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Sean Corcoran
William N. Evans

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ABSTRACT

Using a panel of U.S. 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 on public education. We estimate that 12 to 22 percent of the increase in local school spending over this period is attributable to rising inequality.

Sean Corcoran
New York University
Institute for Education and Social Policy,
Suite 805
665 Broadway
New York, NY 10012
sean.corcoran@nyu.edu

William N. Evans
Keough-Hesburgh Professor of Economics
Department of Economics and Econometrics
447 Flanner Hall
University of Notre Dame
Notre Dame, IN 46556
and NBER
wevans1@nd.edu

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).[1] This surge has prompted two important and related strands of research. The first has sought explanations for the growth in inequality, focusing primarily on changes in the distribution of wages and earnings.[2] 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).

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[1] 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).

[2] 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 mixed 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 or representative democracy in which voting is over a single-dimensional policy space and voters have single-peaked preferences over policy options (Borck 2007). These assumptions are more likely to be approximated in local government than in national ones, yet the Meltzer-Richard hypothesis is most often examined in a national context (Turnbull and Mitias 1999; Fischel 2001; Mueller 2003).

In this paper, we draw upon a balanced panel of more than 10,300 local school districts spanning 1970 – 2000 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 *higher* per-student expenditure in local school districts, driven primarily by an increase in revenues from local sources. Given the redistributive nature of education, our results suggest that some of the potentially negative consequences of rising social inequality may have been counteracted by local government's ability to raise additional funds from growing incomes at the top of the distribution.

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 29 percent of aggregate state and local government expenditure in 2006, a larger share than any other general expenditure category (U.S. Census Bureau 2010). 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 will play 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, theoretical work has highlighted an important mechanism by which inequality can affect educational spending when there are private alternatives: a coalition of the "ends against the middle" that reduces public school spending (Epple and Romano 1996). This hypothesis has received insufficient empirical attention. Finally, public education has an important redistributive aspect to it (Besley and Coate 1991; Hoxby 2003), and heterogeneity in income and hence school spending has implications for the level and dispersion of income in subsequent generations (Benabou 1996).

In our analysis, we begin by examining the relationship between income inequality within school districts and local spending on K-12 education. Given the nature of education spending and our data set, we are able to address omitted variables to a much greater degree than in previous work. Our use of panel data allows us to control for fixed characteristics of localities that would contaminate cross-sectional analyses, and the fact that we have multiple observations per state per year allows us to control for state-specific time varying shocks to school spending. We use variation in spending brought about by court-ordered education finance reform to deal with the potential endogeneity of intergovernmental grants. Further, we exploit the strong predictions of the median voter model to construct arguably exogenous variation in income inequality, generated by changes in the skewness of the income distribution.

Estimating the relationship between within-district income inequality and local education spending introduces several challenges. First, proponents 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 hand, 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 effects to account for the level effects of state finance reforms.

We next consider the relationship between within-district income inequality and the level of expenditure and state aid per student. In making public spending decisions, local households are concerned with both their local tax burden and the overall level of per-student expenditure. In the U.S., education is a shared responsibility between local, state, and federal governments, and in many cases, the structure of state aid programs influence the tax price of local spending (Hoxby, 2001). Our models of expenditure per student thus capture the net effect of growing inequality on overall expenditure per student. Finally, we examine the relationship between within-district income inequality and rates of private school enrollment. We hypothesize that private enrollment will respond to income inequality through several channels: families' enrollment response to fiscal changes brought about by changes in income inequality, to changes in the peer composition of schools that accompanies higher inequality, or both.

2. Theoretical Framework

a. Inequality and Public Finance

In recent years a growing literature in public finance has examined 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 policies 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.[3] 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 income.[4] 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.[5]

An outgrowth of this literature asks whether racial and ethnic heterogeneity influences support for public spending, if perhaps for different reasons. 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 showed 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

[3] 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 the 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.

[4] 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.

[5] 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.

level, and demonstrated that the effects of population characteristics on public spending differ considerably between levels of government (see also Turnbull and Mitias 1999).

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 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) provided an intriguing example of the latter. They argued 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" result, with lower spending on public education and higher rates of private schooling.[6]

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

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

Katz (1997). They found that communities which 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 the growth of 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 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 found 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

The Tiebout (1956) model is the workhorse economic model of local public good provision. In the standard Tiebout framework, households sort into homogenous communities offering their preferred level of public services and taxation. While the pure Tiebout model has restrictive and generally unrealistic assumptions, it has enjoyed much success over the years (Ross and Yinger 1999; Fischel 2001). 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 pure 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 responsive to the level of public spending.

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 varying degrees, in practice, few come close to its ideal, given 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 Metropolitan Statistical Area (MSA)—often held up as an archetype of Tiebout competition—they found 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

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 key to our analysis. We do take the potential endogeneity of the income distribution seriously, and as we outline below, attempt to reduce or eliminate any omitted variables bias through our choice of estimation strategies.

3. Data and Empirical Strategy

a. Data Sources

Our analysis draws upon a balanced panel of demographic and financial data from more than 10,300 local school districts spanning 1970 – 2000. The district panel was constructed 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).[7] 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*

[7] Similar matched panel datasets of school districts were 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.

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). We identified these district changes primarily by contacting all state departments of education where such changes occurred. All of our districts are defined 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 prior years for comparability with 2002.[8] 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 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 10,359 school districts observed in four years, for a total of 41,436 observations.[9] While this panel comprises only 75.6 percent of existing elementary and

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[8] 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).

[9] 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. Alaska, Hawaii, and the District of Columbia are also excluded (the latter two consist of only a single school district).

unified school districts in 2002, these districts account for 95.2 percent of elementary and unified enrollment in that year.[10]

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.[11] 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 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 fit a three-parameter Dagum (1980) distribution, also known as the Burr Type III distribution. In his paper, McDonald 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

10 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. These districts had a total enrollment of 46 million, and 78 percent of these were unified districts. Our sample districts in 2002 had a total enrollment of 44 million. In our panel, 90 percent of districts are unified.

