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Economic Cycle and Wage Premium of Master's Degrees in the Early Years of Transition to the Labour Market

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Research Article

Keywords: Wage returns, master's degree, economic cycle, differences-in-differences, higher education

Posted Date: November 16th, 2026

DOI: https://doi.org/10.21203/rs.3.rs-11282706/v1

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Additional Declarations: No competing interests reported.

Abstract

This study examines wage differentials associated with master's degree completion during the early transition to the labour market in Spain. Using nationally representative microdata from the 2014 and 2019 waves of the Survey on the Labour Insertion of University Graduates (EILU), it compares cohorts entering employment during the Great Recession and its aftermath with those entering during the recovery. The analysis first examines selection into master's completion and employment and then estimates Mincer wage equations and difference-in-differences (DiD) models with degree-program and individual fixed effects and controls. Individual fixed effects account for time-invariant heterogeneity but cannot eliminate selection arising from unobserved factors that change over time. Master's graduates initially display lower employment probabilities in both cohorts. By the final observation, the adjusted employment differential remains negative in 2014 (− 5.4 percentage points) but becomes positive in 2019 (+ 2.3 points). The individual-fixed-effects estimates indicate positive differential wage growth associated with master's completion: 8.7% for the 2014 cohort and 25.6% for the 2019 cohort. Estimates are positive for both sexes, graduates under 35, and most fields. An interval regression using the original bounds of the administrative contribution-base quintiles yields a positive 2019 differential of 4.2%, showing that the result does not depend exclusively on midpoint imputation. The larger differential in the later cohort is consistent with stronger wage progression during the recovery, but it also coincides with the Bologna transition and expansion of the Spanish master's sector. The study therefore provides nationwide evidence on how master's selection, employment access, and early-career wage trajectories differ across cohorts exposed to contrasting conditions, without interpreting these differences as definitive causal effects of postgraduate education or the economic cycle.

Introduction

Tertiary education enrolment has surged over the past decade, especially in advanced economies. According to data from Vieira do Nascimento et al. (2020) the gross enrolment rate in higher education increased globally from 19% to 38%, in the period between 2000 and 2018. Within the European continent, the average enrolment rate rose to 46% in 2018 for the 19–23 age group.

The objective of this paper is to analyse the relationship between the economic cycle and the wage returns of master's degrees in Spain. Postgraduate education is a fundamental stage in the development of human capital in modern economies. The relevance of Master's studies lies in the enhanced employability and tangible economic benefits they provide to higher education graduates[1]. Over the past years Spain has witnessed a dramatic increase in Master's enrolment, rising from 16,609 students in the 2006–2007 academic year to 288,955 in 2023–2024 (SIIU, 2024). This growth is explained, inter alia, by the reform of the curriculum structure in the framework of the Bologna Process (see Fig. 1).

Source: SIIU (Ministry of Universities), 2024.

Despite the expansion in the supply of university graduates, demand for skilled labour has also increased, sustaining positive returns to higher education. As Psacharopoulos and Patrinos (2018) document, the "race between education and technology" has shifted demand outward, with average global private returns of around 9% per year of schooling, although these vary by level, gender, sector, and region. However, beyond this long-term trend, the business cycle plays a critical role in shaping early career outcomes. Evidence shows that entering the labour market during downturns leads to persistent wage and employment penalties. Schwandt and Wachter (2019) find that such effects can last up to a decade, while Minaya et al. (2024) show that economic crises reduce, but do not eliminate, the returns to postgraduate education. This pattern is also observed in Spain, where the Great Recession significantly deteriorated employability and intensified issues such as over-education and underemployment (Fernández et al., 2017; Iriondo, 2022a).

Based on the evidence presented, it is essential to focus the analysis of human capital returns on the financial benefits of postgraduate studies in Spain, particularly during the early years following entry into the labour market. Such an investigation could help mitigate the uncertainty faced by students whose expectations may be inflated by rising enrolment in graduate programs and the substantial costs of exclusive programs—often exceeding 100,000€ at renowned business schools.

This paper examines wage differentials associated with master's degree completion during the early transition to the labour market in Spain. It uses the nationally representative Survey on the Labour Insertion of University Graduates (EILU), conducted by the Spanish National Statistics Institute (INE), to compare the 2014 wave, covering 2009–2010 graduates, with the 2019 wave, covering 2013–2014 graduates. The first cohort entered employment during the Great Recession and its aftermath, whereas the second did so during economic recovery. The comparison also spans the implementation of the Bologna reform and expansion of the *grado–máster* structure. The study therefore examines cohorts exposed to contrasting economic and institutional conditions without seeking to identify the causal effect of the economic cycle itself.

Research specifically examining postgraduate wage returns remains limited. Undergraduate and postgraduate qualifications are often combined into returns per additional year of schooling, while suitable microdata on educational trajectories, employment outcomes, and individual characteristics are scarce (Gándara and Toutkoushian, 2017). These limitations are particularly relevant during early labour market integration, when graduates have little professional experience and postgraduate education may delay employment entry. Observed wage differences may consequently reflect specialization as well as selection into master's education, employment, occupations, and fields.

Evidence for Spain indicates positive but heterogeneous differentials. Lassibille (1993), using the 1980 INE Household Budget Survey, found earnings for long-cycle education to be 29% higher than for short-cycle programmes. Arrazola et al. (2003), analysing men in the 1993–1994 European Community Household Panel, estimated returns between 6% and 9%. Caparrós (2023), using the 2019 EILU and modelling educational selection, reported approximately 15%. Internationally, Minaya et al. (2024) use administrative data and individual fixed effects for graduates in Ohio and estimate an average return of about 14%. They also find substantial heterogeneity and that downturns reduce, but do not eliminate, positive returns. Their study provides a useful external benchmark, although institutional, measurement, and sample differences limit direct comparison with Spain.

This study contributes in four respects. First, it provides nationwide evidence from a European higher-education system, complementing US findings based on single-state records. Second, it distinguishes graduates who completed a master's degree from those who did not, avoiding the conflation of enrolment with qualification attainment. Third, it examines selection into master's completion and employment before analysing wages, which are observed only for employed graduates. Fourth, it compares early-career wage trajectories across crisis and recovery cohorts while considering the simultaneous transformation of Spanish degree structures under Bologna. This combination documents how postgraduate wage differentials vary across cohorts without attributing that variation exclusively to macroeconomic conditions.

