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Conference Paper

Stumbling Blocks or Stepping Stones: Long-Term Effects of Promoting (Youth) Entrepreneurship

Beiträge zur Jahrestagung des Vereins für Socialpolitik 2025: Revival of Industrial Policy

Provided in Cooperation with:
Verein für Socialpolitik / German Economic Association

Suggested Citation: Camarero Garcia, Sebastian; Henao Bermudez, Leandro (2025) : Stumbling Blocks or Stepping Stones: Long-Term Effects of Promoting (Youth) Entrepreneurship, Beiträge zur Jahrestagung des Vereins für Socialpolitik 2025: Revival of Industrial Policy, ZBW - Leibniz Information Centre for Economics, Kiel, Hamburg

This Version is available at:
https://hdl.handle.net/10419/325367

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Stumbling Blocks or Stepping Stones

Long-Term Effects of Promoting (Youth) Entrepreneurship

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February 2025

Abstract: This paper evaluates the effects of Spain's 2013 Entrepreneurship Act subsidy on young entrepreneurs, focusing on gender- and age-specific labor market outcomes. Using regression discontinuity and propensity score matching, we assess the policy's impact on wages, employment probability, and industry upon transitioning to dependent employment. Our findings show that while the subsidy did not extend self-employment spells or improve overall employment probabilities, it significantly boosted wages for women, particularly those with family responsibilities. The policy delayed self-employment exit for high-risk groups, providing them with more time to secure better job matches, especially in the services sector.

JEL codes: J21, J24, J68.

Keywords: entrepreneurship, youth unemployment, regression discontinuity, propensity score matching

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1. Introduction

Entrepreneurial failure may result in a transition to dependent employment, affecting labor supply and labor market flows.[1] During economic downturns, policymakers often respond with wage subsidies, enhanced unemployment benefits, and targeted support for the self-employed—a particularly vulnerable group (Blundell et al., 2021). Labor market policies during crisis typically fall into two categories: active labor market policies, which aim to improve job matches and reemployment prospects, and passive labor market policies, which provide income support but may weaken job search incentives (Calmfors, 1994).

This paper examines how a self-employment subsidy introduced in Spain in 2013 affected employment outcomes, particularly in light of these competing mechanisms. According to standard search theory, self-employment subsidies may alter reservation wages and job acceptance decisions, potentially distorting occupational choices. By lowering the cost of entrepreneurship, such policies may encourage individuals to pursue self-employment instead of continuing their job search—even when wage employment might offer better long-term prospects. This selection mechanism complicates the analysis of survival rates between necessity-driven and opportunity-driven entrepreneurs, a distinction framed by the push-pull dichotomy. This framework distinguishes *pull* entrepreneurs, who pursue opportunities, from *push* entrepreneurs, who enter self-employment due to external constraints such as limited job availability (Caliendo et al., 2023).

Despite substantial literature on self-employment drivers (Camarero Garcia & Hansch, 2024; Camarero Garcia & Murmann, 2024) and the outcomes of established entrepreneurs (Humphries, 2022; Shaw & Sørensen, 2019), fewer studies explore the long-term labor market consequences for individuals who transition from entrepreneurship to wage labor. Understanding these transitions is crucial for evaluating whether self-employment policies improve long-term career trajectories or merely delay entry into wage employment. This paper addresses this gap by analyzing the welfare effects of Spain's 2013 reform, which subsidized young entrepreneurs' social security contributions for up to 30 months.

Relevant to our work, Cueto et al. (2017) assess the effects of a flat-rate subsidy for new entrepreneurs in Spain, which had no significant impact on self-employment survival rates. In contrast, Sánchez-Cañizares et al. (2020) find higher survival rates among subsidized entrepreneurs in Andalusia, where unemployment rates were higher, although these effects diminish over time. Similar studies in Germany (Caliendo et al., 2015) and France (Bourlès & Cozarenco, 2018) explore the impact of entrepreneurship subsidies and reveal mixed results, particularly concerning business survival and income levels.

When considering the long-term consequences of entrepreneurship, Mahieu et al. (2021) find that former entrepreneurs in Belgium earn significantly less upon returning to dependent employment, whereas

[1] New businesses in Europe are relatively short-lived (see Figure **A1** in the Appendix). Even during recovery, only about 40% of newly established enterprises in Spain survived beyond five years. In contrast, survival rates in the United States were slightly higher than in the European Union, with 71% of firms lasting three years in 2013 (OECD, 2022).

Daly (2015) reports no earnings penalty for American workers who return to paid employment. Other studies highlight wage penalties for former entrepreneurs in Denmark (Kaiser & Malchow-Møller, 2011) and Spain (Garcia-Cabo & Madera, 2024). Lastly, Hyytinen & Rouvinen (2008) add that individuals with shorter entrepreneurial spells tend to have lower productivity, contributing to observed wage gaps.

Our study leverages high-quality administrative data from the Spanish Social Security Administration to examine the reform's long-term impacts on employment probability, wage rates, job tenure, and sector. We exploit the natural experiment created by the Spanish Entrepreneurship Act (2013), which introduced a discontinuity in social security contributions for young entrepreneurs based on age and sex. Specifically, the reform made male self-employed individuals eligible for reduced contributions if they were 29 years old or younger, and female self-employed individuals eligible if they were 34 years old or younger, provided they had an active self-employment spell in February 2013. This eligibility criterion allows us to estimate the causal effects of the reform on various employment outcomes.

We complement the regression discontinuity analysis with an evaluation of the effects of short self-employment spells on subsequent labor market outcomes using propensity score matching. This additional approach sheds some light on the mechanisms driving the estimated effects, by comparing individuals with similar observable characteristics who experienced different employment trajectories. Understanding the consequences of short-lived entrepreneurial spells is crucial, as prolonged non-traditional employment can affect future job matches. By examining both the direct effects of Spain's 2013 reform and the broader employment trajectories of former entrepreneurs, this paper analyzes whether self-employment serves as a stepping stone for some sociodemographic groups.

Our findings indicate that the policy did not extend self-employment spells or improve overall employment probabilities for young entrepreneurs on average. Moreover, while male entrepreneurs saw no significant earnings change, women moving to dependent employment within two years experienced a 14% earnings increase, primarily due to transitions into better-paying service sector jobs. These effects were strongest for women with higher unemployment risk, such as those with children or from rural areas.

For women with children, the subsidy delayed their exit from self-employment by three months compared to untreated women. This suggests the policy acted as an active labor market policy, enabling some jobseekers to hold out for better matches—consistent with job search theory (Card et al., 2010). Women who received the subsidy were less likely to accept poor job matches, while untreated women may have been forced into lower-quality employment. By contrast, men who delayed their return to dependent employment suffered income losses. Propensity score matching results show that post-entrepreneurial wages for men are highly dependent on the business cycle. This aligns with evidence that self-employment incentives can yield mixed results, depending on labor market conditions and individual characteristics.

2. Institutional Framework

2.1. Labor Market and Social Security System

Spain is characterized by a dual labor market, where a stark divide exists between permanent workers, who enjoy strong employment protection, and temporary workers, who face frequent job turnover and minimal dismissal costs (Bentolila et al., 2012). This segmentation not only contributed to extreme labor market volatility but also shaped transitions into self-employment, as temporary workers, facing precarious employment prospects, may have been more likely to consider self-employment as an alternative.

Spain's self-employed contribute to the RETA system, which imposes a fixed minimum social security contribution regardless of income, unlike earnings-based systems in the United States and other countries. In 2013, this minimum was EUR 255/month, creating a significant financial burden for low-income entrepreneurs. Unlike the flat-rate scheme, which drastically reduced contributions for new entrepreneurs, the 2013 subsidy we analyze targeted incumbent self-employed individuals, providing a partial relief on mandatory social security payments.

2.2. The Entrepreneurship Act

The Entrepreneurship Act (2013) was introduced to combat high youth unemployment and stimulate self-employment. Unlike other initiatives aimed at encouraging business creation, this policy specifically targeted incumbent entrepreneurs—self-employed individuals who were already operating their businesses. The policy provided financial relief by reducing mandatory social security contributions for a limited period, rather than eliminating them entirely as other schemes, such as the flat-rate policy, had done.

The subsidy's eligibility period, denoted as M, consisted of two phases: the reduction period and the bonus period. The reduction period lasted for the first 15 months following the worker's entry into self-employment. After this, the bonus period provided an additional 15 months of reduced contributions.[2] By October 2015, the subsidy had completely phased out. The monthly subsidy, which was fixed for all self-employed workers, did not vary based on individual contribution bases. It was calculated as 30% of the social security contribution resulting from applying the minimum rate r_t to the minimum contribution base b_t in effect during month t. The accumulated subsidy S naturally depended on the length of the eligibility period M as shown in Equation (1):

$$S = \sum_{t=1}^{M} (0.30 \times r_t \times b_t) \qquad (1)$$

[2] For instance, an entrepreneur who became self-employed in January 2013 would be eligible for the subsidy until September 2015 ($M = 15 + 15$ months). In contrast, someone who started in January 2012 would only qualify until June 2014 ($M = 3 + 15$ months).

Importantly, eligibility was restricted to individuals who had started their self-employment activity after October 2007 and were still actively self-employed as of February 2013. Age restrictions also applied, with male beneficiaries required to be 29 years old or younger and female beneficiaries 34 years old or younger at the time of the policy's introduction.

Figure **A2** in the Appendix illustrates the accumulated subsidy amount in our sample by assignment rules. On average, beneficiaries received a subsidy of approximately EUR 77/month in 2013. About 54% of recipients were entitled to 15 months of reduced contributions, receiving am accumulated subsidy of approximately EUR 1,140. Only 1.7% of beneficiaries qualified for the maximum subsidy period of 30 months, receiving EUR 2,300. The distribution of the total subsidy narrows as individuals approach the eligibility cut-off, with many entitled only to the bonus period of 15 months.

Although there are no official reports on the actual number of beneficiaries, we provide an estimation of the public spending on this policy based on other statistical sources. In 2013 there were about 1.9 million self-employed workers in Spain, of which one third were female.[3] With an estimated 290,000 beneficiaries, the policy represented a EUR 330 million (0.033% of GDP) investment—a relatively modest share of spending on active labor market policies, which was EUR 4,000 million in 2013 according to Cueto & Suárez (2014). However, its effectiveness remains uncertain, as the subsidy's impact on self-employment duration and labor market outcomes varied by gender and individual characteristics.[4]

2.3. Other Policies to Promote Self-Employment

The subsidy targeting young entrepreneurs that we analyse in this paper, while critical, was part of a broader set of policies introduced in response to the severe economic crisis that Spain had been facing since 2008. This crisis underscored the structural weaknesses of the Spanish labor market, which experienced more significant job losses, at a faster pace, than other major European economies. Young people under 25 were disproportionately affected, with their unemployment rate nearing 50%.

In response to these trends, the Spanish government implemented various support measures aimed at mitigating the high youth unemployment rate and encouraging entrepreneurship among younger individuals. In February 2012, the law was modified to allow unemployed individuals to receive a one-time payment of 60% of their remaining contributory unemployment benefits, specifically for the purpose of starting a business. For young men under 30 and women under 35, this cap was increased to 100%.

[3] Spanish legislation differentiates between *own-account* workers and *self-employed* individuals. Our focus is on the latter—those not affiliated with commercial companies, cooperatives, or other corporate entities. We also exclude individuals registered as collaborating family members or as part of a special worker group.

