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J. IGNACIO CONDE-RUIZ
JORGE FERNÁNDEZ-ORELLANA
DANIEL PÉREZ-GUTIÉRREZ

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J. Ignacio Conde-Ruiz†
FEDEA and Universidad Complutense de Madrid

Jorge Fernández-Orellana
Universidad Complutense de Madrid

Daniel Pérez-Gutiérrez
FEDEA

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Abstract

Formally universal labor protection policies may generate unequal outcomes when implemented in segmented labor markets. This paper studies how labor market duality shapes access to protection during large economic shocks, using the COVID-19 pandemic in Spain as an unexpected stress test. We combine administrative microdata with detailed information on employment transitions to examine workers' access to two distinct protection channels: job retention schemes (JRS) and unemployment-related protection. We show that unequal protection during the pandemic operated through two different institutional mechanisms. First, disparities in access to JRS are largely explained by labor market duality. Once contract type and job tenure are taken into account, age-related differences disappear or reverse, indicating that lower coverage among young workers mainly reflects composition effects rather than differential treatment within the scheme. Second, unequal access to unemployment-related protection persists even after controlling for contractual status and tenure. However, this gap is largely absorbed once accumulated labor market experience is taken into account, showing that contribution histories rather than age per se are the main source of disadvantage. Overall, the results show that labor market duality weakens the insurance role of the welfare state during economic crises. More broadly, the paper highlights that the distributional impact of formally universal protection policies depends crucially on the institutional channel through which protection is delivered.

JEL codes: J08, J21, J41, H55.
Keywords: labor market duality; social protection; job retention schemes; unemployment protection; COVID-19; Spain.

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†Corresponding author: `nacho.conderuiz@gmail.com`

1 Introduction

Labor market duality has long been recognized as a central feature of the Spanish employment model and as a major source of inefficient and unequal adjustment during downturns. A large literature has shown that the sharp divide between workers on open-ended and fixed-term contracts amplifies employment volatility, concentrates job destruction on more precarious groups, and generates persistent scarring over the life cycle. Yet much less is known about another consequence of dual labor markets: whether they also shape the effective distribution of protection when governments deploy formally universal labor market policies during large economic shocks.

This question is particularly important because labor protection policies are often designed as broad-based instruments intended to stabilize employment and income in recessions. However, when workers differ markedly in contractual stability, job tenure, and contribution histories, universal policies may generate highly unequal outcomes in practice. In this setting, unequal protection need not arise from explicit exclusion. It may instead emerge from the interaction between policy design and pre-existing segmentation in workers' labor market trajectories. Understanding this interaction is essential for assessing the distributive consequences of labor market institutions.

Spain offers a particularly revealing setting in which to study this issue. Before the 2021 labor market reform, Spain represented a canonical case of labor market duality, with a persistent divide between open-ended and fixed-term contracts, exceptionally high turnover, and large differences in employment stability across workers. This institutional structure has been shown to amplify aggregate labor market fluctuations. Bentolila et al. (2012), for example, show that the large gap in employment protection between fixed-term and permanent contracts substantially worsened Spain's labor market adjustment during the Great Recession. At the micro level, duality has also been linked to persistent employment instability and long-term scarring. Bentolila et al. (2022) document large and lasting earnings and employment losses for cohorts entering the labor market during downturns, while Bentolila et al. (2017) show how segmented career paths contribute to long-term unemployment dynamics. The widespread use of extremely short-duration contracts further underscores the intensity of labor turnover in Spain (Felgueroso et al., 2018), and Cabrales et al. (2017) show that fixed-term contracts are associated with lower on-the-job training and weaker human capital accumulation. Although fixed-term contracts may occasionally operate as stepping stones for some low-skilled young workers (García-Pérez et al., 2019), the broader evidence suggests that duality systematically generates fragile employment trajectories.

This paper studies that duality not only shapes who loses employment in a crisis, but also who gains access to protection once the shock arrives. We study whether unequal protection during the COVID-19 crisis was primarily driven by labor market duality itself or by exclusionary features embedded in the institutional design of protection systems. The pandemic provides an unusually sharp and unexpected stress test for this question. Unlike standard downturns, the COVID-19 shock was sudden, externally imposed, and highly disruptive across sectors, making it especially informative about how pre-existing labor market structures condition the reach of protection policies.

The Spanish case is especially well suited for this purpose because the pandemic activated two distinct protection channels. The first was the large-scale use of —*Expedientes de Regulación Temporal de Empleo*— (ERTEs), the Spanish job retention scheme (hereafter JRS) which allowed firms to temporarily suspend work while preserving the employment relationship. The second consisted of unemployment-related protection outside JRS, including contributory unemployment benefits, means-tested subsidies, and pensions. These two channels differ fundamentally in their institutional logic. JRS depend on the preservation of an ongoing contractual link and on firms' decisions to use the scheme. By contrast, unemployment-related protection is activated once the employment relationship breaks down and depends on workers' eligibility under rules tied to contribution histories, age thresholds, and household circumstances. This distinction allows us to separate two conceptually different sources of unequal protection: composition effects driven by labor market duality, and institutional exclusion embedded in the design of income-support systems.

To do so, we use administrative microdata from the Spanish Continuous Sample of Working Lives (MCVL) —*Muestra Continua de Vidas Laborales*—, which provides detailed information on employment spells, contract type, tenure, contribution records, and benefit receipt. We follow private-sector employees who were employed immediately before the pandemic and track their labor market transitions during the initial phase of the shock. Empirically, we compare workers' access to the two protection channels and examine how observed differences across groups change once we account for contract type, job tenure, and accumulated labor market experience. This empirical design allows us to distinguish whether lower protection among younger and more precarious workers reflects their pre-pandemic contractual position or the eligibility structure of the protection system itself.

Our analysis uncovers a dual mechanism behind unequal protection during the pandemic. First, overall coverage was markedly lower among younger and more precarious workers. However, this disparity operated through two distinct institutional channels. In the case of job retention schemes, age-related gaps largely disappear once contract type and tenure are taken into account. This indicates that lower JRS coverage among young workers mainly reflects composition effects: they were less protected not because they were treated differently within the scheme, but because they were disproportionately concentrated in temporary and short-tenure jobs. By contrast, unequal access to unemployment-related protection persists even after controlling for current contractual status and tenure. Once accumulated labor market experience is introduced, however, the age gradient is largely absorbed, indicating that the key mechanism is not age per se but eligibility rules tied to cumulative contribution histories. Together, these two channels generated a dual vulnerability for workers at the margins of the labor market during the crisis.

Our paper contributes to the literature in three ways. First, it extends the literature on labor market duality by showing that segmentation affects not only employment adjustment, but also the effective distribution of protection during large shocks. Second, it contributes to the literature on job retention schemes by shifting attention from their aggregate stabilization effects to their distributional implications in segmented labor markets. While previous work has shown that ERTEs played an important role in stabilizing employment during the pandemic (OECD, 2024; García-Serrano, 2022; Carrasco et al., 2024), much less is known about how access to these schemes was filtered by pre-existing labor market structures. Third, the paper highlights that formally universal protection policies may fail to provide universal protection in practice when access depends on contractual continuity or contribution histories. This insight has implications that extend beyond the COVID-19 episode and speaks more broadly to the design of labor market institutions in economies with segmented employment structures.

The remainder of the paper is organized as follows. Section 2 describes the institutional setting and develops the conceptual framework. Section 3 presents the data, the measurement of protection outcomes, and the descriptive evidence. Section 4 reports the econometric analysis. Section 5 concludes.

