How do education and human capital affect economic welfare? empirical research from the resources-based perspective

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Abstract

The degree to which workers can perform their jobs efficiently is influenced by their level of education, which allows them to develop competencies that align with the demands of contemporary market dynamics, resulting in higher wages. In this regard, this paper examines the causal effect of higher education on economic welfare from a resource-based perspective, employing a two-stage least squares (2SLS) approach. To achieve this, we analyze how additional years of postsecondary schooling impact wage income at the individual level and how these educational wage premiums vary across different industries. Our analysis is based on a large sample of approximately 20.5 million prime-age workers from the IPUMS USA database, spanning the years 2005 to 2023. To address endogeneity, we employ a Bartik-style instrumental variable constructed from state-level need and merit-based financial aid spending. Our findings indicate that an additional year of higher education has a causal effect on an individual's annual wage income, increasing it by approximately 13.9%. Furthermore, we have found positive and statistically significant returns to higher education across 15 industries. The results found that the magnitude of the income wage premium varies, ranging from 9.4% in Social Services to 17.0% in Agriculture. Our results contribute to the existing evidence that education remains a critical component of human capital development and, subsequently, wage income.

Keywords: *human capital, higher education, personal income, economic welfare, resource-based theory.*

JEL Codes: *I25, I26, J24, O15, O34, O47*

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1. Introduction

In today's highly competitive business world, knowledge-based capabilities have been identified as a critical factor in firms' ability to create value. Innovation, the efficient reorganization of resources, and adaptation to a rapidly changing business environment can only be achieved through effective human capital management. In other words, to address the challenges posed by emerging markets, it is essential to enhance the capabilities through human capital development. Therefore, knowledge has emerged as a critical factor in determining the competitive advantage of enterprises engaged in the production of goods and services.

A thorough review of the resource-based theory (RBT) developed by Barney (1991) reveals a predominant emphasis on valuable, rare, imperfectly imitable, and non-substitutable resources. This theory posits that these indispensable resources are more likely to satisfy customer needs and provide competitive advantages for firms that possess and acquire them. In this context, as demonstrated by several studies (Arik & Penn, 2007; Arik & Dunne, 2014), the competitiveness of nations, regions, and businesses is found to be influenced by investments in education. The enhancement of human capital productivity and efficiency through educational attainment can facilitate the development of the competencies demanded by industry and the market. The existing literature indicates that the return on investment from additional years of schooling can range from 5% to 20% (Psacharopoulos, 1994). When these findings are considered as a whole, it becomes evident that human capital exerts a pivotal influence on both microeconomic and macroeconomic performance.

In the extant literature on RBT, human capital is generally assessed using qualitative criteria. Therefore, the relationship between the microeconomic returns of education and a firm's competitive advantage lacks integration. Furthermore, the returns on education vary across different industries, and the role of education in fostering competitive advantage remains a matter of debate. Consequently, interpreting the value that education creates at both the individual and firm levels is challenging. The objective of this study is to address this emerging gap in the literature.

In this context, the primary research question guiding this study is: How do additional years of schooling affect individual income? The objectives of this study can be outlined as follows: (1) To employ statistical methodologies to assess the microeconomic return on education. (2) To evaluate these returns through RBT constructs in terms of valuable and non-substitutable resources. (3) To provide the implications for firms and institutions operating in different industries.

In the ensuing chapter, the conceptual framework underlying the impact of education on economic prosperity is examined through a review of relevant literature. Subsequently, the study's methodology, findings, and discussion were presented, respectively. Finally, suggestions for policymakers and future implications for subsequent academic studies were presented in line with the study's limitations.

2. Theoretical framework and literature Review

Considering the extant theoretical framework and literature concerning the relationship between education and economic welfare and growth (Romer, 1989; Mankiw et al., 1992; Barro & Sala-i-Martin, 2004; Greiner et al., 2005), the present study provides a detailed analysis of the return of an additional year of schooling to the income earned by individuals.

2.1. The return of an additional year of schooling to the earned income of individuals

In the contemporary business environment, characterized by intense competition, the enhancement of qualifications and the quality of scarce resources, along with their effective allocation, hold paramount importance. According to Barney (1991, 1995, 1996), a firm's resources are pivotal in determining its long-term competitive character, thereby enabling it to supply goods and services, as outlined within the framework of resource-based theory. According to this theory, as long as resources are distributed heterogeneously across firms and these differences remain stable over time, the possession of valuable, rare, imperfectly imitable, and non-substitutable resources will enable the firm to attain a competitive advantage over its rivals in the market. In conclusion, tangible and intangible resources are heterogeneous due to four features that remain stable over time. This heterogeneity allows for variation in firm performance, thereby providing firms with a competitive advantage (Wernerfelt, 1984; Ray et al., 2004).

Given that labor is one of the factors of production, it can be regarded as a primary and essential resource, as it can organize and manage the other factors. In this context, the term "human capital" was initially introduced by Goode (1959) to refer to "knowledge, skills, attitudes, aptitudes, and other acquired traits that contribute to production." Subsequently, human capital was defined by Becker (1962) as a form of investment encompassing a variety of expenses, including, but not limited to, educational expenses, on-the-job training, and medical care. In addition to these ones, a general and comprehensive definition was established by the OECD (1998) as "the knowledge, skills, competencies, and other attributes embodied in individuals that are relevant to economic activity." The cultivation of these skills, knowledge, or health practices has the potential to result in increased income. Along these terms, the cultivation of human capital is enabled by a range of investment channels, including education, on-the-job training, and medical care. These factors have been demonstrated to contribute to an increase in labor productivity, thereby facilitating enhanced income prospects for workers (Becker, 1962, 1964, 1975).

Lucas (1988) advanced the concept that human capital, as opposed to physical capital, plays a predominant role in the production of non-rival, nonexcludable goods. In other words, achieving success in such a competitive environment necessitates a strategic investment in human capital (Wright et al., 2001; Wofford, 2002; Acedo et al., 2006; Evans & Novicevic, 2010). Similarly, Porter (1990) developed a model comprising four major determinants of competitive advantage: "factor conditions, demand conditions, related and supportive industries, and firm strategy, structure, and rivalry." At this juncture, as noted by Wagar & Rondeau (2006), human resource practices, such as education, play a pivotal role in empowering employees. As Boxall (2003) emphasizes, high-performance work systems (HPWSs)-which include human resource practices (HRP) such as employee empowerment through education-play a pivotal role in achieving enhanced company performance and worker well-being. Consequently, Wagar and Rondeau (2006) have indicated that effective HRP, including investments in employee training, has been observed to result in a reduction in voluntary employee turnover. Therefore, it can be posited that investment in employee training has been demonstrated to be a valuable strategy for firms seeking to enhance the retention of their most productive employees and to reduce voluntary turnover rates. In the context of this discussion, as Romer (1990) has previously indicated, the acquisition of knowledge is not only essential for economic growth but also for enhancing regional competitiveness. Given these facts, the focus on acquiring knowledge through educational efforts becomes of paramount importance.

