Inequality and Education Funding: Theory and Evidence from the U.S. School Districts*

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Abstract

We investigate the relationship between inequality and public education funding in a model of probabilistic voting where the private option is available and voting participation differs across income groups. A change in inequality can have opposite effects at different income levels: higher inequality decreases public spending per student and increases enrollment in public schools in poor economies, while the opposite holds in the rich ones. A change in the tax base can also have non-monotonic effects. These novel theoretical predictions, with support in U.S. school district-level data, reconcile previous contradictory results in the political economy literature on redistribution and inequality.

JEL Codes: D72, H42, I21, I22.
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1 Introduction

This paper investigates the relationship between income inequality and the determination of education funding. The worldwide surge in income inequality in the last decades, the importance of human capital for individual and national prosperity, as well as the rising pressures on public budgets, make this a critical issue. However, the literature studying the effects of inequality on public education spending (and, more generally, on public goods provision and income redistribution) has yet to reach a consensus.\footnote{For example, Lindert (1996) finds a negative relationship between inequality and public education spending in a sample of OECD countries. Using U.S. state level data, de la Croix and Doepke (2009) find that higher inequality is positively associated with public spending per student and negatively correlated with public spending per capita. Corcoran and Evans (2009) find that rising inequality within U.S. school districts is associated with higher local revenues per pupil.}

Given these inconclusive results, in this paper we investigate the issue further. In the theoretical part, we provide a model that reconciles the findings of the previous literature, by showing that, in fact, the relationship of inequality and public school funding can be non-monotonic, i.e. negative at low levels of income and positive otherwise. We consider a political economy model of public education provision with a private schooling option and endogenous fertility decisions. We allow household income heterogeneity to be consistent with the skewness of empirical income distributions, where the median is lower than the mean income.

If education is a normal good, the desired education quality increases with income. Tax financed uniform public education quality is insufficient for the rich parents who choose to send their children to a private school. Ceteris paribus, the higher the public school quality, the lower the private enrollment share. The availability of a private choice generates an endogenous income threshold that separates public and private school users. In our model, fertility is connected to the school choice, with the private school users choosing a lower fertility rate than the households that opt for public schooling. For transparency, the fertility is constant within the two groups.

The equilibrium public spending arises as the politically mediated balance between the conflicting interests of these two groups. On the one hand, those opting for private schooling want to minimize the tax burden. On the other hand, those who choose public schooling, want to ensure adequate spending per student.

We study how the political balance and thus the equilibrium education spending and enrollment respond in two counterfactual experiments: a) an increase in the tax base keeping income dispersion constant and b) a mean preserving spread of the income distribution. Our analysis yields two main findings.

First, generalizing results in the previous literature, we show that the relationship between inequality and public spending per student can be non-monotonic, depending on the average income level in the economy. A mean preserving spread decreases public
spending per student but increases tax rates (spending per capita) and public school enrollments in low income economies, while it has opposite effects at high income levels. Furthermore, a marginal increase in the tax base, holding income dispersion constant, can also have non-monotonic effects. Second, we show that both the tax base and inequality effects on redistribution depend critically, not only on the level of the average income per capita in the economy, but also on the parental preferences for quality versus quantity of children.

In order to get more intuition on the mechanisms in place, consider the case of low fertility or a high tax base. In this case, the spending per student in public schools is high and so is the quality of public education. Therefore, only the very rich households use private schools. A mean preserving spread increases support for private education as part of the middle class is replaced with high income households that prefer this option (the shape effect). This replacement also increases spending per student in public schools as resources are spread over fewer children. On the other hand, some middle income households are replaced by poorer families (again, the shape effect). However, since both household types choose public schooling and hence, have identical fertility rates in our model, this replacement neither generates extra support for the public option nor does it decrease its quality. Therefore, marginal households strictly prefer public education, i.e. the indifference income threshold moves to the right of the income distribution (the threshold effect). Nonetheless, since the income threshold is located relatively far in the right tail, its effect is dominated in magnitude by the shape effect. Thus, in this case, a mean preserving spread increases the support for private education and lowers taxes. While both public school enrollment and the tax rate decrease, more resources are available per student, and thus, the quality of public schooling increases.

In contrast, when fertility rates are high or the tax base is low, public schools are of low quality, so both rich and some middle income households use private schools. A mean preserving spread increases support for public education as it replaces middle income households, that opt for private schools, with low income families, which prefer the free alternative (the shape effect). The latter group also chooses high fertility rates, which further reduces (expected) spending per student. While this generates an endogenous shift of the marginal households into private schools (the threshold effect), the shape effect dominates in equilibrium, leading to higher enrollment in public schools and higher taxes, despite lower spending per student.

As a benchmark, we focus on probabilistic voting with households that have uniform political power. Asymmetric distribution of political power is typically associated with authoritarian regimes or partially democratic countries. However, it can also occur in well established democracies if, for example, voter turnout varies systematically with demographic characteristics. Indeed, the literature on political participation in the United States documents large turnout differences across income and age groups in national, state
as well as local elections. While previous evidence applies to political participation in general, we show that education related votes are subject to similar disparities. Using data on school budget votes across school districts in the state of New York, we find a positive and significant correlation between socioeconomic indicators and involvement in local politics related to public education provision. Motivated by these findings, we then extend the model to include an income based index of political power and study its properties.\footnote{Other papers that study asymmetrically distributed political power include Gans and Smart (1996), Bénabou (2000), de la Croix and Doepke (2009).}

In the empirical part, we test predictions arising from the theory using 2000 U.S. school district level data. We investigate the effects of the tax base and the income inequality on three schooling measures: the local public spending per student, the share of private enrollment and the local public spending per capita. The per capita spending can be interpreted as a measure of redistribution through education spending. We show that the aggregate relationship between spending and inequality reveals significant heterogeneities once the data is split into high and low income districts. Across subsamples and estimation methods, our results lend strong support to predictions derived from the theoretical model with respect to enrollment and spending per student in public schools. Allowing for asymmetries in the distribution of political power, in line with U.S. stylized facts, helps explain the behavior of per capita spending in both poor and rich school districts.

We tackle endogenous selection into districts in two ways. First, we use lagged data to generate instrumental variables for the school district demographic, socioeconomic and policy characteristics as well as to capture exogenous shifts in inequality at the level of the school district, as implied by our theoretical model. Second, we provide evidence that selection is less important across larger geographical units. We then use Metropolitan Statistical Area (MSA) definitions, circumscribing locally integrated economies and also a large part of the relocation efforts, to aggregate school districts into units which are less likely to be affected by selection biases.

Finally, given our focus on local education politics, constraints on the local governments’ ability to tax and redistribute are a matter of concern. Therefore, we use data on the severity of local tax and expenditure limitations (TELs) across states in order to test the robustness of our results across different samples.

\subsection{Connections to the literature}

There is a large body of work studying the effects of inequality on public goods provision and income redistribution. However, as mentioned above, on both the theoretical and the empirical front, the literature has often reached inconsistent, even contradic-
tory results. While some political economy papers argue that higher inequality leads to more redistribution through higher taxation (Meltzer and Richard (1981), Persson and Tabellini (1994), Bénabou (1997)), others find that more unequal or more heterogenous societies spend less on public goods (Soares (1998), de la Croix and Doepke (2009)). Glomm (2004) finds that the relationship between inequality and the amount of redistribution through public education services depends on the elasticity of substitution between consumption and the quality of education in the parent’s utility. He finds that for empirically relevant value of this parameter, higher inequality generates less redistribution.

On the empirical side, a number of papers have found that support for redistribution and public goods provision is weaker in more unequal or more heterogenous societies (Goldin and Katz (1997), Alesina et al. (1999, 2001), Luttmer (2001)). Perotti (1996) finds no relationship between inequality and redistribution in democracies. Using data from the U.S. General Social Survey, Lind (2007) finds that inequality between different groups reduces redistribution, while within group inequality increases it. A more recent paper by Bousman et al. (2010) finds that rising inequality in cities and districts is associated with higher local revenue collection and expenditures.

Our paper contributes to the debate by providing more general theoretical results borne out in the data. Bénabou (1997, 2000) and Lee and Roemer (1998) focus on capital market imperfections to show that non-monotonic responses of redistribution to inequality are possible. Fernandez and Levy (2008) also find a non-monotonic effect of increased diversity in a model with income and preference heterogeneity. Complementary to these studies, we obtain a non-monotonic effect of inequality on redistribution at different levels of the average income per capita stemming from endogenous fertility and education choices. Also, in these papers, redistribution occurs through progressive taxation (Bénabou (2000)) or the provision of universal public education (Lee and Roemer (1998)). In the latter case, private and public investments in education are complements, but only the rich households top up.

In contrast, we focus on public education funding when the rich can opt out of the public system. The theoretical endogenous income threshold that separates the education choices has a clear empirical counterpart. This makes the model’s conclusions testable in more dimensions: spending per student, public school enrollment, as well as per capita spending.