[4] We estimate that 140,000 self-employed men under 30 and 150,000 self-employed women under 35 were eligible for the subsidy. These figures are based on official self-employment data by age and gender (Ministry of Labor and Social Economy, 2013), assuming a uniform distribution within the broader age brackets (under 25 and 25–39 years). Given that most beneficiaries qualified for 15 months at EUR 77 per month, total public spending amounted to nearly EUR 330 million.

In September 2013, a law introduced several initiatives aimed at promoting self-employment among young people under 30. These included the introduction of reduced initial social security contributions, the ability to combine unemployment benefits with income from a new business, and an expanded scope for capitalizing unemployment benefits to fund entrepreneurial activities.

The overlap of multiple entrepreneurship policies complicates identifying isolated effects on self-employment entry. However, by focusing on incumbent entrepreneurs, our analysis mitigates this issue, as the policy we evaluate applied exclusively to individuals already in self-employment rather than new entrants.

3. Data and Descriptive Statistics

3.1. Labor Market Trajectories: The CWLS Sample

Our analysis is based on the Continuous Working Life Sample (CWLS), an individual-level dataset compiled by the Spanish Social Security Administration (2021). The dataset integrates fiscal records from the Spanish Tax Authorities and personal information from the National Census, provided by the Spanish Statistics Institute. The CWLS contains detailed administrative information on socioeconomic characteristics and labor market statuses, offering longitudinal data for a non-stratified random sample of 4% of Spain's population from their first registration in the social security system. It provides reliable employment histories retrospectively, extending back to the 1980s.

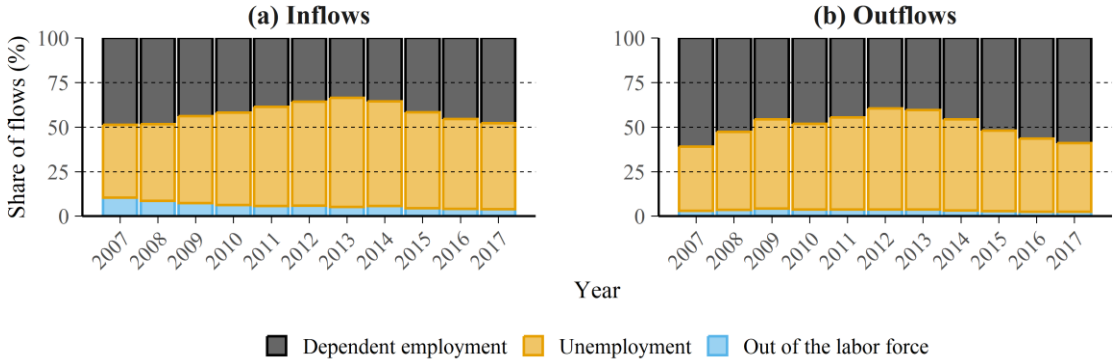
The CWLS has been released in 16 waves, from the first in 2005 to the most recent in 2020. The retention of anonymized individual identifiers across waves allows us to construct a consistent panel dataset, mitigating the sample attrition issues commonly found in survey data. This feature enables us to construct a representative dataset with extensive longitudinal coverage, which is crucial for our analysis of self-employment dynamics. For further details on the CWLS data and variables, as well as the construction of our dataset, see Camarero Garcia & Hansch (2024).

Using CWLS data, we document the yearly transitions into and out of self-employment (Figure **1**). The observed counter-cyclical trend in transitions from unemployment to self-employment aligns with the recession-push hypothesis, suggesting that economic downturns increase necessity-driven entrepreneurship (Fairlie & Fossen, 2020). During the crisis, job scarcity likely compelled many individuals to turn to self-employment as an alternative to unemployment, as evidenced by the sharp rise in transitions from unemployment, peaking at 60% in 2013. In contrast, as the economy recovered, self-employment entries from dependent employment increased, suggesting a shift towards opportunity-driven entrepreneurship.

A similar pattern is observed for transitions out of self-employment in Panel (b). Between 40% and 75% of individuals who exited self-employment moved into dependent employment. This high share

highlights the importance of examining the labor market outcomes of former self-employed individuals, particularly their reintegration into wage employment.

Figure 1: *Inflows to and Outflows from Self-Employment*



Notes: This figure depicts annual (a) inflows to and (b) outflows from self-employment in Spain as a percentage of the total. The sample includes all individuals aged 18 and older. Inflows and outflows are categorized by employment status: Dependent Employment, Unemployment, and Out of the Labor Force.
Source: Author's calculations based on CWLS data.

Figure **A3** in the Appendix further describes the overall distribution of labor market statuses in Spain between 2007 and 2017. The figure presents both relative and absolute numbers of individuals in dependent employment, self-employment, unemployment, and out of the labor force for each year. It shows that the share of self-employed workers remains relatively stable at around 15% throughout most of the sample period. Moreover, the unemployment rate shows a marked increase from 2007 to 2013, coinciding with the global financial crisis, while the share of wage-employed individuals correspondingly declines. Following 2013, there is a steady recovery in wage employment. These trends provide valuable context for understanding the evolving composition of the labor market in Spain, particularly the role of self-employment during periods of economic fluctuation.

3.2. Descriptive Statistics

Table **1** presents the summary statistics for the sample of self-employed workers between 25 and 40 who had an active self-employment spell at the time the Entrepreneurship Act was enacted. In February 2013, at the time of the subsidy's introduction, the average age of self-employed individuals in our sample was 33.5 years. Notably, one-fifth of the sample were non-Spanish nationals, and 45% had no children at the time. About one fourth of the sample had completed a higher education degree, reflecting the diverse educational background of the self-employed population.

Table 1: *Descriptive Statistics*

	All individuals		Female	Male	Mean test	
	Mean	Std. dev.	Mean	Mean	Diff.	Std. err.
Variable	(1)	(2)	(3)	(4)	(5)	(6)
(a) Pre-determined characteristics						
Age in February 2013	33.52	(4.02)	33.39	33.60	-0.22***	(0.06)
Immigrant	0.20	(0.40)	0.21	0.20	0.01*	(0.01)
Family responsibilities						
No children	0.45	(0.50)	0.38	0.50	-0.11***	(0.01)
Formal education						
Higher	0.24	(0.43)	0.31	0.19	0.12***	(0.01)
(b) Previous labor market experience						
Experience in dep. empl. (days)	1,931.50	(1,564.11)	1,670.40	2,100.43	-430.03***	(21.74)
Previous status						
Full-time employed	0.26	(0.44)	0.22	0.28	-0.06***	(0.01)
Part-time employed	0.11	(0.32)	0.15	0.09	0.06***	(0.00)
Non-employed	0.61	(0.49)	0.61	0.60	0.00	(0.01)
Earnings (EUR/month)	1,418.49	(904.01)	1,266.39	1,507.80	-241.41***	(16.45)
Spell duration (days)	330.20	(537.89)	358.91	312.26	46.65***	(8.21)
(c) Self-employment spell						
Self-reported earnings (EUR/month)	907.64	(176.58)	900.84	912.04	-11.19***	(2.44)
Sector						
Agric., extr., prim. manuf.	0.07	(0.26)	0.05	0.08	-0.03***	(0.00)
Manuf. and utilities	0.03	(0.17)	0.02	0.04	-0.02***	(0.00)
Construction	0.11	(0.32)	0.02	0.17	-0.15***	(0.00)
Trade	0.25	(0.43)	0.28	0.23	0.05***	(0.01)
Transport and storage	0.05	(0.22)	0.02	0.07	-0.05***	(0.00)
Acom. and food services	0.13	(0.33)	0.14	0.12	0.02***	(0.00)
I&C, fin., ins., RE, and sc. services	0.16	(0.36)	0.17	0.15	0.01***	(0.01)
Educ., health, soc., aux. and other services	0.20	(0.40)	0.30	0.13	0.17***	(0.01)
Spell duration (days)	2,275.65	(1,369.82)	2,197.66	2,326.11	-128.45***	(19.62)
(d) Macroeconomic context						
Municipality size						
Large city	0.13	(0.34)	0.14	0.13	0.01**	(0.00)
Unemployment rate	21.05	(7.37)	21.11	21.00	0.11	(0.11)
(e) Number of observations	20,174		7,925	12,249		

Notes: This table presents the mean and standard deviation for a selection of variables for the overall sample, as well as for female and male workers separately. It also includes a mean test for differences between groups. All samples are restricted to individuals who were between 25 and 40 years old and had an active self-employment spell in February 2013.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Source: Authors' calculations based on CWLS data.

There are some notable gender differences in the characteristics of self-employed individuals. On average, women were more likely to have family responsibilities and had a higher likelihood of having a higher education degree compared to their male counterparts. Panel (b) of Table **1** highlights differences in previous labor market experience. Male self-employed individuals had accumulated 430 more days of experience in dependent employment than their female counterparts. Interestingly, 61% of the individuals in our sample were non-employed prior to becoming self-employed, which supports the importance of necessity-driven entrepreneurship, particularly during economic downturns.

Those who were employed before transitioning to self-employment were earning an average of EUR 1,400/month. There are statistically significant differences in wages between genders, with men earning an average of EUR 1,500/month compared to EUR 1,270/month for women. This wage gap carried over into

their entrepreneurial ventures. In terms of earnings during the relevant self-employment spell, the average monthly income reported was EUR 907, which is slightly above the legal minimum contribution base at the time (EUR 858.60/month). Male workers reported marginally higher earnings (EUR 912/month) compared to female workers (EUR 900/month).

The majority of self-employed individuals operated in the services sector, with a quarter engaged in trade and 20% in administrative, support, and auxiliary services. Gender differences are notable, particularly in construction, where men were overrepresented (17% vs. 2%).

Additionally, over 43% of the individuals in the sample resided in large or medium-sized cities, indicating a potential urban concentration of entrepreneurial activity. The regional unemployment rate at the time of the subsidy, February 2013, was notably high at 21%, underscoring the difficult economic conditions under which many of these individuals entered or continued their self-employment journey.

4. Empirical Strategy

4.1. Identification

We evaluate a reform that partially subsidized social security contributions for established self-employed workers—individuals who had recently entered the system. The policy targeted the self-employed registered after October 2007, provided they were no older than 34 (women) or 29 (men) as of February 2013. This setup is well-suited for a sharp regression discontinuity (RD) design, where the treatment, denoted by D, is exogenously determined by a threshold function of an observable variable X (age, observed on a monthly basis): $D = \mathbb{1}\{X \leq c\}$. In particular, as the cutoff differs between female and male workers, we estimate these RD models separately.

Naturally, the probability of being eligible for the subsidy changes from one to zero at the cutoff, creating a quasi-experimental setting that allows for causal inference using a RD design. We discuss the formal conditions that guarantee the comparability of treatment and control units at or near the respective threshold, specifically the continuity of the outcome variable and the positive density of the forcing variable around the cutoff (Hansen, 2022). If these assumptions are fulfilled, this discontinuous change introduced by the reform can be used to quantify the causal effects of the policy on outcome Y for units with ages at or near the cutoff. In other words, workers with ages just above the cutoff can be used as a comparison group for units with scores barely below it (Cattaneo & Titiunik, 2022).

For identification of local average treatment effects (ATE), the first condition refers to the regression functions $\mathbb{E}[Y_i(0)|X]$ and $\mathbb{E}[Y_i(1)|X]$, which are assumed to be continuous in $x = c$. In other words, the conditional expectations of both treatment and control groups evolve continuously with the forcing variable X. The second condition relates to the density of the forcing variable, which must be positive near the cutoff.

