Better late than never? Wage effects of delayed baccalaureate graduation in the United States

March 18, 2021

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Abstract

Only 42 percent of students earning baccalaureate degrees in the United States graduate within four years, compared to 53 percent three decades ago. Despite this shift, and plenty of concern over potential harm to students, we know little about whether delayed graduation carries a labor market penalty. Researchers examining time to degree using cross-sectional data report a negative relationship between time to degree and earnings. These findings likely tell us more about ability than about whether delayed graduation imposes a separate cost. Lost earnings from delayed entry into the labor market may be easily countered by higher earnings during school for those taking longer to finish. When we address the endogeneity that arises for student ability and time to degree, we find no evidence of a labor market penalty for delayed graduation. Together, these findings suggest that taking longer to complete college is not necessarily a problem that needs fixing.

JEL Classification: I21, I28, J24

Key Words: human capital, graduation delay, time to degree

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1. Introduction

Most college graduates in the United States spend more than four years earning a baccalaureate degree, a phenomenon predominantly observed in public institutions (Bowen, Chingos, and McPherson, 2009). In the 1970s, 53 percent of college graduates earned degrees within four years. Twenty years later, only 39 percent did so. For non-top 50 public universities, the decline was steeper—from 50 percent in the 1970s to 29 percent in the 1990s. Researchers posit the trend cannot be explained by changes in student preparedness or composition, and instead find that decreased resources for students, paired with increased employment during school, are likely causes (Bound, Lovenheim, and Turner, 2012). Not surprisingly, substantial increases in time to degree have drawn alarm from some researchers, policymakers, and media outlets.

Numerous policies seek to shorten time to degree by raising the cost of delayed graduation. Proposals include increased penalties for withdrawing from courses, credit-hour pricing penalizing students taking fewer than 15 credits per semester, and the endorsement of lockstep programs restricting student choice in courses, ultimately making it more difficult to change majors.[2] In 2016, the Obama administration proposed two significant changes to the federal Pell Grant program. The first provision would have provided approximately 700,000 students "making real progress toward on-time graduation" with an additional $1,915 on average to help pay for college and complete their degrees faster. The second provision, dubbed the "on-

[2]See, for example, the "15 to Finish" policy promoted by Complete College America and other nonprofits. Online at http://completecollege.org/docs/GPS_Summary_FINAL.pdf, accessed 13 November 2016. As of 2013, five statewide higher education systems and at institutions in fifteen states had adopted 15 to Finish. This information is online at http://www.completecollege.org/news.html, accessed 13 November 2016.

track Pell bonus," would have raised the maximum Pell award by $300 for approximately 2.3 million students taking 15 credits per semester in an academic year, a policy meant to encourage the receipt of a bachelor's degree within four years.[3,4] Critics contend that such policies overload students who find it necessary to work during college, or that enter higher education marginally prepared. There is evidence, for example, that excess credit hour penalties do not affect degree completion or time to degree, and instead increase student debt—a result that appears to be driven by first-generation and low-income students (Kramer, Holcomb, and Kelchen, 2018). Notwithstanding, support for such incentives remains broad. Backers cite high costs of delayed entry into the labor market, particularly how lengthened time to degree may encourage students to take on additional debt.[5] We consider the potentially optimizing behavior of students who work more during college and delay graduation, but do so with less debt.

Primarily, we investigate whether delayed graduates incur wage penalties beyond opportunity costs associated with taking longer to obtain a degree. We are interested in whether there is a penalty associated with overshooting the traditional norm of four years. Students working more during college have lower opportunity costs, and additional earnings during college may compensate for potential losses associated with delayed entry into the labor market. Students have always had the option of taking more credits and not changing majors, so it bears exploring whether going through college more slowly might be a good strategy, rather than a

[3] U.S. Department of Education, Fact Sheet: Helping More American Complete College: New Proposals for Success, released 19 January 2016, online at http://www.ed.gov/news/press-releases/fact-sheet-helping-more-americans-complete-college-new-proposals-success, accessed 13 November 2016.
[4]In 2012 alone, Idaho, Indiana, Mississippi, Missouri, and New Mexico passed legislation aimed at reducing time to degree at their public universities.
[5]Complete College America, https://completecollege.org/ (accessed 11 July 2019).

mistake. We approach this question in two steps. First, we examine a simple model of human capital accumulation to explore under what circumstances combining part-time work and a 5- or 6-year path to degree attainment is optimal. Second, we ask whether longer time to degree is penalized in the labor market. For the latter question, human capital holds that additional years of education increases worker productivity, thereby affording higher wages in the labor market (Becker, 1964). The human capital hypothesis predicts that students completing the same amount of credits over a longer period should experience no discernable effect on wages after controlling for work experience. However, according to the screening hypothesis (Spence, 1973), if time to degree serves as a productivity signal to employers, then those finishing sooner may be valued as being more productive in the labor market and should earn higher wages.

Several researchers report a negative association between earnings and time to degree, which they attribute to student ability (Brodaty, Gary-Bobo, and Prieto, 2008; Flores-Lagunes and Light, 2009; Aina and Pastore, 2012). This association alone does not rule out human capital, since the real test is whether workers with the same ability, but different time to degree, are compensated differently. We perform this test by controlling for ability and instrumenting for time to degree with the institutional average. Our instrument is plausible because institutional policies and norms surely affect a student's college trajectory but should have no bearing on labor market rewards apart from the institution's quality, which is also controlled for. To allay concerns that the instrument remains correlated with unobservable aspects of institution quality or student ability, for example, we relax the instrument's exclusion restriction. This method assumes a "plausibly exogenous" that violates strict exogeneity and provides analytical bounds of traditional point estimates.

Our findings suggest that concern over delayed graduation may be misplaced. Under reasonable assumptions regarding hours worked during college, the return to a college degree, and discount rates, students may come out ahead when they work while earning a degree in five or six years. When we address endogeneity in the early-career wage function, we find little evidence of a wage penalty for students taking longer to complete.

The remainder of the study proceeds as follows: section 2 presents a model of human capital that rationalizes a longer, non-traditional path to baccalaureate degree attainment; section 3 discusses relevant literature and motivates the identification strategy using a theoretical model of wages and graduation delay; section 4 introduces the data and discusses long-term trends in time to degree in the United States; section 5 details the empirical strategy; section 6 presents model results; section 7 concludes with policy implications and a suggestion for future research.

2. A simple model of human capital

We appeal to a simplified, discrete multi-stage human capital investment problem similar to Turner (2004). The framework is modified to examine the circumstances under which students rationally prefer a mixture of part-time work (i.e., 30 hours per week) and part-time school (i.e., six years to graduation) to working 10 hours per week, attending college full-time, and graduating in the normal time of four academic years. Define Y_{HS} and Y_C as earnings before and after college completion, respectively. Students enroll in college in year t and work until retirement in year T, with a discount rate of r. For ease of exposition, we assume students in this scenario may work 30 hours per week, earning $\frac{3}{4}Y_{HS}$ each year, and attend college part-time for six years, paying ωF annually in direct costs, where ω is the fraction of direct college costs, F, borne by the student. Alternatively, students may choose to work 10 hours per week, attend college full-time, and graduate in four years. The costs of part-time and full-time enrollment are assumed to be equal. Students prefer to earn a baccalaureate degree in six years while working part-time over the traditional four-year path to degree completion if

$$(1)\quad \frac{3}{4}\sum_{t=1}^{6}\frac{Y_{HS}}{(1+r)^t}+\sum_{t=7}^{T}\frac{Y_C}{(1+r)^t}-\sum_{t=1}^{6}\frac{\omega F}{(1+r)^t}>\frac{1}{4}\sum_{t=1}^{4}\frac{Y_{HS}}{(1+r)^t}+\sum_{t=5}^{T}\frac{Y_C}{(1+r)^t}-\sum_{t=1}^{4}\frac{\omega F}{(1+r)^t}$$

holds, a condition that may be reduced to

$$(2)\quad \frac{4[Y_C+\omega F]}{Y_{HS}}<\frac{2(1+r)^6+(1+r)^2-3}{r(r+2)}.$$ [6]

Whether equation (2) holds depends on parameters in the maximization problem, including risk preferences, direct college costs, and the returns to a baccalaureate degree. All else equal, students are more likely to pursue a nontraditional, longer path to college completion when discount rates are high, returns to a college degree are low, and direct schooling costs are low. In the extreme case, when the student bears no responsibility for direct college costs (i.e., $\omega = 0$), the student is more likely to prefer a longer path to degree completion. In contrast, the traditional path is more likely to be preferred when students are necessarily required to work during college, which is more likely to be the case when $\omega \neq 0$ and F is high. As an example, using figures from the Bureau of Labor Statistics, let Y_{HS} = \$34,600 and Y_C = \$57,800.[7] To demonstrate how time to degree may vary, we choose one non-top 50 public university (University of New Mexico, or UNM) and one top 50 public university (University of Washington, or UW) with F being equal to \$7,146 and \$10,974, respectively.[8] Assuming a standard discount rate of r = .05, and that students bear the full direct costs of college, the

[6] See Appendix A for the derivation of equation (2).
[7] Unemployment rates and earnings by educational attainment, 2016. Bureau of Labor Statistics, online at https://www.bls.gov/emp/ep_chart_001.htm (accessed 22 March 2018).
[8] Figures are for the 2018-2019 academic year and come from official UNM and UW websites.

simplified model results in UNM students preferring to take six years while working part-time over the traditional four-year path working ten hours per week. Because direct college costs are larger, UW students prefer the traditional four-year path under these conditions. However, if the discount rate is assumed to be $r = .10$, students now value current consumption more relative to future consumption, and students at both universities prefer a nontraditional path to degree attainment. This is also the case when students incur zero direct college costs. Note this simplification of human capital assumes no uncertainty regarding the costs and benefits of college, nor does it consider the possibility of binding credit constraints. Nevertheless, the exercise demonstrates that students may rationally choose a longer, nontraditional path to degree attainment under reasonable assumptions.

