INCOME INEQUALITY, THE MEDIAN VOTER, AND THE SUPPORT FOR PUBLIC EDUCATION[1]

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ABSTRACT

Using a panel of U.S. states and school districts spanning 1970 – 2000, we examine the relationship between income inequality and fiscal support for public education. In contrast with recent theoretical and empirical work suggesting a negative relationship between inequality and public spending, we find results consistent with a median voter model, in which inequality that reduces the median voter's tax share induces higher local spending. Nearly 20 percent of the increase in local school spending over the 1970/2000 period is attributable to rising inequality and a falling tax share. These effects appear to operate locally: we find no relationship between inequality and state aid. We find a modest effect of rising income inequality on private school enrollment rates.

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1. Introduction

According to the U.S. Census Bureau, inequality in household income as measured by the Gini coefficient rose more than 20 percent from 1969 to 2006, an increase driven largely by income growth in the top half of the distribution (Gottschalk 1997; Goldin and Katz 2001; Piketty and Saez 2003; Autor, Katz, and Kearney 2008).[2] This surge has fueled two important and related strands of research. The first has sought causal explanations for the growth in inequality, focusing primarily on changes in the distribution of wages and earnings.[3] A second has sought to assess the social and economic consequences of growing inequality, including effects on mortality and health (Kawachi, Kennedy, and Wilkinson 1999; Deaton 2001; Mellor and Milyo 2002) crime (Kelly 2000; Fajnzylber, Lederman, and Loayza 2002), civic engagement and trust (Alesina and La Ferrara 2000, 2002; Costa and Kahn 2003), and economic growth (Benabou 1996; Forbes 2000).

Related to this second strand is a growing literature on the impact of inequality on the demand for public goods and income redistribution. The recent literature in this field has found that inequality and population heterogeneity more broadly defined tend to be associated with a *lower* level of redistribution and support for public services, both across nations and within sub-national jurisdictions (Goldin and Katz 1997; Alesina, Baqir, and Easterly 1999; Alesina, Glaeser, and Sacerdote 2001; Luttmer 2001; Lind 2007; Fernandez and Levy 2008).

[2] The Gini coefficients of household income in 1969, 1979, 1989, 1999, and 2006 were 0.391, 0.404, 0.431, 0.458, and 0.470 (U.S. Census Bureau 2008a). Gini coefficients for family income in these same years were 0.349, 0.365, 0.401, 0.429, and 0.444—consistently smaller, but increasing at a faster rate over time. The U.S. is not unique in this trend—other industrialized nations witnessed a similar rise in inequality over this period—but few experiences match that of the United States in terms of magnitude (Gottschalk and Smeeding 1997; Kenworthy and Pontusson 2005).

[3] One group of results suggests that skill-biased technological change, and—to a lesser degree—globalization and trade are the primary forces driving earnings inequality (Katz and Autor 1999; Autor, Katz, and Kearney 2008) while a competing group of papers argues that institutional factors such as the minimum wage and declining unionization have driven the rise in inequality (Card and DiNardo 2002; Dinardo, Fortin, and Lemieux 1996).

This literature often stands in sharp contrast to the basic predictions of standard voting models. In a classic set of papers, Meltzer and Richard (1981, 1983) proposed that under majority rule, income inequality can result in greater public spending whenever mean income rises relative to that of the median voter. In this model, growing wealth at the top of the income distribution lowers the tax price of raising revenue, allowing the median voter to obtain greater public services at a lower cost to them. Empirically, this model has met with mixed success over the years (Meltzer and Richard 1983; Husted and Kenny 1997; Gouveia and Masia 1998; Alesina, Glaeser, and Sacerdote 2001; Borge and Raatsø 2003; Kenworthy and Pontusson 2005). One plausible explanation for the conflicting evidence in favor of the Meltzer-Richard hypothesis is that most of its empirical tests are applied in settings where the model's assumptions are unlikely to hold. The voting model in these papers presumes direct democracy, or a representative democracy in which voting is over a single-dimensional policy space and voters have single-peaked preferences over policies (Borck 2007). The collective choice process in local government is arguably much more likely to approximate the median voter model assumptions (Fischel 2001), making our focus on local schools spending all the more appropriate.

In this paper, we draw upon a balanced panel of U.S. states and more than 10,300 school districts spanning the 1970 to 2000 period to explore the relationship between rising income inequality and fiscal support for public elementary and secondary education. In contrast with recent theoretical and empirical work suggesting a negative relationship between inequality and public spending, we find results more consistent with the Meltzer and Richard hypothesis—rising income inequality appears to be associated with modestly *higher* per-student expenditure in local school districts. This relationship is not observed at the state level, suggesting that the median voter model may be more a more appropriate representation of the political process at lower levels of government.

The empirical relationship between inequality in income and spending on public education is an important one for several reasons. First, K-12 education is a significant component of the public budget. It comprised upwards of 20 percent of aggregate state and local government expenditure in 2004, a larger share than any other general expenditure category (U.S. Census Bureau 2008b). If any public service were likely to be affected by changes in the income distribution, education should. Second, not all households directly benefit from the quality and quantity of publicly provided education. Households without school age children, the elderly, and families with children in private schools may only indirectly benefit from investments in public education. As a result, the income and demographic composition of the electorate plays an important role in the overall support for public education (Cutler, Elmendorf, and Zeckhauser 1993; Poterba 1997; Harris, Evans, and Schwab 2001). Third, the level and distribution of school spending has historically been tightly linked with income (Goldin and Katz 1997; Hoxby 1998; Fernandez and Rogerson 2001). Much has been written about the effects of income inequality on spending disparities *across* jurisdictions, but less is known about the consequences of rising inequalities *within* districts. Fourth, public education has an important redistributive aspect to it (Besley and Coate 1991; Hoxby 2003), and population heterogeneity in income and hence school spending across districts has important implications for the level and dispersion of income in subsequent generations (Benabou 1996). Fifth, although most of the theoretical literature has considered the role of population heterogeneity on general public goods provisions, a number of theories are specific to education such as the ends against the means hypothesis (Epple and Romano, 1996; Goldin and Katz, 1997). Sixth, given the specific characteristics of education spending and of our data set, we are able to a much greater degree than in previous work deal with omitted variables bias than in much of the previous literature. Our use of panel data allows us to control for the permanent within-district characteristics that would contaminate cross-sections work and the fact we have multiple observations per state per year allows us to capture state and year specific shocks to schools. We use the variation in spending brought about by court-ordered education finance reform to deal with the potential endogeneity of intergovernmental grants. We also exploit the covariance in inequality across geographically distinct districts and the strong predictions of the median voter model to construct arguably exogenous variation in inequality.

In our analysis, we begin by examining the relationship between income inequality within school districts and local spending on K-12 education. Estimating this relationship introduces several challenges. First, students of the Tiebout (1956) model might suspect a relatively low level of income inequality in local districts, to the extent the demand for education is related to income and households have the ability to sort into communities in line with their preferences for school spending. In fact, we show that the lion's share of income inequality in metropolitan areas is *within* school districts, rather than between them, a feature of MSAs that changed little over this period (see also Rhode and Strumpf 2003). Still, we address sorting in part through the use of a within-group estimator that relies on changes in income inequality within districts over time. Second, education finance during the 1970 – 2000 period was characterized by a steady shift away from local funding and toward greater state funding, fueled in part by court-ordered finance reforms (Murray, Evans, and Schwab 1998; Hoxby, 2001; Corcoran and Evans 2008). On the one hand, this shift has diminished the importance of variation in local dollars, as states have sought to equalize school spending. On the other, greater centralization may have weakened pressures to sort by income (Aaronson 1999; Nechyba 2003). Among other things, our models of local spending include controls for intergovernmental grants, as well as state-specific time trends to account for the level effects of state finance reforms.

In the second part of our analysis, we consider the relationship between growing income inequality at the *state* level and state and local education spending per student. Because education in the U.S. is a shared local and state responsibility, a complete picture of the impact of income inequality on school spending requires a look at its effects at both levels of government. Here we examine the link between inequality and two measures of fiscal support for education: state aid to local districts, and combined state and local revenues for education. The former is designed to capture the political economy of spending at the state level, while the latter captures the net effects of inequality at both levels. Finally, in our last section we examine the relationship between within-district income inequality and rates of private school enrollment.

2. Theoretical Framework

a. Inequality and Public Finance

In recent years a growing literature in public finance has examined the how inequality and population heterogeneity impact the demand for public goods. Much of this literature attempts to explain differences in government size and income redistribution across nations and within countries over time. In a classic set of papers, Meltzer and Richard (1981, 1983) proposed a simple model where the electorate votes via majority rule on a system of income redistribution funded by a proportional income tax. They showed that changes in the relative position of the decisive (median) voter in the income distribution can affect the level of redistribution and thus the size of government. Specifically, they showed that growth in mean income relative to the median lowers the "tax price" of redistribution facing the median voter, who rationally votes for greater government spending.[4] The Meltzer-Richard model has been tested empirically in multiple contexts, with mixed results. In the United States, Husted and Kenny (1997) found that extension of the voting franchise led to greater state welfare spending as the income of the median voter fell relative to statewide

[4] This assumes a non-symmetric, positively skewed income distribution. Technically, the tax *price* of local public goods—the cost to a taxpayer of an additional unit of services—is a function of more than just her tax *share* (the proportion of the total tax base held by the taxpayer). In the case of education, the tax price will also include the number of students per taxpayer, the per-student labor cost of teachers, and the like.

income.[5] In contrast, cross-national comparisons of government spending in developed countries have been less supportive of the Meltzer-Richard hypothesis (Perotti 1996; Benabou 1996). Comparing U.S. and European welfare policies, Alesina, Glaeser, and Sacerdote (2001) note that pre-tax income inequality is considerably higher in the U.S. than in European countries, while it is Europe that has larger welfare systems on average.[6]

An outgrowth of this literature asks whether racial and ethnic heterogeneity plays an intervening role in the support for public spending. Alesina, Glaeser, and Sacerdote (2001) found that racial heterogeneity can explain much of the variation in redistributive spending across developed countries. Within the U.S., Luttmer (2001) modeled individual preferences for welfare spending as dependent upon the share of beneficiaries from the same socioeconomic or ethnic group. His analysis of the General Social Survey documents that while individual support for redistribution decreases in the number of area welfare recipients, support *increases* with the fraction of local recipients who are of the same race (see also Lind 2007). Similarly, Alesina, Baqir, and Easterly (1999) showed in a cross-section of U.S. metropolitan areas that public expenditure on education, roads, libraries, sewers, and trash pickup is negatively related to within-MSA ethnic fragmentation. Cutler, Elemendorf, and Zeckhauser (1993) found analogous results at the county level, and demonstrate that the effects of population characteristics on public spending differ considerably between the county and state levels.

b. Inequality and the Support for Public Education

How income inequality and heterogeneity impact spending on public education is a particularly interesting question, for several reasons. First, as mentioned previously, it is the largest

[5] See Gouveia and Masia (1998), Borge and Raatsø (2003), and Perotti (1993) for contrasting evidence. Borck (2007) provides a review of voting models of redistribution and empirical evidence on these models.
[6] More recently, however, Kenworthy and Pontusson (2005) show that nations that experienced the largest *increases* in inequality in the 1980s and 1990s experienced the greatest increases in redistribution.

expenditure category in most state and local government budgets. Second, education is often characterized as a publicly provided private good, where benefits are targeted disproportionately to a minority of the population. To the extent interpersonal preferences for redistribution of the type exemplified in Luttmer (2001) exist, they may be particularly important in education. Finally, the presence of private alternatives to public education may alter the balance of political support for public schooling in more heterogeneous populations.