These assumptions reflect the idea that workers just below and just above the respective cutoff age would display the same average effects absent the treatment.

Consequently, any difference between the average responses of treated and control units at the cutoff can be attributed to the treatment and interpreted as the causal average effect of the treatment at the cutoff, that is, for units with score variable $X_i = c$. Equation (2) describes the canonical treatment effect in a sharp RD design (Cattaneo & Titiunik, 2022). The result asserts that the treatment effect θ_{RD} is identifiable as the vertical distance at the cutoff between two conditional expectations that can be estimated from the data.

$$\theta_{RD} \equiv \mathbb{E}[Y_i(1) - Y_i(0)|X = c] = \lim_{x \downarrow c} \mathbb{E}[Y_i|X_i = x] - \lim_{x \uparrow c} \mathbb{E}[Y_i|X_i = x] \qquad (2)$$

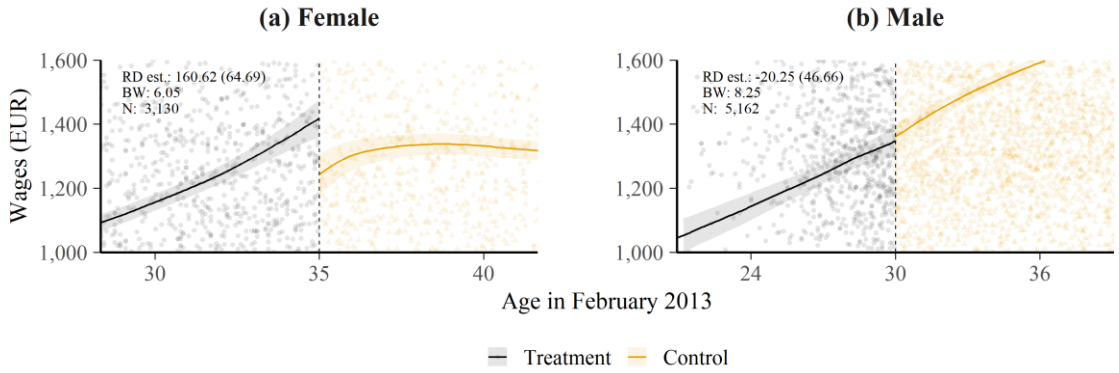
4.2. Estimation

Following standard recommendations (Cattaneo & Titiunik, 2022; Hansen, 2022), we employ local polynomial methods for estimation and inference. Local polynomial analysis for RD designs is implemented by fitting Y_i on a low-order polynomial expansion of X_i, separately for treated and control observations, and in each case using only observations near the cutoff rather than all available observations, as determined by the choice of a kernel function and a bandwidth parameter. We utilize a local polynomial of order 1—i.e., local linear regression (LLR)—with a triangular kernel, which prioritizes observations closer to the threshold to improve local fit.

The choice of bandwidth h is not trivial, as empirical results are usually sensitive to changes in this parameter. The bandwidth introduces a bias-variance trade-off: a larger h increases bias by making the local linearity assumption less plausible, while a smaller h reduces statistical power, increasing variance. To address this trade-off, we adopt the fully data-driven bandwidth selection method introduced by Imbens & Kalyanaraman (2012)—hereafter, I-K, which is asymptotically optimal for RD designs.

In short, given a kernel function $K(\cdot)$ and a bandwidth h, the RD point estimator is obtained by fitting two weighted least-squares regressions of Y_i on a polynomial expansion of X_i of order 1. The RD estimator of the local ATE, θ_{RD}, corresponds to the estimated discontinuity at the cutoff, representing the difference in intercepts. To illustrate these effects, Figure **2** presents the RD design applied to our primary outcome: monthly wages in dependent employment measured one year post-policy introduction. For this illustration, we plot the raw outcomes and superimpose the predicted values from our preferred LLR specification. We also report the estimated θ_{RD}, the bandwidth choice and the effective number of observations. Visual and statistical evidence suggest that the policy increased post-self-employment wages for female entrepreneurs, while no significant effects are observed for men.

Figure 2: *Regression Discontinuity Plot for Wages in Dependent Employment*



Notes: This figure illustrates the relationship between the running variable (age in February 2013) and the outcome (monthly wages in EUR). It is based on local linear regressions performed on either side of the cutoff using the Imbens-Kalyanaraman optimal bandwidth and a triangular kernel. The RD estimate is reported with standard error in parentheses, along with the chosen bandwidth (*BW*) and the effective number of observations (*N*).

Both subsamples are restricted to individuals with an active self-employment spell in February 2013 and an active dependent employment spell one year later.

Source: Authors' calculations based on CWLS data.

5. Policy Evaluation

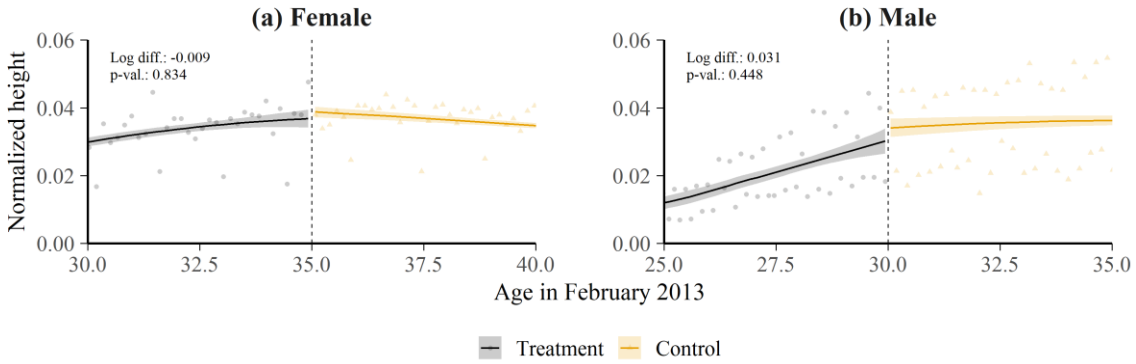
5.1. Strategy Validation

The key identification assumption in a regression discontinuity design is that the conditional expectations of the treated and untreated outcomes are continuous at the cutoff. This assumption can be violated if individuals are able to manipulate the running variable—age, in this case. Individuals might attempt to manipulate eligibility by strategically delaying their self-employment registration to qualify for longer periods. If such behavior were widespread, we would expect to observe excess mass just below the cutoff, leading to a discontinuity in the density of the running variable.

To assess the plausibility of this assumption, we apply the sorting test proposed by McCrary (2008), which examines the density of the running variable at the threshold. This test evaluates whether the density function of X is continuous at $x = c$. The test uses a histogram estimator with narrow bin widths, and a local linear regression model is implemented to detect any evidence of sorting near the cut-off. The test estimates the discontinuity in the frequency of observations near the threshold by regressing histogram heights on the bin midpoints (Hansen, 2022).

Figure **3** depicts the histogram heights and the estimated density for both sides of the cut-off for female in Panel (a) and male workers in Panel (b). Additionally, we report the estimated difference in the density at the cut-off and the associated p-value. The results indicate that there is no significant bunching of workers below the cut-off, suggesting that manipulation of the running variable is unlikely. This supports the validity of the RD model for estimating the treatment effects of the policy.

Figure 3: *McCrary Sorting Test*



Notes: This figure presents the results of the sorting tests for the running variable (age in February 2013), as proposed by McCrary (2008). We report the estimated difference in histogram heights and the p-value for each density discontinuity test in the female (Panel a) and male (Panel b) subsamples.
Both subsamples are restricted to individuals with an active self-employment spell in February 2013.
Source: Authors' calculations based on CWLS data.

To further ensure robustness, we run an alternative manipulation test (Figure **A4** in the Appendix). We perform the data-driven density test suggested by Cattaneo et al. (2020), based on their local polynomial density estimator. This alternative test accounts for potential biases in histogram-based methods, reinforcing that the density of the running variable is continuous at the cut-off for both male and female subsamples.

Beyond manipulation tests, we assess whether predetermined characteristics are balanced across the cutoff, ensuring the comparability of treatment and control groups. In a valid RD design, predetermined covariates should evolve smoothly across the cutoff, as treatment assignment is based solely on the running variable. Figure **A5** in the Appendix provides visual tests of the continuity of selected covariates for female and male workers who had an active dependent employment spell one year after the policy implementation.[5]

The results show that, as expected, socioeconomic and pre-treatment characteristics exhibit a continuous distribution around the cutoff. The estimated RD parameters for these covariates are minimal and statistically insignificant, indicating that these characteristics are balanced across the threshold. While we detect statistically significant differences in higher education among female workers and trade-sector self-employment among male workers, these imbalances do not necessarily compromise the validity of our RD design. Importantly, balance is maintained across a broad set of other key socioeconomic and pre-treatment characteristics, reducing concerns about systematic differences between treated and control units. Finally, RD estimates rely on the assumption that potential outcomes evolve smoothly around the cutoff, and minor imbalances in predetermined covariates do not necessarily indicate a violation of this condition.

[5] Analogous tests for alternative time windows are available upon request.

5.2. Effects on Labor Market Outcomes

Figure **A6** in the Appendix examines whether the subsidy policy extended self-employment durations for treated entrepreneurs relative to their non-subsidized peers. It presents conditional means for the duration (in days) of self-employment spells, measured from start to exit.[6] We would expect that, especially during the first months of self-employment, relaxing financial constraints would help entrepreneurs sustain their businesses for longer periods. In particular, individuals might delay exiting self-employment until they have fully utilized the subsidy (lock-in effect).

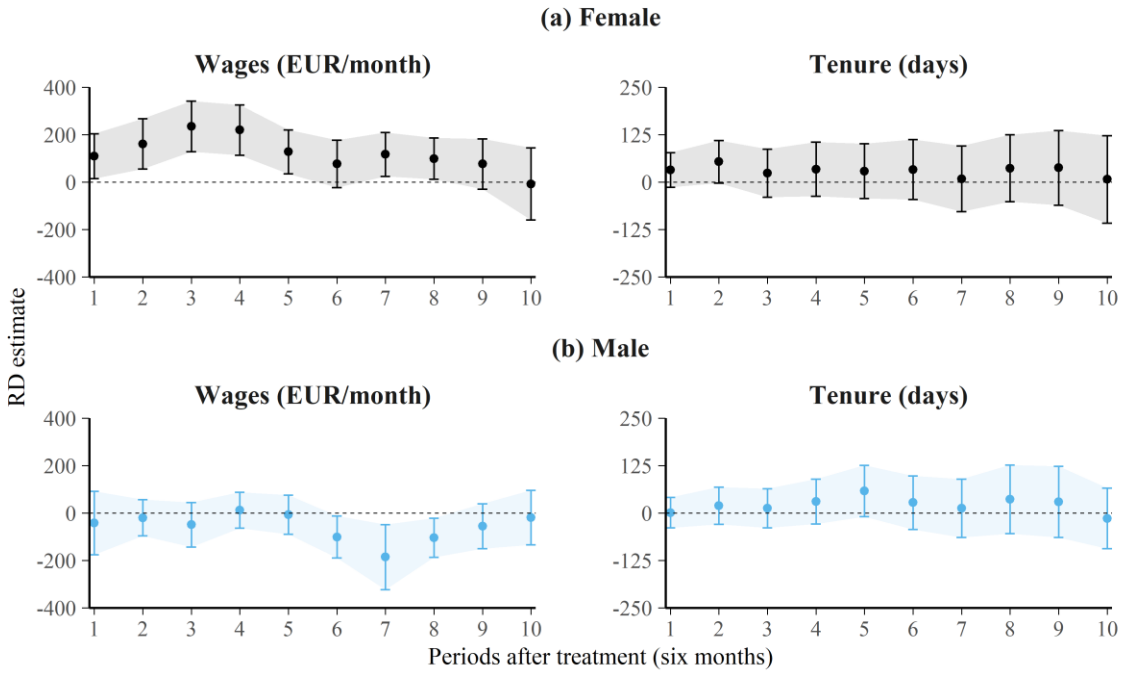
However, the evidence shows no significant extension in average spell duration among male workers. While female workers exhibit a positive effect of almost 60 days, this increase is not statistically significant. These null effects on business survival are consistent with Cueto et al. (2017), who attribute the limited impact of the Spanish program to its broad design. Unlike more targeted start-up initiatives, the Spanish government implemented a program with minimal entry restrictions, making it accessible to large groups. This lack of selectivity may have diluted its effectiveness compared to international programs with stricter eligibility criteria.