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

moments. For reasons described in the next section, our primary inequality measure of interest is the ratio of mean to median family income. Both the mean and median are straightforward to calculate given the estimated parameters of the Dagum distribution. For comparability, we also compute alternative measures of income inequality, including the Gini coefficient, Theil index, and natural log of the ratios of the 95th to 50th, and 50th to 5th percentiles of family income.[12]

Although the income data we use in this project is categorical, the procedure outlined above generates accurate estimates of income inequality. Evans, Hout, and Mayer (2004) used 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 compared 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 were 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 and La Ferrara 2002). 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 1. 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 public school student resides. Several trends are worth noting. Average local education revenues per pupil rose 62 percent in real terms from

[12] 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.

1972 to 2002, at the same time local funds as a share of overall per-student spending fell from an average of 54.5 percent to 42.3 percent. Real median family incomes rose an average of 18 percent over the sample and at the same time, all measures of within-district income inequality increased considerably. Within the standard median-voter model, the inverse tax share is typically measured by the mean to median income; as the dispersion of income increases, the tax price of raising funds falls. In this case the inverse tax share falls by 9 percent. Other measures of within-district inequality in family income show much larger changes. The Gini coefficient 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 95^{th} to 50^{th} percentile of income was 18 percent as compared with an 11 percent increase in the average (log) 50^{th} to 5^{th} ratio, implying the log($95^{th}/5^{th}$) ratio increased 29 percent. We elaborate more on this growth in income inequality within school districts in later sections.

Notably, school districts became considerably more racially diverse from 1970 to 2000, as evidenced by the near doubling of the mean index of race fractionalization, and the doubling of the mean percent nonwhite over this period. The elderly share in the average district rose almost three percentage points, from 9.5 to 12.1.

b. Empirical Strategy

Our goal 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 introduced by Borcherding and Deacon (1972) and Bergstrom and Goodman (1973). In these models, 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. In this instance, the public good measure is local education revenues per pupil. "Taste" for local public services is typically proxied by a vector of population characteristics thought to be associated with demand for these services: age, race, educational attainment, school attendance, homeownership, and median family income are frequent examples (Rubinfeld and Shapiro 1989; Harris, Evans, and Schwab 2001; Hoxby 2001).

Within a median voter framework, the tax share is the fraction of tax revenues that are paid for by the decisive (median) voter. This is a simple measure of the cost of raising an additional dollar in revenues and typically, as costs decline, local spending will rise. In the case of education spending, if the outcome of interest is spending per pupil, all families have one child, and all income is taxed at a constant proportional rate, then the tax share of raising an additional dollar to the median voter is simply the median over mean income. While the true tax price of local spending on education is more complex, this simple model helps illustrate the role of rising income inequality on local spending. For example, suppose that inequality increases because of a rise in income in the top of the distribution. The median will remained unchanged, but because aggregate income (as measured by the mean) has increased, the cost of an additional dollar has declined for the median voter, encouraging greater public spending.

In our initial models, we use the inverse tax share or the within-district mean to median ratio of family income. This is also a measure of inequality, in that distributions of income are more unequal as the inverse tax share rises.

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.[13] 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 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 effect 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

[13] 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 $R^2 = 0.72$).

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.[14] 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 is important to note that public education has traditionally relied heavily on property wealth, rather than income, as its primary tax base. Unfortunately, complete data on property wealth by school district over this period is not available. In a series of robustness checks in Section 4c, we re-estimate equation (1) for a subset of school districts for which we have information about inequality in owner-occupied housing wealth.[15]

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 1 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-

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[14] 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) included 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 were intended to serve as controls, however, and she did not discuss her empirical findings on these measures.

[15] 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 generally not available over this period.

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.

We include state and federal aid per student (IG_{ijt}) as a covariate to capture the effects of intergovernmental aid on local revenues; however, state aid will typically be endogenous to local spending in cases where aid is not administered as a flat grant. Thus, in many 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 initial position in its state's income distribution 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 (Fernandez and Rogerson, 1996). 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 one strategy for addressing such changes in inequality, we 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 falls and the price of local public goods to the median voter declines. To 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 A has a more positively skewed income distribution than District B, 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 A and B, 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. In 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, a change in skewness will only change the burden of who pays for local public goods.

In addition to estimating equation (1) for local revenues per student, we estimate an analogous model for per student current operating expenditures (which excludes capital expenditure) and state aid per student. In these cases the model is identical, but for the exclusion of intergovernmental grants. Here we control directly for the impact of exogenous school finance reforms on expenditure and aid by including interactions between a court-ordered finance reform and the district's quartile of median income in 1970 ($overturn_{jt}$ x $incomeq_{ijk,70}$).

4. Results

a. Income Inequality in School Districts

Table 2 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 mean to median ratio, the Gini coefficient and Theil index, for the 10,358 school districts in our panel. The numbers are weighted by school enrollment which better represents the level of income inequality experienced in the typical student's school district.[16] Panel B shows the average level and growth of income inequality at the metropolitan area level for the same period, again weighted by school enrollment. 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.[17]

We find that most U.S. school districts experienced growth in income inequality between 1970 and 2000. Over 90 percent of all districts saw an increase in income inequality as measured by the mean to median ratio, and this number is around 70 percent for the Gini and Theil indices. The average (median) student was enrolled in a school district that witnessed a 9.0 (8.1) percent rise in income inequality as measured by the mean to median ratio, a 15.6 (15.7) percent rise according to the Gini coefficient, and a 42.5 (39.2) percent rise as measured by the Theil index.[18] More than one in five districts saw increases of 20 percent or more, as measured by the Gini. Thus, 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.