Epple and Romano (1996) provide an intriguing example of the latter. They argue that for public goods like education where private options exist, the likely majority voting equilibrium will be one in which there are two opposing coalitions of voters—one comprised of high- and low-income households who prefer a low level of expenditure on public education, and another made up of middle-income households who prefer a high level of school spending. This coalition of high and low income families oppose greater education spending for different reasons: the low income group prefers lower taxes and a greater level of consumption, while high income families opt for private schools. In this case, greater income inequality increases the likelihood of an "ends against the middle" outcome, with lower spending on public education and higher rates of private schooling.[7]

Evidence favoring the "ends against the middle" hypothesis can be found in the expansion of secondary schooling in the United States in the early 20th century, as demonstrated by Goldin and Katz (1997). They find that communities that supported the expansion of secondary education were more likely to have relatively equal income distributions, as well as populations more homogeneous in religious affiliation or ethnic background. The more heterogeneous communities lagged behind in funding secondary schooling.

Dimensions of community heterogeneity other than income have also been shown to play an important role in the political support for public education. A commonly cited example is the age

[7] See Glomm and Ravikumar (1998) for an alternative model with private schooling options, where the median voter is decisive.

distribution. Poterba (1997) found that states with higher growth in the fraction of residents over age 65 experience slower rates of growth in per-child educational expenditure, an observation consistent with a preference to vote down public programs the elderly do not benefit from themselves. He finds the effect is magnified to the extent the elderly population is racially incongruent with the school-age population. Others who have tested for effects of a growing elderly share on local, as against state, support for education have found a less stark relationship—in part because the elderly benefit from schools through their property values (Ladd and Murray 2001; Harris, Evans, and Schwab 2001; Hilber and Mayer 2007).

c. Local Population Heterogeneity and the Tiebout Model

Tiebout (1956) competition is the workhorse economic model of local public good provision. In the standard Tiebout model, families are assumed to have perfect information, a large number of communities from which they can choose to locate, costless mobility between jurisdictions, and the free entry and exit of local jurisdictions. The outcome of Tiebout competition is that households sort into homogenous communities offering their preferred level of public services and taxation. In its purest form, the Tiebout model "involves a set of assumptions so patently unrealistic as to verge on the outrageous" (Oates 1981). But despite its restrictive assumptions, the model has enjoyed much success over the years (see Ross and Yinger 1999 and Fischel 2001 for surveys). In particular, empirical research has found that local taxes and school quality are capitalized into property values, a key implication of the Tiebout model (Black 1999).

Our effort to estimate the impact of income inequality on the support for local public schools is complicated if the Tiebout model is a correct characterization of the real world. First, one might wonder: given Tiebout sorting, why would local communities have heterogeneous populations? Second, one must be concerned that because of mobility across jurisdictions, within-district income inequality is itself an outcome variable, and including it as a covariate in any type of local spending regression may subject the model to simultaneity bias.

The Tiebout model predicts homogeneity of *demand* for local public goods within communities, not necessarily homogeneity in observable population characteristics (Epple and Platt 1998). In practice, we do not observe preferences but rather correlates with demand for public goods, such as income, wealth, age, race, ethnicity, home ownership, and the like. Yet even if these observable characteristics represent noisy measures of preferences, communities are often much more heterogeneous than Tiebout might predict. This should not be too surprising. While metropolitan areas in the U.S. exhibit features of the Tiebout model to a lesser or greater degree, in practice few come close to its ideal. Labor market and commuting decisions, moving costs, public service bundling, heterogeneous housing, and a fixed supply of communities all act as barriers to perfect sorting. Communities may also be willing to tolerate population heterogeneity in exchange for economics of scale in the production of public goods (Alesina, Baqir, and Hoxby 2000). Ultimately, the weaker are Tiebout forces, the more likely conflicting household demands for public goods will be resolved through the political process.

Rhode and Strumpf (2003) found that Tiebout sorting, that is, movement due to local tax and spending policies, may be of only second order importance in the locational decisions of U.S. households. Juxtaposed against dramatically falling transportation, commuting, and communication costs—all of which in theory should increase sorting and heterogeneity across communities—they found *less* stratification of observable household characteristics and policy outcomes across municipalities over the 1850 – 1990 period. In their analysis of the Boston MSA—often held up as an archetype of Tiebout competition—they find no evidence of increased sorting in the post World War II period. Within-municipality income distributions in suburban Boston have changed little, and the between-community component of income inequality in Boston has risen only slightly since 1949. Along the same lines, Cutler, Glaeser, and Vigdor (1999) and Kremer (1997) found that sorting between neighborhoods has remained constant or declined in recent decades. The former, for example, found that the segregation of blacks within MSAs has declined over time, while the latter estimated that tract-level segregation by educational attainment remained flat between 1960 and 1990 (see also Cutler, Elmendorf, and Zeckhauser 1993).

We should stress that none of this research refutes the Tiebout hypothesis, and it is certainly the case that local public goods and taxation are important factors in household locational decisions. The results above, however, suggest that forces other than Tiebout sorting have been important enough to maintain a relatively high level of within-community heterogeneity—a condition that is a necessary one for our analysis.

We do take the potential endogeneity of population heterogeneity seriously, however, and attempt to reduce or eliminate any omitted variables bias through our choice of estimation strategies. First, we exploit the panel nature of our data, incorporating school district fixed effects into our model. This will eliminate from the analysis any unmeasured local factors that are constant over time which may simultaneously be altering trends in school spending and income inequality. Second, we exploit some characteristics of the geography of income inequality and the specifics of the Meltzer and Richard model to produce some instruments for income inequality.

3. Data and Empirical Strategy

a. Data Sources

Our analysis draws upon two panel datasets. The first is a balanced panel of demographic and financial data from more than 10,300 local U.S. school districts spanning 1970 to 2000, while the second contains similar variables measured at the state level for the same period. We constructed our school district panel by merging eight large national databases: the *Census of Population and Housing* special school district tabulations for 1970, 1980, 1990, and 2000 (U.S. Census Bureau 1973, 1982, 2002a; U.S. Department of Education 1994), the *Census of Governments: School Districts* for 1972, 1982, and 1992 (U.S. Census Bureau 1987, 1992, 1993), and the F-33 *Annual Survey of School Finances* for 2002 (U.S. Census Bureau 2002b).[8] The Census tabulations provide detailed information about household income and demographics within each school district, while the *Census of Governments* and *Annual Surveys of School Finances* represent the primary historical source of school finance data in the United States. These eight databases are supplemented by a number of others, as described in the Data Appendix.

The construction of a matched panel database of school districts spanning more than three decades presents a number of challenges. First, some school district boundaries changed over this period as a result of consolidations, splits, and unifications (a merger of separate elementary and secondary districts). All of our observations on school districts are based on their 2002 geographic definitions, such that if District A and District B merged or unified in 1995, we have combined the data from these two districts in all earlier years for comparability with 2002.[9] Districts involved in splits have been dropped from the panel, though there were very few of these cases. Second, we lose a number of school districts due to missing data in 1970 and 1990. In the 1970 Census, districts with fewer than 300 students were aggregated into one pseudo-district in each of 39 affected states, accounting for a loss of roughly 1,500 mostly rural districts. In the 1990 Census, a small number of

[8] Similar matched panel datasets of school districts have been used by Hoxby (1996, 2001) and Harris, Evans, and Schwab (2001). Prior to 1989-90, school district financial data was only available from the Census of Governments in years ending in a two or five. We thus match the 1970 and 1980 cross-sections of Census data to financial data from the 1972 and 1982 fiscal years. For consistency across the four cross-sections, we match the 1990 and 2000 Census data to financial data from the 1992 and 2002 fiscal years.

[9] For our regression models, we have assigned an indicator variable to be equal to one for all school districts involved in a merger or unification between 1972 and 2002. When an elementary and secondary district covering the same geographic territory consolidate to form a unified (K-12) district, we do not aggregate Census data from the two districts—this would be double-counting—but rather use Census data from the larger area (usually, the secondary district).

counties in California did not participate in the special school district mapping, accounting for a loss of 196 districts. After excluding districts with outlying values of per-pupil expenditure, our balanced panel contains 10,341 school districts observed in four years, for a total of 41,364 observations.[10] While this panel comprises only 75.6 percent of existing elementary and unified school districts in 2002, these districts account for 95.2 percent of elementary and unified enrollment in that year.[11] Information on the construction of our state panel database, as well as further details on our district panel, is provided in the Data Appendix.

Income inequality measures are generally not available at geographic levels smaller than states. Accordingly, we use Census data on the counts of families falling into ordered income categories to calculate inequality measures for every school district in each panel year.[12] To do this, we assume a flexible functional form for the CDF of family income in each district, and use the grouped income data to estimate the parameters of this distribution via maximum likelihood. With these parameters, we can then generate estimates of various measures of income inequality.

The procedure is implemented as follows. Suppose in a particular year there are K income groups and n_{ik} is the number of families in income group k in district i. The K groups are families with incomes $\leq a_1$, ($> a_1$ and $\leq a_2$), …, ($> a_{K-2}$ and $< a_{K-1}$), and $> a_{K-1}$ where $a_1 < a_2 < \ldots < a_{K-1}$. Let y represent income and the CDF of the income distribution $\Pr(y \leq a) = F(a|\beta_i)$, where β_i are the parameters of the assumed distribution of income for district i. Let P_{ik} be the probability of

10 School districts are excluded if their per-student local revenues are more than twice the 95th percentile nationwide, or less than 25 percent of the 5th percentile nationwide, in any year. To avoid double-counting, we include only unified and elementary-level school districts in our sample (secondary-only school districts typically cover the same geographic area as one or more elementary-only school districts). Alaska, Hawaii, and the District of Columbia are excluded (the latter two consist of only a single school district).

11 According to the F-33 *Annual Survey of School District Finances* (U.S. Department of Education 2002b) there were 13,685 elementary and unified districts in operation in 2001-02 with nonzero enrollment. These districts had a total enrollment of 45.98 million, and 77.97 percent of these were unified districts. Our 10,341 districts in 2002 had a total enrollment of 43.79 million. In our panel, 89.91 percent of districts are unified.

12 We use family income as opposed to household income due to the 1970 Census, which only reports the income of families and "unrelated individuals."

observing income in group k, where $P_{i1} = F(a_1|\beta_i)$, $P_{i2}= F(a_2|\beta_i) - F(a_2|\beta_i)$, … $P_{iK-1} = F(a_{K-1}|\beta_i) - F(a_{K-2}|\beta_i)$, and $P_{iK} = 1 - F(a_{K-1}|\beta_i)$. Therefore, the likelihood function for district i in this year is $L_i = \Sigma_k n_{ik} \ln(P_{ik})$ which is maximized through the choice of β_i.