According to Tharenou et al. (1994), the acquisition of human capital—defined as knowledge and skills obtained through educational institutions—is associated with increased productivity, which in turn is often accompanied by higher status and earnings. Similarly, Francesconi (2001) found that individuals with higher levels of education tend to receive a greater number of career outcomes, such as higher income, promotions, and job satisfaction, compared to those with fewer educational opportunities. De Gregorio and Lee (2002) confirmed the association between educational attainment and income distribution, thereby suggesting that education plays a pivotal role in shaping income distribution. Accordingly, Bird and Smith (2005) indicated that individuals with higher education are more likely to be hired at higher wages.

Regarding the return of additional schooling, private return has been predominantly observed in comparison to social return. Social return can be defined as externalities and spillover effects on the environment, such as decreased crime rates, increased tax revenues, and attraction to the environment due to the presence of high-quality human capital. Nevertheless, the social return effect is challenging to evaluate. Therefore, researchers seemed to have attempted to calculate the return on additional years of schooling (Arik & Dunne, 2014). In that context, Lin and Vogt (1996) investigated whether a two-year college education affects the occupational status and earnings gap between some more and less advantaged groups in the U.S., involving 1542 two-year college graduates and 2167 high school graduates. They found that individuals' earnings and/or employment status improved as a result of attending a two-year college and earning a certificate or degree. Arik and Penn (2007) investigated the relationship between an additional year of schooling and the income earned by college-educated workforces. They revealed that a 1% increase in the college-educated workforce resulted in a 9.25% increase in per capita income. Another study by Arik and Dunne (2014), conducted in Tennessee counties, is a notable example of this ongoing research. The findings of this study indicated that an additional year of schooling was associated with an 11.40% increase in income among workers, thereby confirming the conclusions of previous studies. A notable finding in this study is the observed increasing trend in the return to higher educational attainment. In other words, as educational attainment levels increase, the return tends to increase accordingly.

Based on these findings, our first hypothesis is:

Hypothesis 1*: Pursuing additional years of schooling positively affects individual income.*

Given that specific industries demand a distinct array of skills and knowledge, it can be inferred that the gains in various industries may vary due to their unique characteristics (Ahmad

& Schroeder, 2003). Arik and Dunne (2014) demonstrated that the association between additional years of schooling and earned income varies by industry. The effects of additional years of schooling on individual income would be more positive in industries where the percentages of bachelor's degree graduates are high.

In consideration of these findings, we developed the following hypothesis:

Hypothesis 2*: The effect of additional years of schooling on individual income varies according to industry.*

3. Data

This study combines nationally representative time-series cross-sectional microdata from the Integrated Public Use Microdata Series-IPUMS USA (Ruggles et al., 2025) with state-level panel data from the National Association of State Student Grant and Aid Programs (NASSGAP, 2024). The individual-level data are derived from the IPUMS USA database, a harmonized version of the U.S. Census and American Community Survey microdata curated by the University of Minnesota (Ruggles et al., 2025). The extract spans a survey from 2005 to 2023 and was trimmed to include only prime-age adults between 25 and 64 years old. The sample from 2005 to 2023 was chosen to maintain consistency in the IPUMS USA ACS data, as any year prior to 2005 did not include people in group quarters. Additionally, the sample was restricted to individuals between 25 and 65 years old to remove any outliers, which would otherwise bias the estimates of this study. Our final sample comprises approximately 20.5 million person-years of observations across all 50 states.

To ensure data quality and comparability, we imposed several restrictions on the IPUMS sample. We limited the analysis to prime working-age adults (ages 25–64) who were employed at the time of the survey and who reported positive wage and yearly salary income (incwage). Individuals in the sample must have reported working at least one week and one hour per week (wkswork2, uhrswork). Observations with missing, invalid, top-coded, or outlier values greater than $500,000 in yearly wages were excluded. Additionally, we restricted the sample to those with valid education information (educ & educd) and complete demographic identifiers (age, sex, race, hispan). Any observations with missing educational or demographic identifiers were dropped.

To account for the impact of price levels on earnings over our 2005-2023 sample period, we deflated nominal wages using the Consumer Price Index for All Urban Consumers (CPI-U) Retroactive Series, collected from the Bureau of Labor Statistics (Bureau of Labor Statistics, 2025). The CPI-U-RS provides consistent inflation adjustments across the entire time series and includes all items at the U.S. city average level. We rebased the CPI-U-RS index to 2019 (2019=100) to express all wage values in constant 2019 dollars, thereby eliminating the confounding effects of inflation on our estimates of returns to education. This adjustment ensures that observed wage changes reflect real differences in purchasing power rather than nominal price increases over time. Any observations within the sample with missing educational or demographic identifiers were dropped. Summary statistics for the continuous and categorical variables are found in Tables 1 and 2.

Table 1: Summary Statistics of Continuous Variables

Variable	Logged Real Wages	Real Wages	Years of Schooling	Experience	Experience Squared	Hours worked per Week
N	20,526,419	20,526,419	20,526,419	20,526,419	20,526,419	20,526,419
Mean	10.62	59,290.897	14.00	24.20	717.48	40.90
Std. Dev.	0.95	57,999.26	2.77	11.49	568.95	10.75
Min	1.21	3.35	0.00	0.00	0.00	1.00
Median	10.72	45,316.11	13.00	24.00	576.00	40.00
Max	13.36	635,081.06	20.00	58.00	3364.00	98.00

Table 2: Distribution of Categorical Variables

Variables	Frequency	Percent	Cumulative Percent
Male	10,562,246	51.46	51.46
Female	9,964,173	48.54	100
White	15,783,739	76.89	76.89
Non-White	4,742,680	23.11	100
Hispanic	2,584,095	12.59	12.59
Non-Hispanic	17,942,324	87.41	100
Metro	15,651,834	76.25	76.25
Non-Metro	4,874,585	23.75	100

It is important to note that after trimming the outliers, the annual wage income is right-skewed: the mean is \$59,290.897 while the median is \$45,316.11. The average schooling is 14.00 years, with a standard deviation of 2.77, indicating that individuals in the sample are roughly 2 years beyond high school on average in their education, which is consistent with the sample consisting of working-age adults only. Individuals in the sample report an average of 24.20 years of experience, with a standard deviation of 11.49 years, which is also in line with the sample limited to only working-age individuals. The weekly hours worked within the sample cluster are approximately 40.90, indicating that the average individual in the sample works full-time. Our categorical variables appear to be balanced between gender (51.46% male; 48.54% female), while predominantly White (76.89%), with 12.59% identifying as Hispanic. Most individuals in the sample reside in metropolitan areas (76.25%).

