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Are there human capital externalities in U.S. states? Evidence from the Current Population Survey

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Abstract

This paper estimates human capital externalities across U.S. states, using the Current Population Survey (CPS) and state-level data. By directly controlling for individual job characteristics and state labor market conditions, we can identify the human capital externality in an augmented Mincerian model. We find that an extra year of state-level average schooling increases individual wages by five percent above and beyond the private return to education. Subsequent analysis finds that the estimated externality is larger in highly-educated, highly-innovative states. These results imply that the positive coefficient for state-level schooling is in fact an externality and that differences in human capital externalities can help explain "The Great Divergence" in wages between geographic areas with highly-skilled workers versus those with low-skilled workers.

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1. Introduction

Economists have long recognized the possibility of human capital externalities arising as individual workers learn from each other, raising productivity in skilled labor markets (Marshall, 1890). Uncompensated externalities result in individual returns to human capital lower than the total return provided to society. In the aggregate, these externalities can provide an important source of economic growth across cities (Black and Henderson 1999 and Moretti 2011), states (Iranzo and Peri 2009), regions (Yamarik,2010 and Crespo-Cuaresma *et al*. 2012), and nations (Lucas,1988 and Mankiw *et al*.,1992).

Human capital externalities are typically estimated using a Mincerian wage equation augmented with average (city or state) education to capture the external effect. Estimates of the externality (in terms of one more year of average education) range from zero (Rudd, 2000 and Acemoglu and Angrist, 2001) to 3-5 percent (Rauch, 1993) to 25 percent (Moretti, 2004a).[1]

The main empirical issue confronting the augmented Mincerian approach is bias resulting from omitted variables that are correlated with both individual wages and average education. Demand shocks that impact the relative productivity of highly-educated workers and supply shocks that change the relative attractiveness of a region for highly-educated workers will produce a spurious correlation between wages and average education in a state or local labor market. At the individual level, workers in regions with high average education may be more productive than workers with the same observables in regions with low average education, resulting in bias if unobserved worker characteristics are omitted.

The most common strategy to control for this omitted variable bias and enable measurement of the human capital externality is to instrument for average education. Acemoglu and Angrist (2001) use compulsory schooling to instrument for state-level years of schooling, while Moretti (2004a) uses age structure and the presence of a land grant college as instruments for the percentage of college-educated workers in a city. Similarly, Dalmazzo and de Blasio (2007), Liu (2007) and Kirby and Riley (2008) use compulsory schooling and demographic variables to instrument for average schooling in Italian regions, Chinese regions, and English firms, respectively.

However, Lange and Topel (2006) show that these instruments, if not all potential instruments, are likely to be invalid. Building upon a spatial equilibrium model, Lange and Topel demonstrate that an instrument for average education must be orthogonal to productivity differences across regions, unobservable regional human capital, *and* the valuation of local amenities. Additionally, recent work by Stephens and Yang (2014) finds that IV estimation results change considerably when birth-state and birth-year effects are allowed to vary by region.

In this paper, we identify the human capital externality by directly controlling for individual job characteristics and state labor market conditions. We combine the pseudo-panel nature of the Current Population Survey (CPS) with state-level data to produce a rich and comprehensive data set. As opposed to city data, state-level data provide a much broader array of *time-varying*

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[1] Moretti (2004a) used the share of college graduates to measure average education and estimated that a one percent increase in that group raised wages by roughly 1 percent. In his 1990 sample, the average share of college graduates was about 20 percent. Assuming that the average years of schooling in the other 80 percent is 12 years, it takes a change in the college share of 0.25 to increase average years of schooling by one. See Lange and Topel (2006) for details.

controls for labor demand and supply shocks.[2] The combination of state controls and a cohort approach to capturing individual unobserved effects result in estimates of the human capital externality on the order of 5-10%, in line with other studies that incorporate limited geographic controls in their models (Rauch 1993; Rudd 2000; Moretti 2004).

The remainder of the paper proceeds as follows. Section 2 describes the empirical design. Data and estimation techniques are presented in section 3. Section 4 provides the estimation results and section 5 concludes.