We also present RD estimates for the probability of having an active dependent employment spell one year after the treatment (Figure **A8** in the Appendix). Following a human capital and skill development mechanism, prolonged short-term self-employment spells may lead some individuals to acquire skills, industry knowledge, and professional networks that increase their employability (Williams, 2000). However, in line with the effects on self-employment duration, we find no significant impact of the subsidy on average employment probabilities.

We next examine post-subsidy outcomes in dependent employment, focusing on wage rates and job tenure in the subsequent spell. Figure **4** reports estimated treatment effects on monthly wages and job tenure in the first five years following the treatment. Former female entrepreneurs who transitioned into dependent employment within three years after the treatment saw an increase in monthly wages of up to EUR 230—18%, considering earnings in the control group of about EUR 1,300. In contrast, the positive effects on job tenure of up to 53 days are not statistically significant. Notably, wages of former male entrepreneurs remained unaffected by the subsidy during the same period. However, they faced a drop in monthly wages of up to EUR 190 if they extended their self-employment spell beyond that window.

[6] Figure **A7** presents analogous results for the number of days between the treatment and exit dates. The positive effect of 49 days for female entrepreneurs is not statistically significant at conventional levels.

Figure 4: *Effects on Post-Subsidy Outcomes in Dependent Employment*



Notes: This figure presents the regression discontinuity estimates for a selection two outcome variables—wage rates and job tenure—with point estimates depicted alongside 90% confidence intervals. Results are reported for subsamples defined by the number of six-month periods after the treatment in February 2013. Estimates are based on local linear regressions performed on either side of the cutoff using the Imbens-Kalyanaraman optimal bandwidth and a triangular kernel.

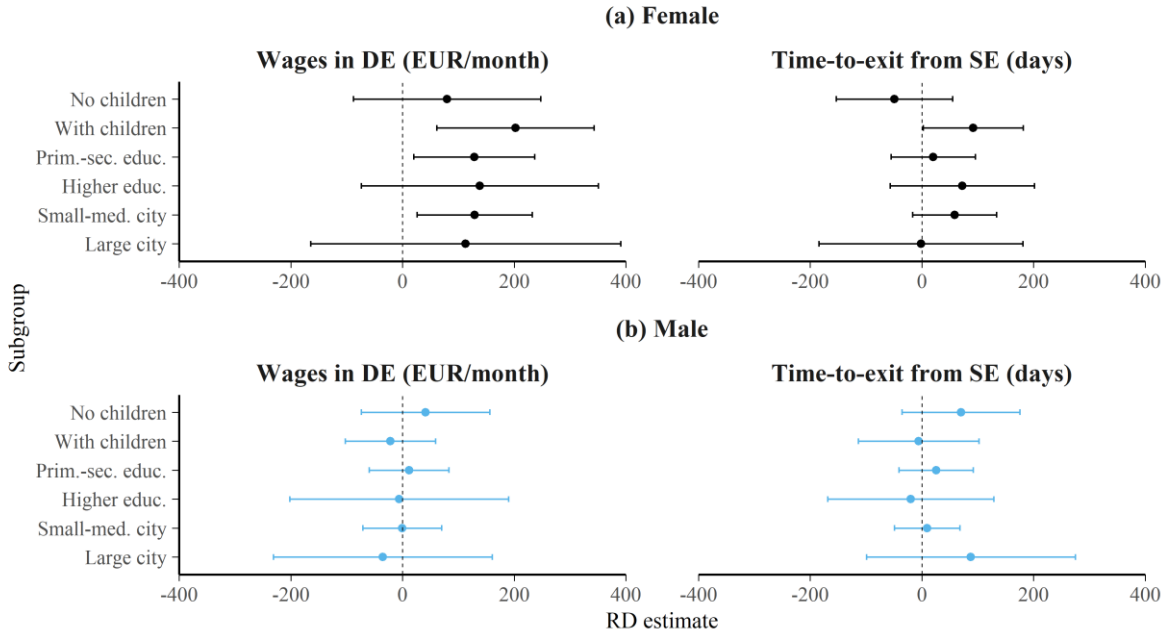
All subsamples are restricted to individuals with an active self-employment spell in February 2013 and an active dependent employment spell in the relevant period.

Source: Authors' calculations based on CWLS data.

To explore the mechanisms behind these effects, we now analyze the effects of the subsidy on the sector in which former entrepreneurs find their jobs. Figure **A9** in the Appendix shows that subsidized women were more likely to transition into dependent employment in the services sector within three years. Consistently, they were also less likely to find jobs in lower-paying sectors like agriculture. Former male entrepreneurs exhibit much smaller and mostly insignificant effects.

Figure **5** provides insights on the policy's heterogeneous impact on self-employed workers. The positive wage effects for former female entrepreneurs are concentrated among those at higher risk of unemployment: women with family responsibilities, lower educational attainment, and those from mid-sized cities and rural areas. Thus, the reform facilitated better job matches for groups experiencing high labor market instability. Although the policy did not increase overall business survival rates, Figure **5** suggests it may have delayed self-employment exit for certain groups. The subsidy provided young women with children with more time—about 90 days—to ensure a successful transition to dependent employment, compared to untreated units.

Figure 5: *Effect Heterogeneity*



Notes: See Figure **4**. This figure focuses on two outcome variables—wage rates in subsequent dependent employment and time-to-exit, which measures the duration from treatment to the end of the self-employment spell.

Results are presented for subgroups of former entrepreneurs, categorized by family responsibilities (with vs. without children), formal education level (primary/secondary vs. higher education), and municipality size (small/medium vs. large city).

Source: Authors' calculations based on CWLS data.

6. Short Self-Employment Spells

6.1. Conceptual Framework

To assess the impact of self-employment promotion among young individuals, we supplement our RD analysis with a matching approach. Short self-employment spells can shape long-term employment outcomes through multiple mechanisms, including human capital development and signaling effects. Williams (2000) highlights that entrepreneurial experience may enhance future wages via skill acquisition (learning-by-doing), sector-specific knowledge, and broader employability signals (Kaiser & Malchow-Møller, 2011). However, self-employment may also lead to diminished future earnings if it signals labor market instability or entrepreneurial failure (Bruce & Schuetze, 2004). This signaling effect likely varies across sociodemographic groups, particularly in terms of gender and prior labor market standing.

Mahieu et al. (2022) provide empirical evidence that former entrepreneurs experience persistent earnings penalties upon re-entering wage employment. Their findings highlight two key mechanisms driving this penalty: (1) a supply-side effect, where former entrepreneurs prefer flexible work arrangements (e.g., part-time jobs or hybrid employment), and (2) a demand-side effect, where employers penalize entrepreneurial experience due to mismatched labor market expectations. Notably, they find that the wage penalty is particularly severe for high-wage individuals who fail as entrepreneurs, as the poor outcome

contrasts with prior positive market expectations. In contrast, young individuals experience less severe penalties, as their pre-entrepreneurial labor market expectations are less established. Thus, the policy implications of promoting self-employment among young individuals should consider these differential signaling effects.

6.2. Matching Approach

We assess whether temporary self-employment affects labor market outcomes for individuals who transition back to dependent employment. The treatment group includes individuals who move from full-time employment to self-employment in a given year and subsequently return to wage employment within a defined period. The control group comprises those who remain in full-time employment throughout the same period. We consider three self-employment duration windows—one, three, and five years—to capture different stages of entrepreneurial experience and labor market reintegration. Shorter spells (one year) may primarily reflect unsuccessful entrepreneurial attempts, while longer durations (three to five years) allow for skill accumulation and business stabilization before transitioning back to wage employment.[7]

A major challenge in this analysis is selection bias, as self-employment decisions are influenced by individual characteristics—such as age—that also affect labor market outcomes. To address this, we employ propensity score matching (PSM) to match individuals who transitioned to self-employment with comparable individuals who remained in full-time employment. This matching process helps ensure that any differences in outcomes can be attributed to the self-employment experience rather than pre-existing differences between individuals. Specifically, we match on immigrant status, family responsibilities, level of formal education, industry, skill level, work experience, pre-treatment earnings, municipality size, and local unemployment rate. These variables capture key determinants of labor market outcomes and self-employment selection.

Tables **B1** (female) and **B2** (male) in the Appendix provide some descriptive statistics for a selection of covariates, where treatment is defined as starting self-employment in 2013 for at most one year. Treatment and control groups of female workers are comparable in terms of nationality, family responsibilities or level of formal education. In contrast, women who remained full-time employed were on average 3 years old at the beginning of their employment spell. Treated women were six percentage points more likely to have low-skilled occupations and their pre-entrepreneurial wages were 23% lower. In addition, control units were more likely to live in large cities. Concerning employment spells one year later, treated women were earning on average about EUR 260 less than their control counterparts. We find similar

[7] For example, in the five-year window, we divide the CWLS data into three subsamples (2009, 2011, 2013), with outcomes measured five years later, when all individuals have returned to full-time employment (respectively, in 2014, 2016, and 2018).

patterns in the male subsample, with larger differences regarding nationality or regional unemployment rates. Notably, treated workers face wages EUR 400 lower than those who remain full-time employed.

Our PSM strategy rests on two key assumptions: (1) unconfoundedness, meaning that, conditional on observed characteristics, no systematic differences in outcomes exist between the treated and control groups beyond the treatment; and (2) common support, meaning that individuals with the same observable characteristics must have a positive probability of belonging to either the treatment or control group.

As exact matching on multiple covariates is challenging due to the large number of variables, we use the propensity score $P(X)$–the conditional probability of receiving treatment given pre-treatment characteristics. As true propensity scores are unobservable in observational studies, we estimate them following Abadie & Imbens (2016). By employing this methodology, we aim to isolate the impact of temporary self-employment on long-term labor market outcomes. Specifically, we aim to determine whether these self-employment spells enhance human capital and lead to higher future wages, or whether they serve as a signal of labor market instability, resulting in lower future earnings.

Naturally, unobservable characteristics—such as risk tolerance, financial constraints or psychological traits—can influence entrepreneurial decisions (Bermejo et al., 2024; Blanchflower & Oswald, 1998; Bönte et al., 2012) and pose a challenge to causal interpretation. Nonetheless, we rely on a rich set of sociodemographic and labor market variables available in the CWLS data, which allows us to approximate a valid comparison.

