When Certification Hurts: Accreditation Signals and Graduate Wages

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Employers cannot directly observe program quality, and institutional reputation provides reliable signals for only a small fraction of graduates. We study whether program-level accreditation in Chile resolves this uncertainty, using a staggered difference-in-differences design applied to administrative earnings records for 2,572 programs and exploiting first-time accreditation decisions between 2007 and 2018. Accreditation raises wages by approximately 3 percent on average, with the premium concentrated among cohorts graduating four to five years after the initial award. This timing reflects a compositional channel, as accreditation attracts higher-ability entrants while simultaneously expanding graduation to marginal completers whose weaker formal-sector attachment reduces the observed employment rate of the graduate pool. Consistent with a Bayesian employer-learning framework, the wage premium reaches 8.9 percent at lower-reputation institutions but turns negative at elite programs, where the certification label adds no new information to already-precise employer priors. Accreditation functions as a corrective signal that redistributes information rents toward graduates whose credentials would otherwise be discounted.

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1 Introduction

Consider the information problem facing a recruiter who receives a resume from a graduate of a program they have never encountered. A business degree from a regional university, perhaps, or an engineering credential from one of the hundreds of private institutions that have entered the market in recent decades. The recruiter cannot inspect the curriculum, cannot verify faculty quality, and cannot observe the underlying productivity of the candidate. Institutional reputation fills this gap, but only partially since it functions as a reliable signal for a handful of long-standing elite universities and provides almost nothing for the rest of the distribution (MacLeod and Urquiola, 2015). For the vast majority of graduates, employers and workers are left in a state of mutual informational deficit. Third-party certification is the institution designed to resolve it. Whether real-world accreditation systems actually do so, however, is far from obvious.

The theoretical prediction from a Bayesian framework of employer learning is precise (Altonji and Pierret, 2001). When employers aggregate noisy proxies for worker productivity, such as institutional prestige, prior hiring experience, alumni networks, their confidence in those estimates varies systematically across the market. At programs where these proxies already yield high-confidence predictions, a new public certification adds little marginal information since the employer's prior was already tight. At programs where priors are weak or dispersed, a verifiable quality label resolves genuine uncertainty and allows the market to update its valuation of the graduate. The implication is that accreditation's value is not uniform. It is an inverse function of the precision of employer prior beliefs. Accreditation should matter most where reputation is weakest. Whether this prediction holds in actual labor markets, and with what magnitude, requires direct empirical evidence.

We investigate three questions. First, what is the causal effect of first-time program accreditation on graduate wages? This establishes the direct return to the signal in the labor market. Second, what is the effect on the formal employment rate of graduates? A wage premium is consistent with two very different labor market stories. Either the signal helps the same graduates get better jobs, or it draws additional graduates into formal employment, changing the composition of who is observed earning a wage. Distinguishing between them is necessary to interpret the wage result. Third, do these effects vary with institutional reputation in the direction the signaling model predicts, largest where employer

uncertainty is highest and negligible or negative where it is lowest? This tests whether accreditation substitutes for brand or merely validates it redundantly.

We answer these questions using Chile's higher education system between 2007 and 2018, which provides an ideal quasi-experimental setting. There is high institutional heterogeneity, a centralized and transparent accreditation registry, and rich administrative data linking graduation records from the Ministry of Education to private-sector earnings reported to the tax authority. This yields a panel of nearly 25,000 program-cohort observations across 2,572 degree programs. The key to identification is the staggered adoption of first-time accreditation decisions. Programs choose when to seek evaluation, but the final decision (both the binary outcome and the duration awarded) is made by an independent panel of commissioners and is largely unanticipated by employers at the time of hiring (Barroilhet et al., 2022). We apply the imputation estimator of Borusyak et al. (2024) to address the negative-weight problems that arise in staggered designs under treatment effect heterogeneity. Pre-trend tests confirm that wages at treated and control programs followed identical trajectories before accreditation ($p > 0.60$ across all three wage horizons), supporting a causal interpretation.

First-time accreditation raises graduate wages by approximately 3 percent on average. The premium is not immediate; it accumulates over the post-accreditation event window and reaches statistical significance by the fifth cohort graduating after accreditation is received, coinciding with the average duration of college programs in Chile. This timing is consistent with the idea that accreditation carries limited pure informational value to the labor market. We show directly that the ability composition of graduates shifts upward following accreditation (the share of high-ability entrants rises significantly) while the socioeconomic composition remains unchanged, suggesting that employers respond primarily to the arrival of better-matched cohorts rather than to the certification signal itself. These results are consistent with the enrollment effects documented in Molina and Valdebenito (2026), and extend that evidence to the labor market.

The employment rate result confirms this picture. Following accreditation, graduation counts rise while the formal employment rate of the graduate pool falls, as the expanded graduation cohort includes marginal completers who do not appear in private-sector earnings records. Wages for those who do find formal employment remain elevated. The decline in the employment rate and the wage premium are therefore two sides of the same expansion: accreditation simultaneously attracts higher-

ability students at the top and draws in marginal completers at the bottom. That 3 percent average effect, however, conceals the paper's most important finding.

When we separate programs by institutional reputation, the results follow the signaling prediction. Graduates of programs at Baseline institutions, where the host institution holds the lowest certification and employer uncertainty is highest, realize a wage premium of 8.9 percent in the first year after graduation. Graduates of programs at Top-tier institutions realize a wage penalty of 5.0 percent. At elite institutions, the employer's prior knowledge about graduate quality is already precise and the accreditation label adds no new information. It may introduce ambiguity about why a highly reputed program required external validation in the first place. The same pattern holds when we stratify by the program's own first-accreditation term, as programs receiving a low-intensity first accreditation (1–3 years) generate a wage premium of 10.7 percent; programs receiving a high-intensity accreditation show no discernible gain. Accreditation generates the largest return precisely where employer priors are weakest. This gradient pins down the signaling channel in a way that a single average estimate cannot.

These findings make three contributions. First, we provide the first causal estimates of the labor market returns to program-level accreditation. The closest prior work, González-Velosa et al. (2015), estimates a 9.2 percentage-point wage premium per additional year of institutional accreditation using cross-sectional OLS, a design that cannot separate the effect of the label from pre-existing quality differences between programs. By exploiting quasi-experimental variation in the timing of first-time accreditation decisions, we provide causal evidence on how institutional reputation shapes graduate labor market outcomes (Dale and Krueger, 2002; Hastings et al., 2013; Hoekstra, 2009; MacLeod et al., 2017; Reyes et al., 2016). Beyond the average effect, we document the underlying mechanism. The ability composition of graduating cohorts shifts upward following accreditation, while the socioeconomic composition remains stable, and the formal employment rate falls as the expanded graduation pool includes marginal completers. The wage premium and the employment rate decline together reflect a single expansion in which programs simultaneously attract higher-ability students and draw in marginal completers at the bottom.

Second, the heterogeneity results provide the first direct empirical test of the employer learning prediction that the value of a quality signal is inversely related to the precision of prior beliefs (Altonji

and Pierret, 2001; Farber and Gibbons, 1996; Lange, 2007; Spence, 1973). The gradient from Baseline to Top-tier institutions, and from low-intensity to high-intensity accreditation terms, is exactly what a Bayesian updating framework predicts. Where institutional reputation already generates tight employer priors, the accreditation label is redundant and may carry a negative valence. Where priors are dispersed, the label resolves genuine uncertainty and is priced accordingly. This connects the accreditation setting to the broader literature on reputation and third-party certification (Dranove and Jin, 2010; Dranove et al., 2003; Elfenbein et al., 2015; Jin, 2005; Lizzeri, 1999; MacLeod and Urquiola, 2015; MacLeod et al., 2017), offering labor market evidence for a mechanism that has been theorized but rarely tested using quasi-experimental variation.

Third, the distributional implication of this mechanism is substantive. Because accreditation generates returns only where institutional reputation is weakest, the wage benefits flow disproportionately to graduates of lower-reputation programs, precisely those whose credentials would otherwise be discounted by employers. In higher education systems characterized by significant institutional heterogeneity, a common feature of education markets across the developing world, formal certification functions as a mechanism for redistributing information rents toward graduates who cannot rely on prestige alone (Dranove and Jin, 2010; Lizzeri, 1999). The policy question is therefore not whether accreditation improves average graduate outcomes, but whether it reduces the informational penalty that graduates at the bottom of the reputation distribution face.

The remainder of the paper proceeds as follows. Section 2 describes the institutional background. Section 3 presents the data sources and sample construction. Section 4 describes the identification strategy. Section 5 presents the main wage effects and the compositional and institutional evidence that supports a signaling interpretation. Section 6 examines how the wage effects vary with institutional reputation and program accreditation intensity, testing the informational mechanism directly. Section 7 concludes.