Based on McDonald (1984), we elected to use a three-parameter Dagum (1980) distribution, also known as the Burr Type III distribution. In his paper, McDonald (1984) fit a series of statistical distributions to U.S. income in 1970 and 1980 and concluded that the Dagum distribution out-performed all other three-parameter models, as well as some four-parameter models, in terms of fit. This distribution also has the advantage of having a straightforward, closed-form solution for its moments. Given the estimated parameters of this distribution, we were able to directly calculate the Gini coefficient, Theil index, coefficient of variation, and the (logged) ratios of the 95th to 5th, 95th to 50th, and 50th to 5th percentiles of family income.[13]

Although the income data we use in this project is only categorical, the procedure outlined above generates accurate estimates of income inequality. Evans, Hout, and Mayer (2004) use categorical data on family income at the state level from various census years to estimate the parameters of the Burr III distribution for each state, and then compare the implied Gini coefficient from these estimates with those reported by the Census from the entire long form sample. The correlation coefficients from the computed state-level Ginis in family income and the reported values for the 1970, 1980, and 1990 census are 0.998, 0.996, and 0.980, respectively.

School district demographics are taken from the special Census tabulations, described in the Data Appendix. From the Census data on race, we calculate a standard index of racial heterogeneity as one minus the sum of the squared population shares of four race categories: white, black, Asian/Pacific Islander, and other (following Alesina, Baqir, and Easterly 1999; Vigdor 2002; Alesina

[13] The number of income categories reported in the Census varies over time. However, as we show in the Data Appendix, changes in the number of categories do not materially affect our estimates of income inequality, nor do they create a systematic bias.

and La Ferrara 2002, and others). This index ranges from a lower bound of zero (perfect homogeneity) to an upper bound of 0.75 (maximum heterogeneity with four groups).

Descriptive statistics for our district panel are provided in Table I.[14] All observations have been weighted by public school enrollment, such that these statistics can be thought of as characterizing the school district in which the average student resides. Several trends are worth noting. First, local education revenues per pupil rose an average of 58 percent in real terms from 1972 to 2002, at the same time local funds as a share of total per-student spending fell from an average of 54.5 percent to 42.2 percent. Second, school districts became considerably more racially diverse from 1970 to 2000, as evidenced by the near doubling of the mean index of race fractionalization. The elderly share in the average district rose almost three percentage points, from 9.5 to 12.1. Finally, the average level of income inequality within school districts rose significantly, by almost every measure. The Gini coefficient of income inequality in the average school district increased 15 percent from 1970 to 2000, while the average Theil index rose a more sizable 39 percent. As was true nationally, income inequality grew more in the top half of the distribution: the rise in the average (log) ratio of the 95th to 50th percentile of income was 20.5 percent as compared with an 8.1 percent increase in the mean (log) 50th to 5th ratio. We elaborate more on this growth in income inequality within school districts in later sections.

b. Empirical Strategy

The goal of this paper is to examine how rising income inequality has affected the fiscal support for public elementary and secondary education. We begin our analysis by examining the relationship between within-school district income inequality, and locally raised revenues for public schools. Our empirical model is similar in spirit to the demand function for local public goods

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[14] Descriptives for our state panel can be found in the Data Appendix, Table A.1.

introduced by Borcherding and Deacon (1972) and Bergstrom and Goodman (1973). In that model, observed expenditure on local public goods reflects the level desired by the median voter, which in turn is a function of the median voter's income, tax share, and taste for public spending. "Taste" for local public services is typically represented by a vector of population characteristics thought to be associated with demand for these services: age, race, educational attainment, school attendance, and homeownership are frequent examples (Rubinfeld and Shapiro 1989; Harris, Evans, and Schwab 2001; Hoxby 2001).

Of course, households choose communities in part based on unobserved preferences for taxes and school quality, and as such, observable proxies may insufficiently control for these preferences. In our model, we exploit the panel nature of our data and incorporate school district fixed effects to capture time-invariant household sorting on the fixed characteristics of school districts.[15] District fixed effects will also account for permanent features of the local tax base that determine the median voter's tax share, such as the presence of taxable commercial property or natural resources.

Our basic empirical specification for local education revenues per student in school district i in state j in year t (y_{ijt}) is given by:

$$(1) \qquad y_{ijt} = \boldsymbol{X}_{ijt}\beta + inequality_{ijt}*\gamma + IG_{ijt}*\theta + \delta_i + \delta_{jt} + u_{ijt}$$

where $\boldsymbol{X}_{ijt}$ is a vector of population and housing characteristics in school district i in year t (which includes median family income, percent of the population below the poverty line, percent of adults who are college graduates, percent school aged (5-17), percent aged 65 and older, percent living in

[15] It is also the case that much of the variation in local revenues is between districts—not within districts over time. In a regression of real local revenues per student on district fixed effects and year dummies alone, the R^2 is 0.79 (adjusted $R^2 = 0.72$).

an urbanized area, percent of housing units that are owner-occupied, percent nonwhite, and the index of racial heterogeneity), $inequality_{ijt}$ is a measure of income inequality in district i in year t, IG_{ijt} is the sum of all intergovernmental grants to district i in year t (from state and federal sources), δ_i is a school district fixed effect, and δ_{jt} is a state-by-year dummy intended to capture state-specific time trends. u_{ijt} is an idiosyncratic error term representing all other time-varying determinants of local spending in district i not accounted for by the model. All regressions are weighted using total public enrollment, so that results can be interpreted for the district attended by the typical public school student.

Our coefficient of interest in equation (1) is γ, the impact of within-school district income inequality on local per-student education revenues, holding constant certain observable district characteristics, time-varying shocks at the state level, and intergovernmental aid. Our use of district fixed effects implies that we are using a within-group estimator, where variation within school districts over time is used to identify γ and other coefficients in the model. Aside from the inclusion of income inequality in the model, equation (1) is a relatively straightforward extension of the Borcherding and Deacon (1972) and Bergstrom and Goodman (1973) approach.[16] As discussed in Section 2, the role of income inequality in this model is as a proxy for the tax share facing the median voter. If the "ends against the middle" voting model dominates, our estimated γ will pick up the effects of these opposing income coalitions on local spending. It should be noted that public education has traditionally relied heavily on property wealth, rather than income, as its primary local tax base. Unfortunately, complete data on property wealth by school district over this period is not

[16] We are not the first include income inequality in a model of local education spending—see, for example, Brown and Saks (1985), Hoxby (2000, 2001), and Urquiola (2000). However, income inequality is rarely an explanatory variable of primary interest in these papers. For example, Hoxby (2000) includes measures of income inequality and racial and ethnic heterogeneity in her district-level regressions assessing the impact of school district competition on school expenditure. These variables are intended to serve as controls, however, and she does not discuss her empirical findings on these measures.

available. In a series of robustness checks in Section 3c, we re-estimate equation (1) for a subset of school districts for which we have information about inequality in owner-occupied housing wealth.[17]

Local revenues represent a sizable fraction of overall spending on public education in the United States, but the local share varies significantly across states and over time (Table I and Corcoran and Evans 2008). School districts in Vermont, for example, provided only 6.1 percent of K-12 education revenues in 2004-05, while local districts in Pennsylvania contributed 53.9 percent to school spending in that year. Thirty years earlier, Vermont localities provided a significantly higher 57.9 percent of revenues, while Pennsylvania districts contributed a lower 46.2 percent (U.S. Department of Education 2008). Among other things, this variation reflects differences in revenue-sharing practices across states, legislative and court-ordered finance reforms altering the state-local balance over time, and economic shocks impacting local districts' ability to raise revenue. Our inclusion of state-specific year effects δ_{jt} accounts for fixed differences across states in the size of the local contribution, and the effects of temporal changes in school funding policy and economic conditions on average local spending in each state. Still, it is likely that temporal changes in state aid policies impacted local spending in ways that varied systematically with district characteristics. For example, school finance reforms in the 1980s and early 1990s often used state aid as a means to equalize spending, or to relieve tax burdens in low-wealth districts. In these cases, low-wealth districts in reform states received more generous infusions of aid than high-wealth districts, or were offered more compensatory aid formulas (Murray, Evans, and Schwab 1998; Hoxby 2001). To the extent districts with growing income inequality were more likely to benefit from finance reforms through greater aid or a lower tax price, we may improperly attribute the effects of these changes to income inequality.

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[17] These models will still fail to incorporate information about local commercial and industrial property wealth. Unfortunately, school district-level data on this form of property wealth is not available over this period.

We include state and federal aid per student (IG_{ijt}) as a covariate to capture the level effects of intergovernmental aid, but state aid will typically be endogenous to local spending in cases where aid is not administered as a flat grant. Thus, in most specifications of model (1), we instrument for intergovernmental aid in district *i* using an interaction of dummy variables for court-ordered school finance reform at the state level and district *i*'s income quartile in 1970. Equation (2) shows the first stage of this regression:

$$(2) \qquad IG_{ijt} = \sum_{k=1}^{3} overturn_{jt} * incomeq_{ijk,70} \phi_k + \boldsymbol{X_{ijt}} \Pi + inequality_{ijt} * \lambda + \theta_i + \theta_{jt} + \varepsilon_{ijt}$$

where $overturn_{jt}$ is a dummy variable that equals one if state *j* experienced a court-ordered school finance reform prior to year *t* and the $incomeq_{ijk,70}$ are dummy variables that equal one if district *i* was in quartile *k* of income in state *j* in 1970 (k = 1, 2, 3). Most existing research has made a strong case for the exogeneity of court-ordered finance reforms: Card and Payne (2002), Figlio, Husted, and Kenny (2004), and Baicker and Gordon (2006) all demonstrate that state supreme court rulings affecting school funding systems are quite difficult to predict. Consequently, most empirical research has treated court-mandated reforms as exogenous events, and we make this assumption here as well.

Finally, one might be concerned that changes over time in local income inequality are itself endogenous to the policies or performance of local school districts. High-income households without children, for example, who perceive school taxes to be too high in one district may relocate to a neighboring district, potentially affecting the income distribution in both the sending and receiving district (see Fernandez and Rogerson 1996 for a theoretical exposition). School finance equalization and the number of area jurisdictions may also influence the level of income sorting between and income inequality within school districts (Aaronson, 1999; Urquiola, 2000).

Identifying exogenous variation in inequality in this context is problematic at best. We need to isolate a factor that alters within-district inequality but has no direct impact on local school spending. Unfortunately, most of the candidate reasons for changing inequality (e.g. skill-biased technical change, globalization, and institutional factors such as the decline in unions or the real minimum wage) directly impact the level of income as well as its distribution. As we demonstrate below, most of the variation in income inequality is within-district rather than between. Likewise, there is wide regional variation in inequality with the South having the highest levels in general and New England the least. These two facts suggest that within geographically distinct areas, there should be high correlation between within-district inequality measures. A viable instrument would then be a measure of inequality from a geographically similar district. However, because schools compete for students (Hoxby 2000) and parents sort into districts on a variety of characteristics, measures of inequality in adjacent districts may contain some of the same omitted factors as its own-district measure of inequality. Therefore, constructing an instrument from geographically similar districts must balance districts that are close enough to capture a first-stage relationship but not too close as to contaminate the estimates with omitted variables bias. To balance these goals, we start by using as an instrument for the Gini coefficient of inequality in a district-year the Gini from the nearest district (as measured by great circle distance) in another county. We also use the Gini from the nearest district in another state. The average great circle distance to the nearest school in .another county is 14.4 miles and this number jumps to 66.1 miles as we move to the nearest district in another state. Given these changes in distance, we expect a large drop in the size and precision of the first-stage estimate as we move from the nearest county to nearest state version of the instrument.