To examine changes in within- versus between-district income inequality, we look at inequality in MSAs in Panel B. We observe a similar overall increase in income inequality in MSAs,

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

[17] 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.

[18] Nationally, the Gini coefficient for family income grew by 17.1 percent (see footnote 1).

with mean increases of 10.1, 17.1 and 44.3 in the mean/median ratio, Gini, and Theil, respectively. Decomposing the Theil index into its within- and between-school district components, we find that—for the average MSA—89 percent of income inequality was within school districts in 2000, while only 11.3 percent was between districts. This latter fraction has increased since 1970, when an average of 7.9 percent of MSA-level inequality was between districts.[19]

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, 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 anywhere from 9 to 42 percent, depending on the index. 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 available 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 Education Revenues and Expenditures

[19] 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).

Table 3 presents our baseline estimates of the relationship between income inequality, local revenues per pupil for K-12 education, current operating expenditures per pupil, and state aid per pupil. 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) we add the ratio of mean to median income as our primary measure of income inequality; we examine alternative measures in Table 5. Column (3) estimates the analogous model for operating expenditures per pupil, and (4) the model for state aid. All models are estimated using ordinary least squares, with district and state-by-year effects, as described in section 3b. Heteroskedasticity-robust standard errors are reported in parentheses.

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 $29 increase in local per student spending.[20] Revenues tend to be higher in districts with high poverty rates, higher proportions of college graduates, higher proportions of renters, and a higher elderly share. Revenues tend to be lower in more urbanized districts, districts with larger school-aged cohorts, and greater racial heterogeneity, though the latter is statistically insignificant.[21] Consistent with the flypaper effect literature, local revenues are lower in districts receiving greater aid, though the reduction is not one-for-one. (Later, we account for the endogeneity of state aid).

As a direct test of the Meltzer-Richard hypothesis, in column (2) we add the ratio of mean to median family income to this model. Consistent with that hypothesis, our estimated coefficient on the mean/median ratio is positive and statistically significant at conventional levels, suggesting that a

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

[21] 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.

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 0.10, with a standard deviation of 0.11. This implies that districts with one standard deviation above-average growth in the mean/median ratio would be predicted to have $197 higher local revenues per student—substantial when compared against the standard deviation of local revenue growth of $1,486. To put these results in a different light, if the average 1970 – 2000 growth in the mean/median ratio generated $179 higher local revenues per student; this would make it responsible for roughly 12 percent ($179/$1,486) of the overall growth in local funding per-student over this period.

Because education is a shared responsibility of local and state government, this increase in local taxes for public schools may not translate into greater spending per student, if these revenues are offset by a similar reduction in state aid. It may also be the case that the increase in local spending associated with inequality is actually a reflection of falling tax prices in those districts, through state aid. If state aid formulas were re-written during this period to provide more generous matching grants to poor districts, for example, and these districts were also experiencing rising inequality, our finding may simply be picking up the effects of a lower tax price.

Column (3) of Table 3 looks at the net effect of income inequality on current expenditure per student. All of the same covariates are included as in the model of column (1), though we also control directly for the effects of school finance reform at each quartile of the district income distribution. We find that, on net, rising income inequality was associated with higher per student spending, though the magnitude is attenuated from that found in model (1). The results of column (4) suggest that the correlation between inequality and local spending is unlikely to be due to the influence of favorable tax price due to state aid. We find that rising income inequality within districts to be associated with *lower* state aid per student rather than more, which would be expected if state aid programs were driving our result. On average, a relative reduction in state aid appears to have accompanied the rise in local taxes.

In Table 4 we address two shortcomings of our OLS models, the endogeneity of intergovernmental aid and the influence of inter-district sorting on our estimate of the coefficient on income inequality. Because local taxes are responsive to state aid programs, and state aid is affected by local spending, our aid variable should not be considered exogenous. In column (2), we present two-stage least squares (2SLS) estimates where the level of real state and federal aid is instrumented using interactions between a dummy variable indicating years following a court-ordered finance reform and the district's 1970 position within its state income distribution. Because our model contains district fixed effects as well as state-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 1. Consistent with Evans, Murray, and Schwab (1997), the first stage demonstrates that court-ordered finance reform increased intergovernmental aid the most among the poorest districts. Our results indicate that after reform, state and federal revenues per pupil increased by $1,659, $702, and $326 in the lowest three income quartiles, relative to what happened in the top income quartile group. All of these results are statistically significant at conventional levels, and the first-stage *F*-test that the excluded instruments are all zero is 70 indicating that finite sample bias is not a concern (Bound, Jaeger, and Baker, 1995). Note that the 2SLS estimates in column (2) differ little from OLS estimates repeated in column (1), save for a lower local elasticity to state and federal grants (a greater "flypaper effect") and a larger standard error.

An additional concern in our OLS model is the possibility that households sort into or out of districts in response to changes in school spending (and in particular, local taxes). To the extent this mobility varies by income, such sorting will affect the income distribution. To illustrate, suppose an initially heterogeneous district witnesses a rise in inequality through growth at the top of the income distribution. The median voter responds to this change in tax price through greater local spending (our observed correlation between inequality and local revenue). High-income households, however, may respond over time to the rise in taxes by moving to another district, which reduces income inequality in the sending district. Such a response would tend to bias our estimate of the coefficient on income inequality toward zero. This example also generates a prediction that the effect of income inequality on local spending should be greatest in districts with the least opportunity for sorting, a point we return to in a later section.

