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The Intergenerational Transmission of Human Capital: Evidence from the Golden Age of Upward Mobility

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School quality affects upward mobility in educational attainment. This conclusion comes from an analysis of families with coresident teenage children in the 1940 census. We study parents in the bottom quartile of the education distribution and define "upward mobility" as a generational move up the educational ladder to the top three quartiles of the child's cohort. At the state level, upward mobility is strongly tied to teacher wages. This relationship holds when we narrow our focus to families on adjacent sides of state borders in the South, where state minimum salary laws created sharp teacher-wage differences between otherwise similar counties.

I. Introduction

Societies aspire to equality of opportunity—the goal that all children have a chance to achieve a prosperous and rewarding life. Public schooling can play a key role in the pursuit of this objective. In the United States, widespread

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availability of public elementary schools in the early twentieth century opened a pathway to prosperity for many children. And the high school movement then expanded access even further, enabling the United States to jump ahead of other nations in the share of children with a secondary education (Goldin and Katz 2008).[1] However, during this "golden age" of expanding educational opportunity, there was remarkable inequality in families' access to high-quality education. Black children in most areas of the South were educated in poorly resourced segregated schools (Margo 1990; Card and Krueger 1992a). Many white children in poorer communities also faced limited opportunities for high-quality schooling.

A long-established body of research suggests that greater school resources induce students to remain in school longer, thus potentially moving children up the education ladder relative to their parents.[2] Even so, there is some skepticism in the literature concerning the effectiveness of input-based education policies (e.g., Hanushek 2006). Moreover, few studies look specifically at the impacts of schooling resources on children from relatively disadvantaged families—a pivotal issue in terms of understanding the impacts of schooling quality on intergenerational mobility. It is conceivable, after all, that the benefits of improvements in local schools accrue mainly to families of higher socioeconomic status and have little effect on children from families at lower rungs of the ladder.

In this paper we use matched parent and child records from the 1940 census to study the impact of school resources on the upward mobility of families in the national distribution of human capital. Specifically, we focus on children of native-born parents in the first quartile of educational attainment (6 years of completed schooling or less). We ask whether these children had completed at least eighth grade by age 16–18, putting them at or above the second quartile of the education distribution for their cohort—and thus ahead of their parents. Descriptively, we show that there was wide variation across

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[1] The fraction of 14–17-year-olds enrolled in high school rose from 10% in 1900 to more than 70% in 1940 (Snyder 1993, table 9).

[2] See Card and Krueger (1996) for an earlier review of the evidence. Some of the most compelling evidence examines the effects of school resources for Black students in the South, building on insights of Bond (1934). Collins and Margo (2006) review some of this evidence. Aaronson and Mazumder (2011) attribute a sizable share of the gains in Black education and literacy in the 1920s and early 1930s to the impact of the Rosenwald school program; see also Carruthers and Wannamaker (2013) and Aaronson et al. (2021).

states in this simple index of upward mobility, with rates as high as 90% for white and Black children in California and several Upper Midwest states and as low as 10% for Black sons in Mississippi and South Carolina.

In models that control for family characteristics and contextual factors, we find that upward mobility rates for families of both races were systematically higher in states with higher average teacher wages and lower average pupil-teacher ratios. While suggestive, these cross-state comparisons may be confounded by unobserved characteristics that are correlated with both state-level average school resources and educational choices of less educated families. We therefore narrow our focus to the effects of teacher salaries on children from adjacent counties along the borders of the southern states. These comparisons reduce or eliminate many local factors that could simultaneously affect educational outcomes and school resources. We also leverage the fact that many southern states had minimum teacher salary regulations that pushed up average teacher wages, leading to large differences in average teacher pay between counties on opposite sides of state borders.

Our analysis begins with an examination of child-parent coresidency and school enrollment among teenagers in the 1940 census. This census—the first to ask about education and labor earnings—administered a "long-form" questionnaire to 100% of the population, providing information on very large samples of families from all parts of the country. We find that most children remained with their parent(s) until at least age 18 and left school before leaving home. These patterns mean that we can use educational outcomes for coresident youth to infer the fractions of children who have completed at least intermediate levels of schooling, as in Goldin and Katz (1999).

Focusing on families where the best-educated parent had at most six grades of schooling, we estimate the fraction of coresident children aged 16–18 who have completed eighth grade. We show that this rate varies widely by race and region, with higher average rates for whites than Blacks and relatively low rates for children of both races in the South. We also show that county-level estimates of the fraction of children from poorly educated families who had achieved at least eight grades of schooling are highly correlated with estimates of the rate of upward mobility in income estimated by Chetty et al. (2014a) for children born in the early 1980s ($\rho = 0.55$), implying strong persistence in local factors that determine intergenerational mobility.

Next we present a series of models of the relationship between school quality and upward mobility at the state level, conducting the analysis separately for white daughters and sons using data from the 49 mainland states (including the District of Columbia) and for Black daughters and sons using data from the 18 segregated southern states. To address concerns that differences in average teacher wage could be due to differences in the fraction of high school versus elementary teachers and might indirectly incorporate a high school access effect, we use statewide minimum teacher wages as an instrumental variable (IV) for the average wage. Ordinary least squares (OLS) and IV models yield similarly sized and highly significant partial correlations between average teacher wages in a state and our measure of upward mobility, suggesting that access to better-paid teachers was an important determinant of schooling for children of lower-educated families, even controlling for detailed parental characteristics and statewide differences in average income, education, and home values.

Finally, we present our analysis of families living in counties on adjacent sides of state borders in the South. These cross-border comparisons hold constant many local factors, including differences in attitudes about the value of education, that might confound a cross-state analysis. Here, we merge data for families with information on mean wages reported by teachers in the same county in the 1940 census. We show that teacher wages varied widely across state borders, despite the similarity of observable economic conditions between adjacent counties. A key factor driving these differences were statewide minimum wage statutes. For example, the minimum teacher wage was $525 per year in Kentucky versus only $320 per year in Tennessee (in today's dollars, approximately $10,300 and $6,300, respectively). In Kentucky and Tennessee, the minimum wage statutes applied to both white and Black teachers, but in many other southern states the laws specified lower teacher wages for Black teachers. In general, these laws led to large differences in teacher wages in communities that were only a few miles apart. Using the cross-border difference in statewide minimum teacher salaries as an IV for the cross-border difference in average teacher wages, we obtain estimated effects of teacher salaries on upward mobility rates that are very similar to those obtained in parallel OLS models and similar as well to those from our cross-state analysis.

Taking the evidence as a whole, we conclude that school resources—and particularly higher teacher salaries—were an important contributor to gains in the relative educational attainment for children of less educated families during the golden age of expanding educational opportunity and rising upward mobility in the United States. Our work similarly shows that the persistent underfunding of Black education in the segregated South—especially in terms of low wages for Black teachers—accounted for much of the Black-white gap in upward mobility rates, helping to explain the persistence of the human capital gap between Blacks and whites.

Our paper contributes to a growing literature on the determinants of upward mobility between generations. Social scientists have studied this issue extensively (for reviews, see Solon 1999; Black and Devereux 2011), and important work by economic historians has traced trends in intergenerational mobility in the United States broadly (Ferrie 2005; Long and Ferrie 2013; Feigenbaum 2018) and more recently for the African American population (Collins and Wanamaker 2017). A prominent series of recent studies by Chetty et al. (2014a, 2014b) has shown that the rate of intergenerational mobility for children born in the early 1980s varies widely across areas of the United States and is correlated with measures of school quality and other local factors.[3] We document similar variation in upward mobility in human capital accumulation in the mid-twentieth century. We extend the literature by isolating the causal effect of local school resources (specifically average teacher salaries) in determining the degree of upward mobility for children from families at the bottom of the human capital distribution.

We also contribute to an important literature on the schooling of African American youth in the segregated South, including Margo (1985, 1986, 1990), Donohue, Heckman, and Todd (2002), Carruthers and Wanamaker (2013, 2017), Collins and Margo (2006), and Aaronson and Mazumder (2011). As in many of these studies, our main analysis relies on local measures of school quality in the South. By focusing on the schooling attainment of youth who were still living with their parents, we extend the literature by showing how local school quality affected the educational choices of children holding constant parental background. Our border-design IV models isolate the effect of differences in teacher wages coming from statewide minimum teacher salary laws, thus strengthening the case for a causal role of school resources on the intergenerational mobility of schooling attainment.

II. Background and Descriptive Analysis

A. Historical Setting

We use data from the 1940 census to study the effects of school quality on intergenerational mobility in human capital investments. A key feature of this setting is the remarkable degree of inequality in school resources available to children in different locations. Between different states there were large differences in the average levels of several measures of school quality, with particularly low levels in the South (see, e.g., Card and Krueger 1992b). Even among the 18 southern states with segregated schooling systems there were large differences in average teacher salaries and class sizes for both white and Black students (Card and Krueger 1992a).

Within states there was also inequality in school resources between richer and poorer school districts, reflecting the limited role of state equalization policies.[4] However, as we discuss in more detail in appendix B, by the late 1930s a

[3] Earlier US-based work (e.g., Solon 1992; Mazumder 2005) relied on relatively small samples, making it impossible to document differences in mobility rates at the local level. A recent body of work uses larger samples of families in the Nordic countries (e.g., Black, Devereux, and Salvanes 2005; Meghir and Palme 2005; Aakvik, Salvanes, and Mazumder 2010; Lundborg, Nilsson, and Rooth 2014; Carneiro et al. 2021). Most of this work focuses on family-specific shocks or on national school reforms.

[4] Nationwide, the average share of school spending derived from local taxes was 83% in 1930 and 68% in 1940. See Bensen and O'Halloran (1987) for an overview of historical trends. Wallis (2000) provides a broad historical overview of fiscal centralization in the United States, while Coen-Pirani and Wolley (2018) give an economic analysis of changes in fiscal centralization during the 1930s.

total of 23 states had adopted statewide minimum teacher salary laws, including 10 in the South. These laws were generally part of broader legislation that provided some additional funding to local school districts that agreed to pay minimum teacher salaries.[5] In some southern states the minimum teacher salary was lower for teachers in Black schools (e.g., Alabama and Georgia), while in others the minimum was the same for both races (e.g., Kentucky and Tennessee). An interesting variant was legislation in North Carolina, where a statewide fund was set up to pay for teacher salaries (with different schedules for white and Black teachers). These laws created relatively large differences in teacher salaries across state borders; we exploit this variation in the analysis presented in section III.

B. Coresidency and Enrollment in 1940

We link parent and child education using information for children who were residing with their parent(s) at the time of the 1940 census. To limit potential selection biases associated with the coresidency restriction, we focus on children no older than 18.[6] This approach is viable because most children remained with their parents until at least age 18. To illustrate, figure 1*A* shows the fraction of children of different ages coresiding with at least one parent. In this figure and throughout the remainder of paper, we limit attention to US-born children whose parents were also born in the United States.[7]

Looking first at coresidency rates for white sons, we see that the fraction residing with a parent is very stable between the ages of 5 and 17 before declining slightly at age 18.[8] A similar pattern is apparent for Black sons, although at all

[5] For example, Tennessee established a state education equalization funding program in 1925 that required local school districts to provide an 8-month school term and a minimum teacher salary schedule. Bergeron et al. (1999) and Fitzgerald (2007) offer interesting histories of this legislation, which was passed after the governor agreed to support legislation banning the teaching of evolution in public schools.

[6] Goldin and Katz (1999) also use children aged 14–18 in their analysis of coresident children in the 1915 Iowa census. Hilger (2017) focuses on educational outcomes among coresident children aged 26–29. The advantage of this approach is that most people complete their education by age 26. The disadvantage, as Hilger discusses, is the potential for selection bias.

[7] To measure coresidency rates for people whose parents were born in the United States (but who were not necessarily living with their parents), we use the subset of "sample line" individuals who were asked supplementary questions in the 1940 census, including the place of birth of their parents. For all subsequent analyses of coresident children, we obtain their parents' place of birth directly from the parents' records.

[8] We use the Integrated Public Use Microdata Series (IPUMS) version of the 1940 census files (Ruggles et al. 2010), which includes a hierarchy of potential links between the records of children and those of potential mothers and fathers in the same household. To minimize errors in the assignment of parental education, we use the most stringent ("unambiguous") link and consequently count many children living in multigenerational households as not living with a parent. The average coresidency

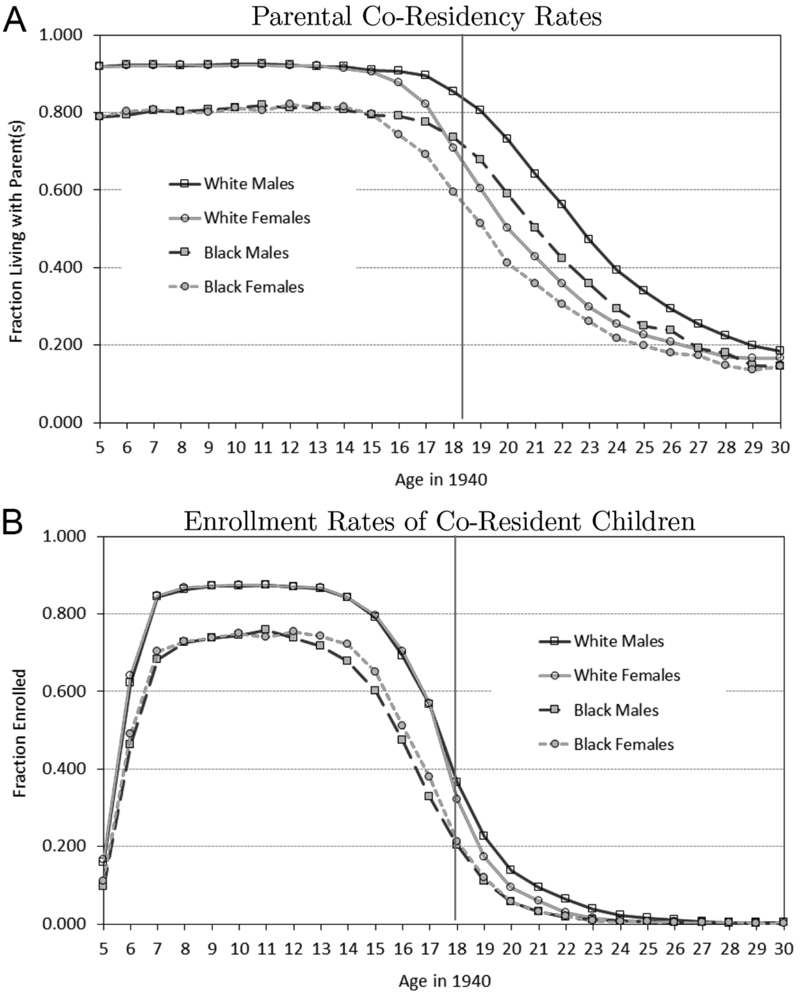

FIG. 1.—Proportion of individuals aged 5–30 living with a parent and enrolled in school. A color version of this figure is available online.

ages the fraction living with at least one parent is about 10 percentage points lower than for white sons.[9] On the basis of these patterns, we infer that there is little threat of selection bias in conditioning on parental coresidency for males up to the age of 18.

rate for white children aged 5–13 rises from 92.7% using only the unambiguous link to 96.9% using all potential links.

[9] Among Black children aged 5–13, 81.4% can be linked to a coresident mother or father using IPUMS's unambiguous intrahousehold linkage criterion. This rate rises to 89.8% including other potential links.

