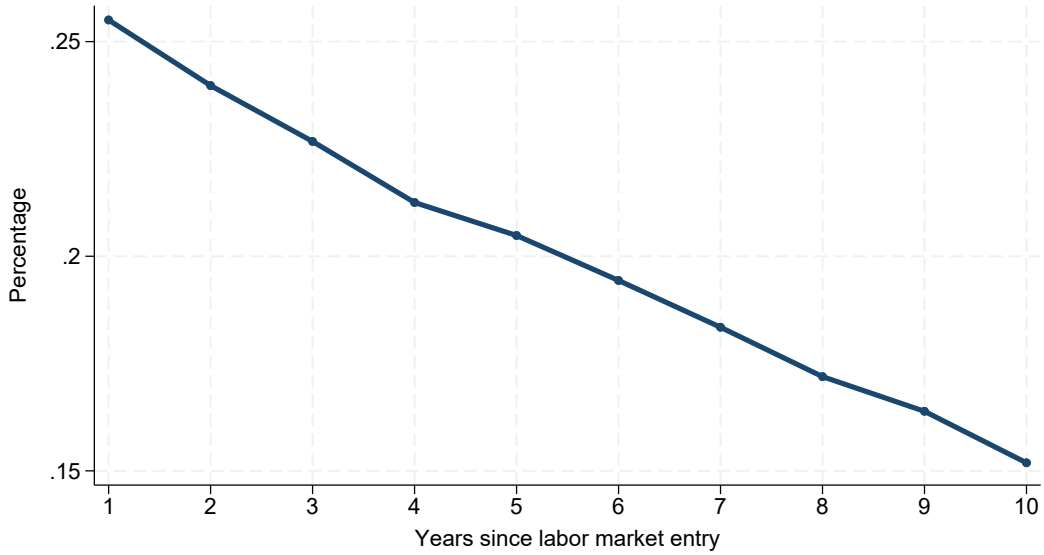
Note: The figure reports the results from the IV model in Equation 2, where the outcome is the cumulative number of firms the person has worked for. It displays the coefficients on the fixed-term entry dummy, scaled by the time-varying average cumulative number of firms for workers who entered with an open-ended contract.

Figure 19: Cumulative number of sectors - Relative difference: FT *vs* OE at entry (IV)



Note: The figure reports the results from the IV model in Equation 2, where the outcome is the cumulative number of sectors the person has worked in. It displays the coefficients on the fixed-term entry dummy, scaled by the time-varying average cumulative number of sectors for workers who entered with an open-ended contract.

Figure 20: Cumulative number of provinces - Relative difference: FT *vs* OE at entry (IV)



Note: The figure reports the results from the IV model in Equation 2, where the outcome is the cumulative number of provinces the person has worked in. It displays the coefficients on the fixed-term entry dummy, scaled by the time-varying average cumulative number of provinces for workers who entered with an open-ended contract.