2 Institutional Background

2.1 The Accreditation System

Chile's higher education system encompasses hundreds of degree programs offered by institutions ranging from long-established public universities with decades of employer relationships to private providers with little observable track record. For graduates of elite programs, the institutional brand functions as a reliable proxy. Employers have decades of hiring experience and can form confident predictions about productivity. For graduates of the many programs where that reputational history is thin or absent, no equivalent shortcut exists. The graduate knows what their training was worth, but the employer does not. Without a way to distinguish quality, the market treats graduates of unrecognized programs as interchangeable and wages compress toward the average (Akerlof, 1970; MacLeod et al., 2017). Programs that invest in quality cannot capture the return because employers cannot observe the investment.

Chile's 2006 Quality Assurance Act created the National Accreditation Commission (CNA) to address this problem. CNA oversees the accreditation of undergraduate programs individually, rather than at the institutional level.[1] A university can have an accredited law program and an unaccredited business program in the same building. When a program passes evaluation, CNA grants an accreditation term of two to seven years based on its quality assessment. A short term signals marginal compliance; a longer term signals stronger performance across the Commission's criteria. The term structure generates a graduated public signal that is updated at each renewal cycle rather than a simple binary indicator. Accreditation is mandatory only for programs in medicine, dentistry, and education, where it is tied to professional licensing. We exclude these three fields from the analysis. For all remaining fields, accreditation is a voluntary strategic decision. Appendix **??** describes the evaluation criteria and process in detail.

[1] Institutional accreditation operates through a separate CNA channel and evaluates the university as an entity rather than any specific program. Its status is strongly correlated with measures of selectivity and prestige, making it a natural proxy for the precision of employer prior beliefs. Our heterogeneity analysis uses the length of institutional accreditation terms as a continuous measure of this prior precision (Barroilhet, 2019; Bordon and Fu, 2015).

2.2 Accreditation as an Information Shock

When CNA grants accreditation, the decision is published immediately in a public registry integrated into the government's official higher education information portal.[2] Employers screening candidates can consult it directly. The ruling, therefore, changes what a recruiter can verify about a program the morning after the Commission's decision. The timing of this change matters for wage outcomes. In Chile, graduation ceremonies occur at the end of the academic year. A program accredited mid-year will have a fraction of that year's graduates already entering the labor market carrying the new label, but the first cohort for whom the accreditation signal is observable throughout their job search is the one graduating in the year following the award.

A key feature of this setting is that employers are not tracking the accreditation pipeline. Programs submit applications and undergo evaluation over several months, but there is no public disclosure of a program's application status or the contents of the peer review process. Employers have no reason to anticipate which programs are under evaluation or when a ruling will arrive. The registry entry, therefore, functions as a genuine information shock from the employer's perspective, even though the program's decision to apply was voluntary and strategic. Furthermore, the presence of accreditation status in official government portals does not guarantee that employers actively search for it. Whether employers update wages in response to the signal at all is not obvious in advance. It is an empirical question our analysis answers.

Because programs decide when to seek their first accreditation based on their own assessment of readiness and competitive pressure, first-time accreditation decisions are distributed across the 2007 to 2018 sample period in a pattern driven by program-level strategy rather than any government-mandated schedule. This staggered voluntary adoption generates variation in the timing of the information shock that, from the employer's perspective, is effectively unpredictable. The endogeneity concern is about timing, as programs apply when they believe they are likely to pass, so treated programs may differ from never-treated ones in ways correlated with wages. This concern motivates robustness specifications that separately employ not-yet-accredited and never-accredited programs as comparison groups, each isolating a different dimension of the selection problem. What is not endoge-

[2] Accreditation status by program is publicly searchable at https://www.mifuturo.cl/buscador-de-acreditacion/ (SIES, Ministerio de Educación de Chile).

nous is the ruling itself. Once an application is filed, an independent panel of CNA commissioners issues a decision on both the binary outcome and the term length, and programs cannot control what that decision will be. The identification strategy, developed in Section 4, exploits the staggered timing of first accreditation as a source of variation that is unpredictable from the employer's side, relying on parallel trends across programs with different accreditation timing.

3 Data

3.1 Labor Market Outcomes

Formal private-sector employers in Chile report annual earnings to the Internal Revenue Service (SII) for every worker on payroll. Under a data access agreement with the Ministry of Finance (*Ministerio de Hacienda*), we obtained access to SII earnings records linked to higher education graduation files. Prior work using this data to study returns to higher education in Chile includes Hastings et al. (2013), González-Velosa et al. (2015), and Bordon and Fu (2015). Graduates employed in the public sector, the self-employed, and the informal sector fall outside the SII universe and are not observed.

For each program–graduation year cell, we observe the number of graduates matched to formal employment and their mean annual taxable earnings one, two, and three years after graduation. The three-horizon structure is a design feature. The signaling model predicts that the information advantage of the accreditation label should decay as employers accumulate direct experience with graduates, so measuring wages at multiple post-graduation horizons allows us to trace that learning curve.

3.2 Accreditation Timing

The CNA maintains a public registry that records, for each program, the exact date of the Commission's ruling, the binary accreditation outcome, and the duration of the term awarded. For each accredited program, we record the year of first CNA accreditation as the treatment date G_p and construct event time $\kappa_{pt} = t - G_p$ to position each graduation cohort relative to the moment the signal became publicly verifiable. First-time accreditations are distributed across all years from 2007 to 2018,

providing the cross-cohort staggering that the imputation estimator exploits.

The registry also records the length of each accreditation term, ranging from two to seven years. We use modal institutional term length to construct the three-tier quality proxy (Baseline: two to three years; Enhanced: four to five years; Top-tier: six to seven years) that drives the heterogeneity analysis in Section 5. Programs' first-time term length serves a parallel role at the program level, measuring the intensity of the quality signal the Commission sent at the moment of first accreditation and allowing us to test whether programs that received longer initial terms generate correspondingly larger wage premiums for their graduates.

3.3 Student Composition

Any effects on wages estimated in this paper could, in principle, reflect two distinct mechanisms. Employers may be updating their beliefs about program quality in response to the accreditation signal, or high-ability students may be selecting into newly accredited programs and mechanically raising average wages through their own human capital. Separating these channels requires data on who attends each program in each cohort. That information comes from DEMRE, Chile's centralized admissions authority, which assigns students to programs through a deferred acceptance algorithm based on standardized test scores (*Prueba de Selección Universitaria*, PSU) and secondary school GPA.

When students register to sit the PSU, they report family income, parental education, and prior academic performance. Because registration is universal among applicants to the centralized system, these records are available at the program–admission cohort level and can be linked to graduation cohorts. The composition test in Section **??** uses this link to ask whether accreditation shifts the socioeconomic profile of students graduating from treated programs. A change in graduate composition beginning at $\kappa \geq 1$ would be consistent with differential enrollment occurring roughly a degree-length earlier, as documented by Molina and Valdebenito (2026) using the enrollment data directly. Because socioeconomic characteristics are measured at registration rather than at graduation, the test assumes that pre-enrollment SES remains informative about students' backgrounds by the time they enter the labor market, a reasonable assumption given that parental education and family income at age 18 are

largely determined before university begins.

3.4 Sample Construction

The full wage panel contains 25,054 program–graduation year observations from 3,429 programs. Two restrictions define the eligible sample. First, we exclude medicine, dentistry, and education, where accreditation is mandatory and tied to professional licensing rather than voluntary quality signaling, as discussed in Section 2. Second, we restrict to programs offered by institutions that held institutional accreditation at some point during the sample period. Employers screening graduates from unaccredited institutions face a different inference problem: the absence of an institutional track record confounds the program-level signal, and students at unaccredited institutions are ineligible for public higher education funding, introducing a selection channel qualitatively distinct from the sorting observed within the accredited sector.

The treatment window covers first accreditations between 2007 and 2018. The lower bound reflects data availability, as 2007 is the first year for which graduation cohorts can be reliably linked to wage outcomes. The upper bound reflects an institutional discontinuity: undergraduate program-level accreditation was suspended in 2018, making it the last year in which a first accreditation can be interpreted as a clean voluntary information shock. Programs with no accreditation event within this window form the never-treated control group. These restrictions yield the estimation sample of 18,826 observations from 2,572 programs: 1,725 treated and 847 never-treated.

3.5 Descriptive Statistics

Table 1 compares pre-accreditation cohort characteristics across treated and control programs, restricting to observations before each program's first accreditation event ($\kappa < 0$). The comparison covers 1,193 of the 1,725 treated programs in the estimation sample. The remaining 532 programs received first-time accreditation in 2007 and have no pre-treatment wage observations in the panel. All 847 never-treated programs are included.