2. Empirical Design

The augmented Mincerian equation posits that the wage of individual *i* living in state *s* in period *t* is determined by:

$$\ln w_{ist} = \beta_0 + \beta_1 ed_{ist} + \alpha ED_{st} + \beta_2 exp_{ist} + \beta_3 exp_{ist}^2 + x_{is}\gamma + v_s + t_t + u_{ist} \qquad (1)$$

where ed_{ist} is individual education, ED_{st} is average state education, exp_{ist} is individual experience, x_{is} are observed demographic characteristics of individual *i* in state *s*; v_s are fixed state effects, and t_t are fixed time effects. The coefficient β_1 is the private return to schooling, while the coefficient α represents the human capital externality. The sum of β_1 and α is the social return to schooling.

Following Moretti (2004a, 2004b), we define the error term as a function of three components:

$$u_{ist} = \mu_s\theta_i + v_{st} + \varepsilon_{ist} \qquad (2)$$

where θ_i is a permanent unobservable component of human capital (i.e. ability and skill) for individual *i*; μ_s is the return on that unobserved skill in state *s*; v_{st} is time-varying shocks to labor demand and supply in state *s* in time *t*; and ε_{ist} is the transitory component of wages which is assumed to be i.i.d. over individuals, states and time. With the exception of the valuation of local amenities, equations (1) and (2) are very similar to those derived by Lange and Topel (2006) using a spatial equilibrium model.

There are two potential sources of bias arising from estimating equation (1). First, the presence of time-varying or transitory shocks to state labor market that are correlated with average education, $E(v_{st}ED_{st}) \neq 0$. For example, a skill-biased technological shock may increase the demand for skilled labor and result in higher state-level education levels. The second source of bias would arise if more productive workers locate in high-education states versus low-education states, $E(\theta_i ED_{st}) > 0$.

As mentioned earlier, the most common strategy to control for this omitted variable bias is to instrument for average education. These instruments must be both exogenous and relevant to generate consistent estimates of the human capital externality, *α*. Relevancy requires that the instruments be correlated with average education. Validity requires that the instruments be orthogonal to the error term u_{ist} - essentially requiring orthogonality between the instruments and

[2] While Rauch (1993) included city-level characteristics and Rudd (2000) and Lang and Topel (2006) include state fixed effects, none of these studies included a comprehensive set of time-varying geographic controls.

both unobservable skills (both individual- and state-level) and transitory shocks to the state labor market.

Our identification strategy is to directly control for observed and unobserved individual job characteristics and *time-varying* state labor market conditions:

$$\ln w_{ist} = \beta_0 + \beta_1 ed_{ist} + \alpha ED_{st} + \beta_2 exp_{ist} + \beta_3 exp_{ist}^2 + x_{ist}\gamma + o_{ist}\kappa + Z_{st}\delta + \mu_c + \nu_s + t_t + \varepsilon_{ist} \quad (3)$$

where o_{ist} are a set of observed job characteristics of individual *i* in state *s* at time *t*; Z_{st} are a set of factors that shift labor demand and supply in state *s* at time *t*; μ_c are fixed cohort effects; ν_s are fixed state effects; and t_t are fixed time effects.

We use Current Population Survey (CPS) data on demographic and job characteristics and construct cohorts to identify differences in observed and unobserved ability in equation (3). Age and its square, gender, race, and ethnicity are included as demographic controls. The job characteristics include controls for union membership and coverage by a collective bargaining agreement, 12 broad occupation codes, and 15 broad NAICS industry indicators. The inclusion of cohort-specific effects, μ_c, arises from the pseudo-panel nature of the CPS. In true panel data individuals are observed multiple times and their individual-specific effects can be controlled through either fixed or random effects. The CPS data we use has only a single observation per individual (more detail on the CPS data is presented in the next section) so it is not possible to include individual-specific fixed effects. Deaton (1985), Verbeek and Nijman (1993), and Nijman and Verbeek (1990) propose to construct cohorts of individuals who are likely to have similar individual effects and include cohort controls in the regression model. We construct our cohorts by 10 age groups (five-year increments from 16-65 years old), 12 occupation groupings and sex, which generates 240 individual cohort controls. These cohorts will proxy for individual unobservables; we assume that these unobserved individual characteristics are similar for those in the same occupation, age range, and gender.[3]

We use state data to control for time-varying labor markets shocks, Z_{st}, in equation (3) (with the exception of Alaska and Hawaii which are excluded from our study). To capture changes in state labor demand, we include state capital-to-labor ratio, log of population density, and the industry share of state output. The capital-to-labor ratio provides the physical capital per average worker. Population density is the ratio of total population to surface area and controls for possible agglomeration effects. The industry share of state output is the percentages of state income generated by the NAICS industry.