2 Institutional Setting and Conceptual Framework

2.1 Labor Market Duality and Protection Mechanisms

At the time of the COVID-19 shock, the Spanish labor market constituted a canonical example of contractual duality. Employment relationships were sharply divided between open-ended contracts, characterized by high job stability and long employment durations, and fixed-term contracts, associated with short tenures, high turnover, and weak attachment to firms. In early 2020, fixed-term contracts accounted for roughly one quarter of dependent employment, more than double the average observed in most European economies. Among young workers, temporary employment rates exceeded 60%, reflecting a strong concentration of labor market duality at early stages of the working life cycle.

This characterization applies to the institutional environment prevailing during the pandemic, prior to the labor market reform enacted in 2021, which substantially altered the contractual composition of employment. While the reform reduced the measured share of fixed-term contracts by more than 10 percentage points, recent evidence shows that this decline largely reflects contractual reclassification rather than a comparable improvement in job stability or a reduction in underlying labor market segmentation (Conde-Ruiz et al., 2025). The COVID-19 shock therefore provides a particularly informative setting to assess how a dual labor market structure conditions the effective reach of formally universal protection policies.

Labor protection during the pandemic operated through two distinct institutional channels. The first channel consisted of JRS, which allowed firms facing temporary shocks to suspend work activity while preserving the employment relationship. Access to JRS was determined at the firm level and required an ongoing contractual link at the time of activation. During the pandemic, these schemes were implemented on an unprecedented scale, with more than 3.5 million workers covered at the peak in April 2020. The program was accompanied by extraordinary measures, including relaxed eligibility requirements, exemptions from social security contributions for firms, and the non-consumption of future unemployment benefit entitlements for covered workers.

The second channel comprised unemployment-related protection outside JRS. This channel included contributory unemployment benefits, means-tested assistance programs, and certain non-contributory pensions. Access to these instruments was governed by individual eligibility rules based on accumulated contribution histories, age thresholds, and household characteristics. The contrast with JRS is central: whereas JRS protect workers whose employment relationship is preserved, unemployment-related protection applies only after that relationship has broken down and therefore depends much more directly on workers' previous labor market trajectories. In a dual labor market, this institutional structure tends to favor workers with long and stable employment careers.

2.2 Implications for Coverage during Economic Shocks

The coexistence of these two protection channels has direct implications for the distribution of coverage in a dual labor market. JRS primarily protect workers who are employed at the time of the shock and whose firms opt into the program. As a result, coverage through JRS is mechanically lower among workers with fixed-term contracts and short job tenures, who are more likely to experience contract termination before firms activate job retention measures. In this case, unequal coverage arises from composition effects driven by labor market duality rather than from differential treatment within the scheme itself.

In contrast, access to unemployment-related protection depends on institutional eligibility criteria that favor workers with long and continuous employment histories. Even conditional on job loss, workers with fragmented careers, limited contribution records, or shorter employment trajectories may fail to meet eligibility thresholds or may only qualify for less generous forms of assistance. Consequently, disparities in coverage through unemployment-related protection may persist even after controlling for current contract type and job tenure, reflecting the contribution-based design of the system rather than differences in treatment across workers with similar employment histories.

These considerations yield two testable predictions. First, differences in access to JRS across worker groups should largely disappear once contractual status and job tenure are taken into account, as unequal coverage reflects composition effects driven by labor market duality rather than differential treatment within the scheme. Second, unequal access to unemployment-related protection should persist after conditioning on contract type and tenure, but this residual gap should be substantially reduced once accumulated labor market experience is taken into account, consistent with eligibility rules that condition access on prior contribution histories rather than on demographic characteristics per se. This second prediction distinguishes between two competing interpretations of the age gap in unemployment protection: one in which younger workers are disadvantaged because of their demographic position, and one in which age merely proxies for weaker and more fragmented labor market attachment. The empirical analysis that follows assesses these predictions using administrative microdata on workers' employment transitions and protection outcomes during the COVID-19 shock.

3 Data and Measurement

3.1 Data

Our empirical analysis is based on administrative microdata from the MCVL, a longitudinal dataset derived from Spanish Social Security records. The MCVL provides detailed daily information on employment spells, contract characteristics, contribution bases, and benefit receipt, allowing us to track labor market transitions and protection outcomes with high precision. Unlike survey data, administrative records identify participation in JRS directly and capture short-lived employment and benefit spells that are difficult to observe in quarterly surveys.

We focus on the onset of the COVID-19 crisis. The baseline sample consists of private-sector employees [1] who were employed in February 2020, immediately prior to the implementation of containment measures. These individuals are followed through April 2020, corresponding to the peak of labor market disruption during the first lockdown[2]. This short observation window captures the immediate labor market effects of the shock and isolates the period in which the distinction between employment preservation through JRS and outright job loss is most clearly observed, avoiding confounding effects from later recovery phases or policy adjustments.

3.2 Labor Market Structure Prior to the Shock

Table 1 documents the strong segmentation of employment relationships in Spain prior to the pandemic, with substantial heterogeneity across demographic groups. While open-ended contracts account for roughly 70% of total employment, fixed-term contracts remain widespread, representing more than one quarter of all employment relationships. This aggregate picture, however, conceals pronounced differences across population groups. Temporary employment is particularly concentrated among younger workers and immigrants, whereas permanent employment becomes dominant only at later stages of the working life cycle.

The age gradient is especially striking. Among workers aged 16–19, nearly 70% hold fixed-term contracts and fewer than one third are employed under open-ended contracts. Although the incidence of permanent employment increases steadily with age, fixed-term contracts remain common well into the late twenties, indicating a prolonged transition toward stable employment. Immigrant workers also display substantially higher temporary employment rates than natives, suggesting that labor market segmentation disproportionately affects groups with weaker attachment or lower bargaining power. These patterns imply that even policies that are formally universal may generate unequal protection across groups, as both eligibility and effective coverage depend crucially on workers' contractual position at the onset of the shock.

[1] Employees in the public sector are excluded, as their hiring processes follow specific institutional mechanisms that are not directly comparable to those of the private sector.

[2] For each period, active labor market statuses are identified. Self-employed workers are also excluded because their employment relationships and protection mechanisms are institutionally distinct from those of dependent employees. We also exclude voluntary separations from employment, as these do not generate eligibility for unemployment-related protection and would mechanically bias the measurement of coverage. When an individual has multiple records within the same period, the last one in chronological order is prioritized, under the assumption that it most accurately reflects their status at the end of the interval. Thus, if an individual first appears as receiving unemployment benefits and later as employed within the same period, the latter status is recorded as it is considered the most representative.

Table 1: Distribution by Contract Type (%)

	Contract Type		
	Fixed-Term	Intermittent Open-Ended	Open-Ended
Female	29.04	3.44	67.52
Male	26.32	2.21	71.47
Ages 16–19	69.71	3.08	27.21
Ages 20–24	54.02	2.58	43.40
Ages 25–29	39.79	2.44	57.77
Ages 30–34	30.30	2.47	67.23
Ages 35–49	23.49	2.58	73.93
Ages 50–65	21.89	3.44	74.67
Immigrants	39.18	4.54	56.28
Natives	26.17	2.58	71.25
Total	**27.61**	**2.79**	**69.60**

Source: Authors' calculations using MCVL data.

The distribution of contract types documented above is not only a structural feature of the Spanish labor market, but also a key determinant of workers' exposure to employment shocks and their access to protection mechanisms. Since fixed-term contracts are easier to terminate and typically provide weaker employment continuity, groups disproportionately concentrated in temporary employment face higher risks of job loss during downturns. Access to income protection, in turn, often depends on employment stability and contribution histories, linking contractual arrangements to effective social protection.