Panel A reveals that graduates of control programs earn approximately 25 percent more per year than graduates of treated programs (CLP 7.5 million versus CLP 6.0 million), while being matched

to formal employment at slightly lower rates (79.8 percent versus 82.1 percent). The wage gap does not reflect a student quality advantage in the control group. Panel B shows that treated programs enter the sample with modestly stronger student bodies. Their graduates score higher on the PSU and on secondary school GPA, and come from slightly higher-SES households, with a larger share of college-educated parents (45.1 percent versus 40.2 percent). The wage gap instead reflects institutional composition. Panel C shows that control programs are disproportionately drawn from Top-tier institutions (16.0 percent versus 10.3 percent), programs at well-established universities whose graduates carry a recognizable brand that substitutes for accreditation in employer priors. Treated programs concentrate in Enhanced institutions (51.6 percent), universities with solid but nationally unrecognized reputations, for whom the CNA stamp closes a genuine information gap. They are also larger (32 versus 24 graduates per cohort) and graduate a substantially higher share of women (59 percent versus 46 percent), reflecting which sectors pursued voluntary accreditation early.

These baseline differences are not a threat to identification. They motivate it. The event-study design absorbs all time-invariant program differences through unit fixed effects and identifies the accreditation effect from within-program wage changes around the treatment event. The descriptive table shows that cross-sectional comparisons between treated and control programs would be confounded by the very reputational differences that the signaling model places at the center of the analysis. The within-program design sidesteps that confound, and the composition data in Panel B provide the baseline against which post-accreditation sorting can be assessed.

4 Empirical Strategy

4.1 Identification and Research Design

The central identification challenge is not establishing that accreditation happened but establishing that its timing was unanticipated. If programs knew years in advance exactly when they would receive their first accreditation, they could invest strategically during the pre-accreditation period, and any post-accreditation wage gains would reinforce the quality signal from earlier quality investments. The staggered difference-in-differences design we use below is built precisely to handle this timing

uncertainty.

We estimate the causal effect of first-time program accreditation on graduate wages using a staggered difference-in-differences design. The treatment is defined as a binary indicator for the year a program receives its initial accreditation from the National Accreditation Commission (CNA). Our primary outcomes are the log annual earnings of graduates measured one, two, and three years after labor market entry.

The signaling mechanism implies that the relevant treatment timing depends on the graduation cohort relative to the public announcement. We define event time κ_{pt} as the difference between graduation year t and the year of first accreditation G_p:

$$\kappa_{pt} = t - G_p \tag{1}$$

where $G_p = \infty$ for never-treated programs. This specification captures the information available to employers at the time of hiring. Because firms observe the accreditation status of the degree-granting program, the wage effects should accrue to cohorts graduating on or after the date the accreditation is publicly awarded ($\kappa \geq 0$). All graduates in the early post-treatment window ($\kappa \in \{1, 2, 3, 4\}$) were enrolled before accreditation was announced, so any wage gains they experience can only reflect employer updating on the new signal, not the compositional effect of students enrolling in the newly accredited program. This feature of the design isolates the pure signaling channel before composition effects can appear.

4.2 Estimation and Inference

The estimation sample is defined in Section 3. The remaining challenge is estimating the effect of accreditation without importing the biases that contaminate the standard two-way fixed effects estimator in staggered settings. When treatment effects are heterogeneous across cohorts, as the signaling model predicts, the traditional estimator uses early-treated units as implicit controls for later-treated units, producing estimates that can be negatively weighted and misleading (Callaway and Sant'Anna, 2021; Goodman-Bacon, 2021; Sun and Abraham, 2021). We avoid this entirely by using the imputation estimator proposed by Borusyak et al. (2024).

The imputation approach proceeds in two steps. We first estimate a unit-and-period fixed effects model on the sample of never-treated and not-yet-treated observations:

$$Y_{pt} = \alpha_p + \lambda_t + \varepsilon_{pt} \tag{2}$$

We then use the estimated parameters to impute counterfactual outcomes $\hat{Y}_{pt}^{(0)}$ for treated observations. The average treatment effect on the treated (ATT) for each event-time κ is the mean difference between observed and imputed outcomes:

$$\widehat{\text{ATT}}(\kappa) = \frac{1}{N_\kappa} \sum_{(p,t):t-G_p=\kappa} \left[Y_{pt} - \hat{Y}_{pt}^{(0)} \right] \tag{3}$$

This estimator is valid under arbitrary treatment effect heterogeneity across cohorts and over time. We cluster standard errors at the program level to account for serial correlation in wages within programs.

The event study profile traces the ATT at each relative period, but the headline estimate reported in Table 2 is a scalar ATT that pools across all cohorts and post-treatment event times. Following Borusyak et al. (2024), this scalar is a weighted average of the cohort-by-event-time specific estimates $\widehat{\text{ATT}}(\kappa)$, where the weights are proportional to the number of treated observations contributing to each cohort-by-event-time estimate. Cohorts with more graduates at a given event time receive more weight, so the scalar ATT reflects where most of the identifying variation lies. The scalar ATT summarizes the average wage premium across all post-accreditation cohorts, while the event study profile reveals whether that premium accrues immediately or builds over time. This distinction matters for distinguishing employer learning from a one-time signaling shock.

4.3 Parallel Trends

Parallel trends requires that, absent accreditation, the wage trajectories of eventually-treated and control programs would have evolved identically over time. In this setting, the assumption is plausible because the timing of first accreditation is not driven by pre-existing wage trends. Programs self-select to apply, but the final decision (including whether accreditation is granted and for how long) is determined by an independent CNA commission after a peer review process whose outcome programs

cannot fully anticipate.

We assess this empirically by examining the pre-treatment coefficients for $\kappa \in \{-4, -3, -2\}$, normalized to $\kappa = -1$, and test their joint significance with an F-test. Flat pre-trends and failure to reject the null provide the main empirical support for the assumption.

4.4 No Anticipation

The design requires that neither programs nor employers adjusted wage-relevant behavior before the accreditation announcement. Programs may have prepared for the CNA review in other ways, but if those preparations did not translate into changes that employers could observe and price into wages, the pre-announcement wage trajectory remains a valid counterfactual.

On the program side, no anticipation does not require that programs be passive before the accreditation decision. Molina and Valdebenito (2026) documents that programs increase graduation rates in the period preceding the CNA review, pushing near-completers to finish on time because graduation rates are an input to the commission's evaluation. However, this behavior is administrative rather than a quality investment. Programs are not upgrading infrastructure, faculty, or curriculum in ways that employers could independently observe and price into wages. Without visible quality improvements, employers have no basis for adjusting wages before the formal CNA ruling is public.

On the employer side, the concern is subtler. CNA peer visits bring in evaluators from academia and industry who observe the program before the commission rules. In principle, these evaluators could start adjusting wages for the program's graduates before the ruling is public. Two features of the Chilean labor market rule this out. Wages for a program's graduates reflect market-wide beliefs, and no single employer can unilaterally move them. And evaluators participate in a regulatory process, not a recruiting one, giving them no systematic incentive to favor specific programs in their own hiring.

5 Results

5.1 Average Effect of Accreditation on Graduate Wages

Table 2 reports the scalar average treatment effect on the treated (ATT) at each wage horizon. The estimated premium is 3.0 percent for wages measured one year after graduation, 2.9 percent for wages measured two years after graduation, and 2.4 percent for wages measured three years after graduation. The first-year estimate is imprecise and not statistically distinguishable from zero. The year-two and year-three estimates are statistically significant at the 10 percent level. In absolute terms, the year-two premium corresponds to roughly CLP 234,000 in additional annual earnings against a control mean of CLP 8.1 million, and the year-three premium corresponds to roughly CLP 235,000 against a control mean of CLP 9.8 million. The modest decline from year one to year three is consistent with employer learning: as hiring managers accumulate direct experience with graduates from newly accredited programs, the informational value of accreditation fades.

The event study in Figure 1 shows the full dynamic profile. Pre-period coefficients at $\kappa \in \{-4, -3, -2\}$ are statistically indistinguishable from zero and jointly insignificant ($p > 0.60$), supporting the parallel trends assumption. For cohorts graduating after accreditation, the wage profile diverges and stays above the pre-trend baseline. In the early post-accreditation window, individual effects at $\kappa \in \{1, 2, 3, 4\}$ are close to zero. At longer horizons ($\kappa \in \{5, 6, 7, 8\}$), the estimates reach 6 to 15 percent, driven jointly by continued employer updating and by the arrival of graduates who enrolled after accreditation was public.

The two phases carry different interpretations. In the early window, graduates at $\kappa \in \{1, 2, 3, 4\}$ enrolled before accreditation was awarded and could not have been drawn to the program by its newly certified status. Any premium they receive reflects employers revising their beliefs in response to the new public signal. In the late window, that ongoing updating is amplified by the arrival of higher-ability cohorts who self-selected into the now-accredited program. Section 5.2 documents when and how the composition shift enters the picture. Section 5.3 shows that institutional gaming around the peer review does not contaminate the wage estimates.