Daughters' coresidency rates for each race are nearly identical to those of sons at ages 5–15. Starting at age 16, however, daughters' rates start to decline, falling by 10 percentage points between ages 16 and 17 and another 10 points between 17 and 18. These profiles suggest that for females there is little threat of selection bias from conditioning on parental coresidency up to age 16 or even 17 but more room for concern in extending the range to age 18.

With that caveat in mind, figure 1*B* shows the age profiles of school enrollment rates of coresident children by gender and race. The rates for white sons are relatively stable between ages 8 and 13, then fall off steadily.[10] Enrollment rates of white daughters are nearly identical to those of sons up to age 17 but fall off a little faster after age 18, reflecting lower college attendance rates of women than men in the pre–World War II period (e.g., Snyder 1993, fig. 14). By age 16 about 30% of white children and 50% of Black children have stopped schooling, and by age 18 more than 60% of white children and 80% of Black children are no longer enrolled. Assuming that the school-leaving age is inversely related to years of completed schooling, these patterns imply that we can infer the probability of having schooling attainment above the level of about the 30th percentile for whites and above the 50th percentile for Blacks by looking at schooling outcomes of children age 16 or older.

For our primary measure of schooling attainment, we therefore use the fraction of 16–18-year-old coresident children who report that they have completed at least eight grades or are enrolled at the time of the census and have completed six or seven grades.[11] For simplicity, we refer to this measure as "having attained eighth grade." As a second, complementary measure, we use years of completed schooling of coresident 14–18-year-olds. Since many 14–18-year-olds were still enrolled when enumerated in the census, we use tobit models to analyze this measure, treating enrollees as right censored. Like our primary measure of eighth-grade attainment, this measure is driven by variation in the lower tail of schooling attainment, although it incorporates high school attainment (whereas our primary measure does not). This gives the alternative measure more statistical power, although as we show the two measures are very consistent with each other across different samples and specifications.

Figure 2*A* and figure 2*B* show how our measure of eighth-grade attainment varies with parental education and region for white sons and daughters. Here, as in the remainder of the paper, we measure parental education by the maximum of mother's and father's education. Outside the South, eighth-grade

[10] There are a number of explanations for nonenrollment of children aged 8–13 who would typically be covered by compulsory schooling laws, including homeschooling, serious disabilities that limited school participation, noncompliance with the law, and miscoding.

[11] Since the enumeration was conducted in April, most of those with seven grades would presumably go on to finish eighth grade. The group with six grades is more uncertain, but as discussed below, dropping this group makes virtually no difference to our findings.

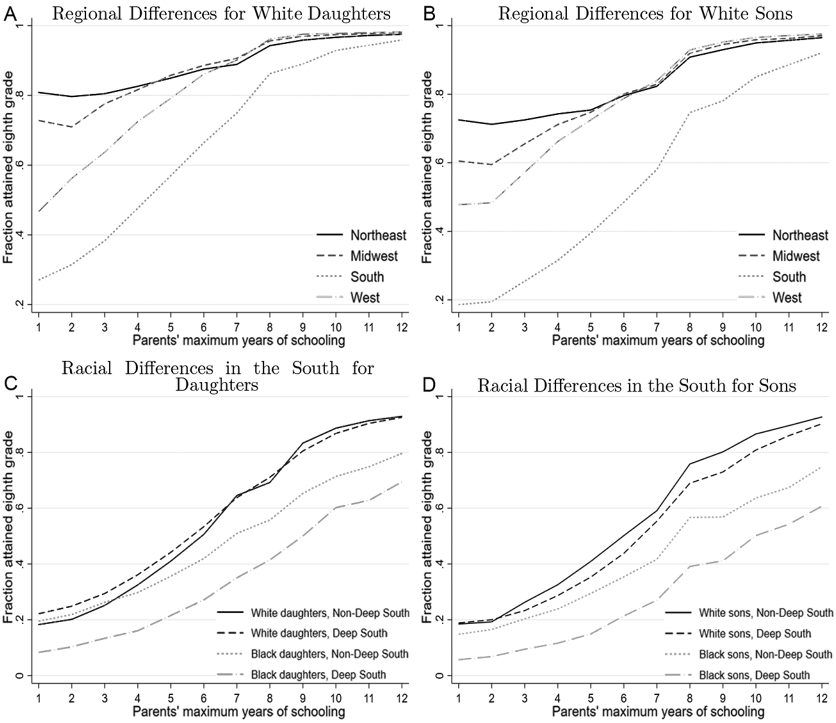

FIG. 2.—Relationship between parent and child education, children aged 16–18 in 1940. A color version of this figure is available online.

attainment rates of children whose best-educated parent has ≥8 years of schooling are very high (95% or higher). For children of less educated parents, however, the fraction attaining eight grades by age 16–18 is lower and varies across regions. Given this pattern, we focus on eighth-grade attainment rates for children whose best-educated parent has 6 years of schooling or less. This cutoff represents the 25th percentile of education among parents of 16–18-year-olds, whereas eighth grade represents the 25th percentile of completed education for their children. Thus, the eighth-grade attainment rate of children whose parents have ≤6 years of schooling represents the fraction of children from families in the bottom quartile of the parental education distribution who achieve a level of schooling in the second or higher quartile for their generation (i.e., is a measure of upward mobility for the bottom quartile of families).

Figure 2*C* and figure 2*D* show eighth-grade attainment rates for white and Black children in the South, distinguishing between families in the Deep South (Alabama, Louisiana, Georgia, Mississippi, and South Carolina) and the other southern states. For whites, eighth-grade attainment rates are not very different in the two sets of states conditional on parental education. For

Black families, however, attainment rates are much lower in the Deep South states, consistent with the much lower quality of schools for Black children in these states. Within both regions of the South, the "gradient" relating parental education and eighth-grade attainment is also flatter for Blacks than for whites, a feature that could be driven in part by the higher rate of mismeasurement of education among Black adults in the 1940 census documented by Margo (1986). Given this potential problem, in the analysis below we mainly focus on the broad group of parents with ≤6 years of schooling without distinguishing between education categories.

Table 1 presents a more detailed summary of alternative measures of educational attainment for teenagers in the 1940 census. Panel A of the table shows data for 16–18-year-olds, while panel B shows data for the broader sample of 14–18-year-olds. Within each age group we show statistics for all children in the relevant age range, regardless of residency status (subpanel a); coresident children (subpanel b); and coresident children whose best-educated parent has at most 6 years of schooling (subpanel c). Finally, to set the stage for our border county analysis, subpanel d shows data for coresident children with poorly educated parents who live in counties along the borders between contiguous southern states.

Consistent with the patterns in figure 1*A*, the first row in panel A(a) shows that approximately 90% of 16–18-year-old white sons and 80% of Black sons were living with at least one parent, while the rates for daughters are roughly 10 points lower. The remaining rows in the panel show four (mutually exhaustive) indicators of schooling status: completed eight grades or more, completed six or seven grades and enrolled, completed less than six grades and enrolled, and not enrolled with less than eight grades completed. We use the sum of the first two as our measure of eighth-grade attainment. The other two groups—those who have left school without completing eighth grade and those who are enrolled but have completed only up to five grades—are classified as nonattainers.

Panel A(b) shows schooling outcomes for the subset of coresident 16–18-year-olds. The schooling outcomes for this subgroup are very similar to those of the overall population in panel A(a), suggesting that selective home leaving does not lead to large biases in measures of schooling attainment. By comparison, coresident children of poorly educated parents, in panel A(c), have substantially worse schooling outcomes, reflecting the strong intergenerational correlation in education documented in figure 2.

Combining students who have actually completed eighth grade plus those who have completed six or seven grades and are enrolled, our measure of "eighth-grade attainment" for children of poorly educated parents has a mean of 70% for white females and 56% for white males, versus a mean of 51% for Black females and 33% for Black males. Part of the disadvantage for Black children is their relative concentration in the South, where schooling completion rates of both races are lower. Even in the South, however, the

Black-white gap in upward mobility rates was 13 percentage points for females (60% for whites vs. 47% for Blacks) and 15 percentage points for males (44% for whites vs. 29% for Blacks).

One concern with our assumption that enrollees with 6 or 7 years of schooling will eventually complete eighth grade is that this has a bigger effect on Blacks than whites, narrowing the Black-white schooling attainment gap relative to the gap in the fraction who have actually completed eight grades in the census. Importantly, however, the models presented below linking school quality to upward mobility are all estimated separately by race, so what matters is whether the relative size of this group is correlated with school quality within race groups. In fact, the correlations are negligible, and all of the estimated school quality effects reported in section III are very similar when we adopt alternative choices, such as assuming that only children who have seven grades will eventually complete eighth grade or that enrollees with five grades will also eventually complete eighth grade.

Panel B of table 1 presents some summary statistics for 14–18-year-old children. As expected, coresidency rates for this broader age group are slightly higher than those for 16–18-year-olds, particularly for females. We show two measures of schooling: median grade attained and enrollment status. For males both outcomes are very similar for all 14–18-year-olds (panel B[a]) and coresident 14–18-year-olds (panel B[b]), but for females there is a bigger gap, particularly in the South. Enrollment rates and median education levels are again substantially lower for children of poorly educated parents (panel B[c]) than among all coresident children. Even among this group, however, a relatively high fraction are enrolled, motivating our use of a censored regression model (tobit) for their years of completed education in section III.

C. Geographic Variation in Eighth-Grade Attainment Rates of Children from Less Educated Families

Figure 3 illustrates the cross-state variation in our measure of upward mobility in human capital for daughters and sons of each race. Figure 3*A* and figure 3*B* show that among white sons, the upward mobility rate varies from 0.27 in Virginia to 0.90 in Utah, while corresponding mobility rates for white daughters are generally slightly higher (and are highly correlated across states with those of sons). As expected, given the patterns in table 1 and figure 2, upward mobility rates for Black daughters and sons (figure 3*C*, 3*D*) are lower than those for whites, particularly in the Deep South. In South Carolina and Mississippi, for example, the upward mobility rate for Black sons is around 0.10. In contrast, upward mobility rates for Black children in some northern and western states are about the same as the rates for white children (e.g., about 0.90 for Black sons in Minnesota, Iowa, and California).

One issue for interpreting these cross-state differences, especially for African American families, is that upward mobility rates may be higher for

Table 1
Characteristics of Youths Age 14–18 in the 1940 Census

	White				Black			
	All States		Southern States		All States		Southern States	
	Females (1)	Males (2)	Females (3)	Males (4)	Females (5)	Males (6)	Females (7)	Males (8)
	A. 16–18-Year-Olds							
	a. All							
Fraction living with ≥1 parent	.84	.91	.78	.90	.69	.79	.68	.79
Fraction with ≥8 grades	.88	.83	.75	.64	.50	.37	.42	.27
Fraction not enrolled, <8 grades	.09	.12	.20	.28	.35	.45	.40	.52
Fraction enrolled, 6–7 grades	.02	.03	.04	.05	.09	.09	.11	.10
Fraction enrolled, <6 grades	.01	.01	.01	.02	.06	.09	.08	.10
	b. Living with One or Both Parents							
Fraction with ≥8 grades	.89	.82	.78	.65	.52	.37	.44	.28
Fraction not enrolled, <8 grades	.08	.13	.16	.27	.29	.43	.33	.50
Fraction enrolled, 6–7 grades	.02	.04	.04	.06	.11	.10	.13	.11
Fraction enrolled, <6 grades	.01	.01	.02	.02	.08	.10	.09	.11
	c. Living with One or Both Parents, Parental Education <6 Grades							
Fraction with ≥8 grades	.64	.50	.53	.37	.38	.23	.33	.19
Fraction not enrolled, <8 grades	.28	.40	.37	.52	.38	.55	.41	.59
Fraction enrolled, 6–7 grades	.06	.06	.07	.07	.13	.10	.14	.10
Fraction enrolled, <6 grades	.02	.03	.03	.04	.11	.12	.12	.13

	d. Border County Sample: Living with One or Both Parents, Parental Education <6 Grades							
Fraction with ≥8 grades			.49	.33			.31	.16
Fraction not enrolled, <8 grades			.40	.55			.43	.60
Fraction enrolled, 6–7 grades			.07	.07			.15	.10
Fraction enrolled, <6 grades			.03	.04			.12	.14
	A. 14–18-Year-Olds							
	a. All							
Fraction living with ≥1 parent	.87	.92	.83	.91	.74	.80	.73	.80
Median grade attained	9	9	9	8	7	6	6	5
Fraction enrolled	.71	.72	.63	.63	.60	.56	.58	.53
	b. Living with One or Both Parents							
Median grade attained	9	9	9	8	7	6	7	5
Fraction enrolled	.76	.72	.71	.65	.67	.58	.65	.55
	c. Living with One or Both Parents, Parental Education <6 Grades							
Median grade attained	8	7	7	6	6	5	6	5
Fraction enrolled	.58	.51	.54	.46	.62	.52	.61	.50
	d. Border County Sample: Living with One or Both Parents, Parental Education <6 Grades							
Median grade attained			7	6			6	5
Fraction enrolled			.54	.46			.60	.49

NOTE.—The sample is restricted to US-born children, and samples with parents exclude those with non-US-born parents. In two-parent households, parental education is defined as years of schooling of the better-educated parent.

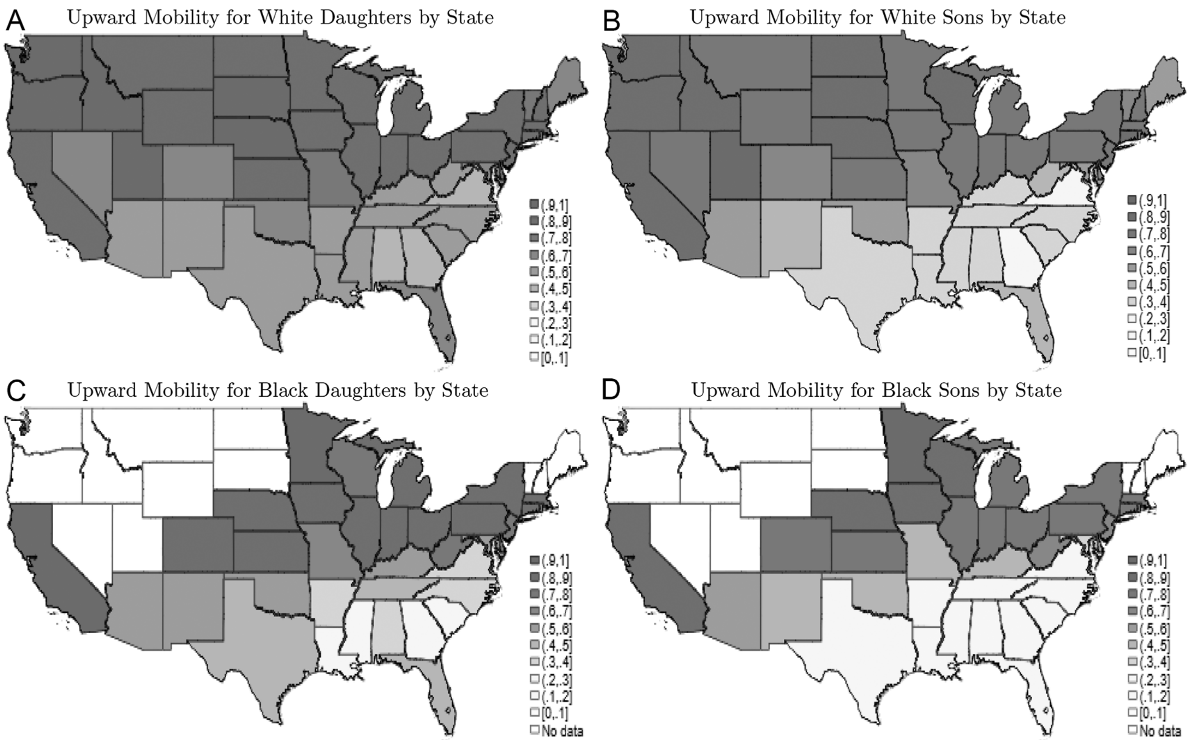

FIG. 3.—Proportion with eight grades of education or more, children aged 16–18 whose parents have 0–6 years of education. A color version of this figure is available online.

children of lower-educated parents who have themselves moved from other parts of the country, reflecting differences in unmeasured skills or attitudes between migrating and nonmigrating parents.[12] We address this in the models below by including controls for whether a child's parents were living in a state other than their state of birth in 1940. Moreover, for Black families we use only comparisons within the South, where a relatively high fraction of parents were born in their state of residence.