3. Is delayed graduation punished in the labor market?

Few studies examine the effect of time to degree on early-career wages, likely due to a clear endogeneity problem. Factors such as ability and institutional quality are expected to affect both time to degree and early-career wages. For example, low ability students generally earn less money than their high ability counterparts yet will still increase their earnings by obtaining a degree. Students that attend lower quality institutions may earn less than those attending higher quality institutions, however it may be the characteristics of the institution that are contributing to the wage differential, not necessarily the time it takes the student to complete. Without adequate controls and well-performing instruments, it is difficult to separate the effects of student ability and institutional quality from time to degree in the wage equation, and ordinary least squares (OLS) offers biased and inconsistent results. We identified four studies explicitly focusing on the relationship between time to degree and wages.

Groot and Oosterbeek (1994) examine the relationship between time to degree and wages using a survey of Dutch students. The main contribution of their study is a clever decomposition of actual years of schooling into five categories: effective years, defined as the number of years required to obtain a certain degree; skipped years; repeated years; inefficient years, defined as years taking courses at a lower level than the student is already qualified; and dropout years, defined as years spent in school not leading to a degree. This decomposition allows for direct testing of the human capital hypothesis against the screening hypothesis. Evidence of the screening hypothesis is expected to be characterized by two findings. First, delayed degree completion should signal low ability, so that inefficient and repeated years would have a negative effect on wages. Second, dropout years should have a nonpositive effect on wages since no ability signal is communicated. OLS results support the human capital hypothesis: skipped years are associated with lower wages and dropout years are associated with higher wages, suggesting a positive relationship between time to degree and wages.

Closely following Groot and Oosterbeek (1994), Aina and Pastore (2012) decompose years of schooling using a cross-section of workers in Italy. Due to data limitations, the decomposition is not as rich as Groot and Oosterbeek, as they are not able to identify inefficient, skipped, or dropout years of schooling. Instead, the authors define actual years of schooling as the sum of effective years, repeated years, and delayed years. The variable of interest, delayed years of schooling, is defined as the number of years (from first grade) a respondent took to earn their first baccalaureate degree, less effective years and repeated years. Defined this way, delayed years range from zero to 12. Because the number of years it takes to earn a baccalaureate degree in Italy ranges from 16 to 19 years, and children starting first grade in Italy are typically six years old, estimates of graduation delay penalties include older workers in their thirties. There is a clear selection issue to be noted here. Workers returning to school to complete degrees in their thirties are likely different from workers that took perhaps one or two extra years to earn a degree. Consequently, OLS estimates which suggest that delayed graduates incur a seven percent wage penalty are difficult to interpret.

Marking a departure from OLS approaches, Brodaty, Gary-Bobo, and Prieto (2008) use three-stage least squares (3SLS) to estimate a system of four structural equations of wages, employment, schooling, and graduation delay using a survey of college graduates in France. The authors use several instruments to identify the model, all constructed using respondent housing information prior to entering college. Years of schooling is not decomposed—instead, the authors define graduation delay as the individual's age at graduation less the average school-leaving age for others earning a similar credential. Results suggest that a one-year delay in college graduation (relative to peers in terms of age) results in a nine percent wage penalty. We also find these results difficult to interpret, but for reasons different than Aina and Pastore (2012). First, the authors use six instruments, all of which are likely themselves endogenous.[9] For example, the authors use vocational school openings and distance to the nearest college when the student was entering sixth grade as instruments for years of schooling and graduation delays. They assert that because instruments are constructed using information predating the survey period, they are predetermined and therefore exogenous (p. 6). However, using predating information to construct instruments only guards against reverse causality—it in no way guarantees instrument exogeneity. As an example, students from high socioeconomic status

[9]Instruments include the stock of vocational schools when the student was in the sixth grade; the change in the stock of vocational school between sixth grade and entering college; distance to the nearest college and its square when the student was in sixth grade; local unemployment rates; and an indicator of living in Paris in the sixth grade.

families may be more likely to locate near universities and have parents that may be more socially connected or have stronger preferences for their children to work in certain occupations. Second, the authors presumably use 3SLS as they suspect error terms across equations to be correlated with one another (due to the presence of unobserved ability in each equation, for example). In this case 3SLS is more efficient than 2SLS, but only if errors are homoskedastic—often an untenable assumption in observational studies. Otherwise, 3SLS estimates are inconsistent (see Cameron and Trivedi, 2005; Wooldridge, 2010 for details). For this reason, it is difficult to know whether estimated wage penalties are statistically different from zero.

The most recent study of graduation delay and early career earnings uses Baccalaureate & Beyond 1993 – 2003 survey data to estimate how different lengths of graduation delay affect employment and earnings in the U.S. The novelty of Witteveen and Attewell's (2019) study is examining how different types of delay (e.g., stopping out versus working full-time during school) impact post-college earnings. Similar to previous studies, the authors find that any delay in graduation results in a meaningful reduction in early career earnings—between eight and 15 percent, depending on the length of delay. Wage penalties appear to be driven by stopping out rather than working full-time during college. Yet, like studies noted above, graduation delay is again assumed to be exogenous in the wage equation, which is unlikely. As a result, findings are descriptive in nature.

Our study makes two contributions to the literature. First, it is the only study we have identified to explicitly explore the causal effect of delayed graduation on early-career wages in the United States. All previous studies attempt to test between human capital and signaling theory while not adequately addressing the endogeneity of time to degree. This topic is particularly relevant given recent policies proposed and enacted at federal-, state-, and institution-levels aimed at reducing baccalaureate time to degree. Second, although a well-studied topic, our results offer an additional test of the human capital hypothesis versus the screening hypothesis, a concept which remains central to studies of labor and education.

To operationalize our study of early-career wages and time to degree, we build on the model developed by Brodaty, Gary-Bobo, and Prieto (2008).[10] We deviate from their approach in our definition of the dependent variable, the set of instruments employed, and assumptions regarding the correlation structure of equation disturbances. Let d be a college graduate's time to degree in months. Subscripts are omitted for ease of exposition. Graduation delay, D, is defined as the graduate's time to degree less normal time in the United States, defined as 45 months, so that $D \equiv d - 45$.[11] We first suppose that worker productivity, q, is given by

(3) $ln(q) = \boldsymbol{S\alpha} + \beta D + \boldsymbol{X\gamma} + \delta_1 + \delta_2$

where $\boldsymbol{S}$ is a vector of variables indicating receipt of a graduate degree, and $\boldsymbol{X}$ is a vector of controls including potential experience and its square, a proxy for ability, race, ethnicity, gender, and institution quality. Note that productivity depends on two different dimensions of student ability, δ_1 and δ_2, with the former observed only by the potential employer and the latter observed by neither the researcher nor the potential employer. Direct productivity effects of schooling are given by $\boldsymbol{\alpha}$, which are expected to be nonnegative. The direct productivity effect of graduation delay, β, is expected to be zero under the human capital hypothesis after controlling for work experience. Assuming log-normally distributed wages, normally distributed

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[10]The reader is referred to Brodaty, Gary-Bobo, and Prieto (2008) for a full derivation of the econometric model.
[11]Forty-five months was chosen as it represents a "four-year" stay from fall in year one to spring in year four. A histogram of time to degree in months is presented below which appears to support this choice.

random error terms, and wages set according to expected worker productivity conditional on the employer's information set, the wages of college graduates, *w*, may be expressed as

(4) $ln(w) = \boldsymbol{S\alpha} + \beta D + \boldsymbol{X\gamma} + \delta_1 + E(\delta_2 | \boldsymbol{S}, D, \boldsymbol{X}, \delta_1)$.