5.2 Enrollment Timing and the Composition of Graduates

A wage increase after accreditation can have two distinct sources, distinguishable only if we can track graduate characteristics over time. Employers may update their beliefs about program quality and raise wage offers broadly, or accreditation may attract higher-ability students who raise average wages through their own human capital. Separating the two requires knowing who graduates from treated programs before and after accreditation.

The timing argument in Section 4 provides a first, indirect answer. Students who enroll because a program is accredited enter a multi-year degree and graduate no earlier than $\kappa = D$, where D is program duration. At $\kappa \in \{1, 2, 3, 4\}$, the composition of the graduate pool is entirely determined by enrollment decisions made before accreditation was public.[3] The composition channel cannot operate in this early window, which is why it provides the cleanest evidence of employer updating. At $\kappa \geq D$, the two channels operate together, and the rising event study profile at those horizons reflects both.

Table 2 and Figure 3 provide a direct test. The ATT for the log-ratio of high-ability graduates (graduates scoring in the top quartile of the PSU) rises by 0.121 log points after accreditation ($p < 0.01$). Treated programs attract more high-scoring students once the accreditation stamp is on the record. The timing of this shift, shown in Figure 3, is informative. Students who enrolled because of the accreditation signal only enter the labor market at $\kappa \geq D$, so the ability shift builds gradually and accounts for a growing share of the wage premium at longer horizons. The pattern is exactly what a signaling equilibrium predicts. High-ability students prefer certified programs because the credential amplifies their own labor market prospects, and the resulting enrollment shift appears only after accreditation is public and only in graduating cohorts several years later.

The income composition measure shows no significant change after receiving the certification (-0.047, $p > 0.10$). Accreditation shifts the ability composition of graduates but not their socioeconomic background, consistent with it functioning as a quality signal rather than a mechanism that alters who, by family background, attends the program.

[3] Figure 6 shows the full distribution of program duration in Chile. The average duration of a degree program is 9.4 semesters, with more than 70 percent of programs lasting between 8 and 11 semesters (four to five years). This confirms that $\kappa = 5$ is the first post-accreditation period at which students who enrolled after the accreditation announcement begin graduating.

5.3 Graduation Counts and Employment Rate

A separate concern is that programs may behave differently around the accreditation event in ways that alter who appears in the graduate pool. We estimate event studies for three supplementary outcomes: the log of total graduates per program–graduation year, the log of graduates found in private-sector formal employment records, and the share of graduates found in those records. Figure 2 presents all three event studies.

Total graduates. Treated programs produce fewer graduates than control programs in the years before accreditation. After the decision, total graduate counts rise. The pre-period pattern violates parallel trends for the graduation count outcome, and the violation points directly to a known mechanism. Programs preparing for a peer review face strong incentives to demonstrate completion rates, which encourages delayed completers to finish before the self-evaluation report is submitted or the peer evaluation visit conducted. Molina and Valdebenito (2026) documents this gaming behavior for the full population of accrediting programs. The anticipation effect operates before the formal decision, not after it.

Employed graduates. The count of graduates found in private-sector formal employment records also rises after accreditation, but less markedly than total graduate counts. This asymmetry is the key to interpreting the employment rate. Accreditation does not leave employment counts unchanged while graduates accumulate; both series increase. The distinction is between the rates at which both grow. The pool of completers expands faster than the pool of formal-sector workers, which is sufficient to pull the employment rate down even when employment counts are rising.

Employment rate. Before accreditation, treated programs exhibit higher formal employment rates than control programs. This positive pre-period selection reflects the same mechanism visible in the graduation counts. When total completions are depressed, the graduates who do finish tend to be the most able and find formal private-sector employment at above-average rates. After accreditation, formal employment rates fall toward or below control levels. This is not evidence that accreditation hurts labor market outcomes. The additional completers who graduate after receiving the certification

could also be more likely to enter self-employment, the informal sector, or public-sector positions that fall outside the SII payroll records, so the formal-employment count rises at a slower pace than total graduates. The wage estimates are unaffected, as they are constructed from the formal-sector sample. As marginal completers exit that sample, the wage analysis becomes more restricted to workers who would have found formal employment regardless of accreditation, making the early-window wage estimates, if anything, conservative.

The 3 percent pooled ATT is therefore a weighted average: a near-zero effect in the early window, where the composition of graduates has not yet shifted, and a growing premium at longer horizons, where higher-ability cohorts account for an increasing share of program graduates. The heterogeneity analysis examines whether this premium varies systematically with how much certification expands the pool of high-ability applicants.

6 Heterogeneity

6.1 Institutional Reputation and the Return to Accreditation

The signaling model predicts that the wage return to accreditation should be largest where employer uncertainty about program quality is greatest. Institutional reputation is a natural proxy for that prior uncertainty. Employers have accumulated information about programs at well-established universities through decades of direct hiring and word-of-mouth. Programs at less prominent institutions are harder to evaluate from reputation alone. If the CNA label resolves genuine uncertainty, its value should be highest at the institutions where employers know the least.

Table 3 tests this prediction by estimating subgroup ATTs stratified by the institution's CNA accreditation tier. The result is not a gradient from positive to zero; it is a reversal from positive to negative.

Graduates of programs at *Baseline* institutions realize a wage premium of 8.9 percent in their first year of employment ($p < 0.01$), falling to 4.6 percent in year two and 2.6 percent in year three. At the control mean of CLP 6.1 million for first-year wages, the 8.9 percent premium amounts to roughly CLP 540,000 per year. The certification converts a noisy prior into a verified quality floor, and the

labor market responds immediately. The effect remains statistically significant across all three years we observe, which suggests that employer uncertainty about these programs is deep enough that a single CNA evaluation does not fully resolve it.

For graduates of *Enhanced* institutions, the first-year premium is -2.7 percent ($p < 0.01$). At the control mean of CLP 6.0 million, this is a reduction of roughly CLP 160,000 per year. For graduates of *Top-tier* institutions, the first-year effect is -5.0 percent ($p < 0.01$), remaining negative through year two (-2.8 percent) and year three (-2.4 percent). At the control mean of CLP 7.1 million for first-year wages among top-tier graduates, the first-year penalty corresponds to roughly CLP 360,000 per year. The signal does not merely fade as institutional prestige rises; it flips sign.

The direction of the reversal is precisely what the signaling framework implies for this part of the quality distribution. At *Top-tier* institutions, employers already hold precise beliefs about graduate quality, formed through years of direct hiring and strong institutional reputation. When a program at such an institution receives its first CNA certification, the announcement carries little information that the market did not already have. But it may carry a different kind of information: a program at a prestigious institution that needed to go through the formal review process, and received a term shorter than its institutional peers, may prompt employers to ask why the program had not been certified earlier. Accreditation, for graduates of top-tier programs, is not a credential that opens doors; it raises the question of why a door was previously half-open. Figure 4 traces the diverging event study profiles: the *Baseline*-tier effect rises over the post-accreditation window, while the *Top-tier* effect remains negative throughout.

6.2 Program-Level Treatment Intensity

Institutional reputation captures the employer's prior about the institution, but the initial accreditation term awarded to the program conveys additional program-specific information. A short initial term certifies quality at the minimum standard. A long initial term certifies quality well above it. If the wage return to accreditation reflects the informational content of the certification, programs whose first accreditation resolves deeper uncertainty should generate larger wage responses.

Table 4 tests this prediction using the duration of the first accreditation term as the stratification

variable. Programs receiving a *Low*-intensity first accreditation (1–3 years) show a first-year wage premium of 10.7 percent ($p < 0.01$). At the control mean of CLP 6.0 million for first-year wages, this amounts to roughly CLP 640,000 in additional annual earnings. The premium remains statistically significant through the third year: 7.3 percent in year two and 5.4 percent in year three. The persistence of the effect across three years after graduation suggests that a short initial term, even after it is renewed or upgraded in subsequent evaluations, leaves a durable informational mark: employers learned from the first evaluation that this program was a borderline case, and the updated belief does not vanish in one hiring cycle.

Programs receiving a *Mid*-intensity first accreditation (4–5 years) show a first-year wage effect of -4.2 percent ($p < 0.01$). Programs receiving a *High*-intensity term (6–7 years) show no statistically significant change (-1.5 percent, $p > 0.10$). The gradient from Low to High is monotone and mirrors the institutional result: the CNA label generates the largest wage returns where the certification resolves the most uncertainty about quality, and it generates negative returns where the evaluation reveals something below what the employer's prior implied. Figure 5 presents the event study profiles for all three intensity groups.

The negative effect at *Mid* intensity is worth pausing on. A program that receives a four- to five-year initial term has demonstrated solid quality, but perhaps not at the level its institutional context led employers to expect. Before accreditation, employers may have assumed that the program, by remaining outside the CNA registry, was either very strong (and therefore needed no external validation) or very weak (and therefore failed the evaluation). A moderate first-time term rules out both extremes and may lower wage offers for graduates from programs where employers had maintained a high-quality prior.