Figure 4 presents a finer-grained geographical comparison, giving maps of county-level upward mobility rates for sons and daughters. As a point of departure, figure 4*A* shows upward mobility rates in human capital for all 16–18-year-olds with low-educated parents, pooling children of both races. Average upward mobility rates are clearly lower in the South than in the Northeast or Midwest, with a relatively sharp boundary between Pennsylvania, Ohio, Indiana, and Illinois (to the north) and Maryland, Virginia, and Kentucky (to the south). For comparative purposes, figure 4*B* shows a county-level map of intergenerational mobility rates in income for children of all races born during the period 1980–83, using measures constructed by Chetty et al. (2014a).[13] Visually, the maps in figure 4*A* and 4*B* are quite similar, with lower mobility rates in the South, Arizona, and New Mexico and high mobility rates in the Midwest. Indeed, the cross-county correlation between the two measures of mobility is 0.57, suggesting a high degree of persistence in local factors affecting intergenerational mobility in the United States.

A concern with mobility rates based on averages for whites and Blacks is that the variation is driven by differences in the local fraction of Black children. To address this, we show upward mobility rates at the county level for Black (fig. 4*C*) and white (fig. 4*D*) children (omitting counties with fewer than 30 children of the associated race group). In fact, the geographic variation in overall mobility rates is very similar to the variation for whites only. Moreover, rates for whites and Blacks generally move together, although there are exceptions (e.g., West Virginia appears to have been relatively better for upward mobility of Black children than for white). One can also see a visually clear set of higher-mobility counties for Black children in a belt from Chicago to Detroit.

[12] In an earlier version of this paper (Card, Domnisoru, and Taylor 2018), we showed that upward mobility rates for children of parents who were born outside the United States are substantially higher than those of children with native-born parents (see also Card, DiNardo, and Estes 2000). Butcher (1994) presents an interesting comparison between Black immigrants to the United States and native-born African Americans who have or have not migrated across states, concluding that mobility status, rather than immigration status, is a key distinguishing feature of Black adults in the 1980 census.

[13] These maps use a measure by Chetty et al. that gives the predicted income percentile (at age 26) for children born to parents at the 25th percentile of the income distribution.

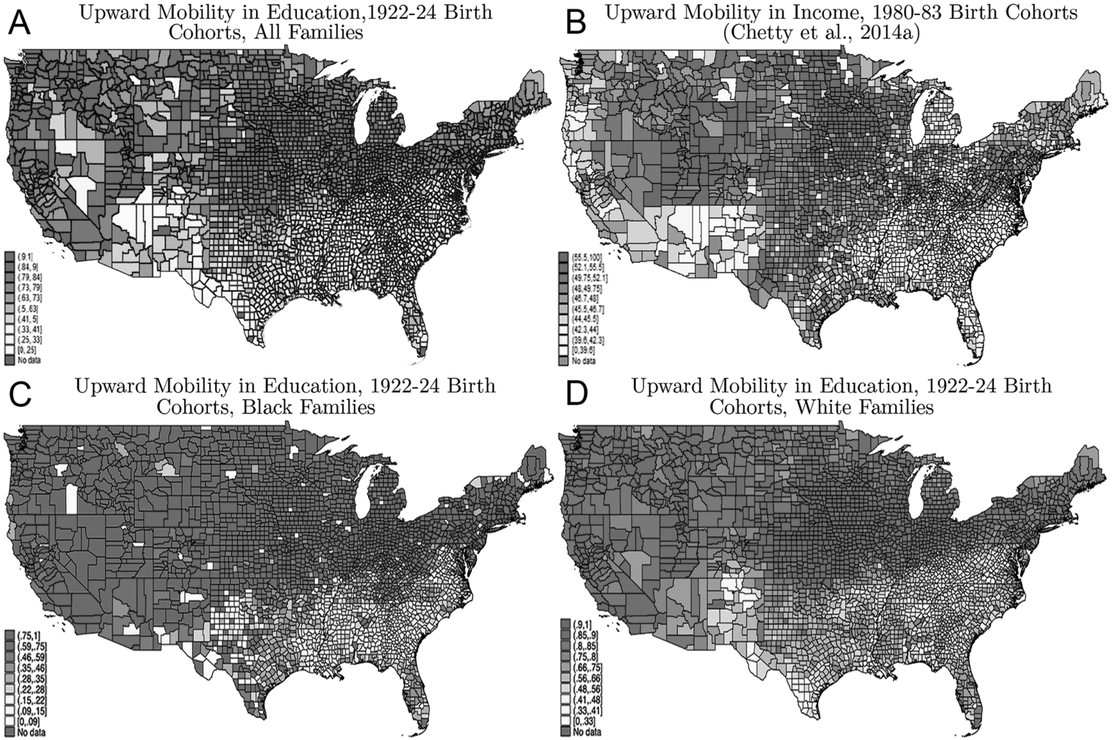

FIG. 4.—Proportion with eight grades of education or more, children aged 16–18 whose parents have 0–6 years of education. A color version of this figure is available online.

III. Upward Mobility and Schooling Quality

In this section we analyze the connection between upward mobility in human capital and the quality of local public schools. We begin with a series of models at the state level. We then move to an analysis at the county level, using data for families living in counties adjacent to a border between two of the segregated southern states.

A. State-Level Analyses

Measures of School Quality

For our state-level analysis we use information on average teacher wages and pupils per teacher originally assembled by Card and Krueger (1992a, 1992b).[14] Card and Krueger's data set includes biannual observations for each state. We assign an average teacher wage and an average pupil-teacher ratio to each child by using an unweighted average of the observations for his or her state of residence over the period from age 6 to age 14. For example, a child aged 16 in 1940 is assigned an average of the quality variables for reports available in 1930, 1932, . . . , 1938.

In their cross-state analysis, Card and Krueger (1992b) adjusted teacher wages in different states and years by dividing by the regional average wage of workers hired in federal construction projects. Since our focus is on a narrower cohort of students, we adopt a different approach and deflate mean teacher wage assigned to each student by a state-level index of earnings for nonteachers with at least some college education, derived from the 1940 census.[15] Figure A1*A* plots the adjusted mean teacher wage for 1940 against the unadjusted mean for 1940. Although the adjustment has some effect, the two series are very highly correlated ($\rho = 0.89$), and all of the estimates reported below are changed only slightly if we use unadjusted teacher wages rather than the adjusted data. The high correlation reflects the fact that most of the variation in average teacher wages is driven by variation in the wages of teachers relative to other workers with some college education. Consistent with this observation, teacher wages and average teacher education were

[14] Card and Krueger also compiled data on average term lengths by state. As in their analysis, however, we find that term length has no significant effect on upward mobility rates or mean schooling levels once we condition on the other two measures. For simplicity, we therefore ignore term length.

[15] To be more precise, we begin with the sample of white workers aged 22–65 who (i) had at least 1 year of college education, (ii) reported earnings in the 1940 census, and (iii) had an occupation other than "teachers, n.e.c." (category 18). We then fit a regression model for log earnings, including a dummy variable for female, dummy variables for each category of education, a cubic in potential years of experience, and unrestricted state dummies, with New York state as the omitted state. Denote the estimated fixed effect for state *s* as $\hat{\delta}_s$. These provide estimates of the deviation in mean wages for a representative worker relative to earnings in New York. Our adjustment factor for each state is then $\exp(\hat{\delta}_s)$.

relatively highly correlated across states in 1940, suggesting that higher pay results in a higher-quality pool of teachers.[16]

Across states our simple measures of upward mobility for 16–18-year-old coresident teenagers are strongly correlated with indicators of school quality. Figure 5*A* and figure 5*B* plot upward mobility rates of white daughters and sons against the average pupil-teacher ratios and average teacher wages (measured in hundreds of dollars per year) assigned to each child. Figure 5*C* and figure 5*D* repeat this exercise for Black daughters and sons, using school quality measures for the Black schools in the 18 southern states with de jure segregation. The graphs for Black children are particularly striking, with low-quality schools and low mobility rates in states like South Carolina, Louisiana, and Mississippi and higher-quality schools and higher mobility rates in states like Delaware, Maryland, and Missouri.

In comparing the graphs for white and Black students, it is worth noting that the horizontal axes—which provide school quality metrics—are on different scales. In figure 5*A* and 5*B*, for white students, state-level student-teacher ratios never exceed 40, and teacher wages are never less than $500 per year. In contrast, in figure 5*C* and 5*D*, for Black students, there are many states with student-teacher ratios that exceed 40 and many states in which average teacher wages are less than $500.

School Quality and Upward Mobility

While suggestive, the simple plots in figure 5 ignore two key sets of factors that could confound the relationship between school quality and upward mobility. The first is family background variables, such as parental education or rural versus urban location. The second is other statewide variables, such as average income or average education that could affect both average levels of school funding and the schooling choices of low-educated families. We address the effects of these two sets of variables using a two-step approach. In the first step we estimate a linear probability model for our measure of upward mobility that includes a rich set of family background variables and unrestricted state dummies. These estimated dummies can be interpreted as “adjusted” statewide mobility rates, accounting for differences in family variables across states. In the second step, we then relate the estimated state dummies to measures of average school quality and other statewide contextual factors. An advantage of this two-step approach is that we can use relatively simple inference procedures in the second step, particularly if we want to estimate the second-step model by IVs and account for the relatively small number of states.

[16] See fig. A1*B*. For this graph we calculate the fraction of teachers in the 1940 census with a college degree in each state and plot this variable against the mean state teacher wage in 1940 from the Card-Krueger data.

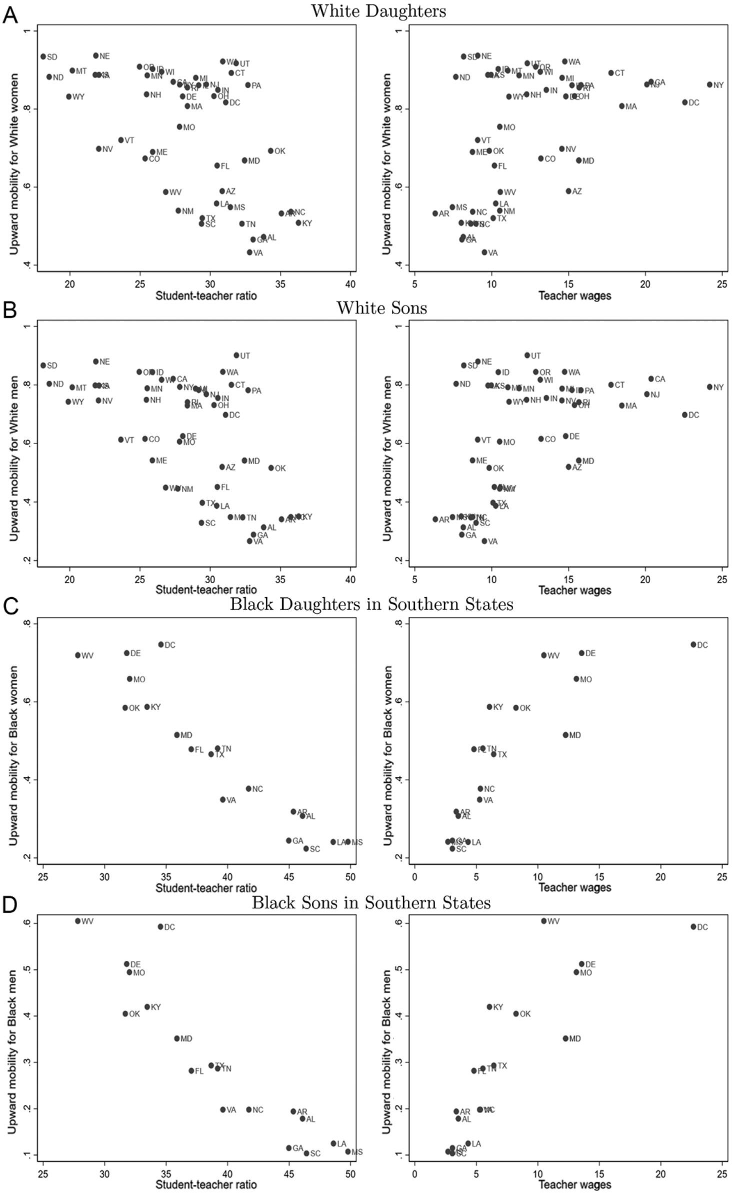

FIG. 5.—Relationship between upward mobility in education and school quality measures. Teacher wages are in hundreds of dollars per year. A color version of this figure is available online.

Specifically, we begin with the sample of coresident 16–18-year-old children of native born parent(s) with ≤6 years of schooling. We define P_i as an indicator for the event that child i has completed eight grades of education (or has completed six or seven grades and is currently enrolled). We then estimate a linear probability model,

$$P_i = A_i\beta_A + F_i\beta_F + \sum_s D_i^s\alpha_s + \epsilon_i, \tag{1}$$

where A_i is a vector of age dummies, F_i is a vector of family-level control variables, and D_i^s is a dummy indicating whether child i lives in state s. We include the following variables in F_i: (1) indicators for only mother present or only father present, (2) indicators for one parent or both parents born in a different state, (3) indicators for whether the father (in a two-parent family) or the mother (in a one-parent family) had moved within the current state or moved between states in the period from 1935 to 1940, (4) indicators for living in an urban area or living on a farm, (5) indicators for 5-year intervals summarizing the age of the oldest parent, and (6) indicators for the level of education of the best-educated parent. In estimation we normalize the state of residence coefficients (α_s) by assuming that their weighted sum is zero. We fit the model separately by gender and race, using data for white children in every state and data for Black children in the segregated southern states.