While our analysis builds on de la Croix and Doepke (2009), there are several important differences. First, a more flexible parametrization of the income distribution allows us to recover their results as a particular case. Second, by including the tax base in the model, we are able to uncover non-trivial insights related to the non-monotonic response of redistribution to an increase in inequality as a function of the economy mean income level. Moreover, these new predictions are supported by the data. Thus, we provide a simple mechanism that helps reconcile the previous contradictory results in the political
economy literature on redistribution and inequality. Another distinct theoretical contribution is the introduction of a parsimonious and tractable mapping between income and political influence. In this extension we show that a unique equilibrium exists for empirically relevant values of the political power parameter. Finally, in our model there are no tax deductions for private education spending and revenues are raised through a consumption tax. On the empirical side, they use U.S. state level data to document correlations between inequality and education spending whereas we focus on a richer dataset at school district level.

In contrast to models that study how sorting across communities affects public goods provision and inequality\(^3\), in this paper we study how education funding responds to exogenous changes in inequality. In our model fertility decisions are made well in advance of the political process that establishes the quality of public education. Since deciding on the number of children also implies selecting their quality, whenever school choices (private or public) are conditioned by residential choices, location and fertility are jointly pinned down before public spending is chosen. Thus, in our analysis, households’ commitment regarding fertility and school choices imply the income distribution in each community already incorporates any relocation incentives.

The remainder of the paper is structured as follows. Section 2 presents the model. Section 3 defines the equilibrium and derives the main analytical results. Section 4 documents significant participation differences in local politics related to public education provision and extends the benchmark model by incorporating political power. Section 5 is devoted to the empirical evidence. Section 6 concludes. Proofs are relegated to Appendix A, while Appendix B describes the data used in the empirical part.

2 The model

The economy is populated by a large number of households, which are heterogenous in income. The mass of households is normalized to one. Each household consists of an adult and a number of children. Children are educated either in public schools, which are financed by a consumption tax, or in private schools, financed by parental spending. Household income is distributed according to a Pareto distribution, with p.d.f. \( f \) and c.d.f. \( F \), with parameters \( y_l > 0 \) and \( \alpha > 2 \), and support \( y \in [y_l; \infty) \).\(^4\) The mean and


\(^4\)The p.d.f. is given by \( f(y) = \alpha y_l^{\alpha} y^{\alpha - 1} \), for \( y > y_l \) and zero otherwise. The c.d.f. is \( F(y) = 1 - (y_l/y)^{\alpha} \). The Pareto distribution is used for tractability reasons (see also Lee and Roemer (1998)). Other distributions used in the literature yield similar qualitative results. Similar qualitative results obtain using a log-normal income distribution.
standard deviation of the income distribution are given by:

\[ \mu = \frac{\alpha}{\alpha - 1} y_l \text{ and } \sigma = \frac{y_l}{\alpha - 1} \frac{1}{\alpha - 2}. \] (1)

Adults derive utility from net of tax consumption \( c \), the number of children \( n \) and the quality of their education \( E \), which can be private or public. Private education has a unit price. Let \( q \) denote the quality of public schools. Households can opt out of publicly provided education and send their children to a private school of quality \( e^r \). The preferences are given by:

\[ u(c, n, E) = \ln(c) + \gamma [\ln(n) + \theta \ln(E)], \] (2)

where \( E = q, e^r \), \( \gamma > 0 \) and \( \theta \in (0, 1) \).

Besides providing tractability, the assumption of logarithmic utility is consistent with the empirical evidence, which suggests that income and substitution elasticities of education spending have similar magnitudes.

The government taxes the consumption of all households at the constant rate \( \tau \). Tax revenues are used to finance public education of uniform quality for all children. For simplicity, we assume that quality of schooling is equal to the spending per student. The public policy is determined through a probabilistic voting mechanism described below. Besides being tractable, this mechanism formalizes in a general way preference aggregation in local education politics.

Private choices on fertility and education are made before voting on the quality of public education takes place. Agents have perfect foresight regarding the outcome of the voting process. Thus, in equilibrium, the expected spending per student in public education equals the level chosen by voting.

This timing reflects the sizeable differences in the relative costs and time horizons of the decisions involved. While public education spending is usually decided through yearly budget votes, fertility and child rearing decisions cannot be easily adjusted at this frequency and depend largely on "pre-determined" characteristics, such as income, education level, race, religion, etc. A similar argument applies to the choice between public and private schooling, which in the U.S. is tightly connected to residential choice and therefore can entail substantial switching costs.

Furthermore, notice that under perfect foresight, a quantity-quality trade-off maps

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5 As we explain later, private decisions are taken before public policy is implemented. This implies that households have to form expectations over policy variables, such as the quality of public education, \( q \). Since we assume perfect foresight, and in order to avoid clutter, we use the same notation for expected and realized values of a variable.

6 See Gradstein et al. (2005), pg. 50-51 for a discussion.

7 de la Croix and Doepke (2009) also conclude that in countries where the educational and residential segregation are correlated, private decisions generate strong lock-in effects. Thus, in the case of the U.S. school districts, the focus of our empirical analysis, having households make their decisions before the policy is chosen seems the most appropriate assumption.
fertility decisions into consistent school choices. Therefore, even if households decide on private vs. public education after policies are set, as long as fertility decisions occur before the vote on public education quality, the same (sub-game perfect) equilibrium will obtain as under the original timing.\textsuperscript{8}

2.1 Household’s problem

Household consumption can be interpreted in this framework as housing services and is subject to a constant tax rate $\tau$. As we show later, this assumption implies that the decisions regarding the quantity and quality of children are not affected by taxation. Furthermore, in this case endogenous fertility generates a constant tax base in equilibrium, which does not depend on the aggregate enrollment in public education.\textsuperscript{9}

Rearing children involves a time cost. Denote by $2 = (0, 1)$ the fraction of the parent’s time spent raising a child, and with $U^p$ and $U^r$ the utility of households whose children are educated in the public and private schools, respectively. Given the expected quality of publicly provided education $q$ and the tax rate $\tau$, a household with income $y$ that chooses public education solves the following problem:

$$\max_{c \geq 0, n \geq 0} U^p(c, n, q) = \ln(c) + \gamma \ln(n) + \gamma \theta \ln(q),$$

s.t. $c(1 + \tau) \leq y(1 - \phi n).$ (3)

The solution of problem (3) is $n^p = \gamma / [\phi(1 + \gamma)].$

On the other hand, a household choosing private education solves:

$$\max_{c \geq 0, n \geq 0, e \geq 0} U^r(c, n, e) = \ln(c) + \gamma \ln(n) + \gamma \theta \ln(e),$$

s.t. $c(1 + \tau) + ne^r \leq y(1 - \phi n).$ (5)

The solutions to the problem (5) are $n^r = [\gamma(1 - \theta)] / [\phi(1 + \gamma)]$ and $e^r = \phi y / (1 - \theta)$. Comparing $n^p$ and $n^r$ we see that households that choose private schooling have a lower fertility than those sending the children to public schools. Also note that consumption for both household types is a constant share of income $c = y / ((1 + \gamma)(1 + \tau)).$

Substituting $n^p$ in (3) and $n^r$ and $e^r$ in (5) we obtain the indirect utilities of households that choose public and private schooling, respectively:

$$V^p(y, q) = \ln \left[ \frac{y}{(1 + \gamma)(1 + \tau)} \right] + \gamma \ln \left[ \frac{\gamma}{\phi(1 + \gamma)} \right] + \gamma \theta \ln(q)$$

\textsuperscript{8}See Dottori and Shen (2009) for a related discussion.
\textsuperscript{9}de la Croix and Doepke (2009) obtain similar results assuming an income tax and tax deductibility of private education spending.
\[ V^r(y) = \ln \left[ \frac{y}{(1 + \gamma)(1 + \tau)} \right] + \gamma \ln \left[ \frac{\gamma(1 - \theta)}{\phi(1 + \gamma)} \right] + \gamma \theta \ln \left[ \frac{\phi \theta y}{1 - \theta} \right]. \tag{8} \]

A household will choose public education if and only if \( V^p(y, q) \geq V^r(y). \) This inequality is satisfied for households with income lower than a threshold \( \tilde{y}, \) given by:

\[ \tilde{y} = \frac{q}{\phi \theta \delta}, \text{ where } \delta = (1 - \theta)^{\frac{1}{\alpha}} \in (0, 1). \tag{9} \]

Households choose the school type taking the other households’ decisions as given. Denote by \( \Psi \) the fraction of households that choose public schooling. In equilibrium, the individual choices must be consistent with the aggregate outcome, that is, the share of households with income lower than the threshold \( \tilde{y} \) should be equal to \( \Psi. \) The consistency condition is:

\[ \Psi(q) = F(\tilde{y}(q)) = \int_{y_t}^{\tilde{y}(q)} f(y)dy = 1 - \left( \frac{y_t}{y(y)(q)} \right)^{\alpha}. \tag{10} \]

An alternative way to state this rationality condition is that the expected quality of public education equals the implemented level: \( E(q) = q. \)