To control for changes in state labor supply, we use state labor force participation rate, unionization rate, effective minimum wage, average public assistance and average workers compensation. The effective minimum wage, average public assistance and average workers compensation are intended to control for the reservation wage. We also include the state unemployment rate to measure the slack in the labor market. Time invariant state effects are controlled through state fixed effects.

3. Data and Estimation

Microdata were obtained from the CPS Outgoing Rotations Groups (ORG) files for 1992-2005. A respondent in the CPS is interviewed each month for four months, leaves the

[3] In an alternate set of regressions, we construct age-state-gender cohorts. Using these cohorts did not change estimation results appreciably from the age-occupation-gender cohorts.

sample for eight months, and then returns to the sample for four additional months. Respondents are asked detailed economic questions in months four and eight of their stay in the sample and these variables are contained in the ORG files. For our study, we use an individual's first instance in the sample (the economic data from the fourth month) and restrict our sample to employees aged 16 to 65 years.

The state-level data are drawn from a variety of sources. Capital data are from Yamarik (2013), while the population and labor force participation data are from the U.S. Census Bureau (2012a). The NAICS share of state income is constructed using data from the U.S. Bureau of Economic Analysis (2012). Unionization, minimum wage and unemployment data are from Hirsch and Macpherson (2012), U.S. Department of Labor (2012) and the U.S. Bureau of Labor Statistics (2012), respectively. Average public assistance and average workers compensation are from the U.S. Census Bureau (2012b).

We estimate variations on (3) using least squares, clustering standard errors on state-year. The dependent variable is the natural logarithm of the hourly wage rate in 2011 dollars. We use four specifications. The base specification includes individual controls for education, age, age-squared, sex, marital status, race and ethnicity plus state-level average education.[4] The second specification adds individual dummies for union membership, coverage by a collective bargaining agreement, industry, and occupation as well as state fixed effects. The third specification includes the NAICS share of output at the state level and the cohort fixed effects.[5] The final, preferred, specification, based upon equation (3), adds the state-level *time-varying* controls for labor demand and supply. For clarity, we present individual-level variables in lower-case and state-level variables in upper-case.

4. Results

Table 1 presents the estimates of the private and external returns to schooling (the complete results for specification 4 are presented in Appendix A). The coefficients for the individual controls have the expected sign and are significant at the one percent level. Of particular note, the coefficients imply that union membership and collective bargaining agreement raise individual wages by 15.2 and 6.0 percent, respectively. For the state-level variables, higher levels of capital per worker, population density, NAICS share, workers compensation, public assistance and unemployment increase the individual wage; while labor force participation, unionization and minimum wage have no impact.

We estimate a human capital externality in the range of 2.4 to 5.5 percent in columns 1-4. Using a limited set of individual controls in column 1, we find that an additional year of average schooling raises the individual wage by 2.4 percent. After controlling for individual unionization, industry and occupation effects and, finally, time-varying state effects, these estimates increase to 3.6 and 5.5 percent in columns 2 and 4, respectively. At the same time, as one would expect, the private return to schooling falls from 8.3 percent in column 1 to 5.0 percent in columns 3 and 4 as controls for job characteristics and cohort are incorporated into the model.[6]

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[4] The CPS records individual schooling by categories of primary school completed or highest secondary and tertiary degree obtained. We use the imputation method of Jaeger (1997) to convert this into a years of schooling variable.

[5] The inclusion of the cohort fixed effects requires the removal of the individual gender and occupation variables due to perfect multicollinearity.

[6] Including occupation and industry controls in the model is typical of these studies, but may lead to downward bias in the private returns to education as workers with higher education sort into higher-paid occupations and industries.

We next use the share of college-educated workers as the measure of state-level aggregate schooling in columns 5-8. This alternate measure of state-level schooling allows us to directly compare our estimates to Moretti (2004a). Understanding the particular role of college investment is important given ever-decreasing state subsidies for college education (as a share of total cost) in the past two decades.

If we re-estimate specification 4 without industry controls and excluding occupation from the cohort construct then the private return to education returns to an 8 percent estimate, comparable to specification 1. The estimate of external returns to education rises to 6 percent.