Note that the employer may use graduation delay, *D*, as a signal conveying information about unobserved ability, δ_2.[12] Conditional expectations are linear due to assuming normality of variables, giving the following expression

(5) $E(\delta_2 | \boldsymbol{S}, D, \boldsymbol{X}, \delta_1) = \boldsymbol{S\alpha}' + \beta' D + \boldsymbol{X\gamma}' + \theta\delta_1$

where $\boldsymbol{\alpha}', \beta', \boldsymbol{\gamma}'$, and θ are theoretical regression coefficients. This result is substituted into wage equation (4) to obtain a modified version of Mincer's (1974) human capital earnings function with primed coefficients indicating screening effects as in Spence (1973).

(6) $ln(w) = \boldsymbol{S}(\boldsymbol{\alpha} + \boldsymbol{\alpha}') + (\beta + \beta')D + \boldsymbol{X}(\boldsymbol{\gamma} + \boldsymbol{\gamma}') + (1 + \theta)\delta_1$.

Graduation delay now appears in the earnings function for two reasons: there is a direct productivity effect on earnings, β, and a screening effect on earnings, β'. These two components are not identified individually without making stronger assumptions. Accordingly, we infer which mechanism dominates based on the sign of $(\beta + \beta')$. Because we expect β to be zero, an overall negative effect of graduation delay on early career wages favors the screening hypothesis. In contrast, a nonnegative coefficient on graduation delay is considered evidence that the human capital hypothesis is the dominant mechanism.

4. Data

[12] It is arguable whether employers observe undergraduate time to degree in the job screening/interviewing process. However, given that both education and work experience are observed, it seems reasonable that an employer may deduce time to degree using these two, or simply ask during the interview process, which is not uncommon. Moreover, employers may also request academic transcripts.

Our primary data source is the restricted-use Education Longitudinal Survey of 2002 (ELS:2002), which provides students' secondary, postsecondary, and subsequent labor market outcomes for a nationally representative sample of the 2004 graduating high school cohort.[13] The dependent variable in our analysis is the natural log of hourly wages in 2011 dollars. This measure is obtained at the third and final follow-up, approximately eight years after the cohort graduation date, defined as June 1, 2004.

We limit the sample to college graduates who earned a high school diploma and enrolled in college within two years, including students who did not graduate high school in normal time. Because the survey only follows students eight years after their expected high school graduation, it does not permit analysis of nontraditional students such as those who matriculate in their late 20s or later. This is not necessarily a limitation as our focus is on early-career wages when employers are more likely to use graduation delay as a negative productivity signal. Older job applicants typically have several years of work experience (and other information) on their application materials that employers can use to estimate their productivity. The sample is not conditional on employment at the third follow-up, so that some respondents have zero wages at the third follow-up. All descriptive statistics and regression estimates are constructed using the respondent's third follow-up panel weight.

Using the National Longitudinal Study of 1972 (NLS72) and the National Education Longitudinal Study of 1988 (NELS:88), Bound, Lovenheim, and Turner (2012) find that trends in time to degree across 1972 and 1992 graduating high school cohorts vary significantly

[13]ELS:2002 documentation explicitly notes the survey is intended to inform policymakers of the "rate of progress through postsecondary curriculum" and the "social and economic rate of return on education to both the individual and society" (Ingalls et al. 2014, pp. 10).

according to the student's first institution type. The authors classify students' first institutions into five categories: non-top 50 public colleges, top 50 public colleges, less selective private colleges, highly selective private colleges, and community colleges. This categorization is based on 2005 *U.S. News & World Report* college rankings. Highly selective private colleges include the top 50 ranked private colleges, the top 65 ranked liberal arts colleges, and four U.S. Armed Services Academies: the U.S. Military Academy at Westpoint, the U.S. Naval Academy, the U.S. Coast Guard Academy, and the U.S. Air Force Academy. We adopt the same categorization scheme to facilitate comparison, except that we exclude students who started at community colleges since they do not have institution selectivity data. Appendix B details which colleges fall in each category.

Several variables are included in wage equations to help isolate the effect of time to degree on earnings. Standardized test scores, measured in terms of composite ACT score, are included to capture observed student ability. Higher standardized test scores are expected to be positively correlated with future earnings (Betts and Grogger, 2003). Potential work experience and its square are included, as workers earn more as they acquire additional labor market experience, but at a diminishing rate (Mincer, 1974; Heckman *et al.*, 2003; Lemieux, 2006). Variables capturing whether a post-baccalaureate degree has been earned are included, as additional credentials are expected to increase future earnings (Card, 2001). Family characteristics, such as parents' highest educational attainment and household income, are also included as they are strong predictors of college completion (Bowen, Chingos, and McPherson, 2009). Models include controls for demographic characteristics, including respondent gender, race, and ethnicity. Institution quality is proxied using restricted 2004 Barron's Admissions Competitiveness Index data. This index considers several aspects of the incoming college cohort, including high school grades, standardized test scores, acceptance rates, and class rank. Additional proxies for institution quality include expenditures per full-time equivalent student and student-faculty ratios. We hypothesize that some employers may interpret time to degree differently depending on their familiarity with norms at the student's degree-granting institution. To account for this, we control for the distance between where a student earned their degree and where they are observed working in the final follow-up, measured in miles. We also include a dummy variable equal to one if the respondent worked during the first two years of college, and zero otherwise. This is included to capture work experience accrued during school, which is expected to generate returns to early career wages.[14]

A concern with the analysis is that our study period overlaps with the Great Recession. Changes in macroeconomic conditions may affect earnings. Students may be inclined to delay college graduation when facing a soft labor market. To address this concern, seasonally adjusted employment rates from the U.S. Bureau of Labor Statistics' Local Area Unemployment Statistics program are included for the month the student earned his or her degree to capture broad labor market conditions in the state the student graduated. Local unemployment rates at four years from enrollment are also included to help capture the effect of macroeconomic conditions on graduation delay. Estimates are qualitatively similar if these local unemployment rates are excluded.

Our innovation lies in how we address the endogeneity of graduation delay in the wage equation. Estimating a system of early-career wages and graduation delay using equation-by-

[14]Unfortunately, ELS:2002 does not ask about work experience for all years between the baseline and third follow-up, so actual work experience has multiple gaps in it. likewise, there are indications of working during the 2004-2006 academic years, but no actual work experience variables for the 2006-208 academic years.

equation OLS produces biased and inconsistent results. This is due to confounding factors affecting both graduation delay and early-career earnings, such as institution quality and unobserved student ability. For example, higher quality institutions provide better academic training and advising, resulting in better job placement and a higher likelihood of timely degree receipt. Higher-ability students are better equipped to complete college in four years and command higher wages in the labor market. We control for institution quality and observed ability to help isolate a causal path from the student's time to degree to their early-career wages. To address endogeneity from unobservable factors, we use an instrument for graduation delay which is plausibly unrelated to early-career wages except through its relationship with the student's own time to degree, after controlling for institution- and individual-level characteristics.

The instrument, the ratio of six- to four-year graduation rates at the student's first institution, is constructed using the Integrated Postsecondary Education Data System (IPEDS). This ratio captures the prevalence of graduation beyond 45 months at the institution-level, and is closely related to average time to degree, which is unavailable at the institution-level in the IPEDS.[15] We suspect that time to degree at the student's institution affects the student's own time to degree through what could be considered a peer effect. If a large proportion of one's peers in college are planning on overshooting normal time, the student may be more likely to consider this a valid path to degree attainment.[16] There are other reasons why this relationship

[15]This ratio is constructed from two variables in the 2004 IPEDS Graduation Rates survey. The 6-year graduation rate (BAGR150) is divided by the 4-year graduation rate (BAGR100).
[16]As one respondent told Bowen, Chingos, and McPherson (2009), graduating in four years is like "leaving the party at 10:30 p.m." (p. 237).

may hold. It may broadly reflect institution quality, the resources available to the student, or additional direct costs associated with delayed graduation. Other measures of institution quality, including student-faculty ratios and expenditures per student at the student's first institution, are meant to further isolate a causal path from the student's time to degree to their early-career earnings.[17] These measures broadly capture the amount of resources available to the student in their first institution, and are expected to be negatively correlated with the student's own time to degree. IPEDS data are from 2004, when ELS:2002 respondents were in their last year of high school. Thus, students in the ELS:2002 do not directly affect these measures. To examine the sensitivity of results to the year of the institution-level instrument, estimates using 2003 and 2005 IPEDS data are also reported. An additional sensitivity analysis uses a different construction of the time to degree ratio instrument—the ratio of eight- to four-year graduation rates at the student's first institution. This instrument may better reflect completion patterns for students attending institutions with particularly long time to degree, such as non-top 50 public institutions, for example. Please refer to Appendix C for results of this sensitivity analysis. Table C1 illustrates that choice of the cohort year and the definition of the time to degree ratio do not qualitatively change the story.