As an alternate strategy we also use higher moments of the local income distribution as instruments for the mean-to-median ratio of income. In the strict form of the median voter hypothesis, the median voter's tax share is defined as the ratio of median to mean income. As inequality in the top half of the distribution increases, the tax share decline and the price of local public goods to the median voter declines. For the median voter, the distribution of income that determines the tax share is irrelevant—what is important is the tax price of local services. To illustrate this, consider two districts with the same mean and median income, but district 1 has a more positively skewed income distribution than district 2, perhaps driven by a larger share of income coming from the top few percentiles of income. In this simple case, holding all else constant the median voter model would predict the same level of spending on local goods in districts 1 and 2, since the tax share is identical for the decisive voter. However, in a cross-section of districts we would anticipate that as the skewness of income rises, the mean rises faster than the median and the tax share will fall. Subsequently, if the median voter model is correct, then the skewness of income should be a valid instrument for the tax share. Within the median voter model (and holding mean income constant to eliminate income effects) an increase in skewness in the income distribution will only increase spending through a reduction in the tax share. In this situation, the changing skewness will only change the burden of who pays for local public goods.

The second part of our analysis considers the relationship between growing income inequality at the *state* level and education spending per student. The empirical model we estimate is similar to that in equation (1), with a few modifications, including explicit controls for court-ordered finance reforms. With local revenues accounting for 45 percent of overall education spending, it may be the case that changes in local spending are augmented or offset by changes at the state level. Our goal here is to examine how—on net—income inequality has affected spending on education. Finally, our last section examines the relationship between income inequality and private schooling rates at the local school district level.

4. Results

a. Income Inequality in School Districts

Table II provides some descriptive statistics for income inequality within and between U.S. school districts over the 1970 – 2000 period. Panel A shows the distribution of 1970 – 2000 income inequality growth as measured by percent changes in the Gini coefficient and Theil index, for the 10,340 school districts in our panel. We provide both an unweighted distribution—where the unit of observation is the school district—and an enrollment-weighted distribution, which better represents the level of income inequality experienced in the typical student's school district.[18] Panel B shows the average level and growth of income inequality at the metropolitan area level for the same period, with enrollment weights and without. In this panel, we also decompose the MSA Theil index of income inequality into its within- and between- district components, and present average within and between shares over MSAs.[19]

We find that most U.S. school districts experienced growth in income inequality between 1970 and 2000. The average (median) school district witnessed an 8.2 (7.6) percent rise in income inequality as measured by the Gini coefficient, and a 24.2 (18.7) percent rise as measured by the Theil index. Close to 70 percent of all districts saw an increase in income inequality, and more than one in five saw increases of 20 percent or more, as measured by the Gini. When weighting by enrollment—which downplays the influence of sparsely populated rural districts—growth in within-

[18] We use 2000 K-12 enrollment in public schools as weights. Weighting using 1970 enrollment does not substantially affect these empirical distributions.

[19] School districts are mapped to MSAs based on their 2002 MSA assignment in the NCES Common Core of Data (N=329 MSAs). Thus a district *i* that was not a part of MSA *m* in 1970 but joined that MSA before 2002 would be counted as part of that MSA for the entire period. For a given MSA, our Theil decomposition is calculated as follows:

$$T = T_W + T_B = \sum_{k=1}^{m}\left[\frac{n_k}{n}\frac{\bar{y}_k}{\bar{y}}\right]T_k + \sum_{k=1}^{m}\left[\frac{n_k}{n}\frac{\bar{y}_k}{\bar{y}}\right]\ln\left[\frac{\bar{y}_k}{\bar{y}}\right]$$

where T_W is the within-school district component of income inequality in that MSA and T_B is the between-school district component. m is the number of districts within the MSA, n and n_k are total enrollment in the MSA and district k respectively. $\bar{y}_k$ is mean income in district k, while $\bar{y}$ is mean income in the MSA.

district income inequality appears even more substantial. In this case, the average (median) student resided in a district where the growth in inequality was 15.6 (15.7) percent as measured by the Gini and 42.5 (39.2) percent as measured by the Theil.[20] The vast majority of students were in districts where income inequality rose, and a substantial fraction lived in districts where inequality increased by 25-30 percent or more.

This increase in income inequality is also reflected in MSAs (Panel B), where we find a mean increase of 13.0 and 32.3 percent in the Gini and Theil, respectively (18.1 percent and 47.1 percent when weighting by enrollment). Decomposing the Theil index into its within- and between-school district components, we find that—for the average MSA—93.5 percent of income inequality was within school districts in 2000, while only 6.5 percent was between districts. The latter fraction was only slightly higher than in 1980, when an average of 6.0 percent of MSA-level inequality was between districts.[21] When weighting by enrollment, the between-district share is almost twice as large (11.5 percent in 2000) but arguably smaller than one might expect if there were intensive income sorting between districts. Here again the average between-district share was virtually unchanged from 1980, when an average of 11.7 percent of MSA-level inequality existed between districts.[22]

Taken together, we find that the surge in income inequality documented at the national level is reflected in the income distributions of most local school districts. This finding is not tautological. With close to 15,000 local districts in the United States, Tiebout sorting by income could have had,

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[20] Nationally, inequality in family income grew by 17.1 percent (see footnote 1).

[21] The between-district share is likely understated in 1970 due to the large number of small districts missing from the Census in that year (see Section 3a and the Data Appendix). Our 1980 data is much closer to the universe of unified and elementary districts.

[22] In 2000 the MSA with the highest share of inequality between districts was Newark, NJ, where 31.7 percent of the overall Theil index was due to between-district inequality. As one might expect, the between-district component and the number of school districts in the MSA are highly correlated, reflecting greater opportunities to sort by income. For example, in 2000, Cleveland (75 area districts), Chicago (258), and St. Louis (113) had relatively high between-district Theil indices, while Miami, Las Vegas, and Shreveport—all comprised of only one or several districts—had low between-district inequality. Unlike Rhode and Strumpf (2003), we do find a small but steady increase over time in the between-school district component in Boston and several other large MSAs. However—like those authors—we estimate the between-district component to be small relative to the within-district component (¼ the size or less).

in theory, a strong moderating effect on the growth of income inequality within local jurisdictions. Instead, we find that the average student attended a district in which income inequality rose about 16 percent. A substantial fraction of students attended districts in which income inequality rose 25 percent or more. Even in metropolitan areas—where the greatest opportunities for Tiebout sorting exist—we observe relatively low between-district income inequality over the full 1970 to 2000 period. There is little doubt that Tiebout sorting by income exists, and to a greater degree in MSAs with a larger number of districts. But forces other than income sorting appear to have been sufficiently important to maintain a relatively high level of within-community heterogeneity, an observation key to our analysis that follows.

b. The Relationship between Income Inequality and Local Education Expenditure

Table III presents our baseline estimates of the relationship between income inequality and local per-student spending on K-12 education. We begin in the first column by estimating a standard demand function for local spending per student, excluding any measure of income inequality. Then, in columns (2) and (3) we add two alternate measures of income inequality to our baseline model: the ratio of mean to median income and the Gini coefficient. Finally, in column (4) we replicate column (3) but instead instrument for intergovernmental aid using exogenous changes in school funding systems.

Our estimated coefficients in column (1) are generally of the expected sign, and similar to those found in other empirical estimates of local demand functions for education. Per-student revenues increase with median family income, with a $1,000 rise in income associated with a $28 increase in local per student spending.[23] Revenues tend to be higher in districts with high poverty rates, higher proportions of college graduates, higher proportions of renters, and a higher elderly

[23] This implies an approximate income elasticity for local spending of 0.5 at the mean.

share. Revenues tend to be lower in more urbanized districts, districts with larger school-aged cohorts, and greater racial heterogeneity.[24] The latter effect is statistically insignificant when allowing for heteroskedasticity-robust standard errors, but is modest in size: we estimate that a one standard deviation increase in the racial fractionalization index (about 0.20) is associated with a $40 lower level of per-pupil revenues, an effect size roughly the equivalent of $1,400 lower median income.

As a direct test of the Meltzer-Richard hypothesis, in column (2) we add the ratio of mean to median family income to our baseline empirical model. Consistent with that hypothesis, our estimated coefficient on the mean/median ratio is positive and statistically significant at all conventional levels, suggesting that a lower tax share can induce greater local spending. Our estimated effect is also economically significant. The average 1970 to 2000 growth in the ratio of mean to median income was 0.13, with a standard deviation of 0.19. This implies that districts with one standard deviation above-average growth in the mean/median ratio would be predicted to have $344 higher local revenues per student—substantial when compared against an overall standard deviation in the growth of local revenues of $1790. To put these results in a different light, the growth in the mean to median income over the 1970 to 2000 period generated $235 higher local revenues per student. Therefore, rising income inequality is responsible for roughly 16 percent ($235/41478) of the growth in local school spending over this period.

Columns (3) and (4) replace the mean to median ratio with our within-district Gini coefficient. The relationship between income inequality and local revenues per student is similar to that found in column (2): higher income inequality is found to be associated with higher local spending on education. Districts where the 1970-2000 increase in the Gini coefficient was one

[24] See Poterba (1997) for evidence on cohort size effects on school spending, and Oates (2005) for a survey of the "renter effect" literature. The renter effect found here is sizable: we estimate that a one standard deviation decrease in the fraction of housing units that are owner-occupied (about 0.146) is associated with a $154 increase in per-student revenues, almost 1/10 of a standard deviation in local revenues per student.

standard deviation higher than average (4.6) are predicted to have increased local per-student revenues by roughly $161, on a baseline rise in revenues of $1,478.

In column (5), we present two-stage least squares estimates where the level of intergovernmental aid is instrumented using school finance reform/income quartile interactions. The instruments are interactions between a dummy variable indicating years after court-ordered finance reform and the district's initial position in within-state income distribution. Because our model contains district fixed effects as well as (state x year) interactions, only three of the income quartile interactions are uniquely identified. The first-stage estimates for this model are reported in column (1) of Appendix table#. The dependent variable is the real state and federal revenues per pupil and that consistent with the results in Evans, Murray, and Schwab (1997), the first-stage demonstrates that court-ordered finance reform increased spending among the poorest districts. Our results indicate that after reform, real per pupil state+local revenues increased by $1531, $602, and $511 in the lowest three within-state income quartile groups, relative to what happened in the top income quartile group. All of these results are statistically significant at conventional levels and th first-stage F-test that the excluded instruments are all zero is 74 indicating that finite sample bias is not a concern (Bound, Jaeger, and Baker, 1995). Note that the 2SLS estimates in column (5) differ little from OLS estimates in column (3), save for a much lower local elasticity to state and federal grants (a larger "flypaper effect") and a near tripling of the standard error.

In Table IV, we experiment with alternative measures of within-district income inequality in columns (1) – (5): the Theil index, coefficient of variation, natural logarithm of the ratio of 95^{th} to 5^{th} percentiles of income, and a measure that divides overall income inequality into two components—inequality in the top half of the distribution (the log 95^{th} to 50^{th} ratio) and inequality in the bottom half of the distribution (the log 50^{th} to 5^{th} ratio). In columns (6) – (8) we present our two-stage least squares estimates of the effects of income inequality on local revenues for education, based on three different instruments for income inequality: the Gini coefficient in the closest district in another county or state (as an instrument for the Gini measure of income inequality), and the skewness of income (as an instrument for the mean-to-median ratio of income). As described in Section 3b, the former two are intended to identify the effect of income inequality using only that portion of income inequality attributable to regional economic conditions. Of course, the challenge is identifying districts that are close enough to capture a first stage relationship but not too close as to contaminate the estimates. We present results using a relatively close district (in another county) and a relatively distant district (in another state).