In the second step, we relate the estimated state coefficients from equation (1) to W_s, the average level of teacher wages assigned to children from state s used in the first-stage model, and PT_s, the average pupil-teacher ratio for the children in state s, along with additional controls (X_s):

$$\hat{\alpha}_s = \pi_0 + W_s\pi_W + \mathrm{PT}_s\pi_{\mathrm{PT}} + X_s\pi_X + \xi_s. \tag{2}$$

We present a simplified version of this model (specification a) that includes only the teacher wage variable, as well as a more general version (specification b) that includes the two school quality measures and three other state-level controls: the mean years of education of whites aged 25–55 in the state, the mean income of whites aged 25–55 in the state, and the mean value of homes in the state (all obtained from 1940 census tabulations). We estimate these models by weighted least squares, weighting the data for state s by the inverse sampling variance of the estimated state dummy.[17]

Coefficient estimates from the first-stage models are reported in columns 1–4 of table A2. In brief, these models suggest that upward mobility is more likely for families living in urban areas, less likely for those living on farms, more likely for families where the parents have moved from their state of birth, and less likely if the family has moved in the past 5 years. Within the set of households

[17] If there were no other control variables, this weighting would lead to second-stage estimates that are numerically identical to those from a one-step procedure in which we included the school quality measures directly in the first-stage model (Hanushek 1974).

we study—those with parents who have no more than 6 years of education—upward mobility is substantially more likely in families with higher parental education. Interestingly, the coefficient estimates are not too different for males versus females of either race or for Blacks versus whites of either gender.

The key coefficient estimates from the second-stage models for each race and gender group are presented in columns 1–4 of panel A in table 2. We note that these models are estimated on 49 observations for whites (the 48 mainland states plus the District of Columbia) and on 18 observations for Blacks (the 18 segregated southern states, including the District of Columbia). Inspection of the coefficient estimates from alternative specifications leads to two main conclusions. First, the coefficient on teacher wages in specification a with no other controls is positive, highly statistically significant, and very similar in magnitude across gender and race groups. The similarity of the effects for sons and daughters is reassuring, given concerns about potential selection bias in conditioning on coresidency for females. The similarity between white and Black students is also interesting because the level of teacher wages in the segregated Black schools of the South was less likely to reflect the preferences of Black families and was instead largely driven by the decisions of white politicians.[18]

The estimated effects are large: W is scaled in \$100s, so a coefficient of 0.02 implies that a \$500 salary difference—comparable to the gap between a typical southern state and a typical state outside the South—is associated with a 10 percentage point difference in the probability of upward intergenerational mobility for children from the lowest parental education quartile.

Second, the estimated effect of teacher wages is only slightly attenuated when we fit specification b, which includes the average pupil-teacher ratio (PT) and controls for average income, education, and home values in the state. In fact, most of the attenuation is attributable to the addition of PT, which is mildly correlated with teacher wages ($\rho \approx -0.2$) and exerts its own independent effect on mobility. The magnitude of the estimated pupil-teacher effects is around -0.01, which implies that a five-student reduction in the pupil-teacher ratio raises the intergenerational mobility rate by about 5 percentage points.

While our main models pool all families in which the best-educated parent has no more than 6 years of schooling, one could ask whether the effects of school quality are larger or smaller for children whose parents have different levels of education. To evaluate this, we reestimated our first- and second-stage models separately for different parental education groups up to a maximum of 8 years, pooling the relatively small set of families with parental education of 0–2 years. Figure 6 shows the estimated effects of teacher wages on upward mobility rates of the different parental education groups

[18] Black communities could exert some influence on local school resources through the Rosenwald program (see Aaronson and Mazumder 2011). We investigate the impact of Rosenwald schools directly in our border county analysis below.

Table 2
Cross-State Models of Educational Attainment in Families with Parental Education of Six Grades or Less

	Linear Probability Model for Eighth-Grade Completion				Tobit Model for Highest Completed Grade			
	Whites		Blacks		Whites		Blacks	
	Females (1)	Males (2)	Females (3)	Males (4)	Females (5)	Males (6)	Females (7)	Males (8)
	A. OLS models							
Number of state observations	49	49	18	18	49	49	18	18
	a. State Average Teacher Wage Only							
Coefficient of teacher wage	.020	.029	.020	.023	.173	.224	.146	.179
	(.004)	(.004)	(.004)	(.005)	(.024)	(.027)	(.031)	(.038)
	b. State Average Teacher Wage and Average Pupil-Teacher Ratio, with Added Controls							
Coefficient of teacher wage	.017	.027	.015	.021	.137	.206	.113	.190
	(.006)	(.007)	(.006)	(.007)	(.034)	(.040)	(.042)	(.050)
Coefficient of pupil-teacher ratio	−.010	−.014	−.009	−.007	−.036	−.078	−.060	−.043
	(.002)	(.003)	(.002)	(.002)	(.024)	(.023)	(.021)	(.018)
	II. IV Models, Using State Minimum Teacher Wage as Instrument for Teacher Wage							
Number of state observations	23	23	10	10	23	23	10	10
OLS coefficient of teacher wage in IV sample	.020	.027	.025	.028	.180	.221	.182	.242
	(.004)	(.004)	(.009)	(.010)	(.031)	(.031)	(.060)	(.062)
First-stage coefficient	1.372	1.385	1.290	1.270	1.395	1.386	1.317	1.304
	(.139)	(.134)	(.265)	(.268)	(.145)	(.142)	(.265)	(.265)
Reduced-form coefficient	.028	.037	.038	.038	.270	.316	.281	.337
	(.007)	(.008)	(.008)	(.011)	(.058)	(.060)	(.060)	(.083)
IV coefficient of teacher wage	.020	.027	.030	.030	.193	.228	.214	.258
	(.005)	(.005)	(.006)	(.007)	(.032)	(.032)	(.044)	(.043)

NOTE.—Standard errors are in parentheses. In cols. 1–4, the dependent variable is the estimated state dummy from a linear probability model fit to children aged 16–18 who are observed in the 1940 census living with at least one parent, with parental education of six grades or less. In cols. 5–8, the dependent variable is the estimated state dummy from a tobit model for years of completed education fit to youth aged 14–18 living with at least one parent, with parental education of six grades or less. See the main text for other controls included in the first-stage models. Second-step models are fit by weighted least squares or weighed IV, using as a weight the inverse sampling variance of the state dummy that is the dependent variable. Models with controls in panel A include mean education of whites aged 25–55 in the state, mean income of whites 25–55 in the state, and mean value of homes in the state, all obtained from 1940 census tabulations, as well as the number of years of schooling mandated by compulsory schooling and child labor laws, calculated as in Stephens and Yang (2014).

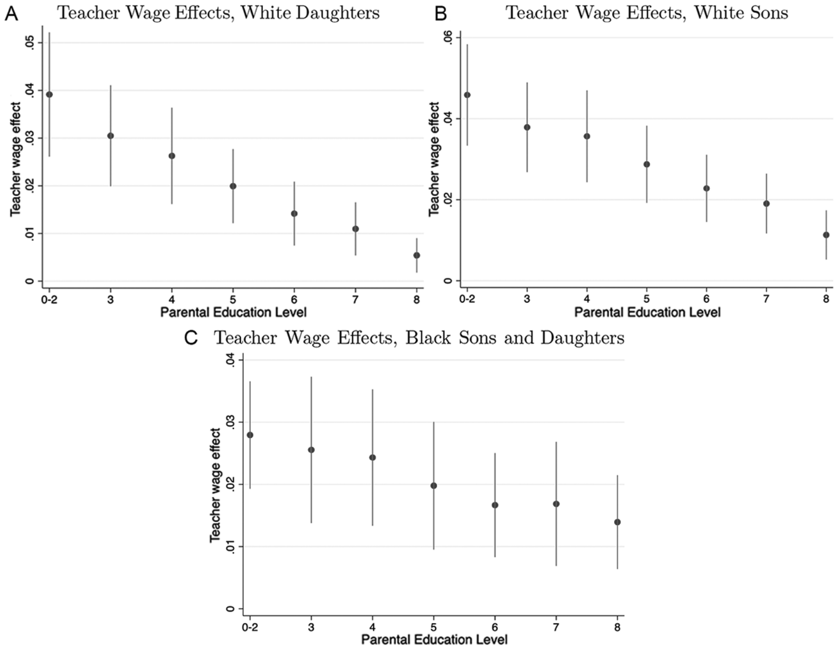

FIG. 6.—Relationship between teacher wage and eighth-grade attainment at the state level—parameter estimates for bivariate models. Each point represents the parameter estimate from a bivariate regression (as in table 2, specification a) that restricts the sample by parental education. Lines indicate 95% confidence intervals. A color version of this figure is available online.

using our second-stage specification a, which has no other controls.[19] Among white families we see a pattern of declining teacher wage effects across education groups, with a coefficient of around 0.04 for children with very poorly educated parents and closer to 0.01 for children whose parents have 8 years of schooling. Among Black families the gradient across parental education groups is less pronounced, perhaps reflecting the greater degree of mismeasurement in education for Black adults noted by Margo (1986) as well as the possibility that each year of schooling for Black parents was less valuable in producing human capital, given the low quality of the schools they attended.[20]

[19] We combine genders for Black families to build sample sizes, which seems reasonable given the results in table 2.

[20] Card and Krueger (1992a) find that the labor market returns to southern-educated Blacks born in the 1920s and working in outside the South in 1980 were quite low and interpret this as evidence of the low quality of their schools.

School Quality and Completed Education

To complement our analysis of the effects of school quality on the probability of attaining at least eighth grade, we present a parallel analysis using an alternative measure of educational achievement—actual years of completed schooling for children 14–18, treating those who are still enrolled as right censored. We replace the linear probability model of equation (1) with a tobit model that includes all of the same control variables. Then, in the second step, we regress the estimated state effects from the tobit on our school quality measures and other controls.

In general the estimates from these models, reported in columns 5–8 of table 2, tell the same story as our analysis of eighth-grade completion. In particular, we see that teacher wages have significant positive effects on educational attainment of 14–18-year-olds, with similar magnitudes for both genders and both race groups. Moreover, as in the models for eighth-grade attainment, the effect of teacher wages is only slightly attenuated when we add controls in the second-stage model for the pupil-teacher ratio and mean education, income, and house values in the state.

To get a sense of the magnitudes implied by the models for completed education, note that the estimated state effects α_s from the first-step tobit model are scaled in years of education. Thus, a coefficient of 0.2 on annual teacher salary—as is approximately the case for sons—implies that a $100 increase in average teacher salaries is associated with a 0.2-year increase in completed education. This is a relatively large effect. Interpreted causally, it implies that raising teacher wages by $500 per year would lead to one extra year of education among 14–18-year-olds from less educated families. The estimated effects of the average pupil-teacher ratio are more variable across the four race/gender groups, and the point estimate for white females is insignificant at conventional levels. It is worth noting, however, that an effect of −0.04 implies that a five-student reduction in the pupil-teacher ratio would lead to a 0.2-year increase in completed education among 14–18-year-olds.

Controlling for the Composition of the Teaching Workforce

A potential issue with the simple models in panel A of table 2 is that the variation in average teacher wages across states is driven by differences in the composition of the teaching workforce rather than by differences in the pay for the teachers at each grade. Specifically, high school teachers are generally paid a significant premium relative to elementary school teachers.[21] Thus, in states with a higher fraction of high school teachers the average pay of teachers would tend to be higher. Any associated impact on student choices could

[21] For example, detailed tabulations in National Education Association (1923, tables 57–58) suggest that in the early 1920s teachers at senior high schools were paid about 30% more than elementary school teachers, controlling for city size.

reflect a high school accessibility effect rather than an effect of having higher-paid teachers at each grade.

One way to eliminate workforce composition effects is to make use of minimum teacher salary laws. As noted earlier, by the late 1930s a total of 23 states had established statewide floors for teacher salaries. With this in mind, panel B of table 2 presents simple IV estimates of specification a, restricting the sample to states that had statutory minimum teacher wages and using the minimum salary as an instrument for the state average teacher wage. We note that the sample sizes here (especially for Black children) are quite modest, so the results should be interpreted cautiously.

The first row in the panel shows that the OLS coefficients for this smaller sample are quite similar to those found with the full sample. In the second row we show that the first-stage models reveal a large positive effect of a state's minimum salary on average teacher salaries, with a typical point estimate of 1.3 to 1.4. The IV estimates in the fourth row are quite similar to the OLS estimates and are about as precisely estimated. Overall, the estimates suggest that the observed correlation between average teacher salaries and educational outcomes of children from lower-educated families was attributable to variation in state's choices about average salary per teaching position rather than to policies or practices affecting the composition of the teaching workforce.

B. County-Level Analysis: A Cross-Border Design

While the state-level models in table 2 are suggestive, a concern is that there may be unobserved factors (like tastes for education) that are orthogonal to the family characteristics included in F_i and to the state-level controls included in X_s but that affect both school resources and the schooling choices of less educated families in a state. To address this concern, we turn to an analysis of schooling outcomes for children from matched pairs of counties lying on opposite sides of the borders of the southern states. There are three key advantages of such an approach. First, by focusing on within-pair comparisons we eliminate many local factors—like labor market opportunities for better-educated workers—that vary across broader areas. Second, the number of border counties is relatively large, enabling us to expand the list of local control variables included in our models. Third, in most southern states the counties served as school districts. Thus, it is natural to take the county as the unit of analysis for studying the effect of local schools. By narrowing our focus to local schools, we address concerns that have been raised about the use of more aggregated school quality measures (e.g., Hanushek, Rivkin, and Taylor 1996).

Despite the similarity of adjacent cross-border counties, average teacher wages were often very different, owing in part to statewide minimum teacher salary laws. Appendix B reports the minimum teacher wage in each of the

10 southern states that had a minimum salary law as well as a brief legislative history for each state. The highest minimums were in Delaware ($1,000 per year for white and Black teachers), Maryland ($1,000 per year for whites, $585 per year for Blacks), and Oklahoma and West Virginia (both $585 per year for both race groups). The lowest minimum salaries were in Alabama ($350 for whites and $262.50 for Blacks), Georgia ($280 for whites and $175 for Blacks), and Mississippi ($80 for Blacks).[22] These salary laws exerted a significant impact on salaries. In figures B1 and B2 we plot mean teacher wages in each country (estimated from the 1940 census) against the minimum teacher wage in the corresponding state.[23] The figures show a strong positive relationship, with correlations of 0.69 for Black teachers and 0.60 for white teachers.

Although a majority of southern states had minimum salary laws, eight others did not, posing a challenge for the strategy of using the difference in state minimum wages within a border pair as an IV for the difference in average salaries. One possibility we explore in our robustness analysis below is to use a minimum wage of $0 for states with no law. After conducting a visual analysis of the distributions of teacher wages in each southern state, however, we elected to use the 10th percentile of annual teacher earnings in the state as our preferred proxy minimum wage. Figure B3 presents the distributions of annual earnings for Black teachers in each of the 18 southern states, indicating the state minimum salary for states with a law and the 10th percentile of annual salaries for other states. Despite measurement error in the reporting of income in the census,[24] there are clear "shoulders" in the distribution of earnings at the statewide minimum in states like Alabama, Georgia, and Mississippi. Interestingly, there are similar shoulders at around the 10th percentile in states like Arkansas, Florida, Louisiana, and South Carolina that had no statewide minimum.

[22] As explained in more detail in app. B, some states had multiple minimums. For example, Delaware had a lower minimum salary ($800) for teachers holding an older certification. Mississippi's constitution set the minimum wage to $80 for both Black and white teachers, but as a practical matter this wage was binding only for Black teachers. Mississippi's equalization program offered districts the incentive to adopt higher minimum salaries for white teachers, but we do not use these standards either because many districts in Mississippi opted out of the equalization program.

[23] To construct teacher salaries in each county, we calculate mean income reported by individuals age 16 or older who have at least six grades of completed education and report their occupation as "teachers, n.e.c." and their industry as "education services."