To examine how these pre-existing differences translated into labor market outcomes at the onset of the pandemic, we analyze transitions across labor market states between February and April 2020. Using daily employment and benefit records, individuals are classified into five mutually exclusive states: (i) Employment (Employ.), (ii) employment under a job retention scheme (JRS), (iii) receipt of unemployment benefits or assistance (Benefits), (iv) unemployment without benefit receipt (Unemp. No Ben.)[3] and (v) retirement (Retired). Transitions over this period allow us to identify employment disruptions and subsequent access to protection.

A worker is considered affected by the shock if the employment relationship observed in February 2020 is no longer active in its original form by April, either because it is suspended under a JRS or because it has been terminated. Among affected workers, we distinguish between protection through job retention schemes and protection through unemployment-related instruments (benefits), reflecting the fundamentally different institutional logics governing employment preservation and income replacement.

[3]Individuals are classified as *"Unemployed without benefits"* if, in a given period, they do not appear in the administrative records as employed, under JRS, or as recipients of unemployment benefits or pensions. This category may encompass different situations, such as active job search without entitlement to benefits or temporary withdrawal from the labour market. The category is an imputation required due to a limitation in the data. When an individual loses their job and does not receive any type of benefit or pension, their record disappears from the administrative data until they establish a new employment relationship.

Table 2: Transitions from Employment Status

	Transitions from Employment				
	Employ.	JRS	Benefits	Unemp. No Ben.	Retired
Female	69.66	21.72	4.68	3.84	0.10
Male	68.71	22.16	5.40	3.59	0.14
Ages 16–19	37.05	27.43	4.74	30.78	0.00
Ages 20–24	47.77	30.47	7.77	13.99	0.00
Ages 25–29	60.12	27.07	7.12	5.69	0.00
Ages 30–34	65.47	24.93	6.08	3.52	0.00
Ages 35–49	71.02	21.89	4.75	2.34	0.00
Ages 50–65	76.88	17.25	3.92	1.58	0.37
Immigrants	61.30	23.33	6.64	8.72	0.01
Natives	70.35	21.77	4.86	2.89	0.13
Total	**69.32**	**21.93**	**5.06**	**3.57**	**0.12**

Source: Authors' calculations using MCVL data.

Table 2 shows a strong age gradient in employment stability at the onset of the pandemic. Employment retention rises sharply with age: fewer than 40% of workers aged 16–19 remain employed between February and April 2020, compared with nearly 80% among those aged 50–65. Younger workers are not only more likely to exit employment but, conditional on job loss, are less likely to transition into protected states such as JRS or benefit receipt, and more likely to move directly into unemployment without benefits. This pattern highlights the weaker effective protection faced by younger workers during the initial phase of the shock.

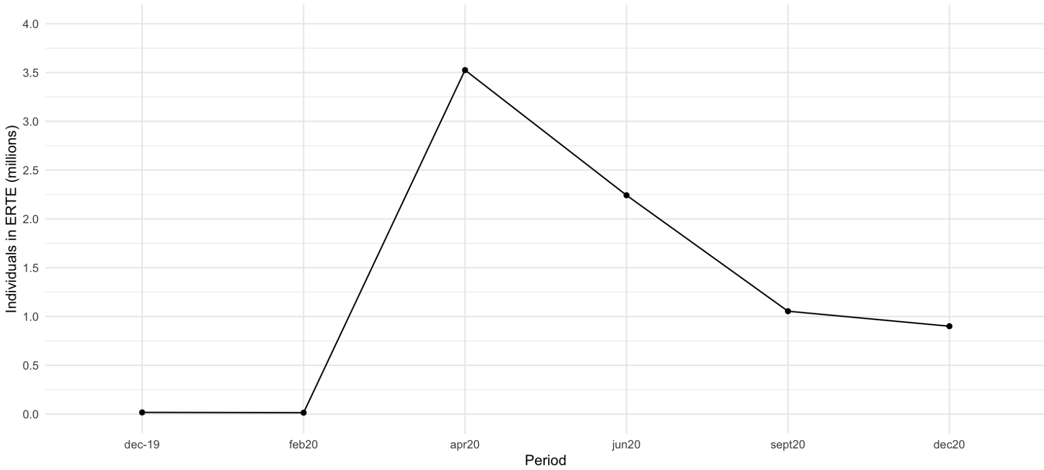
These differences closely mirror the distribution of contract types across age groups. Fixed-term contracts are concentrated among young workers, whereas open-ended contracts dominate among older cohorts. Since open-ended contracts are associated with greater employment continuity and broader access to protection mechanisms, the concentration of temporary employment among young workers helps explain both their greater exposure to job loss and their lower effective coverage during the crisis.

3.3 Employment Protection through the JRS Program

To assess the extent to which JRS protected employment during the initial phase of the pandemic, we construct a measure of effective coverage that compares the use of JRS with employment exits not covered by this mechanism.

The underlying logic is straightforward. Faced with a sudden collapse in economic activity, firms could either preserve employment relationships by placing workers on JRS—temporarily suspending contracts or reducing working hours—or adjust employment through dismissals or the non-renewal of fixed-term contracts. From the worker's perspective, these responses represent alternative adjustments to the same shock, but they imply very different outcomes in terms of employment continuity and income protection.

Figure 1: Individuals on JRS by period



Note: Authors' own elaboration using MCVL data.

Figure 1 illustrates the extent to which firms relied on this adjustment channel at the onset of the crisis. The number of workers covered by JRS rose from virtually zero to more than 3.5 million by April 2020, reflecting the massive use of this mechanism in response to the sudden collapse in economic activity. This peak shows that the initial adjustment was, to a large extent, channelled through the temporary suspension of ongoing employment relationships rather than through immediate job destruction, highlighting the central role of JRS in cushioning the first impact of the shock on employment.

To summarize this margin of adjustment, we define for each group i the following protection rate:

$$\text{JRS Protection Rate}_i = \frac{JRS_i}{JRS_i + Exit_i} \tag{1}$$

where JRS_i denotes workers placed on JRS and $Exit_i$ employment exits not covered by JRS.

Table 3: JRS Protection Rate (%)

	Contract Type			
	Fixed-Term	Intermittent Open-Ended	Open-Ended	**Total**
Female	38.27	71.60	91.26	**72.40**
Male	32.78	57.59	91.69	**70.80**
Ages 16–19	31.54	48.67	79.44	**43.59**
Ages 20–24	39.21	60.44	88.24	**58.42**
Ages 25–29	40.17	71.20	90.44	**67.95**
Ages 30–34	36.41	67.75	91.27	**72.18**
Ages 35–49	32.78	64.95	92.75	**75.45**
Ages 50–65	31.98	69.15	91.12	**74.63**
Immigrants	32.01	57.69	83.04	**60.23**
Natives	36.09	68.04	92.60	**73.42**
Total	**35.34**	**66.65**	**91.50**	**71.56**

Source: Authors' calculations using MCVL data.

Table 3 reveals substantial heterogeneity in effective employment protection during the initial phase of the pandemic. This heterogeneity, however, is driven primarily by contract type rather than by workers' demographic characteristics.

Protection rates increase sharply with contractual stability. Workers holding open-ended contracts exhibit protection rates above 90 percent, whereas coverage among temporary workers remains close to one third, indicating that employment adjustment for this group largely occurred through contract termination rather than through temporary suspension mechanisms such as JRS.

Once comparisons are made within contract types, age differences become much smaller. Among workers with open-ended contracts, protection rates remain both high and relatively uniform across age groups, and a similar pattern holds for temporary and intermittent open-ended contracts. The lower aggregate protection rates observed among younger workers therefore reflect primarily their greater concentration in temporary employment rather than an age-specific difference in access to JRS protection.