The analysis proceeds in three stages. First, descriptive comparisons and binary-response models identify the characteristics associated with master's completion. Second, employment probit models examine whether completion is associated with employment at the beginning and end of each period. Third, wage equations are estimated using Mincer OLS models and difference-in-differences (DiD) specifications with degree-program and individual fixed effects. The models control personal and sociodemographic characteristics, educational background, employment conditions, field, and autonomous community. Individual fixed effects account for time-invariant characteristics that may jointly influence completion and earnings, mitigating bias from stable unobserved heterogeneity but not from unobserved factors that change over time.

The wage measure is derived from Social Security contribution-base quintiles and is administrative but grouped rather than continuous. The main semilogarithmic models assign a representative midpoint to each interval. As a robustness check, an interval regression uses the original bounds and accommodates left-, interval-, and right-censored observations. Additional models examine heterogeneity by gender, age, and field. These exercises assess whether the central pattern depends on midpoint imputation or a particular subgroup.

The results show non-random selection into master's completion and lower initial employment probabilities among master's graduates in both cohorts. By the final observation, the adjusted employment differential remains negative in 2014 (− 5.4 percentage points) but becomes positive in 2019 (+ 2.3 points). Cross-sectional OLS wage coefficients are initially negative, whereas the longitudinal estimates indicate positive differential wage growth associated with completion. With individual fixed effects, this differential is 8.7% for the 2014 cohort and 25.6% for the 2019 cohort. Positive estimates are observed for both sexes, graduates under 35, and, particularly in 2019, across fields. The interval regression also produces a positive 2019 association of approximately 4.2%, supporting the direction of the findings without relying on midpoint imputation.

The larger differential in the later cohort is consistent with stronger wage progression during the recovery and the changing role of master's qualifications following Bologna. However, the data cannot disentangle these mechanisms or remove all time-varying selection. The estimates are therefore interpreted as differential wage changes associated with master's completion, not definitive causal effects of postgraduate education or macroeconomic recovery.

The remainder of the article is organized as follows. Section 2 reviews the literature. Section 3 describes the data, variable construction, and empirical strategy. Section 4 presents the selection, wage, heterogeneity, and robustness results. Section 5 concludes and discusses policy implications, limitations, and avenues for future research.

Materials and methods

This study uses the Survey on the Labour Insertion of University Graduates (EILU), conducted nationwide by the Spanish National Statistics Institute (INE), to analyse the 2014 and 2019 waves, comprising 30,379 and 31,651 graduates, respectively. Social Security records allow employment status to be observed in 2011 and 2014 for the first wave and in 2015 and 2019 for the second. The surveys share a consistent structure covering employment conditions, education–job match, contract type and mobility. Spanish graduates represent over 97% of respondents, women constitute the majority in both waves, and the distribution across broad fields of study remains broadly stable.

The 2019 wave nevertheless provides more detailed information on family background, educational trajectories and labour-market outcomes. Its respondents are also slightly older, with fewer graduates under 30 and more aged 35 or above. General scholarship coverage increases moderately, whereas excellence scholarships decline. Raw survey responses indicate that master's participation rises from 34.56% in 2014 to approximately 47% in 2019.

For empirical analysis, master's status was reconstructed to preserve the temporal ordering of treatment. It equals one only when a respondent completed a master's degree after the reference undergraduate qualification and zero when no master's degree was completed. Respondents who completed a master's before or concurrently with the reference degree, or whose timing could not be established, were excluded. Thus, the raw participation rates above differ from the treatment shares in the final estimation samples. The resulting indicator identifies postgraduate completion during the post-graduation window and is not a time-varying measure of enrolment. Estimation samples are further restricted by the availability of the covariates required by each specification.

Labour-market outcomes improve across waves, the endpoint employment rate rises from 74.65% to 85.70%, the average monthly Social Security contribution base used as the wage proxy increases from €1,606 to €2,215, and average work experience expands from 2.45 to 3.87 years. Employment is defined from recorded Social Security affiliation and therefore does not capture employment without an observable affiliation record. Likewise, the wage measure is available only for affiliated graduates with a valid contribution base. The wage difference-in-differences analysis consequently uses a balanced sample with valid contribution-base information at both observation points. Its estimates should therefore be interpreted as conditional on Social Security-covered employment rather than as effects for all graduates. Differences across waves may reflect both macroeconomic conditions and cohort-specific labour-market dynamics.

In order to explore the returns associated with postgraduate higher education during the initial years of labour market engagement and drawing on the core principles of human capital theory, the methodological strategy employed two complementary estimation tools. Firstly, return rates were assessed through a semilogarithmic wage function, which is grounded in the traditional Mincer (1974) model. This estimation, carried out using ordinary least squares, enables

an examination of how the labour market values the postgraduate education of graduates. The starting point for these specifications is encapsulated in the following income equation:

$lnyi = \beta 0 + \beta 1 HCi + \beta 2 EXPi + Xi\gamma + \varepsilon i$ (i)

where $lnyi$ represents the natural logarithm of the labour income of individual i, HC_i denotes the stock of human capital (which, in this study, corresponds to holding a university master's degree), $EXPi$ indicates labour market experience, Xi is a vector representing other observable characteristics of the individual, and εi is the random error term, which is assumed to be distributed as N(0,1).

As mentioned above, master's status was defined as one only when the respondent completed the degree after the reference undergraduate qualification and as zero when no master's degree was completed. Degrees obtained before or concurrently with the reference qualification, as well as observations with indeterminate timing, were excluded. Observable differences by master's status were first assessed using unequal-variance t-tests and Pearson chi-squared tests. We then estimated cohort-specific logit models of master's completion and year-specific probit models of Social Security-covered employment, with robust standard errors. The specifications include demographic and educational characteristics, scholarships, institution type, mobility, field-of-study effects, regional effects and the additional background variables available in each cohort. We report average marginal effects; these models document observable selection and should not be interpreted as a correction for unobserved selection or as causal estimates.

The control variables are organized into three distinct categories, personal and sociodemographic data, graduate education and learning data, and variables pertaining to the graduate's employment situation. In this framework, dichotomous variables are defined—such as the graduate's gender (coded as 1 for female and 0 otherwise), the presence of any disability (coded as 1 if such a condition is reported, and 0 otherwise), and the age group to which the individual belongs. It is important to note that there is a degree of differentiation in the number of variables included in the EILU 2014 compared to its subsequent cohort. For instance, parental academic background— namely, whether either the father or the mother has attained a university education (coded as 1 if so, and 0 if not)—is information provided exclusively in the second wave of the survey.