[24] Miller and Paley (1958) conducted a matched comparison of incomes reported to the Internal Revenue Service and in the 1950 census and found notable discrepancies. Typically only 40%–45% of people who reported income to the census in a given $500 cell had taxable income in the same cell, and only 65% had taxable income in the same range or plus or minus one cell.

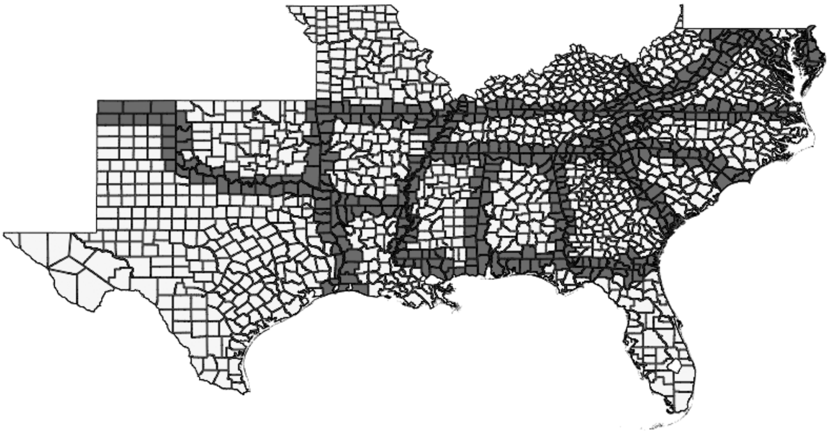

FIG. 7.—Border counties used in the county-level analysis. A color version of this figure is available online.

Cross-Border Pairs

We use the county adjacency file published by the US Census Bureau (2010) to identify contiguous counties along the borders between southern states. Figure 7 shows the borders in question, highlighting the counties that can be paired. Since a given county on one side of the border often can be paired with more than one county on the other side, we selected the "best match" from among all possible pairs by minimizing the difference in mean education of white adults residents (calculated using in the 1940 census) in the two counties. We also drop all pairs with more than a 1-year absolute difference in mean white education levels, leading to a sample of 261 border county pairs. Although all border counties have relatively large numbers of white families, some have relative few Black families. We exclude counties with fewer than 10 Black residents between 16 and 18 years of age and those with no resident Black teachers, narrowing the sample for Black families to 190 county pairs along 27 distinct border segments.

Table 3 shows differences in average adult incomes, teacher wages, and upward mobility rates for pooled samples of counties on either side of the border segments between Deep South and border states (e.g., Mississippi vs. Tennessee).[25] In each case the first state listed has the lower statutory minimum wage for Black teachers (or 10th percentile of Black teacher wages in the entire state in cases where there is no statewide minimum).[26] The

[25] To streamline the table, we present only border segments for which there are at least 250 Black 16–18-year-olds on each side of the border, although we retain the less populated segments in our analysis.

[26] In the case of comparisons involving Mississippi, note that although it had a (very low) minimum wage for Black teachers, it had no minimum for white teachers.

Table 3
Comparisons of Cross-Border Counties, Borders between a Deep South State and a Peripheral South State

	Black			White			
	Average Income	Teacher Wage	Upward Mobility	Average Income	Teacher Wage	Upward Mobility	Counties
MS-TN:							
Mississippi	230	268	.14	480	608	.35	5
Tennessee	301	656	.32	769	819	.38	5
Gap	−72	−388	−.17	−290	−212	−.03	
SC-NC:							
South Carolina	384	422	.19	887	853	.35	9
North Carolina	405	624	.28	872	932	.40	12
Gap	−21	−202	−.09	15	−79	−.05	
AL-TN:							
Alabama	257	417	.17	635	756	.32	2
Tennessee	351	613	.38	688	679	.39	3
Gap	−94	−197	−.21	−53	77	−.07	
MS-AR:							
Mississippi	264	193	.09	1,007	1,054	.23	3
Arkansas	308	343	.18	885	780	.34	5
Gap	−43	−149	−.09	122	274	−.11	
GA-FL:							
Georgia	319	350	.17	796	751	.41	8
Florida	316	483	.22	785	916	.43	8
Gap	3	−133	−.05	11	−165	−.02	
AR-LA:							
Arkansas	332	388	.26	930	693	.49	4
Louisiana	355	418	.19	1,004	978	.56	5
Gap	−23	−30	.06	−74	−285	−.07	
LA-TX:							
Louisiana	404	467	.21	971	981	.50	6
Texas	348	478	.29	812	822	.50	6
Gap	56	−11	−.08	159	159	.00	

NOTE.—The state with the lower minimum Black teacher salary (or 10th percentile) is listed first. Summary statistics are displayed for border county pairs for which the difference in the average educational attainment of whites aged 25–55 is less than 1 year. We list cases with at least 250 Black 16–18-year-olds on each side of the border. Upward mobility is the fraction of 16–18-year-olds attaining eighth grade in families with parental education of six grades or less.

first three columns provide statistics for the Black populations in each border region, while the next three columns show parallel statistics for the white populations.

Four main features stand out from the table. First, average adult incomes are generally similar for Blacks on either side of the border, although there are a few larger gaps for whites. In light of this, in our models we include controls for average income and other local characteristics in each county. Second, for each border segment, Black teacher wages are always lower in the state with the lower minimum teacher salary for Black teachers. In some instances these differences are large; for example, along the Mississippi-Tennessee border, Black teachers in Mississippi earned less than half as much as their Tennessean counterparts. Third, upward mobility rates for Black children of less educated parents are typically lower in border counties where teachers are poorly paid. Here again, the differences are sometimes large; for example, along the Mississippi-Tennessee border upward mobility is much lower in Mississippi (0.14) than in Tennessee (0.32). Finally, the cross-border differences in teacher salaries and mobility rates are generally smaller for whites than for Blacks.

Model Specification and Main Results

As in our state-level studies, we focus on two outcome measures for coresident children whose parents had ≤6 years of education: the eight-grade attainment rate for 16–18-year-olds and completed years of schooling among 14–18-year-olds. Given the results from our state-level models, which are very similar for males and females, we pool sons and daughters, which is helpful because of the small samples of age-appropriate children in some counties.

As in our cross-state analyses, we use a two-step methodology. For our analysis of eighth-grade attainment the first step is a linear probability model for the event of having at least eighth-grade schooling, containing all of the same control variables used in our cross-state models plus a gender dummy and unrestricted county dummies.[27] For our analysis of completed education among 14–18-year-olds the first-stage model is a tobit, treating children who are still enrolled as right censored. In either case, we then proceed with a second-step model in which the dependent variable is the within-pair difference in estimated county dummies from the first-stage model. In particular, in the second stage we form the difference $\Delta\hat{\alpha}_p$ in the estimated dummies for the two counties in matched pair p and fitted the following model:

$$\Delta\hat{\alpha}_p = \pi_0 + \Delta W_p \pi_W + \Delta X_p \pi_X + \epsilon_p, \tag{3}$$

where ΔW_p is the within-pair difference in average teacher wages for border pair p and ΔX_p is a vector of within-pair differences in county-level controls. We fitted the model by weighted least squares (or weighted two-stage least squares) using as weights the inverse of the estimated sampling variance of the dependent variable.

In the vector of controls in our models for Black children we include the differences in the fractions of Black families living on a farm or in urban areas;

[27] Since we are evaluating counties in bordering states, we might be concerned that there is some cross-border migration that contaminates our evaluation. Thus, we also tried including a dummy variable indicating that a parent was born in the neighboring state. This addition led to virtually no difference for the estimated coefficients reported below.

the differences in mean parental earnings and parental education of Black families; the difference in mean education of white adults; the difference in the county-average number of Rosenwald teachers per Black student in the 1931 birth cohort (from Aaronson et al. 2021); the difference in the number of years of schooling mandated by compulsory schooling and child labor laws, calculated as in Stephens and Yang (2014); and the difference in the fraction of Blacks in the country population. In the parallel set of models for white children we substitute the differences in mean parental earnings and mean parental education for white families and drop the Rosenwald teacher variable.

In our baseline specifications we use average teacher salaries calculated from the 1940 census as a measure of local school quality, although we present some models in our robustness analysis that include a control for the within-county difference in pupil-teacher ratios and other models that use administrative data on mean teacher salaries by county (i.e., data derived from various state reports on average teacher wages in a given county). There are three reasons for this. First, we can obtain average teacher salaries from the census for every border county, whereas administrative data on teacher salaries are unavailable for several states (including Alabama, Louisiana, and Mississippi). Second, using teacher wages as the measure of quality, we can readily compare OLS models to IV models that use the difference in state minimum teacher salaries as an IV for the within-pair difference in teachers wages. There is no comparable way to address potential endogenous differences in pupil-teacher ratios within county pairs, arising, for example, from enrollment responses to the quality of teachers in different counties. Third, as noted in the discussion of table 2, the measured effect of teacher salaries across states appears to be relatively robust to the inclusion of the pupil-teacher ratio. We will see below that this is also the case in our cross-border models.

Table 4 presents estimates of our main second-stage models for the schooling outcomes of Black children. Panel A contains the estimates for our model of upward mobility in education, while panel B presents the estimates for schooling attainment. In each case we present a model estimated using schooling outcomes for all children in each set of counties, along with a second model based on within-pair differences in outcomes for children in rural areas only (i.e., excluding those in cities and towns). Rural Black families in the South were among the most disadvantaged groups in the country in the late 1930s, so it is interesting to see whether higher teacher pay at their local schools had any effect on their children. Standard errors for all models in this table are clustered at the border segment level to accommodate potential correlations across border pairs from the same segment. Given the small numbers of border clusters (reported in the last column of the table) we also report wild bootstrap confidence intervals (Roodman et al. 2019).

Row 1 of panel A gives our baseline specification for the effect of teacher wages on upward mobility. The OLS estimate of 0.027 is close to the estimates of 0.020 and 0.023 obtained for Black daughters and sons in our cross

Table 4
Effect of Teacher Wages on Educational Attainment among Black Children, Border County Analysis

	OLS	First Stage	Reduced Form	2SLS	F-statistic	n	n_c
	A. Effects on Eighth-Grade Attainment, Ages 16–18, Parental Education Six Grades or Less						
1. Baseline	.027	.807	.028	.034	100.49	180	27
	(.007)	(.080)	(.007)	(.008)			
	[.012, .046]	[.635, 1.045]	[.013, .046]	[.016, .056]			
2. Rural areas	.028	.730	.029	.040	62.86	163	24
	(.007)	(.092)	(.006)	(.009)			
	[.012, .044]	[.544, .970]	[.015, .045]	[.019, .061]			
	B. Effects on Years of Schooling, Ages 14–18, Parental Education Six Grades or Less						
1. Baseline	.210	.810	.240	.297	88.12	190	27
	(.032)	(.086)	(.038)	(.043)			
	[.141, .284]	[.626, 1.043]	[.163, .350]	[.208, .406]			
2. Rural areas	.217	.729	.238	.326	59.59	176	26
	(.037)	(.094)	(.033)	(.050)			
	[.127, .289]	[.530, .987]	[.172, .320]	[.207, .444]			

NOTE.—The sample is restricted to county border pairs for which the difference in the education of whites is less than 1 year and county sample sizes are at least 10. Our instrument is the mandated minimum salary (or the 10th percentile of earnings for states with no minimum salary). Teacher salaries are measured from the census. Controls include differences between counties in Rosenwald Fund exposure, fraction urban, fraction living on farm, average Black parental income and education, and average education of whites as well as differences in the number of years of schooling mandated by compulsory schooling and child labor laws, calculated as in Stephens and Yang (2014). n gives the number of border county pairs; n_c gives the number of borders. Standard errors (in parentheses) are clustered at the border level; 95% wild bootstrap confidence sets for the null hypothesis are reported in brackets.

state models (shown in cols. 3 and 4 of panel A[a] in table 2), suggesting that the cross-state models are not in fact significantly affected by aggregation biases (Hanushek, Rivkin, and Taylor 1996) or by cross-state differences in family preferences or local opportunities that are eliminated (or at lease substantially reduced) within adjacent county pairs. To evaluate the potential role of any remaining omitted factors, we reestimated the OLS model with alternative subsets of the county-level control variables. None of the individual controls has a noticeable effect on the estimated teacher wage coefficient, nor does the sequential addition of controls (in any order) lead to large changes in the effect. Altonji, Elder, and Taber (2005) and Oster (2017) suggest that this invariance can be taken as evidence that other unobserved determinants of educational attainment within county pairs are unlikely to lead to substantially different estimates.

Nevertheless, to address any remaining concerns about potential endogeneity of teacher wages, in columns 2–4 we show the first-stage, reduced-form, and IV estimates resulting from specifications in which we use the within-pair difference in state minimum teacher salaries as an instrument for the difference in mean teacher wages. In these specifications we use the statewide

10th percentile of teacher wages as a proxy for the minimum wage in states with no legislative minimum: we show the stability of the results under alternative strategies in the robustness analysis below. The first-stage estimate under this assumption is 0.807 and is precisely estimated, and the reduced-form estimate of 0.028 is also relatively precisely estimated. The resulting IV estimate, 0.034, is somewhat larger than the OLS estimate, although the difference is not significant.[28] There is certainly no evidence that the correlation between teacher wages and eighth-grade attainment of Black children from less educated families is upward biased by unobserved local factors.

Row 2 shows a parallel set of OLS and IV estimates for the upward mobility of Black children from rural areas. The OLS estimate of the effect of teacher wages is virtually the same as the one for all children, the IV estimate is slightly higher for this group, and both are very close to the IV estimates from our cross-state models for Black daughters and sons (0.032 and 0.030, respectively).

Taking the estimates from the state and border-pair analysis as a whole, we conclude that higher teacher wages exerted a relatively strong effect on upward mobility rates of Black children, with each \$100 per year in additional teacher wages leading to a 2–3 percentage point gain in the probability that a Black child whose parents had ≤6 years of schooling would reach at least eighth grade. To put this in perspective, the Tennessee-Mississippi gap in average teacher wages for Black families living along the state border was about \$390 per year (see table 3). Our models imply that this gap can account for 45%–65% of the 18 percentage point gap in eighth-grade attainment rates between students on the Tennessee side and those on the Mississippi side.

Panel B of table 4 presents estimates for models in which the dependent variable is years of completed schooling among 14–18-year-olds. Here, OLS estimates of the effect of teacher salaries are very similar for all children (0.190) and those in rural areas (0.193). Again, both are similar to OLS estimates from cross-state data, which yield estimated coefficients of 0.147 for Black daughters and 0.181 for Black sons (see table 2). As is true for our OLS models of eighth-grade attainment, we find that adding or removing controls from our second-stage model leads to little or no change in the magnitude of the estimated effect of teacher wages on completed education.

The IV estimates in panel B of table 4 are somewhat larger than the corresponding OLS estimates, suggesting that if anything the OLS estimates may be conservative.[29] Given the modest numbers of independent border

[28] We note that even using the wild bootstrap confidence interval the IV estimate is statistically different from zero at conventional levels. Dividing the confidence interval by four yields a value of 0.01, which is 20% larger than the nominal (clustered) standard error.

[29] Note that the weights assigned to different county pairs are slightly different in panel A and panel B. Hence, the first-stage model estimates are slightly different in the two panels, even though the dependent and explanatory variables are identical.

segments, we believe a safe conclusion is that higher teacher salaries helped to raise the amount of education obtained by Black children of less educated parents, with each $100 increase in teacher salaries leading to something like 0.2 to 0.3 extra years of completed schooling.