Overall, these results are consistent with the view that the effective reach of the program depended mainly on pre-existing contractual structures. Groups more exposed to unprotected employment exits were those disproportionately employed under fixed-term contracts, among whom young and immigrant workers are overrepresented. In this sense, rather than creating new inequalities, the crisis and the associated policy response largely amplified pre-existing patterns of labor market segmentation.

3.4 Income Protection outside JRS: Unemployment Benefits, Assistance, and Pensions

The pandemic not only triggered a sharp contraction in employment but also tested the capacity of social protection mechanisms to cushion income losses associated with job separations. While JRS played a central role in preserving employment relationships during the initial phase of the crisis, a substantial fraction of affected workers lost their jobs without being covered by this mechanism. For these workers, income protection depended on access to alternative public transfers, primarily unemployment benefits, means-tested assistance programs, or, for older workers, pensions.

In this subsection, we examine the extent to which workers who exited employment without being placed on JRS were able to access alternative forms of income support. To summarize coverage, we define the protection rate outside JRS for group i as:

$$\text{Income Protection Rate}_i = \frac{Benefits_i}{Exit_i} \tag{2}$$

where $Benefits_i$ denotes the number of workers in group i receiving unemployment benefits, assistance, or pensions after leaving employment, and $Exit_i$ represents total employment exits not covered by JRS.

Table 4: Income Protection Rate (%) (Unemployment Benefits, Assistance, or Pensions)

	Contract Type			
	Fixed-Term	Intermittent open-ended	Open-Ended	**Total**
Female	56.07	77.55	57.77	**57.68**
Male	58.95	73.46	65.37	**60.70**
Ages 16–19	13.46	12.07	12.62	**13.35**
Ages 20–24	35.26	38.89	39.40	**35.84**
Ages 25–29	55.22	64.81	56.93	**55.82**
Ages 30–34	62.80	76.30	62.85	**63.37**
Ages 35–49	66.08	82.70	64.50	**66.69**
Ages 50–65	72.86	83.55	70.55	**73.38**
Immigrants	43.64	62.59	38.49	**43.23**
Natives	60.92	78.32	69.28	**63.28**
Total	**57.66**	**75.74**	**61.82**	**59.37**

Source: Authors' calculations using MCVL data.

Table 4 shows that, unlike JRS coverage, protection through unemployment benefits, assistance programs, or pensions depends much more strongly on workers' age than on contract type. Protection rates increase sharply along the age profile, from barely 13% among workers aged 16–19 to more than 70% among those aged 50–65.

Differences across contract types are comparatively modest once age is taken into account.

Within each age group, protection rates are broadly similar across fixed-term, intermittent open-ended, and open-ended contracts, indicating that contractual modality plays a secondary role in determining access to income support after job loss outside JRS.

This pattern is consistent with the institutional design of unemployment-related protection, where access depends strongly on prior contribution histories, accumulated employment spells, and other eligibility criteria linked to workers' previous labor market trajectories. Older workers, having typically accumulated longer and more stable employment records, are therefore more likely to qualify for benefits following job separation. Conversely, younger workers, whose employment histories are shorter and more fragmented, face substantially lower access to income protection once employment is lost.

Overall, the results indicate that vulnerability following employment loss during the pandemic was driven less by contract type per se and more by workers' accumulated attachment to the labor market over the life cycle.

3.5 Total Protection during the Pandemic

To obtain a more comprehensive measure of the scope of employment protection mechanisms during the pandemic, we define the *Total Protection Rate*. This measure combines both coverage provided through JRS and the protection subsequently obtained via unemployment insurance, assistance benefits, or pensions. Formally, for each group i:

$$\text{Total Protection Rate}_i = \frac{JRS_i + Benefits_i}{JRS_i + Exit_i} \tag{3}$$

with $Benefits_i$ representing workers who, after losing employment, access unemployment benefits, assistance programs, or pensions, and $Exit_i$ those who experience employment loss without accessing subsequent protection mechanisms.

Table 5: Total Protection Rate during the Pandemic (%)

	Contract Type			
	Fixed-Term	Intermittent Open-Ended	Open-Ended	**Total**
Female	72.87	93.59	96.29	**88.33**
Male	72.43	88.75	97.13	**88.58**
Ages 16–19	40.74	54.87	82.04	**51.12**
Ages 20–24	60.48	75.82	92.82	**73.18**
Ages 25–29	73.12	89.86	95.85	**85.77**
Ages 30–34	76.34	92.35	96.76	**89.79**
Ages 35–49	77.20	93.95	97.42	**91.82**
Ages 50–65	81.56	94.94	97.38	**93.20**
Immigrants	61.67	84.14	89.55	**77.45**
Natives	75.01	93.05	97.75	**90.22**
Total	**72.63**	**91.90**	**96.85**	88.44

Source: Authors' calculations using MCVL data.

Table 5 reports total protection rates during the pandemic, combining workers protected through JRS and those who, after job loss, accessed unemployment benefits, assistance programs, or pensions. Overall, protection mechanisms covered nearly 88% of workers affected by the shock, indicating that public intervention substantially mitigated employment and income losses during the crisis.

However, this aggregate figure masks sharp differences across demographic groups. Protection rises steeply with age: only about half of workers aged 16–19 were ultimately protected, compared with more than 90% among workers aged 35 and above. This pattern reflects both lower JRS coverage and weaker access to subsequent income support among younger workers, likely linked to shorter contribution histories and weaker labor market attachment. As a result, young workers remained the group most exposed to unprotected employment loss during the pandemic.

Differences across contract types are also substantial, but they appear to reflect underlying workforce composition to a large extent. Open-ended workers exhibit the highest protection rates, exceeding 96% on average, while temporary workers remain substantially less protected, with overall coverage around 73%. Intermittent open-ended contracts occupy an intermediate position but still display protection levels much closer to open-ended contracts than to temporary ones. Importantly, once age is taken into account, protection rates within each contract type become considerably more similar, suggesting that accumulated employment histories and eligibility conditions play a central role in determining effective protection.

Finally, sizable gaps also emerge by migrant status. Immigrant workers exhibit protection rates roughly thirteen percentage points below natives, a pattern consistent with weaker attachment to stable employment relationships and more limited access to contributory benefits. Taken together, these results indicate that although pandemic protection mechanisms reached a large share of affected workers, coverage remained uneven and systematically lower among younger and more weakly attached segments of the workforce.

4 Determinants of Access to Protection during the Pandemic

This section examines the factors shaping effective access to the main employment protection mechanisms deployed during the pandemic: JRS, on the one hand, and unemployment benefits, assistance allowances, or pensions on the other. Although both instruments form part of the social protection system, their institutional logic, activation rules, and eligibility requirements differ substantially. JRS are firm-initiated measures that preserve the employment relationship and therefore depend crucially on contract type and on the stability of the pre-shock employment match. By contrast, access to unemployment benefits, assistance allowances, or pensions requires the termination of the employment relationship and the fulfillment of pre-existing eligibility conditions, including sufficient contribution histories, age, or other institutional requirements.

To capture these differences, we estimate two separate probability models—both using a logit specification—to identify how workers' contractual and sociodemographic characteristics influenced access to each protection channel. Estimating these models separately allows us to isolate more clearly the structural determinants of protection and to evaluate the extent to which labor market segmentation in Spain conditioned the inclusiveness of the policy response.