[17]Student-faculty ratios are calculated using the 2004 IPEDS Instructional Staff/Salaries survey and the 12-Month Enrollment survey. The numerator is the grand total of undergraduates enrolled for credit during the 12-month reporting period (FYRACE24). The denominator is the total employee count (EMPCNTT) for those with academic rank employed at least nine months out of the reporting year. Expenditures per student use data from the 2004 IPEDS Finance survey. For not-for-profit public institutions the numerator is the total amount of education-oriented expenditures, which is found by subtracting hospital services (F1C121) and research (F1C021) from total operating expenditures (F1C151). For not-for-profit private institutions total operating expenditures are F2E131, hospital services are F2E091, and research is F2E021. For for-profit private institutions total operating expenses are F3E07 and research is F3E02 (there are no hospital expenditures on behalf of these institutions).

As mentioned above, we also use methods that relax the exclusion restriction of the instrument following Conley, Hansen, and Rossi, (2012). This is meant to address the case in which we have an instrument that is only approximately exogenous—that is, there remains unobserved factors such as student ability or institution quality which are correlated with the instrument. Such methods allow one to construct analytical bounds of point estimates that assume approximate exogeneity of the instrument. Results of these tests are informative as to how much the exclusion restriction drives results employing traditional IV assumptions.

Table 1 presents the cumulative distribution of time to degree as well as its mean for ELS:2002 data. We report measures from Bound, Lovenheim, and Turner (2012) to examine whether trends in time to baccalaureate degree have persisted using this most recent NCES survey. Table 1 suggests that the overall mean time to degree has not changed significantly across 1992 and 2004 graduating high school cohorts, standing most recently at 4.83 years. Average time to degree for those at non-top 50 public colleges held steady at 4.93 years across 1992 and 2004 cohorts. At top 50 public schools there was a marked decrease in time to degree from 4.66 to 4.42 years, an average difference of approximately three months. The percent of graduates completing in four years or less at these schools increased from approximately 40 to 57 percent. There were small increases in time to degree for students starting at highly selective private institutions, and small decreases for those starting at less selective institutions. Overall, ELS:2002 data indicate that increases in time to baccalaureate degree discussed by Bound, Lovenheim, and Turner (2012) may have slowed or perhaps even reversed course.[18]

[18]We do not report whether differences in time to degree across the three surveys (NLS72, NELS:88, and ELS:2002) are significantly different from zero because we do not possess restricted-use versions of each survey.

Table 2 presents descriptive statistics by institution type for variables included in the analysis. The average hourly wage for graduates by first institution type ranges from $19.06 for non-top 50 public institutions to $24.49 for highly selective private institutions. Graduation delay follows an opposite pattern according to first institution type: students attending non-top 50 public universities exceeded normal time by an average of over ten months, while the average at highly selective private universities was just over two months. The ratio of six- to four-year completion rates at the student's institution was approximately 2.3 at non-top 50 public universities, meaning that students were nearly two-and-a-half times more likely to earn a degree within six years relative to four years. Not surprisingly, student-faculty ratios were higher at non-top 50 public universities, and expenditures per student were nearly twice as high at top 50 public universities relative to their non-top 50 counterparts. The average distance between where students graduated college and where they worked at the third follow-up is substantially larger for highly selective private institutions. Composite ACT scores are highest at highly selective private universities and lowest at non-top 50 public universities.

Figure 1 shows the distribution of graduation delay, revealing nearly half of college graduates in the sample earned their degrees on-time. Approximately seven percent of graduates took an additional two to three months beyond normal time to complete, roughly the length of an additional summer term. As expected, there are regular spikes in graduation delay every six or seven months beyond normal time. A very small number of students completed a baccalaureate degree in less than normal time.[19] The distribution of the time to degree ratio is shown in Figure

[19]Note minor differences in academic calendars across institutions may lead to a slightly different definition of "normal time." As a result, we consider 44-46 months to be normal time to degree.

2. This figure indicates that approximately 40 percent the sample attended an institution with a ratio of 6- to 4-year graduation rates between one and 1.5. There are a significant number of respondents attending institutions that graduated at least twice as many students within six years compared to four.

5. Empirical model

We estimate a system of structural equations for early-career wages and graduation delay using full information, multiple equation maximum likelihood methods akin to Zellner and Theil (1962).[20] Full information methods are necessary as some exogenous regressors are necessarily excluded from the first stage. For example, early-career wages postdate time to degree at the student's institution, so receipt of a master's or doctoral degree should be excluded from the graduation delay equation. A similar argument can be made for the distance between where the student graduated and where they worked at the final follow-up. The system may be expressed as

(7) First Stage: $Delay_{is} = \theta + \boldsymbol{X_{is}\zeta} + \pi Z_{is} + \mu_{is}$

(8) Second Stage: $ln(w_{is}) = \alpha + \beta \widehat{Delay}_{is} + \boldsymbol{X_{is}\gamma} + \boldsymbol{Y_{is}\delta} + \varphi Z_{is} + \sigma_s + \varepsilon_{is}$

where graduation delay and wages for student *i* in state *s* are functions of a common set of control variables included in the vector $\boldsymbol{X_{is}}$, such as potential experience and its square, ability, gender, race, ethnicity, family characteristics, and institutional characteristics.[21] $\boldsymbol{Y_{is}}$ includes determinants of wages that cannot plausibly be included in the student's time to degree equation

[20]Note this is not identical to the standard single equation, limited information instrumental variables estimator, although results are similar using this method.

[21]This is calculated by subtracting years of schooling plus six from the respondent's age. We use potential experience because actual years of experience cannot be precisely calculated at the third follow-up in ELS:2002.

for timing issues. The wage equation includes state fixed effects to account for heterogeneity across labor markets. μ_{is} and ε_{is} are idiosyncratic error terms. Z_{is} is an instrument assumed directly correlated with graduation delay, but not early-career wages, so that $\varphi = 0$ under strict instrument exogeneity. We report results of tests for weak instruments following Kleibergen and Paap (2006).

To account for a "plausibly exogenous" instrument, we relax the exclusion restriction and assume $\varphi \neq 0$ following methods in Conley, Hansen, and Rossi (2012). We provide point estimates by making assumptions regarding the support φ. This strategy allows the exclusion restriction to be violated in various ways. We allow symmetric violation of instrument exogeneity assuming $\varphi \in [-.i, .i]$ for i = 1, 2, …, 5. We next consider asymmetric violations of instrument exogeneity that assume the instrument is positively correlated with wages, where $\varphi \in [0, .i]$ for i = 1, 2, …, 5. Lastly, we assume that the IIV is negatively correlated with wages, so that $\varphi \in [-.i, 0]$ for i = 1, 2, …, 5.

In addition to employing the strategy proposed by Conley, Hansen, and Rossi (2012), we considered constructing analytical bounds following Nevo and Rosen (2012). However, for Nevo and Rosen bounds to be informative there must be negative correlation between the instrument and endogenous regressor. We considered simply taking the inverse of the time to degree ratio to force this negative correlation, however the support of the transformed instrument would be substantially different due to some observations being undefined due to division by zero.

Lastly, we are concerned that any causal interpretation may be confounded by two other early career outcomes: the likelihood of being employed and the likelihood of having received a graduate degree at the third follow-up. Graduation delay, for example, may directly impact the likelihood of being employed, and employment may subsequently impact early career wages, creating a "backdoor" path from graduation delay to wages. A similar argument may be made for earning a graduate degree. Ruling out these backdoor paths is thus an important step in isolating a causal path from graduation delay to early career wages. We control for these two outcomes in regressions, but we also consider them as outcomes in 2SLS models as presented in equations (7) and (8). We are less concerned about these potential backdoor paths if we find no evidence of a direct link between graduation delay and employment, and graduation delay and earning a graduate degree.

6. Results

6.1 Validity of the instrument

Valid estimation via 2SLS requires two conditions be met. First, instruments must be highly correlated with endogenous variables. This can be assessed in various ways. A naive approach involves conducting simple *t*-tests on instrument coefficients in the first stage. Evidence of instrument relevance is given by estimated coefficients which are statistically different from zero at conventional significance levels. A more robust approach is to estimate partial *F*-statistics for instruments in the first stage. For this approach, we follow Kleibergen and Paap (2006) and perform a generalized reduced rank test that is robust to heteroskedasticity. Partial *F*-statistics which test the null hypothesis of weak instruments are considered sufficiently large if they are approximately ten in magnitude or larger, with larger values presenting stronger evidence against weak instruments (Cragg and Donald, 1993; Stock and Yogo, 2005).