6.3. Results

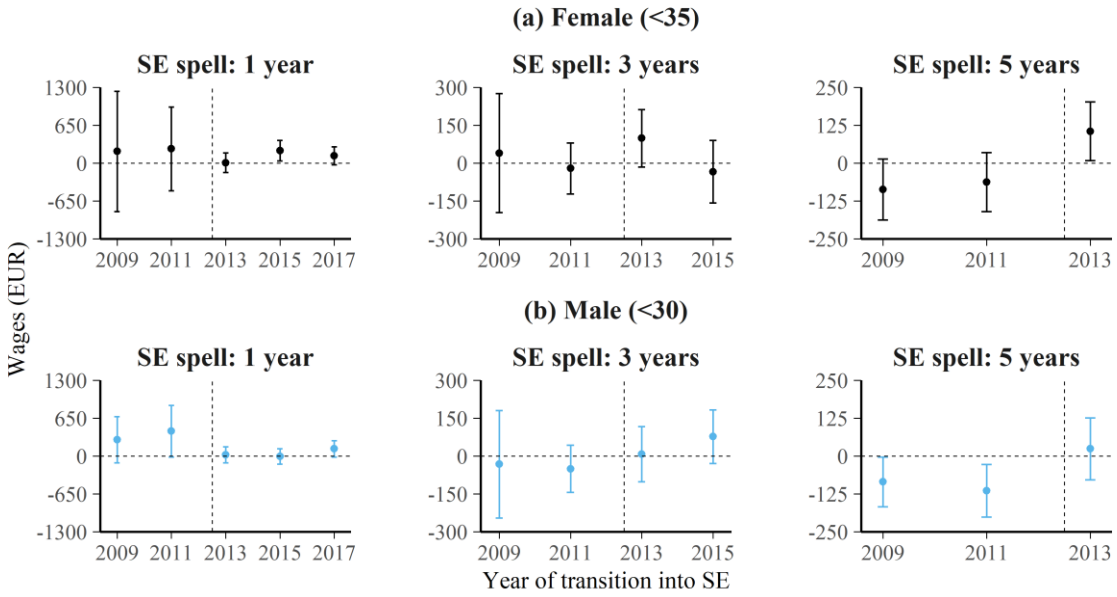
Our baseline results are obtained using k-nearest neighbors matching. Based on the standardized mean bias criterion, these algorithms achieve the best covariate balance between the treatment and control groups across all subsamples. These methods are also more computationally efficient than kernel or radius matching, which offer similar quality. We tested different values of k, the number of control group neighbors used to estimate the average treatment effect on the treated (ATT). As the choice of k had minimal impact on quality criteria, we opted for simplicity and set $k = 1$.

We present the distribution of estimated propensity scores for the 2013 subsample to assess the balance between treated and control groups (Figures **A10** and **A11** in the Appendix). The results confirm that nearest-neighbor matching successfully balances the distributions, enhancing comparability between groups. To rule out systematic differences in observables, we also examine covariate balance. Figures **A12** and **A13** in the Appendix provide a graphical overview of standardized mean differences across covariates before and after matching. Before matching, the treatment and control groups exhibited some differences, with standardized mean differences exceeding the 0.1 threshold. However, after applying nearest-neighbor matching, these biases were significantly reduced, improving balance across all covariates.

We proceed to examine the impact of short self-employment spells on wage rates (Figures **6** and **7**), which highlight heterogeneity depending on the spell duration and the business cycle. Entrepreneurial experience lasting up to one year produces null to positive effects of up to EUR 270 (15% with respect to the control group) on post-entrepreneurial wages of female workers who became self-employed in 2013 onward compared to women who remained full-time employed. Longer self-employment spells led to insignificant effects among women over 35, whereas younger women who entered entrepreneurship in 2013 show small but positive wage effects after five-year spells—about EUR 100.

In contrast, former male entrepreneurs show mixed results: negative effects up to EUR 160 (8% with respect to the control group) if they became self-employed during the recession, and positive effects of about EUR 110 (5% with respect to the control group) during the recovery period. Entrepreneurial periods lasting up to five years led to significant negative wage effects of up to EUR 265 (13% with respect to the control group) when compared to workers who remained in full-time employment.

Figure 6: *Effects on Wages in Dependent Employment (Below Age Cutoff)*

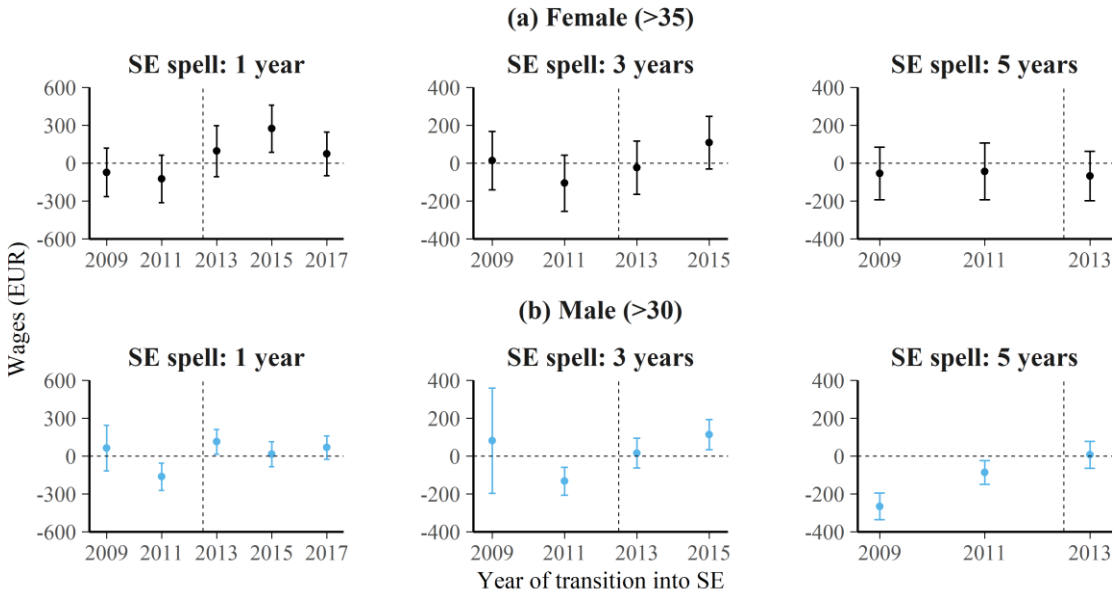


Notes: This figure illustrates the propensity score matching estimates for wages in full-time employment after one, three, and five years. The estimates were obtained using *k*-nearest neighbor matching with *k = 1*. Point estimates are shown with 90% confidence intervals. The treated group consists of individuals who switched from full-time to self-employment (SE) in the year indicated on the horizontal axis and were full-time employed after the given SE spell duration. The control group includes individuals who remained full-time employed during the same period.

Samples are restricted to females aged 34 or younger and males aged 29 or younger.

Source: Authors' calculations based on CWLS data.

Figure 7: *Effects on Wages in Dependent Employment (Above Age Cutoff)*



Notes: See Figure **6**.
Samples are restricted to females aged 35 or older and males aged 30 or older.
Source: Authors' calculations based on CWLS data.

7. Conclusion

This paper evaluates the effects of Spain's 2013 Entrepreneurship Act subsidy on young entrepreneurs, focusing on gender- and age-specific labor market outcomes. Using a regression discontinuity design, we examine the policy's impact on wages, employment probability, and industry when switching to dependent employment. We also assess whether the subsidy acted as an active labor market policy rather than a mere transfer, identifying the groups most affected. Additionally, employing propensity score matching, we explore how temporary self-employment influences labor market outcomes for individuals transitioning back to dependent employment.

The subsidy did not significantly extend self-employment spells or improve employment probabilities for young entrepreneurs. While men showed no significant post-subsidy wage improvements, women who transitioned to dependent employment within two years experienced an 18% increase in wages, with a notable shift into the services sector. The positive wage effects are concentrated among women at higher risk of unemployment, particularly those with family responsibilities. While the policy did not increase overall business survival rates, it delayed self-employment exit by up to 90 days for groups facing high labor market instability. Thereby, the subsidy provided them with time to ensure a successful transition to dependent employment, compared to unsubsidized workers.

We also evaluate the decision to start a business in the context of the support measures promoting the transition to self-employment in 2013. Post-entrepreneurial wages for men are highly dependent on the

business cycle. Men who became self-employed during the financial crisis faced earnings losses of up to 13% upon returning to dependent employment, while women who self-employed after 2013 and re-entered wage employment three years later saw a 15% increase in monthly earnings. For younger workers, self-employment spells generally have minimal or no significant impact on earnings, regardless of gender, duration, or business cycle. However, young women tend to experience positive earnings effects upon returning to wage employment, especially after short self-employment spells.

For men over 30, results are generally negative, particularly for those who became self-employed during the crisis and returned to wage employment after three to five years. These men face significant earnings penalties compared to their continuously employed peers. In contrast, older women either remain unaffected or benefit from positive outcomes if they return to full-time employment within a year.

Our findings highlight that self-employment signals different attributes depending on prior labor market positioning. Men, typically expected to follow stable career paths, face a stronger penalty for instability if self-employment is seen as a deviation from an expected progression. Employers view their entrepreneurial experience as unstable, especially if they return to wage work after a short or unsuccessful self-employment spell. Conversely, women, often expected to prioritize flexibility, may have their self-employment experience viewed more favorably, aligning with perceptions of proactivity and adaptability.

Policymakers should complement broad subsidy programs with business development support, mentoring, and networking opportunities to improve entrepreneurial sustainability. These subsidies should target women and other groups at higher risk of unemployment, with a focus on supporting their transition to higher-paying sectors in dependent employment. Policies should also encourage early re-employment opportunities for male entrepreneurs and provide re-entry incentives after short self-employment spells. New regulations should provide stronger re-employment incentives and skill development programs for entrepreneurs over these age thresholds. Targeted retraining and upskilling initiatives could help older cohorts mitigate skill atrophy or sectoral mismatches.

To prevent negative signaling effects, policymakers should design mechanisms to assess entrepreneurial talents early, enabling timely interventions that encourage business growth or support re-entry into wage employment before long-term negative effects arise. Additionally, well-designed entrepreneurship incentives could have positive long-term economic impacts, but policymakers must focus on the consequences of failed self-employment and prepare individuals for sustainable employment or alternative labor market paths.