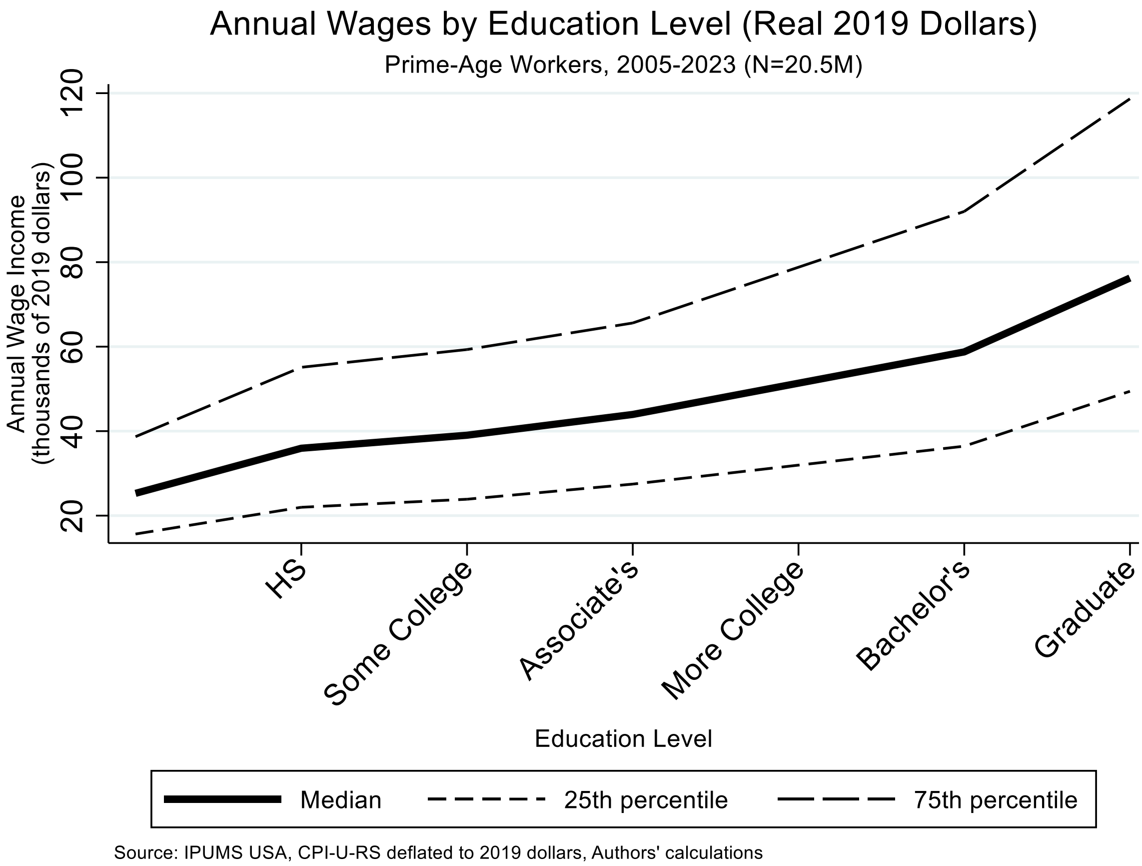

Figure 1. Annual Wages by Educational Level

Within our sample, we find that median annual wages rise monotonically with education for prime-age workers during the sample period of 2005–2023. The median increases from roughly \$30k for high school graduates to about \$70k in 2019 dollars for graduate degrees, with steady gains at "some college," "associates", and "more college", and a clear step up once the individuals in the sample obtain a bachelor's degree. The 25th and 75th percentiles follow the same pattern, but the variance increases with higher educational attainment, albeit skewed upward. The widening interquartile range indicates greater wage dispersion among more educated workers, consistent with higher returns and increased heterogeneity associated with advanced degrees.

To estimate a linear effect of education on wages, we constructed a continuous education variable (yrs_school) from the discrete educational attainment categories provided by IPUMS (educ & educd). IPUMS USA provides two measures of educational attainment with the general measure (educ) that classifies respondents into broad categories such as “high school graduate,” “two years of college,” “five or more years of college,” or “bachelor’s degree” while the more detailed measure (educd) offers finer distinctions “Grade 12” or “1 year of college.” Unfortunately, many observations in the IPUMS dataset lack a value associated with the educd variable, and it does not contain a variable for continuous years of schooling. Because IPUMS USA does not supply a variable for continuous years of schooling, it was necessary to translate the categorical responses of educ and educd into a linear measure suitable for regression analysis. We mapped each grade directly to its expected year of schooling whenever detailed codes were available. For example, respondents who reported completing grades 5 or 6 were assigned 5 or 6 years of schooling, respectively. Given the limitations of the data, some assumptions had to be made when assigning a numerical value to potentially overlapping categories. For instance, Individuals in the sample labeled as “12th grade, no diploma”, “High school graduate or GED”, “Regular high school diploma”, and “GED or alternative credential” were all coded as having 12 years of education, since the distinction between the categories is unclear.

In further cases where educational attainment was not explicitly stated, we turned to the broader educ variable and made additional assumptions about grouping educational attainment levels into ranges. For these grouped categories, we assigned the average value of education. For instance, we assigned 2.5 years for individuals in the sample labeled as “Nursery school to grade 4” and 6.5 years for those labeled as “Grade 5, 6, 7, or 8”, to approximate the individuals' education. This approach preserved the representativeness of the sample but inevitably introduced some measurement errors by compressing variation within the grouped categories. This issue extends to higher education when a member of a sample is labeled as having “1 or more years of college credit, no degree” and “5+ years of college,” since the precise number of years completed was unknown. We coded these as 13 and 18 years of education, respectively.

Because our research question focuses on the impact of postsecondary schooling on wage income, rather than primary or secondary education, we further decomposed total schooling into two components: years up to high school completion (yrs_school_to_hs) and years beyond high school (yrs_highered). This specification allowed us to separately identify the marginal effect of higher education. Although our coding conventions carry limitations, especially in the treatment of ambiguous categories, they represent the best feasible harmonization of IPUMS data into a linear years-of-schooling measure suitable for econometric analysis.