Both models rely on labor market information observed in February 2020, immediately prior to the economic disruption caused by the pandemic. The analysis starts from individuals who were employed at that time and then follows their labor market status through April 2020. We estimate the probability of accessing each protection channel as a function of workers' pre-shock contractual and sociodemographic characteristics.

Specifically, the model for JRS access is estimated among workers affected by the shock, defined as those whose February employment relationship was either suspended under JRS or terminated by April 2020. Within this group, the dependent variable identifies whether the worker was placed on JRS rather than transitioning into other states, including unemployment benefits, pensions, or inactivity. In turn, the model for access to unemployment benefits, assistance allowances, or pensions is estimated only among workers whose employment relationship was terminated and who were not protected through a JRS. Within this subgroup, the dependent variable identifies whether the worker obtained unemployment-related protection rather than remaining without formal coverage. This sequential approach reflects the institutional structure of the protection system: JRS operate on the margin of employment preservation, whereas unemployment-related protection operates after the employment relationship has broken down.

Given the binary nature of the outcomes of interest, both models are estimated using a logit specification of the following form:

$$\log\left(\frac{\mathbb{P}(Y_i=1)}{1-\mathbb{P}(Y_i=1)}\right) = \beta_0+\beta_1\text{Age}_i+\beta_2\text{Contract}_i+\beta_3\text{EmploymentDuration}_i+\beta_4X_i+\varepsilon_i, \quad (4)$$

where Y_i denotes an indicator of access to the protection mechanism under consideration. The vector Age_i includes indicators for age groups (35–50 and 51 or older), with workers aged 16–34 serving as the reference category. Contract_i includes indicators for contract type (intermittent open-ended and regular open-ended), with fixed-term contracts as the baseline. $\text{EmploymentDuration}_i$ captures tenure in the main job as of February 2020, measured in tenure brackets. The vector X_i includes additional controls for education, gender, nationality, sector of activity, contribution base, and working-time arrangement.

4.1 Determinants of Access to JRS

To examine the determinants of access to job retention schemes during the pandemic, we estimate a logistic regression model that identifies the individual and contractual characteristics associated with a higher probability of being covered by JRS. The analysis relies on workers' labor market status in February 2020, immediately prior to the onset of the health crisis, thereby mitigating endogeneity concerns arising from adjustments induced by the shock itself.

The dependent variable, Y_i, is a binary indicator equal to one if individual i was placed on JRS in April 2020, and zero otherwise. Consistent with the empirical design described above, the estimation sample is restricted to workers affected by the shock, that is, workers whose February employment relationship was either suspended under JRS or terminated by April 2020. The purpose of this specification is therefore to estimate the conditional probability of being placed on JRS rather than transitioning out of employment, controlling for observable characteristics that proxy employment stability and attachment to the labor market.

Table 6: Average Marginal Effects on Probability of JRS

	(1) Access to JRS	(2) Access to JRS	(3) Access to JRS
Age (ref.: 16–34)			
35–50 years	0.0403***	-0.0014	-0.0159***
	(0.0022)	(0.0020)	(0.0016)
51 years or more	0.0182***	-0.0137***	-0.0484***
	(0.0029)	(0.0025)	(0.0023)
Contract Type (ref.: Temporary)			
Intermittent open-ended		0.2041***	0.0651***
		(0.0063)	(0.0026)
Regular open-ended		0.4364***	0.0114***
		(0.0033)	(0.0026)
Employment Duration (ref.: ≤3 months)			
4–6 months			0.1148***
			(0.0052)
7–12 months			0.5103***
			(0.0053)
13–24 months			0.6312***
			(0.0046)
25–36 months			0.7404***
			(0.0048)
More than 36 months			0.8557***
			(0.0035)
Individual controls	✓	✓	✓
Contract type controls		✓	✓
Employment duration controls			✓
Pseudo R^2	0.17	0.32	0.48
Observations	154,803	154,803	154,803

Notes: Average marginal effects reported. Robust standard errors in parentheses. Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. Effects measure percentage-point changes in probability relative to the reference category. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 6 reports average marginal effects from logistic regressions estimating the probability of being covered by JRS in April 2020 among workers affected by the shock.[4] The three columns progressively introduce controls in order to disentangle the role of age, contract type, and employment stability in shaping access to the scheme.

Column (1), which includes only age indicators, suggests that older workers were more likely to be protected through JRS. Workers aged 35–50 display a 4.0 percentage-point (pp) higher probability of coverage than those aged 16–34, while those aged 51 or more exhibit a

[4]For comparison, Table A.1 in the Appendix reports the corresponding log-odds coefficients from the baseline logit specifications.

1.8 pp higher probability relative to the reference group. Taken at face value, this pattern would suggest that protection increased with age during the initial phase of the crisis.

However, this relationship changes sharply once differences in contract type are taken into account. Column (2) introduces controls for contract type and shows that age differences largely disappear, while contract type emerges as a central determinant of coverage. Workers with regular open-ended contracts exhibit a 43.6 pp higher probability of being covered compared with temporary workers, while those with intermittent open-ended contracts show a 20.4 pp higher probability.

Column (3), which additionally controls for employment duration, reveals that a substantial part of the advantage associated with open-ended contracts operates through longer employment relationships. Conditional on both contract type and employment duration, workers with regular open-ended contracts still display a modest positive marginal effect of 1.1 pp, whereas intermittent open-ended contracts show a smaller but still significant effect of 6.5 pp. Employment duration also strongly increases coverage: for example, workers with more than 36 months of tenure display an increase of approximately 85.6 pp in the probability of JRS coverage relative to workers with very short tenure (three months or less).

Taken together, these results provide strong evidence of a composition effect. The higher raw coverage observed among older workers reflects primarily their concentration in more stable employment relationships—especially open-ended contracts and longer tenure—rather than differential treatment within the scheme itself. Once these structural differences are taken into account, the age gradient in access to JRS largely disappears and, in the fully specified model, reverses.

The robustness of the main findings is assessed in the Appendix. The Linear Probability Model in Table A.2 confirms the qualitative patterns observed in the baseline marginal effects. In particular, the conditional age profile remains stable across models: while older workers initially appear more likely to access JRS, this gradient reverses once contract type and employment duration are included.

The robustness check using a finer age grouping (Table A.3) further confirms that the main results are not driven by the specific age categorization. Although raw age gradients remain large, they largely disappear after conditioning on tenure.

As an additional robustness check, we re-estimate the main specifications excluding sectors that were particularly exposed to the employment shock during the pandemic, including commerce, hospitality, low-skilled services, and temporary work agencies (Table A.4). The results remain qualitatively unchanged. In particular, the apparent age gradient in access to job retention schemes disappears once contract type and employment duration are taken into account, while contractual stability and job tenure continue to be the main determinants of access. This suggests that the baseline results are not driven by the sectoral composition of employment.

Overall, the results indicate that JRS coverage was strongly conditioned by workers' contractual stability and prior attachment to firms. The lower aggregate coverage observed among young workers largely reflects their structural concentration in temporary, intermittent, and short-tenure jobs. In this sense, the operation of JRS tended to reinforce existing dualities in the Spanish labor market, disproportionately benefiting workers in more stable employment relationships while leaving those in more precarious positions less protected against the employment shock.

4.2 Determinants of Access to Income Protection

In contrast to JRS coverage, access to unemployment benefits, assistance allowances, or pensions is only relevant for workers whose employment relationship is terminated. For this reason, the estimation sample for this specification is restricted to workers who lose their job during the observation window and are not protected through JRS.

Within this group, we estimate the probability of receiving income protection as a function of workers' pre-shock characteristics. In this specification, the dependent variable Y_i equals one if the worker receives unemployment benefits, assistance allowances, or pensions after job loss, and zero otherwise.

