LONG-RUN SUBSTITUTABILITY BETWEEN MORE AND LESS EDUCATED WORKERS: EVIDENCE FROM U.S. STATES, 1950–1990

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Abstract—We estimate the aggregate long-run elasticity of substitution between more educated workers and less educated workers (the slope of the inverse demand curve for more relative to less educated workers) at the U.S. state level. Our data come from the (five) 1950–1990 decennial censuses. Our empirical approach allows for state and time fixed effects and relies on time- and state-dependent child labor and compulsory school attendance laws as instruments for (endogenous) changes in the relative supply of more educated workers. We find the aggregate long-run elasticity of substitution between more and less educated workers to be around 1.5.

I. Introduction

The aggregate, long-run elasticity of substitution between more educated workers and less educated workers (the slope of the inverse demand curve for more educated workers relative to less educated workers) plays an important role in several areas of economics. For instance, the extent to which differences in average labor productivity across countries can be explained by differences in levels of education depends on this substitution elasticity (for example, Klenow & Rodriguez-Clare, 1997; Hendricks, 2002). The impact of an increase in the share of more educated workers on the average return to education is also determined by the elasticity of substitution between more and less educated workers. And understanding whether technological change is biased toward more or less educated workers and quantifying the impact of biased technological change on the return to education also requires knowledge of this substitution elasticity (for example, Autor & Katz, 1999; Katz & Murphy, 1992; Acemoglu, 2002). Our main contribution in this paper is to provide estimates of the long-run elasticity of substitution between more and less educated workers using data on U.S. states for the period 1950–1990.

The literature estimating the elasticity of substitution between workers with different levels of education using aggregate data stretches from the 1970s (for example, Griliches, 1969; Johnson, 1970; Dougherty, 1972; Fallon & Layard, 1975) to the 1990s (for example, Katz & Murphy, 1992; Angrist, 1995). One of the main difficulties faced by researchers in this area is that the education wage premium and the number of more relative to less educated workers are determined simultaneously by demand and supply. Estimating the slope of the (inverse) demand curve for more educated workers relative to less educated workers therefore requires solving the standard identification problem. Empirical work that does not address identification explicitly is likely to lead to misinterpretations of the data. For example, consider an economy experiencing rapid technological change favoring educated workers. This could lead to an increase in the relative demand as well as the relative supply of more educated workers (for example, Fallon & Layard, 1975; Acemoglu, 1998). The relative wage of more educated workers could therefore be increasing at the same time as relative employment of more educated workers is rising, even if the relative demand curve for more educated workers is decreasing in their relative wage (because firms substitute away from more educated workers as these become relatively more expensive). Hence one needs to be careful in inferring the slope of the relative demand curve for more educated workers from data on the relative employment and the relative wage of more educated workers [see Hamermesh (1993) for a summary of the identification problem in the context of labor demand estimation].

We identify the long-run elasticity of substitution between more and less educated workers at the U.S. state level using data from the (five) 1950–1990 decennial censuses. Our empirical approach allows for state and time fixed effects and relies on time- and state-dependent child labor and compulsory school attendance laws as instruments for the (endogenous) relative supply of more educated workers [data on these laws have been collected by Acemoglu and Angrist (2000)]. Our identifying assumption is that changes in these laws are independent of expected shifts in the relative demand for more educated workers. Our principal conceptual framework adapts the constant-elasticity-of-substitution approach of Katz and Murphy (1992), but we also consider the so-called translog framework as an alternative. The main difference between the two approaches is that the translog framework allows the elasticity of substitution between workers with different education levels to vary with their relative supply.

We estimate the long-run elasticity of substitution between more and less educated workers with a variety of methods, ranging from two-stage least squares to Fuller-modified limited-information maximum likelihood, which has been shown to be more robust to instrument weakness than two-stage least squares (for example, Stock, Wright, & Yogo, 2002; Hahn & Hausman, 2002). Our estimates of the long-run elasticity of substitution between workers with high and low education levels range between 1.2 and 2, and

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our preferred estimate is 1.5. These estimates are similar to several other estimates that try to correct for the endogeneity of average schooling attainment (using approaches that differ from ours).

Estimates of the elasticity of substitution between workers with different levels of education in production [that is, estimates of the slope of the demand for more educated workers relative to less educated workers] can be used to determine to what extent changes in the education wage premium are driven by shifts in the relative demand for more educated workers. This question also dates back to the 1970s. For example, Fallon and Layard (1975) ask why the secular increase in the supply of more educated workers in the 1950s and 1960s did not decrease the education wage premium, and Griliches (1969), Bowles (1970), and Dougherty (1972) previously analyzed very similar issues. The increase in the education wage premium during the 1980s and 1990s revived interest in this question (for example, Katz & Murphy, 1992; Acemoglu, 2002). We quantify shifts in the relative demand for more educated workers, which we interpret as skill-biased technological change, across U.S. states between 1950 and 1990 using both the constant-elasticity-of-substitution framework and the translog framework.

The rest of the paper is organized as follows. Section II presents the constant-elasticity-of-substitution framework and our main estimating equation. Section III discusses the data and instruments. Section IV presents and discusses our estimates of the long-run elasticity of substitution between more and less educated workers obtained using the constant-elasticity-of-substitution framework. Section V presents the translog specification and the implied elasticity estimates. Section VI presents and discusses our estimates of skill-biased technological change for U.S. states between 1950 and 1990. Section VII summarizes and concludes.

II. The Constant-Elasticity-of-Substitution Framework

Our simplest model assumes that output Y in state s in year t is produced according to a constant-returns-to-scale, constant-elasticity-of-substitution production function

$$Y_{st} = A_{st}(L_{st}^{(\sigma-1)/\sigma} + B_{st}H_{st}^{(\sigma-1)/\sigma})^{\sigma/(\sigma-1)}, \tag{1}$$

where L_{st} denotes efficiency units of less educated workers, and H_{st} efficiency units of more educated workers employed in production; A_{st} and B_{st} capture Hicks-neutral and skill-biased shifts in technology, respectively; and the parameter $\sigma > 0$ determines the substitutability between more and less educated workers [see Katz and Murphy (1992) and Acemoglu (2002) for very closely related approaches]. We have eliminated physical capital from the production function for simplicity. Including physical capital in the analysis is straightforward and does not lead to changes in the specification or interpretation of our results under assumptions that we defend as reasonable in the Appendix.

The production function in equation (1), combined with cost minimization and pricetaking in the labor market, leads to the following relative demand curve for more educated workers:

$$\ln(H_{st}^D/L_{st}^D) = -\sigma \ln(w_{st}^H/w_{st}^L) + \sigma \ln B_{st}. \tag{2}$$

Hence, the long-run elasticity of substitution between more and less educated workers (the percentage decrease in the relative demand for more educated workers, H^D/L^D, in response to a 1% increase in their relative wage, w^H/w^L) is equal to σ. It is a defining feature of the constant-elasticity-of-substitution production function that this elasticity is constant along the relative demand curve. In section V we implement a (translog) specification that allows the substitution elasticity to vary along the demand curve.