Table 3: Educational Attainment Variable - educd

Educational Attainment (educd)	Frequency	Percent	Cumulative Percent
Nursery school to grade 4	12,474	0.06%	0.06%
Nursery school, preschool	2,620	0.01%	0.07%
Kindergarten	2,412	0.01%	0.09%
Grade 1	5,018	0.02%	0.11%
Grade 2	10,372	0.05%	0.16%
Grade 3	21,721	0.11%	0.27%
Grade 4	16,264	0.08%	0.35%
Grade 5 or 6	33,569	0.16%	0.51%
Grade 5	24,301	0.12%	0.63%
Grade 6	120,666	0.59%	1.22%
Grade 7 or 8	29,811	0.15%	1.36%
Grade 7	30,501	0.15%	1.51%
Grade 8	87,426	0.43%	1.93%
Grade 9	178,644	0.87%	2.81%
Grade 10	204,257	1.00%	3.80%
Grade 11	246,421	1.20%	5.00%
12th grade, no diploma	291,357	1.42%	6.42%
High school graduate or GED	845,499	4.12%	10.54%
Regular high school diploma	3,473,742	16.92%	27.46%
GED or alternative credential	582,964	2.84%	30.30%
Some college, but less than 1 year	1,379,828	6.72%	37.02%
One or more years of college credit, no degree	2,888,219	14.07%	51.10%
Associate's degree, type not specified	2,040,562	9.94%	61.04%
Bachelor's degree	4,901,132	23.88%	84.91%
Master's degree	2,220,200	10.82%	95.73%
Professional degree beyond a bachelor's	520,927	2.54%	98.27%
Doctoral degree	355,512	1.73%	100.00%

Table 4: Educational Attainment Variable - educ

Educational Attainment (educ)	Frequency	Percent	Cumulative Percent
Nursery school to grade 4	70,881	0.35%	0.35%
Grade 5, 6, 7, or 8	326,274	1.59%	1.93%
Grade 9	178,644	0.87%	2.81%
Grade 10	204,257	1.00%	3.80%
Grade 11	246,421	1.20%	5.00%
Grade 12	6,573,390	32.02%	37.02%
1 year of college	2,888,219	14.07%	51.10%
2 years of college	2,040,562	9.94%	61.04%
4 years of college	4,901,132	23.88%	84.91%
5+ years of college	3,096,639	15.09%	100.00%

Table 5: Educational Attainment Variable - yrs_school

Years of Schooling (yrs_school)	Frequency	Percent	Cumulative Percent
0	17,506	0.0009	0.0009
1	5,018	0.02%	0.11%
2	10,372	0.05%	0.16%
2.5	63,380	0.31%	0.47%
3	21,721	0.11%	0.57%
4	16,264	0.08%	0.65%
5	24,301	0.12%	0.77%
6	120,666	0.59%	1.36%
7	30,501	0.15%	1.51%
8	87,426	0.43%	1.93%
9	178,644	0.87%	2.81%
10	204,257	1.00%	3.80%
11	246,421	1.20%	5.00%
12	8,081,781	39.37%	44.37%
13	1,379,828	6.72%	51.10%
14	2,040,562	9.94%	61.04%
16	4,901,132	23.88%	84.91%
18	2,220,200	10.82%	95.73%
19	520,927	2.54%	98.27%
20	355,512	1.73%	100.00%

Most workers within our sample bunch are at two levels of education: high school (12 years) and bachelor's (16 years). In Table 5, individuals with 12 years of education account for 39.37% of the sample, and those with 16 years of education account for 23.88%, together totaling 63.25% of the sample. Those with graduate degrees skew the distribution towards higher education, with 18, 19, and 20 years of schooling totaling 15.09% of the distribution. The bridge between a high school degree and a bachelor's degree is nontrivial. Those with 13 years of education (some college) are 6.72% and those with 14 years (two years/associate's) are 9.94%.

In addition to the IPUMS USA CPS microdata, we include state-level financial aid data derived from the Annual Survey of State-Sponsored Student Financial Aid, conducted each year by the National Association of State Student Grant & Aid Programs (NASSGAP). These data provide coverage of state-funded financial aid expenditures for undergraduate higher education from 2004 to 2023. Financial Aid categories within the sample include need-based grants, merit-based grants, combined need-and-merit programs, special-purpose awards, and programs not otherwise classified. For the purposes of this study, we focus on undergraduate aid expenditures, aggregated by state and year, and then construct measures of total need-based and merit-based aid. These data capture the policy environment that shapes the affordability of higher education and the incentives of students to attend university across states and over time.

The instrumental variable used in our analysis is a Bartik-style, or shift-share, instrument designed to isolate exogenous variation in state-level financial aid policy. Following the logic of Bartik (1991), the instrument interacts with a state's historical reliance on need-based financial aid, calculated from a stable base period (2004-2006) to serve as the local "share," and the cumulative national growth in need-based aid over time, which serves as the aggregate "shift." Similar to traditional Bartik measures that use historical industry mixes to predict local economic growth, this instrument predicts a state's financial aid environment. This predicted environment is then matched to each individual based on their state and the year they turned eighteen, thereby capturing the policy context relevant to their initial college enrollment decision.

To assess the long-term impact of financial aid policies on educational attainment, we merged our individual-level IPUMS data with the state-year panel of financial aid. We avoided a simple merge based on the survey year. This is because individuals beyond the typical college enrollment years (ages 25 and up) are likely not impacted by financial aid in the same way as those aged 18 to 24. To account for this, we implemented a timing correction to link each individual to the financial aid environment when the individuals in the sample were aged 18 to 24. For each person in the sample, we calculated a "college-entry year" by adding 18 years to their birth year. This calculated year, along with their state of residence, was then used as the key to merge each individual with the state-level financial aid data from that specific historical period. This procedure ensures that an individual's total educational outcome, observed later in life, is matched with the financial aid policy environment that was relevant during their formative college decision-making period.

A key limitation of our Bartik instrumental variable approach, however, is the lack of panel data at the individual level. We instead create a time lag between the instrument's measurement and the observation of the educational outcome later in life for individuals in the sample. This lag makes our instrument potentially susceptible to long-term divergent economic trends between states over time, as the initial aid policy "share" may be correlated with other persistent socio-economic characteristics at the state level. Consequently, there is a risk that the instrument does not fully satisfy the exclusion restriction, as it may be capturing the cumulative effects of these long-run state-specific trends rather than the sole causal effect of financial aid.