Table 7: Average Marginal Effects on Access to Income Protection after Job Loss

	(1) Income Protection	(2) Income Protection	(3) Income Protection	(4) Income Protection
Age (ref.: 16–34)				
35–50 years	0.1083***	0.1062***	0.1075***	-0.0017
	(0.0052)	(0.0052)	(0.0052)	(0.0063)
51 years or more	0.1208***	0.1182***	0.1182***	-0.0663***
	(0.0068)	(0.0068)	(0.0068)	(0.0099)
Contract Type (ref.: Temporary)				
Intermittent open-ended		0.1161***	0.1128***	0.1023***
		(0.0257)	(0.0257)	(0.0260)
Regular open-ended		0.0488***	-0.0079	-0.0111
		(0.0082)	(0.0097)	(0.0095)
Employment Duration (ref.: ≤3 months)				
4–6 months			0.0683***	0.0721***
			(0.0064)	(0.0063)
7–12 months			0.1294***	0.1321***
			(0.0074)	(0.0073)
13–24 months			0.1674***	0.1653***
			(0.0089)	(0.0089)
25–36 months			0.1690***	0.1667***
			(0.0120)	(0.0120)
More than 36 months			0.1241***	0.1137***
			(0.0122)	(0.0123)
Years since entry into labor market				0.0083***
				(0.0003)
Individual controls	✓	✓	✓	✓
Contract type controls		✓	✓	✓
Employment duration controls			✓	✓
Experience control				✓
Pseudo R^2	0.25	0.26	0.27	0.28
Observations	32,404	32,404	32,404	32,404

Notes: Average marginal effects reported. Robust standard errors in parentheses. Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. Effects measure percentage-point changes in probability relative to the reference category. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table 7 reports average marginal effects from logistic regressions estimating the probability of receiving income protection following job loss.[5] The four columns progressively

[5]For comparison, the Appendix reports the corresponding log-odds coefficients from the baseline logit specifications.

introduce controls to disentangle the roles of age, contract type, employment duration, and accumulated labor market experience in shaping access to protection.

Column (1), which includes only age indicators, shows that older workers were substantially more likely to receive income protection. Workers aged 35–50 display a 10.83 pp higher probability of receiving benefits relative to workers aged 16–34, while those aged 51 or more show a 12.08 pp higher probability.

Once contract type is accounted for in column (2), age differences remain largely stable. Contract type also emerges as an important correlate of access. Workers with intermittent open-ended contracts have an 11.61 pp higher probability of receiving income protection compared with temporary workers, while those with regular open-ended contracts show a 4.88 pp higher probability. However, these differences should not be interpreted as purely contractual effects, since they may partly reflect more stable prior employment histories.

Column (3), which additionally controls for employment duration, highlights the strong influence of prior tenure. Conditional on contract type and tenure, the marginal effect for regular open-ended contracts becomes slightly negative (-0.79 pp), while intermittent open-ended contracts retain a significant positive effect of 11.28 pp. Longer employment durations also strongly increase the probability of receiving income protection: compared with workers with 3 months or less, those with 4–6 months have a 6.83 pp higher probability, those with 7–12 months a 12.94 pp higher probability, those with 13–24 months a 16.74 pp higher probability, those with 25–36 months a 16.90 pp higher probability, and those with more than 36 months a 12.41 pp higher probability. These effects are consistent with the role of accumulated contributions in determining eligibility for unemployment-related protection.

A potential concern is that the positive age gradient observed in columns (1)–(3) may simply proxy accumulated labor market attachment, as older workers have had more time to build contribution histories. Column (4) addresses this mechanism by including a measure of labor market experience—years since entry into the labor market.[6] Once this variable is introduced, the positive age gradient largely disappears. The coefficient for workers aged 35–50 becomes statistically indistinguishable from zero, while the effect for those aged 51 or more turns negative. By contrast, accumulated experience emerges as a strong and highly significant predictor of access to income protection.

These results indicate that the previously observed age gradient primarily reflects differences in cumulative contribution histories rather than age per se. In other words, access to unemployment-related protection is driven mainly by accumulated labor market attachment, not by demographic characteristics alone. The institutional design of the system, which ties eligibility to prior contributions and employment continuity, therefore generates systematic differences across workers depending on the length and continuity of their employment trajectories.

The robustness of the main findings is assessed in the Appendix. The Linear Probability Model in Table A.6 confirms the qualitative patterns observed in the baseline marginal effects. The robustness check using a finer age grouping in Table A.7 further supports the

[6]Controlling for accumulated experience is conceptually appropriate in the case of unemployment-related protection, as eligibility is explicitly conditioned on prior contribution histories. In contrast, access to JRS depends on the preservation of an ongoing employment relationship rather than on cumulative contributions. For this reason, accumulated experience does not constitute a direct eligibility criterion in the case of JRS and is therefore not included in those specifications.

stability of the results across alternative age definitions. The qualitative age pattern is preserved after conditioning on tenure, indicating that the findings are not driven by the specific age categorization adopted in the baseline specification.

As an additional robustness check, we examine whether the results reflect the uneven sectoral impact of the pandemic. During the crisis, employment losses were particularly concentrated in sectors such as commerce, hospitality, low-skilled services, and temporary work agencies, raising the possibility that the estimated relationships could partly capture sector-specific shocks. To address this concern, we re-estimate the baseline models after excluding these sectors. The corresponding estimates, reported in Table A.8, remain largely unchanged.

Overall, these results indicate that access to income protection after job loss is strongly conditioned by prior labor market attachment rather than by contract type alone. Young workers face lower coverage primarily because they are disproportionately concentrated in short-tenure positions and have accumulated fewer contribution years. Once tenure and accumulated experience are taken into account, differences by contract type largely diminish, and the apparent age gradient is substantially reduced. This pattern underscores that eligibility rules tied to cumulative and continuous contribution histories systematically disadvantage workers with fragmented employment trajectories, particularly younger cohorts whose attachment to the labor market is still in its early stages.

5 Conclusions

Young workers were the group most exposed to the labor market consequences of the COVID-19 crisis. Workers aged 16 to 34 accounted for more than half of total employment losses between the fourth quarter of 2019 and the second quarter of 2020, despite representing only about one quarter of salaried employment. At the same time, their access to protection mechanisms was substantially lower than that of older workers. Coverage through JRS reached around 75% among prime-age and older workers but only 65% among workers aged 16 to 34, and barely exceeded 40% among teenagers. Alternative protection channels also displayed large gaps, so that total coverage remained well below that observed for older workers. Taken together, these figures indicate that young workers experienced both larger employment losses and weaker institutional protection during the pandemic.

These differences largely reflect structural features of the Spanish labor market. Young workers are disproportionately employed under fixed-term contracts, which are both less likely to be covered by JRS and more likely to terminate during downturns. In addition, shorter and more fragmented employment histories reduce access to unemployment-related protection once the employment relationship is lost. The crisis therefore revealed a broader feature of dual labor markets: they shape not only who is more exposed to job loss, but also who is effectively protected when formally universal policies are deployed under stress.

The econometric analysis clarifies the mechanisms behind these patterns. Differences in access to JRS largely disappear once contract type and tenure are controlled for, indicating that lower protection among young workers mainly reflects their contractual position rather than differential treatment within the scheme. By contrast, differences in access to unemployment-related protection persist after conditioning on current contract type and tenure, but they are substantially reduced once accumulated labor market experience is taken into account. This suggests that unequal protection outside JRS is driven less by age per se than by eligibility rules tied to cumulative contribution histories. Taken together, these findings point to a dual mechanism of under-protection: one rooted in contractual segmentation, and another embedded in the design of income-support institutions.