As in Table III, all estimated coefficients on income inequality have positive signs, excepting the log 50th to 5th ratio, which has a negative relationship with spending. However, only the Theil index and two-part income inequality measures have statistically significant relationships with spending, likely a reflection of these measures emphasizing varying dimensions of income inequality. Our point estimate based on the Theil index suggests an effect comparable to that found using the Gini in Table III: districts where growth in income inequality was one standard deviation higher than average (10.0) spent about \$180 more per student, on average.

The pattern of estimated coefficients in columns (4) and (5)—where income inequality is measured using two components—is also consistent with the Meltzer and Richard model. Changes in income inequality that increase mean income relative to the median should lower the tax share of the median income voter (promoting higher spending), while changes that decrease mean income relative to the median should raise the tax share (promoting lower spending). We find exactly this pattern here. As the 95th percentile of income rises relative to the median within a district, we observe mean increases in per-pupil spending. As the 5th percentile of income falls relative to the median, we observe decreases. Specifically, a one percentage point (0.01) increase in the 95th percentile relative to the median is associated with a \$9.09 increase in real per-student revenues; a one percentage point fall in the 5th percentile relative to the median is associated with a $1.89 decline in per-student revenues. Given standard deviations of 0.18 and 0.31 (respectively) in these variables over all years, a log 95/50 ratio one standard deviation above the mean is associated with $164 higher spending per pupil, and a log 50/5 ratio one standard deviation above the mean is associated with $59 lower spending per pupil. Our estimates in column (5) utilizing instrumental variables for intergovernmental grants are very similar to those estimated via OLS in column (4). In this case, the first stage is very similar to that reported in column (1) of Appendix Table #.

Of course, a one percentage point rise in the log 95/50 ratio will not have an equivalent effect on the tax share as a percentage point fall in the log 50/5 ratio. Because households at the top of the income distribution earn a disproportionate share of aggregate income, an increase in the log 95/50 ratio will do more to increase mean income than an equivalent decline in the log 50/5 ratio. To compare the magnitudes of our coefficient estimates on these inequality measures, we did the following: using our estimated parameters of the income distribution in each district, we compute the fraction of total district income earned by the bottom quartile ($\ell(0.25)$, where $\ell(x)$ is the Lorenz curve for an individual district), and the fraction of total district income comprised by the top quintile ($1 - \ell(0.75)$). In the average district over all years, the top quartile earned 48.5 percent of total income, while the bottom quartile earned 7.9 percent of income. Thus, for the average district, a hypothetical increase in income of 10 percent in the top quartile (resulting in a 10 percent rise in the 95/50 ratio) will increase mean income by 4.9 percent.[25] Similarly, a rise in income of 10 percent in the bottom quartile (resulting in a 10 percent *fall* in the 50/5 ratio) will increase mean income by 0.8 percent. Thus a proportional income change at the top of the distribution will have an effect on mean income that is roughly 6.1 times that of an proportionally equivalent income change at the

[25] Mean income can be written as a weighted average of income across the four quartiles, where the weights are the income shares. Holding constant the 0.485 income share for the top quartile, a 10 percent increase in top quartile income translates into a 4.85 percent increase in mean income (0.485*0.10 = 0.0485).

bottom of the distribution. This ratio is not far from the 4.8 ratio of our two regression coefficients in column (4), suggesting that our coefficient estimates, in relative terms, are reasonable.

Finally, columns (6) – (8) present our instrumental variables estimates of the relationship between income inequality and local school revenues per student. Columns (6) and (7) use the Gini coefficient in the nearest school district in another county or state (respectively) to instrument for the local Gini, while column (8) uses the skewness of income as an instrument for the local mean to median income ratio. The first-stage estimates for these three models are reported in columns (2)-(7) of Appendix Table #. In these models, we have two endogenous variables (real state + federal revenues per pupil and the inequality measure in family income) so we include as instruments in all models the three interactions of court-ordered reform and the within-state quartiles of family income interactions plus the instruments for the inequality measure. In columns (2), (4), and (6) of the Appendix Table, the dependent variable in the first-stage is the real state+federal revenues per student and in these models, the three reform x income quartile dummies generate estimates nearly identical to the baseline estimates in column (1) of the table. In columns (3) and (5), the dependent variable in the first-stage is the within-district Gini coefficient for family income while in column (7), the dependent variable is the within-district mean/median ratio of family incomes. As expected, there is positive correlation in the Gini coefficient within broadl-defined geographic areas but as we move from the nearest district in another county to the nearest district in another state, the coefficient on the instrument drops and the standard error rises. In column (7), the within-district skewness in family income is positively correlated with the mean to median ratio in the t-statistic on this instrument is greater than 23. In columns (2)-(7), the F-statistic for the null hypothesis that the coefficients on the identifying instruments in the first-stage are all zero are of a size so as to make finite sample bias not a concern.

The resulting two-stage least squares estimate in column (6) is large, perhaps implausibly so: for districts one standard deviation above the mean in income inequality we estimate that per-student revenues are almost $1,300 higher. The large size is of a concern in that the p-value for test of overidentifying restrictions is 0.03. Of course, the closest district in another county may not be sufficiently distant to avoid contamination from omitted variables bias, particularly in metropolitan areas. Our use of the closest district in another state in column (7) may alleviate this concern, as most districts will not be competing with schools that distance away. Unfortunately, in this case our two-stage least squares estimate becomes statistically insignificant, while remaining positive (and about half the size of our point estimate in Table III). In this case, the test of overidentifying restrictions results suggest we cannot reject the null the model is correctly specified. In column (8), where we use the skewness of the local income distribution as an instrument for the mean to median income ratio, the two-stage least squares estimate is more reasonable (1.66 times the point estimate in Table III), positive, and statistically significant.[26] Our coefficient estimate of $3,017 implies that districts with one standard deviation above average growth in income inequality (0.19) are predicted to have spent about $570 more on average per student. In this case, the first-stage is large and precise, the p-value on the test of overidentifying restrictions is large as well, the results with and without controlling for endogeneity are very similar.

c. Extensions and Robustness Checks

Empirical tests of the median voter model as an explanation for the growth of government and redistribution almost exclusively use income inequality as a measure of the median voter's tax

[26] In this case, the first stage coefficient on the skewness statistic (in thousands) is 0.005 with a *t*-statistic of 23.6.

share. Public education, however, has traditionally relied heavily on the property tax.[27] To the extent school districts rely on property taxes, it is unclear how an increase in income inequality can significantly lower the tax price of public spending without a corresponding change in the distribution of housing wealth. Econometric specifications of demand functions for public education sometimes include a property-wealth measure of tax price, such as the ratio of mean to median housing wealth (e.g. Bergstrom and Goodman 1973). Unfortunately, consistent data on property wealth by school district is not available for all years of our analysis. The 1970, 1990, and 2000 Census school district tabulations do, however, provide counts of owner-occupied homes falling into ordered valuation categories, which permit us to calculate measures of housing wealth inequality comparable to the income inequality measures used above. (They do not, unfortunately, provide any information about local commercial and industrial property wealth). We again fit the 3-parameter Dagum distribution to the housing value data in each district-year, in the same manner outlined in Section 3.

Our calculated measures of inequality in owner-occupied housing wealth (as measured by the Gini coefficient) are consistently lower on average than inequality in income, by 21 to 37 percent, although the standard deviation across districts is larger for property wealth inequality than for income inequality (Table I). The two inequality measures are also highly correlated, at 0.63, 0.63, and 0.62 in 1970, 1990, and 2000 respectively.[28]

Columns (1) and (3) of Table V present the results of our baseline regression, with our income inequality measure replaced by the Gini coefficient of housing wealth inequality. (Due to missing data in 1980 and for select districts in other years, our sample size drops to 29,536). For

[27] Among fiscally independent school districts in 2001-02, an average of 72 percent of locally raised revenues were raised from property taxes (author's calculations using U.S. Census Bureau 2002b). Aggregating to the state level, the proportion of local revenues accounted for by the property tax ranged from a low of 35 – 40 percent in Alabama, Louisiana, and Vermont, to a high of 90 percent in New Jersey and Texas.

[28] Consistent with earlier sections, we use enrollment weights in calculating these correlation coefficients. The unweighted correlations are 0.64, 0.57, and 0.48 in these three years.

comparison purposes, columns (2) and (4) estimate our baseline model of Table III using the same sample of districts used in columns (1) and (3). We find that growth in housing wealth inequality has a positive effect on per-student spending comparable with income inequality. Based on our point estimate of 15.1, districts one standard deviation above the mean in housing wealth inequality (7.2) are found to spend $109 more per student, on average. For the sample of districts for which we have both inequality measures, our point estimate for the coefficient on income inequality (23.4) implies a remarkably similar effect size: a one standard deviation increase in income inequality (5.7) is associated with $132 more spending per student. These coefficients are nearly the same when using instrumental variables for state and federal grants (columns (3) – (4)).[29]

The presence of rigid equalization programs that in some states impose a high tax price on local school spending (Hoxby 2001) may cast doubt that our observed relationship between income inequality and revenues reflects changes in the median voter's preferred level of spending. The textbook example is California's *Serrano v. Priest* ruling (1971), which effectively ended the practice of local finance in that state and centralized spending decisions at the state level (Brunner and Sonstelie 2006). Since *Serrano*, the California legislature has fixed the level of local school expenditure through the use of "revenue limits," to which the state and local school districts contribute. Due to property tax limitations set by Proposition 13 (1978), local districts collect property taxes at a fixed rate of 1

[29] We have also estimated a model comparable to that in columns (3) and (4) in which housing wealth inequality and income inequality are included simultaneously as regressors. In this case, both point estimates are positive and comparable to those in Table V, though the point estimate is smaller here for income inequality (14.0 for the housing wealth Gini and 16.8 for the income Gini). Only the estimate on housing wealth inequality is statistically significant (p=0.03), although the other is close (p=0.14). These results are suggestive that our income inequality measure in general corresponds to an underlying local inequality in property wealth, but that income inequality may play an additional role in explaining variation in local school taxes.

percent, while the state fills the remaining gap between property tax collections and the revenue limit.[30]

Given restrictions imposed by the California school finance system, one might expect to find no relationship between growing income inequality and local school spending in that state. That is, because localities have little to no leeway in determining per-student expenditure, median voters have no opportunity to respond to changes in the local income distribution through higher (or lower) taxes. On the other hand, it may be the case that the rigid system used in California *amplifies* the relationship between inequality and spending. Under a fixed local tax rate, changes in the income distribution that increase mean income will *mechanically* increase local tax collection (which in California is offset by a lower state contribution). Median voters have no opportunity to mediate a surge in local income through lower tax rates.

In fact, this is exactly what we observe in our data. Column (5) of Table V shows the results of our baseline regression model applied only to California school districts. Here our point estimate for the coefficient on income inequality is 74.1, more than twice that found in Table III. Given a standard deviation of 6.1 for the Gini coefficient across California districts, this point estimate implies an increase of $452 per student for every one standard deviation increase in income inequality. Of course, the mean level of local expenditure in California ($2,136) is higher than the national average, but even when expressed as a proportion of spending, this effect is a large one.