Table 1: Private and External Returns to Schooling

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
years of schooling$_{i,s,t}$	0.0834***	0.0566***	0.0499***	0.0499***	0.0833***	0.0566***	0.0499***	0.0500***
	(0.000212)	(0.000219)	(0.000268)	(0.000268)	(0.000212)	(0.000219)	(0.000268)	(0.000268)
AVG. YEARS OF SCHOOL$_{s,t}$	0.0243***	0.0361***	0.0341***	0.0551***				
	(0.00638)	(0.00584)	(0.00597)	(0.000268)				
COLLEGE SHARE$_{s,t}$					0.356***	0.350***	0.335***	0.403***
					(0.0535)	(0.0490)	(0.0510)	(0.0522)
Observations	868,035	868,035	773,381	773,381	868,035	868,035	773,381	773,381
R-squared	0.383	0.483	0.509	0.509	0.383	0.483	0.509	0.509
Individual Controls:								
age & demographics	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
union & collective bargain	No	Yes	Yes	Yes	No	Yes	Yes	Yes
industry dummies	No	Yes	Yes	Yes	No	Yes	Yes	Yes
occupation dummies	No	Yes	No	No	No	Yes	No	No
cohort fixed effects	No	No	Yes	Yes	No	No	Yes	Yes
State Controls:								
state fixed effects	No	Yes	Yes	Yes	No	Yes	Yes	Yes
NAICS share	No	No	Yes	Yes	No	No	Yes	Yes
time-varying state factors	No	No	No	Yes	No	No	No	Yes

Notes: The dependent variable is the natural logarithm of hourly earnings in 2011 dollars. Each regression includes individual years of schooling and an aggregate average measure of schooling. The additional individual and state controls used are shown at the bottom with details provided in Appendix A. The standard errors in parenthesis are corrected for clustering on state-year where *** significant at the 1% level, ** significant at the 5% level and * significant at the 10% level.

Table 2: Estimates of Private and Social Returns to Education

Author	**Data**				**Individual Controls**				**Aggregate Controls**						**Results**	
Author	**Time Period**	**Source of data on individuals**	**geography level**	**Estimator**	**Education**	**demographic controls**	**Occupation, industry, union controls**	**Individual Effects**	**Education**	**Labor Market**	**Labor Demand**	**Labor Supply**	**Geographic**	**Other Factors**	**Private return**	**Externality**
Rauch	1980	Census	SMSA	GLS	years	sex, race	occupation, union		mean value	unionization		population	region FE, climate, coast		0.05	0.03-0.05
Acemoglu and Angrist	1950-90	Census, 40-49 white males	states	OLS	years	No			mean value						0.07-0.09	0.06-0.07
Acemoglu and Angrist	1950-90	Census, 40-49 white males	states	IV	years	No			mean value						0.06-0.07	zero
Rudd	1978-91	CPS	48 + DC	WLS	years	sex, race	industry		lagged mean	unemployment rate		agglomer index		non-labor % of income	N/A	0.05
Rudd	1978-91	CPS	48 + DC	WLS	years	sex, race	industry		lagged mean	unemployment rate		agglomer index	state FE	non-labor % of income	N/A	0-0.03
Moretti	1979-94	NLSY	SMSA	OLS	years	sex, race		fixed effects	% of college graduates	unemployment rate	Katz-Murphy index		city FE		0-0.10	1.00 implies a 0.25 return
Moretti	1980, 90	Census	SMSA	OLS & IV	years	sex, race			% of college graduates	unemployment rate	Katz-Murphy index		city FE		0.06-0.09	1.25 implies a 0.30 return
Monaco and Yamarik	1992-2005	CPS	48 + DC	OLS	years	sex, race	industry, occupation, union	cohort effects	mean value, % college graduates	unemployment rate, unionization	K/L, NAICS share	labor force participation, per capita workers comp, per capita assistance	state FE		0.05-0.08	0.02-0.05

The coefficient estimates for the college share range from 0.34 to 0.40 and follow a similar pattern as columns 1-4. Under the complete specification, a one percentage point increase in the share of college-educated workers raises the individual wage by 0.40 percent, which translates into a human capital externality of 10.0 percent per year of schooling.[7,8] Thus we find external returns to college education of higher magnitude than average years of education. This supports Rosenthal and Strange (2008) who find that proximity to college-educated workers raises individual wages, while proximity to less-than-college workers lowers individual wages.