In labor market equilibrium, the relative demand for more educated workers is equal to the relative supply, H_{st}/L_{st}. Hence, equation (2) implies that equilibrium wages are linked to the relative supply of more educated workers by

$$\ln\left(\frac{w_{st}^H}{w_{st}^L}\right) = -\frac{1}{\sigma}\ln\left(\frac{H_{st}}{L_{st}}\right) + \alpha_t + \alpha_s + u_{st}, \tag{3}$$

where we have written the skill-biased technology, $\ln B_{st}$, as the sum of a state fixed effect, a time effect, and a residual state-time effect, $\alpha_t + \alpha_s + u_{st}$. This is our main estimating equation.

As the long-run relative supply of more educated workers at the state level is likely to be positively correlated with shifts in relative labor demand at the state level (captured by u_{st}), the coefficient $1/\sigma$ cannot be estimated consistently using least squares (the positive correlation may arise because of interstate migration or extended studies in response to higher wage premia for more educated workers). We therefore use instrumental variables estimation. Our instruments are constructed using information on compulsory attendance and child labor laws gathered by Acemoglu and Angrist (2000) (who also show that these laws affect average levels of education of U.S. states). Our identifying assumption is that changes in compulsory attendance and child labor laws are unrelated to the expected skill-biased technology shock.

III. Data and Instruments

A. Labor Supply and Wages

Our wage and labor supply data come from the U.S. Census Integrated Public Use Microdata Sample (IPUMS) and refer to the (five) 1950–1990 decennial censuses. All wage data used in our empirical work refers to U.S.-born white males between 40 and 50 years of age. This ensures that changes in average wages are not driven by age, gender, or race composition. Our data identify the highest schooling degree obtained by each person in the sample. This allows us to group workers in four education categories: high school dropouts (HSD) are workers without a high school degree, high school graduates (HSG) are workers with a high school degree who did not go to college, college dropouts (CD) are workers with at least one year of schooling after high school but no college degree, and college graduates (CG) are workers with a four-year college degree. We measure the supply of workers with different education levels in each state, either as the share of white male workers between 21 and 59 years of age in the four education categories or as the share of all workers between 16 and 65 years of age in the four education categories. With these data we obtain the relative supply of more educated workers, H_{st} / L_{st}, on the right-hand side of equation (3) in the following way:

- We treat HSD as less educated workers: $L_{st} \equiv L_{st}^{\text{HSD}}$.
- We treat the three categories HSG, CD, and CG as more educated workers and aggregate them using $H_{st} \equiv L_{st}^{\text{HSG}} + L_{st}^{\text{CD}}(\bar{w}_t^{\text{CD}}/w_t^{\text{HSG}}) + L_{st}^{\text{CG}}(\bar{w}_t^{\text{C}}/\bar{w}_t^{\text{HSG}})$, where $\bar{w}^{\text{CD}}$, $\bar{w}^{\text{CG}}$, $\bar{w}^{\text{HSG}}$ denote average national wages for college dropouts, college graduates, and high school graduates, respectively (we therefore treat the three categories of more educated workers as perfect substitutes and measure the aggregate supply of more educated workers in high school equivalence units). The wage data used to aggregate the three categories of more educated workers refer to white males when we calculate supply using the shares of white males in the four education categories, and to all workers when we calculate supply using the shares of all workers in the four education categories.

We consider two measures of the relative supply of more educated workers—one based on the labor supply of white male workers only, and the other based on the labor supply of all workers—to ensure that our empirical findings are robust to labor market segmentation by gender and race. If markets are very segmented in this dimension, the supply measure using white males only is more appropriate. If there is no segmentation at all, the supply measure using all workers is preferable. In practice, the two supply measures yield very similar estimates of the elasticity of substitution between more and less educated workers.

We obtain relative wage of more educated workers, w_{st}^H/w_{st}^L, on the left-hand side of equation (3) by dividing the average weekly wage of high-school-equivalent workers in the white-male wage sample, w^H, by the average weekly wage of workers without a high school degree, w^L, in the same sample (details are given in the Appendix). As robustness check we also measure more educated workers in college equivalence units.

We associate the cutoff between more and less educated workers with high school graduation for three reasons. First, between 1950 and 1990, the most important aspect of increased schooling attainment was the rising share of workers with at least a high school diploma. Table 1 (based on data for white male workers between 21 and 60 years) shows that the proportion of workers without a high school degree decreased from 60% in 1950 to 12% in 1990. The increase of college graduates, in comparison, was much smaller (from 8% in 1950 to 25% in 1990). Second, associating the cutoff between more and less educated workers with high school graduation is in line with the cross-country literature on the role of education in economic development (for example, Mankiw, Romer, & Weil, 1992; Bils & Klenow, 2000; Caselli & Coleman, 2000; Hendricks, 2002). Third, our instruments for changes in the relative supply of more educated workers (changes in compulsory attendance and child labor laws) mainly affect the high school graduation margin.

Table 1.—Evolution of Schooling in the U.S. Working Population

Year	Share of HS dropouts (average)	Share of HS graduates (average)	Share of college dropouts (average)	Share of college graduates (average)
1950	0.60	0.22	0.10	0.08
1960	0.50	0.28	0.11	0.11
1970	0.35	0.35	0.15	0.15
1980	0.22	0.37	0.20	0.21
1990	0.12	0.33	0.30	0.25

Source: Authors' calculations on U.S. Census IPUMS data 1950, 1960, 1970, 1980, and 1990.
Sample: U.S.-born, white, male workers between 21 and 59 years of age in 48 U.S. contiguous states.

Table 2 shows the evolution of the wage premium of college graduates relative to high school dropouts between 1950 and 1990 and compares it with the wage premium of college graduates relative to high school graduates. The wage premium of college graduates relative to high school dropouts increased by 90% over the whole period, which exceeds the increase of the wage premium of college graduates relative to high school graduates. The qualitative behavior of the two education wage premia in each decade is similar.