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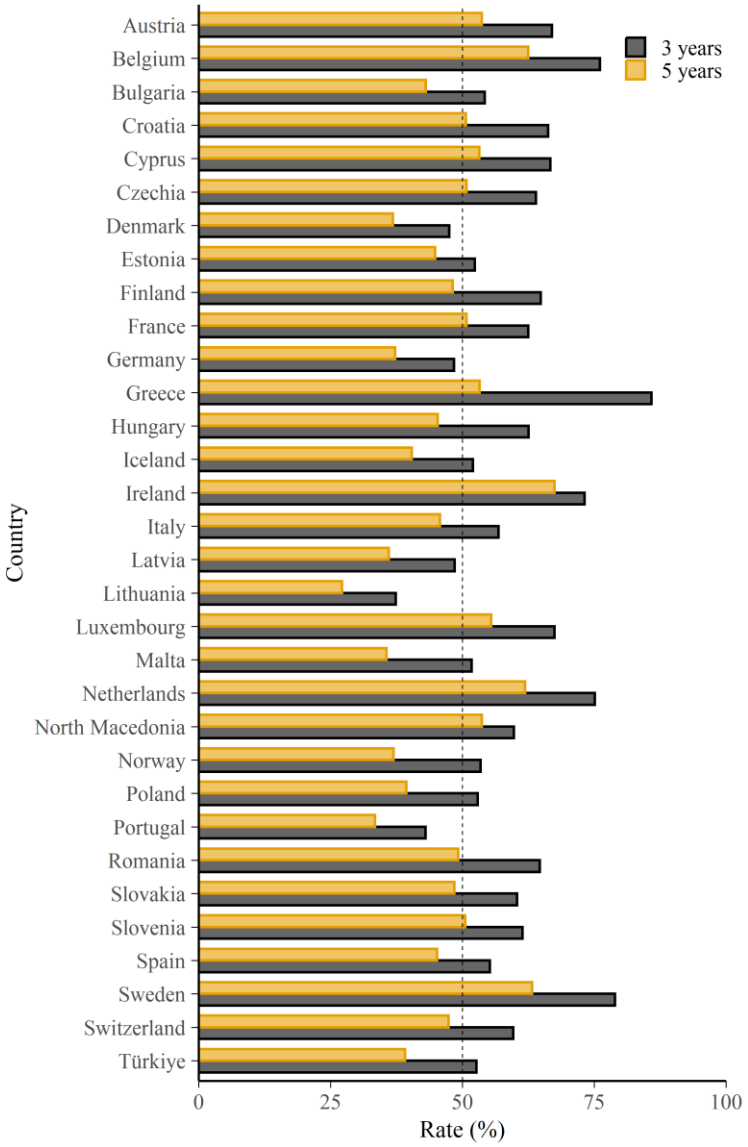
Appendix A. Figures

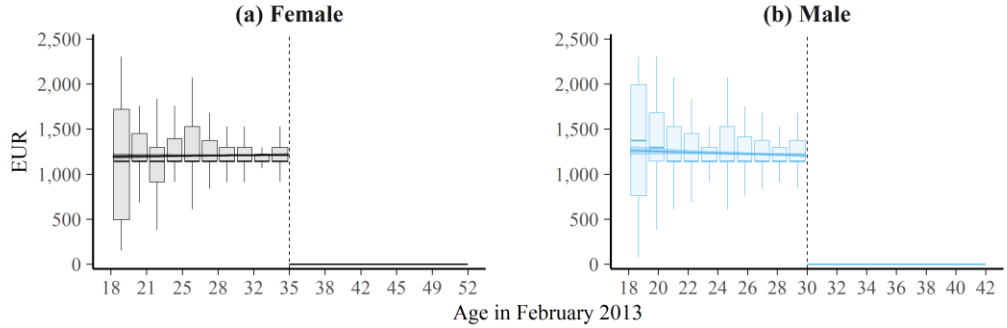
Figure A1: *Business Survival Rates (2020)*



Notes: This figure illustrates the business survival rates in 2020 for a selection of countries. Survival is defined as being active and identifiable, regardless of changes in economic activity, size, ownership, or location. For example, the three-year enterprise survival rate for newly founded enterprises is the percentage of businesses that were established in 2017 and remained active in 2020, relative to all enterprises founded in 2017.

Source: Eurostat (2024)

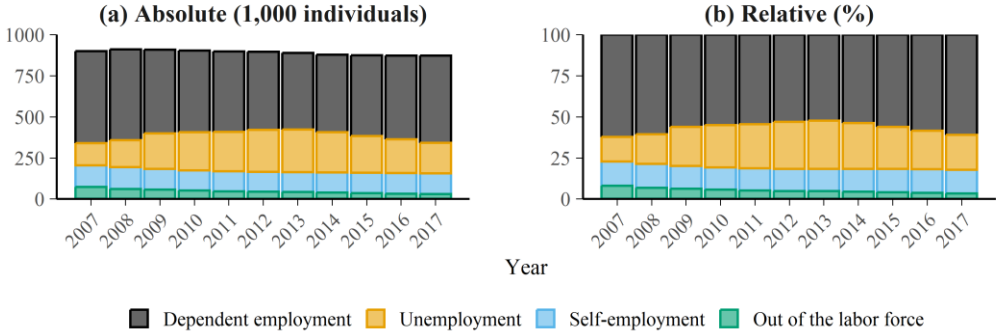
Figure A2: *Accumulated Subsidy*



Notes: This figure illustrates the distribution of the estimated total subsidy amount perceived by beneficiaries, categorized by gender and age group. The total subsidy amount is based on the eligibility period and the monthly subsidy, calculated as 30% of the contribution, which in turn was determined by applying the minimum rate to the minimum contribution base effective during each month. We show a trend line representing the subsidy as a function of age in February 2013.

Source: Authors' calculations based on CWLS data.

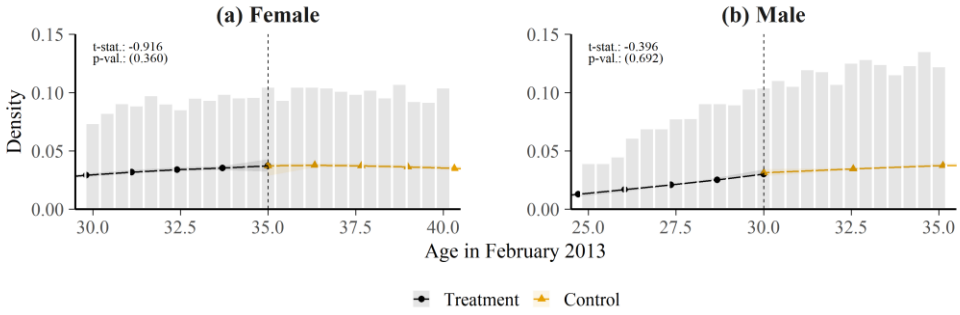
Figure A3: *Labor Market Status Distribution*



Notes: This figure illustrates the yearly occupational distribution in Spain from 2007 to 2017, presented in both (a) absolute and (b) relative terms. The sample is limited to individuals aged 18 and older. Workers are categorized according to their relevant statuses: Dependent Employment, Out of the Labor Force, Self-Employment, and Unemployment.

Source: Authors' calculations based on CWLS data.

Figure A4: *Density Plot for Manipulation Testing*

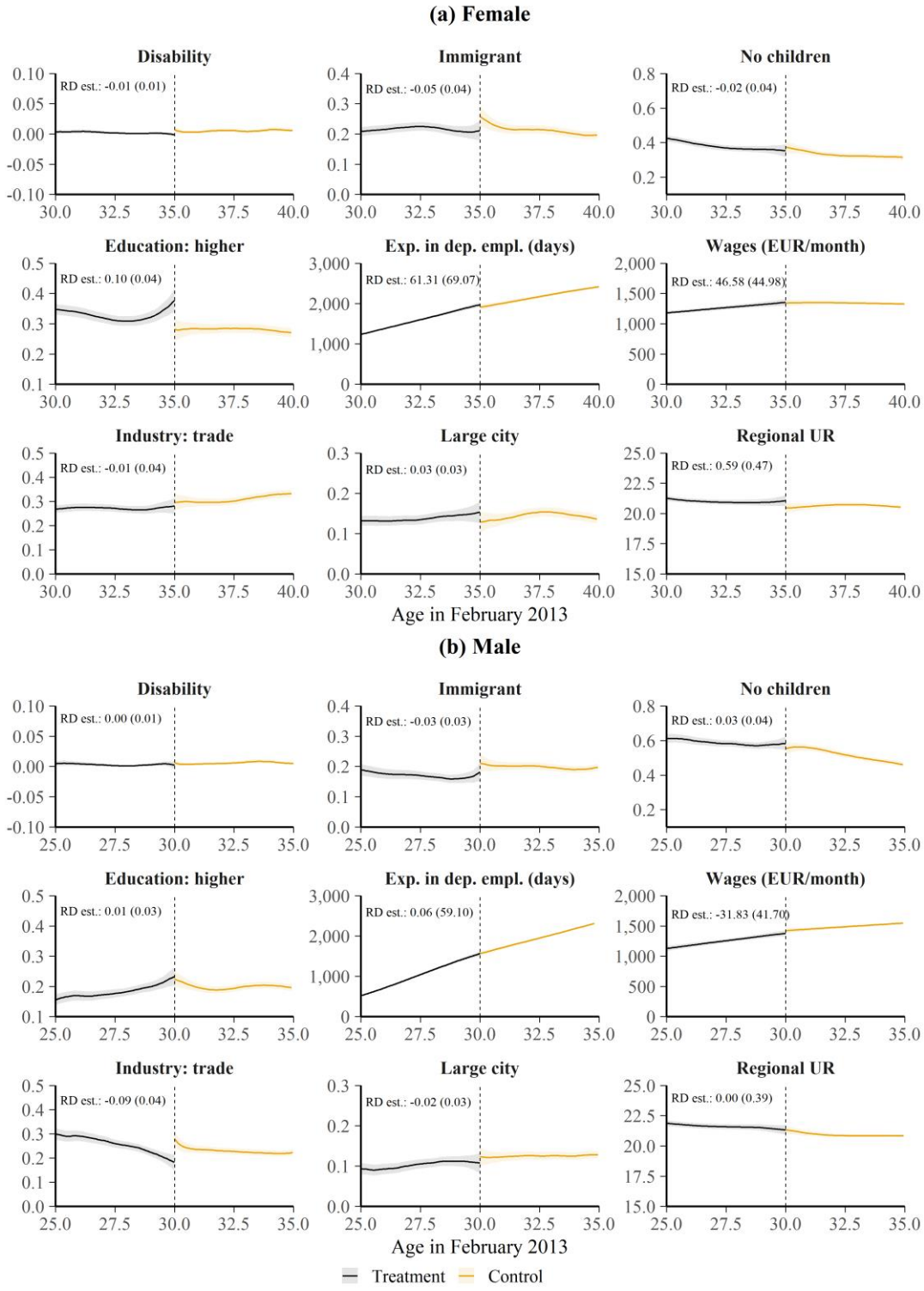


Notes: This figure presents the results of the manipulation tests of the running variable (age in February 2013) for each subsample. First, a histogram of the running variable is shown. Second, local second-order polynomial density estimates, as suggested by Cattaneo et al. (2020), are displayed. The t-statistic and p-value for each density discontinuity test are reported. The subsamples are restricted to individuals who were engaged in active self-employment in February 2013.

Robust bias-corrected confidence intervals are indicated as shaded areas.

Source: Authors' calculations based on CWLS data.

Figure A5: *Covariate Balance*

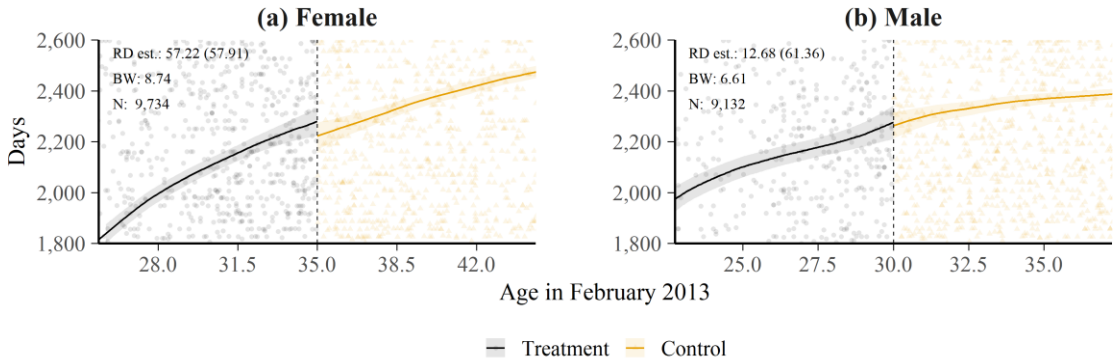


Notes: This figure illustrates the relationship between the running variable (age in February 2013) and a selection of covariates. It is based on local linear regressions conducted on either side of the cut-off, using the Imbens-Kalyanaraman optimal bandwidth calculation and a triangular kernel.