6.3 Prior Uncertainty and the Value of the Certification Signal

The institutional and program-level results, taken together, provide a consistent account of how the labor market processes a new quality signal. The return to accreditation is not a fixed premium attached to the CNA label. It is a function of how much the label tells an employer about program quality beyond what they already knew.

Employers enter every hiring cycle with beliefs about the quality of each program's graduates. Those beliefs are more precise for well-established institutions and less precise for newer or regionally concentrated programs. The CNA certification is a public, standardized evaluation. Its effect on wages depends on the gap between the employer's prior and what the certification reveals.

Where that gap is large, as for *Baseline* institutions receiving a short initial term, the certification resolves substantial uncertainty and wages respond sharply upward. The year-one premium of 8.9 percent at *Baseline* institutions, combined with the year-one premium of 10.7 percent for *Low*-intensity accreditations, places the upper end of the informational return in the range of 9 to 11 percent for the group of programs where employer uncertainty was greatest before the CNA decision. This is a large effect. At the control mean, it represents CLP 540,000 to CLP 640,000 in additional annual earnings in the first year alone.

Where the gap is small or reversed, as for *Top-tier* institutions or *High*-intensity programs, the certification adds little and may invite unfavorable comparisons with other programs whose absence from the registry previously signaled either high quality or irrelevance to the formal accreditation system. The wage response at the top of the prior-precision distribution is negative and persistent.

This pattern follows directly from a model of Bayesian employer updating. Define employer prior precision as the inverse of the variance in beliefs about program quality. The informational value of the CNA signal is decreasing in that precision: the more the employer already knows, the less the signal changes their beliefs, and the smaller the wage response. The CNA label functions as an information transfer from a regulator who has conducted a direct evaluation to employers who must otherwise rely on indirect signals. Where the indirect signals are plentiful and credible, the regulatory evaluation adds little. Where they are sparse or noisy, the regulatory evaluation generates a wage premium proportional to the uncertainty it resolves.

At *Baseline* institutions, the two channels documented in Section 5 reinforce each other. The accreditation signal resolves the most uncertainty for employers, producing an immediate and large wage premium. The same signal also attracts higher-ability students who had previously avoided the program, and as those cohorts graduate, the premium grows further. At *Top-tier* institutions, neither channel operates positively: the signal adds no information to already-precise employer beliefs, and the self-selection of high-ability students into the program was already occurring through institutional

reputation alone. The gradient in wage returns across institutional tiers reflects the gradient in prior uncertainty, amplified at the bottom of the distribution by the composition effect and compressed at the top where that effect was already exhausted before accreditation.

7 Conclusion

We study what happens to graduate wages when a university program receives its first accreditation from Chile's National Accreditation Commission. The average treatment effect, pooled across all post-accreditation cohorts, is approximately 3 percent over the first three years of employment, with statistical significance in years two and three. That number understates both what the data show and what it means. The wage return to accreditation is not a uniform feature of the certification system. It is a direct function of how much the label tells employers about program quality beyond what they already knew.

Two dimensions of heterogeneity make this point precisely. Stratified by institutional accreditation tier, graduates of programs at *Baseline* institutions gain 8.9 percent in year-one wages after their program is first certified, an effect that persists, at 2.6 percent, three years later. Graduates of programs at *Top-tier* institutions lose 5.0 percent in year-one wages and the negative effect remains through year three. Stratified by the duration of the first accreditation term awarded to the program, graduates of *Low*-intensity programs (1–3 year first term) gain 10.7 percent in year one and 5.4 percent in year three. Graduates of *High*-intensity programs (6–7 year first term) show no statistically significant wage change. Both gradients point in the same direction: the return to accreditation is large and positive where employer uncertainty was high before the certification decision, and it is small or negative where employers already held precise beliefs about program quality.

This pattern is exactly what a Bayesian model of employer updating predicts. Employers enter each hiring cycle with prior beliefs about program quality, beliefs formed through institutional reputation, alumni networks, and accumulated experience. The CNA certification is a public, standardized signal. Where employer priors are noisy, as for programs at less prominent institutions receiving short initial terms, the signal resolves substantial uncertainty and wages adjust sharply. Where priors are precise, the signal adds little, and at the top of the quality distribution it may invite unfavorable comparisons:

a prestigious program that required external certification to validate what reputation had not already established may prompt employers to revise their beliefs downward relative to the unconditional prior. The CNA label is not inherently valuable. Its value is entirely in the information it resolves.

The causal interpretation rests on two design features. First, accreditation timing is staggered across programs in a way that the imputation estimator of Borusyak et al. (2024) exploits without importing the negative-weighting biases of standard two-way fixed effects. Second, the wage effects are concentrated in the early post-accreditation window ($\kappa \in \{1, 2, 3, 4\}$), when all graduates in the data enrolled before accreditation was announced. Compositional selection cannot explain the early-window premium: higher-ability students who chose the program because of its accredited status only graduate four to five years after enrollment, outside this window. The paper confirms this directly: the log-ratio of high-ability graduates rises significantly after accreditation ($p < 0.01$), but the timing places this shift outside the identifying window. Supplementary event studies for graduation counts, formal employment rates, and on-time completion rates further show that institutional gaming around the CNA review operates through anticipation effects that contaminate only the graduation-count pre-trend, leaving the wage pre-trends statistically flat.

These findings carry direct implications for how voluntary accreditation systems should be designed. Chile's CNA architecture gives programs the choice of whether to seek certification, and the regulatory agency has limited capacity to pursue programs that do not apply. This creates a predictable misallocation. The programs for which accreditation generates the largest wage returns to graduates, those at *Baseline* institutions with the weakest employer priors, tend to have the least organizational capacity to prepare for an external review and, historically, the weakest incentive to pursue a credential that may be less legible to their local employer networks. A voluntary system without targeted support will concentrate certification among programs where the informational gap is smallest. Policymakers who want accreditation to function as an information subsidy for graduates who cannot rely on institutional prestige alone must actively encourage or subsidize lower-tier programs to seek evaluation. The distributional benefit that our heterogeneity results identify is the largest return the system generates. Leaving it to self-selection forfeits it.

Two questions that this paper does not answer remain open. First, we measure wages in the first three years after graduation, a window where the early-window signaling effect dominates. Whether

the wage gap between treated and control graduates narrows, stabilizes, or widens at longer horizons, as enrollment composition shifts and as employers accumulate direct experience with newly accredited programs, determines whether the return to accreditation is a temporary information premium or a more durable labor market advantage. Second, the paper identifies the signaling channel and rules out the composition channel for the early window, but it does not test whether the accreditation process itself induces quality-improving investments in faculty, curriculum, or infrastructure that eventually show up in wages beyond what the label alone explains. Separating that channel from pure employer updating at longer horizons requires tracking both graduate characteristics and institutional inputs over the full post-accreditation period. The data exist to answer both questions; doing so would complete the picture of how accreditation functions as a quality assurance institution and not just as a signal.

Table 1: Descriptive Statistics: Pre-Accreditation Cohorts

	Treated	Control	Difference
Panel A: Labor Market Outcomes			
Graduation cohort size	32.395	23.659	8.736***
	(37.949)	(38.795)	(0.826)
Employment rate, year 1	0.821	0.798	0.023***
	(0.233)	(0.245)	(0.005)
Annual earnings, yr. 1 (CLP)	5959.493	7459.020	-1499.526***
	(3613.547)	(4572.991)	(89.660)
Panel B: Student Characteristics			
PSU score (math-verbal)	520.714	516.325	4.389**
	(89.589)	(72.766)	(1.832)
GPA score	582.142	559.617	22.525***
	(66.470)	(61.832)	(1.448)
Male	0.408	0.539	-0.131***
	(0.320)	(0.310)	(0.007)
High income	0.203	0.180	0.024***
	(0.263)	(0.223)	(0.005)
College-educated parents	0.451	0.402	0.049***
	(0.288)	(0.264)	(0.006)
Panel C: Institutional Quality			
Baseline Inst.	0.355	0.347	0.009
	(0.479)	(0.476)	(0.010)
Enhanced Inst.	0.516	0.420	0.095***
	(0.500)	(0.494)	(0.011)
Top-tier Inst.	0.103	0.160	-0.056***
	(0.304)	(0.366)	(0.007)
N (Programs)	1193	847	

Note: This table displays the means of program-level characteristics for the Control (programs never accredited during the sample period) and Treated (eventually-accredited programs, observed in pre-accreditation cohorts) groups. Standard deviations are reported in parentheses beneath the means. The unit of observation is the academic program. The "Difference" column reports the coefficient from an OLS regression of the variable on a treatment status indicator. The N (Programs) row counts programs with at least one pre-accreditation wage observation. Institution quality tiers are defined by CNA accreditation duration: Baseline (2–3 years), Enhanced (4–5 years), Top-tier (6–7 years). Significance levels are denoted as follows: * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

Table 2: Effect of First-Time Accreditation on Graduate Wages and Graduate Composition

	Labor Market Outcomes (logs)			Grads composition	
	Wage yr 1	Wage yr 2	Wage yr 3	Ability	Income
Accreditation effect	0.030	0.029*	0.024*	0.121***	-0.047
	(0.022)	(0.017)	(0.014)	(0.046)	(0.042)
Num. Obs	18672	18707	18708	13443	13217
Control mean	6012.6	8078.4	9806.1	0.4	0.2

Note: Standard errors (clustered at the program level) are in parentheses. The control outcome mean for wage columns is the average annual wage among never-accredited programs, reported in thousands of CLP. Composition outcomes are log-ratios: $\log(s/(1-s))$ where s is the program-cohort share of high-PSU graduates (top quartile, year-specific cutoff) or high-income graduates. Composition outcomes are available only for graduates matched to DEMRE centralized admissions records. All programs are offered by accredited institutions. The treatment group includes programs accredited for the first time. Average treatment effects on the treated follow Borusyak et al. (2024). * p<0.1, ** p<0.05, *** p<0.01.