Table V also presents a number of alternative specifications of our baseline model. For example, the regression used in column (6) defines education spending as local revenues *per school aged child*, rather than per public school student. To the extent the rise in spending associated with

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[30] The system is actually more complex than this, and there is considerably more spending inequality in California than is commonly believed. This is due in part to a local override option, which allows districts to exceed the revenue limit, and the practice of raising funds through private educational foundations (see Brunner and Sonstelie 2006). Still, the vast majority of spending on public education is dictated by the revenue limits.

income inequality is a reflection of differential migration into private schools, we may be observing the effects of declining public school enrollment. That is, if enrollment declines and resources do not adjust at once, fewer public school students per capita can also produce a lower tax price, and an increase in spending. We do find a smaller point estimate here for our coefficient on income inequality, but the change in dependent variable prevents these coefficients from being directly comparable. Across all years, local revenues per school aged child averaged $2,605 with a standard deviation of $1,701. Districts where the increase in inequality was one standard deviation higher than average (+9.6) are predicted to have increased local revenues per child by roughly $244. Relative to the standard deviation of local spending, this effect size is only slightly smaller than that found in Table III. We directly examine the effects of income inequality on private schooling rates in our final section.

As further robustness checks, we have estimated our baseline regression models in a number of additional ways, including (1) restricting the national sample to unified (K-12) districts, (2) with monetary variables measured in natural log units, (3) without enrollment weights, (4) using a measure of household income inequality as opposed to inequality in family income (which precludes the use of 1970 data), and (5) restricting our panel to 1980 – 2000 (which recaptures approximately 2,600 districts). The results of these alternatively specified models are qualitatively similar to those found in our baseline model.[31]

d. The Relationship between Income Inequality and State-level Education Expenditure

[31] These results are available from the authors upon request. In the case of the model estimated using natural logs, we find that one standard deviation above average growth in the Gini coefficient to be associated with a 9.6 x 0.0072 = 6.9 percent rise in local revenues per student. On a baseline over all years of $2,977, this represents about $205 per pupil. When our baseline model (Table III columns (3) – (4)) is estimated without enrollment weights, our estimated coefficient on the Gini coefficient is smaller: 15.6 and 16.7, respectively. This translates into roughly $78 – 84 per pupil higher spending in districts one standard deviation above the average (5.0) in income inequality.

In Table VI we present the results of a model equivalent to that applied to our panel of school districts in Tables III to V, but estimated using a panel of state-level observations from 1970, 1980, 1990, and 2000. The state-level perspective builds upon our earlier results in several ways. Most importantly, educational expenditure in the United States is a shared local/state responsibility. A complete picture of the impact of income inequality on the fiscal support for public education requires an understanding of its effects at both levels of government. Second, income is more relevant as a tax base at the state level. Changes in the state-wide income distribution may have a direct impact the level of state aid, affecting the overall level of spending in local districts (de Bartolome 1997; Figlio, Husted, and Kenny 2004). On the other hand, the median voter model is arguably less likely to apply in state, as against local, politics.

Our model specification in Table VI is nearly identical to that used in Table III, with a few exceptions. Columns (1) and (2) use real state and local revenues per pupil as the dependent variable, while columns (3) and (4) uses only state aid to local districts. Our intergovernmental grant variable is restricted to federal grants per pupil, which we treat as exogenous. We also include a dummy variable that equals one for state *j* in year *t* if a high court in that state ruled the school finance system unconstitutional in any year prior to year *t*. Assuming that court rulings take time to have an effect on spending (Murray, Evans, and Schwab 1998) we also include a variable that indicates the number of years since a court overturn. Finally, an additional dummy variable equals one for state *j* in year *t* if a high court ruled in favor of the state prior to year *t* (i.e. upholding the existing system). This variable is intended to capture the effects of legislated school finance reforms provoked by litigation. We include separate state and year effects in these models, and—as in our district regressions—we weight observations using public school enrollment. Alaska, Hawaii, and the District of Columbia are excluded.

We find that within-state income inequality is positively related to the sum of state and local expenditure per student (column (2)), but unrelated to state aid per student (column 4), which suggests that the effects of income inequality are concentrated at the local level. Aside from median family income, few other covariates have a statistically significant relationship with education spending at the state level. States with higher rates of poverty have somewhat greater state aid on average (columns (3) – (4)), but no higher overall spending (columns (1) – (2)), and states with higher elderly populations have higher state and local spending but no higher state aid (roughly consistent with the findings of Poterba 1997 and Harris, Evans, and Schwab 2001).[32] As expected, states that experienced court-ordered finance reforms saw statistically significant increases in state aid (Murray, Evans, and Schwab 1998). [*Is the effect size found here in the state table plausible? How should it be compared with the coefficient estimated at the school district level?*]

Taken together, our state-level regressions suggest that the positive relationship between income inequality and school spending is strongest at the local level, where the assumptions of the median voter model are most likely to hold. When restricting our analysis to state aid alone, we do not observe higher aid in states where growth in income inequality is above average. That is, income inequality did not operate on education spending through additional state aid.

e. The Relationship between Income Inequality and Private Schooling

In our look at the impact of rising income inequality on public school expenditure, we have found little evidence favoring the "ends against the middle" hypothesis in which rich and poor households jointly oppose educational spending. Rather, we find results more consistent with a median voter model in which a lower tax price stimulates higher public spending. However, it may be still be the case that growing income inequality affects the support for public education in other

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[32] We find qualitatively similar results when estimating these models in logarithms, or when using current expenditures per school-aged child as the dependent variable, as in Poterba (1997).

ways—in particular, through enrollment in private schools. In Table VII we use our school district panel to examine directly the relationship between within- district income inequality and enrollment in private school. Our empirical model for the fraction of elementary and secondary-aged children enrolled in private school ($private_{ijt}$) roughly mirrors that presented earlier in equation (1):

$$(3) \qquad private_{ijt} = \boldsymbol{X}_{ijt}\beta + inequality_{ijt}{*}\gamma + EXP_{ijt}{*}\theta + \delta_i + \delta_{jt} + u_{ijt}$$

where $\boldsymbol{X}_{ijt}$ and $inequality_{ijt}$ are defined as in Section 3b and EXP_{ijt} is the level of current operating expenditures for public K-12 education in school district i in year t. As before, we include school district (δ_i) and state-by-year fixed effects (δ_{jt}) to capture fixed differences across districts in private schooling rates and temporal changes at the state level.[33] Of course, it is likely that private schooling and public expenditures are simultaneously determined. Recognizing this, in column (2) we instrument for current operating expenditure using our school finance reform/income quartile interactions introduced earlier, relying solely on the variation in expenditure that occurs through exogenous changes in school funding formulas.

Our OLS estimates in column (1) find a positive but statistically insignificant relationship between within-district income inequality and private schooling rates. Private schooling is found to rise on average with median family income, and fall with district poverty, the percent nonwhite, and the size of the school-aged population. Holding these variables constant, we observe a positive and statistically significant relationship between racial fractionalization and private school enrollment that is quite meaningful in size. A standard deviation higher 1970-2000 growth in racial fractionalization (0.157) is associated with a 0.5 percentage point higher rate of private schooling—fairly significant given a baseline private enrollment rate of 10 percent.

[33] In a regression of the proportion of K-12 enrollment in private school on district fixed effects and year dummies alone, the R^2 is 0.85 (adjusted $R^2 = 0.80$).

Our estimated positive coefficient on expenditures per student in column (1) is somewhat counter-intuitive if we believe expenditure is linked to public school quality—admittedly a tenuous assumption. But expenditures may in part be a response to past school performance. Under-performing school districts may receive an infusion of new spending designed to improve outcomes. These districts also may be more likely to have high rates of private schooling. As might be expected, our point estimate changes sign in our two-stage least squares estimates (column (2)): here, districts with higher spending on average appear to have lower rates of private schooling. Further, our point estimate on income inequality increases in size and becomes statistically significant. In this case, our coefficient estimate of 0.052 implies that districts with one standard deviation above average growth in income inequality (4.6) are estimated to have private schooling rates that are 0.24 percentage points higher, on average. While a modest-sized effect, it falls short of that implied by our coefficient on racial fractionalization. Here we find that a standard deviation higher 1970-2000 growth in racial fractionalization is associated with a 0.42 point higher private schooling rate—an effect nearly twice as large.

f. Conclusion

As national populations have become more racially and ethnically diverse, and as incomes across families has become more unequal, scholars have begun to ponder whether this growing population heterogeneity may alter the extent to which governments provide basic services such as pubic goods and the social safety net. The recent theoretical and empirical work suggests that public goods provisions and the generosity of welfare benefits are lower in more racially and ethnically diverse populations. Models specific to public education suggest a similar outcome. Growing income inequality may encourage a battle of the ends against the middle where high income people opt out of public sector into the private schools while lower income groups prefer private rather than public consumption and hence lower taxes. These forces at the ends of the income distribution may therefore lower support for public schools in more economically diverse populations.

In contrast, growing income inequality may have some unintended consequences that could spur on spending. In a simple voter model a la Meltzer and Richards, growing wage inequality at the top of the distribution reduces the price that the median voter must pay for additional public goods, thereby encouraging a greater provision of government goods and services. We examine the impact of growing income inequality on local support for public schools using panel data for over 10,000 schools over the 1970-2000 period. In contrast to the more recent literature, our results suggest that the median voter model is a more accurate description of the experience in this governmental sector. As income inequality has grown, so too has the local dollars flowing to K-12 education. Our results indicate that 20 percent of the growth in local per student K12 finance over the past 30 years have come from a drop in the tax share of the median voter, a result of rising income inequality.

The Meltzer and Richards model has been extensively tested in the past yet the results are all over the map. The strength of the results I this paper may be driven by a number of factors. First, many of the previous tests have used national or state level data while our analysis focuses on local school districts. Fischel (2001) suggests that the collective choice process in local government is arguably much more likely to approximate the median voter model assumptions. Second, we analyze from over 10,000 school districts during a period of rapidly changing income distribution, giving us tremendous statistical power. Third, the panel nature of our data allows us to sweep out a tremendous amount of potentially damaging omitted variables bias from the model.

Given the important redistributive nature of education, our results suggest that some of the potentially negative consequences of rising social inequality may he been counteracted by local governments ability to raise additional funds from growing incomes at the top of the distribution.

The long term benefits of such a transfer are beyond the scope of this paper, but are obviously a topic for future discussion.

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Data Appendix

A.1. Sources – District Panel

Our balanced panel of school districts consists of matched demographic and financial data on U.S. school districts from 1970 to 2000. We constructed this panel by merging eight large national databases: the *Census of Population and Housing* school district tabulations for 1970, 1980, 1990, and 2000 (U.S. Census Bureau 1973, 1982, 2002a; U.S. Department of Education 1994), the *Census of Governments: School Districts* for 1972, 1982, and 1992 (U.S. Census Bureau 1987, 1992, 1993), and the F-33 *Annual Survey of School Finances* for 2002 (U.S. Census Bureau 2002b).

These eight databases were supplemented by a number of others. First, because the 2000 Census school district tabulation (U.S. Census Bureau 2002a) failed to include a table for public and private school enrollment, we used Census tract level data to compute private enrollment in every school district. This procedure required overlaying boundary files for census tracts with those for unified and elementary school districts, and aggregating enrollment counts to the district level.[34] While tracts are almost always smaller than school districts, they are not necessarily contained entirely within the boundaries of one district. In cases where tracts crossed district boundaries, we allocated public and private enrollment to districts based on the fraction of tract land area in each district. Of course, this method only works well when enrollment is uniformly distributed over the tract—a less plausible assumption in rural, suburban or geographically diverse areas. In densely populated urban areas, tracts are usually contained in only one district. Thus, measurement error in our private school variable is likely to be highest in districts with the smallest populations. The use of enrollment weights in our analysis should ameliorate at least some of this error.