Subsample is restricted to individuals who had an active self-employment spell in February 2013 and an active dependent employment spell one year later.

Source: Authors' calculations based on CWLS data.

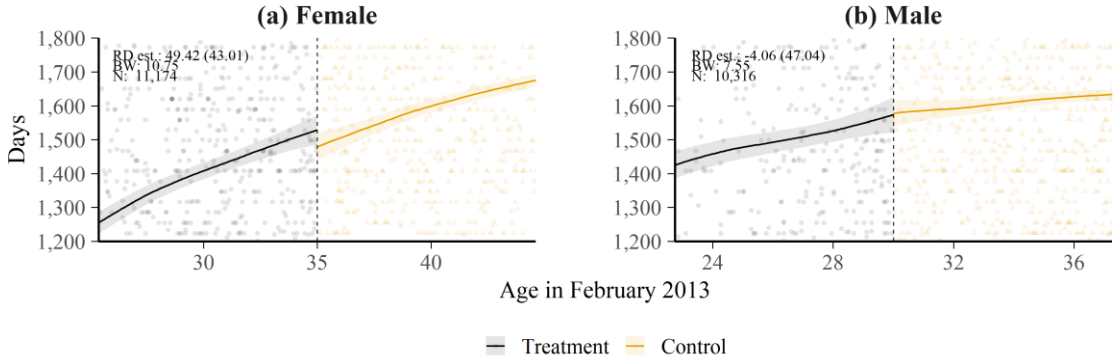
Figure A6: *Effects on Self-Employment Spell Duration*



Notes: See Figure **2**. This figure illustrates the relationship between the running variable (age in February 2013) and the total duration of self-employment spells.

Source: Authors' calculations based on CWLS data.

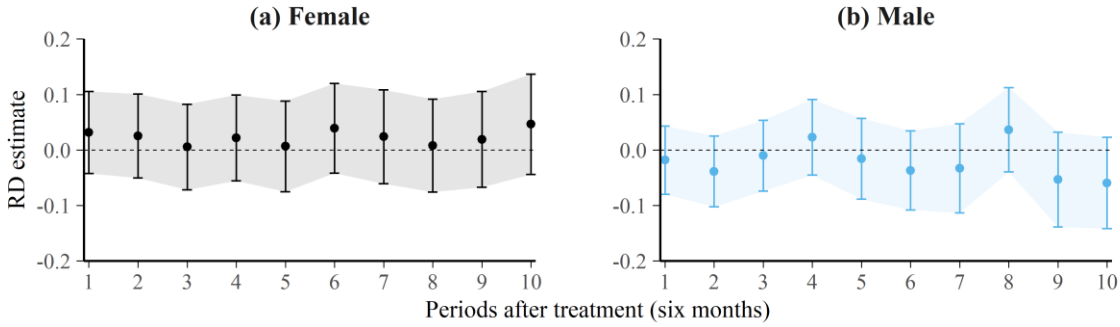
Figure A7: *Effects on Time to Exit from Self-Employment*



Notes: See Figure **2**. This figure illustrates the relationship between the running variable (age in February 2013) and the time difference between treatment (February 2013) and exit from self-employment.

Source: Authors' calculations based on CWLS data.

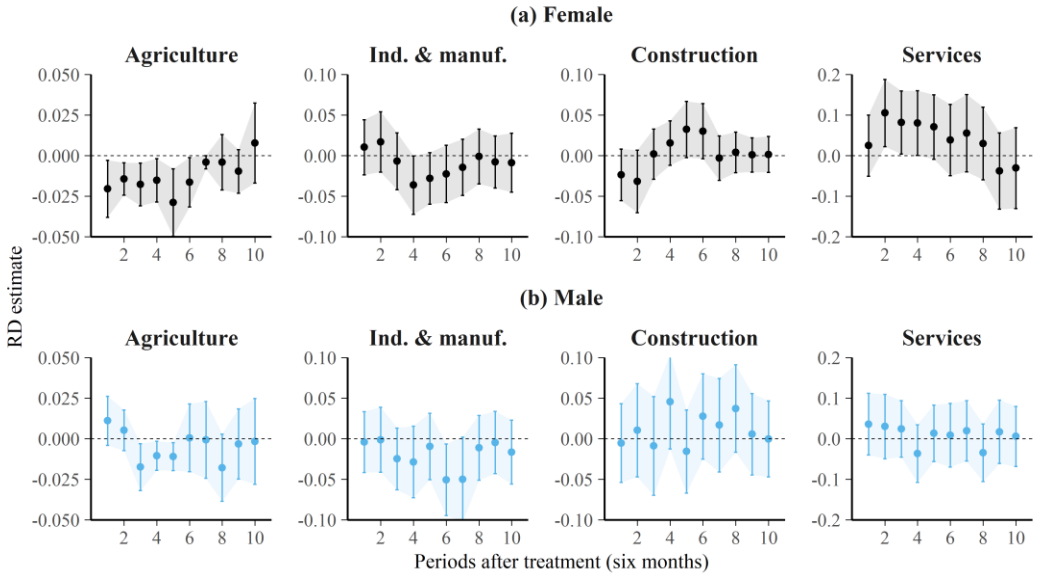
Figure A8: *Effects on Employment Probabilities*



Notes: See Figure **4**. This figure presents the regression discontinuity estimates for the probability of having a dependent employment spell after a given number of six-month periods following the treatment in February 2013.

Source: Authors' elaboration based on CWLS data.

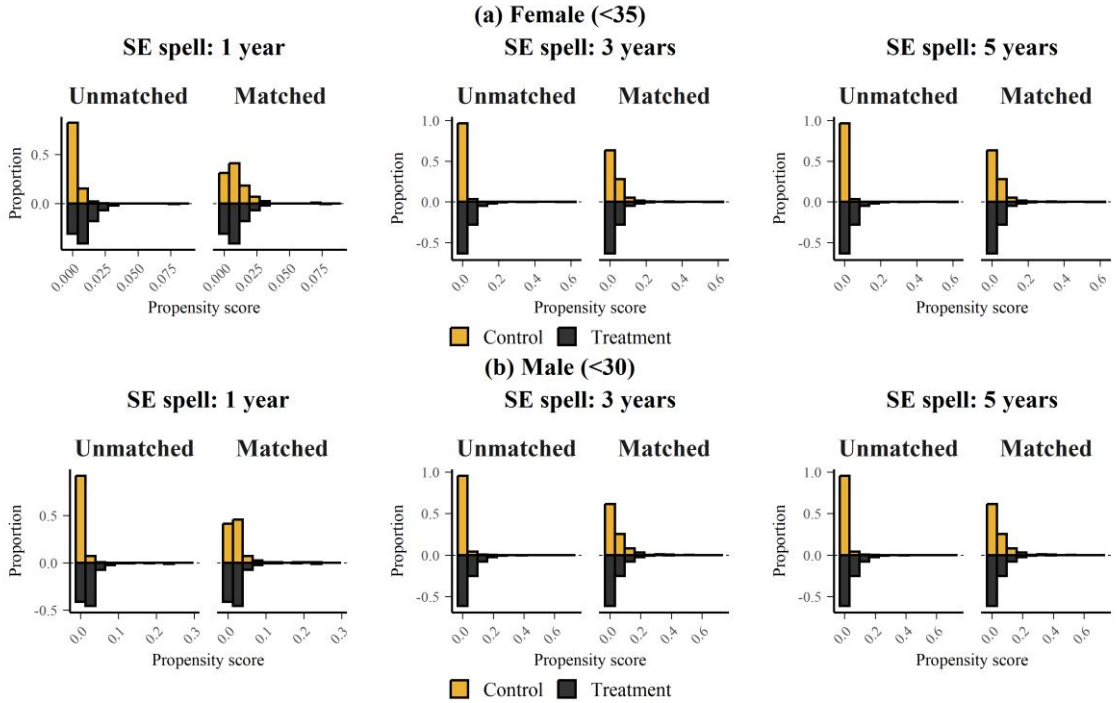
Figure A9: *Effects on Dependent Employment Sector Probabilities*



Notes: See Figure **4**. This figure presents the regression discontinuity estimates for the probability of having a dependent employment spell in a specific industry, conditional on being in dependent employment after a given number of six-month periods following the treatment in February 2013.

Source: Authors' elaboration based on CWLS data.

Figure A10: *Propensity Score Balance (Below Cutoff)*

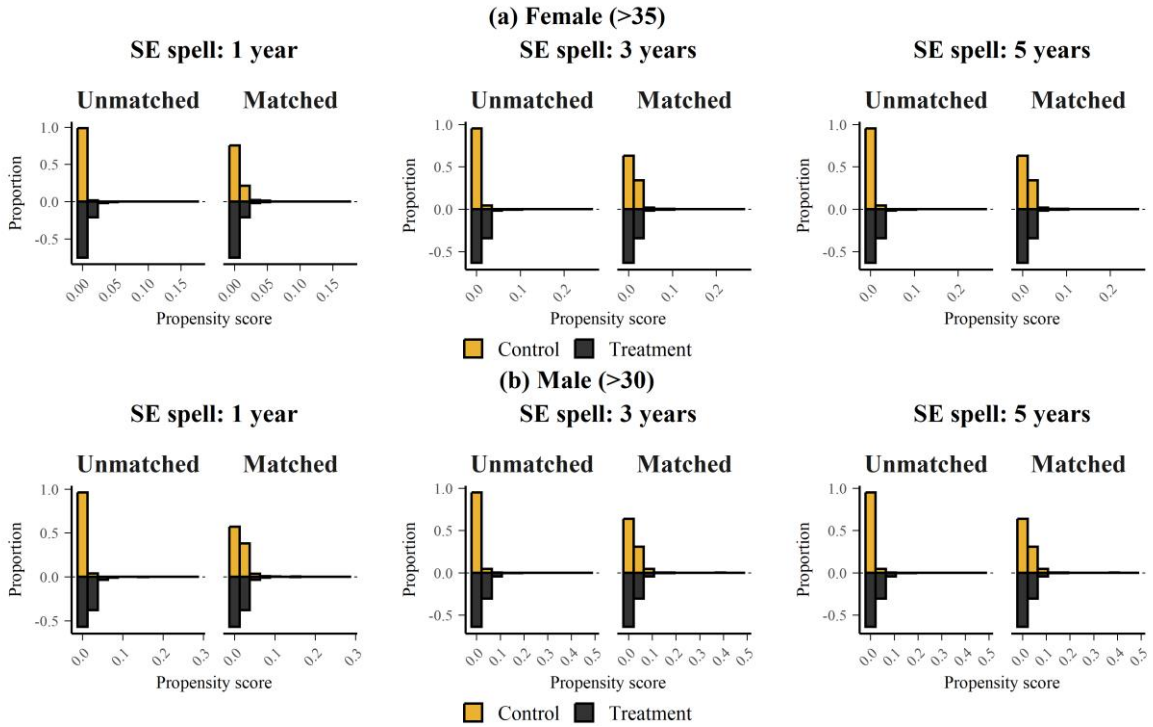


Notes: This figure illustrates the distribution of the estimated propensity scores in the 2013 subsample, both before and after implementing *k*-nearest neighbor matching with *k=1*. The treated group consists of individuals who switched from full-time to self-employment in 2013 and were full-time employed after one, three, or five years. The control group consists of individuals who remained full-time employed during the same time window.