Although the primary focus of our cross-border design is Black families, we present a parallel analysis for white families in table 5. In the models for eighth-grade completion (panel A), neither the OLS or IV coefficients on teacher wage are statistically significant, but the IV point estimates are similar in magnitude to the corresponding OLS and IV estimates at the state-level analysis (and to the estimates in table 4 for Black children). In models for completed education of 14–18-year-olds (panel B), the OLS and IV coefficients are all statistically significant, and the IV estimates in particular are quite close in magnitude to those we obtained for children of both races in our cross-state models (with an estimated effect of around 0.20). As with our estimates for Black families, it appears that any upward biases in the OLS specification caused by endogeneity of local teacher wages may be small, and indeed OLS estimates may be downward biased due to measurement error. Nevertheless, on the basis of IV estimates, it appears that the impacts of teacher wages in the southern border counties may have been moderately lower

Table 5
Effect of Teacher Salaries on Years of Education among White Children, Border County Analysis

	OLS	First Stage	Reduced Form	2SLS	F-statistic	n	n_c
	A. Effects on Eighth-Grade Attainment, Ages 16–18, Parental Education Six Grades or Less						
1. Baseline	.010	.502	.007	.014	60.22	260	32
	(.006)	(.065)	(.006)	(.012)			
	[−.003, .025]	[.362, .653]	[−.008, .023]	[−.013, .045]			
2. Rural areas	.009	.527	.008	.015	55.32	259	32
	(.006)	(.071)	(.006)	(.012)			
	[−.005, .022]	[.373, .696]	[−.007, .025]	[−.012, .046]			
	B. Effects on Years of Schooling, Ages 14–18, Parental Education Six Grades or Less						
1. Baseline	.114	.504	.115	.228	55.94	261	32
	(.038)	(.067)	(.045)	(.088)			
	[.028, .195]	[.355, .655]	[.007, .246]	[.026, .451]			
2. Rural areas	.105	.532	.119	.223	51.83	260	32
	(.034)	(.074)	(.047)	(.087)			
	[.028, .172]	[.363, .711]	[.016, .26]	[.023, .455]			

Note.—The sample is restricted to county border pairs for which the difference in the education of whites is less than 1 year and county sample sizes are larger than 10. Our instrument is the mandated minimum salary (or the 10th percentile of earnings for states with no minimum salary). Teacher salaries are measured from the census. Controls include differences between counties in fraction urban, fraction living on farm, average parental income of whites, and average education of whites. n gives the number of border county pairs; n_c gives the number of borders. Standard errors (in parentheses) are clustered at the border level; 95% wild bootstrap confidence sets for the null hypothesis are reported in brackets.

for white children of poorly educated parents than for comparable Black children.

In sum, the cross-border design leads to the same broad conclusion as the state-level analysis: higher teacher pay has a meaningful impact on the schooling attainment of children in families with relatively low parental education. We suspect that improved student outcomes are driven by the fact that higher pay improved the capacity of school districts to attract and retain effective teachers. Evidence favoring this supposition is provided in figure B4, which shows that in our border counties there is a clear and strong positive relationship between county-level teacher pay and teacher education, as measured by the share of teachers with a college degree. This observation is true of both Black and white teachers. A similar point can be made at the state level; as we have seen, in figure A1 college completion rates among white teachers are higher in states that have higher average teacher pay.

Robustness Checks

In this section we summarize the results of a series of checks that probe the robustness of our cross-border analysis of Black families. Table 6 shows a series of alternative specifications for the second-step models reported in table 4. The first line of each panel presents our baseline specification, showing the OLS and IV estimates of the effect of higher teacher salaries on the eighth-grade attainment rate of 16–18-year-olds (panel A) and on the completed years of schooling of 14–18-year-olds (panel B). Subsequent rows show the estimates under various alternatives.

Row 2 in each panel presents results when we assume that the minimum teacher wage in states with no legislation is $0 rather than setting the minimum to the 10th percentile of Black teacher salaries in the state. Not surprisingly, this alternative leads to a smaller first-stage coefficient (0.33 vs. 0.81 in our baseline specification), but the estimated effect of the within-county-pair difference in minimum salaries on the within-pair difference in average salaries is still highly significant. And reassuringly, the IV estimates are very similar to the ones from our baseline model.

Row 3 takes an alternative approach to the problem of states with no minimum teacher salary and restricts the sample to countries on the borders between states that both have a minimum salary law. This reduces our sample of county pairs by a factor of two-thirds and limits the number of border segments (which determines the number of clusters for our standard errors) to only 10. To help conserve degrees of freedom in the second-stage models, we drop four of the control variables (differences in the fraction of Black residents, compulsory schooling ages, and fractions living on farms and in urban areas) that are all insignificant in our main specification. In the smaller sample there is some attenuation in the OLS coefficients for both dependent variables, but the estimated first-stage coefficient of minimum salaries is still highly significant. The implied IV estimate for our model of eighth-grade attainment is somewhat larger than the estimate from our baseline specification, while the IV estimate for our model of years of schooling is smaller in magnitude than our baseline estimate. In both cases, however, the differences are relatively modest.

Rows 4 and 5 present specifications that include two other potential indicators of local school resources: the presence of a high school for Black students in the county (which is available for all counties) and the pupil-teacher ratio among Black schools (which is available for most counties).[30] Adding controls for these variables has little impact on our estimates of the effect of teacher salaries, although as would be expected on the basis of our state-level models we find that higher pupil-teacher ratios are associated with worse schooling outcomes for children of less educated parents.

Finally, in row 6 we present an alternative specification where we use average teacher wages collected from state education reports rather than from the 1940 census. Since county-average teacher salary data are not available for several states, the sample of county pairs is reduced to only 82, and the number of border segments falls to 10. Again, to conserve degrees of freedom we drop four of the control variables from our main specification that are not significantly related to student outcomes. The OLS estimates using this alternative wage measure and the much smaller sample are both very close to the OLS estimates from our baseline specification, and the first-stage coefficients are also similar to those from our baseline specification. The reduced-form and IV estimates of the effect of teacher salaries are also relatively close to the corresponding estimates from our baseline specification, although the estimates from the model for eighth-grade attainment are not statistically different from zero at conventional levels.

Overall, we believe that the robustness checks presented in table 6 are strongly supportive of the key conclusions from our baseline specifications. The OLS and IV estimates for both education outcomes are relatively stable across specifications, suggesting that each 100 dollars in teacher salaries raises eighth-grade attainment rates by approximately 2 percentage points and mean completed education of 14–18-year-olds by around 0.2 years.

As a final check on our conclusions, we investigate the potential impact of selective out-migration of Black families from counties in the South with relatively low school quality. Our sample period (1940) coincides with the late

[30] We use a variety of archival sources to measure the presence of a high school, including lists of high schools and measures of high school enrollment in the state reports of education. Of the 190 county pairs in our border analysis, 63% have a Black high school in both counties, 27% have a Black high school in only one of the two counties, and 10% have no high school in either county. We construct the pupil-teacher ratio by dividing the number of enrolled Black students in a county in the 1940 census by the number of Black teachers. We winsorize the ratio at 10 and 70 and drop counties in which the estimated pupil-teacher ratio is more than 300.

Table 6
Effect of Teacher Wages on Educational Attainment among Black Children, Border County Analysis: Robustness Checks

	OLS	First Stage	Reduced Form	2SLS	F-statistic	n	n_c
	A. Effects on Eighth-Grade Attainment, Ages 16–18, Parental Education Six Grades or Less						
1. Baseline	.027	.807	.028	.034	100.49	180	27
	(.007)	(.080)	(.007)	(.008)			
	[.012, .046]	[.643, 1.045]	[.012, .048]	[.016, .058]			
2. Using zero as minimum wage in states with no law	.027	.337	.013	.040	53.60	180	27
	(.007)	(.046)	(.004)	(.013)			
	[.013, .048]	[.240, .429]	[.003, .025]	[.011, .071]			
3. Using only county pairs with state minimum wage on each side of border	.024	.822	.037	.044	11.91	58	10
	(.008)	(.238)	(.007)	(.015)			
	[−.028, .053]	[.247, 1.511]	[.022, .068]	[.009, .111]			
4. Adding control for presence of high school in county	.026	.798	.026	.032	113.6	180	27
	(.007)	(.075)	(.007)	(.008)			
	[.011, .045]	[.636, 1.012]	[.011, .045]	[.014, .055]			
5. Adding control for pupil-teacher ratio in county	.025	.807	.025	.031	86.01	175	27
	(.008)	(.087)	(.007)	(.008)			
	[.008, .045]	[.634, 1.061]	[.010, .046]	[.013, .055]			
6. Using administrative data on teacher wages	.023	.829	.024	.029	37.35	82	10
	(.004)	(.136)	(.006)	(.009)			
	[.013, .041]	[.285, 1.280]	[.012, .047]	[.006, .067]			

	B. Effects on Years of Schooling, Ages 14–18, Parental Education Six Grades or Less						
1. Baseline	.210	.810	.240	.297	88.12	190	27
	(.032)	(.086)	(.038)	(.043)			
	[.141, .292]	[.631, 1.068]	[.161, .341]	[.207, .400]			
2. Using zero as minimum wage in states with no law	.210	.351	.094	.269	59.15	190	27
	(.032)	(.046)	(.024)	(.063)			
	[.140, .286]	[.241, .442]	[.031, .144]	[.101, .416]			
3. Using only county pairs with state minimum wage on each side of border	.182	.796	.226	.284	12.47	60	10
	(.060)	(.225)	(.052)	(.052)			
	[−.170, .325]	[.243, 1.477]	[.145, .394]	[.184, .414]			
4. Adding control for presence of high school in county	.191	.801	.218	.273	94.91	190	27
	(.032)	(.082)	(.033)	(.038)			
	[.120, .268]	[.614, 1.030]	[.156, .305]	[.198, .358]			
5. Adding control for pupil-teachers ratio in county	.197	.804	.227	.282	74.83	185	27
	(.036)	(.093)	(.041)	(.040)			
	[.117, .273]	[.597, 1.097]	[.154, .347]	[.206, .379]			
6. Using administrative data on teacher wages	.145	.838	.215	.257	36.81	82	10
	(.018)	(.138)	(.035)	(.057)			
	[.103, .188]	[.307, 2.037]	[.155, .402]	[.104, .433]			

NOTE.—The sample is restricted as in table 4 (which gives the baseline results). The variables and instrument are also as in table 4, except as noted. Specifications with $n_c < 27$ drop detailed county-level controls. n gives the number of border pairs; n_c gives the number of borders. Standard errors (in parentheses) are clustered at the border level; 95% wild bootstrap confidence sets for the null hypothesis are reported in brackets.

stages of the Great Migration, during which many families left the South, in part because of poor schooling opportunities for their children.[31] To the extent that migration flows were greater from the counties in our paired sample with lower teacher salaries, there could be a sample selection bias in our analysis. To assess the potential size of such a bias we constructed a simple measure of net out-migration for Black families in our border counties—the ratio of Black 14–18-year-olds in 1940 relative to 4–8-year-olds in 1930—and look for a relationship between out-migration and Black teacher wages. Figure 8 shows no relationship between out-migration and teacher earnings for Black or white families. We therefore conclude that selective out-migration is unlikely to lead to a a serious bias in our estimates.[32]

IV. Conclusion

This paper provides new evidence on the link between school quality and upward mobility of children from poorly educated families during the golden age of expanding access to public education in the United States (Goldin and Katz 2008) and improving upward mobility in educational attainment (Hilger 2017). We document substantial systematic geographic variation in upward mobility in schooling attainment, and we provide evidence that location-specific differences are due in part to differences in public school resources, especially average teacher salaries. In a state-level analysis we find that educational outcomes for children of poorly educated parents are strongly tied to school quality measures, controlling for family background characteristics and statewide levels income and education. We confirm this conclusion using data on matched pairs of adjacent counties across state borders in the South, where there were often wide differences in average teacher salaries driven by statewide minimum salary laws.

Our work documents important consequences of inequalities in public schooling in the United States, especially disparities due to racial segregation in education. In many southern states, Black public school teachers earned less than half of what white teachers earned—a disparity that is all the more striking given that white teachers in the South were relatively poorly paid. In 1940, the median school-age Black child lived in a state in which the cost of living–adjusted annual salary of Black teachers was only $649 (in Virginia), while the corresponding median white student had a teacher with a cost of living–adjusted

[31] See Boustan (2016) for a recent economic history of the Great Migration and an overview of previous work on the topic. Collins and Wanamaker (2014), Black et al. (2015), and Aaronson et al. (2021) study selection into migration during the Great Migration.

[32] A regression analysis shows that the relationship between the two variables is slightly negative, although far from statistically significant. We also looked at selective migration of white families but also found no evidence of a correlation with the relative wages of white teachers.

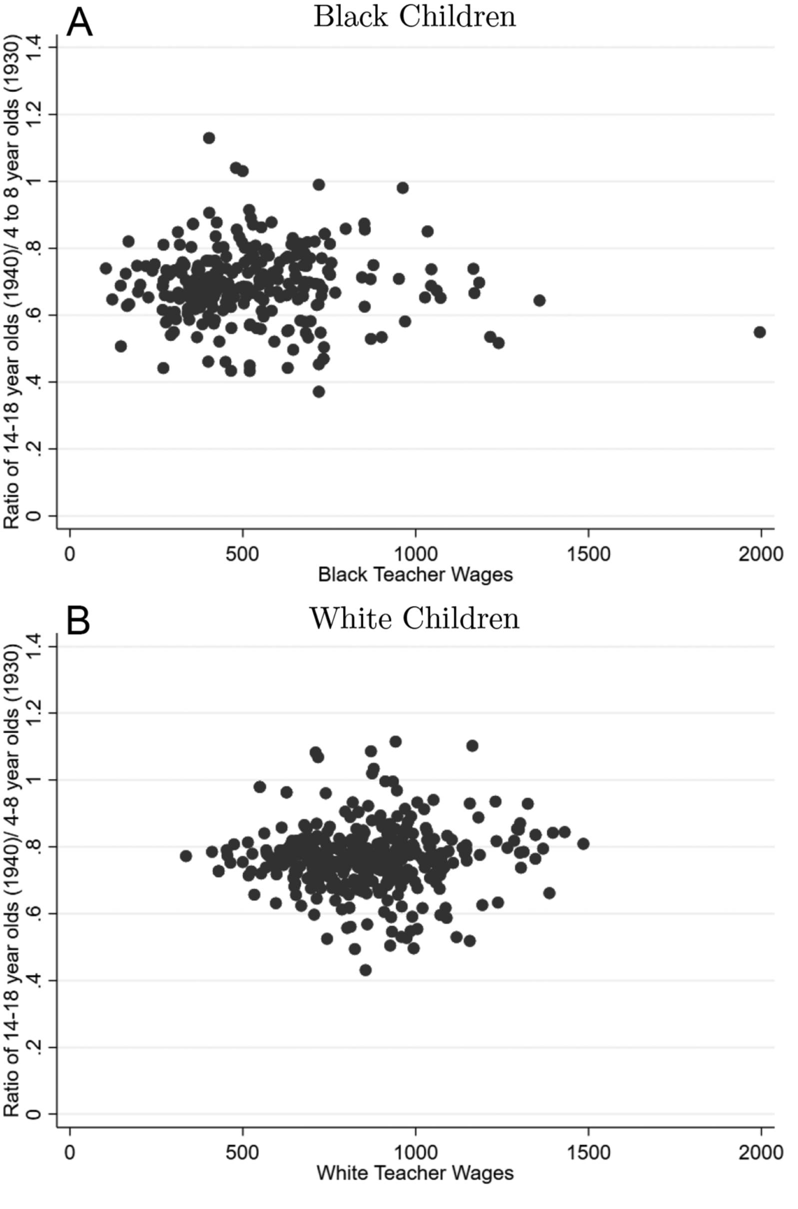

FIG. 8.—Relationship between county child-cohort population changes 1930–1940 and teacher wages. Data are from the 1930 and 1940 US census. Teacher earnings are from the 1940 census. The sample is restricted to counties included in the border analysis, with more than 20 children aged 14–18 in 1940. A color version of this figure is available online.

salary of $1,727 (in Wisconsin).[33] Using the estimated teacher salary effects on eighth-grade attainment from either our cross-state or border designs, this gap translates to a difference of approximately 0.22 in the probability of eighth-grade completion for a child whose parents had no more than 6 years of schooling. This estimate is about two-thirds of the actual observed Black-white gap in eighth-grade attainment for such children.