Regarding the graduate's education and learning information, this section encompasses dichotomous variables such as whether the student received a general scholarship or an excellence scholarship (coded as 1 if the support was received, and 0 otherwise) and proficiency in a foreign language (coded as 1 if the skill is present, and 0 otherwise). It also covers international mobility processes for master's studies, the use of information technologies, and the type of institution attended (public/private). In addition, fixed effects for the graduate's field of study were included. Finally, regarding the graduate's employment situation, dichotomous variables are included indicating whether the graduate works part-time (coded as 1 if so, and 0 if full-time) and whether they were employed while pursuing their degree.

It is important to note that the wage variable was constructed based on the median of the contribution-base quintiles of employees in the Social Security system. Given that this variable is not strictly continuous, the present study adopts a pragmatic semilogarithmic approximation, consistent with empirical practices in the economics of education literature that employ categorical variables with multiple levels. This approach allows the coefficients to be interpreted in terms of the average elasticity of the approximated log-wage, maintaining comparability across cohorts and avoiding information losses associated with dichotomous recoding. Recent literature supports this strategy[2].

Nevertheless, while acknowledging that the discretization of wage bands entails a loss of information within each earnings quintiles, the robustness of this approximation was verified by estimating an ordered probit model (table VII), given the ordinal nature of the wage variable in the EILU.

The difference-in-differences estimator was subsequently employed to isolate the "causal" effect of postgraduate education on graduates' salaries. Numerous empirical studies have used this technique in the field of education[3].

In this study, the treatment group comprises university graduates who completed a master's degree between their first and last observed years of labour market insertion following graduation from their bachelor's program. The control group, by contrast, consists of graduates who did not pursue a master's degree over the same period. The pre-treatment period corresponds to the wage reported in the first year after completing undergraduate studies, while the post-treatment period refers to the wage observed in the last year of the insertion window captured by the EILU survey. Consequently, the DID estimator compares the evolution of wages between both groups across these two points in time, attributing any differential change to the effect of postgraduate education.

The identification strategy relies on a Difference-in-Differences (DID) framework that compares wage dynamics between graduates who completed a master's degree (treatment group) and those who did not (control group), before and after postgraduate education. Under the parallel trends assumption, the DID estimator captures the average treatment effect of obtaining a master's degree on wages by netting out time-invariant differences between groups and common shocks over time.

The empirical specification is given by:

$yi\,t = \alpha + \beta 1 \text{Master}i + \gamma Postt + \delta (\text{Master}i * \text{Post}t) + Xit\,\theta + \varepsilon it$ (ii)

where Master_i identifies individuals in the treatment group, *Postt* denotes the post-treatment period, and the interaction term captures the DID estimator. The vector Xit includes a rich set of controls, and εit is an error term. The parameter of interest δ measures the differential change in wages between treated and control individuals over time.

The model is estimated using two alternative specifications, one including degree-program fixed effects and a full set of covariates, and another incorporating individual fixed effects to account for time-invariant unobserved heterogeneity.

Although the proposed model constitutes a powerful tool for evaluating the returns to postgraduate education, it is important to recognize its limitations concerning the availability of a homogeneous dataset, which must include information on the outcome variable both before and after the treatment, and for individuals who received the treatment as well as for those who did not. Additionally, it is necessary to consider the presence of other unobserved factors. Thus, the term X_i represents a vector of covariates that may influence the outputs of the educational process. These variables correspond to those previously included in the first model presented in this section.

Another potential consideration within the employed methodology lies in validating the parallel trends assumption, which states that the trajectory of the outcome variable for both groups would have been identical in the absence of treatment. In this particular case, the EILU faces two limitations for verification: (i) On the one hand, there is no data available on salaries in the period prior to the completion of university studies; ii) Additionally, the sample of postgraduate students has an over-representation of individuals whose characteristics are correlated with factors that do not allow comparing, prior to the insertion of cofactors, the trend of salaries[4].

In this sense, and as previously noted, a rich set of covariates is included to reduce heterogeneity across individuals and improve identification. The treatment is assumed to occur after the graduate's first employment. To mitigate potential endogeneity in the decision to pursue a master's degree, the DID framework accounts for time-invariant unobserved heterogeneity and common shocks across cohorts, although it cannot fully eliminate selection into postgraduate education. The validity of the empirical strategy is further assessed through robustness checks, including alternative specifications that account for the ordinal nature of the wage variable, stability tests of the estimated coefficients, and placebo regressions using non-wage outcomes.

Results

Selection into master's education and employment

Table 1 shows substantial observable selection into master's completion. In the 2014 cohort, general and excellence scholarships were associated with increases of 11.5 and 14.6 percentage points, respectively, in the probability of completing a master's degree. In 2019, age became particularly relevant. Relative to graduates younger than 30, the probability was 18.4 percentage points lower among those aged 30–34 and 28.2 points lower among older graduates. Excellence scholarships, previous work, mobility, distance education and parental higher education were also positively associated with master's completion. Field-of-study and regional effects were jointly significant in both cohorts, confirming that master's completion was not independent of the observed characteristics.

Table 1
Observable selection into master's degree completion[5]

Variable	EILU 2014	EILU 2019
Female	0.31 (0.49)	0.07 (0.57)
Disability	—	1.95 (2.52)
Age 30–34	−2.43*** (0.63)	−18.39*** (0.66)
Age 35 or older	−0.94 (0.84)	−28.22*** (0.81)
General scholarship	11.46*** (0.97)	0.67 (0.60)
Excellence scholarship	14.61*** (5.28)	3.13** (1.27)
Public university	0.51 (0.64)	−2.01** (0.92)
Previous work	—	3.95*** (0.77)
Foreign-language indicator	—	0.14 (0.54)
Mobility	0.74 (0.60)	6.26*** (0.69)
Distance modality	—	4.65** (2.15)
Father with higher education	—	4.12*** (0.63)
Mother with higher education	—	4.19*** (0.64)
Family-size measure	—	−0.37* (0.21)
Joint Wald tests		
Age categories	$\chi^2(2) = 17.02$***	$\chi^2(2) = 1{,}339.80$***
Field-of-study effects	$\chi^2(29) = 170.06$***	$\chi^2(31) = 2{,}290.65$***
Regional effects	$\chi^2(17) = 149.31$***	$\chi^2(17) = 98.48$***
Observations	10,039	30,278

Table 2 documents marked differences in employment trajectories. Master's graduates had lower adjusted employment probabilities at the beginning of both periods, 7.1 percentage points in 2011 and 8.6 points in 2015. In the 2019 cohort, this association reversed by the end of the period, when master's graduates exhibited a 2.3-percentage-point higher adjusted employment probability. Transition rates indicate delayed labour-market entry rather than persistent exclusion, 46.7% of master's graduates entered Social Security-covered employment between 2015 and 2019, compared with 26.8% of non-master's graduates. Consequently, the wage results are interpreted as conditional on Social Security-covered employment and the availability of contribution-base information.