We constructed Bartik-style instruments using the following equation (1).

$$B_{it} = \sum_{k\in\{need,merit\}} s_i^{aid} * \Delta A_t^{aid} \tag{1}$$

Where s_i^{aid} denote the average share of financial aid (need and merit-based) expenditures in state i during the chosen base period of (2004–2006), and ΔA_t^{aid} denotes the national change in total expenditures on financial aid in year t relative to the base period. In practice, our instrument captures the interaction between a state's level of fixed historical undergraduate aid programs and exogenous changes in overall aid generosity at the national level. The share term, s_i^{aid} is time-invariant and therefore not affected by contemporary shocks to state economies or wages. The shift term ΔA_t^{aid} reflects aggregate changes in aid across all states, driven by national fiscal conditions or policy, and is assumed to be exogenous to any single state.

Our initial model combines merit- and need-based financial aid, and as a robustness check, we separate the instrument into need- and merit-based financial aid individually. This allows us to avoid a model where the instrument is exactly identified with the potentially endogenous variable. We justified this choice by arguing that need-based and merit-based aid respond to different policy drivers and therefore provide distinct sources of exogenous variation. Testing the instruments separately enables us to statistically validate the instruments using a Sargan-Hansen test (Sargan, 1958; Hansen, 1982). This disaggregated approach allows us to verify that our findings are not driven by any single type of financial aid.

Table 6: Bartik IV Summary Statistics

Variable	Observation	Mean	Min	Median	Max	Std. Dev.	Variance	Skewness	Kurtosis
Need-Based	20,526,419	0.026	0	0	0.859	0.098	0.010	4.553	25.363
Merit-Based	20,526,419	0.019	0	0	1.201	0.097	0.009	6.688	53.783
Combined	20,526,419	0.045	0	0	2.008	0.172	0.030	5.297	37.519

The most prominent characteristic of the Bartik instrument is that each instrument has a value of zero for at least 75% of the sample. This is an expected feature of the shift-share design, as the instrument is only "active" for individuals in states with a non-zero historical aid share and in years with positive national aid growth. Despite the prevalence of zeros, the instruments exhibit substantial variation, with a standard deviation of 0.172 for the combined instrument and a maximum value of over 2.0. This confirms there is sufficient variation for identification. The distribution is also right-skewed, with a mean of 0.045, which is considerably larger than the median of zero. This indicates that the instrument's variation is driven by a minority of observations with large, positive values.

4. Methodology

This study is reliant upon the basis of Mincer's (1974) model. The following equation (2) was established to test Hypothesis 1 and Hypothesis 2.

$$lnY = \alpha + \beta_1 S_i + \beta_2 AGE + \beta_3 AGE^2 + \varepsilon \tag{2}$$

In this equation, $\boldsymbol{lnY}$ denotes the natural log of the individual's earnings, $\boldsymbol{S_i}$ indicates the number of years of schooling for the specified individuals, ***AGE*** and ***AGE²*** represent the person's experience and the square of that experience, respectively. To address variation in earnings due to the number of work hours, we added the number of hours worked as a variable. This resulted in the formulation of the following equation (3):

$$lnY = \alpha + \beta_1 S + \beta_2 MAGE + \beta_3 MAGE^2 + \beta_4 HW + \varepsilon \tag{3}$$

where ***HW*** is the number of hours worked per week.

To account for the marginal returns to specific school attainment levels, regional characteristics, and industrial differences, Model 2 is expanded to include dummy variables for the individual's race, sex, ethnicity, and whether they reside in a metropolitan location.

We expand on the Mincer model by incorporating two-stage least squares to obtain causal estimates of the effect of education on wage income. The second stage of this approach can be expressed as:

$$lnY = \alpha + \beta_1 \hat{S}_{it} + \beta_2 Z'_{it} + \lambda_s + \tau_t + u_{it} \tag{4}$$

where lnY is still wage income, $\hat{S}_{it}$ represents the predicted value of schooling from the first stage of the model, Z'_{it} represents a vector of exogenous control variables which include kindergarten through Highschool education, experience, experience squared, the number of hours worked by the individual per week, sex, race, and metropolitan status. λ_s and τ_t represent state and year fixed effects while u_i is our error term.

To determine $\hat{S}_{it}$ the first stage of the modified Mincer equation can be expressed as:

$$\hat{S}_i = \pi_0 + \pi_1 B_{it} + \pi_2 Z'_{it} + \lambda_s + \tau_t + v_{it} \tag{5}$$

The variables in this model are equivalent to model 3, with the exception of B_{it} which measures the value of the combined Bartik variable, or the separated Bartik variables for need-based and merit-based financial aid.

5. Results

To argue for the causal effects of our first hypothesis, we must first verify that the Bartik instrumental variable is a relevant instrument. We estimated the first-stage regression of years of higher education on the Bartik instrumental variable, using the variables of years of K–12 schooling, experience and experience squared, weekly hours worked, and indicators for gender, race, ethnicity, and metropolitan residence of the individual in the sample. To ensure robustness and eliminate variation caused by time and location, we also include state and year fixed effects, while clustering the standard errors by state. Given that our main model assumes only one endogenous regressor and one instrument, that model is exactly identified. To address concerns about weak instruments under heteroskedasticity and clustering, we computed the Kleibergen–Paap rank Wald F statistic and compared it to the Stock–Yogo 10% critical value for the maximal IV size. Additionally, for our second model, which categorizes the instrument by need and merit-based financial aid, we perform a Sargan-Hansen test to determine the validity, as our model is over-identified.

Table 7: First Stage Mincer Equation Results Using Combined Bartik IV

yrs_highered	Coefficients	Robust Std. Err.	t	P>t	[95% Conf.	Interval]
Bartik_IV	-1.8938***	0.386	-4.91	0.000	-2.650	-1.137
yrs_school_to_hs	0.1796***	0.018	10.10	0.000	0.145	0.214
experience	-0.1112***	0.004	-25.24	0.000	-0.120	-0.103
experience_sq	0.0011***	0.000	15.75	0.000	0.001	0.001
uhrswork	0.0213***	0.001	24.72	0.000	0.020	0.023
female	0.3374***	0.017	20.33	0.000	0.305	0.370
nonwhite	-0.3062***	0.053	-5.75	0.000	-0.411	-0.202
hispanic	-1.0269***	0.150	-6.86	0.000	-1.320	-0.733
metro_area	0.6645***	0.029	22.70	0.000	0.607	0.722
Observations			20,526,419			
Kleibergen–Paap rk LM			10.78			
Kleibergen–Paap rk Wald F			24.08			
Number of clusters			50			
Year FE			Yes			
State FE			Yes			

Notes: Standard errors clustered by State
Hansen J Statistic not included due to the model being exactly identified
Demographic & Geographic references are Male, White, Non-Hispanic, Non-Metro
All regressions use data and population weights from IPUMS USA. *** p<0.001, ** p<0.01, * p<0.05.