Second, all of our court rulings on state school finance systems are taken from Corcoran and Evans (2008), which updates Murray, Evans, and Schwab (1998) and others. Only rulings from the highest state court on the constitutionality of school funding are included. Finally, all school district consolidations, splits, and unifications (a merger of separate elementary and secondary districts) between 1970 and 2002 were researched individually, using dozens of sources, including official state documents, news accounts, minutes from school board meetings, and school district websites. All of school district observations are based on their 2002 geographic definitions, such that if District A and District B merged or unified in 1995, we have combined the data from these two districts in all earlier years, for comparability with 2002. Data was combined by aggregating across districts, or by taking a weighted average of component district characteristics as appropriate. Districts involved in splits have been dropped from the panel, although there were very few of these cases.

A.2. Sources– State Panel

For our state panel, we combined state demographic data from the decennial *Census of Population and Housing* from 1970, 1980, 1990 and 2000 [cites] with fiscal data from the *Digest of*

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[34] Boundary files can be downloaded from http://www.census.gov/geo/www/cob/ bdy_files.html (Access Date: August 9, 2003). The 2000 School District Tabulation does include one table (PCT23) that reports school enrollment for males and females age three and older, by various age categories: 3-4, 5-9, 10-14, 15-17, 18-19, etc., with no public-private distinction. Together with K-12 public enrollment counts from the *Common Core of Data*, it would be possible to use the residual (census enrollment – CCD public enrollment) as an estimate of the number of children enrolled in private school. This would, however, require the use of two different data sources in the construction of one variable, as well as critical assumptions about the fraction of five-year olds in kindergarten and the fraction of 15-17 and 18-19 year olds in secondary school.

Education Statistics in 1971, 1981, 1991, and 2001. (The *Digest* is an annual series, spanning 1969-70 to 2005-06; see U.S. Department of Education 2008). Data from the *Digest* includes local, state, and federal revenues, as well as total membership. Gini coefficients of inequality in family income were taken from Census Table S4, "Gini Ratios by State: 1969, 1979, 1989 and 1999."[35] Summary statistics for our state panel are provided in Table A.1.

Table A.1: Descriptive statistics, Panel of U.S. states, 1970 to 2000

	Means				Standard deviations			
	1970	1980	1990	2000	1970	1980	1990	2000
Real local revenues per student	2,390	2,430	3,449	3,859	949	1,044	1,522	1,379
Real state and local revenues per student	4,157	5,066	6,912	8,302	1,176	1,106	1,699	1,627
Real federal revenues per student	379	511	451	645	123	105	84	110
Real current expenditure per student	3,938	4,960	6,663	7,772	969	993	1,598	1,539
Percent of revenues from local sources	51.3	42.5	45.6	42.5	13.9	13.5	13.5	10.6
Percent of enrollment in private school	11.0	10.6	9.7	10.6	4.9	3.8	3.5	2.8
Gini coefficient of family income (x 100)	35.8	36.3	40.6	42.9	2.4	1.7	2.1	2.5
Real median family income, in thousands	44.3	43.4	48.6	52.5	6.4	4.6	7.3	6.4
Percent of households below poverty	13.9	12.5	13.4	12.4	5.6	3.0	3.5	2.8
Percent college graduates or higher	10.6	16.1	20.1	24.3	2.0	2.8	3.5	3.8
Percent school aged (5 – 17)	26.1	21.0	18.4	18.9	1.3	1.1	1.7	1.1
Percent aged 65 and older	9.9	11.2	12.4	12.3	1.5	1.8	2.0	2.0
Percent of homes owner occupied	63.5	64.8	64.3	66.1	7.0	6.6	5.8	6.0
Percent of population nonwhite	12.0	20.1	24.1	31.1	7.2	10.3	12.2	14.2
Index of race fractionalization (x 100)	-	27.5	31.7	39.0	-	11.3	12.5	13.0
Percent of population in urbanized areas	73.1	73.2	74.7	79.0	13.2	13.1	13.2	12.7
=1 if system overturn prior to this year	0.0	16.6	20.3	40.1	0.0	37.6	46.4	49.5
=1 if system upheld prior to this year	0.0	26.3	41.4	65.7	0.0	44.5	49.8	48.0
Years since court overturn	0.0	0.9	3.1	6.9	0.0	2.1	5.9	10.0

Notes: authors' calculations using a balanced panel of states (N=48 in each year; Alaska, Hawaii, and D.C. are not included). States are weighted by public K-12 enrollment. All monetary values measured in 2002 dollars.

A.3. Inequality Measures

As described in section 3, we fit the 3-parameter Dagum (or Burr Type III) distribution to Census grouped income data in local school districts to calculate measures of income inequality. For a random variable z, the cumulative distribution function for the three-parameter Dagum distribution is as follows, for $z \geq 0$ and $(a, b, p) > 0$:

$$(4) \qquad F(z) = \left[1 + \left(\frac{b}{z}\right)^{a}\right]^{-p}$$

[35] http://www.census.gov/hhes/www/income/histinc/state/state4.html (Access date: July 27, 2008).

Given the parameters *a, b,* and *p,* the r^{th} moments of this distribution are defined as:

$$(5) \qquad E[z^r] = pb^r \beta\left(1-\left(\frac{r}{a}\right), p+\left(\frac{r}{a}\right)\right),$$

where $\beta(.)$ is the complete beta function. The *c*th percentile of the income distribution is found using:

$$(6) \qquad x_c = b\left[c^{-(1/p)} - 1\right]^{-(1/a)}.$$

Dagum (1980) showed that the Gini coefficient can be calculated directly as:

$$(7) \qquad Gini = -1 + \frac{\beta(p,p)}{\beta(p, p+(1/a))},$$

and values of the Lorenz curve $\ell(x)$ calculated as:

$$(8) \qquad \ell(x) = I\left(p+\frac{1}{a}, 1-\frac{1}{a}, x^{1/p}\right),$$

where *I(.)* is the incomplete beta function. Finally, for the Dagum distribution the Theil index (or generalized entropy 1) is calculated as (McDonald, 1984):

$$(9) \qquad Theil = \frac{\psi\left(p+\frac{1}{a}\right)}{a} - \frac{\psi\left(1-\frac{1}{a}\right)}{a} - \Gamma\left(p+\frac{1}{a}\right) - \Gamma\left(1-\frac{1}{a}\right) + \Gamma(p) + 1,$$

where $\psi(z)$ is the digamma function $\psi = \frac{\Gamma'(z)}{\Gamma(z)}$.

In addition to its high level of accuracy at the state level (Evans, Hout, and Mayer 2004), we were also interest in seeing how this procedure would perform in smaller geographic areas. Using the same maximum likelihood procedure, we estimated county-specific parameters of the Dagum distribution for 1970, 1980 and 1990. While the Census does not report Gini coefficients at the county level, they do report several other aggregate measures of the income distribution. For example, in 1990 the Census reports the fraction of families in each county earning $50,000 or more. We compared these fractions to the same fraction calculated with our estimated Dagum parameters (i.e. 1 – F(50,000; *a,b,p*)). Again, the correlation between these values is high: 0.996 for 1990. Analagously, we calculated average family income in each county using the moment generating function for the Dagum distribution, and compared these to the average family income reported by the Census. The correlation between the actual and predicted values in this case was 0.997. On the whole, it appears our maximum likelihood procedure performs remarkably well.

Finally, we were concerned that changes in the number of income categories reported over time in the Census might affect our estimates of income inequality. To test for this possibility, we collapsed the 16 income groups reported in the 2000 Census to 8, and re-estimated the Dagum parameters and Gini coefficients. In a regression of *Gini_16* (the Gini coefficient estimated using 16 income groups) on *Gini_8* (the Gini estimated using 8 groups), we estimate an intercept of 0.01 and a slope of 0.96, but cannot reject the null hypotheses that the intercept is zero and the slope is one. This implies that there is no systematic bias in using a smaller number of income groups (a smaller number of groups creates classical measurement error).

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Table I: Descriptive Statistics, Panel of U.S. School Districts 1970 – 2000

	Mean					Standard deviation				
	1970/72	1980/82	1990/92	2000/02	70-00 Change	1970/72	1980/82	1990/92	2000/02	70-00 Change
Real local revenues per student	2,381	2,392	3,313	3,765	1,478	1,264	1,432	2,198	2,216	1,789
Real state and federal aid per student	1,826	2,611	3,672	4,918	3,085	760	961	1,490	1,886	1,803
Percent of revenues from local sources	54.5	45.8	45.1	42.2	-11.1	17.6	18.7	20.4	18.1	18.7
Percent of enrollment in private school	10.5	8.8	10.2	10.4	1.0	8.4	5.8	6.7	5.5	5.6
Gini coefficient of family income (x 100)	33.8	35.1	37.3	38.8	5.0	4.9	4.5	5.4	5.7	4.6
Theil index of family income (x 100)	20.1	21.7	25.1	28.0	7.8	6.7	6.3	9.0	11.6	10.0
Mean to median ratio of income	1.10	1.15	1.20	1.25	0.13	0.16	0.08	0.11	0.12	0.19
Log(95/50) ratio of family income	0.88	0.91	0.99	1.06	0.18	0.15	0.15	0.17	0.18	0.15
Log(50/5) ratio of family income	1.49	1.57	1.65	1.61	0.11	0.28	0.28	0.36	0.31	0.28
Log(95/5) ratio of family income	2.37	2.48	2.64	2.67	0.29	0.40	0.39	0.49	0.45	0.38
Coefficient of variation of family income	0.70	0.73	0.96	0.97	0.27	0.32	0.28	3.49	0.59	0.59
Real median family income, in thousands	47.28	49.72	51.62	55.55	8.64	12.50	12.46	16.19	18.18	11.57
Percent of households below poverty	13.7	12.4	13.5	12.4	-1.3	9.1	7.1	8.3	7.3	6.9
Gini coefficient of housing wealth (x 100)	28.0	-	28.7	28.3	0.02	7.6	-	6.5	7.4	6.4
Percent college graduates or higher	10.6	13.8	21.0	23.6	12.9	6.7	7.5	11.6	12.3	7.7
Percent school aged (5 – 17)	26.6	34.9	26.4	19.3	-7.6	3.8	4.7	3.9	2.8	3.3
Percent aged 65 and older	9.5	11.4	12.2	12.1	2.7	3.9	4.4	4.4	4.2	3.8
Percent of homes owner occupied	65.8	66.9	65.6	67.2	0.3	15.1	14.6	14.1	14.4	7.8
Percent of population nonwhite	15.7	19.8	24.0	30.9	15.4	16.9	20.4	22.7	25.0	15.3
Percent of population in urbanized areas	71.3	57.7	72.8	78.3	11.0	35.5	45.3	34.5	30.6	20.1
Index of race fractionalization (x 100)	16.6	22.1	26.2	33.0	17.5	16.7	18.7	19.9	20.3	15.7

Notes: authors' calculations using a balanced panel of elementary and unified school districts (N=10,341 in each year). District observations are weighted by public K-12 enrollment. School districts with a per-student local revenue more than twice the nationwide 95th percentile or less than 25 percent of the 5th percentile of per-student local revenues have been excluded. Districts in Alaska, Hawaii, and D.C. are excluded. All monetary values measured in 2002 dollars.