Notice that \( \Psi(q) \) is not equal to the fraction of children that go to public schools since the model incorporates fertility decisions. Thus, the fraction of children in public schools is given by:

\[ N(q) = \frac{n^p \Psi(q)}{n^p \Psi(q) + n^r (1 - \Psi(q))}. \tag{11} \]

Substituting the expressions for \( n^p \) and \( n^r \) we obtain:

\[ N(q) = \frac{\Psi(q)}{(1 - \theta) + \theta \Psi(q)} > \Psi(q). \tag{12} \]

### 2.2 Government budget constraint

The government budget is balanced:

\[ \int_{y_t}^{\tilde{y}} qn^p f(y)dy = \tau \int_{y_t}^{\infty} \frac{y}{(1 + \gamma)(1 + \tau)} f(y)dy, \tag{13} \]

where the left-hand side is the total public education spending, and the right-hand side the collected tax revenues from both types of households (public and private school users, respectively). The right-hand side of (13) shows clearly that the fraction of income that is taxable is constant across income groups and is equal to \( 1/((1 + \gamma)(1 + \tau)). \) As a result, the total tax base is constant and does not depend on the fraction of households choosing private schooling. Integrating over the support of the distribution, the right-hand side of the government budget constraint becomes \( \tau \mu/((1 + \gamma)(1 + \tau)), \) where \( \mu = \int_{y_t}^{\infty} yf(y)dy \) is the average income and also the tax base. Using the expression for \( n^p \) in the left-
hand side, we can express the quality of public schooling as a function of the fraction of households that choose public schools, $\Psi$, and the tax rate, $\tau$:

$$q\Psi(q)\frac{\gamma}{\phi} = \frac{\tau \mu}{1 + \tau}. \quad (14)$$

### 2.3 Voting on public education funding

The public policies are determined through probabilistic voting. The voting problem is unidimensional, i.e. once the tax rate is chosen, the spending per student $q$ is determined from (14). Consider a set-up with two political parties, each proposing a program. Voters care about the education policy proposed but also about a second dimension of the electoral platform, called "ideology". The probability that an individual votes for a party thus depends on her ideological bias toward the party’s proposed platform. The results of the elections are a random event, each party having a probability of winning.

The ideological preferences are assumed to be orthogonal to those on public policy. Thus, the probability that a person votes for a certain party (and the party vote share) is a smooth function of the distance between the two platforms. This framework has a unique equilibrium in which both parties converge to the same platform (see Persson and Tabellini (2002)), which maximizes the following social welfare function:

$$W(\tau) = \int_{\tilde{y}}^{\bar{y}} U^P(y, n^p, q, \tau)p(y)f(y)dy + \int_{\tilde{y}}^{\infty} U^R(y, n^r, e^r, \tau)p(y)f(y)dy,$$

subject to the government budget constraint (14).

The first and second terms of the welfare function are the aggregate utilities of the households that choose public and private education, respectively. The term $p(y)$ captures the political power of the group. We first assume $p(y) = 1$, that is, all voters have the same political power. We relax this assumption in section 4.

Note that the income threshold $\tilde{y}$ is taken as given in the maximization, in keeping with the assumption that fertility and education choices are predetermined when the vote takes place. While making the analysis more tractable, this assumption is still consistent, in equilibrium, with perfect foresight: the expected and the actual shares of households that choose public schooling are equal.

Substituting the indirect utility functions, (7) and (8), in (15) and grouping terms, we get:

$$W(\tau) = \ln \left( \frac{1}{(1 + \gamma)(1 + \tau)} \right) + \gamma \ln \left( \frac{\gamma}{\phi(1 + \gamma)} \right) + \gamma \theta \ln(q(\tau)) \int_{\tilde{y}}^{\bar{y}} f(y)dy +$$

$$\int_{\tilde{y}}^{\infty} \left\{ \gamma \ln(1 - \theta) + \gamma \theta \ln \left( \frac{\phi \theta y}{1 - \theta} \right) \right\} f(y)dy.$$
Since only the first and the third term are functions of the policy variables, the welfare can be rewritten as

\[ W(\tau) = -\ln(1 + \tau) + \gamma \theta \Psi \ln(q(\tau)), \]  

(16)

where \( \Psi(y) \) is taken as given. Substituting \( q \) from (14) and taking the first order condition with respect to \( \tau \) yields:

\[ \tau = \gamma \theta \Psi. \]  

(17)

Everything else equal, the tax increases with the households’ concern for children as well as with the fraction of households using public education. In the next section we define the equilibrium and study its properties.

3 Equilibrium analysis

**Definition 1.** A politico-economic equilibrium is an income threshold \( \tilde{y} \), private allocations \((c^p, n^p)\) if \( y \leq \tilde{y} \), \((c', n', c')\) if \( y > \tilde{y} \), and a public policy \((q, \tau)\) such that:

(i) household’s decisions solve problems (3) or (5), given public policy \((q, \tau)\);

(ii) the government budget is balanced, i.e. it satisfies (14);

(iii) the tax rate \( \tau \) solves the social welfare maximization problem (15);

(iv) the consistency condition (10) is satisfied.

Next, we solve for the equilibrium threshold \( \tilde{y} \). To minimize clutter, we drop functional dependencies where possible. We use the expression of \( q \), (14), and \( \tau \), (17) in (9) to obtain:

\[ \tilde{y} = \frac{\mu}{\delta} \frac{1}{1 + \gamma \theta \Psi(y)}. \]  

(18)

Using the consistency condition, (10) yields the following expression in \( \tilde{y} \):

\[ \tilde{y} = \frac{\mu}{\delta} \frac{1}{1 + \gamma \theta \left[ 1 - \left( \frac{y_1}{\tilde{y}} \right)^\alpha \right]}. \]  

(19)

**Proposition 1.** There exists a unique and interior equilibrium income threshold \( \tilde{y}^* \in (y_1, \infty) \) that solves equation (19) (proof in the Appendix).

Note that the equilibrium threshold \( \tilde{y}^* \) is always interior because the support of the income distribution does not have an upper bound. When \( \tilde{y}^* \to \infty \), the fraction of students in public schools goes asymptotically to 1. Equilibrium uniqueness also owes to the endogenous fertility, which ensures that the tax base is independent of public education enrollment and thus the right hand side of equation (19) is decreasing in \( \Psi \).
Proposition 1 implies there is a unique equilibrium public spending per student:

$$q^* = \bar{y} \phi \theta \delta = \frac{\phi \theta \mu}{1 + \gamma \theta \left[1 - \left(\frac{y_l}{\bar{y}}\right)^{\alpha}\right]}.$$  \hfill (20)

We use equations (10) and (20) to express $\Psi^*$ as a function of $q^*$:

$$\Psi^* = \frac{1}{\gamma \theta} \left(\frac{\phi \theta \mu}{q^*} - 1\right).$$  \hfill (21)

Using (10) in (12), we obtain the equilibrium enrollment in public schools:

$$N^* = \frac{\Psi^*}{(1 - \theta) + \theta \Psi^*}, \text{ where } \Psi^* = 1 - \left(\frac{y_l}{\bar{y}}\right)^{\alpha}. \hfill (22)$$

In the following, we investigate how changes in the income distribution of a school district affect the main policy variables. We focus on two experiments: a) a change in the average income per capita $\mu$, keeping the standard deviation $\sigma$ constant and b) a mean preserving spread in the income distribution (change $\sigma$ while keeping $\mu$ constant).

### 3.1 A change in the mean income (tax base)

Now we analyze the effects of changing the mean income, $\mu$, on the equilibrium public spending per student $q^*$, the tax rate $\tau^*$, and enrollment in public schools $N^*$. Recall that in our model $\mu$ also represents the tax base.

Denote by $f(\mu, \sigma) = [y_l(\mu, \sigma)/\bar{y}(\mu, \sigma)]^{\alpha(\mu, \sigma)}$. The derivative of $N^*$ with respect to $\mu$ is:

$$\frac{\partial N^*}{\partial \mu} = \frac{1 - \theta}{[(1 - \theta) + \theta \Psi^*]^2} \frac{\partial \Psi^*}{\partial \mu} = \frac{1 - \theta}{[(1 - \theta) + \theta \Psi^*]^2} \phi \frac{q^* - \mu \frac{\partial q^*}{\partial \mu}}{\gamma (q^*)^2},$$

where

$$\frac{\partial q^*}{\partial \mu} = \frac{\phi \theta \left\{1 + \gamma \theta (1 - f) + \mu \gamma \theta \frac{\partial f(\mu, \sigma)}{\partial \mu}\right\}}{[1 + \gamma \theta (1 - f)]^2}. \hfill (24)$$

Using (17) and (21) we obtain the change in the equilibrium tax rate with respect to $\mu$:

$$\frac{\partial \tau^*}{\partial \mu} = \gamma \theta \frac{\partial \Psi^*}{\partial \mu}. \hfill (25)$$

Thus, $\text{sign}(\partial \tau^*/\partial \mu) = \text{sign}(\partial \Psi^*/\partial \mu) = \text{sign}(\partial N^*/\partial \mu)$. Studying the properties of the function $\partial N^*/\partial \mu$ yields the following results.
Proposition 2. Let $\gamma = \left[\frac{2(\delta e)}{(\delta e) - 1}\right] / \left\{\theta \left[1 - e^{-2}\right]\right\}$ and $\tau = \left[\frac{1}{\delta} - 1\right] / \left\{\theta \left[1 - (1/e)\right]\right\}$, where $e$ is the Euler’s constant.