Table 2

Master's status and Social Security-covered employment[6] Panel A. Employment rates and adjusted associations

Cohort	Year	No master (%)	Master (%)	Raw gap (pp)	Adjusted AME (pp)
EILU 2014	2011	40.41	33.58	−6.83***	−7.10*** (1.89)
	2014	54.66	66.43	11.77***	−5.41** (2.09)
EILU 2019	2015	57.06	37.02	−20.05***	−8.60*** (0.55)
	2019	78.60	79.21	0.61	2.29*** (0.47)
Panel B. Employment transitions					

Master's degree wage differentials across cohorts

Tables 3 and 4 present estimates of the wage differential associated with completing a master's degree for the 2014 and 2019 EILU cohorts. For each cohort, the first two specifications estimate conventional Mincer equations at the beginning and end of the observation period, while Model 3 pools both years. Models 4 and 5 introduce a difference-in-differences (DiD) specification, first with degree-program fixed effects and the covariates described in the methodological section and then with individual fixed effects. The treatment variable identifies graduates who completed a master's degree, whereas the comparison group comprises graduates who did not complete one. The estimates are interpreted in relation to evidence from advanced countries with high higher-education enrolment, where the returns to master's education vary with labour market conditions, field of study, and the comparison population (Song et al., 2008; Okahana and Hao, 2019; Kakizawa et al., 2014).

Table 3

Estimates of the wage premium of studying a master's degree (EILU − 2014)[7]

Employment transition	2011–2014 No master	2011–2014 Master	2015–2019 No master	2015–2019 Master
Not employed in either year	32.53	27.35	16.15	16.32
Entered employment	27.05	39.07	26.78	46.66
Left employment	12.81	6.22	5.25	4.47
Employed in both years	27.60	27.36	51.82	32.55
Pearson test	$\chi^2(3) = 137.63$***		$\chi^2(3) \approx 1{,}500$***	

Table 4

Estimates of the salary premium for studying a master's degree (EILU – 2019)[8]

VARIABLES	(1)	(2)	(3)	(4)	(5)
	OLS 2011	OLS 2014	OLS pooled	DID deg FE	DID ind FE
master	-0.111***	-0.055***	-0.072***	-0.125***	
	-0.017	-0.009	-0.009	-0.014	
period2			0.259***	0.233***	0.246***
			-0.008	-0.009	-0.009
interaction				0.089***	0.083***
				-0.015	-0.02
partial	-0.345***	-0.447***	-0.451***	-0.451***	
	-0.018	-0.01	-0.01	-0.01	
exper	0.025***	0.026***	0.027***	0.027***	
	-0.002	-0.001	-0.001	-0.001	
female	-0.055***	-0.065***	-0.064***	-0.063***	
	-0.014	-0.008	-0.008	-0.008	
age3034	0.245***	0.112***	0.131***	0.131***	
	-0.016	-0.009	-0.009	-0.009	
age35hi	0.518***	0.301***	0.327***	0.325***	
	-0.019	-0.011	-0.011	-0.011	
general_scholars.	-0.120***	-0.100***	-0.099***	-0.099***	
	-0.015	-0.008	-0.008	-0.008	
excellence scholarship	0.043	0.039*	0.044**	0.045**	
	-0.042	-0.023	-0.022	-0.022	
languages	0.020**	0.018***	0.019***	0.018***	
	-0.008	-0.005	-0.004	-0.004	
mobility	-0.061***	-0.026**	-0.040***	-0.041***	
	-0.023	-0.012	-0.012	-0.012	
public univ.	-0.058***	-0.032***	-0.031***	-0.028***	
	-0.019	-0.011	-0.011	-0.011	
work&study	-0.026*	-0.011	0.003	0.003	
	-0.015	-0.008	-0.008	-0.008	
constant	6.861***	7.117***	6.939***	6.949***	7.029***
	-0.044	-0.025	-0.025	-0.025	-0.006
Observations	9,994	14,570	24,564	24,564	25,412
Adjusted R^2	0.308	0.296	0.331	0.332	0.554
Robust standard errors in parentheses					
*** p < 0.01, ** p < 0.05, * p < 0.1					

For the 2014 cohort (Table 3), the cross-sectional estimates show a negative association between master's completion and earnings. Master's completers earned approximately 10.5% less than non-completers in 2011 and 5.4% less in 2014, while the pooled specification indicates a 6.9% wage disadvantage. All three coefficients are statistically significant. Although these estimates appear inconsistent with the positive returns predicted by human capital theory, they should not be interpreted as evidence that completing a master's degree reduces wages. Rather, they describe conditional differences between graduates who followed distinct educational and employment trajectories. The negative coefficients may reflect differences in initial labour market attachment, the timing of postgraduate enrolment, occupational choices, or other characteristics associated with master's participation and completion. Potential selection arising from unobserved characteristics and sample heterogeneity has been widely discussed in this literature (Heckman and Li, 2004; Suga, 2019; Flores and Vigier, 2020).

Models 4 and 5 use the structure of the data to compare wage changes over time. In the specification with degree-program fixed effects, the interaction between master's completion and the post-treatment period is 0.089, equivalent to an additional wage increase of approximately 9.3% between 2011 and 2014 for master's completers relative to non-completers. The adjusted (R^2) is 0.332. This coefficient captures the differential wage change associated with completing a master's degree after accounting for the observed controls, common temporal changes, and time-invariant differences across degree programs. It does not, however, establish an unconditional causal return.

The result remains similar when individual fixed effects are introduced. Model 5 produces an interaction coefficient of 0.083, corresponding to an 8.7% differential wage increase, and an adjusted (R^2) of 0.554. Individual fixed effects control for time-invariant personal characteristics—such as persistent ability, motivation, or family background—that may influence both postgraduate completion and earnings. They therefore mitigate bias arising from stable individual heterogeneity, although they cannot eliminate selection associated with unobserved factors that vary over time. The proximity of the estimates in Models 4 and 5 nevertheless indicates that the central result is not driven exclusively by permanent differences across individuals or degree programs.