Table 3: Heterogeneous Effects by Institutional Accreditation Quality

	Log wage yr 1	Log wage yr 2	Log wage yr 3
Accreditation × Baseline inst.	0.089***	0.046***	0.026***
	(0.017)	(0.011)	(0.009)
Accreditation × Enhanced inst.	-0.027***	-0.003	0.008
	(0.010)	(0.009)	(0.008)
Accreditation × Top-tier inst.	-0.050***	-0.028***	-0.024***
	(0.011)	(0.010)	(0.008)
Num. Obs	18672	18707	18708
Control Mean (CLP thousand)			
Baseline inst.	6057.3	7936.2	9592.7
Enhanced inst.	5996.4	8124.3	9966.6
Top-tier inst.	7135.4	9316.7	11080.8

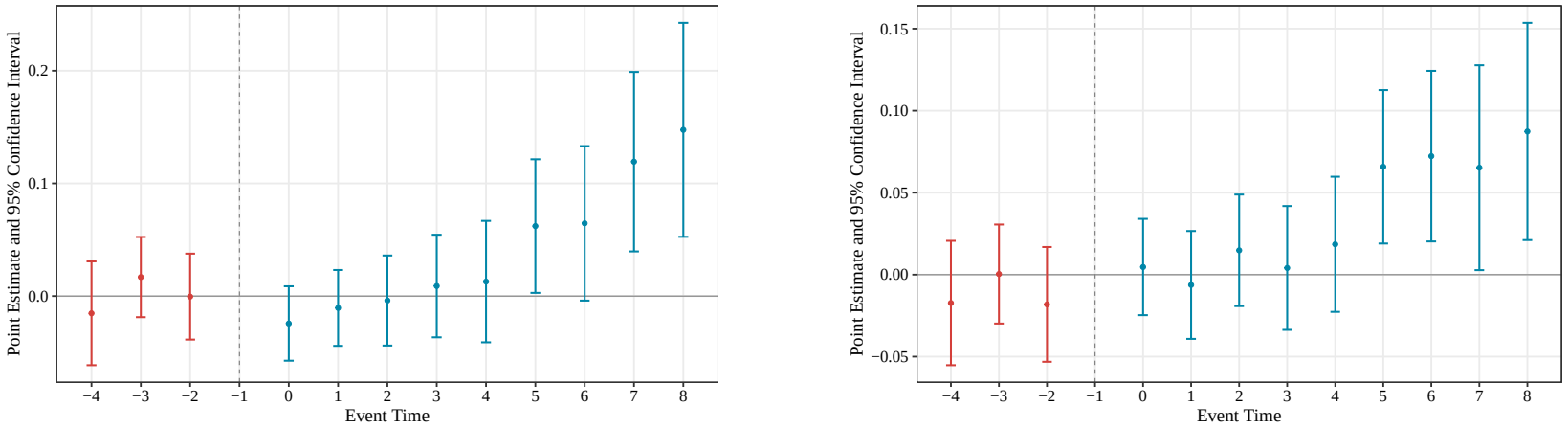
Note: This table displays heterogeneous effects of first-time accreditation on graduate wages by institutional quality. Standard errors (clustered at the program level) are in parentheses. The unit of observation is the academic program. Each row reports the average treatment effect weighted by the share of treated observations within that institutional tier, estimated following Borusyak et al. (2024) with program and graduation-year fixed effects. Institutional quality is defined by the modal CNA accreditation duration of the host institution: Baseline (1–3 years), Enhanced (4–5 years), Top-tier (6–7 years). Control means are in CLP thousands and correspond to never-accredited programs within each institutional tier. Significance levels are denoted as follows: * p<0.1, ** p<0.05, *** p<0.01.

Table 4: Treatment Intensity: Effects by Years Awarded at First Accreditation

	Log wage yr 1	Log wage yr 2	Log wage yr 3
Treatment Intensity			
Accreditation × Low (1-3 years)	0.107***	0.073***	0.054***
	(0.026)	(0.018)	(0.014)
Accreditation × Mid (4-5 years)	-0.042***	-0.014	-0.008
	(0.012)	(0.011)	(0.009)
Accreditation × High (6-7 years)	-0.015	-0.024**	-0.012
	(0.014)	(0.012)	(0.012)
Num. Obs	18672	18707	18708
Control mean (thousands CLP)	6012.6	8078.4	9806.1

Note: This table displays the effect of first-time accreditation on graduate wages by treatment intensity. Standard errors (clustered at the program level) are in parentheses. The unit of observation is the academic program. Treatment intensity is defined by the number of years awarded at first accreditation: Low (2–3 years), Mid (4–5 years), High (6–7 years). Each row reports the average treatment effect weighted by the share of treated observations within that intensity group, estimated following Borusyak et al. (2024) with program and graduation-year fixed effects. The control mean is the average wage among never-accredited programs, reported in CLP thousands. Significance levels are denoted as follows: * p<0.1, ** p<0.05, *** p<0.01.

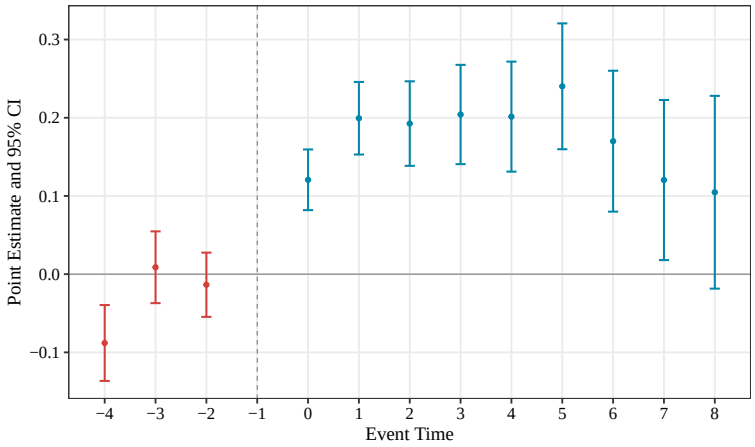
Figure 1: Event study estimates: Wages of employed grads



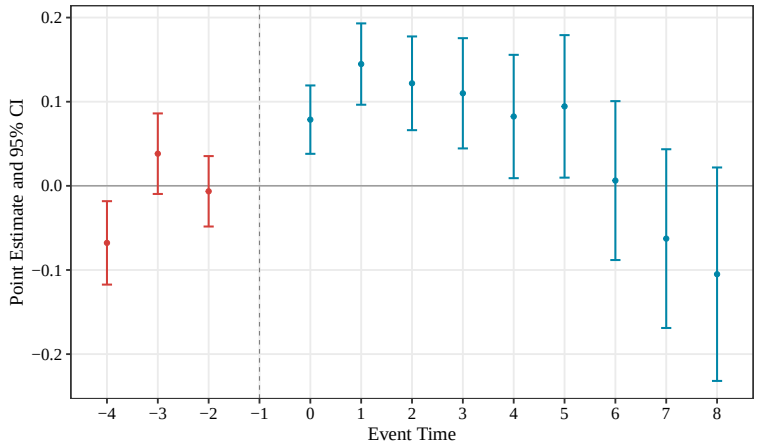
(a) $1^{st}Year$ (b) $2^{nd}Year$

Note: This figure plots event study estimates of the effect of first-time accreditation on log wages of graduates (conditional on being employed): (a) 1 year, and (b) 2 years after graduation. $\kappa = -1$ is the reference period. Vertical bars represent 95% confidence intervals. Estimates follow Borusyak et al. (2024) with standard errors clustered at the program level.