Second, instruments must be exogenous. In other words, instruments should only affect the dependent variable through the endogenous variable. Because verifying instrument exogeneity requires knowledge of the true model error, this requirement cannot be tested and instead must be maintained. In our case, instrument exogeneity requires that time to degree at the student's institution only affect early-career earnings through the student's own time to degree, after controlling for ability and institution quality. Although we argue for instrument exogeneity, it is impossible to know for certain whether this requirement holds. For this reason, we present results assuming $\varphi \neq 0$, which allows us to partially identify how much the imposition of strict exogeneity drives the main results.

Column (1) in Table 3 presents first-stage results. Results suggest a strong positive relationship between time to degree ratios and graduation delay. A one-unit increase in the time to degree ratio is expected to increase the student's own graduation delay by approximately 10 weeks. A simple *t*-test shows that we can reject the null hypothesis that the estimated first-stage coefficient is not different from zero at the one percent-level. The Kleibergen-Paap *F*-statistic is 24.08—strong evidence against weak instruments.

6.2 Main Results

As mentioned before, Table 3 suggests a strong positive relationship between the student's own time to degree and the average at their institution. Several other variables are economically significant in the first stage. A one percentage point increase in the local unemployment rate at the time of graduation is associated with nearly 3.5 additional months of graduation delay. This supports the hypothesis that some students attempt to "wait out" soft labor markets by delaying college graduation. The local unemployment rate four years after first enrollment is meant to capture any macroeconomic effects of the Great Recession. Results reveal a large negative relationship between time to degree and the local unemployment rate at normal time to degree. During the Great Recession students may have felt pressure to start working as soon as possible, or were possibly less willing to overshoot the four-year mark and incur the direct costs of delayed graduation. Estimates also suggest that higher ability translates into shorter time to degree, consistent with Flores-Lagunes and Light (2010). A ten-point increase in the composite ACT score results in a two-month decrease in graduation delay, roughly the length of an eight-week summer course. Women are found to be less likely to delay graduation compared to men.

Column (2) in Table 3 replicates previous work which likely does not break the endogeneity of time to degree in the wage equation. Equation-by-equation OLS suggests a wage penalty of 4.8 percent for a one-year graduation delay, similar to findings in Brodaty, Gary-Bobo, and Prieto (2009), Aina and Pastore (2012), and Witteveen and Attewell (2019), although smaller in magnitude. Column (3), which instruments for graduation delay, finds no such an effect. The 2SLS point estimate of $(\beta + \beta')$ is positive and not statistically different from zero. Other notable findings emerge from 2SLS results. Higher observed ability, as measured by standardized test scores, is associated with higher early-career earnings—a ten-point increase in the composite ACT score predicts a nine percent increase in wages. Results suggest a 7.5 percent wage penalty for women compared to men.

Although point estimates for $(\beta + \beta')$ are not statistically significant, it is still possible for time to degree to serve as a productivity signal for certain subgroups within the sample. To test for wage penalties on various subgroups, we conduct two subanalyses. First, we run models by institution type (i.e., non-top 50 public, top 50 public, less selective private, and highly selective private). Table 4 presents results by institution type. Cell sizes are now much smaller, but significant wage penalties are detected for top 50 public institutions (10.8 percent) and less selective private institutions (9.6 percent) for each year of delay using equation-by-equation OLS. As in our main results, 2SLS produces no evidence of a wage penalty for delayed college graduation. Second, we run models including interactions of fitted graduation delay with various characteristics such as gender, race, institution type, and the state unemployment rate when the student graduated college. We include gender and race as interactions to see whether there is any evidence of a wage penalty for women and students of color, respectively. Employers may be more sensitive to graduation delay from students attending less prestigious institutions, which motivates our inclusion of interaction terms involving the student's institution type. Additionally, we include interactions with local unemployment rates at graduation in order to test whether employers discriminate based on time to degree in markets with the weakest labor demand.

Because time to degree is itself an endogenous explanatory variable in the wage equation, including interactions of graduation delay with other variables results in additional endogenous regressors. A straightforward way to deal with additional endogenous regressors is to instrument for them using interactions of our preferred instrument, the time to degree ratio at the student's institution, with the variable of interest. The intuition is simple: we provide evidence above that the time to degree ratio is an appropriate instrument for the student's own graduation delay. Accordingly, the interaction of the time to degree ratio and the female dummy may serve as an instrument for the interaction between the student's own graduation delay and the female dummy variable, for example. This approach is dealt with in Wooldridge (2010).[22] In addition to the

[22]See Example 6.2 in Wooldridge (2010) for an application of this strategy (pp. 133-134).

interaction effect, the female main effect is also included in the regression to avoid spurious statistical significance of interaction terms.

Results of models including interaction effects are presented in Table 5. Similar to results without interaction effects, there is evidence of a wage penalty when endogeneity is ignored. Main effects of delayed graduation are estimated between 4.8 and 6.0 percent for each year of delay beyond normal time. Those graduating from top 50 public institution experience additional penalties of 8.4 percent for each year of delay. As before, 2SLS offer little evidence of main effect of graduation delay on early career wages. We do, however, find evidence that delayed graduation results in higher early career wages for students graduating in areas with high labor demand, although wage premia are erased in areas of high unemployment. 2SLS estimates suggest no wage penalty for female students, non-white students, or those attending various institution types.

As evident in Table 6, we find no evidence of any meaningful direct relationship between graduation delay and employment or graduation delay and receipt of a graduate degree by the third follow-up. This suggests that we do not need to be overly concerned about confounding backdoor paths from time to degree to early career earnings as discussed earlier. Point estimates are small in magnitude and not statistically different from zero. In terms of the likelihood of having earned a master's or doctorate by the third follow-up, women are 2.3 percentage points more likely to have done so compared to men. Those earning a graduate degree are significantly more likely to be employed at the third follow-up (12.2 percentage points for those with a master's and 46.7 percentage point more for those with a doctorate). Higher levels of potential work experience are positively correlated with employment at the third follow-up.

6.3. Relaxing the exclusion restriction

Table 7 reports results when relaxing the exclusion restriction in the main analysis. Analytical bounds constructed using the strategy in Conley, Hansen, and Rossi (2012) are presented first. When we assume that the support of φ is asymmetrically nonzero, 95 percent confidence intervals include zero, with larger bounds associated with a large overall support region. When we assume asymmetric bounds with $\varphi > 0$, we now find 95 percent confidence intervals that are negative, with upper bounds near zero. Estimated lower bounds become quite large in magnitude as the upper bound of φ increases. It is important to note, as argued in the theoretical section, that the plausibly exogenous instrument cannot be reasonably expected to be positive. This would imply that taking longer to graduate college increases your early career earning—we theoretically expect nonzero returns to time to degree rather than strictly positive returns. In the last case we assume that φ is asymmetrically negative, which is in line with previous studies finding that graduation delay reduces wages for early career college graduates. Violating the exclusion restriction in this manner also results in 95 percent confidence intervals that contain zero. We thus interpret the results of sensitivity analysis following Conley, Hansen, and Ross (2012) to support results in the main analysis. This suggests that the exclusion restriction in the main analysis is not driving results to a large degree.

7. Conclusions

This paper examines the effect of delayed baccalaureate graduation on early-career wages in the United States. OLS estimates not adequately addressing endogeneity in the wage equation suggest large, significant wage penalties for students taking longer to graduate, a finding consistent with previous literature and in support of delayed graduation as a negative productivity signal. We proxy for institution quality, student ability, and instrument for the student's own time to degree using the institutional average. This approach reveals a pattern we find repeatedly using different instrument sets and subpopulations—while OLS estimates find a one-year delay in graduation results in a wage penalty of approximately five percent, we find no evidence of any wage penalty using a 2SLS model of early-career wages. This result holds when we include interaction effects to test for wage penalties for women, students of color, by institution type, and for students graduating college in soft local labor markets. We reject the hypothesis of weak instruments and offer evidence in favor of instrument relevance.

Time to degree may be costly for students and institutions for various reasons. At the student-level, delayers may incur additional direct costs in the form of increased tuition, books, and room and board. However, as demonstrated in our model of human capital, these costs may easily be countered by lower opportunity costs for students under reasonable assumptions. Not addressed in this study is how taking a slower path to degree completion may reduce the psychic costs of attending college. Given the findings of our theoretical and empirical exercises, it is not clear that taking longer to complete college is a problem that needs fixing. Proposed reforms aimed at returning to the four-year college degree must instead rely on other arguments, such as strained institutional resources or the crowding out of new college entrants by delayers—arguments deserving further scrutiny in the literature.