B. *Instruments*

Acemoglu and Angrist (2000) have collected data on state- and year-specific compulsory attendance and child labor laws. We use these laws as instruments for changes in the relative supply of more educated workers at the state level. The basic information is summarized in eight dummies, *CL*6, *CL*7, *CL*8, *CL*9 and *CA*8, *CA*9, *CA*10, *CA*11, associated with each individual in our sample. The dummy *CLX* (with $X = 6, 7, 8, 9$) is equal to 1, and all other child-labor-law dummies are equal to 0, if the state where the individual is likely to have lived when aged 14 had child labor laws imposing a minimum of *X* years of schooling. And the dummy *CAX* (with $X = 8, 9, 10, 11$) is equal to 1, and all other compulsory attendance law dummies are equal to 0, if the state where the individual is likely to have lived when aged 14 had compulsory attendance laws imposing a minimum of *X* years of schooling. The eight dummies are used to calculate the share of individuals for whom each of the *CL*6–*CL*9 and *CA*8–*CA*11 dummies is equal to 1 in each state. Six out of these eight shares (we omit *CL*6 and *CA*8,

TABLE 2.—THE EVOLUTION OF RELATIVE WAGES IN THE UNITED STATES

Year	$\bar{w}^{CG}/\bar{w}^{HSD}$	$\bar{w}^{CG}/\bar{w}^{HS}$
1950	1.34	1.20
1960	1.69	1.36
1970	1.95	1.45
1980	1.98	1.45
1990	2.55	1.76
Percentage change over whole period	+90%	+46%

Source: Authors' calculations on U.S. Census IPUMS 1950, 1960, 1970, 1980, and 1990. Wages are measured as weekly wages of full-time U.S.-born, white, male workers between 40 and 50 years of age.

as both sets of variables add up to 1) are used as instruments for the relative supply of more educated workers. The data do not include precise information on where individuals lived when aged 14, which is why we follow Acemoglu and Angrist (2000) in assuming that at age 14 individuals either all lived in the current state of residence (state-of-residence approach) or all lived in the state where they were born (state-of-birth approach). Each method has drawbacks and advantages. For example, the state-of-birth approach probably approximates better the residence at age 14, which should translate into better explanatory power of the instruments for the relative supply of more educated workers. But if interstate migration responds to differences in education premia, states that experience upward shifts in the relative labor demand for more educated workers may attract relatively more workers from states requiring longer schooling. And this may induce a correlation between the instruments and relative labor demand shifts. The state-of-residence approach, on the other hand, generates correlation between the instruments and the relative supply of more educated workers only through the group of people who were affected by the compulsory attendance and child labor laws at 14 and did not migrate to another state. This minimizes concerns regarding the endogeneity of the instruments but at the same time reduces their explanatory power for the relative supply of more educated workers.

Our identifying assumption is that changes in child labor and compulsory attendance laws are not affected by expected shifts in the relative demand for more educated workers. This assumption seems reasonable. Acemoglu and Angrist (2000) argue that changes in these laws were determined by sociopolitical forces operating at the time of their implementation. It seems unlikely that these forces were related to future shifts in the relative demand for more educated workers. Moreover, Acemoglu and Angrist (2000) show that changes in child labor and compulsory attendance laws affected schooling primarily in those grades that were directly targeted, which is unlikely to be consistent with changes in laws being driven by future shifts in the labor demand for more educated workers in general. In addition, Lochner and Moretti (2004) report that changes in child labor and compulsory attendance laws preceded increases in schooling. The correlation between changes in child labor and compulsory attendance laws and subsequent changes in the relative supply of more educated workers is therefore unlikely to be driven by omitted factors such as tastes for schooling or family background variables.

Table 3 reports first-stage regression results for state-of-residence and state-of-birth instruments using different approaches to the measurement of the relative supply of more educated workers. The regressions include state as well as time fixed effects. Comparing the results using the state-of-residence approach [specifications (1) to (3)] and the state-of-birth approach [specifications (4) to (6)] confirms that the instruments have more explanatory power when constructed using the state-of-birth approach. This can be seen either looking at the *F*-statistic for the joint significance of all child labor and compulsory-attendance-law instruments or at the partial R^2. It can also be seen that the explanatory power of the instruments varies according to how the relative supply of more educated workers is constructed. Generally speaking, instruments work best when used to predict the (raw) ratio of high school graduates to high school dropouts [specifications (1) and (4)]. Differences across specifications are relatively small when using the state-of-birth approach, however. In this case, the *F*-statistic for the joint significance of all child labor and compulsory attendance law instruments is similar whether we predict the (raw) ratio of high school graduates to high school dropouts, the ratio of more educated workers in high

TABLE 3.—FIRST-STAGE REGRESSIONS

	Instruments Obtained Using State-of-Residence Approach			Instruments Obtained Using State-of-Birth Approach		
Specification	(1)	(2)	(3)	(4)	(5)	(6)
*CL*7	0.17 (0.09)	0.06 (0.08)	0.06 (0.08)	0.04 (0.13)	0.13 (0.16)	0.13 (0.16)
*CL*8	0.21 (0.09)	0.11 (0.11)	0.12 (0.11)	0.20 (0.14)	0.45 (0.17)	0.42 (0.17)
*CL*9	0.22 (0.10)	0.10 (0.11)	0.07 (0.11)	0.20 (0.13)	0.14 (0.21)	0.16 (0.21)
*CA*9	0.01 (0.08)	0.06 (0.08)	0.06 (0.08)	0.35 (0.10)	0.03 (0.19)	0.03 (0.18)
*CA*10	0.07 (0.09)	0.19 (0.11)	0.19 (0.11)	0.38 (0.14)	0.11 (0.17)	0.12 (0.17)
*CA*11	0.06 (0.10)	0.11 (0.11)	0.10 (0.12)	0.45 (0.15)	0.12 (0.12)	0.12 (0.12)
Partial R^2	0.058	0.056	0.056	0.065	0.061	0.064
F-test	2.56	1.84	1.84	3.91	3.70	3.75
p-value	0.02	0.08	0.08	0.001	0.003	0.002

Dependent variable: $\ln(H_{st}/L_{st})$ calculated using white male workers between 21 and 59 years of age. All first-stage regressions include state fixed effects and time fixed effects. Heteroskedasticity-robust standard errors are reported in parenthesis.

Specifications (1) and (4): $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{HSG}$.

Specifications (2) and (5): $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{HSG} + L_{st}^{CD}(\bar{w}^{CD}/\bar{w}^{HSG}) + L_{st}^{CG}(\bar{w}^{CG}/\bar{w}^{HSG})$ (high school equivalence units obtained using weights from relative average wages).

Specifications (3) and (6): $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{CG} + L_{st}^{CD}(\bar{w}^{CD}/\bar{w}^{CG}) + L_{st}^{HSG}(\bar{w}^{HSG}/\bar{w}^{CG})$ (college equivalence units obtained using weights from relative average wages).