Similarly, on the basis of the estimated impacts on completed education, the Black-white gap in teacher wage translates to a disadvantage in completed schooling of 2–3 years. Assuming a 7% return to each year of education, an increase in resources allocated to the median Black child (to the median level of white children) would have resulted in 14%–21% higher earnings per year of work. But this may well understate the disadvantage to Black cohorts born in the 1920s because low schooling quality may have also reduced the return per year of schooling. A rough calculation suggests that our counterfactual increase in teacher pay for Black students might have increased annual earnings by 20%–30%.[34] Of course, increased education has many other potential benefits, including a longer work life, lower lifetime unemployment, higher social status, and gains in education for the grandchildren's generation. The low mobility in human capital experienced by Black children during the first half of the twentieth century was likely an important precursor to the persistence of racial inequality in labor markets over the remainder of the century (Bayer and Charles 2018) and the similarly disadvantageous pattern of income mobility experienced by African Americans in the late twentieth century (Chetty et al. 2020).

[33] In calculating these statistics, we assigned state average teacher wages from administrative records for all children aged 6–18 in the state in which the child lived, making adjustments as described in footnote 22. Given that more than half of Black children lived in the South, the median teacher salary for Black students was in a segregated state, Virginia, that paid lower salaries to Black public school teachers than their white public school teachers.

[34] Card and Krueger (1992a) find that a 10% increase in teachers' pay is associated with a 0.1 percentage point increase in the return to schooling. So a doubling of teacher pay is associated with a 1-point increase in the return to schooling. If we evaluate this gap for an individual with 8 years of schooling, this amounts to differential in annual earnings on the order of 8%.

Appendix A

Additional Tables and Figures

Table A1
Parental Education in Households with Children Aged 14–18 in the 1940 Census

	Grade Attainment by Percentile				
	10th	25th	50th	75th	90th
All	4	6	8	10	12
Foreign-born	0	4	7	8	12
Native-born	5	7	8	11	12
White parents:					
All	5	7	8	11	13
Northeast	7	8	8	12	13
Midwest	7	8	8	12	13
South	4	6	7	10	12
West	7	8	10	12	14
Black parents:					
All	2	4	5	8	10
Northeast	4	6	8	9	12
Midwest	4	6	8	9	12
South	2	4	5	7	9
West	5	7	8	11	12

NOTE.—In two-parent households, educational attainment is defined as years of schooling of the better-educated parent.

Table A2
Summary of Estimated Cross-State Models for Educational Attainment of Young Women and Men Whose Parents Have 0–6 Years of Education

	Linear Probability Model for Eighth-Grade Completion				Tobit Model for Highest Completed Grade			
	Whites		Blacks		Whites		Blacks	
	Females (1)	Males (2)	Females (3)	Males (4)	Females (5)	Males (6)	Females (7)	Males (8)
Urban	.067	.107	.123	.131	.640	.963	.962	1.122
	(.003)	(.003)	(.005)	(.005)	(.019)	(.017)	(.031)	(.028)
Farm	−.014	−.057	−.068	−.107	−.319	−.641	−.493	−.874
	(.003)	(.003)	(.004)	(.004)	(.018)	(.015)	(.029)	(.024)
Mother only	.051	.039	−.017	−.025	.406	.303	−.267	−.186
	(.003)	(.003)	(.006)	(.005)	(.021)	(.019)	(.036)	(.030)
Father only	.030	.052	.015	.017	.137	.373	.041	.242
	(.004)	(.004)	(.007)	(.006)	(.026)	(.023)	(.045)	(.038)
Moved within state since 1935	−.083	−.093	−.086	−.067	−.781	−.793	−.727	−.648
	(.002)	(.002)	(.003)	(.003)	(.014)	(.013)	(.022)	(.017)
Moved to a new state since 1935	−.146	−.173	−.173	−.151	−1.447	−1.403	−1.586	−1.379
	(.007)	(.006)	(.014)	(.012)	(.043)	(.037)	(.081)	(.072)

Table A2 (*Continued*)

	Linear Probability Model for Eighth-Grade Completion				Tobit Model for Highest Completed Grade			
	Whites		Blacks		Whites		Blacks	
	Females (1)	Males (2)	Females (3)	Males (4)	Females (5)	Males (6)	Females (7)	Males (8)
One parent born in a different state	.014	.011	−.015	−.001	.159	.135	−.145	−.065
	(.003)	(.003)	(.006)	(.005)	(.021)	(.019)	(.037)	(.030)
Both parents born in a different state	.036	.037	.010	.024	.423	.380	.008	.149
	(.004)	(.003)	(.006)	(.006)	(.024)	(.021)	(.043)	(.037)
Age 15					−.245	−.314	−.391	−.392
					(.025)	(.024)	(.038)	(.031)
Age 16					−.716	−.949	−.912	−.909
					(.024)	(.022)	(.036)	(.030)
Age 17	−.019	−.033	−.019	−.032	−.985	−1.327	−1.343	−1.352
	(.002)	(.002)	(.004)	(.003)	(.024)	(.022)	(.036)	(.029)
Age 18	−.036	−.059	−.061	−.069	−1.277	−1.613	−1.787	−1.649
	(.002)	(.002)	(.004)	(.003)	(.024)	(.021)	(.035)	(.028)
Parental education grade 1	−.003	.016	−.004	.003	.203	.186	.190	.191
	(.009)	(.008)	(.009)	(.007)	(.060)	(.052)	(.060)	(.049)
Parental education grade 2	.041	.037	.043	.025	.581	.504	.657	.631
	(.008)	(.006)	(.007)	(.006)	(.048)	(.042)	(.049)	(.040)
Parental education grade 3	.112	.091	.092	.054	1.192	1.071	1.161	1.018
	(.007)	(.006)	(.007)	(.006)	(.043)	(.037)	(.046)	(.038)
Parental education grade 4	.169	.138	.140	.084	1.724	1.513	1.567	1.369
	(.006)	(.005)	(.007)	(.005)	(.040)	(.035)	(.044)	(.036)
Parental education grade 5	.252	.215	.228	.147	2.462	2.227	2.367	2.100
	(.006)	(.005)	(.007)	(.005)	(.039)	(.034)	(.044)	(.036)
Parental education grade 6	.299	.265	.290	.203	2.925	2.650	2.854	2.521
	(.006)	(.005)	(.007)	(.006)	(.039)	(.034)	(.046)	(.038)
Constant	.354	.305	.250	.265	6.785	6.273	6.598	5.854
	(.025)	(.021)	(.027)	(.021)	(.156)	(.125)	(.175)	(.138)
Parental age fixed effects	Y	Y	Y	Y	Y	Y	Y	Y
Observations	179,665	217,897	92,061	100,646	309,515	353,649	160,288	167,212

NOTE.—Standard errors are in parentheses. In cols. 1–4, the dependent variable is eighth-grade completion in a linear probability model fit to youth age 16–18 who are observed in the 1940 census living with at least one parent with maximum parental education between 0 and 6 years. In cols. 5–8, the dependent variable is years of schooling in a tobit model fit to youth age 14–18 living with at least one parent with maximum parental education between 0 and 6 years.

Table A3
Summary of Estimated Cross-Border County Models for Educational Attainment of Young Women and Men Whose Parents Have 0–6 Years of Education

	Linear Probability Model for Eighth-Grade Completion		Tobit Model for Highest Completed Grade	
	Whites (1)	Blacks (2)	Whites (3)	Blacks (4)
Urban	.056	.101	.493	.674
	(.006)	(.007)	(.038)	(.044)
Farm	−.028	−.068	−.294	−.557
	(.004)	(.006)	(.024)	(.035)
Mother only	.045	−.027	.391	−.294
	(.004)	(.005)	(.025)	(.029)
Father only	.052	.004	.414	−.024
	(.006)	(.007)	(.034)	(.041)
Moved within state since 1935	−.103	−.083	−.829	−.762
	(.004)	(.004)	(.021)	(.025)
Moved to a new state since 1935	−.145	−.128	−1.177	−1.353
	(.009)	(.014)	(.051)	(.080)
Age 15			−.323	−.319
			(.035)	(.043)
Age 16			−.748	−.770
			(.033)	(.041)
Age 17	−.033	−.023	−1.044	−1.185
	(.004)	(.005)	(.033)	(.040)
Age 18	−.063	−.064	−1.273	−1.511
	(.004)	(.005)	(.033)	(.039)
Parental education grade 1	−.020	.029	.024	0.372
	(.013)	(.010)	(.079)	(.068)
Parental education grade 2	.028	.041	.449	.618
	(.011)	(.009)	(.065)	(.057)
Parental education grade 3	.064	.080	.899	.992
	(.010)	(.008)	(.059)	(.054)
Parental education grade 4	.125	.107	1.375	1.354
	(.009)	(.008)	(.055)	(.052)
Parental education grade 5	.224	.184	2.247	2.115
	(.009)	(.008)	(.054)	(.051)
Parental education grade 6	.291	.244	2.745	2.560
	(.009)	(.008)	(.055)	(.054)
Constant	−.140	−.147	3.533	2.966
	(.066)	(.037)	(.364)	(.234)
Parental age fixed effects	Y	Y	Y	Y
Observations	80,461	53,640	137,074	91,558

NOTE.—Standard errors are in parentheses. The sample is restricted to border county pairs analyzed in tables 4 and 5. In cols. 1 and 2, the dependent variable is eighth-grade completion in a linear probability model fit to youth aged 16–18 who are observed in the 1940 census living with at least one parent with maximum parental education between 0 and 6 years. In cols. 3 and 4, the dependent variable is years of schooling in a tobit model fit to youth aged 14–18 living with at least one parent with maximum parental education between 0 and 6 years.

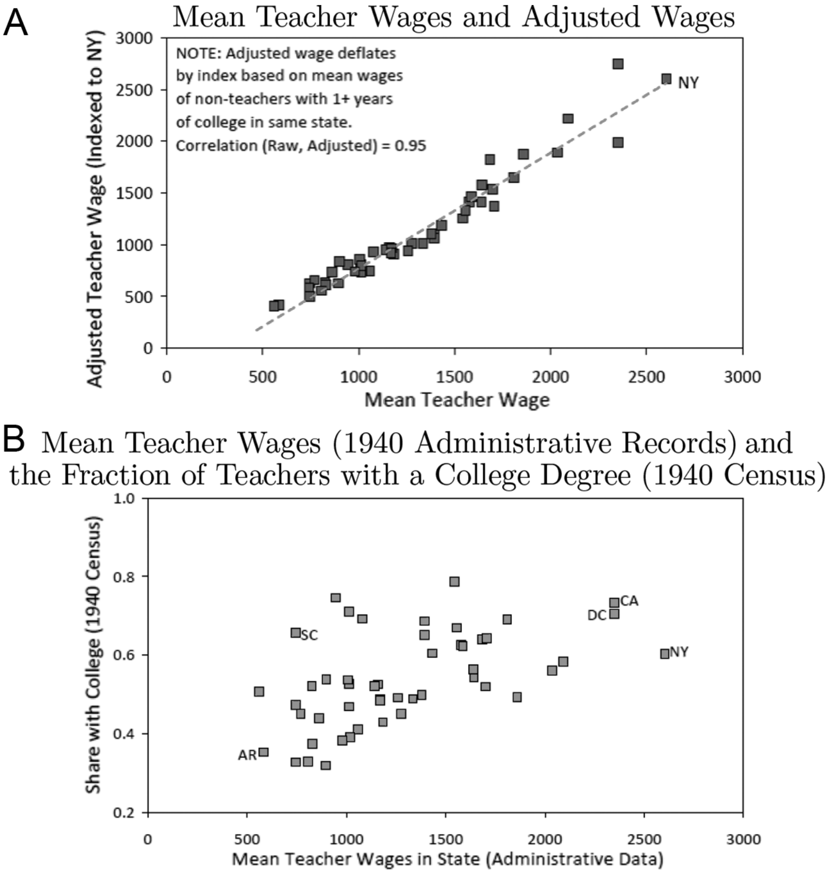

FIG. A1.—Observations about teacher wages in 1940 (whites only). A color version of this figure is available online.

Appendix B

Data Description and Data Sources

Summary of Minimum Teacher Salary Legislation in the South

Cross-state variation in teacher salaries—for both Black and white teachers—was in part the consequence of differences in state policy. In 1940, minimum teacher salaries were set according to administrative schedules in 24 states nationwide, including 10 of the southern states in our analysis.[35] Minimum salary provisions were generally part of broader legislation through which

[35] Much of our discussion draws from a research report of the National Educational Association (1940).

state boards of education provided funds to counties to supplement local expenditure for schooling. The supplementary funding was generally intended to finance the lengthening of the school term and increases in teacher pay. In exchange for state funds, counties were required to abide by state standards. Such minimum salary standards also aimed to reduce inequalities in teacher pay that resulted from differences in local tax revenues.

As shown in table B1, in most southern states salary schedules set minimum salaries that were lower for Black teachers, even for comparable levels of education, experience, and teacher certification. Such practices had not yet been successfully challenged in court as of 1939.[36] Outside the Deep South, several states with segregated schools set minimum salary standards that were the same for Black and white teachers, including Delaware, Kentucky, Oklahoma, Tennessee, and West Virginia. We briefly describe below the minimum salary standards in the southern states included in our analysis.

- *Alabama*.—In 1919, Alabama passed legislation mandating that the State Board of Education establish a standardized salary schedule in counties benefiting from state funds. An explicit minimum salary schedule appears in subsequent regulation (e.g., the 1927 School Code; Davis 1927). By 1940, all counties receiving state funding under the Minimum Program Fund were required to comply with the teacher minimum salary schedule and were required to provide a 7-month school term. Salaries of Black teachers were set to be 75% of the minimum for white teachers. The minimum for whites for a class E certificate (1 year of college or less) was \$50 per month, or \$350 for the 7-month required term. For Black teachers, this translated to \$262.50 for the 7-month term. All counties in Alabama received funding under the Minimum Program Fund in 1940 and were therefore required to comply with the minimum salary schedule.[37]
- *Delaware*.—In 1917, Delaware established a commission that surveyed its educational system and recommended a new school code, subsequently adopted. The report found that high teacher turnover and poor training were due to the low annual salaries. The new school code set

[36] As discussed in Coleman (1947), Black teachers and the National Association for the Advancement of Colored People challenged race-based salaries for teachers; the first case to reach federal courts was *Mills v. Anne Arundel County Board of Education*. In 1939, Walter Mills, a teaching principal in Anne Arundel County, sued the Maryland State Board of Education for providing lower minimum salaries for Black teachers. The federal court ruled the practice discriminatory, and in 1941 the Maryland legislature responded by equalizing minimum salaries for Black teachers. Similar lawsuits were filed during the 1940s in what came to be known as the “salary equalization movement.”