The subsamples are restricted to females aged 34 or younger and males aged 29 or younger.

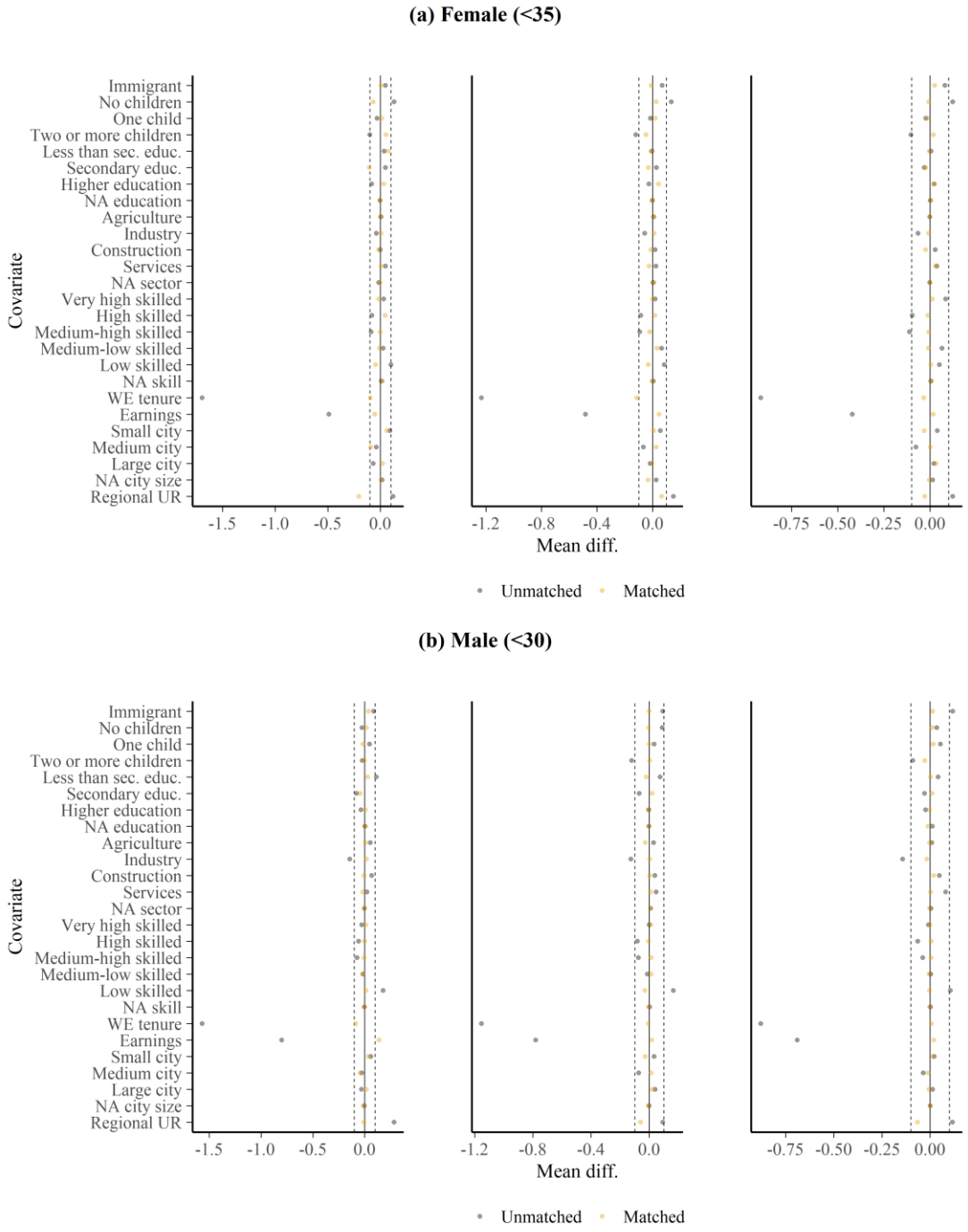
Source: Authors' elaboration based on CWLS data.

Figure A11: *Propensity Score Balance (Above Cutoff)*



Notes: See Figure **A10**.
The subsamples are restricted to females aged 35 or older and males aged 30 or older.
Source: Authors' elaboration based on CWLS data.

Figure A12: *Covariate Balance (Below Cutoff)*

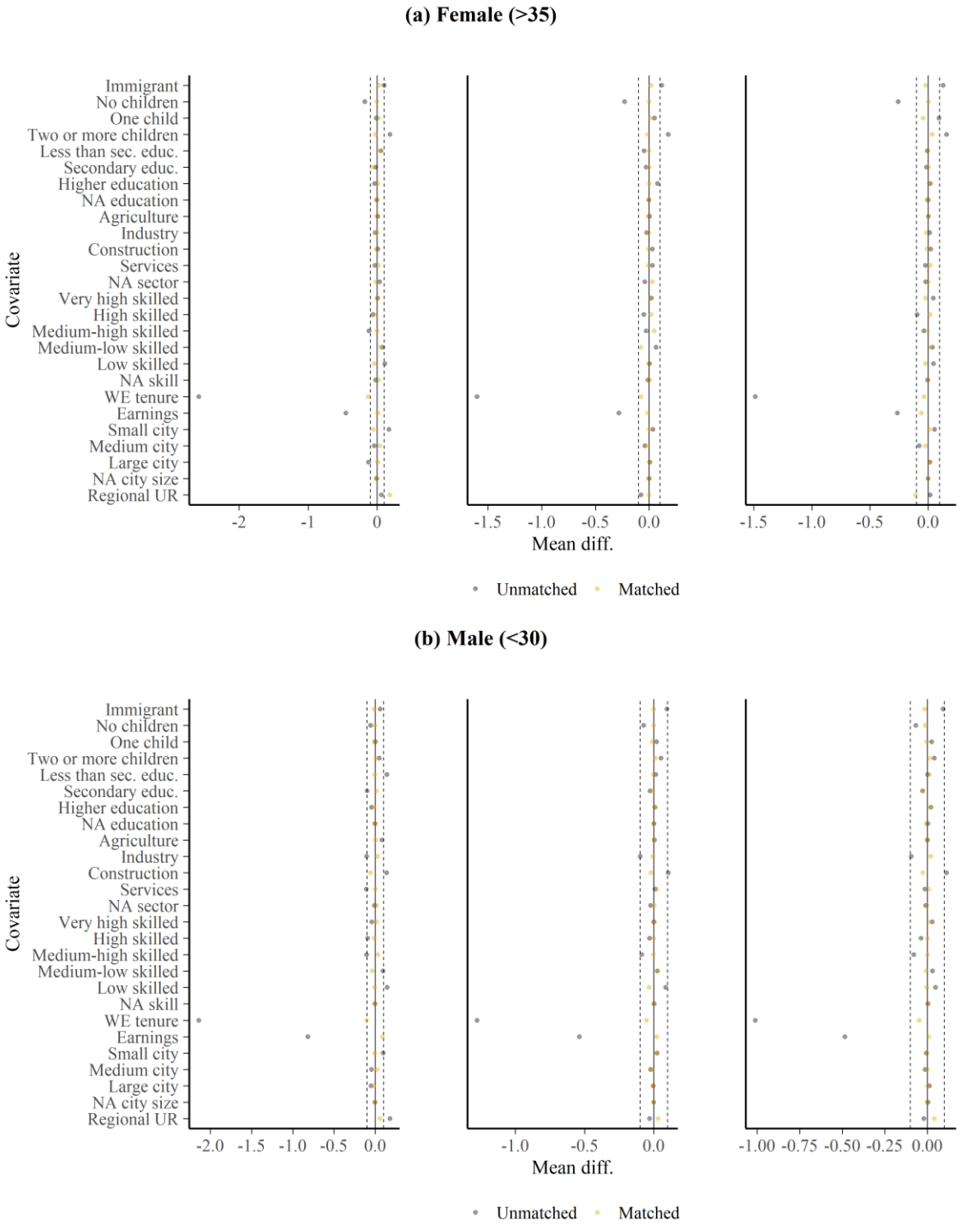


Notes: This figure illustrates the standardized mean differences across covariates in the 2013 subsample, both before and after implementing *k*-nearest neighbor matching with *k=1*. The vertical dashed lines represent a 0.1 threshold.

The subsamples are restricted to females aged 34 or younger and males aged 29 or younger.

Source: Authors' calculations based on CWLS data.

Figure A13: *Covariate Balance (Above Cutoff)*



Notes: See Figure **A12**.
The subsamples are restricted to females aged 35 or older and males aged 30 or older.
Source: Authors' calculations based on CWLS data.

Appendix B. Tables

Table B1: *Descriptive Statistics (Female)*

	Treatment	**Control**	**Mean test**	
	Mean	Mean	Diff.	Std. err.
Variable	(1)	(2)	(3)	(4)
(a) Pre-determined characteristics				
Age at entry	34.69	37.88	-3.19***	(0.66)
Immigrant	0.10	0.09	0.01	(0.02)
Family responsibilities				
No children	0.49	0.51	-0.02	(0.03)
Formal education				
Higher education	0.27	0.28	-0.01	(0.03)
(b) Pre-treatment job				
Sector				
Services	0.86	0.83	0.03	(0.02)
Occupation type				
Low-skilled	0.17	0.11	0.06**	(0.03)
Tenure	547.04	2,018.73	-1471.70***	(58.58)
Wage (log)	7.10	7.33	-0.23***	(0.04)
(c) Macroeconomic context				
Municipality size				
Large city	0.10	0.18	-0.08***	(0.02)
Regional UR	25.89	25.55	0.34	(0.45)
(d) Post-treatment job				
Wage	1,536.07	1,798.55	-262.48***	(56.65)
Spell duration	1.61	3.37	-1.76***	(0.14)
Firm size	826.14	884.75	-58.62	(200.27)
Services	0.61	0.63	-0.02	(0.03)

Notes: This table presents the mean for a selection of variables for the sample of female workers in the 2013 subsample, for treatment and control units separately. It also includes a mean test for differences between groups.

* p < 0.1, ** p < 0.05, *** p < 0.01

Source: Authors' calculations based on CWLS data.

Table B2: *Descriptive Statistics (Male)*

	Treatment	**Control**	**Mean test**	
	Mean	Mean	Diff.	Std. err.
Variable	(1)	(2)	(3)	(4)
(a) Pre-determined characteristics				
Age at entry	36.36	37.96	-1.61***	(0.42)
Immigrant	0.13	0.09	0.03**	(0.01)
Family responsibilities				
No children	0.48	0.52	-0.04*	(0.02)
Formal education				
Higher education	0.18	0.18	0.00	(0.02)
(b) Pre-treatment job				
Sector				
Services	0.58	0.61	-0.03	(0.02)
Occupation type				
Low-skilled	0.24	0.15	0.09***	(0.02)
Tenure	501.88	2,210.66	-1708.78***	(50.12)
Wage (log)	7.21	7.49	-0.28***	(0.02)
(c) Macroeconomic context				
Municipality size				
Large city	0.12	0.15	-0.03**	(0.01)
Regional UR	26.78	25.89	0.90***	(0.29)
(d) Post-treatment job				
Wage	1,649.51	2,058.87	-409.36***	(35.99)
Spell duration	1.29	3.30	-2.01***	(0.08)
Firm size	209.72	580.65	-370.93***	(54.19)
Services	0.33	0.36	-0.04*	(0.02)

Notes: See Table **B1**.
Source: Authors' calculations based on CWLS data.