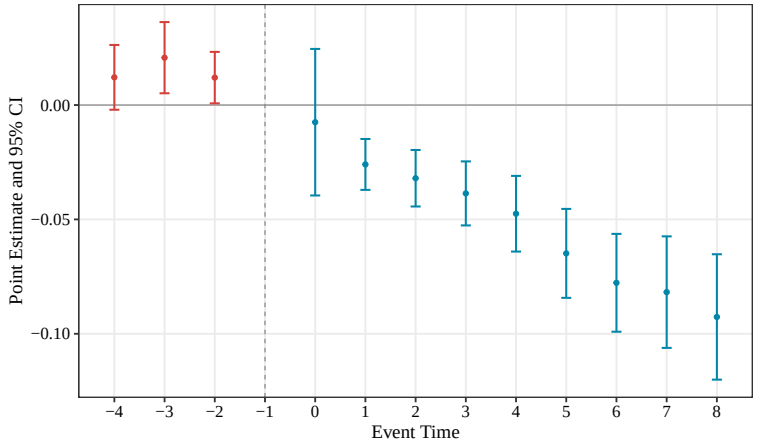
Figure 2: Event study estimates: Mechanism outcomes



(a) Graduation cohort size



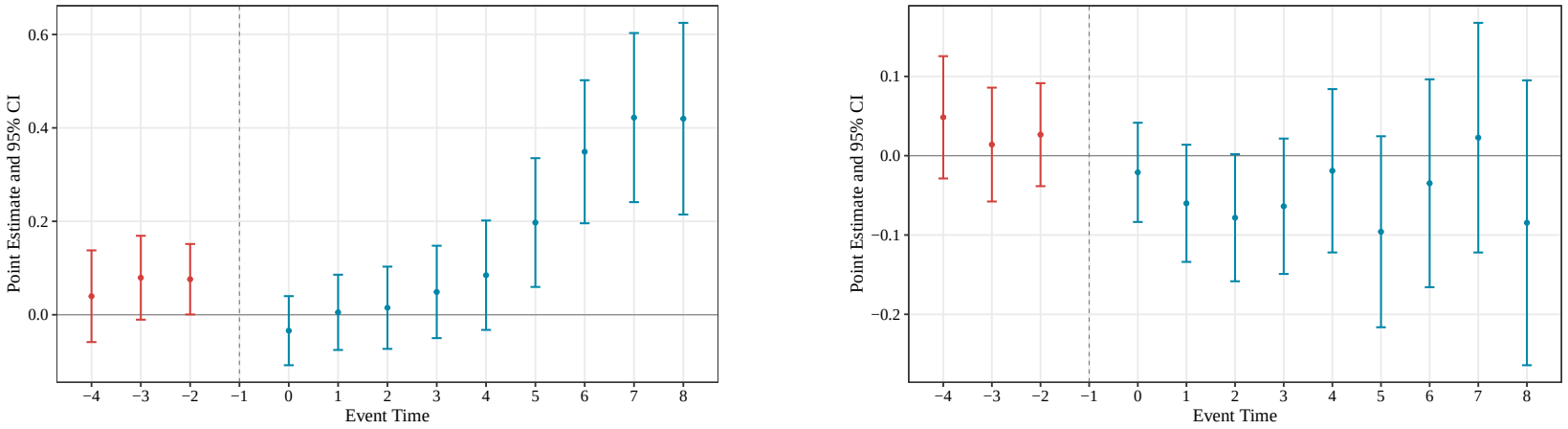
(b) Employed cohort size



(c) Employment rate

Note: This figure plots event study estimates of the effect of first-time accreditation on three mechanism outcomes: (a) log graduation cohort size (graduates per program–graduation year), (b) log employed cohort size (graduates matched to private-sector earnings records), and (c) formal employment rate two years after graduation. These are descriptive estimates as pre-trend tests reject parallel trends for all three outcomes, consistent with anticipatory increases in on-time graduation rates documented in Molina and Valdebenito (2026). $\kappa = -1$ is the reference period. Vertical bars represent 95% confidence intervals. Estimates follow Borusyak et al. (2024) with standard errors clustered at the program level.

Figure 3: Event study estimates: Graduate composition following accreditation

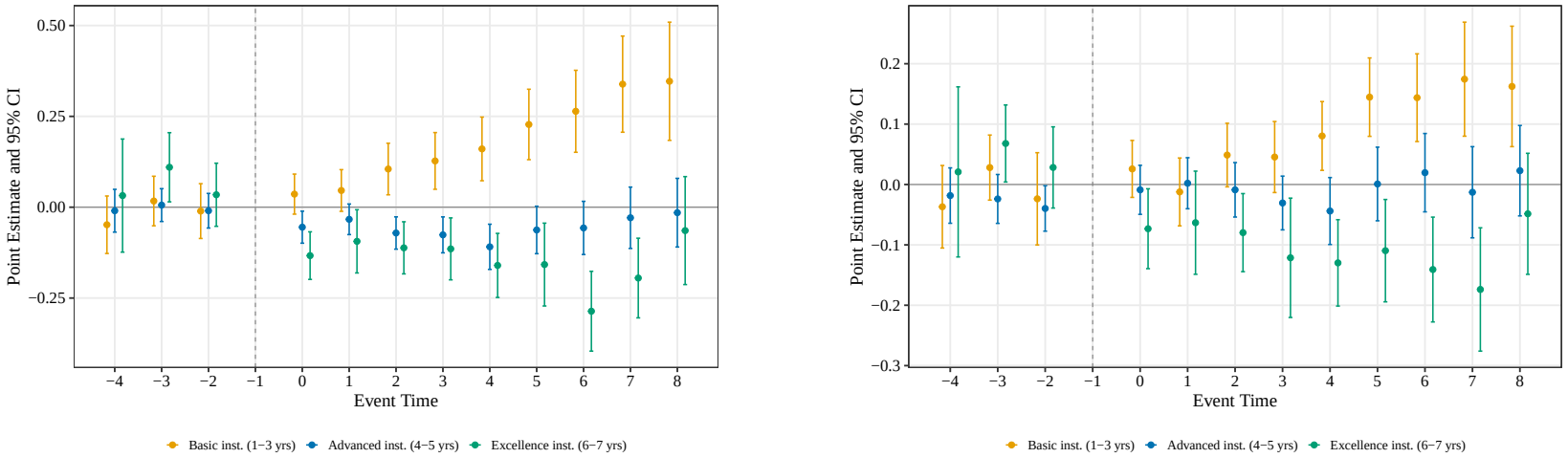


(a) Academic ability composition

(b) Socioeconomic composition

Note: This figure plots event study estimates of the effect of first-time accreditation on the log-ratio of high- to low-group graduates within each program–graduation year. Panel (a) defines high ability as scoring at or above the year-specific 75th percentile of the distribution of average DEMRE verbal and math scores. Panel (b) defines high socioeconomic status as belonging to a household with an above-median household income. A positive estimate indicates that the graduating cohort tilts toward higher-ability or higher-income graduates relative to the counterfactual. $\kappa = -1$ is the reference period. Vertical bars represent 95% confidence intervals. Estimates follow Borusyak et al. (2024) with standard errors clustered at the program level.

Figure 4: Event study estimates: Wages of employed grads by institutional tier

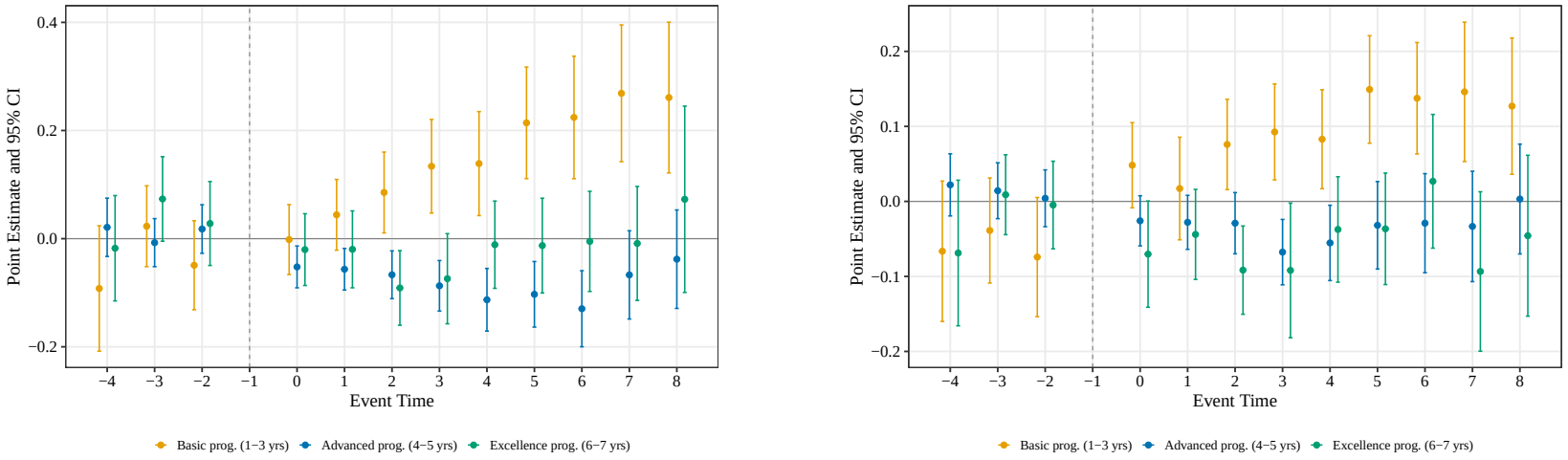


(a) $1^{st}Year$

(b) $2^{nd}Year$

Note: This figure plots event study estimates of the effect of first-time accreditation on log wages of graduates by institutional quality tier. Panels (a) and (b) show estimates for wages after 1 and 2 years after graduation. Colors indicate institutional quality tiers: baseline (yellow), enhanced (blue), top tier (green). $\kappa = -1$ is the reference period. Vertical bars represent 95% confidence intervals. Estimates follow Borusyak et al. (2024) with standard errors clustered at the program level.