Column (3) presents two stage least squares using the skewness of family income as an instrument for the mean to median income ratio. The first stage estimates for these three models are reported in columns (2) - (3) of Appendix Table 1. In this model, we have two endogenous variables (real state and federal revenues per pupil, and income inequality) so we include as instruments in all models the finance reform/income quartile dummies and the instrument for inequality. In the appendix table, the first stage dependent variable is real state and federal revenues per student, and the three reform/income quartile instruments generate estimates nearly identical to the baseline in column (1). In column (3), the dependent variable is the within-district mean/median ratio of income. As expected, the skewness in income is positively correlated with the mean/median ratio, and the *t*-statistic on this instrument is greater than 21. In columns (2) – (3), 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 (3) of Table 4 is nearly twice as large as that found in columns (1) and (2), and is statistically significant. We cannot, however, reject the null hypothesis that the OLS estimate in column (1) and the 2SLS estimate in column (3) are the same. For districts one standard deviation above the mean in income inequality, we estimate that local revenues per student are almost $327 higher, on average, or 22% of growth in local spending per student over this period.[22] In this case, the first stage is large and precise; the *p*-value on the test of over-identifying restrictions is large as well, and the results with and without controlling for endogeneity are statistically indistinguishable.

In Table 5, we experiment with alternative measures of within-district income inequality: the Gini coefficient, Theil index, natural logarithm of the ratio of 95th to 5th percentiles of income, and a measure that divides overall income inequality into two components—inequality in the top half of the distribution (the log 95/50 ratio) and inequality in the bottom half of the distribution (the log 50/5 ratio). Estimates in columns (2) and (4) are two-stage least squares estimates of the effects of income inequality on local revenues for education, using the Gini and Theil index respectively, and the skewness of income as an instrument for inequality.

As in Table 3, all estimated coefficients on income inequality have positive signs, excepting the log 50/5 ratio, which has a negative relationship with spending. All are statistically significant, with the exception of the log 95/5 ratio. Our point estimate based on the Gini coefficient an effect comparable to that found using the mean to median ratio in Table 3: districts where growth in income inequality was one standard deviation higher than average (5.0) spent about $172 more per student on average ($689 based on the 2SLS estimate).

The pattern of estimated coefficients in column (6)—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 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 increases in per-pupil spending, on average. As the 5th percentile of income falls relative to the median, we observe

[22] The first stage coefficient on the skewness statistic (in thousands) is 5.1E-6 with a *t*-statistic of 21.15.

decreases. Given the standard deviation of growth in these variables of 15 and 28 points, respectively, one standard deviation higher growth in the log 95/50 ratio is associated with $136 higher spending per pupil, and one standard deviation higher growth in the log 50/5 ratio is associated with $53 lower spending per pupil.

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.[23] 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 a proportionally equivalent income change at the bottom of the distribution. This ratio is not far from the 4.2 ratio of our two regression coefficients in column (4), suggesting that our coefficient estimates, in relative terms, are reasonable.

[23] 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).

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 share. Public education, however, has traditionally relied heavily on the property tax.[24] To the extent school districts rely on property taxes, it is unclear how an increase in income inequality will 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 often include a property wealth measure of the tax share when available, such as the ratio of mean to median housing wealth (e.g. Bergstrom and Goodman 1973). Unfortunately, consistent data on property values 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 information about local commercial and industrial property wealth). To do this, we again fit the three-parameter Dagum distribution to the housing value data in each district-year, in the manner described in Section 3.

Our calculated measures of inequality in owner-occupied housing wealth (as measured by the mean to median housing value ratio and the Gini coefficient) are consistently lower on average than inequality in income, although the variation across districts is larger for property wealth inequality than income inequality (Table 1). The two inequality measures are strongly correlated, at 0.45, 0.53, and 0.70 in the case of the mean to median ratio in 1970, 1990, and 2000, and 0.60, 0.62, and 0.61 in in the case of the Gini coefficient.

24 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.

Columns (1) and (2) of Table 6 present the results of our baseline regression, with our income inequality measure replaced by the mean to median ratio of housing wealth. (Due to missing data in 1980 and for select districts in other years, our sample size drops to 29,391). We find that growth in housing wealth inequality is also positively related to local spending per student, and its coefficient is also comparable with that for income inequality. Here, districts one standard deviation above the mean in 1970 – 2000 growth in housing wealth inequality (0.137) are found to spend \$108 to \$120 more per student on average. The estimated coefficient is quite similar when using instrumental variables for state and federal grants (column (2)).

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 case 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 percent, while the state fills the remaining gap between property taxes and the revenue limit.[25]

Given restrictions imposed by the California school finance system, one would not expect the median voter model to apply at the local level in that state. 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. Because California

[25] 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.

represents a sizable share of the districts in our panel, we re-estimated our baseline OLS and 2SLS models excluding California. The results are presented in columns (3) and (4) of Table 6. Their exclusion has a very minor effect on our estimates.

In an additional specification, shown in column (5) of Table 6, we interact our income inequality measure with an indicator of the extent of between-district sorting opportunities: the number of districts per student within a 25-mile radius (see Hoxby, 2000). We divide districts into quartiles based on the number of available districts per student, as measured in 2000, with the lowest (omitted) quartile representing the least district choice. Districts in the lowest quartile tend to be in states with a relatively small number of geographically large districts (such as the South and West), while those in the highest quartile tend to be in states with many community-based districts (such as the Northeast and Midwest). We find that the largest coefficient on income inequality is indeed among districts with the fewest sorting opportunities. The effect for districts in all other quartiles is substantially smaller, and these differences, compared to the lowest quartile, are statistically significant at conventional levels.