While the use of an extensive set of time-varying state controls may raise concerns of collinearity, we find limited levels of collinearity among our state-level variables. A related concern is whether the time-varying controls are endogenous. To assess whether this biases the results, we re-estimate specification 4 with lagged values of the state-level controls. Using a one year lag leads to an estimate of private return to education of 5 percent and an external return of 5 percent, lagging 3 and 5 years results in the external returns falling to 4 percent. At the extreme, using only pre-estimate values for the state controls leads to an external return of 2 percent. Thus we conclude that endogeneity among time-varying state controls is not driving our results.

To place our findings in perspective, table 2 presents information on related studies that use worker-level microdata, including data source, estimation approach, individual- and state-level controls, and private and social returns to education. Our estimates of the external return are similar to Rauch and Rudd as well as the OLS estimates of Acemoglu and Angrist. The externalities associated with the percent college educated are smaller in magnitude than Moretti, though he notes that when he re-estimates his model at the state-level, the estimates are roughly one-third smaller.

Ciccone and Peri (2006) assert that the positive link between individual wages and state-level education does not necessarily imply an externality. Assuming that workers of different education levels are imperfect substitutes, an increase in the average level of schooling will have two distinct effects on the wage distribution. First, the increase in the relative supply of college-educated workers will lower the wages of high-education workers and raise the wages of low-education workers. Second, human capital externalities will raise the wage of both groups.

We estimate our augmented Mincerian model for different terminal levels of education, following Moretti (2004b). If externalities exist, then there should be a positive and significant coefficient on state-level education for each of the education subsamples. If, in addition, the returns to state-level schooling are higher for the low-education subsamples then this suggests some substitutability between low-skill and high-skill workers (a rightward shift in the demand curve for low-skill workers).

Table 3 presents the estimates of the private and external returns to schooling for different levels of education using the full specification from (3). We divide our sample into four

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7 As with Moretti (2004a), it takes a change in the college share of 0.25 to increase average years of schooling by one in our sample. As a result, the external return of an additional average year of schooling is 0.25 x 0.40 = 0.10

8 We perform a few checks on our model to examine whether the cohorts appear to do an adequate job of controlling for the possibility that high-productivity people self-select into higher education areas. Assuming that this selection may vary systematically by age, we re-estimate the models for “young” workers and “old” workers. We also construct alternative cohorts using state-occupation-gender groupings. The results of these alternative specifications do not vary considerably from our preferred specification Full estimation results are available from the authors upon request.

sub-samples based on individual education attainment: (i) less than high school (ii) high-school graduates, (iii) some college and two-year college degrees and (iv) four-year college graduates and higher. We measure aggregate schooling as average years of schooling in columns 1-4 and as the share of college-educated workers in columns 5-8. The individual years of schooling variable is dropped as the samples are based on education level.

We find that the effect of average education is greater for low-education workers. A one year increase in average education raises the wages of high school drop-outs by 8.3 percent and high-school graduates by 7.8 percent. Likewise, a ten percentage point increase in the share of college-educated workers raises the wages of high school drop-outs by 3.7 percent (a 9.3 percent externality) and high-school graduates by 4.2 percent (a 10.5 percent externality). For the higher-educated groups, a one year increase in average schooling raises the wages of those with some college by 3.5 percent and 4-year college graduates by 4.4 percent. A ten percentage point increase in the share of college-educated workers raises the wages of workers with some college by 1.9 percent (corresponding to a 4.3 percent externality) and 4-year college graduates by 3.0 percent (a 6.9 percent externality). The positive coefficient for state-level education across the education sub-groups is indicative of a human-capital externality and aligns with the findings of Moretti (2004b).