Table 8: First Stage Mincer Equation Results Using Separated Bartik IV

yrs_highered	Coefficients	Robust Std. Err.	t	P>t	[95% Conf. Interval]	
Bartik_IV_Need	-2.8306***	0.233	-12.14	0.000	-3.287	-2.374
Bartik_IV_Merit	-1.0739	0.559	-1.92	0.055	-2.169	0.022
yrs_school_to_hs	0.1786***	0.018	9.79	0.000	0.143	0.214
experience	-0.1163***	0.004	-28.87	0.000	-0.124	-0.108
experience_sq	0.0012***	0.000	18.74	0.000	0.001	0.001
uhrswork	0.0213***	0.001	24.85	0.000	0.020	0.023
female	0.3368***	0.167	20.21	0.000	0.304	0.370
nonwhite	-0.3071***	0.053	-5.82	0.000	-0.411	-0.204
hispanic	-1.0263***	0.150	-6.84	0.000	-1.320	-0.732
metro_area	0.6642***	0.029	22.82	0.000	0.607	0.721
Observations		20,526,419				
Kleibergen–Paap rk LM		15.74				
Kleibergen–Paap rk Wald F		75.80				
Number of clusters		50				
Year FE		Yes				
State FE		Yes				

Notes: Standard errors clustered by State
Demographic & Geographic references are Male, White, Non-Hispanic, Non-Metro
All regressions use data and population weights from IPUMS USA. *** p<0.001, ** p<0.01, * p<0.05.

The results for Table 7 present the first stage of our two-stage least squares specification, which incorporates the Bartik instrument that combines both need- and merit-based aid. Given that the model includes one instrument and one presumed endogenous variable, we cannot statistically test the exclusion restriction through a Sargan-Hansen test. Rather, we justify the validity of the instrument by arguing that the financial aid policy is exogenous to new students. We assume that the financial aid policy is correlated with students' engagement in higher education and is directly unrelated to the students' wage income. The Bartik instrumental variable scales common national shifts in state-funded undergraduate aid by pre-period state aid shares and is estimated with the previously mentioned, assumed exogenous controls for experience, gender, race, ethnicity, and metropolitan status, along with state and year fixed effects. Under this design, and given the first-stage results, if the exclusion restriction holds for the instrument, then the Bartik instrumental variable functions as a valid instrument for identifying the causal effect of postsecondary schooling on wages. The results for Table 8 also show the first stage of our two-stage least squares specification, where we separate financial aid into need-based and merit-based components. Since there are two instruments, we can statistically determine the validity of the instruments using a Sargan-Hansen test.

The first-stage regression results, presented in Table 7, confirm the strength and relevance of the Bartik instrumental variable in predicting years of higher education. The Kleibergen–Paap Wald F statistic of 24.08 exceeds the Stock–Yogo critical value of 16.38, indicating that the instrument is a strong predictor and not susceptible to weak instrument bias. Notably, the coefficient on the Bartik instrument is -1.89 and highly significant, which indicates a negative relationship between the predicted financial aid and educational attainment. The model's control variables perform largely as expected, with a one-year increase in K-12 schooling associated with a 0.18-year increase in postsecondary education, suggesting that educational attainment prior to college is correlated with more postsecondary education. The coefficients for experience reveal a parabolic relationship, where each additional year of work is initially associated with a 0.11-year decrease in higher education, which increases over time. Residing in a metropolitan area is associated with a 0.66-year increase in higher education, likely reflecting greater access to postsecondary institutions. The model also shows that females, on average, attain 0.34 more years of higher education than males, while nonwhite and Hispanic individuals attain 0.31 and 1.03 fewer years, respectively, relative to the white, non-Hispanic reference group.

Table 8 presents the results for the over-identified first-stage model, where the Bartik instrument is disaggregated into need-based and merit-based financial aid to examine the distinct influence of each type of aid. The coefficient for the need-based instrument, Bartik_IV_Need, is -2.83 and is statistically significant, indicating that a one-unit increase in the predicted shock to need-based aid is associated with a 2.83-year decrease in higher education. Alternatively, the coefficient for the merit-based instrument, Bartik_IV_Merit, is smaller in magnitude at -1.07 and is not statistically significant at the 95% confidence interval. These results indicate that the negative association observed in the combined model is driven predominantly by the need-based financial aid instrument, while the merit-based component has a weaker and statistically insignificant effect.

Table 9: Second Stage Mincer Equation Results Using Combined Bartik IV

ln_real_wage	Coefficients	Robust Std. Err.	t	P>t	[95% Conf. Interval]	
yrs_highered	0.1389***	0.003	40.16	0.000	0.132	0.146
yrs_school_to_hs	0.0508***	0.002	31.07	0.000	0.048	0.054
experience	0.0353***	0.000	71.66	0.000	0.034	0.036
experience_sq	-0.0005***	0.000	-63.28	0.000	-0.001	0.000
uhrswork	0.037***	0.000	77.57	0.000	0.036	0.038
female	-0.2308***	0.007	-32.14	0.000	-0.245	-0.216
nonwhite	-0.1114***	0.010	-11.50	0.000	-0.131	-0.092
hispanic	-0.1063***	0.009	-11.28	0.000	-0.125	-0.087
metro_area	0.1215***	0.007	17.34	0.000	0.107	0.136
Observations			20,526,419			
Centered R-squared			0.3651			
Uncentered R-Squared			0.3651			
Number of Clusters			50			
Year FE			Yes			
State FE			Yes			

Notes: Standard errors clustered by State
Demographic & Geographic references are Male, White, Non-Hispanic, Non-Metro
All regressions use data and population weights from IPUMS USA. *** p<0.001, ** p<0.01, * p<0.05.

Table 10: Second Stage Mincer Equation Results Using Separated Bartik IV

ln_real_wage	Coefficients	Robust Std. Err.	t	P>t	[95% Conf. Interval]	
yrs_highered	0.1370***	0.004	35.66	0.000	0.129	0.145
yrs_school_to_hs	0.0512***	0.002	31.45	0.000	0.048	0.054
experience	0.0352***	0.001	69.94	0.000	0.034	0.036
experience_sq	-0.0005***	0.000	-62.73	0.000	-0.001	-0.000
uhrswork	0.0370***	0.000	77.26	0.000	0.036	0.038
female	-0.2302***	0.007	-32.33	0.000	-0.244	-0.216
nonwhite	-0.1120***	0.010	-11.33	0.000	-0.132	-0.092
hispanic	-0.1083***	0.009	-11.93	0.000	-0.127	-0.090
metro_area	0.1228***	0.007	17.51	0.000	0.109	0.137
Observations			20,526,419			
Centered R-squared			0.3651			
Uncentered R-Squared			0.3651			
Hansen J Statistic			2.769			
Number of Clusters			50			
Year FE			Yes			
State FE			Yes			

Notes: Standard errors clustered by State
Demographic & Geographic references are Male, White, Non-Hispanic, Non-Metro
All regressions use data and population weights from IPUMS USA. *** p<0.001, ** p<0.01, * p<0.05.