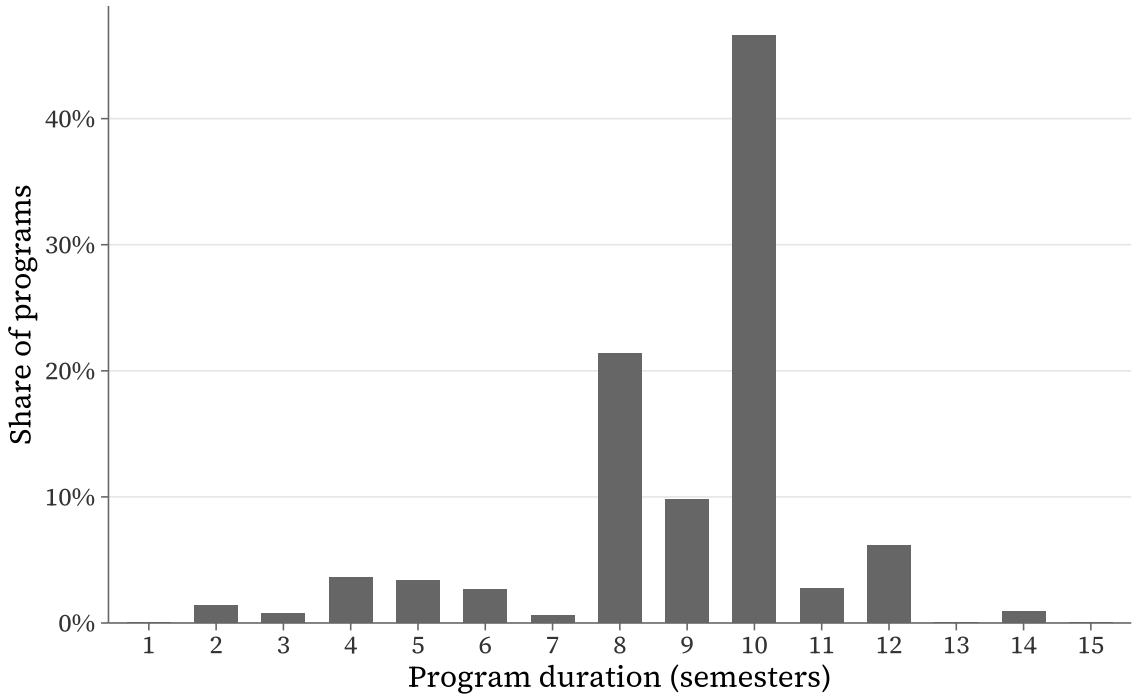
Figure 5: Event study estimates: Wages of employed grads, by treatment intensity



(a) $1^{st}Year$ (b) $2^{nd}Year$

Note: This figure plots event study estimates of the effect of first-time accreditation on log wages of graduates by program treatment intensity. Panels (a) and (b) show estimates for wages after 1 and 2 years after graduation. Colors indicate first-accreditation intensity tier: low (yellow), mid (blue), high (green). $\kappa = -1$ is the reference period. Vertical bars represent 95% confidence intervals. Estimates follow Borusyak et al. (2024) with standard errors clustered at the program level.

Figure 6: Distribution of programs duration



Note: This figure plots the full distribution of programs' duration in the Chilean higher education system. more than 70% of programs have a duration of between 4 and 5 years.

References

Akerlof, George A., "The market for "lemons": Quality uncertainty and the market mechanism," *The Quarterly Journal of Economics*, 1970, *84* (3), 488–500.

Altonji, Joseph G. and Charles R. Pierret, "Employer Learning and Statistical Discrimination," *Quarterly Journal of Economics*, 2001, *116* (1), 313–350.

Barroilhet, Agustín, "Problemas estructurales de la acreditación de la educación superior en Chile: 2006-2012," *Revista Pedagogía Universitaria y Didáctica del Derecho*, June 2019, *6* (1), 43–78.

_ , Rocío Ortiz, Bernardo F. Quiroga, and Mónica Silva, "Exploring conflict of interest in university accreditation in Chile," *Higher Education Policy*, June 2022, *35* (2), 479–497.

Bordon, Paola and Chao Fu, "College-major choice to college-then-major choice," *The Review of Economic Studies*, October 2015, *82* (4), 1247–1288.

Borusyak, Kirill, Xavier Jaravel, and Jann Spiess, "Revisiting event study designs: Robust and efficient estimation," *Review of Economic Studies*, January 2024.

Callaway, Brantly and Pedro H.C. Sant'Anna, "Difference-in-Differences with multiple time periods," *Journal of Econometrics*, 2021, *225* (2), 200–230.

Dale, Stacy Berg and Alan B. Krueger, "Estimating the Payoff to Attending a More Selective College: An Application of Selection on Observables and Unobservables," *Quarterly Journal of Economics*, 2002, *117* (4), 1491–1527.

Dranove, David and Ginger Zhe Jin, "Quality disclosure and certification: Theory and practice," *Journal of Economic Literature*, December 2010, *48* (4), 935–963.

_ , Daniel Kessler, Mark McClellan, and Mark Satterthwaite, "Is more information better? The effects of "report cards" on health care providers," *Journal of Political Economy*, 2003, *111* (3), 555–588.

Elfenbein, Daniel W., Raymond Fisman, and Brian McManus, "Market structure, reputation, and the value of quality certification," *American Economic Journal: Microeconomics*, November 2015, *7* (4), 83–108.

Farber, Henry S. and Robert Gibbons, "Learning and Wage Dynamics*," *The Quarterly Journal of Economics*, November 1996, *111* (4), 1007–1047.

González-Velosa, Carolina, Graciana Rucci, Miguel Sarzosa, and Sergio Urzúa, "Returns to Higher Education in Chile and Colombia," 2015.

Goodman-Bacon, Andrew, "Difference-in-differences with variation in treatment timing," *Journal of Econometrics*, December 2021, *225* (2), 254–277.

Hastings, Justine S., Christopher A. Neilson, and Seth D. Zimmerman, "Are some degrees worth more than others? Evidence from college admission cutoffs in Chile," July 2013.

Hoekstra, Mark, "The Effect of Attending the Flagship State University on Earnings: A Discontinuity-Based Approach," *Review of Economics and Statistics*, 2009, *91* (4), 717–724.

Jin, Ginger Zhe, "Competition and disclosure incentives: An empirical study of HMOs," *The RAND Journal of Economics*, 2005, *36* (1), 93–112.

Lange, Fabian, "The Speed of Employer Learning," *Journal of Labor Economics*, 2007, *25* (1), 1–35.

Lizzeri, Alessandro, "Information Revelation and Certification Intermediaries," *The RAND Journal of Economics*, 1999, *30* (2), 214–231.

MacLeod, W. Bentley and Miguel Urquiola, "Reputation and school competition," *American Economic Review*, November 2015, *105* (11), 3471–3488.

_ , Evan Riehl, Juan E. Saavedra, and Miguel Urquiola, "The big sort: College reputation and labor market outcomes," *American Economic Journal: Applied Economics*, July 2017, *9* (3), 223–261.

Molina, Cristhian and Rocıo Valdebenito, "Signaling Quality or Gaming the System? Evidence from College-Major Accreditation," 2026.

Reyes, Loreto, Jorge Rodríguez, and Sergio Urzúa, "Heterogeneous Economic Returns to Postsecondary Degrees: Evidence from Chile," *Journal of Human Resources*, 2016, *51* (2), 416–460.

Spence, Michael, "Job market signaling," *The Quarterly Journal of Economics*, 1973, *87* (3), 355–374.

Sun, Liyang and Sarah Abraham, "Estimating dynamic treatment effects in event studies with heterogeneous treatment effects," *Journal of Econometrics*, December 2021, *225* (2), 175–199.