1) If $\gamma \leq \gamma$, then $\partial N^*/\partial \mu > 0$ and $\partial \tau^*/\partial \mu > 0$;
2) If $\gamma > \gamma$, then $\partial N^*/\partial \mu < 0$ and $\partial \tau^*/\partial \mu < 0$;
3) If $\gamma \in (\gamma, \gamma)$, then there exist a unique $\tilde{\mu} \in (0, \infty)$ such that
   3.1) if $\mu \in (0, \tilde{\mu})$, then $\partial N^*/\partial \mu \leq 0$ and $\partial \tau^*/\partial \mu \leq 0$;
   3.2) if $\mu \in (\tilde{\mu}, \infty)$, then $\partial N^*/\partial \mu > 0$ and $\partial \tau^*/\partial \mu > 0$;

(Proof in the Appendix).

The next corollary establishes sufficient conditions under which the equilibrium spending per student $q^*$ varies positively with the mean income.

Corollary 1. 1) If $\gamma \geq \gamma$, then $\partial q^*/\partial \mu > 0$;
2) If $\gamma \in (\gamma, \gamma)$ there exists $\tilde{\mu} > \mu$ such that $\partial q^*/\partial \mu > 0$ on the interval $\mu \in (0, \tilde{\mu})$ (Proof in the Appendix).

As it is apparent from Proposition 2, the effects of an increase in the tax base depend on $\gamma$. Equilibrium fertility allocations $n^p$ and $n^r$ are increasing functions of $\gamma$, while private education spending $e^r$ does not depend on $\gamma$.\(^{10}\) We therefore interpret $\gamma$ as a relative weight of fertility in the parental preferences.

Everything else equal, a marginal increase in the tax base keeping dispersion constant has two effects. As $y_l$ increases, the right tail of the income distribution becomes thicker. The increase in the mass of relatively richer households has a positive effect on the demand

\(^{10}\)As $\gamma$ increases, parents prefer fertility ($\gamma$) over quality ($\gamma \theta$) since since $\theta < 1$. 
for private education. Call this the (exogenous) *shape effect*. Second, it increases the resources available for public education. This makes the households that were previously indifferent between private and public education to always choose the latter. Call this the (endogenous) *threshold effect*. The two movements have opposite effects on the tax rate and equilibrium enrollment. The net effect depends on the quality of public education (defined as spending per student) relative to the private option.

Public education quality is low when few resources are available (low $\mu$) or when there are many children enrolled (high $\gamma$, i.e. high fertility), corresponding to case 2 and 3.1 in Proposition 2. Panel a in figure 1 shows this case. This implies a relatively large mass of rich households in the right tail choosing, in equilibrium, private education. An increase in $\mu$ further increases this mass, generating a large increase in the support for private education (the shape effect). It dominates the higher enrollment in public education by some middle income families caused by the threshold effect. Therefore the equilibrium tax and public enrollment decrease. However, the equilibrium spending per student can increase as the withdrawal of rich households from public education frees some resources.

Panel b in figure 1 shows the case when the tax base ($\mu$) is high or fertility preference ($\gamma$) is low (regimes 1 and 3.2 in Proposition 2). In this case, the public education resources are high, so only the very rich households prefer private education. Thus, when the tax base increases, the shape effect generates a more modest boost of demand for private education than in the case above. Again, the threshold effect implies borderline households choose public education when average income increases marginally. However, the threshold effect dominates the shape effect in this case. Increased support for public education generates higher enrollment and taxes. Nonetheless, equilibrium spending per student can decrease if the increase in enrollment outpaces that in revenues.

### 3.2 A mean preserving spread

Next, we analyze the relationship between public policies and inequality - proxied by $\sigma$, the standard deviation of the income distribution. We perform a mean-preserving spread and study its implications on equilibrium public spending per student $q^*$, the tax rate $\tau^*$, and the enrollment in public schools $N^*$. Taking the derivative of $q^*$ with respect to $\sigma$ while keeping $\mu$ constant yields:

$$\frac{\partial q^*}{\partial \sigma} = \frac{\phi \theta \mu}{\{1 + \gamma \theta [1 - f(\mu, \sigma)]\}^2} \frac{\partial f(\mu, \sigma)}{\partial \sigma}$$

(25)
and

\[
\frac{\partial N^*}{\partial \sigma} = \frac{1 - \theta}{[(1 - \theta) + \theta \Psi^*]^2} \frac{\partial \Psi^*}{\partial \sigma} = -\frac{1 - \theta}{[(1 - \theta) + \theta \Psi^*]^2} \frac{\phi \mu}{\partial q^*} \frac{\partial q^*}{\partial \sigma}.
\]

Thus, \(\text{sign}(\partial r^*/\partial \sigma) = \text{sign}(\partial \Psi^*/\partial \sigma) = \text{sign}(\partial N^*/\partial \sigma) = -\text{sign}(\partial q^*/\partial \sigma)\). Next, we study the properties of functions \(\partial q^*/\partial \sigma\), \(\partial N^*/\partial \sigma\), and \(\partial r^*/\partial \sigma\). The results are summarized in the following proposition:

**Proposition 3.** Let \(\gamma = [(2/(\delta e)) - 1] / \{\theta [1 - e^{-2}]\} \) and \(\zeta = [(1/\delta) - 1] / \{\theta [1 - (1/e)]\}\), where \(e\) is the Euler’s constant.

1) If \(\gamma \leq \gamma\), then \(\partial r^*/\partial \sigma < 0\), \(\partial N^*/\partial \sigma < 0\), \(\partial q^*/\partial \sigma > 0\);
2) If \(\gamma > \gamma\), then \(\partial r^*/\partial \sigma > 0\), \(\partial N^*/\partial \sigma > 0\), \(\partial q^*/\partial \sigma < 0\);
3) If \(\gamma \in (\gamma, \gamma)\), then there exist a unique \(\mu \in (0, \infty)\) such that
   3.1) if \(\mu \in (0, \hat{\mu}]\), then \(\partial r^*/\partial \sigma > 0\), \(\partial N^*/\partial \sigma > 0\), \(\partial q^*/\partial \sigma \leq 0\);
   3.2) if \(\mu \in (\hat{\mu}, \infty)\), then \(\partial r^*/\partial \sigma < 0\), \(\partial N^*/\partial \sigma < 0\), \(\partial q^*/\partial \sigma > 0\);

(Proof in the Appendix).

The intuition of these results is the following. A mean preserving spread decreases the size of the middle class, adding mass to the tails of the income distribution (poor and rich households). This is the shape effect. Whether support for public education increases or not following this change in the shape of the distribution depends on the initial location of the indifference threshold. Moreover, the endogenous response of this threshold to higher inequality generates an additional effect.

Again, consider the case of low public education quality (low \(\mu\) or high \(\gamma\)), corresponding to cases 2 and 3.1 in Proposition 3, and shown in panel a of figure 2. This implies that many rich and middle income households choose the private option. Thus, the indifference threshold lies relatively far from the right tail, in some middle income range. First, there are two opposing shape effects that arise under a mean preserving spread. On the one hand, the middle class shrinks and so does the support for private education. On the other hand, the mean preserving spread increases the mass of rich households in the right tail who send their children to private education. The overall effect on demand for public education thus depends on the relative magnitude of these opposing effects.

Second, when public education is of low quality, an increase in inequality prompts the threshold households to switch to private education, as the mean preserving spread adds more poor, high fertility households in the left tail, which further reduce spending per student. This is the threshold effect. In this case, the negative effect on the demand for private education caused by the reduction of middle class dominates the positive effects stemming from the extra mass of rich households as well as the endogenous shift in the income threshold towards private schooling. As a result, the enrollment in public schools
Figure 2: A mean preserving spread, indicated by dot variables (e.g. $\sigma^* > \sigma$) and solid lines. Panel a: high fertility preference ($\gamma$) or low tax base ($\mu$). Panel b: low fertility preference or high tax base. The arrow indicates the endogenous change in the indifference threshold. Dark (light) shaded areas represent increases (decreases) in the support for private education.

goes up and so does the tax rate. Despite the increase in revenues (and the extra resources made available by households who left public schools), spending per student is lower in equilibrium as middle income households (who were choosing lower fertility and private schooling before) have been replaced by low income and high fertility households that benefit from public education.