The remaining coefficients in Model 4 reveal familiar wage patterns. Women earned approximately 6.1% less than men, while each additional year of professional experience was associated with a 2.7% increase. Relative to graduates under 30, those aged 30–34 earned 14.0% more and those aged 35 or above earned 38.4% more, consistent with greater accumulated experience and labour market attachment. Part-time employment displayed the largest negative association, wages were approximately 36.3% below those of full-time workers.

Several educational characteristics were also significantly associated with earnings. A general scholarship, mobility during university, and attendance at a public university were associated with wage differences of − 9.4%, − 4.0%, and − 2.8%, respectively. By contrast, an excellence scholarship was associated with a 4.6% premium and knowledge of an additional language with a 1.8% premium. Working while studying was not statistically significant. Significant wage differences were also observed across autonomous communities, although these coefficients are omitted from the table to conserve space.

For the 2019 cohort (Table 4), the initial cross-sectional pattern is again negative, but the subsequent change is more pronounced. In 2015, master's completion was associated with wages approximately 8.2% lower than those of non-completers. By 2019, the coefficient had become positive and statistically significant, indicating a 3.3% wage advantage. In the pooled specification, however, the coefficient is close to zero and statistically insignificant. This progression suggests that the positive differential emerged during the graduates' early labour market trajectories rather than being present at the initial observation.

The DiD specifications indicate a substantially larger differential wage change for this cohort. Model 4 produces an interaction coefficient of 0.229, equivalent to an additional wage increase of approximately 25.8% between 2015 and 2019 for master's completers relative to non-completers. The estimate is almost unchanged when individual fixed effects are introduced. Model 5 yields a coefficient of 0.228, corresponding to a 25.6% differential increase. The adjusted (R^2) values are approximately 0.42 and 0.48, respectively. The close correspondence between both estimates supports the stability of the result across alternative fixed-effects structures, while retaining the same caution regarding unobserved time-varying differences.

The covariates display patterns broadly consistent with the earlier wave, although their magnitudes differ. Women earned 4.2% less than men, each additional year of professional experience was associated with a 1.8% increase, and part-time workers earned approximately 25.0% less than full-time workers. Graduates aged 35 or above earned 17.1% more than those under 30, whereas the coefficient for the 30–34 group was small and statistically insignificant. Working while studying was associated with wages 2.8% higher. Excellence scholarships, marriage, paternal higher education, and distance education were positively associated with earnings, with estimated differentials of 7.5%, 6.7%, 1.7%, and 4.1%, respectively. Conversely, general scholarships, attendance at a public university, belonging to a large family, and knowledge of additional languages were associated with wage differences of − 6.8%, − 7.1%, − 1.8%, and − 1.7%. Maternal higher education and mobility were not statistically significant. Regional wage differences remained substantial, supporting the inclusion of autonomous-community controls.

These results are consistent in direction with previous studies that use alternative strategies to address educational selection. Using the same EILU 2019 wave, Caparrós (2023) also obtains a small negative OLS coefficient and a positive estimate of approximately 15% after accounting for selection. Minaya et al. (2024), using administrative data and individual fixed effects for graduates in Ohio, report an average return of approximately 14% and find that economic downturns reduce, but do not eliminate, positive returns, with substantial variation across fields and demographic groups. The estimate obtained here for Spain is larger, but the common positive direction provides a useful external-validity comparison. Direct equivalence should nevertheless be avoided because the institutional setting, labour market, outcome construction, and composition of the samples differ.

The comparison between cohorts therefore reveals an increase in the individual-fixed-effects estimate from 8.7% in the first wave to 25.6% in the second. This pattern coincides with the overall evolution of earnings during the early transition to the labour market. Earnings for the 2009–2010 graduating cohort increased by only 13.0% during its observation period, whereas the 2013–2014 cohort experienced growth of 49.6% between the initial and final employment observations. The larger master's differential in the second wave is thus consistent with the more favourable wage trajectory observed during the economic recovery. It does not, by itself, identify a causal effect of the economic cycle, since the two cohorts also differ in institutional setting, educational composition, and other potentially relevant circumstances.

The macroeconomic context offers one plausible interpretation. Sans-de-Galeano and Terskaya (2017) observe that the onset of the Great Recession initially increased average earnings in Spain, partly because younger, less-qualified, and foreign workers were disproportionately displaced from employment. From 2010 onward, however, wages declined markedly and continued falling until 2013, followed by only a slight recovery in 2014. This period closely overlaps with the labour market entry and observation window of the first EILU cohort. The second cohort, by contrast, entered the labour market under more favourable aggregate conditions. The evidence is therefore consistent with differences across cohorts exposed to contrasting macroeconomic environments, rather than demonstrating that the cycle itself caused the change in the estimated master's premium.

Institutional change provides an additional explanation. During the analysed period, the Bologna reform reorganized Spanish university education into bachelor's, master's, and doctoral cycles. The 2014 EILU cohort still contained many graduates from the former three-year *diplomatura* and five-year *licenciatura* programs, whereas graduates from the new four-year *grado* structure were underrepresented (Fig. 1). The 2019 cohort was considerably more concentrated in the *grado–máster* structure. Consequently, in the later cohort, a master's degree may capture part of the specialization previously incorporated into the fourth and fifth years of the *licenciatura*. The simultaneous expansion of master's education, the Bologna transition, and improving macroeconomic conditions prevents the cross-cohort difference from being attributed exclusively to any single mechanism.

The potential heterogeneity of the estimates was examined by gender, age, and field of study using the DiD model with individual fixed effects in each subsample. Tables 5 and 6 report the results by gender and age. In the 2014 cohort, the differential associated with master's completion was positive and statistically significant for women (9.7%), men (6.8%), graduates under 30 (6.3%), and those aged 30–34 (7.3%), but not for graduates aged 35 or above. In the 2019 cohort, the estimates were again significant for women (20.0%), men (19.1%), graduates under 30 (15.8%), and those aged 30–34 (14.5%). The coefficient for graduates aged 35 or above was small and statistically insignificant. For this older group, wages may depend more strongly on accumulated career experience and earlier occupational trajectories than on recently completed postgraduate education.