We have estimated a number of alternative specifications not shown here. For example, we re-estimated our baseline model of Tables 3 and 4 using local revenues *per school aged child*, rather than per public school student. This allows for the possibility that increases in spending associated with income inequality are a reflection of movement into private schools which mechanically reduces public enrollment and inflates public spending per student, if resources are slow to adjust. We find a point estimate that is only marginally smaller than that found in our original model. In addition, we have estimated models that (1) restrict the sample to unified (K-12) districts, (2) use monetary variables measured in natural log units, (3) omit enrollment weights, (4) use a measure of household income inequality as opposed to inequality in family income (which precludes the use of 1970), and

(5) restrict our panel to 1980 – 2000 (which recaptures approximately 2,600 districts). The results of these models are qualitatively similar to those found in our baseline case.[26]

d. The Relationship between Income Inequality and Private School Enrollment

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 support for public education in other ways—in particular, through enrollment in private schools. In Table 7 we use our school district panel to look directly the relationship between within- district income inequality and enrollment in private school. Our empirical model for the percent of school-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.[27] Of course, it is likely that public expenditure is responsive to private schooling rates, so in column (2) we instrument for current operating expenditure per student using our school finance reform/income quartile interactions

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[26] Results are available from the authors upon request.

[27] 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).

introduced above, relying solely on variation in expenditure that occurs through exogenous changes in school funding formulas. We present results using two measures of income inequality: the mean to median ratio of income, and the Gini coefficient.

Our OLS estimates in column (1) and (3) 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 also observe a positive and statistically significant relationship between racial fractionalization and private school enrollment that is 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—significant given a baseline private enrollment rate of 10 percent.

Our estimated positive coefficient on expenditures per student in columns (1) and (3) is somewhat counter-intuitive if we believe expenditure is linked to public school quality (admittedly a tenuous assumption). However, expenditures may in part be a response to past school performance, where poor or under-performing school districts 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 2SLS estimates (columns (2) and (4)): here, districts with higher spending on average have lower rates of private schooling.

In our 2SLS estimates, both point estimates of the coefficient on inequality increase in size, and in the case of the Gini coefficient becomes statistically significant. In the latter case, our estimate of 0.055 implies that districts with one standard deviation above average growth in the Gini coefficient (5.0) are estimated to have private schooling rates that are 0.28 percentage points higher on average. While modest in size, it falls short of the effect implied by our coefficient on race fractionalization. Here we find that a standard deviation higher 1970 – 2000 growth in racial fractionalization is associated with a 0.43 to 0.45 point higher private schooling rate—an effect nearly twice as large.

Finally, in columns (5) and (6) of Table 7 we interact our income inequality measure with the quartile measures of district sorting opportunities (as measured by districts per student within a 25-mile radius). One would expect the relationship between inequality and private schooling to be moderated in districts were greater sorting opportunities exist. In the case of the Gini coefficient (column (6)), we do find that the point estimate on income inequality is largest among districts in the lowest quartile of district choice (0.082), but the difference is only statistically significant between the 1st and 2nd quartile. A similar pattern exists for the mean to median measure, but the point estimates do not statistically differ from zero.

5. Conclusion

As income inequality has risen in the U.S. and its population has grown more racially and ethnically diverse, scholars have begun to ponder whether this growing heterogeneity will alter the extent to which governments provide basic public services and a social safety net. Recent theoretical and empirical work suggests that public goods provision and the generosity of welfare benefits are lower in more racially and ethnically diverse jurisdictions. Models specific to public education suggest a similar outcome. With respect to income, growing income inequality may encourage a battle of the "ends against the middle," where high income families opt out of public schooling into the private sector and lower income groups choose lower taxes and greater private consumption over investments in public education. As a result, forces at the ends of the income distribution may reduce support for public schools in economically diverse populations.

In contrast, growing income inequality may have unanticipated effects on local public good provision. In a simple voting model like that of Meltzer and Richard, growing wage inequality at the top of the distribution reduces the tax price of public goods to the median voter, thereby encouraging greater spending on government services. We examined the impact of growing income inequality on local support for public schools using panel data for over 10,000 school districts over the 1970 – 2000 period. In contrast to other 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 in local school districts, so too have local dollars into elementary and secondary education. Our results indicate that 12 percent of the growth in local per-student revenues over the past 30 years can be explained by a decline in the tax share facing the median voter, a consequence of rising income inequality concentrated in top of the distribution.

The Meltzer-Richard hypothesis has been extensively tested in the past, in a number of settings, with conflicting results. The strength of the results in this paper may be driven by a number of factors. First, much of the previous work has used national or state level data while our analysis focuses on local school districts. Common wisdom in the public finance literature suggests that the collective choice process in local government is much more likely to approximate the assumptions of the median voter model assumptions (Fischel 2001; Mueller 2003). Second, we examine changes in spending in more than 10,000 school districts during a period of rapidly changing income inequality, giving us tremendous statistical power. Third, the panel nature of our data allows us to limit our exposure to potentially contaminating omitted variables bias.

Given the important redistributive nature of education, our results suggest that some of the potentially negative consequences of rising social inequality may have been counteracted by local government's 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 an important topic for future discussion.

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.[28] 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. 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

[28] Boundary files can be downloaded from http://www.census.gov/geo/www/cob/ bdy_files.html (last accessed 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.

a random variable z, the cumulative distribution function for the three-parameter Dagum distribution is as follows, for $z \geq 0$ and $(a, b, p) > 0$:

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

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

	Means					Full sample	
	1970/ 72	1980 /82	1990 /92	2000 /02	70-00 Change	Mean	Standard Deviation
Real local revenues per student	2385	2400	3323	3776	1486	2985	1957
Real state and federal aid per student	1824	2608	3665	4909	3077	3269	1798
Real current expenditures per pupil	3476	4424	6166	7494	3841	5501	2216
Percent of revenues from local sources	54.5	45.9	45.2	42.3	-11.0	47.0	19.3
Percent of enrollment in private school	10.5	8.8	10.2	10.4	1.1	10.0	6.8
Real median family income (x1000)	47.3	49.8	51.7	55.6	8.7	51.1	15.5
Percent of households below poverty	13.7	12.4	13.5	12.4	-1.3	13.0	8.0
Mean to median ratio of family income	1.13	1.14	1.19	1.23	0.10	1.17	0.11
Gini coefficient of family income (x 100)	33.8	35.1	37.3	38.8	5.0	36.3	5.5
Theil index of family income (x 100)	20.1	21.7	25.1	28.0	7.8	23.8	9.3
Log(95/5) ratio in family income	2.37	2.48	2.64	2.67	0.29	2.54	0.45
Ln(95/50) ratio in family income	0.88	0.91	0.99	1.06	0.18	0.96	0.18
Ln(50/5) ratio in family income	1.49	1.57	1.65	1.61	0.11	1.58	0.31
Mean to median housing value	1.13	--	1.15	1.16	0.03	1.15	0.12
Gini coefficient of housing value	28.0	--	28.7	28.3	0.02	28.3	7.2
Percent college graduates or higher	10.6	13.8	21.0	23.7	13.0	17.3	11.2
Percent school aged (5 – 17)	26.7	34.9	26.4	19.3	-7.6	26.5	6.7
Percent aged 65 and older	9.5	11.4	12.2	12.1	2.7	11.3	4.3
Percent of homes owner occupied	65.9	66.9	65.7	67.3	0.3	66.4	14.6
Percent of population nonwhite	15.7	19.8	23.9	30.9	15.3	22.7	22.2
Percent of population in urbanized areas	71.3	57.7	72.8	78.3	11.0	70.4	37.4
Index of race fractionalization (x 100)	16.5	22.1	26.2	32.9	17.5	24.5	19.9

Notes: authors' calculations using a balanced panel of elementary and unified school districts (N=10,359 in each year). District observations are weighted by public K-12 enrollment. Districts in Alaska, Hawaii, and D.C. are excluded. All monetary values measured in 2002 dollars.

Table 2: Income Inequality in School Districts and Metropolitan Areas – 1970 to 2000

A. Change in income inequality within school districts, 1970 to 2000 (weighted by district size)

Percent change in income inequality:	Mean/ Median Ratio	Gini Coefficient	Theil Index
Mean	8.96	15.60	42.50
10th percentile	0.42	-2.30	-4.20
25th percentile	4.22	6.25	15.86
Median	8.09	15.69	39.17
75th percentile	12.23	24.93	63.77
90th percentile	18.95	32.66	89.03
Mean income inequality in 1970	1.13	33.8	20.1
Mean change in income inequality, 1970-2000	0.10	5.0	7.8
Standard deviation of change in inequality, 1970-2000	0.11	4.6	10.0

B. Income inequality within metropolitan areas (weighted by MSA enrollment)

	1970	1980	1990	2000	% Change
Mean/median ratio	1.14	1.15	1.20	1.25	10.1
Gini coefficient (x 100)	34.7	36.1	38.5	40.6	17.1
Theil index (x 100)	20.8	22.6	26.2	30.1	44.3
% within school districts	92.1	92.1	89.1	88.7	-3.7
% between school districts	7.9	7.9	10.9	11.3	43.6
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=10,358 in each year; 331 MSAs in Panel B, except 1990 where 328 MSAs are included). Notes: Panel A uses same sample selection criteria as Table 1; Panel B makes use of the full (unbalanced) panel of school districts.

Table 3: OLS Estimates of Real Local Revenues per Pupil, Real Expenditures per Pupil, and Real State Revenues per Pupil

Dependent Variable:	Local revenues per pupil		Expenditures per pupil	State revenues per pupil
	(1)	(2)	(3)	(4)
Real median family income, in thousands	28.51***	33.51***	-10.63**	-42.15***
	(5.34)	(4.70)	(3.95)	(3.536)
Ratio of mean to median family income		1787.9***	1198.10***	-702.8***
		(449.7)	(299.7)	(213.0)
Percent of population below poverty line	12.98***	3.46	7.65*	18.04***
	(3.68)	(3.61)	(3.26)	(3.77)
Real state and federal revenues per pupil	-0.372***	-0.366***		
	(0.022)	(0.021)		
Percent college graduates or higher	45.80***	35.98***	21.48***	-22.66***
	(5.99)	(5.01)	(3.91)	(3.53)
Percent school aged (5-17)	-57.64***	-59.12***	-58.63***	0.618
	(5.05)	(4.97)	(4.08)	(4.391)
Percent aged 65 and older	15.77*	15.31*	-28.33***	-65.59***
	(7.14)	(7.04)	(6.50)	(6.00)
Percent of housing units owner-occupied	-10.47*	-12.21*	6.89*	16.99***
	(5.24)	(5.25)	(2.80)	(2.74)
Percent nonwhite	2.13	0.743	11.18***	-0.927
	(2.77)	(2.694)	(2.14)	(3.352)
Index of race fractionalization	-234.8	15.28	-294.6	-24.39
	(400.9)	(400.1)	(197.0)	(343.8)
Percent living in urbanized area	-2.68***	-2.17***	-1.36***	0.796*
	(0.43)	(0.41)	(0.31)	(0.320)
Observations	41,436	41,422	41,422	41,422
R-squared	0.908	0.908	0.952	0.891

Notes: Robust standard errors in parentheses (*** p<0.001, ** p<0.01, * p<0.05). All models include school district fixed-effects plus state x year fixed-effects. In models (3) and (4), we also include interactions between the within-state quartiles in median family income in 1970 times a dummy for whether the state has a court-ordered education finance reform in effect that year.