TABLE 4.—THE SUBSTITUTION ELASTICITY BETWEEN MORE AND LESS EDUCATED WORKERS, 1950–1990
(CES SPECIFICATION; SUPPLY CONSTRUCTED USING WHITE MALES ONLY)

Estimation Method	Measured Relative Supply of More Educated Workers		
	(1) Supply: All Groups[a]	(2) Supply: All Groups[b]	(3) Supply: 2 Groups Only[c]
Panel A			
(i) LS	−6.25*** (0.40)	−6.66*** (0.40)	−5.55*** (0.30)
(ii) LS with state dummies and time fixed effects	2.85*** (0.57)	3.44*** (0.71)	3.12*** (0.72)
Panel B			
(i) 2SLS with state dummies and time fixed effects (using state-of-residence instruments)	1.38** (0.63)	1.75* (0.90)	1.56* (0.85)
(ii) LIML with state dummies and time fixed effects (using state-of-residence instruments)	1.20*** (0.48)	1.63* (0.72)	1.72** (0.69)
(iii) Fuller LIML, *F*-constant = 1, with state dummies and time fixed effects (using state-of-residence instruments)	1.30** (0.59)	1.72** (0.84)	1.78** (0.77)
(iv) Fuller LIML, *F*-constant = 4, with state dummies and time fixed effects (using state-of-residence instruments)	1.50** (0.44)	1.96** (0.92)	2.00** (0.84)
Panel C			
(i) 2SLS with state dummies and time fixed effects (using state-of-birth instruments)	1.36*** (0.47)	1.78*** (0.71)	1.96*** (0.61)
(ii) LIML with state dummies and time fixed effects (using state-of-birth instruments)	1.28*** (0.40)	1.69*** (0.61)	1.92*** (0.69)
(iii) Fuller LIML, *F*-constant = 1, with state dummies and time fixed effects (using state-of-birth instruments)	1.33*** (0.42)	1.75** (0.62)	1.96** (0.65)
(iv) Fuller LIML, *F*-constant = 4, with state dummies and time fixed effects (using state-of-birth instruments)	1.42*** (0.45)	1.85*** (0.63)	2.00** (0.64)

Years: 1950, 1960, 1970, 1980, and 1990, 48 contiguous U.S. states, total of 240 observations.

The parameters presented and their standard errors are obtained from the estimates of equation (3) using heteroskedasticity-robust standard errors and applying the delta method. The dependent variable in the regression is the natural logarithm of the ratio between the weekly wage of more educated full-time white male workers 40 to 50 years of age and the wage of less educated full-time white male workers 40 to 50 years of age. Relative supplies are constructed including only white male workers between 21 and 59 years of age.

[a] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{CG} + L_{st}^{CD}(\bar{w}^{CD}/\bar{w}^{CG}) + (L_{st}^{HSG}(\bar{w}^{HSG}/\bar{w}^{CG})$.

[b] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{HSG} + L_{st}^{CD}(\bar{w}^{CD}/\bar{w}^{HSG}) + L_{st}^{CG}(\bar{w}^{CG}/\bar{w}^{HSG})$.

[c] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{HSG}$.

* = significant at 10%, ** = significant at 5%, *** = significant at 1%.

school equivalence units to high school dropouts, or the ratio of more educated workers in college equivalence units to high school dropouts.

Table 3, shows that the effect of the child-labor- and compulsory-attendance-law instruments on the decennial changes of the relative supply of more educated workers is of the expected sign. Their joint level of significance varies between 0.1% and 8%. To ensure that our estimates of the long-run elasticity of substitution are as robust as possible to weak-instrument concerns, we implement the limited-information maximum likelihood estimator recommended by Chao and Swanson (2002) as well as the Fuller-modified limited-information maximum likelihood estimator recommended by Stock et al. (2002) and by Hahn and Hausman (2002) in addition to the two-stage least squares estimator.

IV. Estimates

A. Elasticity of Substitution

Table 4 summarizes our estimates of the long-run elasticity of substitution σ between more and less educated workers using measures of supply relying on white male workers only. Standard errors (in parentheses) are obtained by applying the delta method (for example, Ruud, 2000, p. 367) to the distribution of the original estimate $(1/\sigma)$ obtained by estimating equation (3). The three panels correspond to results obtained using least squares estimation (panel A), instrumental variables estimation using the state-of-residence approach (panel B), and instrumental variables estimation using the state-of-birth approach (panel C). The columns correspond to different ways of measuring the supply of more educated workers. Column (1) measures more educated workers in high school equivalence units, column (2) measures more educated workers in college equivalence units, and column (3) measures more educated workers by the (raw) number of high school graduates.

The results in row (i) of panel A refer to least squares estimates of the long-run elasticity of substitution between more and less educated workers and do not take account of state or time fixed effects. The results indicate that a higher relative supply of more educated workers is associated with higher relative wages for more educated workers (because the point estimate of the coefficient is negative). The results in row (ii), obtained using least squares with state and time fixed effects, make clear that the finding of a positive correlation between the relative supply of more educated workers and the education wage premium in row (i) is driven by omitted fixed effects. Once these effects are included in the empirical analysis, a higher relative supply of more educated workers is associated with lower relative wages for more educated workers. The long-run elasticity of substitution between more and less educated workers in row (ii) is around 3 with a standard error around 0.65 (with relatively small variations depending on how the supply of more educated workers is measured). We refer to this estimate as the *long-run* elasticity because estimation relies on 10-year changes in the relative supply of more educated workers and their relative wage.

As the relative supply of more educated workers is likely to be positively correlated with outward shifts in relative labor demand, instrumental variables estimation is preferable to least squares estimation. Panel B gives the results of estimating the long-run elasticity of substitution between more and less educated workers, using compulsory attendance and child labor laws as instruments for the relative supply of more educated workers. The instruments are constructed following the state-of-residence approach. Row (i) contains two-stage least squares estimates of the long-run elasticity of substitution, controlling for state and time fixed effects. It can be seen that the value is less than half of the corresponding least squares estimate, whereas the estimated standard errors are similar in the two cases. This confirms the suspicion that the least squares estimator of the long-run elasticity of substitution is biased upward. As our empirical specification is overidentified, we can test the exogeneity of the instruments [using a version of the Hausman test that allows for heteroskedasticity of the residuals; see Woolridge (2001, p. 123)]. The test does not reject the null hypothesis that all instruments are exogenous at the 5% confidence level no matter how we measure the supply of more educated workers.

Panel B, rows (ii)–(iv), implement three instrumental variables estimators that have been shown to be more robust to weak-instrument concerns than two-stage least squares. The limited-information maximum likelihood estimate of the long-run elasticity of substitution is somewhat smaller but more precise than two-stage least squares estimates. The two Fuller limited-information maximum likelihood estimates are calculated for Fuller constants 4 and 1. The Fuller constant 1 results in the most unbiased estimator and is recommended when one wants to test hypotheses; the Fuller constant 4 minimizes the mean squared error of the estimator (Fuller, 1977). Both Fuller limited-information maximum likelihood estimates are similar to two-stage least squares estimates.