Table 5
Heterogeneity of estimates according to sex and age (EILU – 2014)

VARIABLES	(1)	(2)	(3)	(4)	(5)
	OLS 2015	OLS 2019	OLS pooled	DID deg FE	DID ind FE
master	-0.086***	0.032***	-0.003	-0.148***	
	-0.011	-0.005	-0.006	-0.011	
period2			0.588***	0.500***	0.486***
			-0.006	-0.007	-0.008
interaction				0.229***	0.228***
				-0.012	-0.015
partial	-0.192***	-0.342***	-0.290***	-0.288***	
	-0.016	-0.008	-0.008	-0.008	
exper	0.020***	0.014***	0.018***	0.018***	
	-0.001	-0.001	0	0	
female	-0.040***	-0.070***	-0.044***	-0.043***	
	-0.011	-0.005	-0.005	-0.005	
age3034	0.126***	-0.027***	0.007	0.009	
	-0.013	-0.006	-0.006	-0.006	
age35hi	0.391***	0.029***	0.157***	0.158***	
	-0.017	-0.009	-0.009	-0.009	
public	-0.103***	-0.053***	-0.074***	-0.074***	
	-0.016	-0.009	-0.009	-0.008	
distance edu	0.054*	0.008	0.042**	0.040**	
	-0.028	-0.018	-0.017	-0.017	
large_family	-0.014***	-0.019***	-0.017***	-0.018***	
	-0.004	-0.002	-0.002	-0.002	
father_hiedu	0	0.025***	0.017***	0.017***	
	-0.012	-0.006	-0.006	-0.006	
mother_hiedu	-0.011	0.011*	0.007	0.006	
	-0.013	-0.006	-0.006	-0.006	
married	0.093***	0.040***	0.065***	0.065***	
	-0.012	-0.007	-0.007	-0.007	
general scholars.	-0.084***	-0.056***	-0.069***	-0.070***	
	-0.011	-0.006	-0.006	-0.006	
excellence scholarship	0.051*	0.059***	0.073***	0.072***	
	-0.028	-0.011	-0.012	-0.012	
languages	-0.020**	-0.018***	-0.017***	-0.017***	
	-0.01	-0.005	-0.005	-0.005	
mobility	-0.008	-0.004	-0.003	-0.005	
	-0.017	-0.007	-0.007	-0.007	
work&study	0.056***	0.028***	0.026***	0.028***	
	-0.012	-0.006	-0.006	-0.006	
constant	7.093***	7.876***	7.223***	7.274***	7.061***
	-0.061	-0.033	-0.016	-0.017	-0.005

Table 6
Heterogeneity of estimates according to sex and age (EILU – 2019)

VARIABLES	(1)	(2)	(3)	(4)	(5)
	DID women	DID men	DID age < 30	DID age 30–34	DID age > = 35
period2	0.249***	0.240***	0.376***	0.214***	0.054***
	(0.013)	(0.013)	(0.018)	(0.014)	(0.011)
interaction	0.093***	0.066**	0.061*	0.070**	0.029
	(0.028)	(0.027)	(0.033)	(0.035)	(0.025)
constant	6.932***	7.169***	6.788***	7.101***	7.522***
	(0.008)	(0.008)	(0.011)	(0.009)	(0.006)
Observations	15,210	10,347	13,880	6,935	4,742
Adjusted R^2	0.497	0.608	0.448	0.552	0.701
Robust standard errors in parentheses					
*** p < 0.01, ** p < 0.05, * p < 0.1					

Tables 7 and 8 present the estimates by field of study. In the 2014 cohort, the interaction was positive and statistically significant for Arts and Humanities (27.6%), Sciences (14.0%), and Engineering and Architecture (5.9%, significant at the 10% level), but not for Social and Legal Sciences or Health Sciences. In the 2019 cohort, the differential was positive and significant across all five fields: Arts and Humanities displayed the largest estimate (39.8%), followed by Sciences (30.7%), Engineering and Architecture (23.2%), Social and Legal Sciences (18.9%), and Health Sciences (8.4%). These results show that the larger differential in the later cohort was not confined to a single discipline, although its magnitude varied considerably across fields.

Table 7
Heterogeneity of estimates according to branch of knowledge (EILU – 2014)

VARIABLES	(1)	(2)	(3)	(4)	(5)
	DID women	DID men	DID age < 30	DID age 30–34	DID age > = 35
period2	0.472***	0.502***	0.717***	0.542***	0.251***
	(0.011)	(0.011)	(0.018)	(0.013)	(0.009)
interaction	0.182***	0.175***	0.147***	0.135***	0.018
	(0.019)	(0.021)	(0.026)	(0.025)	(0.016)
constant	7.033***	7.151***	6.795***	7.018***	7.538***
	(0.007)	(0.007)	(0.010)	(0.008)	(0.005)
Observations	19,609	14,722	15,537	10,031	8,763
Adjusted R^2	0.446	0.499	0.452	0.477	0.606
Robust standard errors in parentheses					
*** p < 0.01, ** p < 0.05, * p < 0.1					

Table 8
Heterogeneity of estimates according to branch of knowledge (EILU – 2019)

VARIABLES	(1)	(2)	(3)	(4)	(5)
	DID arts_hum	DID sciences	DID socsc	DID eng_arq	DID heathsc
period2	0.145***	0.311***	0.264***	0.245***	0.180***
	(0.038)	(0.038)	(0.014)	(0.013)	(0.032)
interaction	0.244***	0.131**	0.042	0.057*	0.068
	(0.064)	(0.065)	(0.031)	(0.033)	(0.066)
constant	6.873***	6.870***	6.935***	7.258***	7.095***
	(0.023)	(0.025)	(0.009)	(0.008)	(0.019)
Observations	1,959	2,187	11,514	6,374	3,523
Adjusted R^2	0.640	0.472	0.525	0.617	0.411
Robust standard errors in parentheses					
*** p < 0.01, ** p < 0.05, * p < 0.1					

The subgroup intercepts and period coefficients provide additional descriptive evidence on initial wage levels and wage progression. In the 2014 cohort, the exponentiated intercepts imply baseline monthly wages of approximately €1,298 for men and €1,025 for women. By age, the corresponding values range from approximately €888 for graduates under 30 to €1,849 for those aged 35 or above. The post-period coefficients imply wage growth of 45.7% among graduates under 30, 23.9% among those aged 30–34, and 5.5% among those aged 35 or above. The same age gradient appears more strongly in the 2019 cohort. Baseline monthly wages range from approximately €894 for the youngest group to €1,878 for the oldest, while the corresponding increases over the observation period are 104.8%, 71.9%, and 28.5%. These estimates indicate faster wage progression among younger graduates, who began from substantially lower earnings levels.