These results have implications that extend beyond the COVID-19 episode. They show that formally universal labor protection policies do not necessarily generate universal protection in practice. In segmented labor markets, the distributive reach of crisis-response instruments depends crucially on the institutional channel through which protection is delivered. Policies based on preserving employment relationships tend to favor workers already attached to stable jobs, while policies based on prior contributions disadvantage those with shorter and more fragmented careers.

In this context, the 2021 labor market reform may have important implications for protection in future downturns. Existing evaluations have emphasized that the reform sharply reduced measured temporary employment, but did not produce an equally strong improvement in effective job stability, as many employment relationships remain highly intermittent or short-lived (Conde-Ruiz et al., 2025). Our results point to an additional and less explored implication. Even if the reform has not substantially reduced underlying instability, the decline in fixed-term hiring may still increase the reach of protection mechanisms that depend on preserving an employment relationship.

This is particularly relevant in the case of JRS. During the COVID-19 crisis, workers under fixed-term contracts were much less likely to be protected through JRS, largely because these contracts were easier to terminate before firms activated internal adjustment mechanisms. By contrast, workers under intermittent open-ended contracts exhibited substantially higher protection rates. To the extent that the reform shifted workers away from fixed-term contracts and toward regular open-ended or intermittent open-ended arrangements, it may have increased the share of workers who remain contractually attached to firms and can therefore be covered by instruments such as JRS during future shocks.

This implication should, however, be interpreted with caution. First, the expansion of Intermittent open-ended contracts may have extended this category to employment relationships with lower stability than previously observed. Second, recent evidence suggests that early termination rates among open-ended contracts may have increased after the reform, particularly during probation periods. Consequently, the gains in formal contractual continuity may not fully translate into longer-lasting employment relationships. Still, even without a commensurate improvement in effective stability, a labor market with fewer fixed-term contracts may be better positioned to channel adjustment through internal flexibility mechanisms rather than through outright job destruction.

More broadly, reducing contractual segmentation while strengthening internal adjustment mechanisms—such as the RED Mechanism introduced after the pandemic—may improve the capacity of the Spanish labor market to absorb future shocks in a more equitable manner. The broader lesson of the paper is therefore not only that young workers were less protected during COVID-19, but that the insurance role of the welfare state is itself conditioned by the contractual structure of the labor market through which that protection is delivered.

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Appendix

Table A.1: Log-Odds Coefficients on Access to JRS

	(1) Access to JRS	(2) Access to JRS	(3) Access to JRS
Age (ref.: 16–34)			
35–50 years	0.2802***	0.9248	-0.2050***
	(0.0154)	(0.0304)	(0.0213)
51 years or more	0.1222***	2.5898***	-0.5869***
	(0.0193)	(0.0176)	(0.0270)
Contract Type (ref.: Temporary)			
Intermittent open-ended		0.9248***	0.8664***
		(0.0304)	(0.0362)
Regular open-ended		2.5898***	0.1355***
		(0.0176)	(0.0297)
Employment Duration (ref.: ≤3 months)			
4–6 months			0.9887***
			(0.0427)
7–12 months			2.8967***
			(0.0369)
13–24 months			3.4998***
			(0.0412)
25–36 months			4.2177***
			(0.0493)
More than 36 months			5.7254***
			(0.0500)
Individual controls	✓	✓	✓
Contract type controls		✓	✓
Employment duration controls			✓
Pseudo R^2	0.17	0.32	0.48
Observations	154,803	154,803	154,803

Notes: Log-odds coefficients reported. Robust standard errors in parentheses. Column (1) includes only age indicators; Column (2) adds contract type; Column (3) further includes employment duration. Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.2: Linear Probability Model Coefficients on Access to JRS

	(1) Access to JRS	(2) Access to JRS	(3) Access to JRS
Age (ref.: 16–34)			
35–50 years	0.0429***	-0.0037*	-0.0199***
	(0.0023)	(0.0021)	(0.0018)
51 years or more	0.0183***	-0.0151***	-0.0469***
	(0.0029)	(0.0026)	(0.0022)
Contract Type (ref.: Temporary)			
Intermittent open-ended		0.2263***	0.1428***
		(0.0062)	(0.0054)
Regular open-ended		0.4655***	0.0317***
		(0.0029)	(0.0043)
Employment Duration (ref.: ≤3 months)			
4–6 months			0.1025***
			(0.0048)
7–12 months			0.5007***
			(0.0048)
13–24 months			0.6115***
			(0.0051)
25–36 months			0.7176***
			(0.0056)
More than 36 months			0.8280***
			(0.0050)
Individual controls	✓	✓	✓
Contract type controls		✓	✓
Employment duration controls			✓
R-squared	0.18	0.33	0.53
Observations	154,803	154,803	154,803

Notes: Linear Probability Model coefficients reported. Robust standard errors in parentheses. Column (1) includes only age indicators; Column (2) adds contract type; Column (3) further adds employment duration. Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.3: JRS Access with Alternative Age Groupings

	(1) Access to JRS	(2) Access to JRS	(3) Access to JRS
Age (ref.: 20–24)			
25–29 years	0.0534***	0.0075**	-0.0020
	(0.0045)	(0.0035)	(0.0026)
30–34 years	0.0680***	-0.0034	-0.0158***
	(0.0044)	(0.0036)	(0.0027)
35–49 years	0.0801***	-0.0019	-0.0220***
	(0.0039)	(0.0031)	(0.0023)
50–65 years	0.0601***	-0.0130***	-0.0517***
	(0.0042)	(0.0034)	(0.0027)
Contract Type (ref.: Temporary)			
Intermittent open-ended		0.2041***	0.0665***
		(0.0064)	(0.0026)
Regular open-ended		0.4371***	0.0140***
		(0.0033)	(0.0026)
Employment Duration (ref.: ≤3 months)			
4–6 months			0.1134***
			(0.0054)
7–12 months			0.5042***
			(0.0055)
13–24 months			0.6230***
			(0.0048)
25–36 months			0.7330***
			(0.0050)
More than 36 months			0.8501***
			(0.0037)
Individual controls	✓	✓	✓
Contract type controls		✓	✓
Employment duration controls			✓
Pseudo R^2	0.17	0.32	0.48
Observations	152,196	152,196	152,196

Notes: Average marginal effects reported. Robust standard errors in parentheses. Column (1) includes only age indicators; Column (2) adds contract type; Column (3) further includes employment duration. Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. Effects measure percentage-point changes in probability relative to the reference category (20–24 years). $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.4: Access to JRS Excluding Sectors Highly Affected by COVID-19

	(1) Access to JRS	(2) Access to JRS	(3) Access to JRS
Age (ref.: 16–34)			
35–50 years	0.0389***	-0.0121***	0.0005
	(0.0037)	(0.0032)	(0.0037)
51 years or more	0.0063	-0.0247***	-0.0140**
	(0.0046)	(0.0039)	(0.0057)
Contract Type (ref.: Temporary)			
Intermittent open-ended		0.1821***	0.0630***
		(0.0097)	(0.0047)
Regular open-ended		0.4852***	0.0519***
		(0.0047)	(0.0047)
Employment Duration (ref.: ≤3 months)			
4–6 months			0.0934***
			(0.0076)
7–12 months			0.5021***
			(0.0079)
13–24 months			0.5594***
			(0.0075)
25–36 months			0.6814***
			(0.0080)
More than 36 months			0.8231***
			(0.0058)
Individual controls	✓	✓	✓
Contract type controls		✓	✓
Employment duration controls			✓
Pseudo R^2	0.18	0.36	0.50
Observations	66,326	66,326	66,326