Panel C presents instrumental variables estimates of the long-run elasticity of substitution when the child-labor- and compulsory-attendance-law instruments are constructed following the state-of-birth approach. Row (i) contains the two-stage least squares estimate of the long-run elasticity of substitution, controlling for state and time fixed effects. The point estimates are very similar to or larger than in the corresponding specification using the state-of-residence approach (depending on how we measure the supply of more educated workers), whereas the standard errors are somewhat smaller. This is consistent with the state-of-birth approach being preferable to the state-of-residence approach for predicting the relative supply of more educated workers, but also being more likely to be affected by interstate migration. Implementing the Hausman test of overidentifying restrictions yields that instrument exogeneity cannot be rejected at the 5% confidence level except in column (2), where the supply of more educated workers is measured in college equivalence units (the *p*-value is 7% in this case).

Panel C, rows (ii)–(iv), implement the three instrumental variable estimators that have been shown to be more robust to weak-instrument concerns than two-stage least squares (limited-information maximum likelihood and Fuller limited-information maximum likelihood with Fuller constants equal to 1 and 4, respectively). The estimates are very close to two-stage least squares values, and the standard errors are somewhat smaller. The point estimates of the long-run elasticity of substitution obtained using different instrumental variables specifications and measures of the supply of more educated workers are therefore rather similar and range from 1.2 to 2.

Our preferred estimator is the Fuller limited-information maximum likelihood estimator minimizing the mean squared error using state-of-residence instruments [panel B, row (iv), column (1)], which yields a highly significant long-run elasticity of substitution between more and less educated workers of 1.5, close to the middle of the range of estimates obtained using other instrumental variables estimation methods.

Table 5 reports our estimates of the long-run elasticity of substitution σ between more and less educated workers when we calculate the supply using data on all workers in our sample. Our instrumental variables estimates are now somewhat larger than in table 4, but the differences are always small and statistically insignificant.

B. *Stability of the Elasticity of Substitution over Time*

So far we have assumed the long-run elasticity of substitution between more and less educated workers to be constant over time. We now test this assumption by allowing the elasticity of substitution to differ between the 1950–1970 period and the 1970–1990 period. Using the state-of-residence instruments and measuring more educated workers in high school equivalence units yields a two-stage least squares estimate of the elasticity of substitution of 1.61 with a standard error of 0.85 for the 1950–1970 period and 1.47 with a standard error of 0.71 for the 1970–1990 period. Using the state-of-birth instruments, the two-stage least squares estimate is 1.92 with a standard error of 0.92 for the

TABLE 5.—THE SUBSTITUTION ELASTICITY BETWEEN MORE AND LESS EDUCATED WORKERS, 1950–1990 (CES SPECIFICATION; SUPPLY CONSTRUCTED USING ALL WORKERS)

	Measured Relative Supply of More Educated Workers		
Estimation Method	(1) Supply: All Groups[a]	(2) Supply: All Groups[b]	(3) Supply: 2 Groups Only[c]
Panel A			
(i) LS	−5.55*** (0.31)	−6.25*** (0.39)	−6.25*** (0.39)
(ii) LS with state dummies and time fixed effects	2.04*** (0.33)	2.50*** (0.50)	2.56*** (0.60)
Panel B			
(i) 2SLS with state dummies and time fixed effects (using state-of-residence instruments)	1.53*** (0.66)	1.81** (0.86)	2.38*** (0.53)
(ii) LIML with state dummies and time fixed effects (using state-of-residence instruments)	1.53*** (0.39)	1.85*** (0.41)	2.01*** (0.5)
(iii) Fuller LIML, F-constant = 1, with state dummies and time fixed effects (using state-of-residence instruments)	1.61*** (0.39)	1.81*** (0.41)	2.22*** (0.55)
(iv) Fuller LIML, F-constant = 4, with state dummies and time fixed effects (using state-of-residence instruments)	1.61*** (0.39)	1.88*** (0.42)	2.19*** (0.52)
Panel C			
(i) 2SLS with state dummies and time fixed effects (using state-of-birth instruments)	1.47*** (0.45)	1.69** (0.70)	2.19*** (0.99)
(ii) LIML with state dummies and time fixed effects (using state-of-birth instruments)	1.42** (0.67)	1.64** (0.73)	2.12* (1.30)
(iii) Fuller LIML, F-constant = 1, with state dummies and time fixed effects (using state-of-birth instruments)	1.49** (0.70)	1.66** (0.72)	2.12* (1.30)
(iv) Fuller LIML, F-constant = 4, with state dummies and time fixed effects (using state-of-birth instruments)	1.58** (0.69)	1.69* (0.70)	2.09* (1.12)

Years: 1950, 1960, 1970, 1980, and 1990; 48 U.S. contiguous states; total of 240 observations. The parameters presented and their standard errors are obtained from the estimates of equation (3) using heteroskedasticity-robust standard errors and applying the delta method. The dependent variable in the regression is the natural logarithm of the ratio between the weekly wage of more educated full-time white male workers 40 to 50 years of age and the wage of less educated full-time white male workers 40 to 50 years of age. Relative supplies are constructed including all workers between 16 and 65 years of age.

[a] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{CG} + L_{st}^{CD}(\bar{w}^{CD}/\bar{w}^{CG}) + L_{st}^{HSG}(\bar{w}^{HSG}/\bar{w}^{CG})$.

[b] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{HSG} + L_{st}^{CD}(\bar{w}^{CD}/\bar{w}^{HSG}) + L_{st}^{CG}(\bar{w}^{CG}/\bar{w}^{HSG})$.

[c] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{HSD}$, $H_{st} \equiv L_{st}^{HSG}$.

* = significant at 10%, ** = significant at 5%, *** = significant at 1%.

1950–1970 period and 1.72 with a standard error of 0.63 for the 1970–1990 period. Hence, the point estimates are very similar to those obtained for the 1950–1990 period, and the standard errors are somewhat larger. The hypothesis that the long-run elasticity of substitution has remained approximately constant cannot be rejected at any standard level of significance, and we therefore conclude that the assumption is reasonable. The other instrumental variables estimators yield very similar results.

C. *Comparisons with Previous Estimates of the Elasticity of Substitution*

Table 6 summarizes estimates of the aggregate elasticity of substitution between more and less educated workers obtained in previous studies, focusing on work that attempts to take into account the simultaneity of labor demand and labor supply. Johnson (1970) estimates the elasticity of substitution between more and less educated workers to be 1.34, using a cross section of U.S. states in 1960; he uses average age, share of black residents, and share of urban residents as an instrument for the relative supply of more educated workers. Fallon and Layard (1975) estimate the elasticity of substitution between

TABLE 6.—COMPARISON OF ESTIMATES OF THE SUBSTITUTION ELASTICITY IN THE LITERATURE

Authors and Method	Preferred Estimate	Standard Error
Ciccone and Peri Panel of U.S. states, 1950–1990	1.50	0.44
Johnson (1970) Cross section of U.S. states	1.34	n.a.
Fallon and Layard (1975) Cross section of countries	1.49	0.15
Katz and Murphy (1992) U.S. time-series data	1.41	0.30
Angrist (1995) Palestinian labor force data	2.00	0.19
Murphy, Riddle, and Romer (1998) Canadian time-series data	1.36	0.24
Krusell et al. (2000) U.S. time-series data	1.66	0.63
Caselli and Coleman (2000) Cross section of countries	1.31	0.12

Note: As in most of the literature, the estimated parameter is the reciprocal of the elasticity of substitution ($1/\sigma$); we used those estimates and the delta method to calculate the point estimate and standard deviation of the elasticity of substitution, σ.