Conversely, when the tax base ($\mu$) is large or fertility preference ($\gamma$) is low, such as in cases 1 and 3.2 in Proposition 3 (panel b of figure 2), the resources for public schooling are higher and, compared with the case above, the mass of middle income households that prefer private education is lower. Thus, the negative effect on the demand for private education generated by a reduction of middle income class is weaker and it is likely to be dominated by the positive effect generated by an increase in the mass of rich households (the shape effects). Second, there is again a threshold effect. In this case, the marginal households strictly prefer public education when inequality increases. Since the indifference threshold is far in the tail, the increase in demand for private education from the extra mass of rich households dominates, generating a decrease in public enrollment and the tax rate. In equilibrium, public school enrollment decreases faster than tax revenue, resulting in an increase in public spending per student.

To sum up, when inequality increases, the size of the poor and rich class increases at the expense of the middle class. When the tax base is low enough, the need for public education spending goes up steeply as a large share of mid income families choosing low fertility and private schooling are now replaced by high fertility low income families that choose public education. Thus, the relatively poorer households steer the political process in their favor, raising the tax rate. As the tax base is a constant share of the mean income,
this increases the public spending per capita, or the size of redistribution. When the tax base is high, the interests of the rich households dominate as the shifts in fertility and education choices associated with the mean preserving spread are now weaker. Thus, the tax rate and the size of redistribution go down. Interestingly, the per student spending in public education, being driven by the endogenous response of enrollment, decreases in the first case and increases in the second.

4 Political power

So far we have assumed that each household carries the same weight in the political process. Next, relying on previous literature, as well as new empirical evidence on public education politics, we document that even in a well-established democracy, like the United States, political participation indicators, such as voter registration and turnout, are positively correlated with income across age and education levels. We then investigate the effects of this bias on public education provision within the theoretical framework.

4.1 Political participation patterns across income groups

Political participation patterns across income groups in the United States have been well documented in the literature (Verba et al. (1995), Rosenstone and Hansen (1993)). At national level, there are striking differences. The 2000 Voter Supplement of the Current Population Survey (CPS) reveals that among those in the 25-44 age group, less than 30% of those in the last income category (under $5000) voted in the 2000 presidential elections while in the highest bracket ($75,000 and over) the turnout was of 70%. The propensity to vote is increasing with household income across all age groups. However, it is much lower in the 18-24 age group, not exceeding 40% in the richest income bracket. Those of age 64 and over vote in proportions ranging from around 50% to almost 90% between the lowest and the highest income groups. Similar patterns are revealed in the 2006 Congressional elections: 50.7% in the lowest income group (less than $10,000) registered but only 24.3% voted, compared to 82.1% registration and 64.6% turnout in the highest bracket ($150,000 and over).

Despite an increasing weight of the federal outlays, public education funding in the United States is still largely decided at state and local level. One may assume that voters understand better, and thus are more concerned with local policies, which they can more easily tweak in their favor. Perhaps surprisingly, a few studies (Morlan (1984), Hajnal and Lewis (2003)) find the contrary: turnout in local elections is on average half that of national elections, with some cities performing much worse. These low figures suggest that disadvantaged segments of the population might be even less represented at local level. While little data is available at this level, in a recent study of mayoral and
city council elections in Californian cities, Hajnal and Lewis (2003) find that an index of socioeconomic status - summarizing income, education, poverty and home ownership measures - is indeed a positive and highly significant correlate of voter turnout.

In the following, we use participation rates in school budget votes across school districts in the state of New York to document differences in political participation at school district level. The New York State Education Department provides vote counts from 2003-2004 budget votes. In order to obtain a measure of the turnout, we divide the number of voters to the number of adult persons in each school district. Comprehensive demographic data at district level is only available from the 2000 Census. However, these characteristics move slowly over time.\(^\text{11}\) We therefore use the Census numbers to compute voter turnout indicators. Across the 628 school districts in NY, the mean turnout is 14\%, with a maximum of 46\% and a minimum of 3\%. Table 1 shows least squares regressions of the turnout variable on school district characteristics. Results confirm that participation in local public education politics is associated positively with household income, the share of population with college degrees and the share of population of age 65 or superior and negatively to the share of population living in poverty and the share of non-white population. While these correlations describe school district aggregates, taken together with the previous evidence on the determinants of political participation, they suggest that income and education differences generate asymmetric propensities to vote even in local political processes, like those related to the provision of public education.

### 4.2 Political power and public education provision

We use the benchmark model to implement and study a general, yet parsimonious political power function that assigns more clout to the rich. Next, we show that under fairly general conditions the equilibrium continues to be unique. Finally, we analyze the effects of uneven political representation on the public education budget, enrollment and spending per student.

To model the direct dependence between income and political power, we define

\[
p(y) = y^\nu, \tag{26}\]

where \(y\) is the income level and \(\nu > 0\). The welfare function (15) becomes

\[
W(\tau) = \int_{y}^{\bar{y}} \left\{ \ln \left[ \frac{y}{(1 + \gamma)(1 + \tau)} \right] + \gamma \ln \left[ \frac{\gamma}{\phi(1 + \gamma)} \right] + \gamma \theta \ln(q) \right\} p(y)f(y)dy + \\
\int_{\bar{y}}^{\infty} \left\{ \ln \left[ \frac{y}{(1 + \gamma)(1 + \tau)} \right] + \gamma \ln \left[ \frac{\gamma(1 - \theta)}{\phi(1 + \gamma)} \right] + \gamma \theta \ln \left[ \frac{\phi \theta y}{1 - \theta} \right] \right\} p(y)f(y)dy.
\]

\(^{11}\)Adult population counts for a representative sample of NY school districts from the 2005 American Community Survey show a correlation of 0.99 with the 2000 figures.
Table 1: Voter turnout correlates

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>log(mean income)</td>
<td>3.10***</td>
<td>1.31</td>
<td>-1.59</td>
<td>2.70**</td>
</tr>
<tr>
<td></td>
<td>(4.66)</td>
<td>(1.05)</td>
<td>(-1.24)</td>
<td>(2.04)</td>
</tr>
<tr>
<td>Share 65+</td>
<td>0.13**</td>
<td>0.12**</td>
<td>0.06</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.17)</td>
<td>(2.12)</td>
<td>(1.10)</td>
<td></td>
</tr>
<tr>
<td>Share college</td>
<td>6.41*</td>
<td>4.68</td>
<td>-0.79</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(1.81)</td>
<td>(1.44)</td>
<td>(-0.26)</td>
<td></td>
</tr>
<tr>
<td>Share poverty</td>
<td>-36.80***</td>
<td>-20.25***</td>
<td>(-6.15)</td>
<td>(-3.33)</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Share non-white</td>
<td>-18.23***</td>
<td></td>
<td></td>
<td>(-8.39)</td>
</tr>
<tr>
<td>Adj. $R^2$</td>
<td>0.035</td>
<td>0.048</td>
<td>0.094</td>
<td>0.171</td>
</tr>
<tr>
<td>N. obs.</td>
<td>628</td>
<td>628</td>
<td>628</td>
<td>628</td>
</tr>
</tbody>
</table>

Notes: The dependent variable is the voter turnout, measured as percentage, in 2003-2004 school budget votes in the state of New York. Mean district income is expressed in dollars. Share variables, such as the age, education and race controls are expressed in percentages. Robust standard errors within parantheses. * indicates significance at the 10 percent level, ** indicates significance at the 5 percent level, *** indicates significance at the 1 percent level.

Then, using (26) and retaining the relevant terms simplifies the expression to

$$W(\tau) = -\ln(1 + \tau) + \gamma \theta \Psi^p \ln(q).$$

(27)

where $\Psi^p = 1 - (y_l/\bar{y})^{1-\nu}$. 

Notice that the only difference relative to (16), the aggregate welfare in the benchmark model, is the weight assigned to public education spending, which here is $\Psi^p$ rather than $\Psi = 1 - (y_l/\bar{y})^\alpha$. It is easy to see that $\Psi^p < \Psi$. Thus, when political power is directly proportional to income, the interests of the rich (lower taxes) have a higher weight in the aggregate welfare. Since they are using mostly private education, the social welfare function reflects the new political balance by assigning a lower weight to public education provision.

The definition of equilibrium is similar to that in the benchmark model. The optimal tax rate is

$$\tau^p = \gamma \theta \Psi^p,$$

while the private education income threshold is given by

$$\bar{y} = \frac{\mu}{\delta \Psi \Psi^p + \gamma \theta \Psi^p}.$$

(28)

Proposition 4. Let $\bar{y}^p = (e^{-\ln(1/2)-\ln\delta} - 1)/\theta$. If $\gamma > \bar{y}^p$, there exists a unique equilibrium
income threshold $\tilde{y} \in (y_l, \infty)$ that solves equation (28), $\forall \nu > 0$. Moreover, uniqueness is ensured $\forall \gamma > 0$, for sufficiently small $\nu$. (Proof in the Appendix)

In the benchmark model, higher public education enrollment translates into higher tax revenues as the tax rate increases with the propensity to choose public education and the tax base stays constant. However, now the chosen tax rate reflects the taste of rich households for private education. In the following we study how the main results in the previous section change when we allow for political power.