[37] See the Alabama Department of Education 1939 report (pp. 96–197).

the lowest minimum salary for a provisional elementary third-grade certificate at $400. This minimum was strongly binding for Black teachers, as the median salary of Black teachers was only $315 dollars (General Education Board 1919).

- *Georgia*.—Georgia's 1926 Equalization Act disbursed education funding to counties according to a formula developed by the State Board of Education. While there was considerable support for a minimum salary schedule, Governor Eugene Talmadge stood in active opposition.[38] However, his successor, Governor E. D. Rivers, endorsed a minimum salary schedule for teachers, and in 1937 the state passed legislation funding counties so that they could provide a minimum school term of 7 months and meet a minimum salary schedule for teachers. Minimum salaries were set lower for Black teachers than for white teachers. As of 1940, all counties in our analysis were receiving equalization funding and were thus required to comply with minimum salary schedules.
- *Kentucky*.—Legislation introduced in 1912 ended the practice of paying teachers on the basis of the number of students in the district and instead made pay conditional on the number of students in attendance. The law set wages at a minimum of $35 a month. Conditioning pay on the number of students in attendance provided incentives for teachers to keep students in attendance, but the law also provided a cap of $70 per month on salaries.
- *Maryland*.—The first minimum wage for teachers was introduced in Maryland in 1904, but it pertained only to white teachers. A minimum standard for Blacks was later introduced in 1918, at $280 per year (while the minimum for whites that year stood at $600 per year). Over the 1920s and 1930s, the minimum standards for Black teachers remained lower than those for whites, for teachers holding the same level of education and experience. Under court order, the Maryland legislature eventually equalized minimum salaries for Black and white teachers in 1941.
- *Mississippi*.—In 1924, Mississippi passed legislation mandating an $80 minimum salary for all teachers—$20 per month for a 4-month minimal school term required by the state constitution. As a practical matter, this minimum pertained only to Black teachers. Counties that received state equalization funds were required to pay white teachers a minimum of $532 for an 8-month term (and a minimum for Blacks of $161.50 for a 6-month term). However, these higher minimum standards did not apply to school districts independent of county boards. Thus, we consider the constitutional minimum standard of $80 to be applicable for Black

[38] Governor Talmadge also vehemently opposed any form of racial integration and opposed activities of the Rosenwald Fund.

teachers and consider Mississippi to be a state for which there was no binding minimum annual salary for white teachers.[39]

- *North Carolina*.—Legislation in North Carolina established a Teacher's Salary Fund in 1919. This legislation extended the constitutional minimum term length from 4 to 6 months and fixed a minimum teacher salary. By 1940, North Carolina provided funds for an 8-month school term and set teacher salaries according to a statewide schedule. The requirement for counties to abide by the minimum teacher salary schedule was clarified in communication between the state superintendent and the attorney general.[40] In 1940, the minimum salary was a relatively generous $504 for Black teachers and $656 for white teachers.
- *Oklahoma*.—Under its 1939 equalization program, the state of Oklahoma disbursed state funds to local districts maintaining an 8-month school term. In exchange, districts were required to comply with a teacher pay schedule that set the minimum at $50 per month for a first-grade elementary certificate.
- *Tennessee*.—Tennessee established a state education equalization funding program in 1925. To receive state funding, local school districts were required to provide an 8-month term and had to meet a minimum teacher salary schedule. In elementary schools the salary schedule was the same for white and Black teachers. According to Bergeron et al. (1999), the 1925 General Education Bill was hotly contested by conservatives, especially rural politicians who opposed state intervention at the local level and also opposed taxes to support the state system of higher education. Teachers, on the other hand, very much favored the law, to such extent that State Teacher's Association lobbyists, who had packed the state capitol building, were ordered off the floor of the senate. It seems that Governor Austin Peay achieved the necessary political support for this bill through a political compromise, gaining favor with fundamentalists by agreeing to not veto the Butler Act—legislation banning the teaching of evolution in public schools (Fitzgerald 2007).
- *West Virginia*.—In 1882, West Virginia became the first state to adopt a minimum salary law for teachers. The minimum for the lowest certificate

[39] Because salaries of Black teachers in Mississippi were so low, teachers often sought out other earnings opportunities. In a survey conducted by Wilson (1947), Mississippi teachers indicated that they also held the following jobs: "beautician, dental assistant, farming, hotel maid, insurance collector, kindergarten work, laundress, merchant, ministry, nurse's aid, . . . and seamstress."

[40] The *Biennial Report of the Attorney-General of the State of North Carolina* (North Carolina Department of Justice 1922) provides the following quote from the Honorable E. C. Brooks, State Superintendent of Public Instruction, Raleigh, North Carolina: "Dear Sir: You ask whether or not a county board of education may adopt a salary schedule for the teachers in the county less than that adopted by the State Board of Education. We think not."

was set at $18 per month. Local boards of education were compelled to pay Black teachers the same as white teachers with the same training, experience, and credentials. In 1909, West Virginia Superintendent Thomas Miller commented on the minimum wage legislation in response to an inquiry from Illinois educators: "The minimum salary law has produced good results in the state and while the average salary is considerably above the minimum, our enactment has prevented many districts from reducing wages below a respectable standard" (Illinois Educational Commission 1909).

- *Other states*.—Other southern states did not establish minimum salary laws prior to 1940 but sometimes had legislation regarding teacher pay. For example, legislation in South Carolina in 1924 established maximum amounts the state would allow counties to pay teachers (but no minimum) under the equalization funding as part of a plan meant to ensure a 6-month term. Table B2 includes the dates of adoption of minimum salary legislation in the southern states and the name of the relevant legislation.

Table B1
Teacher Wages and State Minimum Wages in Southern States

	Black Teachers		White Teachers	
	Census State	Minimum/ 10th Percentile[a]	Census State	Minimum/ 10th Percentile[a]
Alabama	457	262.5	864	350
Arkansas	416	210[a]	644	320[a]
Delaware	1,223	1,000	1,388	1,000
Florida	548	360[a]	1,023	640[a]
Georgia	438	175	879	280
Kentucky	835	525	902	525
Louisiana	522	245[a]	1,063	670[a]
Maryland	1,185	585	1,471	1,000
Mississippi	295	80	751	392[a]
Missouri	1,130	450[a]	997	530[a]
North Carolina	670	504	968	656
Oklahoma	787	585	940	585
South Carolina	464	260[a]	923	657[a]
Tennessee	676	320	873	320
Texas	584	330[a]	1,035	640[a]
Virginia	635	400[a]	992	540[a]
West Virginia	1,048	585	1,056	585

NOTE.—Authors' analysis. Census earnings are from the 1940 census. Minimum statutory salaries for states with minimum mandated teacher salaries are described in app. B.
[a] These states have no statutory minimum salary. Instead we give the 10th percentile of the race-specific statewide earnings among public school teachers working at least 16 weeks the previous year.

Table B2
Minimum Salary for Teachers, Southern States

State	Year Introduced	Legislative Reference
West Virginia	1882	West Virginia 15th Legislature, Adjourned Session 3(ii), Ch. 101
Maryland	1904	Maryland General Assembly 1904, Ch. 584
Kentucky	1912	Kentucky General Assembly, Regular Session, Ch. 139
Delaware	1919	97th Session, General Assembly, School Code, Art. 9
North Carolina	1919	North Carolina Public Laws and Resolutions, General Assembly 37–604, Ch. 114
Mississippi	1924	Mississippi Regular Session Appropriations, General Legislation and Resolutions 1–627
Tennessee	1925	Tennessee 64th General Assembly, Public Acts 1–708
Alabama	1927	1927 School Code, "Minimum Program Fund"
Georgia	1937	Acts and Resolutions 7-2244, 1937, Title VII, p. 882, "Equalizing Opportunities"
Oklahoma	1939	Oklahoma 17th Legislature, Regular Session, Ch. 34, Art. 14
South Carolina	1945	South Carolina General Assembly, Regular Session 1–1302, Part II, No. 223, Sec. 76
Virginia	1946	Virginia General Assembly, Extraordinary Session 3-126, House Committee Substitute for Senate Joint Resolution No. 6
Louisiana	1948	Louisiana Regular Session, Act No. 155
Texas	1949	Minimum Foundation School Laws (Gilmer-Aikin Laws): Senate Bills 115, 116, and 117
Florida	1955	Florida 35th Regular Session, General Acts 186–187, Ch. 29, 698
Arkansas	1957	Act 39 of 1957
Missouri	1985	"Excellence in Education Act"

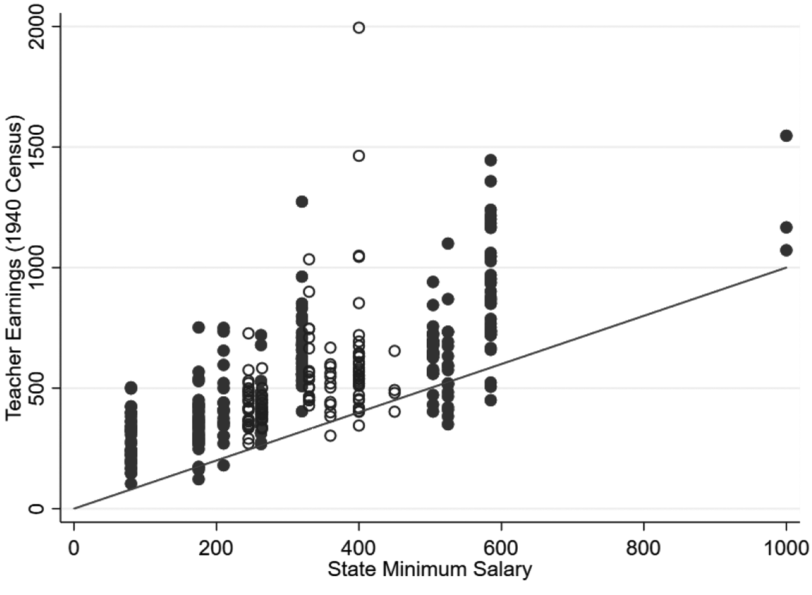

FIG. B1.—Minimum wages and county-average earnings of black teachers, southern border counties. Teacher earnings are calculated for southern state border counties using 1940 US census data. Hollow circles are observations from states for which there is no state minimum teacher salary. For these states, we use the 10th percentile of teacher earnings as the de facto minimum. A color version of this figure is available online.

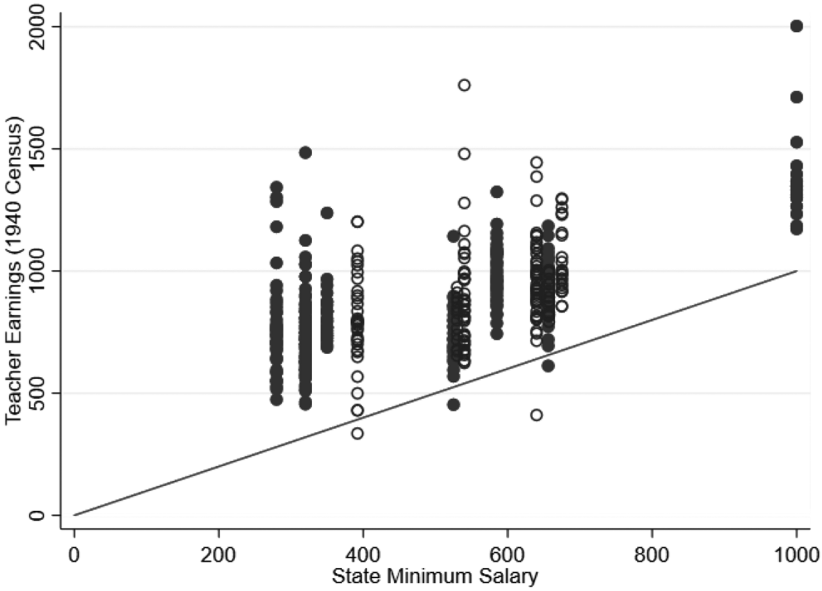

Fig. B2.—Minimum wages and county-average earnings of white teachers, southern border counties. Teacher earnings are calculated for southern state border counties using 1940 US census data. Hollow circles are observations from states for which there is no state minimum teacher salary. For these states, we use the 10th percentile of teacher earnings as the de facto minimum. A color version of this figure is available online.

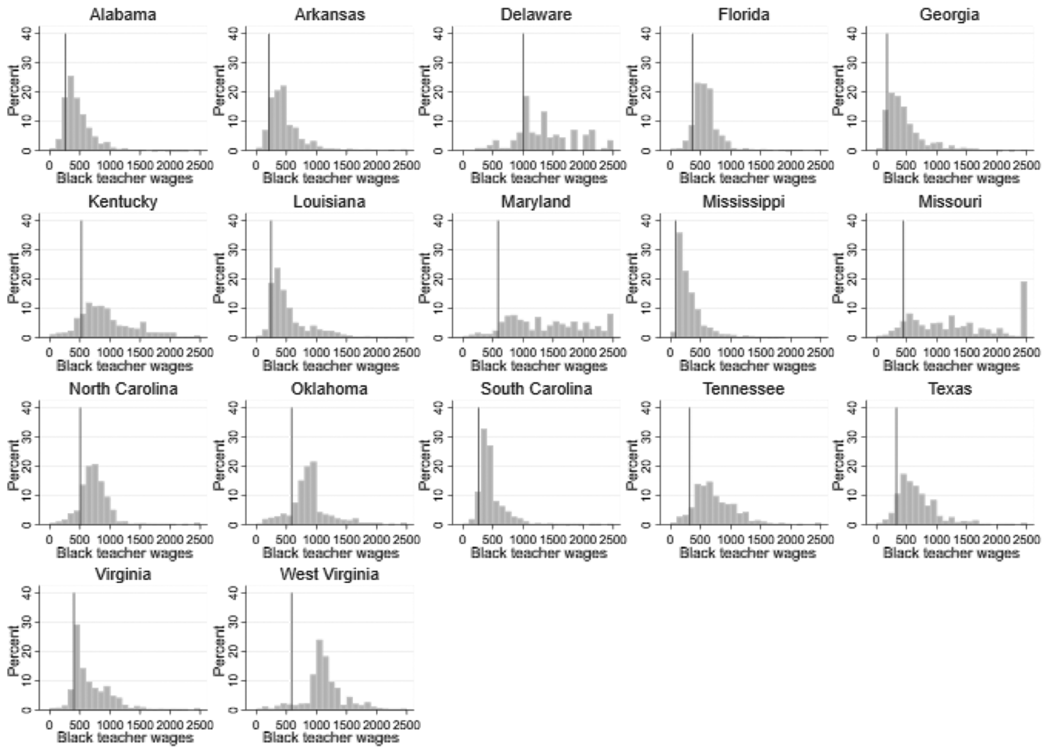

FIG. B3.—Distribution of black teacher earnings, border counties. A color version of this figure is available online.

FIG. B4.—Relationship between teacher earnings and teacher education in southern border counties. A color version of this figure is available online.

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