Table 4: OLS and 2SLS Estimates Real Local Revenues per Pupil Equation

	(1)	(2)	(3)
Real median family income, in thousands	33.51***	37.647***	41.96***
	(4.70)	(5.601)	(6.58)
Ratio of mean to median family income	1787.9***	1825.8***	3271.0***
	(449.7)	(449.86)	(603.1)
Percent of population below poverty line	3.46	3.26	-4.46
	(3.61)	(3.67)	(4.22)
Real state and federal revenues per pupil	-0.366***	-0.301***	-0.293***
	(0.021)	(0.057)	(0.057)
Percent college graduates or higher	35.98***	38.10***	30.32***
	(5.01)	(5.28)	(6.56)
Percent school aged (5-17)	-59.12***	-59.10***	-60.23***
	(4.97)	(4.98)	(4.97)
Percent aged 65 and older	15.31*	21.44**	21.44**
	(7.04)	(8.22)	(8.26)
Percent of housing units owner-occupied	-12.21*	-13.92*	-15.42**
	(5.25)	(5.48)	(5.84)
Percent nonwhite	0.743	0.329	-0.791
	(2.694)	(2.784)	(2.704)
Index of race fractionalization	15.28	55.71	261.0
	(400.1)	(405.0)	(395.1)
Percent living in urbanized area	-2.17***	-2.20***	-1.78***
	(0.41)	(0.41)	(0.41)
Instrument for:			
State and federal revenues	No	Yes	Yes
Mean/median income ratio	No	No	Yes
1st stage F-test: Instruments=0			
Real state+federal revenues		74.8	61.8
Mean to median family income			115.7
P-value, test of over-identifying restrictions (DOF of test)		0.163	0.106
		(2)	(2)
Observations	41,422	41,422	41,422
R-squared	0.909		

Notes: Robust standard errors in parentheses (*** p<0.001, ** p<0.01, * p<0.05). All models include school district fixed-effects plus state x year fixed-effects. In models (3) and (4), the instruments for real state + federal revenues are interactions between the within-state quartiles (2nd through 4th) in median family income times a dummy for whether the state has a court-ordered education finance reform in effect that year. In model (4) the instrument for the mean/median ratio is family income is the skewness in within-district family income.

Table 5: OLS and 2SLS Estimates of Real Local Revenues per Pupil Equation
with Alternate Measures of Income Inequality

	(1)	(2)	(3)	(4)	(5)	(6)
Income inequality measure:	Gini	Gini	Theil	Theil	Ln(95/5)	Ln(95/50) and ln(50/5)
Income inequality	34.43***	137.8***	17.92***	37.08***	28.67	
	(9.19)	(22.73)	(5.43)	(7.24)	(104.6)	
Ln(95/50) ratio of family income						904.2***
						(223.6)
Ln(50/5) ratio of family income						-213.0*
						(99.60)
Real median family income, in thousands	31.51***	46.40***	31.39***	38.63***		
	(4.90)	(7.00)	(4.87)	(6.32)		
Instrument for:						
State and federal revenues	No	Yes	No	Yes	No	No
Income inequality measure	No	Yes	No	Yes	No	No
1st stage F-test: Instruments=0						
Real state+federal revenues		98.9		83.3		
Mean to median family income		61.8		61.8		
P-value, test of over-identifying restrictions (DOF of test)		0.244		0.010		
		(2)		(2)		
Observations	41,422	41,422	41,422	41,422		41,422
R-squared	0.908		0.908		0.908	0.908

Notes: Robust standard errors in parentheses (*** p<0.001, ** p<0.01, * p<0.05). All models include school district fixed-effects, state x year fixed-effects, plus the demographic covariates used in Table 3. In models (2) and (4), the instruments for real state + federal revenues are interactions between the within-state quartiles (2nd through 4th) in median family income times a dummy for whether the state has a court-ordered education finance reform in effect that year. Th instrument for the inequality measures is the skewness in within-district family income.

Table 6: Extensions and Robustness Checks.
OLS and 2SLS Estimates of Real Local Revenues per Pupil Equation

			Exclude CA	Exclude CA	
	(1)	(2)	(3)	(4)	(5)
Real median family income, x 1000)	25.47***	38.73***	34.50***	39.0***	32.19***
	(7.20)	(8.32)	(5.20)	(6.75)	(4.30)
Mean/median family income (x100)			1557.9**	3103.8***	2257***
			(510.6)	(671.5)	(356.2)
Mean/median of housing value (x100)	794.8*	879.9*			
	(386.7)	(402.9)			
Mean/median family income (x 100)					-2142.3***
x 2nd quartile of districts per student					(631.8)
x 3rd quartile of districts per student					-1815.1***
					(384.0)
x 4th quartile of districts per student					-1350.1***
					(374.7)
Instrument for:					
State and federal revenues	No	Yes	No	Yes	No
Mean/median income ratio	No	No	No	Yes	No
1st stage F-test: Instruments=0					
Real state+federal revenues		53.0		70.0	
Mean to median family income				91.0	
P-value, test of over-identifying restrictions (DOF of test)		0.059		0.244	
		(2)		(2)	
Observations	29,391	29,062	39,505	39,505	40,710
R-squared	0.854		0.880		0.911

Notes: Robust standard errors in parentheses (*** p<0.001, ** p<0.01, * p<0.05). All models include school district fixed-effects, state x year effects, and the demographic covariates used in Table 3. Columns (1) – (2) replace the measure of income inequality with the within-district mean/median ratio in self-reported housing values. Columns (3) and (4) estimate models excluding data for California.