TABLE 7.—TRANSLOG ESTIMATES OF α_{HL} AND IMPLIED SUBSTITUTION ELASTICITY

		Measured Relative Supply of More Educated Workers		
Estimation Method	Parameter	(1) Supply: All Groups[a]	(2) Supply: All Groups[b]	(3) Supply: 2 Groups Only[c]
2SLS with state dummies and time trend (using state-of-residence instruments)	$\underline{\alpha}_{HL}$	0.13** (0.06)	0.22*** (0.08)	0.22*** (0.06)
	$\bar{\sigma}_{st}$	1.54** (0.25)	1.93** (0.25)	1.93** (0.25)
2SLS with state dummies and time trend (using state-of-birth instruments)	$\underline{\alpha}_{HL}$	0.12** (0.06)	0.25*** (0.07)	0.24*** (0.05)
	$\bar{\sigma}_{st}$	1.50*** (0.25)	2.03*** (0.29)	2.01** (0.21)

$\underline{\alpha}_{HL}$: Estimated from equation (5).
$\bar{\sigma}_{st}$: Elasticity of substitution between more and less educated workers, calculated using equation (6) evaluated at the national value of the wage share of more educated workers (0.62).
Sample: 1950, 1960, 1970, 1980, and 1990; 48 U.S. contiguous states; total number of observations: 240; heteroskedasticity-robust standard errors in parentheses.
[a] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{\mathrm{HSD}}$, $H_{st} \equiv L_{st}^{\mathrm{CG}} + L_{st}^{\mathrm{CD}}\,(\bar{w}^{\mathrm{CG}}/\bar{w}^{\mathrm{CG}}) + L_{st}^{\mathrm{HSG}}\,(\bar{w}^{\mathrm{HSG}}/\bar{w}^{\mathrm{CG}})$.
[b] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{\mathrm{HSD}}$, $H_{st} \equiv L_{st}^{\mathrm{HSG}} + L_{st}^{\mathrm{CD}}\,(\bar{w}^{\mathrm{CD}}\bar{w}^{\mathrm{HSG}}) + L_{st}^{\mathrm{CG}}\,(\bar{w}^{\mathrm{CG}}/\bar{w}^{\mathrm{HSG}})$.
[c] $\ln(H_{st}/L_{st})$ calculated using $L_{st} \equiv L_{st}^{\mathrm{HSD}}$, $H_{st} \equiv L_{st}^{\mathrm{HSG}}$.
* = significant at 10%, ** = significant at 5%, *** = significant at 1%.

more and less educated workers to be 1.49 using cross-country data; their instrument for the relative supply of more educated workers is income per capita. Angrist (1995) analyzes the relationship between the return to schooling and the supply of more educated workers among Palestinians in the West Bank and the Gaza Strip during the 1980s; his approach exploits the fact that the increase in the supply of more educated workers was mainly driven by the creation of new local institutions of higher education. The elasticity of substitution between more and less educated workers implied by his estimates is 2.[1] Caselli and Coleman (2000) estimate the aggregate elasticity of substitution between more and less educated workers using cross-country data and find a value of 1.31. Katz and Murphy (1992) estimate the aggregate elasticity of substitution between more and less educated workers using U.S. time-series data for the 1963–1987 period. Their identifying assumption is that year-by-year variations in the relative supply of more educated workers are independent of skill-biased technology shocks. Their estimate, which is probably best interpreted as a short-run substitution elasticity, is 1.41. Krusell et al. (2000) also use U.S. time-series data to estimate the short-run aggregate elasticity of substitution between more and less educated workers and find a value of 1.66. Murphy et al. (1998) apply Katz and Murphy's (1992) approach to Canadian time series data and obtain an estimate of 1.36. Hence, our preferred estimate of the aggregate elasticity of substitution between more and less educated workers (1.5) lies in the middle of the range of estimates obtained in previous studies. It is interesting to note that our estimate of the long-run elasticity of substitution is rather similar to estimates of the short-run elasticity of substitution available for the United States. This may be an indication that it is not much easier to substitute less educated workers for more educated workers in the long run than in the short run.

V. Translog Estimates of the Elasticity of Substitution

The constant-elasticity-of-substitution aggregate production function assumes that the elasticity of the relative demand for more educated workers with respect to relative wage of more educated workers is constant along the relative demand curve. This assumption can be relaxed by using a translog specification instead.[2] The translog production function is

$$\ln Y_{st} = \ln \alpha + \alpha_L \ln L_{st} + \alpha_H \ln H_{st} + \frac{\alpha_{LL}}{2}(\ln L_{st})^2 2 + \frac{\alpha_{HH}}{2}(\ln H_{st})^2 + \alpha_{LH} \ln L_{st} \ln H_{st} + \alpha_{BH} \ln B_{st}^{TR} \ln H_{st} + \alpha_{BL} \ln B_{st}^{\mathrm{TR}} \ln L_{st}. \tag{4}$$

Our constant-returns-to-scale assumption implies the following parameter restrictions: $\alpha_L + \alpha_H = 1$, $\alpha_{LL} + \alpha_{LH} = 0$, $\alpha_{HH} + \alpha_{LH} = 0$, and $\alpha_{BL} + \alpha_{BH} = 0$.

Cost minimization and pricetaking in the labor market imply that the share of total wages going to more educated workers, which will be denoted by β_{st}, is equal to the elasticity of output with respect to the efficiency units of more educated workers,

$$\beta_{st} = \frac{w_{st}^H H_{st}}{w_{st}^H H_{st} + w_{st}^L L_{st}} = \frac{\partial \ln Y_{st}}{\partial \ln H_{st}} = \alpha_H + \alpha_{HL} \ln(H_{st}/L_{st}) + \alpha_{BH} \ln B_{st}^{\mathrm{TR}}, \tag{5}$$

where the last equality makes use of the translog production function in equation (4). This is our basic estimating equation for the translog specification. The key parameter, α_{HL}, can be estimated consistently using the same instruments and the same identifying assumptions as in the constant-elasticity-of-substitution case. The elasticity of substitution between more and less educated workers σ_{st} in the translog case can then be obtained as

$$\sigma_{st} = 1 + \frac{\alpha_{HL}}{(1 - \beta_{st})\beta_{st}}, \tag{6}$$

[1] Though Angrist does not estimate the elasticity of substitution between more and less educated workers directly, it is straightforward to back the elasticity out from his estimates [using, for example, the formulas on pp. 26–28 of Hamermesh (1993) for the two-factor CES production function].