Table 9 presents the main finding of this paper for hypothesis 1, the second-stage estimate of the causal return to schooling from our primary model. The coefficient on yrs_highered is 0.1389 and is statistically significant. This indicates that each additional year of higher education is estimated to increase an individual's wage and salary income by 13.9%. This estimate can be interpreted causally if our instrument is deemed to be strong and valid, as indicated by the results in Table 10. The disaggregated, over-identified specification model yields a nearly identical return to schooling of 0.1370, demonstrating that the finding is not sensitive to the construction of the instrument. Crucially, the second model allows for a formal test of the exclusion restriction with the Hansen J statistic of 2.769, which fails to reject the null hypothesis that the instruments are valid, providing evidence that our instrument is valid. Therefore, the combination of a consistent and significant coefficient of yrs_highered across the models, along with the first-stage results indicating that the instruments are not weak, and the Sargan-Hansen tests provide evidence that an additional year of higher education causes a substantial wage increase of approximately 13.7% to 13.9%.

Table 11: Industry-Specific Returns to Higher Education

Industry	Coef (IV)	Std. Err (IV)	Coef (OLS)	F-stat	Observations
Agriculture	0.1701***	0.0159	0.1029	13.38	229,131
Mining	0.1673***	0.0137	0.1402	111.16	119,251
Transportation	0.1620***	0.0036	0.1247	24.84	604,757
Personal Services	0.1572***	0.0052	0.1125	25.80	1,451,521
Trade	0.1492***	0.0058	0.0957	21.97	1,955,702
Health	0.1379***	0.0091	0.1184	20.76	732,229
Education	0.1316***	0.0087	0.1144	26.37	5,201,545
Manufacturing	0.1296***	0.0038	0.1386	30.40	3,725,801
Construction	0.1290***	0.0073	0.1021	25.28	224,026
Finance	0.1157***	0.0097	0.0915	25.12	926,689
Professional Services	0.1150***	0.0182	0.0928	25.40	463,803
Entertainment	0.1149***	0.0095	0.0983	23.53	1,511,355
Other	0.1120***	0.0039	0.0908	23.56	1,320,073
Public Administration	0.0968***	0.0079	0.0723	21.86	808,523
Social Services	0.0944***	0.0058	0.0806	21.62	1,252,013

Each row represents a separate regression of log wages on years of higher education for workers in that industry.
All regressions include the same controls as Model 3 in Tables 7 through 10, with year and state fixed effects.
Standard errors clustered by state.
All OLS coefficients were statistically significant at the 0.001 level
All regressions use data and population weights from IPUMS USA. *** p<0.001, ** p<0.01, * p<0.05.

We test the second hypothesis that the returns to schooling vary by industry by estimating 15 separate 2SLS models for each industry category in the IPUMS database. The results of are presented in Table 11. For these regressions, we employ the same controls as in previous models, and we utilize the combined Bartik instrument as the sole instrument. This means that for these results, the models are just identified. The results indicate a positive, statistically significant, causal relationship between higher education and each of the 15 industries. The findings suggest that the economic benefits of postsecondary schooling are robust and translate into higher wages across all industry sectors. All first-stage F-statistics for these models exceed the critical threshold, confirming the strength of the instrument within each subsample.

The estimated wage increases from an additional year of higher education range from a low of 9.4% in Social Services and 9.7% in Public Administration to a high of 16.7% in Mining and 17.0% in Agriculture. The range between industries confirms that the economic value of an additional year of education is not uniform, but rather contingent on the industry of employment. While the results confirm that returns vary, they do not perfectly align with the prediction that returns would be strongest in industries with the highest concentration of bachelor's degrees. For instance, sectors such as Finance (11.6%) and Professional Services (11.5%), which are highly demanding of bachelor’s degree holders, show relatively modest returns relative to the highest returns found in sectors like Agriculture (17.0%), Mining (16.7%), and Personal Services (15.7%). This suggests that the relationship may be more complex and potentially reflects a scarcity premium, where a degree provides a larger marginal benefit in industries where college-educated labor is less common.

6. Discussion

A recent Gallup poll found that only 35% of U.S. adults consider a college education “Very important”, 24% claim that college is “Not too important”, and 40% of the population finds education “Fairly important”, which is a decline and increase from roughly 75, 21, and 4 percent, respectively (Gallup, 2025). Among 18 to 34-year-olds, those who believe education is “Very important” declined from 74% in 2013 to 35% in 2025, which is particularly concerning as, in theory, they have the most to gain from increased wages over time. Additionally, a Pew Research Center report found that only 22% of adults believe a four-year degree is worth the cost if it requires student debt (Pew Research Center, 2024).

Our findings provide robust causal evidence that higher education remains a powerful driver of individual economic welfare, despite the increasing perception that it has lost value. Our central finding of this paper, that an additional year of higher education causally increases an individual's wages by approximately 13.9%, challenges the notion that a college education is losing its value. A return of 13.9% is substantial and continues to be evidence that investing in human capital through formal schooling provides a significant competitive advantage in the labor market. Our findings are also consistent with the 5% to 20% range established in previous economic literature (Becker, 1964; Mincer, 1974; Psacharopoulos, 1994). It surpasses the 9.25% return (Arik & Penn, 2007) and the 11.40% return (Arik & Dunne, 2014), suggesting that the economic premium for postsecondary education remains a consistent driver of wage premiums for workers, rather than diminishing.

Furthermore, our analysis of industry-specific returns adds nuance to our findings. While a wage premium for higher education was found in every industry studied, the variation in returns is revealing. The finding that some of the highest returns are in sectors like Agriculture (17.0%) and Mining (16.7%) suggests a "scarcity premium" in fields where a college degree is less common. This outcome is inconsistent with the industry-specific findings of Arik and Dunne (2014), who found that the positive impact of additional years of schooling on individual income is more pronounced in industries with a high percentage of bachelor's degree holders. This research has demonstrated that additional years of schooling have the greatest impact on the following fields, respectively: education (13.7%), financial services (14.8%), healthcare

(14.9%), utilities (15.2%), manufacturing (15.4%), and professional services (17%). Given the results obtained, it can be posited that a college degree in these fields serves as a valuable signal to employers, leading to high-paying roles. Conversely, in sectors like Finance (11.6%) and Professional Services (11.5%), where a degree is often a baseline requirement, there is less of a scarcity premium and thus less value. While the marginal return for an additional year may seem more modest in these fields, the degree itself offers higher lifetime earnings than not having a degree.