Table 3: Private and External Returns to Schooling by Education Level

VARIABLES	(1) less than high school	(2) high school graduates	(3) some college	(4) 4-year college and higher	(5) less than high school	(6) high school graduates	(7) some college	(8) 4-year college and higher
AVG. YEARS OF SCHOOL$_{s,t}$	0.0831***	0.0784***	0.0341**	0.0434*				
	(0.0182)	(0.0141)	(0.0167)	(0.0259)				
COLLEGE SHARE$_{s,t}$					0.374***	0.416***	0.194**	0.296**
					(0.110)	(0.0812)	(0.0944)	(0.129)
Observations	103,794	258,409	227,051	184,127	103,794	258,409	227,051	184,127
R-squared	0.399	0.365	0.421	0.331	0.399	0.365	0.421	0.331

Notes: The dependent variable is the natural logarithm of hourly earnings in 2011 dollars. Each regression includes individual controls for age & demographics, union and collective bargaining, industry and cohort fixed effects and state-level controls for state fixed effects, NAICS share and time-varying state factors. See Appendix A for details. The standard errors in parenthesis are corrected for clustering on state-year where *** significant at the 1% level, ** significant at the 5% level and * significant at the 10% level.

We next examine whether differences in educational externalities could be a source of regional divergence across U.S. states. City-level analysis by Moretti (2004b), Shapiro (2006) and Diamond (2012); and state-level analysis by Ciccone and Peri (2006) and Lindley and Machin (2012) have documented that geographic areas with high-skill workers experience higher wage and housing price growth relative to those with low-skill workers. Moretti (2012) termed this "the great divergence". At the very least, states with high-skill workers will generate larger *total* externalities due to their larger stock of education attainment. However, individual workers in these same states may also experience a larger *marginal* externality if their fellow workers are more innovative.

We test the role of education externalities in explaining this divergence by dividing our sample along two state-level dimensions: education and innovation. We use the share of college educated workers in the 2000 U.S. Census to rank educational attainment and the State New Economy Index of Atkinson (2002) to rank innovation.[9] We split the 50 states into halves to produce four sub-samples: (i) low-education, low-innovation; (ii) low-education, high-innovation; (iii) high-education, low-innovation; and (iv) high-education, high-innovation. Since the sub-samples are defined based on 2000 data, we restrict our estimation to the years 2001-5. The results using average years of schooling are presented in columns 1-4 of Table 4 and those using the share of college-educated workers are shown columns 5-8.

We find that education externalities are larger in the high-education, high-innovation states. In the low-education, low-innovation states, an extra year of average state-level schooling increases the individual wage by 2.9 percent. The externality increases roughly two-fold, to 5.6, for high-education, high-innovation states. However, a ten percent increase in the share of college-educated workers in the low-education, low-innovation states raises the individual wages by 5.6 percent (a 14 percent externality), while that same ten percent increase in the high-education, high-innovation states increases the individual wage by 5.5 percent (a 13.8 percent externality). These results suggest that there are larger *marginal* externalities accruing from increased college education in low-education low-innovation areas, consistent with earlier results for education sub-groups.

[9] The State New Economy Index is a composite measure of (i) knowledge jobs, (ii) export orientation and FDI, (iii) economic dynamism and competition, (iv) internet use by consumers and business, and (v) research and development (Atkinson, 2002).

Table 4: Private and External Returns to Schooling by State-Level Education and Innovation

VARIABLES	(1) low education, low innovation states	(2) low education, high innovation states	(3) high education, low innovation states	(4) high education, high innovation states	(5) low education, low innovation states	(6) low education, high innovation states	(7) high education, low innovation states	(8) high education, high innovation states
years of schooling$_{i,s,t}$	0.0555***	0.0552***	0.0513***	0.0518***	0.0555***	0.0552***	0.0513***	0.0518***
	(0.000783)	(0.00113)	(0.00185)	(0.000558)	(0.000846)	(0.000113)	(0.00195)	(0.000558)
AVG. YEARS OF SCHOOL$_{s,t}$	0.0288**	0.0688	0.00585	0.0564***				
	(0.0125)	(0.0493)	(0.122)	(0.00546)				
COLLEGE SHARE$_{s,t}$					0.562***	0.0691	1.657	0.557***
					(0.0796)	(0.333)	(1.078)	(0.0440)
Observations	94,907	50,584	16,330	155,659	94,907	50,584	16,330	155,659
R-squared	0.466	0.482	0.487	0.490	0.467	0.482	0.488	0.490

Notes: The dependent variable is the natural logarithm of hourly earnings in 2011 dollars. Each regression includes individual controls for age & demographics, union and collective bargaining, industry and cohort fixed effects and state-level controls for state fixed effects, NAICS share and time-varying state factors. See Appendix A for details. The standard errors in parenthesis are corrected for clustering on state-year where *** significant at the 1% level, ** significant at the 5% level and * significant at the 10% level.