[2] Detailed treatments of the translog production function can be found in, for example, Berndt and Christensen (1973), Greene (1997), and Hamermesh (1993).

TABLE 8.—AVERAGE ANNUAL SKILL-BIASED TECHNOLOGICAL PROGRESS, 1950–1990, BY STATE

State	Annual SBTP CES Specification	Annual SBTP Translog Specification	State	Annual SBTP CES Specification	Annual SBTP Translog Specification
Arizona	0.087	0.067	Michigan	0.060	0.048
New Mexico	0.073	0.058	Kentucky	0.060	0.050
Maine	0.071	0.059	Tennessee	0.060	0.053
Arkansas	0.070	0.057	Texas	0.060	0.049
Colorado	0.070	0.048	Massachusetts	0.059	0.040
New Hampshire	0.068	0.053	Nebraska	0.059	0.038
Montana	0.067	0.050	Alabama	0.059	0.050
North Carolina	0.067	0.056	Virginia	0.058	0.047
Pennsylvania	0.067	0.053	Georgia	0.057	0.048
New Jersey	0.066	0.046	Ohio	0.055	0.046
California	0.066	0.046	Wisconsin	0.055	0.042
Florida	0.066	0.053	Missouri	0.055	0.046
Wyoming	0.065	0.052	Iowa	0.055	0.041
Connecticut	0.065	0.043	Vermont	0.054	0.045
Idaho	0.065	0.051	Louisiana	0.052	0.045
Utah	0.064	0.044	Delaware	0.052	0.041
West Virginia	0.064	0.054	Kansas	0.052	0.034
Minnesota	0.063	0.043	Mississippi	0.050	0.044
New York	0.063	0.045	Nevada	0.049	0.039
Oregon	0.062	0.046	Oklahoma	0.048	0.039
Illinois	0.062	0.046	Indiana	0.048	0.041
Maryland	0.062	0.047	Rhode Island	0.047	0.037
South Carolina	0.061	0.052	North Dakota	0.046	0.034
Washington	0.061	0.042	South Dakota	0.039	0.027

where the subscript st makes explicit that the elasticity of substitution varies across states and over time.

Table 7 summarizes estimates of the parameter α_{HL} [obtained estimating equation (5) with two-stage least squares controlling for state and time fixed effect] and of the implied elasticity of substitution evaluated at the U.S. average value for the wage share of more educated workers, $\bar{\sigma}_{st}$. It can be seen that α_{HL} is significantly positive, whether we use the state-of-residence or the state-of-birth approach to construct the instruments. Combined with equation (6), this implies that the aggregate long-run elasticity of substitution between more and less educated workers is greater than unity in all states. The implied values for $\bar{\sigma}_{st}$ are close to the long-run estimates obtained using the constant-elasticity-of-substitution specification. Estimates obtained using the limited-information maximum likelihood and Fuller modified limited-information maximum likelihood methods are similar to two-stage least squares estimates.

VI. An Application: Quantifying Shifts in the Relative Demand for More Educated Workers, 1950–1990

Our constant-elasticity-of-substitution and translog estimates of the slope of the relative demand curve for more educated workers allow us to identify relative labor demand shifts at the U.S. state level for the period 1950–1990. Our conceptual framework associates such shifts with skill-biased technological progress (SBTP). We first identify demand shifts using the constant-elasticity-of-substitution specification, and then using the translog specification.

Combining equation (3) with estimates of the aggregate elasticity of substitution between more and less educated workers allows us to estimate shifts of the relative labor demand for more educated workers (SBTP) for each state, $\Delta \ln B_{st}$, where Δ denotes the difference between adjacent decennial censuses. Table 8 summarizes our estimate of average annual SBTP for the 48 contiguous U.S. states over the period 1950–1990 using our preferred estimate of the substitution elasticity (1.5). It can be seen that many western U.S. states experienced large increases in the relative demand for more educated workers, to the point that SBTP was as fast as 8% per year. Several southern states in contrast had rates of SBTP lower than 5% per year. As U.S. states have access to the same technology, these differences are likely due to the pattern of sectoral specialization. Most of the states that experienced larger SBTP started out with a greater supply of more educated workers in 1950 and have seen fast growth in high-tech sectors since.

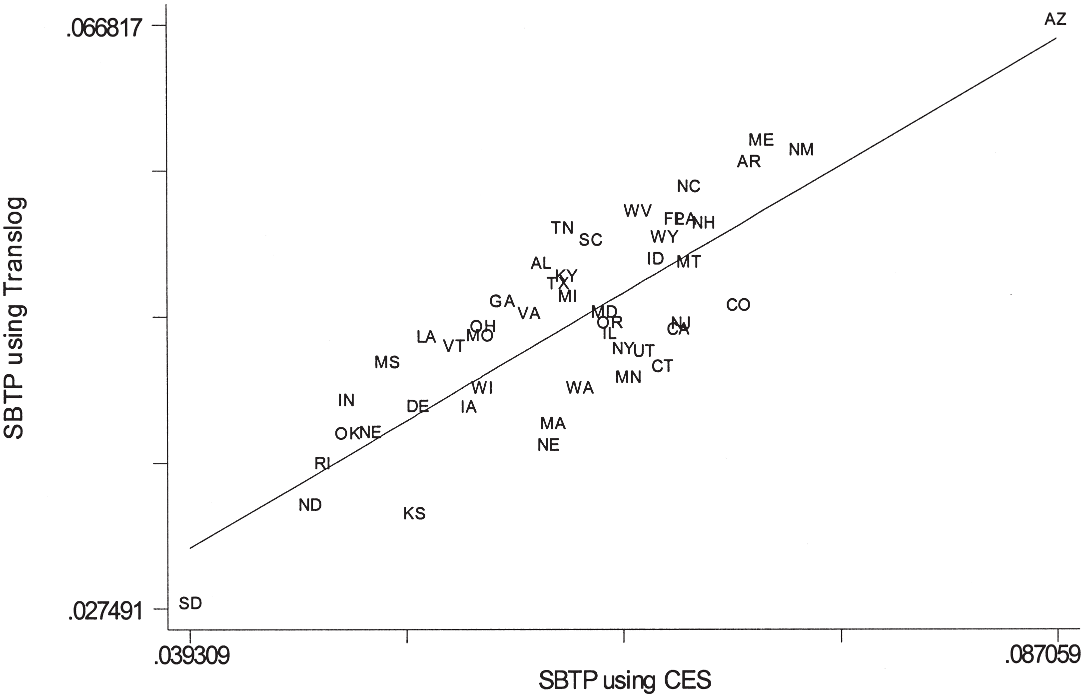
The relative labor demand shifts implied by the translog estimates of the long-run elasticity of substitution between more and less educated workers can be calculated as

$$\Delta \ln\left(\frac{w_{st}^{H}}{w_{st}^{L}}\right) - \frac{1}{\sigma_{st}}\Delta \ln\left(\frac{H_{st}}{L_{st}}\right). \tag{7}$$

where σ_{st} is the state-time specific elasticity of substitution implied by the translog production function [defined in equation (6)].