6.1. Theoretical Implications

What are the theoretical implications of these findings? First, the present study directly combines the classic "returns to an additional year of education" framework of education economics (Mincer, 1974; Psacharopoulos, 1994) with the resource-based view (Barney, 1991).

Second, the finding that each additional year of higher education increases wages by approximately 13.7%–13.9% provides not only conceptual but also quantitative support for the firm's view that human capital is a " valuable, rare, imperfectly imitable, and non-substitutable" resource (Barney, 1991). Consequently, human capital is no longer considered a "qualitative" element, as previously discussed within the RBT, instead, it is now regarded as a distinct economic value.

Third, the strong validation of Becker's human capital theory (1964, 1975) and Mincer's income function approach (1974) using US data from 2005 to 2023 demonstrates the theory's continued relevance in contemporary labor markets. The returns obtained are close to the upper limit of the 5%–20% range found by Psacharopoulos (1994) on a global scale, providing new evidence that reduces the debate between the "declining returns to higher education" narrative and the theoretical literature.

Fourth, by clearly measuring the impact of education on individual wages, the study provides a micro-based basis for the human capital-driven growth models of Romer (1989, 1990) and Mankiw, Romer, & Weil (1992). This provides data to support the "education-human capital-productivity-wage-economic welfare" chain, connecting the micro and macro literatures.

Fifth, the returns, which are positive across all 15 sectors, ranging from 9.4% to 17%, indicate that returns to education are not a single effect but vary depending on the sector's technological level, skill mix, and supply-demand balance.

Sixth, the higher returns in sectors with relatively low rates of higher education, such as agriculture (17.0%) and mining (16.7%), imply that human capital is a “scarcer” resource in areas where educated labor is scarce. This highlights the often-overlooked dimension of "context-specific scarcity" in RBT: the same degree manifests different values across sectors.

Finally, despite the common perception that education is not worth the debt, as evidenced by the findings of Gallup (2025) and Pew (2024) surveys, this research consistently demonstrates that the economic returns to higher education remain strong. In other words, the distinction between perceived returns and actual returns is a matter of speculation. These findings underscore the necessity for integration with behavioral economics perspectives.

6.2. Practical Implications

What are some policy implications of the findings of this study? First, the approximately 13.9% wage premium for an additional year of higher schooling rationalizes the allocation of public resources to higher education in terms of economic welfare. This rate is comparable to the return on many physical capital investments, and therefore, education expenditure can be considered an investment, not a consumption.

Second, the high returns to education in agriculture and mining suggest that private scholarships, tax credits, or work-study programs that encourage higher education in these sectors can be a powerful lever for regional development and increased productivity. On the other hand, in relatively low-return sectors, increasing returns through additional skills programs could be enhanced.

Third, these findings can be used to counter the skeptical perspective in higher education, as suggested by Gallup (2025) and Pew (2024) surveys, thereby mitigating irrational withdrawals.

Fourth, universities can enhance the curricula for low-return sectors by offering additional skill sets, such as data analytics, project management, and digital competencies, to improve graduates' wage profiles.

Finally, it is more logical for students to focus on long-term returns rather than short-term costs (debt, opportunity costs) when making educational decisions. Furthermore, it would be sound to consider profitability not only in terms of department selection but also in terms of sector.

6.3. Limitations and directions for future research

While the present study offers substantial evidence regarding the value of higher education, it is essential to acknowledge the limitations as discussed below.

A primary concern for any instrumental variable strategy is the validity of the exclusion restriction. Our Bartik-style instrument is designed to isolate exogenous shifts in financial aid policy, but it may not fully satisfy this crucial assumption. Students aged 17 to 18 have little say over financial aid policy, so from their perspective, the policy appears random; however, states with a history of generous financial aid may also possess other unobserved, long-term characteristics. Those idiosyncratic characteristics may be related to higher taxation from higher incomes within the state, or if a state runs financial aid through a lottery system, that could be a signal of some other economic factors that may affect a student's choice of higher education aside from the financial aid itself. Some states will have different economic trajectories or a stronger commitment to the quality of public universities, which would not be exclusionary to individual wage income. There is a risk that our instrument is capturing the effects of confounding factors, which would impact an individual's wage income rather than the effect of education. Additionally, the negative coefficient for the instrument in the first-stage regression suggests that increases in financial aid are associated with reduced years of schooling for those in our sample, which warrants further investigation.

Our research is also limited by not being a true panel dataset, but rather a repeated cross-section. Although our sample is very large, we cannot follow the same individuals over time, which prevents the use of individual or household-level fixed effects to control for unobserved, time-invariant personal characteristics such as innate ability, motivation, or family background. While we control for a rich set of demographic variables, some of the estimated return to schooling is likely influenced by this unobserved heterogeneity, and our results would suffer from omitted variable bias.

Finally, our analysis is limited by a lack of granularity in both the education and outcome variables. The primary independent variable, "years of schooling," was constructed from categorical data and inevitably introduces some measurement error by being unable to disentangle the actual number of years an individual attended university, but rather makes assumptions about the likely number of years a degree should take to obtain, i.e., a bachelor's degree taking 4 years to complete. We are also unable to differentiate between the returns to a degree in engineering versus one in the humanities. Similarly, while we control for broad industry categories, we cannot control for an individual's specific profession or occupation, network support, and psychometric criteria such as personality traits, which would require eliminating the variation from our data.

The limitations we encountered provide several directions for future research. Studies could use an alternative identification strategy, or a quasi-experiment could validate the causal estimates presented in this research. True longitudinal panel data would enable researchers to implement individual fixed-effects models, allowing them to better isolate the impact of education from innate ability.

6. Conclusion

We employ a 2SLS estimation technique using educational and demographic data from IPUMS USA, as well as a Bartik-inspired instrumental variable approach that utilizes need- and merit-based university financial aid by state. Our analysis draws on a national sample of approximately 20.5 million prime-age workers from the IPUMS USA database, spanning the years 2005 to 2023. Our results indicate that an additional year of post-secondary schooling has a causal effect on an individual's annual wage income, increasing it by 13.9%. Furthermore, the wage premium is positive and statistically significant across all industries, although its magnitude varies between industries. Our findings offer a counter-narrative to recent public skepticism about the value of higher education and reaffirm that investing in human capital through higher education remains a highly rewarding pathway to individual economic welfare.

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