We hope to inform policymakers concerning the costs and benefits of lengthened time to degree, especially those at institutions currently considering or actively enforcing policies discouraging alternative paths to degree completion taking longer than four years. Utility-maximizing students may be better off pursuing a longer path to degree attainment. It is important to then consider the growing proportion of nontraditional students in higher education in the United States, and to promote policies accommodating rational decisions to work during school while taking fewer credit hours per semester. Policies penalizing students for not remaining on a four-year path to degree completion may ultimately hurt their very chances of completing college at all.

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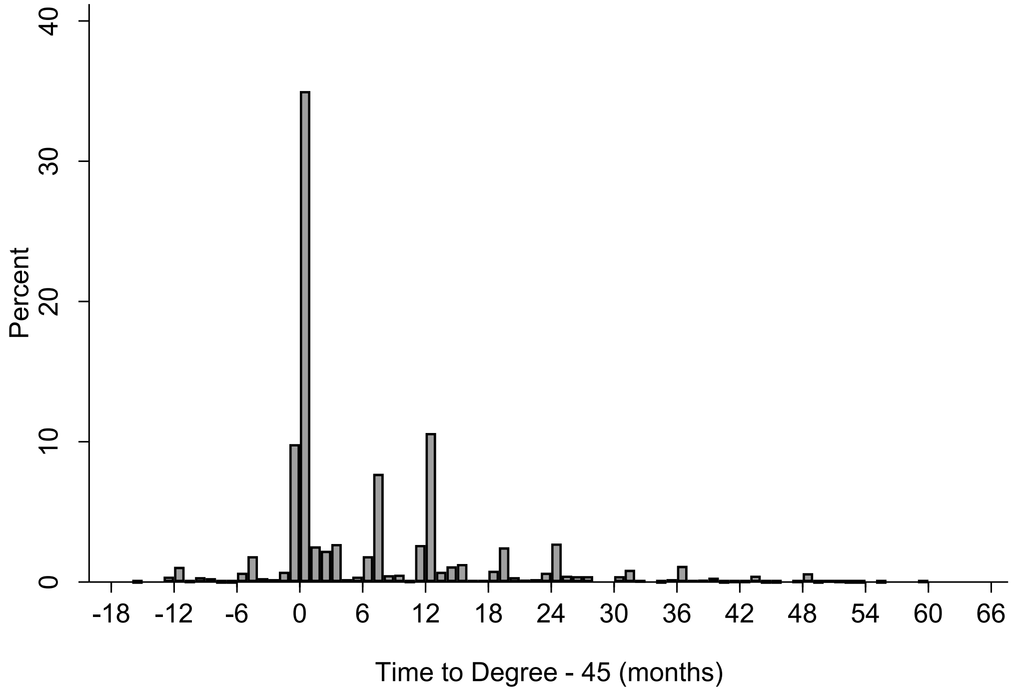
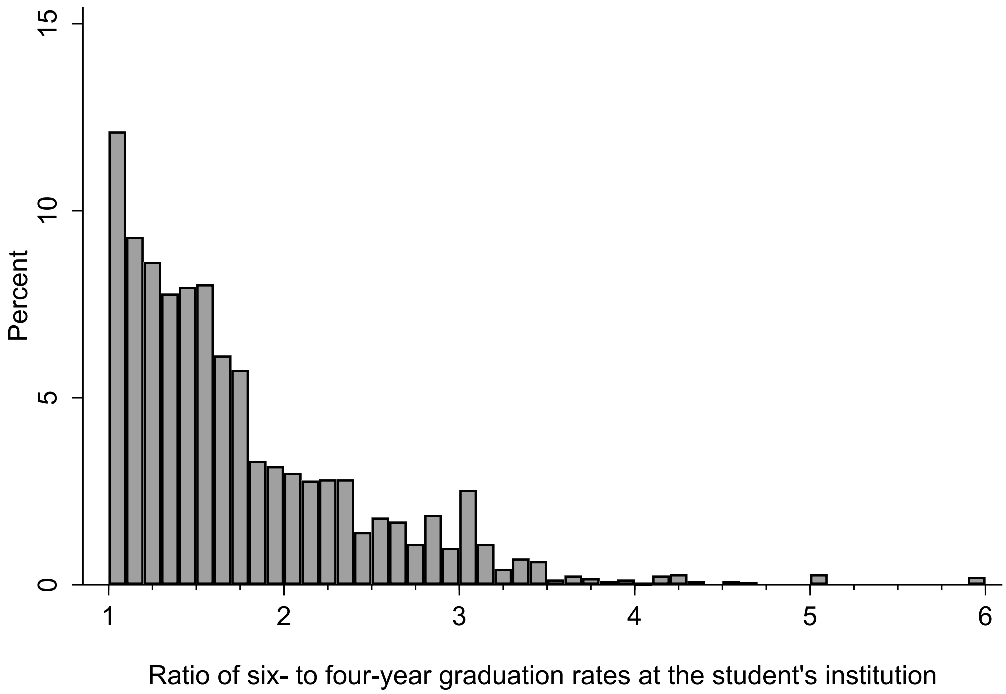


Source: Education Longitudinal Study of 2002.

Figure 1. Distribution of graduation delay, baccalaureate earners, ELS: 2002.



Source: Authors' calculations, IPEDS 2004, and ELS:2002. Outliers significantly above six, which constitute 0.41 percent of the sample, are not shown.

Figure 2. Distribution of the ratio of six- to four-year graduation rates

Table 1. Time to degree distributions for all graduates by first institution type

	time to degree distribution				
	4	5	6	7	mean
full sample:					
NLS72	53.1	81.8	90.6	96.3	4.48
NELS:88	39.4	72.7	88.3	94.7	4.81
ELS:2002	42.3	72.1	85.7	93.5	4.83
non-top 50 public:					
NLS72	49.7	82.3	91.1	96.3	4.49
NELS:88	29.1	68.8	87.8	95.1	4.93
ELS:2002	34.2	68.5	85.0	94.1	4.93
top 50 public:					
NLS72	52.7	81.5	89.2	96.4	4.49
NELS:88	39.7	82.0	93.7	96.6	4.66
ELS:2002	56.7	85.2	95.2	98.1	4.42
less selective private:					
NLS72	66.7	87.3	94.0	98.7	4.28
NELS:88	58.0	84.6	93.4	98.6	4.60
ELS:2002	56.1	83.4	92.5	96.1	4.51
highly selective private:					
NLS72	65.2	88.2	93.8	96.8	4.31
NELS:88	73.1	91.9	98.1	99.8	4.20
ELS:2002	68.6	91.7	96.3	98.2	4.28
community college:					
NLS72	36.5	67.8	83.0	92.6	4.90
NELS:88	15.5	44.2	70.8	83.6	5.58
ELS:2002	16.5	43.9	64.4	81.6	5.69

Note: NLS72 and NELS:88 figures reproduced from Bound, Lovenheim, and Turner (2012). ELS:2002 calculations were made using third follow-up panel weights. In each survey the sample includes baccalaureate-earners enrolling in a postsecondary institution within two years of their high school cohort graduation month. High school cohort graduation month is assumed to be June 1972 for NLS72, June 1992 for NELS:88, and June 2004 for ELS:2002. Students were followed for eight years following the expected high school cohort graduation month.

Table 2. Sample characteristics of employed college graduates in the ELS:2002

	non-top 50 public	top 50 public	less selective private	highly selective private
hourly wage (2011 USD)	19.06 (9.06)	21.43 (11.31)	20.06 (11.60)	24.49 (14.00)
graduation delay	10.51 (12.28)	4.77 (9.57)	3.39 (9.54)	2.24 (7.54)
time to degree ratio	2.29 (.81)	1.52 (.26)	1.33 (.35)	1.27 (.43)
student-faculty ratio	13.79 (4.35)	9.62 (1.93)	11.48 (8.12)	6.51 (2.66)
expenditure per student ($1,000s 2004 USD)	15.13 (7.43)	32.63 (12.92)	20.28 (8.50)	70.22 (87.94)
distance college-work (1,000s miles)	.23 (.50)	.32 (.61)	.22 (.45)	.52 (.75)
master's	.15	.17	.18	.16
doctorate	.02	.05	.04	.07
unemployment rate at graduation	7.65 (2.31)	6.85 (2.15)	6.46 (2.13)	6.13 (1.73)
unemployment rate 4 years after enrollment	5.90 (1.19)	6.07 (1.08)	6.00 (1.22)	5.94 (1.09)
experience	3.50 (.93)	3.41 (1.06)	3.35 (1.36)	3.39 (1.15)
ACT composite	22.60 (3.96)	25.92 (3.77)	24.00 (4.15)	28.50 (3.62)
female	.53	.53	.59	.53
white	.75	.75	.80	.77
Hispanic	.07	.06	.09	.08
black	.11	.05	.06	.02
American Indian	.003	.01	.005	.002
Asian	.03	.08	.03	.10
two or More Races	.04	.05	.02	.02
Hawaiian/pacific islander	.002	.001	.00	.00
Barron's – most competitive	.001	.12	.00	.50
Barron's - highly competitive	.03	.30	.10	.27
Barron's - very competitive	.20	.45	.40	.22
Barron's - competitive	.58	.13	.38	.004
Barron's - less competitive	.12	.00	.05	.00
Barron's - non-competitive	.05	.00	.01	.00
Barron's - special designation	.001	.00	.01	.00
observations	990	510	550	340

Source: ELS:2002, Barron's Admissions Competitiveness Index 2004, Bureau of Labor Statistics, and IPEDS. Standard deviations are in parentheses. Final follow-up panel weights are used. Seasonally-adjusted unemployment rates are at the state-level. Time to degree ratios are defined as the ratio of six- to four-year graduation rates at the student's own institution.