Table 8 reports our estimates of SBTP as implied by the translog specification of the production function. The results are rather similar to those obtained using the constant elasticity of substitution specification. Figure 1 plots SBTP for each state obtained using the constant elasticity of

FIGURE 1.—CORRELATION OF CES AND TRANSLOG SBTP ACROSS U.S. STATES, 1950–1990



substitution framework against SBTP obtained using the translog framework. It can be seen that the correlation is high (the correlation coefficient is 0.75, and the two methods yield very similar sets of states with slow SBTP and sets of states with rapid SBTP). The main differences arise during the 1980s, where the translog specification yields smaller relative labor demand shifts than the constant-elasticity-of-substitution specification. This is because the wage share of more educated workers has been increasing over time and the translog specification implies that increases in this share (once it is above 0.5) raise the elasticity of substitution. The higher the long-run elasticity of substitution (the flatter the relative labor demand for more educated workers), the smaller the reduction in the education wage premium implied by increases in the relative supply of more educated workers. Hence, smaller shifts in the relative-labor-demand curve for more educated workers are necessary to explain rising education wage premia. As the long-run elasticity of substitution implied by the translog specification for the 1980s (2.33) is considerably larger than the value obtained with the constant-elasticity-of-substitution specification, the implied relative-labor-demand shifts are substantially smaller. As this finding is supported neither by previous studies nor by our constant-elasticity-of-substitution estimates for the 1970–1990 period, we put more weight on the constant-elasticity-of-substitution results for the 1980s.

Table 9 presents our estimates of average annual SBTP across states for each decade between 1950 and 1990 [formally this estimate is obtained as $(\Delta\alpha_t + \Delta u_{st})/10$; see equation (3)], using our preferred constant-elasticity-of-substitution estimate of the long-run elasticity of substitution between more and less educated workers. It can be seen that SBTP accelerated in the 1980s [this finding is consistent with Caselli and Coleman (2002)]. A less well-known result is that there has been rapid SBTP since the 1950s.

VII. Summary

Our main contribution is to provide estimates of the long-run elasticity of substitution between more and less educated workers using data on U.S. states for the period 1950–1990. Our estimates rely on state–time-specific child labor and compulsory attendance laws as instruments for changes in the relative supply of more educated workers and control for state- and time-specific fixed effects. Our preferred estimator yields a point estimate of the long-run elasticity of substitution of 1.5. This implies that a 1% increase in the relative wage of more educated workers reduces relative demand by 1.5%. Or, taking a different perspective, a 1% increase in the relative supply of more educated workers reduces their relative wage by 0.66%.

This estimate of the long-run elasticity of substitution between more and less educated workers is rather robust to

TABLE 9.—AVERAGE ANNUAL SKILL-BIASED TECHNOLOGICAL PROGRESS BY DECADE

Decade	CES Specification
1950s	0.051
1960s	0.061
1970s	0.054
1980s	0.075

a series of variations in the measurement of the relative supply of more educated workers, the construction of the instruments for changes in relative labor supply, and the (instrumental variables) estimation method. Our elasticity estimate is in the middle of the range obtained in previous studies (using either U.S. time series data or cross-country data), despite substantial differences in the estimation methods.

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APPENDIX

1. Physical Capital in the Production Function

Our framework can easily accommodate physical capital as a separate input, as long as this input and the constant-elasticity-of-substitution composite of more and less educated workers enter the production function in a weakly separable way, or formally, as long as the aggregate production function can be written as

$$Y_{st} = F[K_{st}, A_{st}(L_{st}^{(\sigma-1)/\sigma} + B_{st}\, H_{st}^{(\sigma-1)/\sigma})^{\sigma/(\sigma-1)}], \quad \text{(A1)}$$

where K_{st} is physical capital. It is straightforward to show that equation (A-1), combined with cost minimization and pricetaking in the labor market, implies that the relative demand for more educated workers is given by equation (2).

A particular case of equation (A-1) is the (Cobb-Douglas) production function

$$Y_{st} = A_{st} K_{st}^{\alpha_s} (e_{st}^{(\sigma-1)/\sigma} + B_{st}\, H_{st}^{(\sigma-1)/\sigma})^{(1-\alpha_s)\sigma/(\sigma-1)}.$$

This function has the property that the (state-specific) income shares going to capital and to labor (of all education levels) are constant over time and equal to α_s and to $1 - \alpha_s$, respectively. The constancy of labor shares over time implied by this specification turns out to be a reasonable description of U.S. state data for the 1975–2000 period, as we show in the next section.

2. Labor Shares in U.S. States

We adopt the procedure proposed by Gollin (2002) to calculate labor income shares at the U.S. state level. The first step is to impute as labor income all the wage and salary income of employees. Then we calculate the average labor income of employees, and we impute to the self-employed the same average labor income. The sum of measured labor income of employees and imputed labor income of the self-employed is used as a measure of total labor income. Dividing total labor income by total income gives us an estimate of the labor income share at the state level. State-level data on total income, employees' wages, and income of the self-employed are available from the Bureau of Economic Analysis (2004), *National Income and Production Accounts* for 1975–2000. We then use the state-level labor income shares over this period to check whether labor income shares have trended upward or downward. We cannot reject the hypothesis that labor income shares have no such trend at the 5% level for 45 out of 48 states. Though there are a few outliers (Alaska and Wyoming with low labor shares and the District of Columbia with high labor share), 40 states have labor shares between 0.67 and 0.72 over the whole period. Details are available upon request.

3. Data on Workers and Wages

The paper uses data from the 1950, 1960, 1970, 1980, and 1990 IPUMS files in order to calculate the relative supply of skills and relative wages. The sample used is exactly the same as in the work by Acemoglu and Angrist (2000) and kindly provided to us by the authors. We exclude the noncontiguous states (Alaska and Hawaii) and the District of Columbia. The wage observations are weighted by the IPUMS weighting variable in order to obtain state averages. The schooling attainment of individuals are divided into four groups (high school dropouts, high school graduates, college dropouts, and college graduates) using the variable *HIGRADED* for the 1950–1980 data and the variable *YEARSCH* for the 1990 census. The wage variable used is the weekly wage, in current dollars, obtained by dividing the yearly wage (wage and salary income) by the number of weeks worked. Wages are top-coded uniformly across census years (the censoring is at the 98th percentile times 1.5). The wage of a high school (college) efficiency unit of labor is measured as total wages of workers with at least a high school degree in state s and year t divided by the supply of more educated workers in high school (college) efficiency units. The data on child labor and compulsory attendance laws are described in detail in Acemoglu and Angrist (2000).