Table II: Income Inequality in School Districts and Metropolitan Areas – 1970 to 2000

A. Change in income inequality within school districts, 1970 to 2000

	Percent change in Gini		Percent change in Theil index	
	Weighted	Unweighted	Weighted	Unweighted
Mean	15.63%	8.22%	42.53%	24.18%
10th percentile	-2.31	-10.82	-4.23	-21.66
25th percentile	6.25	-2.38	15.86	-4.02
Median	15.71	7.55	39.23	18.71
75th percentile	24.94	18.11	63.94	45.20
90th percentile	32.66	27.58	89.02	71.45
Mean income inequality in 1970	33.75	34.14	20.12	20.92
Mean change in income inequality	5.00	2.28	7.82	3.43

B. Income inequality within metropolitan areas

	1970	1980	1990	2000	% Change
Enrollment weighted:					
Gini coefficient (x 100)	34.3	35.9	38.4	40.5	18.1
Theil index (x 100)	20.4	22.4	26.0	30.0	47.1
% within school districts	91.0	88.3	88.7	88.5	-2.7
% between school districts	9.0	11.7	11.3	11.5	27.8
Not weighted:					
Gini coefficient (x 100)	34.6	35.5	37.6	39.1	13.0
Theil index (x 100)	20.8	21.9	24.9	27.5	32.3
% within school districts	95.6	94.0	94.0	93.5	-2.2
% between school districts	4.4	6.0	6.0	6.5	47.7
Mean number of districts	16.3	18.8	18.6	18.9	-

Source: authors' calculations using a balanced panel of elementary and unified school districts (N=10,340 in each year; 329 MSAs in Panel B, except 1990 where 327 MSAs are included). Notes: Panel A uses same sample selection criteria as Table II; Panel B makes use of the full (unbalanced) panel of school districts.

Table III: Baseline Regressions, Real Local Revenues per Pupil

	(1)	(2)	(3)	(4)
Real median family income, in thousands	27.785***	34.167***	30.819***	36.874***
	(5.290)	(4.699)	(4.849)	(5.607)
Ratio of mean to median family income		1,808.1***		
		(416.241)		
Gini coefficient of family income (x100)			35.071***	35.750***
			(9.139)	(9.109)
Percent of population below poverty line	12.326***	3.225	-0.798	-1.015
	(3.681)	(3.529)	(4.085)	(4.174)
Real state and federal revenues per pupil	-0.368***	-0.362***	-0.367***	-0.270***
	(0.022)	(0.021)	(0.021)	(0.058)
Percent college graduates or higher	46.631***	36.161***	39.624***	42.949***
	(5.957)	(5.132)	(5.024)	(5.434)
Percent school aged (5-17)	-58.250***	-59.254***	-58.011***	-57.892***
	(4.993)	(4.885)	(4.960)	(4.990)
Percent aged 65 and older	14.487**	14.582**	10.218	19.236**
	(7.070)	(6.897)	(7.618)	(8.631)
Percent of housing units owner-occupied	-10.584**	-12.733**	-10.991**	-13.448**
	(5.158)	(5.149)	(5.201)	(5.358)
Percent nonwhite	1.744	0.009	0.246	-0.388
	(2.768)	(2.737)	(2.794)	(2.943)
Index of race fractionalization	-201.577	91.457	-150.885	-93.784
	(395.952)	(393.143)	(394.320)	(404.174)
Percent living in urbanized area	-2.654***	-2.178***	-2.370***	-2.426***
	(0.426)	(0.401)	(0.408)	(0.412)
Instrument for state + fed	NO	NO	NO	YES
P-value, test of overidentifying restrictions				0.265
Observations	41,364	41,350	41,351	41,351
R-squared	0.908	0.910	0.909	0.711

Notes: Robust standard errors in parentheses (*** p<0.01, ** p<0.05, * p<0.1). All models include school district fixed-effects plus state x year fixed-effects.

Table IV: Alternate Measures of Income Inequality and IV Estimates for Gini Coefficient

Dependent variable: real local revenues per pupil	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Income inequality measure:	Theil (x100)	Coefficient of Variation	Log(95/5)	Log(95/50) and Log(50/5)	Log(95/50) and Log(50/5)	Gini (IV)	Gini (IV)	Mean/Median (IV)
Income inequality	17.957***	13.796	48.001	909.259***	945.104***	240.059**	14.587	3,017.1***
	(5.406)	(17.306)	(103.562)	(222.481)	(221.983)	(94.354)	(76.461)	(539.3)
Log(50/5) ratio of family income	-	-	-	-189.437*	-244.952**	-	-	-
				(98.711)	(104.244)			
Real median family income, in thousands	30.620***	27.061***	27.892***	31.655***	37.878***	55.482***	34.547***	44.051***
	(4.820)	(5.418)	(5.132)	(4.797)	(5.577)	(13.037)	(7.937)	(6.516)
Instrument for state + fed	NO	NO	NO	NO	YES	YES	YES	YES
Instrument for inequality measure	NO	NO	NO	NO	NO	YES	YES	YES
P-value, test of overidentifying restrictions						0.031	0.731	0.209
Observations	41,351	41,328	41,351	41,351	41,351	39,440	38,949	41,350
R-squared	0.91	0.91	0.91	0.91	0.71	0.65	0.72	0.72

Notes: Robust standard errors in parentheses (*** p<0.01, ** p<0.05, * p<0.1). All models include school district fixed-effects plus state x year fixed-effects. Demographic covariates shown in Table III are also included in these regression models. In column (6), the Gini coefficient in the closest school district located in another county is used as the instrumental variable; in (7), the Gini coefficient in the closest school district in another state; in (8), the skewness of income is used as the instrumental variable for the mean to median ratio of income.

Table V: Extensions and Robustness Checks

Dependent variable: Real local revenues per pupil	(1)	(2)	(3)	(4)	(5) California only	(6) Per child
Gini coefficient of housing wealth (x100)	15.122**		16.203**			
	(6.991)		(7.265)			
Gini coefficient of family income (x100)		23.386*		24.002*	74.100***	25.504***
		(13.456)		(13.446)	(15.522)	(7.504)
Real median family income, in thousands	27.636***	29.204***	44.780***	45.066***	69.269***	30.089***
	(7.510)	(7.017)	(8.603)	(8.363)	(7.496)	(4.405)
State x year effects	YES	YES	YES	YES	Year only	YES
Instrument for state + fed	NO	NO	YES	YES	NO	NO
Observations	29,536	29,534	29,260	29,258	1,446	36,919
R-squared	0.91	0.91	0.69	0.70	0.92	0.88

Notes: Robust standard errors in parentheses (*** p<0.01, ** p<0.05, * p<0.1). All models include school district fixed-effects. Demographic covariates shown in Table III are also included in these regression models. Columns (1) – (4) only use district observations from 1970, 1990, and 2000 (the only years for which we have Census housing values). Column (5) uses only observations on California districts from 1980 – 2000 (post-*Serrano*).

Table VI: Effect of Income Inequality at the State Level

Dependent variable:	(1) State + local	(2) State + local	(3) State	(4) State
Real median family income, in thousands	121.408** (54.682)	120.696** (53.968)	99.592** (45.756)	99.659** (45.976)
Gini coefficient of family income (x100)		231.551*** (72.398)		-21.578 (98.374)
Percent of population below poverty line	49.825 (58.280)	-37.945 (69.416)	84.910 (51.756)	93.090 (58.901)
Real federal revenues per pupil	0.883 (0.831)	0.943 (0.836)	0.404 (1.012)	0.398 (1.018)
Percent college graduates or higher	134.493 (86.209)	119.171 (83.677)	-27.879 (68.744)	-26.452 (70.194)
Percent school aged (5-17)	-89.217 (79.811)	-91.548 (77.592)	-139.124 (96.043)	-138.907 (95.869)
Percent aged 65 and older	310.890*** (106.989)	280.281** (107.815)	78.815 (123.983)	81.667 (125.698)
Percent of housing units owner-occupied	-0.956 (45.896)	-39.131 (43.257)	-27.853 (48.703)	-24.295 (50.294)
Percent nonwhite	-26.075 (16.890)	-36.231** (15.908)	-4.146 (13.678)	-3.199 (14.396)
Percent living in urbanized area	-75.350** (29.885)	-67.787** (28.820)	-13.419 (23.124)	-14.124 (23.128)
=1 if state court overturn	250.949 (225.647)	235.509 (214.730)	400.817* (225.028)	402.256* (224.813)
Years since overturn	4.171 (18.110)	-5.771 (16.811)	1.942 (17.591)	2.868 (18.582)
=1 if state court upheld	171.106 (151.313)	108.582 (146.671)	-192.792 (185.814)	-186.966 (183.917)
Observations	192	192	192	192
R-squared	0.96	0.97	0.89	0.89

Notes: Robust standard errors in parentheses (*** p<0.01, ** p<0.05, * p<0.1)

Table VII: Effect of Income Inequality on Private Schooling

	(1)	(2)	(3)	(4)
Real median family income, in thousands	0.110***	0.109***	0.110***	0.111***
	(0.012)	(0.016)	(0.011)	(0.011)
Gini coefficient of family income (x100)	0.013	0.052**	0.023	-0.014
	(0.017)	(0.024)	(0.022)	(0.023)
Percent of population below poverty line	-0.058***	-0.044**	-0.060***	-0.056***
	(0.014)	(0.017)	(0.015)	(0.014)
Percent college graduates or higher	0.014	0.032*	0.015	0.016
	(0.014)	(0.018)	(0.014)	(0.013)
Percent school aged (5-17)	-0.049***	-0.132***	-0.047***	-0.048***
	(0.015)	(0.027)	(0.015)	(0.015)
Percent aged 65 and older	0.057**	0.027	0.052**	0.050**
	(0.024)	(0.031)	(0.023)	(0.022)
Percent of housing units owner-occupied	-0.010	-0.001	-0.007	-0.006
	(0.012)	(0.014)	(0.011)	(0.011)
Percent nonwhite	-0.054***	-0.039***	-0.057***	-0.051***
	(0.011)	(0.011)	(0.012)	(0.010)
Index of race fractionalization	3.186***	2.690**	3.566***	2.920***
	(1.112)	(1.079)	(1.201)	(1.081)
Percent living in urbanized area	0.001	-0.001	0.001	0.001
	(0.001)	(0.002)	(0.001)	(0.001)
Real per student current expenditures – public schools	0.0002***	-0.001***	0.0002**	0.0002**
	(0.0001)	(0.000)	(0.0001)	(0.0001)
2nd quartile competition (dists/stud or eherf)			-0.068	0.058**
			(0.044)	(0.027)
3nd quartile competition (dists/stud or eherf)			0.048	0.104***
			(0.031)	(0.032)
4nd quartile competition (dists/stud or eherf)			0.041	-0.023
			(0.029)	(0.035)
Instrument for expenditures	NO	YES	NO	NO
Observations	38,052	37,852	37,196	37,852
R-squared	0.89	0.20	0.889	0.889

Notes: Robust standard errors in parentheses (*** p<0.01, ** p<0.05, * p<0.1). Dependent variable measured as the percent of K-12 enrollment in private schools. Regressions weighted using total K-12 enrollment (public + private).