Table 7: Effect of Income Inequality on Private Schooling

Inequality measure:	Mean/ Median (1)	Mean/ Median (2)	Gini (3)	Gini (4)	Mean/ Median (5)	Gini (6)
Real median family	0.109***	0.109***	0.110***	0.108***	0.107***	0.105***
income, in thousands	(0.012)	(0.017)	(0.012)	(0.017)	(0.016)	(0.015)
Inequality measure	0.053	1.866	0.013	0.055*	2.114	0.082**
	(0.600)	(0.964)	(0.017)	(0.024)	(1.413)	(0.031)
Inequality measure x 2nd quartile of					-3.932	-0.119*
districts/student					(2.457)	(0.057)
Inequality measure x 3rd quartile of					2.034	-0.008
districts/student					(1.747)	(0.037)
Inequality measure x 4th quartile of					0.526	-0.010
districts/student					(1.601)	(0.035)
Percent of population	-0.053***	-0.032	-0.058***	-0.044*	-0.030	-0.041*
below poverty line	(0.013)	(0.017)	(0.014)	(0.017)	(0.017)	(0.017)
Real current expenditures per pupil	0.195*	-1.580***	0.193*	-1.543***	-1.575***	-1.489***
(in thousands)	(0.086)	(0.448)	(0.085)	(0.438)	(0.436)	(0.420)
Percent college graduates	0.016	0.036	0.014	0.034	0.037*	0.032
or higher	(0.014)	(0.018)	(0.014)	(0.018)	(0.018)	(0.017)
Percent school aged	-0.049**	-0.139***	-0.049**	-0.135***	-0.138***	-0.130***
(5-17)	(0.015)	(0.028)	(0.015)	(0.027)	(0.028)	(0.026)
Percent aged 65 and older	0.059*	0.033	0.057*	0.027	0.028	0.021
	(0.023)	(0.030)	(0.024)	(0.031)	(0.029)	(0.029)
Percent of housing units	-0.010	-0.000	-0.010	0.001	0.003	0.006
owner-occupied	(0.011)	(0.014)	(0.012)	(0.015)	(0.014)	(0.013)
Percent nonwhite	-0.054***	-0.038***	-0.054***	-0.039***	-0.043***	-0.048***
	(0.011)	(0.011)	(0.011)	(0.011)	(0.013)	(0.013)
Index of race	3.168**	2.891**	3.186**	2.717*	3.442**	3.499**
Fractionalization	(1.103)	(1.055)	(1.112)	(1.076)	(1.125)	(1.155)
Percent living in	0.001	-0.001	0.001	-0.001	-0.001	-0.001
urbanized area	(0.001)	(0.002)	(0.001)	(0.002)	(0.002)	(0.002)
Instrument for expenditures	No	Yes	No	Yes	Yes	Yes
1st stage F-test: Instruments=0						
P-value test of overid. restrictions						
Observations	38,052	37,867	38,052	37,867	37,191	37,192
R-squared	0.851		0.851			

Notes: Robust standard errors in parentheses (*** p<0.001, ** p<0.01, * p<0.05). Dependent variable measured as the percent of K-12 enrollment in private schools. Regressions weighted using total K-12 enrollment (public + private). All models include school district fixed-effects and state x year effects.

Appendix Table 1: First-Stage Estimates for 2SLS Models from Tables 4, 5, 6 and 7

First-stage estimates corresponding to:	Table 4 Col. (2)	Table 4, Column (3)		Table 5, Column (2)	Table 5, Column (4)
Endogenous covariate:	S+F	S+F	Mean/ Median	Gini	Theil
Instruments:	(1)	(2)	(3)	(4)	(5)
Court-ordered finance reform x within-state quartile of mean district family income					
x quartile 1	1546.8***	1495.3***	0.0183**	0.164	1.815***
	(114.8)	(114.8)	(0.0047)	(0.138)	(0.356)
x quartile 2	573.7***	526.7***	0.0130**	-0.215	0.973**
	(74.4)	(74.3)	(0.0043)	(0.124)	(0.333)
x quartile 3	508.4***	477.0***	0.0125*	-0.188	0.839
	(120.8)	(126.0)	(0.0062)	(0.161)	(0.463)
Skewness of family income		-0.010**	5.1E-6***	1.2E-4***	4.4E-4***
		(0.0022)	(2.4E-7)	(6.2E-6)	(2.5E-5)
Observations	41,422	41,422	41,422	41,422	41,422
1sr stage F-test: Instruments are jointly zero	74.8	61.8	115.7	98.9	83.3
R^2	0.856	0.857	0.743	0.833	0.769

First-stage estimates corresponding to:	Table 6 Col. (2)	Table 6, Column (4)		Table 7 Col. (2)
Endogenous covariate:	S+F	S+F	Mean/ Median	S+F
Instruments:	(6)	(7)	(8)	(9)
Court-ordered finance reform x within-state quartile of mean district family income				
x quartile 1	1685.4***	1659.2***	0.014**	618.8***
	(137.1)	(120.2)	(0.0051)	(52.8)
x quartile 2	496.2***	702.2***	0.011*	431.2***
	(102.6)	(84.2)	(0.046)	(63.0)
x quartile 3	442.9***	326.5***	0.006	431.5
	(126.6)	(70.3)	(0.0044)	(138.4)
Skewness of family income		-0.013***	5.0E-6***	
		(0.0024)	(2.6E-7)	
Observations	29,062	39,505	39,505	37,867
1sr stage F-test: Instruments are jointly zero	53.0	70.0	91.0	50.0
R^2	0.880	0.857	0.713	0.917

Notes: Robust standard errors in parentheses (*** p<0.001, ** p<0.01, * p<0.05). S+F represents real state + federal revenues per pupil. Mean/media in columns (3), (4), (5) and (8) and the within-district family income mean/median ratio. In column (6), mean/median is within-district mean/median housing price ratio. The first-stage estimates for S+F from Table 5 columns (2) and (4) are identical to those from Table 4, column (3). All models include school district fixed-effects, plus state x year effects. Demographic covariates shown in Table 3 are also included in these regression models. Columns (1) – (5) only use district observations from 1970, 1990, and 2000 (the only years for which we have Census housing values). Column (6) uses only observations from 1970, 1990 and 2000. Columns (7) and (8) use exclude data from California districts from 1970 – 2000. Column (9) includes the sample from the private schooling results in Table 7.