Notice that the political power specification (26) is a monotonic and continuous function of income. Furthermore, as $\nu \to 0$, the income weights in the social welfare function vanish, yielding the benchmark model. Thus, in the limit, the results derived in Propositions 2 and 3 continue to hold.

Moreover, (26) provides a tractable way of determining the parameter $\nu$ based on the income level and the propensity to vote. Thus, knowing that the income groups $y_l$ and $y_h$ have propensities to vote $p_l$ and $p_h$ respectively, $\nu = \ln(p_h/p_l)/\ln(y_h/y_l)$. Using the median income levels in the top and the bottom brackets of the CPS data described above, together with the respective turnout figures, yields $\nu = 0.26$.

For illustration, we replicate the exercises in Propositions 2 and 3 with and without political power. We use $\nu = 0.26$, $\phi = 0.075$, $\theta = 0.4$ and $\gamma = 2.7$ in the benchmark model, corresponding to the case of intermediate fertility rates (case 3).

Figure 3 graphs the three policy variables - public school enrollment, public spending per capita and the tax rate - as functions of the average income per capita, keeping dispersion constant. The thin lines represent the benchmark model and the thick lines the model with political power.

As expected, adding income correlated political weights lowers the tax rates at all income levels. However, lower taxation determines some households to switch to private education and thus enrollment in public schools also declines. Thus, public spending per student declines much less than revenues. Besides these level effects, political power induces tax rates to strictly increase with the mean income. In the benchmark model the tax rates follow a U-shaped pattern as a function of mean income for intermediate values of $\gamma$.

The thin lines in figure 4 display, from left to right, changes in the main variables, for a range of mean incomes when the standard deviation of the distribution increases by 10%. Thus, in the leftmost panel, public school enrollment increases with inequality in poor districts but declines in more unequal rich districts, as already shown in Proposition 3. Then, we allow for political power. The thick lines depict similar changes when inequality increases. Rich households now have more power in setting the tax rate, such that higher inequality leads to lower tax rates in all districts as well as more abrupt declines in spending per student in poor districts. Case 3 in Proposition 3 shows that for intermediate values of the altruism coefficient $\gamma$, the equilibrium tax rate increases with

20
Main education variables as a function of the mean income (tax base), keeping dispersion constant, under political power ($\nu = 0.26$, thick line) versus benchmark ($\nu = 0$, thin line).

inequality in poor districts, where the welfare of the relatively more numerous disadvantaged households depends on the quality of public schooling. This effect is overturned by allowing richer households to enjoy political power.

We have shown that augmenting the model to include political power preserves the uniqueness of the politico-economic equilibrium under fairly general conditions and induces the tax rate and the public spending per student to decrease more strongly with inequality. Moreover, while comparative statics results in the benchmark model are preserved for small asymmetries in political power ($\nu \to 0$), for values of $\nu$ consistent with turnout levels by income categories, the tax rate can respond in a monotonic fashion to changes in inequality.\(^{12}\)

5 Empirical evidence

In the previous section, we have used a political economy model of public education provision with endogenous fertility and private education decisions to investigate the interactions between income inequality and redistribution. Our analysis shows how public education provision depends on the mean income of the economy (the tax base) and the level of income inequality (the dispersion around the mean).

One important result is that an increase in inequality, defined as a mean preserving spread, can have opposing effects on spending per student in poor and rich economies. The model also predicts that such changes are accompanied by non-monotonic movements

\(^{12}\)One can show that $\partial \tau^*/\partial \sigma < 0, \forall \mu > 0$ for $\nu$ large enough.
Changes in the main education variables from a 10 percent increase in income dispersion, for a given mean income, under political power (ν = 0.26, thick line) versus benchmark (ν = 0, thin line).

in public school enrollments.

Specifically, more unequal poor districts vote higher education taxes, but have a higher share of students enrolled in public schools, so the education quality, measured as spending per student decreases. In contrast, an increase in inequality in a rich district generates a decrease in the tax rate, while reducing the share of enrollment in public education which results in an increase of per student spending. Allowing for income based political power implies that even in poor districts, an increase in inequality can decrease the tax revenues, which further depresses spending per student.

In this section we use U.S. data at different aggregation levels to test these predictions. First, we ask whether the predicted non-monotonic relationship between household income inequality and education public spending per student finds support in the data. Second, we test whether the associated changes in public school enrollments and public spending per capita are consistent with the theory.

5.1 Data and methodology

Our analysis draws on school district (SD) demographic and financial data in 1990 and 2000 from the National Center for Education Statistics (NCES) of the U.S. Department of Education. Details on the data sources and summary statistics are provided in Appendix B.

In line with the theory, the empirical exercises below allow the effects of inequality to vary with the income level by splitting the sample of districts into rich and poor. Since in
the model a household includes only one parent, we use household level data to construct the empirical counterparts of the model’s tax base and income dispersion. Thus, in all regressions we include the average household income and its dispersion in the district. This allows us to test the theoretical effects of a mean preserving spread (Proposition 3), as well as those derived from a change in mean income (Proposition 2 and Corollary 1).

First, we estimate for each SD the mean and the standard deviation of the household income distribution using the 16 income brackets provided in the School District Tabulation data.\textsuperscript{13}

The literature on inequality and public spending usually considers other measures such as the Gini coefficient or the mean/median ratio. However, given the particular hypotheses we aim to test, these measures are inadequate as they do not distinguish between changes in the mean and changes in the variance of the distribution.

We estimate the following equations:

\[
L = \beta_0 + \beta_\mu \text{Mean} + \beta_\sigma \text{St.Dev.} + \beta_\mu^R \text{Mean} \ast \text{Rich} + \beta_\sigma^R \text{St.Dev.} \ast \text{Rich} + \beta_X X
\] (29)

where \(L\) is local public spending per student, \(\text{Mean}\) and \(\text{St.Dev.}\) are estimates of the first two moments of the household income distribution and \(\text{Rich}\) is a categorical variable which takes value one for districts above some income threshold and zero otherwise. Finally, \(X\) is a vector of control variables described below. The income threshold for the categorical variable is set at 1.44 times the median value for the SD average income per capita, or around $25,600.\textsuperscript{14} We estimate similar models for the share of public enrollment and the local spending per capita.\textsuperscript{15}

**Control variables.** To isolate the effect of local politics on education funding, we include the state and federal revenues per student in the regression.\textsuperscript{16} State fixed effects remove further (unmodeled) idiosyncratic biases. We also control for the SD type and size using a set of eight categorical variables spanning the rural-urban axis while also accounting for size (e.g. small town, mid-size city, large city).

Other types of heterogeneity beside income have been shown to shape public spending decisions in systematic ways. Racial diversity is one of them (see Alesina et al. (1999),

\textsuperscript{13}All households are assigned an income equal to the mid-point of their bracket. The average income of households in the last income bracket is directly available in the data. Alternative estimates that assume, for each bracket, median income levels estimated from micro-data yield very similar results. To check the accuracy of estimates based on grouped data, we compute Gini coefficients at the state level as population weighted averages of the corresponding SD Gini coefficients. Comparing our estimates against microdata Census based estimates we obtain a correlation of 0.99%.

\textsuperscript{14}Results are robust to different income threshold values.

\textsuperscript{15}In order to obtain the empirical counterpart of the model spending per capita (which is also the tax rate), we divide local public spending by the total number of households with kids. This indicator accounts both for the single parent assumption in the model as well as for the presence of households without children in the data.

\textsuperscript{16}Spending controls also capture other potential biases in the state level policies, such as for example correlations induced by yardstick competition.
Boustan et al. (2010)). We therefore include a Herfindahl index of population shares as well as the share of non-white population to account for such biases. Another factor that is likely to play an important role in the political support for public education is the population age structure (see Poterba (1997), Harris et al. (2001)). Since this aspect is not explicitly addressed in the theoretical model, we add two demographic controls: the share of residents over the age 65 and the share of under 18 in total SD population. We also control for other characteristics that may alter the spending patterns such as college education attainment and share living in poverty.

**Tax and Expenditure Limitations.** Finally, we address the issues associated with the tax and expenditure limitations (TELs) at local level. In the context of education finance, these limitations have emerged, starting with the 1971 ruling of the California Supreme Court in *Serrano vs. Priest*, as part of a wider redistribution effort aimed at equalizing the quality and availability of educational programs and services across school districts within a state.

These TELs include limits on overall and specific property tax rates and property tax revenue, caps on property assessment values, restrictions on general revenue and expenditure levels as well as full disclosure clauses. Previous literature studying the effects of such constraints on education finance (e.g. Shadbegian (2003), Figlio (1997)) has shown that their stringency varies across states and moreover, in many cases these limitations can be overridden, e.g. by majority vote of the electorate. However, in those instances where the TELs are more severe, school financing is no longer determined by local politics, as implied by our model.