Overall, master's completion was associated with a positive differential wage trajectory in both cohorts, with a markedly larger estimate in the 2019 wave. The consistency between degree-program and individual fixed-effects specifications, together with the broadly positive subgroup results, strengthens the descriptive evidence. Nevertheless, the estimates should be interpreted as differential wage changes associated with master's completion, not as definitive causal effects of postgraduate education or macroeconomic recovery.

Robustness checks

Two complementary exercises assess the robustness of the main findings. First, the heterogeneity estimates reported in Tables 5–8 show that the positive differential is observed for both sexes and for graduates under 35 in both cohorts. By field of study, it is concentrated in selected fields in 2014 but becomes positive and significant across all five branches in 2019. These results suggest that the aggregate estimate is not driven by a single demographic or disciplinary group, while confirming meaningful variation in its magnitude.

Second, Table 9 addresses the interval-censored nature of the EILU wage measure. Rather than assigning the midpoint of each contribution-base quintile to every respondent, the interval regression uses the corresponding lower and upper bounds directly. Observations in the first quintile are treated as left-censored, those in the three intermediate quintiles as interval-censored, and those in the highest quintile as right-censored. The model includes 21,172 observations: 3,923 left-censored, 12,857 interval-censored, and 4,392 right-censored.

Table 9
– Interval regression of the administrative contribution base, 2019 (EILU 2019)

VARIABLES	(1)	(2)	(3)	(4)	(5)
	DID arts_hum	DID sciences	DID socsc	DID eng_arq	DID heathsc
period2	0.399***	0.503***	0.503***	0.497***	0.401***
	(0.039)	(0.046)	(0.011)	(0.013)	(0.028)
interaction	0.335***	0.268***	0.173***	0.209***	0.081**
	(0.063)	(0.068)	(0.020)	(0.027)	(0.041)
constant	6.892***	6.887***	7.014***	7.247***	7.229***
	(0.027)	(0.029)	(0.007)	(0.008)	(0.015)
Observations	2,461	2,639	15,950	8,407	4,874
Adjusted R^2	0.475	0.292	0.477	0.470	0.465
Robust standard errors in parentheses					
*** p < 0.01, ** p < 0.05, * p < 0.1					

Variable	Coefficient	Robust SE	z	p-value	95% CI	Implied change
Master's degree	0.0408***	0.0060	6.78	< 0.001	[0.0290, 0.0526]	4.17%

Conditional on the complete set of controls, master's completion is positively associated with the latent monthly contribution base (coefficient = 0.0408; robust SE = 0.0060; p < 0.001; 95% CI: 0.0290–0.0526), corresponding to an approximate 4.2% differential. This supports the positive direction of the 2019 estimates without relying on midpoint imputation. Its magnitude is not directly comparable with the DiD interaction because the interval regression uses the final-period cross-section. Overall, the exercise strengthens the robustness of the wage measurement, but does not eliminate time-varying unobserved selection or establish a causal effect.

Conclusions

This study examined wage differentials associated with master's degree completion during early labour market entry, using the nationally representative 2014 EILU cohort (2009–2010 graduates) and 2019 cohort (2013–2014 graduates). It contributes evidence for Spain and a European setting in which postgraduate expansion coincided with the Bologna transition and contrasting macroeconomic conditions. It also compares master's completers with graduates who did not complete one and follows their employment and wage trajectories during the first years after graduation.

The empirical strategy first examined selection into master's completion and employment and then estimated OLS and difference-in-differences (DiD) wage models with degree-program and individual fixed effects. The models include personal, educational, employment, field-of-study, and regional controls. Individual fixed effects account for time-invariant characteristics that may influence both postgraduate completion and earnings, mitigating bias from stable unobserved heterogeneity. However, they cannot eliminate selection related to unobserved factors that change over time. The findings are therefore interpreted as differential wage changes associated with master's completion, not definitive causal effects.

The selection analysis confirms that master's completion is non-random and that its predictors differ across cohorts. In 2014, scholarships and language skills were among its principal educational correlates. In 2019, completion was more strongly associated with age, excellence scholarships, mobility, working while studying, parental education, university type, field, and region. Master's graduates initially displayed lower employment probabilities in both cohorts. This disadvantage remained in 2014, although it narrowed, whereas in 2019 it disappeared and became a small positive differential by the final observation. Specifically, the adjusted final-period differential was − 5.4 percentage points in 2014 and + 2.3 percentage points in 2019. These patterns underline the importance of examining employment selection before interpreting wages observed only among employed graduates.

The wage estimates reveal a consistent temporal pattern. For the 2014 cohort, the OLS specifications produced negative conditional associations between master's completion and earnings. Once wage changes were compared longitudinally, the differential became positive: 9.3% with degree-program fixed effects and 8.7% with individual fixed effects. For the 2019 cohort, the cross-sectional estimate changed from an 8.2% disadvantage in 2015 to a 3.3% advantage in 2019. The corresponding DiD estimates reached 25.8% and 25.6%, respectively. Master's completers therefore experienced stronger early-career wage growth in both cohorts, with a considerably larger differential in the later wave.

Complementary analyses support this pattern while revealing heterogeneity. Positive differentials were found for both sexes and graduates under 35 in both cohorts, whereas estimates for those aged 35 or above were not significant. Across fields, positive results were confined to selected branches in 2014 but appeared in all five branches in 2019. The interval regression addresses the grouped and censored nature of the administrative contribution-base measure. Using the original quintile bounds rather than imputed midpoints, master's completion was associated with a 4.2% higher latent monthly contribution base in 2019 (p < 0.001). Although this final-period estimate is not directly comparable with the DiD interaction, it confirms that the positive association does not depend exclusively on midpoint imputation.

The increase from 8.7% to 25.6% coincides with contrasting macroeconomic environments. Average earnings grew by only 13.0% during the observation period of the first cohort, compared with 49.6% for the later cohort. This is consistent with the possibility that entering the labour market during the Great Recession constrained early wage progression, while the second cohort benefited from recovery. The results align in direction with Caparrós (2023), who reports a 15% estimate for Spain, and provide a European comparison with Minaya et al. (2024), who estimate an average return of approximately 14% in Ohio and find that downturns reduce, but do not eliminate, positive returns. However, the EILU comparison cannot isolate macroeconomic conditions from concurrent institutional changes. The transition from the former *diplomatura* and *licenciatura* structure to the *grado–máster* system altered both graduate composition and the meaning of postgraduate specialization.