5. Conclusions and Discussion

This paper examines the role of human capital externalities in explaining individual wages. We identify the externality by directly controlling for individual job characteristics and state labor market conditions. Predictably, controlling for individual unobservables through cohort fixed effects lowers the private return to education. Perhaps surprisingly, including a broad set of time-varying state controls results in larger estimates of the externality from schooling. The importance of transitory labor market shocks is referenced by Rudd (2000, p. 11), but prior studies based on city-level estimates do not have a rich set of time-varying controls to incorporate into their models. Under our preferred specification, we estimate that an extra year of state-level schooling increases individual wages by 5.5 percent, roughly equivalent to the private return of 5.0 percent.

The estimated magnitude of the human capital externality is consistent with prior research that used a similar identification scheme. Rauch (1993), Rudd (2000) and Lange and Topel (2006) use individual worker characteristics and city- or state-level conditions for identification and estimate a human capital externality between 3 and 7 percent.[10] It is also notable that the externality on the specification which uses the state's share of college graduates is larger (roughly twice) in absolute magnitude than the state's average years of education, which is in line with Iranzo and Peri (2009) and Moretti (2004a).

We next estimate the human capital externality across different samples and find that it is larger for individuals (and states) with low levels of education and for states with high levels of both education and innovation. These two results imply that the estimated externality is capturing spillover effects and is explaining some of the wage divergence across U.S. states. The larger externality associated with post-secondary education is consistent across workers of different education levels and has important policy implications for state governments. Our results support the hypothesis that investing in education is as important for state-level economic performance as it is for the individual students who do so. The continued transfer of cost from the state to the individual student could lead to underinvestment in human capital relative to the socially optimal level.

[10] The lone exception is Rudd (2000) who estimated an insignificant human capital externality when state fixed effects are included.

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Appendix A: Full Estimation Results from Table 1, Full Specification (Column 4)

Dependent variable: log of hourly wages	Coefficient	Standard error
<u>*Individual and Aggregate Schooling:*</u>		
Years of schooling	0.0499	0.000268
AVG YEARS of SCHOOLING	0.0551	0.000268
<u>*Individual Age and Demographics:*</u>		
Age	0.0431	0.00102
Age-squared	-0.000455	1.37e-05
Married	0.0551	0.00139
Separated, Divorced, Widowed	0.0204	0.00179
Black	-0.0877	0.00164
American Indian	-0.0513	0.00484
Asian	-0.0353	0.00293
Other race	-0.0177	0.00565
Hispanic	-0.0577	0.00176
<u>*Union and Collective Bargaining Status:*</u>		
Union member	0.152	0.00173
Collective bargaining coverage	0.0598	0.00504
<u>*NAICS Individual Industries:*</u>		
Mining	0.171	0.00645
Utilities	0.190	0.00585
Construction	0.0348	0.00325
Wholesale Trade	0.00567	0.00319
Retail Trade	-0.180	0.00247
Transportation and Warehousing	0.0403	0.00353
Information	0.0605	0.00334
Finance, Insurance, and Real Estate	0.0192	0.00252
Prof., Scientific, Tech. plus Management	0.0223	0.00286
Administrative Services	-0.0660	0.00360
Educational Services	-0.142	0.00423
Health Care and Social Assistance	-0.0552	0.00231
Accommodation and Food Service	-0.231	0.00297
Other Services (including Arts)	-0.142	0.00361
<u>*State Controls:*</u>		
NAICS SHARE	0.171	0.00645
ln(*CAPITAL per WORKER*)	0.0558	0.0147
ln(*POPULATION DENSITY*)	0.150	0.0157
LABOR FORCE PARTICIPATION RATE	0.0565	0.0736
UNIONIZATION RATE	0.0108	0.0240
MINIMUM WAGE	0.00150	0.00141
PUBLIC ASSISTANCE PER CAPITA	0.223	0.0521
WORKERS COMP. PER CAPITA	0.0396	0.0194

UNEMPLOYMENT RATE	0.235	0.0825
Observations	773,381	
R-squared	0.509	

Notes: The results are for column (4) of Table 1. The coefficients for the individual demographic variables are interpreted relative to the omitted categories of *Single* and *White*. The coefficient for each NAICS individual industry is interpreted relative to the omitted *Manufacturing* industry.