Notes: Average marginal effects reported. Robust standard errors in parentheses. The specification excludes sectors most directly affected by COVID-19 containment measures (sectors 4, 5, 6, and 10). Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.5: Log-Odds Coefficients on Access to Income Protection after Job Loss

	(1) Income Protection	(2) Income Protection	(3) Income Protection	(4) Income Protection
Age (ref.: 16–34)				
35–50 years	0.6423***	0.6314***	0.6515***	-0.0113
	(0.0311)	(0.0311)	(0.0317)	(0.0417)
51 years or more	0.7237***	0.7092***	0.7223***	-0.4212***
	(0.0423)	(0.0424)	(0.0431)	(0.0633)
Contract Type (ref.: Temporary)				
Intermittent open-ended		0.7665***	0.7682***	0.7060***
		(0.1877)	(0.1924)	(0.1946)
Regular open-ended		0.3028***	-0.0486	-0.0704
		(0.0520)	(0.0596)	(0.0601)
Employment Duration (ref.: ≤3 months)				
4–6 months			0.4001***	0.4336***
			(0.0379)	(0.0384)
7–12 months			0.7938***	0.8308***
			(0.0477)	(0.0481)
13–24 months			1.0669***	1.0733***
			(0.0630)	(0.0637)
25–36 months			1.0790***	1.0841***
			(0.0871)	(0.0882)
More than 36 months			0.7578***	0.7038***
			(0.0799)	(0.0803)
Years since entry into labor market				0.0531***
				(0.0022)
Individual controls	✓	✓	✓	✓
Contract type controls		✓	✓	✓
Employment duration controls			✓	✓
Experience control				✓
Pseudo R^2	0.25	0.26	0.27	0.28
Observations	32,404	32,404	32,404	32,404

Notes: Log-odds coefficients reported. Robust standard errors in parentheses. Column (1) includes only age indicators; Column (2) adds contract type; Column (3) further includes employment duration; Column (4) additionally controls for labor market experience (years since entry into the labor market). Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.6: Linear Probability Model Coefficients on Access to Income Protection after Job Loss

	(1) Income Protection	(2) Income Protection	(3) Income Protection	(4) Income Protection
Age (ref.: 16–34)				
35–50 years	0.1155***	0.1139***	0.1158***	0.0081
	(0.0053)	(0.0053)	(0.0053)	(0.0068)
51 years or more	0.1221***	0.1203***	0.1226***	-0.0727***
	(0.0068)	(0.0068)	(0.0067)	(0.0103)
Contract Type (ref.: Temporary)				
Intermittent open-ended		0.1153***	0.1137***	0.1016***
		(0.0235)	(0.0237)	(0.0235)
Regular open-ended		0.0325***	-0.0158*	-0.0182**
		(0.0073)	(0.0080)	(0.0079)
Employment Duration (ref.: ≤3 months)				
4–6 months			0.0898***	0.0916***
			(0.0066)	(0.0065)
7–12 months			0.1521***	0.1533***
			(0.0072)	(0.0071)
13–24 months			0.1800***	0.1772***
			(0.0083)	(0.0083)
25–36 months			0.1790***	0.1748***
			(0.0106)	(0.0106)
More than 36 months			0.1222***	0.1106***
			(0.0102)	(0.0101)
Potential Experience				
Years since entry into labor market				0.0087***
				(0.0003)
Individual controls	✓	✓	✓	✓
Contract type controls		✓	✓	✓
Employment duration controls			✓	✓
Experience control				✓
R-squared	0.29	0.29	0.30	0.32
Observations	32,640	32,640	32,640	32,640

Notes: Linear Probability Model coefficients reported. Robust standard errors in parentheses. Column (1) includes only age indicators; Column (2) adds contract type; Column (3) further adds employment duration; Column (4) additionally controls for potential labor market experience (years since entry into the labor market). Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.7: Access to Income Protection with Alternative Age Groupings

	(1) Income Protection	(2) Income Protection	(3) Income Protection	(4) Income Protection
Age (ref.: 20–24)				
25–29 years	0.1356***	0.1340***	0.1324***	0.0965***
	(0.0090)	(0.0090)	(0.0089)	(0.0082)
30–34 years	0.1723***	0.1698***	0.1699***	0.0996***
	(0.0094)	(0.0094)	(0.0092)	(0.0094)
35–49 years	0.1917***	0.1886***	0.1901***	0.0802***
	(0.0079)	(0.0079)	(0.0078)	(0.0097)
50–65 years	0.2033***	0.1998***	0.2002***	0.0373***
	(0.0088)	(0.0088)	(0.0087)	(0.0129)
Contract Type (ref.: Temporary)				
Intermittent open-ended		0.1146***	0.1105***	0.1050***
		(0.0248)	(0.0248)	(0.0253)
Regular open-ended		0.0387***	-0.0132	-0.0136
		(0.0082)	(0.0097)	(0.0096)
Employment Duration (ref.: ≤3 months)				
4–6 months			0.0718***	0.0718***
			(0.0065)	(0.0065)
7–12 months			0.1277***	0.1277***
			(0.0074)	(0.0074)
13–24 months			0.1567***	0.1567***
			(0.0088)	(0.0088)
25–36 months			0.1585***	0.1585***
			(0.0117)	(0.0117)
More than 36 months			0.1105***	0.1105***
			(0.0120)	(0.0120)
Potential Experience				
Years since entry into labor market				0.0058***
				(0.0004)
Individual controls	✓	✓	✓	✓
Contract type controls		✓	✓	✓
Employment duration controls			✓	✓
Experience control				✓
R-squared	0.25	0.25	0.26	0.27
Observations	31,154	31,154	31,154	31,154

Notes: Average marginal effects from logit models reported. Robust standard errors in parentheses. The reference age group is 20–24 years. Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Table A.8: Average Marginal Effects on Access to Income Protection after Job Loss

	(1) Income Protection	(2) Income Protection	(3) Income Protection	(4) Income Protection
Age (ref.: 16–34)				
35–50 years	0.1140***	0.1118***	0.1139***	0.0171*
	(0.0075)	(0.0075)	(0.0075)	(0.0093)
51 years or more	0.1398***	0.1378***	0.1388***	-0.0211
	(0.0093)	(0.0093)	(0.0092)	(0.0141)
Contract Type (ref.: Temporary)				
Intermittent open-ended		0.1113***	0.0931***	0.0866***
		(0.0331)	(0.0334)	(0.0335)
Regular open-ended		0.0797***	0.0107	0.0064
		(0.0126)	(0.0150)	(0.0150)
Employment Duration (ref.: ≤3 months)				
4–6 months			0.0728***	0.0789***
			(0.0088)	(0.0087)
7–12 months			0.1272***	0.1312***
			(0.0106)	(0.0105)
13–24 months			0.1909***	0.1899***
			(0.0119)	(0.0120)
25–36 months			0.2049***	0.2044***
			(0.0163)	(0.0164)
More than 36 months			0.1750***	0.1667***
			(0.0173)	(0.0177)
Years since entry into labor market				0.0068***
				(0.0005)
Individual controls	✓	✓	✓	✓
Contract type controls		✓	✓	✓
Employment duration controls			✓	✓
Experience control				✓
Pseudo R^2	0.2	0.24	0.26	0.27
Observations	15,916	15,916	15,916	15,916

Notes: Average marginal effects reported. Robust standard errors in parentheses. Additional controls include education, gender, nationality, sector, region, working-time arrangement, and reason for job separation. Effects measure percentage-point changes in probability relative to the reference category. $^{*}p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$.