The study therefore documents nationally representative differences in master's selection, employment access, and early-career wage progression across two cohorts exposed to contrasting economic and institutional conditions. This national coverage and the comparison with non-completers provide an institutional complement to US evidence based on single-state administrative records. The results support closer coordination between postgraduate provision and labour market demand, stronger career guidance, and transition support during downturns. They also caution against treating master's degrees as producing a uniform premium across graduates, fields, or periods. Longer panels and designs capable of addressing time-varying selection are required before attributing the observed differentials causally to master's completion or macroeconomic recovery.

Declarations

Author Contribution

S.T.R. contributed to the conceptualization of the study, methodology, data curation, formal analysis, investigation, and preparation of the original manuscript. I.I. contributed to the conceptualization and methodology, supervised the research, validated the results, and critically reviewed and edited the manuscript. Both authors read and approved the final manuscript.

Data Availability

The microdata from the 2014 and 2019 waves of the Survey on the Labour Insertion of University Graduates (EILU) are available from the Spanish National Statistics Institute (INE) at the following URL: https://www.ine.es/dyngs/INEbase/es/operacion.htm?c=Estadistica_C&cid=1254736176991&menu=ultiDatos&idp=1254735976597.

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31. Annex: 1 – Econometric estimates by OLS and difference-in-difference estimator

Footnotes

1. See: (Frederiksen & Kato, 2011), (Mendoza, 2016) and (Biasi et al., 2021), for an extend explanation of the human capital benefits.
2. See: Gándara and Toutkoushian (2017); Okahana and Hao (2019); and Caparrós (2023).
3. See: Duflo (2001); Blundell et al. (2001); Lechner (2010); Fredriksson and Magalhaes de Oliveira (2019); Schwerdt and Woessmann (2020) and Greene and Liu (2020).
4. See: Jaeger, Joyce, and Kaestner (2018), Kanh-Lane and Lang (2020) and Schwerdt and Woessmann (2020), for similar studies.
5. Notes: Entries in the first panel are average marginal effects from cohort-specific logit models, with robust standard errors in parentheses. The omitted age category is younger than 30. All models include field-of-study and regional effects. *p < 0.10, **p < 0.05, ***p < 0.01.
6. Employment is defined from Social Security affiliation. Raw rates and transitions use the available descriptive samples (10,677 and 30,473 observations). Adjusted AMEs are obtained from year-specific probit models with robust standard errors and the controls described in Table X; their estimation samples contain 9,738 and 30,278 observations, respectively. The raw and adjusted estimates therefore need not coincide exactly. *p < 0.10, **p < 0.05, ***p < 0.01.
7. Controls of degree/field of study and Autonomous Community were included in models 1 and 2, however, for better visualization, these coefficients were not included in the table.
8. Controls of degree/field of study and autonomous community were included in models 1 and 2, however, for better visualization, these coefficients were not included in the table.

Figures

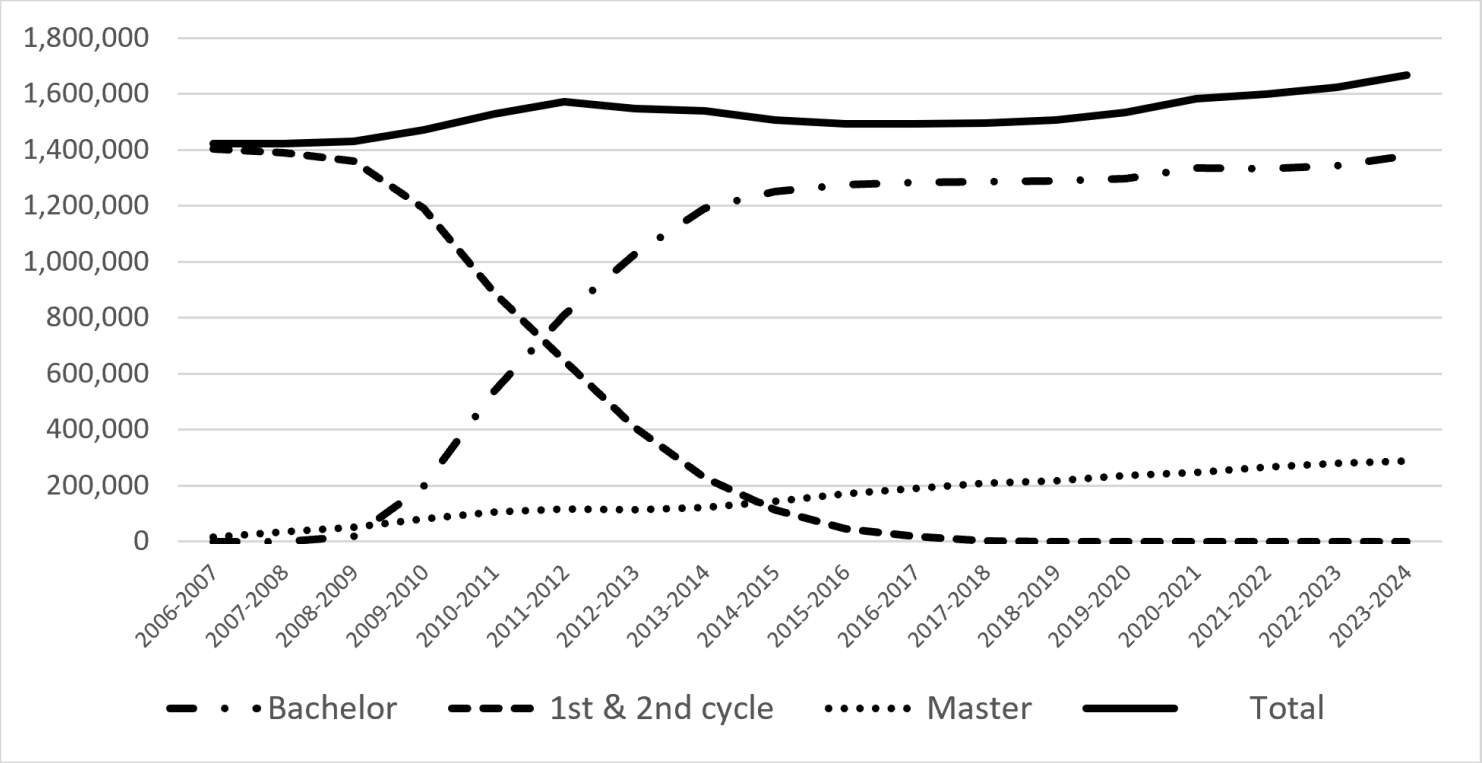


Figure 1

Students enrolled in the Spanish University System according to the level of studies (excluding PhD)