Table 3. Wage models of graduation delay penalty, all institutions

	OLS	OLS	2SLS
	(1)	(2)	(3)
variable	Delay	Wages	Wages
graduation delay (months)		-.004***	.015
		(.001)	(.010)
time to degree ratio	2.228***		
	(.263)		
student-faculty ratio	.014	-.001	-.003
	(.035)	(.002)	(.003)
expenditures per student	.005	.0007**	.0007**
	(.005)	(.0003)	(.0003)
unemployment rate at graduation	3.331***	.001	-.066*
	(.089)	(.007)	(.036)
unemployment rate 4 years after enrollment	-3.418***	.003	.073*
	(.163)	(.014)	(.039)
master's		.066	.154
		(.087)	(.102)
doctorate		.386**	.280
		(.179)	(.195)
experience	1.634***	.064	.012
	(.590)	(.123)	(.131)
experience^2	.140	-.006	-.002
	(.111)	(.016)	(.017)
ACT composite	-.220***	.005	.009**
	(.048)	(.003)	(.004)
female	-1.417***	-.107***	-.075***
	(.333)	(.022)	(.028)
Hispanic	.804	-.033	-.055
	(.681)	(.044)	(.047)
black	.886	-.051	-.074
	(.659)	(.043)	(.047)
American Indian	-.356	.044	.058
	(2.492)	(.162)	(.169)
Asian	.622	.022	.006
	(.764)	(.050)	(.052)
two or more races	-.285	-.022	-.021
	(.879)	(.057)	(.059)
Hawaiian/Pacific Islander	-2.914	.146	.205
	(4.379)	(.280)	(.294)
institution selectivity fixed effects	YES	YES	YES
state fixed effects	NO	YES	YES
college-work distance	NO	YES	YES
parents' education	YES	YES	YES
family income	YES	YES	YES
Kleibergen-Paap rk *F*-statistic			24.08
observations			2,340

Source: ELS:2002, IPEDS, Barron's Admissions Competitiveness Index of 2004. The dependent variable in equation (1) is the total time, measured in months, elapsed between first entering college and earning the first undergraduate degree, centered at 45 months. The dependent variable in equations (2) through (4) is the natural log of hourly wages at the third follow-up. The Kleibergen-Paap rk *F*-statistic tests the null hypothesis of weak instruments. Robust standard errors are reported in parentheses.

Table 4. Wage models of graduation delay penalty by institution type

	OLS	2SLS	obs.
	(1)	(2)	
institution type			
non-top 50 public	-.003	< .001	980
	(.002)	(.029)	
top 50 public	-.009**	.017	490
	(.004)	(.036)	
less selective private	-.008**	.006	530
	(.003)	(.033)	
highly selective private	-.007	-.014	340
	(.007)	(.031)	

Source: ELS:2002, IPEDS, Barron's Admissions Competitiveness Index of 2004. The dependent variable is the natural log of hourly wages at the third follow-up. All models include the same controls as listed in Table 3. Robust standard errors are reported in parentheses.

Table 5. Estimates of graduation delay with interaction effects

main/interaction effect	OLS (1)	(2)	(3)	(4)	2SLS (5)	(6)	(7)	(8)
graduation delay	-.003	-.004*	-.005*	-.003	.442**	-.009	.022	.023
	(.005)	(.002)	(.002)	(.002)	(.224)	(.025)	(.015)	(.015)
x unemployment rate	< .001				-.044*			
	(.001)				(.023)			
x female		< .001				.039		
		(.002)				(.042)		
x white			.002				-.021	
			(.003)				(.029)	
x top 50 public				-.007**				-.039
				(.003)				(.036)
x less selective private				-.004				-.019
				(.003)				(.023)
x highly selective private				.005				-.021
				(.006)				(.030)
observations				2,370				2,344

Source: ELS:2002, Barrons Admissions competitiveness Index 2004, and the IPEDS. The dependent variable is the natural log of hourly wages at the third follow-up. Models include both main effects and interaction effects. See Appendix B for definitions of institution type. All models include the same controls as listed in Table 3. Unemployment rate is measured at the month of graduation. Robust standard errors are reported in parentheses.

Table 6. The effect of delayed graduation on graduate degree attainment and employment

	(1)	(2)
	earned graduate degree	employed at third follow-up
graduation delay	< .001	-.006
	(.003)	(.006)
student-faculty ratio	< .001	.001
	(.001)	(.001)
expenditures per student	> -.001	< .001
	(< .001)	(< .001)
unemployment at graduation	< .001	.20
	(.010)	(.021)
unemployment in year 4	.005	.031
	(.011)	(.023)
master's	.952***	.122**
	(.029)	(.058)
doctorate	.008	.467***
	(.056)	(.112)
experience	.012	.269***
	(.038)	(.075)
experience^2	-.004	-.032***
	(.005)	(.010)
ACT composite	< .001	-.002
	(.001)	(.002)
female	.023***	-.030*
	(.008)	(.016)
hispanic	.013	-.010
	(.014)	(.027)
black	-.020	.004
	(.013)	(.027)
American Indian	-.031	.001
	(.049)	(.097)
Asian	-.007	-.052*
	(.015)	(.030)
two or more races	.008	-.035
	(.017)	(.034)
Hawaiian/Pacific Islander	.007	.014
	(.085)	(.169)
Barron's Admissions Competitiveness Index	YES	YES
college-work distance	YES	YES
state fixed effects	YES	YES
parents' education	YES	YES
family income	YES	YES
observations		2,340

Source: ELS:2002, Barrons Admissions Competitiveness Index 2004, and the IPEDS. Kleibergen-Paap rK *F*-statistics test the null hypothesis of weak instruments. Robust standard errors are reported in parentheses.

Table 7. Analytical bounds relaxing the exclusion restriction following Conley, Hansen, and Rossi (2012) and Nevo and Rosen (2012)

Model	95% CI of $\hat{\beta}$
OLS	[-.002, -.006]
2SLS	[-.005, .034]
Conley, Hansen, and Rossi (2012)	
$\varphi \in [-.5, .5]$	[-.114, .095]
$\varphi \in [-.4, .4]$	[-.093, .075]
$\varphi \in [-.3, .3]$	[-.073, .054]
$\varphi \in [-.2, .2]$	[-.052, .034]
$\varphi \in [-.1, .1]$	[-.033, .014]
$\varphi \in [0, .5]$	[-.114, -.004]
$\varphi \in [0, .4]$	[-.093, -.004]
$\varphi \in [0, .3]$	[-.073, -.004]
$\varphi \in [0, .2]$	[-.052, -.004]
$\varphi \in [0, .1]$	[-.033, -.004]
$\varphi \in [-.5, 0]$	[-.014, .095]
$\varphi \in [-.4, 0]$	[-.014, .075]
$\varphi \in [-.3, 0]$	[-.014, .054]
$\varphi \in [-.2, 0]$	[-.014, .034]
$\varphi \in [-.1, 0]$	[-.014, .014]

Note: $\boldsymbol{\varphi}$ is the coefficient on the instrumental variable (time to degree ratio) in the second stage (wage equation).