We address this issue by testing our results in three different samples. First we consider the complete sample, denoted by *All*, covering the entire U.S. except Alaska, Washington D.C. and Hawaii. We then consider three criteria to eliminate the SDs where local politics is significantly constrained by state imposed limitations: property tax rate limits, caps on revenues and expenditures and finally limitations applied to property reassessment. We use data from Yuan et al. (2007) to classify states according to the stringency of these TELs. Thus, in our second sample, denoted by *No-TELs 1*, we consider states that either do not have any limitations on local tax rates, revenues or expenditures in place in 2000 or, if they have any, these restrictions can be overridden by a local vote. Finally, we construct a third sample, denoted by *No-TELs 2*, removing from sample *No-TELs 1* the states that have adopted property assessment increase rates lower that 5%. This threshold has been used in the literature on local public finance to distinguish between stringent and non stringent TELs (e.g. Shadbegian (1999)). Appendix B contains more details on the composition of these samples.

**Endogeneity concerns.** We first estimate equation (29) using OLS. However, least squares estimates may be biased due to reverse causality stemming from endogenous sorting across SDs. If this is the case, the demographic as well as the economic characteristics
of the school districts are likely to be endogenous. The same applies to state and federal funding levels which are, by construction, functions of these characteristics (e.g. poverty level).

To deal with this potentially important issue, we follow a two-step strategy.

First, we address the potential endogeneity of demographic and policy variables by instrumenting all these SD characteristics in 2000 including the state and federal funding levels by the corresponding values in 1990. Regarding the SD mean income, \( \mu \), and its dispersion, \( \sigma \), we construct instrumental variables that are correlated with the SD tax base and inequality, respectively, but not directly linked to local education funding, as discussed further below. We then obtain a first set of estimates using instrumental variable techniques on the various subsamples of the SD data.

The second step we take to mitigate self-selection concerns relies on the observation that household relocation decisions are subject to significant spatial frictions. For example, the 2000 census reveals that 55.7\% of population was living in the same house as in 1995, 25.7\% changed house within the same county and 9.9\% moved to a different county within the same state. While this aggregate stylized fact can hide a large degree of spatial heterogeneity, it implies nonetheless that household mobility, and hence selection, is less important across larger geographical units. In our case, a natural observation unit to consider is the Metropolitan Statistical Area. According to the Census Bureau, these areas include “the counties containing the core urban area, as well as any adjacent counties that have a high degree of social and economic integration (as measured by commuting to work) with the urban core.” Thus, MSAs are good approximations for actual labor market areas, which, in turn, are likely to circumscribe the relocation possibilities and efforts of a large share of households. Indeed, comparing annual migration rates across different geographical units during 1980-2000, Molloy et al. (2011) find that while average inter-county rates are almost 6\% of total population, the inter-MSA rates are around 3\%.

The NCES school district data for 2000 records the MSA to which an SD belonged in 1999. We use this information to construct population weighted averages of all our variables at MSA level. We consider the SDs that do not belong to an MSA as freestanding labor market areas and keep them in the data set. Since some MSAs span more than one state, we assign the artificial SDs to the state where the largest share of the population resides. Also, we assign it the set of geographic characteristics of the largest SD (usually located in the core urban area). We denote the original data at school district level by SD and the MSA clustered data by Metro-SD. Finally, we replicate the No-TELs fiscally unrestricted subsamples in the Metro-SD data.

In conjunction with the instrumental variable strategy described before, the artificial school districts created by this procedure internalize some of the households’ relocation potential and thus further mitigate potential endogeneity concerns.

We now describe the procedure used to construct instruments for the 2000 SD income
distribution moments. Following the approach in Boustan et al. (2010) we create for each SD a synthetic income distribution that replaces the actual frequencies across income levels in 2000 with the corresponding 1990 shares. These are constructed by converting the endpoints of the 1990 SD income bins into percentiles of the national income distribution, obtained from 1990 Census micro-data from.\(^\text{17}\) The percentiles are then projected onto the 2000 national income distribution, again obtained from Census micro-data, generating new, synthetic income cutoffs. SD level population shares from 1990 are then assigned to the synthetic 2000 income brackets. This artificial distribution effectively freezes the population distribution of an SD across income brackets at 1990 levels and changes the representative income in each bracket according to national trends.\(^\text{18}\) Thus, it captures the exogenous component of income inequality (due to broader trends in economic activity that have shifted the national income distribution), which individual SDs are too small to alter.

The mean and standard deviation of this synthetic distribution are used to instrument the actual 2000 income distribution moments. There is a strong correlation between the synthetic and the actual moments: 0.96 for the mean and 0.84 for the standard deviation.

In the following we discuss the effects of inequality on the local spending per student. We then analyze the effects on enrollment and spending per capita.

### 5.2 Spending per student

For brevity, we omit all the coefficients associated with control variables. We retain \(\beta_\mu\) and \(\beta_\sigma\), the coefficients associated with the mean and respectively the standard deviation of the local household income distribution, and \(\beta_\mu^R\) and \(\beta_\sigma^R\), their interactions with the income indicator variable Rich. We compute tests to assess whether poor and rich districts respond in different ways to inequality changes by checking whether the sign of the inequality coefficient in poor districts \(\beta_\sigma\) is different from the sign of the rich districts’ coefficient, \(\beta_\sigma + \beta_\sigma^R\). For example, in the case of local public spending per student, we test \(\beta_\sigma < 0\) in poor districts and \(\beta_\sigma + \beta_\sigma^R > 0\) in rich districts. We set the null hypothesis for rich districts as \(H_0^R: \beta_\sigma + \beta_\sigma^R \leq 0\) and verify if it is rejected by the data. A similar test is performed for tax base changes. We report the statistics of these tests; the corresponding \(p\)-values are enclosed in square brackets.

The first column of table 2 contains the least squares estimates from the benchmark sample (All). Columns (2)-(4) contain the instrumental variable estimates (IV) using the

---

17The IPUMS-USA database is provided by Ruggles et al. (2010).

18For example, suppose that in a given SD, 12% of the households had an income between $10,000 and $20,000 in 1990. Assume that in the 1990 U.S. income distribution, $10,000 and $20,000 correspond to the 7th and the 10th percentiles, respectively. In the 2000 U.S. distribution, the 7th percentile now corresponds to $15,000 while the 10th percentile corresponds to $30,000. Thus, in the synthetic 2000 income distribution of the SD, 12% of the households are assumed to have incomes between $15,000 and $30,000.
school district samples (All, No-TELs 1 and No-TELs 2) and columns (5)-(7) contain the equivalent IV estimates in the Metro-SD samples.

Table 2: Inequality and redistribution: effects on spending per student.

<table>
<thead>
<tr>
<th>Sample</th>
<th>SD</th>
<th>Metro-SD</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>All</td>
<td>All</td>
</tr>
<tr>
<td></td>
<td>LS</td>
<td>IV</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>( \beta_{\mu} )</td>
<td>(1.6)</td>
<td>(2.7)</td>
</tr>
<tr>
<td>Rich*Std.Dev.</td>
<td>37.9***</td>
<td>44.5***</td>
</tr>
<tr>
<td>( \beta_{\mu}^R )</td>
<td>(4.5)</td>
<td>(5.1)</td>
</tr>
<tr>
<td>Mean Income</td>
<td>29.7***</td>
<td>39.1***</td>
</tr>
<tr>
<td>( \beta_{\mu}^R )</td>
<td>(3.2)</td>
<td>(3.9)</td>
</tr>
<tr>
<td>Rich*Mean</td>
<td>–25.5***</td>
<td>–35.0***</td>
</tr>
<tr>
<td>( \beta_{\mu}^R )</td>
<td>(4.0)</td>
<td>(4.6)</td>
</tr>
</tbody>
</table>

Hypotheses tests

\( H^*_a : b_0 + \beta_{\mu} > 0 \)

<table>
<thead>
<tr>
<th></th>
<th>All</th>
<th>All</th>
<th>No-TELs 1</th>
<th>No-TELs 2</th>
<th>All</th>
<th>All</th>
<th>No-TELs 1</th>
<th>No-TELs 2</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>58.94</td>
<td>38.57</td>
<td>24.34</td>
<td>20.92</td>
<td>1.76</td>
<td>2.33</td>
<td>1.86</td>
<td></td>
</tr>
<tr>
<td></td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.09]</td>
<td>[0.06]</td>
<td>[0.09]</td>
<td></td>
</tr>
</tbody>
</table>

\( H^*_a : b_0 + \beta_{\mu}^R > 0 \)

|          | 1.02 | 0.90 | 1.15 | 1.41 | 7.08 | 6.43 | 5.59 |
|          | [0.16] | [0.17] | [0.14] | [0.12] | [0.00] | [0.01] | [0.01] |

Adj. \( R^2 \)

|          | 0.61 | 0.32 | 0.30 | 0.29 | 0.14 | 0.15 | 0.14 |
|          | 12865 | 12392 | 9156 | 8580 | 6798 | 5391 | 4995 |

Notes: The dependent variable is the local public spending per student, expressed in dollars. Coefficients associated with the control variables (described in the text) are not reported. The school district mean household income and its standard deviation are expressed in thousand dollars. For data sources and summary statistics see Appendix B. Robust standard errors within parantheses. The hypotheses test sign restrictions derived from the theoretical model. The table reports the size of the tests and, within square brackets, the p-values at which the corresponding null hypotheses are rejected. * indicates significance at the 10 percent level, ** indicates significance at the 5 percent level, *** indicates significance at the 1 percent level.