Appendix A: Mathematical derivations

Recall r is the discount rate, Y_H and Y_C are earnings with a high school diploma and a baccalaureate degree, respectively, and F is direct full-time schooling costs per year. Students prefer a six-year, .75 FTE (i.e., 30 hour per week) employment approach to traditional baccalaureate degree attainment (i.e., .25 FTE, four-year path) when

(A1) $\frac{3}{4}\sum_{t=1}^{6}\frac{Y_H}{(1+r)^t}+\sum_{t=7}^{T}\frac{Y_C}{(1+r)^t}-\sum_{t=1}^{6}\frac{F}{(1+r)^t}>\frac{1}{4}\sum_{t=1}^{6}\frac{Y_H}{(1+r)^t}+\sum_{t=5}^{T}\frac{Y_C}{(1+r)^t}-\sum_{t=1}^{4}\frac{F}{(1+r)^t}.$

To simplify this expression, we pull out constant terms from the summation operators, combine like terms, and apply the rule of finite geometric series which states that

(A2) $\sum_{k=1}^{n}a^k=\frac{a(1-a^n)}{1-a}$,

giving the following equation:

(A3) $\frac{Y_H}{4}\left[\frac{\frac{3}{1+r}\left[1-\left(\frac{1}{1+r}\right)^6\right]}{1-\frac{1}{1+r}}-\frac{\frac{1}{1+r}\left[1-\left(\frac{1}{1+r}\right)^4\right]}{1-\frac{1}{1+r}}\right]+Y_C\left[\sum_{t=7}^{T}\frac{1}{(1+r)^t}-\sum_{t=5}^{T}\frac{1}{(1+r)^t}\right]+F\left[\sum_{t=1}^{4}\frac{1}{(1+r)^t}-\sum_{t=1}^{6}\frac{1}{(1+r)^t}\right]>0.$

Removing summation notation for the second and third terms allows the expression to be reduced to:

(A4) $\frac{Y_H}{4}\left[\frac{\frac{3}{1+r}\left[1-\left(\frac{1}{1+r}\right)^6\right]}{1-\frac{1}{1+r}}-\frac{\frac{1}{1+r}\left[1-\left(\frac{1}{1+r}\right)^4\right]}{1-\frac{1}{1+r}}\right]-(Y_C+F)\left[\frac{1}{(1+r)^5}+\frac{1}{(1+r)^6}\right]>0$

Simplifying the second term and adding it to both sides produces:

(A5) $\frac{Y_H}{4}\left[\frac{\frac{3}{1+r}\left[1-\left(\frac{1}{1+r}\right)^6\right]}{1-\frac{1}{1+r}}-\frac{\frac{1}{1+r}\left[1-\left(\frac{1}{1+r}\right)^4\right]}{1-\frac{1}{1+r}}\right]>(Y_C+F)\left[\frac{r+2}{(1+r)^6}\right]$

Moving all the discount rate parameters to the left-hand side and simplifying yields the solution in equation (2):

(A6) $\frac{2(1+r)^6+(1+r)^2-3}{r(r+2)}>\frac{4[Y_C+F]}{Y_H}.$

Appendix B: U.S. News & World Report Classification

Table B1. U.S. News & World Report College Rankings, by institution type, 2005

	Highly Selective Private Schools	
Top 50 Public Schools	Top 65 Private Schools	Top 50 Liberal Arts Schools
University of California–Berkeley	Harvard University	Amherst College
University of Virginia	Princeton University	Williams College
University of Michigan–Ann Arbor	Yale University	Swarthmore College
University of California–Los Angeles	University of Pennsylvania	Wellesley College
University of North Carolina–Chapel Hill	Duke University	Carleton College
College of William and Mary	MIT	Middlebury College
University of Wisconsin–Madison	Stanford University	Pomona College
University of California–San Diego	California Institute of Tech.	Bowdoin College
University of Illinois	Columbia University	Davidson College
Georgia Institute of Technology	Dartmouth College	Haverford College
University of California–Davis	Northwestern University	Claremont-McKenna
University of California–Irvine	Washington Univ. of St. Louis	Wesleyan University
University of California–Santa Barbara	Brown University	Grinnell College
University of Texas–Austin	Cornell University	Vassar College
University of Washington	Johns Hopkins University	Harvey Mudd College
Pennsylvania State University	University of Chicago	Washington and Lee
University of Florida	Rice University	Smith College
University of Maryland–College Park	Notre Dame University	Hamilton College
Rutgers University–New Brunswick	Vanderbilt University	Colgate University
University of Georgia	Emory University	Oberlin College
University of Iowa	Carnegie Mellon University	Colby College
Miami University (Ohio)	Georgetown University	Bates College
Ohio State University	Wake Forest University	Bryn Mawr College
Purdue University	Tufts University	Colorado College
Texas A&M–College Station	Univ. of Southern California	Macalester College
University of Connecticut	Brandeis University	Scripps College
University of Delaware	New York University	Mt. Holyoke College
University of Minnesota–Twin Cities	Case Western Reserve	Barnard College
Indiana University	Lehigh University	Kenyon College
Michigan State University	Univ. of Rochester	College of the Holy Cross
Clemson University	Tulane University	Trinity College

Note: Adopted from Bound *et al.* (2012) and the 2005 U.S. News & World Report College Rankings. Highly selective private colleges also include the four U.S. Armed Services Academies: the U.S. Military Academy at Westpoint, the U.S. Naval Academy, the U.S. Coast Guard Academy, and the U.S. Air Force Academy.

Table B1. U.S. News & World Report College Rankings, by Institution Type, 2005 (continued)

	Highly Selective Private Schools	
Top 50 Public Schools	Top 65 Private Schools	Top 50 Liberal Arts Schools
SUNY at Binghamton	Rensselaer Polytechnic	Lafayette College
University of California–Santa Cruz	Yeshiva University	Occidental College
University of Colorado–Boulder	George Washington Univ.	Bard College
Virginia Tech.	Pepperdine University	Furman University
University of California–Riverside	Syracuse University	Whitman College
Iowa State University	Worcester Polytechnic	Union College
North Carolina State University	Boston University	Franklin and Marshall
University of Alabama	University of Miami	Sewanee College
University of Missouri–Columbia	Fordham University	University of Richmond
Auburn University	Southern Methodist Univ.	Connecticut College
University of Kansas	Brigham Young University	Centre College
University of Tennessee–Knoxville	Clark University	Dickinson College
University of Vermont	Stevens Inst. of Technology	Skidmore College
Ohio University	St. Louis University	Gettysburg College
University of Arizona	Baylor University	Pitzer College
University of Massachusetts–Amherst	American University	DePauw University
University of Nebraska–Lincoln	Howard University	Rhodes College
University of New Hampshire	Marquette University	Reed College
	University of Denver	
	University of Tulsa	
	Texas Christian University	
	University of Dayton	
	Drexel University	
	Illinois Institute of Technology	
	University of San Diego	
	Catholic University	
	Loyola University	
	Univ. of San Francisco	
	University of the Pacific	
	New School	
	Northeastern University	
	Seton Hall University	
	University of St. Thomas	

Note: Adopted from Bound *et al.* (2012) and the 2005 U.S. News & World Report College Rankings. Highly selective private colleges also include the four U.S. Armed Services Academies: the U.S. Military Academy at Westpoint, the U.S. Naval Academy, the U.S. Coast Guard Academy, and the U.S. Air Force Academy.

Appendix C: Sensitivity of results to cohort year and definition of instrument

Table C1. Sensitivity analysis using alternative IPEDS cohorts and an alternative instrument

	(1)	(2)	(3)
	2003 IPEDS	2005 IPEDS	The ratio of eight- to four-year graduation rates
graduation delay	.015*	.014	.012
	(.009)	(.009)	(.009)
student-faculty ratio	-.002	-.003	.002
	(.002)	(.003)	(.002)
expenditure per student	.0007**	.0007**	.0007**
	(.0003)	(.0003)	(.0003)
unemployment at graduation	-.068**	-.063**	-.056*
	(.031)	(.032)	(.033)
unemployment in year 4	.075**	.070**	.063*
	(.034)	(.035)	(.037)
master's	.157	.151	.141
	(.099)	(.099)	(.099)
doctorate	.248	.281	.296
	(.193)	(.193)	(.192)
experience	-.003	.012	.020
	(.130)	(.130)	(.129)
experience^2	> -.001	-.001	-.002
	(.017)	(.017)	(.017)
ACT composite	.009**	.009**	.009**
	(.004)	(.004)	(.004)
female	-.078***	-.078***	-.080***
	(.027)	(.027)	(.027)
hispanic	-.060	-.053	-.052
	(.047)	(.047)	(.047)
black	-.081*	-.074	-.071
	(.046)	(.046)	(.046)
American Indian	.054	.056	.056
	(.169)	(.168)	(.167)
Asian	.003	.007	.008
	(.052)	(.052)	(.052)
two or more races	-.026	-.020	-.021
	(.059)	(.059)	(.058)
Hawaiian/Pacific Islander	.207	.201	.196
	(.293)	(.292)	(.290)
Barron's Admissions Competitiveness	YES	YES	YES
college-work distance	YES	YES	YES
state fixed effects	YES	YES	YES
parents' education	YES	YES	YES
family income	YES	YES	YES
Kleibergen-Paap rk *F*-Statistic			
observations	2,340	2,330	2,340

Source: ELS:2002, Barrons Admissions Competitiveness Index 2004, and the IPEDS. The dependent variable is the natural log of hourly wages at the third follow-up. Kleibergen-Paap rK *F*-statistics test the null hypothesis of weak instruments. Robust standard errors are reported in parentheses.