The coefficients have stable signs and are in general statistically significant across subsamples and estimation methods. Relative to the least squares estimates, the IV coefficients increase in absolute value. Computing the LM test of the Kleibergen-Paap rk statistic\(^{19}\) as well as the related weak identification Wald F statistic, we reject the weak instruments hypothesis at any customary confidence level.

Based on these estimates one can reject the null hypothesis that poor and rich districts react in the same way to an increase in income dispersion, with a high degree of confidence in all SD level samples. The Metro-SD estimates yield a similar conclusion. Thus, an

\(^{19}\)We compute the test using \textit{ivreg2} command in Stata 12 and obtain values above 200. The test statistic is distributed \( \chi^2 \) with \( L_1 - k_1 + 1 \) degrees of freedom where \( L_1 \) is the number of excluded instruments and \( k_1 \) the number of endogenous regressors. For more details, see Kleibergen and Paap (2006). The Wald F statistics is above 20 in all samples.
increase in inequality lowers local spending per student in poor districts but increases it in rich districts, as predicted by case 3 of Proposition 3.

Moreover, in line with the theoretical results derived in Corollary 1, the mean income has a positive effect on spending per student, at least in the poor districts. While in the SD samples an increase in the tax base actually appears to depress spending per student in the rich districts, the result vanishes in all the Metro-SD subsamples. Based on these estimates, we can conclude that spending per student increases with the tax base in all school districts although it does so faster in poorer districts.

In order to further test the validity of the theoretical model, we investigate the implications for public school enrollment and public spending per capita.

### 5.3 Public education enrollment

<table>
<thead>
<tr>
<th>Sample</th>
<th>SD</th>
<th>Metro-SD</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>LS</td>
<td>IV</td>
</tr>
<tr>
<td>Std.Dev</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$[\beta_{\text{Std.Dev}}]$</td>
<td>(0.005)</td>
<td>(0.009)</td>
</tr>
<tr>
<td>Rich*Std.Dev</td>
<td>-0.1***</td>
<td>-0.1***</td>
</tr>
<tr>
<td>Mean Income</td>
<td>(0.02)</td>
<td>(0.02)</td>
</tr>
<tr>
<td>Rich*Mean</td>
<td>-0.04***</td>
<td>-0.05***</td>
</tr>
<tr>
<td>Hypotheses tests</td>
<td>(0.01)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>$b_n + b_{\beta_{\text{Std.Dev}}}$ &lt; 0</td>
<td>55.83</td>
<td>66.27</td>
</tr>
<tr>
<td>Hypotheses tests</td>
<td>(0.00)</td>
<td>(0.00)</td>
</tr>
<tr>
<td>$b_n + b_{\beta_{\text{Std.Dev}}}$ &gt; 0</td>
<td>18.33</td>
<td>31.92</td>
</tr>
<tr>
<td>Hypotheses tests</td>
<td>(0.00)</td>
<td>(0.00)</td>
</tr>
<tr>
<td>Adj. $R^2$</td>
<td>0.31</td>
<td>0.07</td>
</tr>
<tr>
<td>N. obs.</td>
<td>12865</td>
<td>12392</td>
</tr>
</tbody>
</table>

**Notes:** The dependent variable is the enrollment share in public schools, expressed in percentages. Coefficients associated with the control variables (described in the text) are not reported. The school district mean household income and its standard deviation are expressed in thousand dollars. For data sources and summary statistics see Appendix B. Robust standard errors within parantheses. The hypotheses test sign restrictions derived from the theoretical model. The table reports the size of the tests and, within square brackets, the p-values at which the corresponding null hypotheses are rejected. * indicates significance at the 10 percent level, ** indicates significance at the 5 percent level, *** indicates significance at the 1 percent level.

Table 3 shows how a change in inequality affects public education enrollment. Simi-
larly to the results on spending per student, these effects vary with the income level. In the SD samples, the baseline dispersion coefficient $\beta_\sigma$ is positive but not significantly different from zero, implying small and/or noisy enrollment changes in response to changes in income inequality in poor SDs. However, the coefficient of the interaction term $\beta^{R}_\sigma$ is negative and highly significant. Therefore, we reject the null hypothesis of monotonous response across income levels and conclude that more unequal school districts rely less on public education only when the average income is high enough. This result becomes even stronger in the Metro-SD data where the baseline coefficients become significant in all subsamples.

The tax base effects are also in line with the theory. In all SD samples an increase in the average household income lowers public school enrollment in poor districts and increases it in rich districts. Looking at the Metro-SD results, one can notice that while the baseline tax base coefficient and the interaction term have opposite signs, the hypothesis $\beta_{\mu} + \beta^{R}_\mu > 0$ is not supported. An explanation for this result is that the Metro-SD data has by construction clustered urban labor markets into larger synthetic areas and maintained the more remote and poor districts as separate observations, thus increasing the relative importance of the latter group. If in poor districts higher average income lowers public school enrollment, as predicted by the model, then the opting out channel is stronger in the Metro-SD samples. Indeed, when comparing the SD against the Metro-SD estimates, the baseline coefficients on both dispersion and mean income increase in absolute value by an order of magnitude while the interaction terms’ coefficients, characterizing the extra response in rich districts, are rather similar.

5.4 Spending per capita

Finally, table 4 describes the effects on the overall redistribution through public education provision, as measured by per capita spending. Note that in our model spending per capita is different from the tax rate. While it can be shown that inequality affects redistribution in the same way as the tax rate, the response to a change in tax base is different. In particular, numerical simulations suggest that redistribution increases with the tax base in all school districts.\textsuperscript{20}

In the SD samples, the coefficients associated with income dispersion are negative and significant in the poor districts. This goes against the benchmark theoretical results. Moreover, the interaction coefficients are positive but small and less precisely estimated. Looking at the Metro-SD estimates, the interaction terms are larger but only marginally significant. The hypothesis of lower redistribution in rich districts ($H_{\sigma} : \beta_\sigma + \beta^{R}_\sigma < 0$)

\textsuperscript{20}Spending per capita is the right-hand side of (14). Thus, under a mean preserving spread, \[\frac{\partial [\mu^\tau/(1+\tau)]}{\partial \sigma} = \frac{1}{(1+\tau)^2} \frac{\partial \tau}{\partial \sigma}.\] A tax base increase generates \[\frac{\partial [\mu^\tau/(1+\tau)]}{\partial \mu} = \frac{\tau + \tau^2 + \mu \frac{\partial \tau}{\partial \mu}}{(1+\tau)^2}.\] Intuitively, since \(\frac{\partial \tau}{\partial \mu} < 0\) at most for small \(\mu\) values (see figure 3), the negative term is likely dominated.
Table 4: Inequality and redistribution: effects on spending per capita.

<table>
<thead>
<tr>
<th>Sample</th>
<th>SD Metro-SD</th>
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<td>All</td>
<td>All</td>
<td>No-TELs 1</td>
<td>No-TELs 2</td>
<td>All</td>
<td>No-TELs 1</td>
<td>No-TELs 2</td>
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<td></td>
<td>LS</td>
<td>IV</td>
<td>IV</td>
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<td>(6)</td>
<td>(7)</td>
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<td></td>
</tr>
<tr>
<td>Std.Dev</td>
<td>–1.5</td>
<td>–7.3**</td>
<td>–6.6**</td>
<td>–7.0**</td>
<td>–10.0*</td>
<td>–10.5*</td>
<td>–11.6*</td>
<td></td>
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<td></td>
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</tr>
<tr>
<td>[βₐ]</td>
<td>1.2</td>
<td>2.4</td>
<td>2.8</td>
<td>2.9</td>
<td>5.3</td>
<td>6.1</td>
<td>6.4</td>
<td></td>
<td></td>
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<tr>
<td>Rich*Std.Dev</td>
<td>9.3***</td>
<td>9.3***</td>
<td>8.0</td>
<td>6.9</td>
<td>22.6**</td>
<td>21.6**</td>
<td>21.4*</td>
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<tr>
<td>[βᵣᵣ]</td>
<td>2.8</td>
<td>4.3</td>
<td>4.9</td>
<td>5.1</td>
<td>9.5</td>
<td>10.1</td>
<td>12.5</td>
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<td></td>
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<tr>
<td>Mean Income</td>
<td>14.6***</td>
<td>17.1***</td>
<td>14.1***</td>
<td>14.2***</td>
<td>40.9</td>
<td>45.4</td>
<td>47.6</td>
<td></td>
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<tr>
<td>[βᵣᵢ]</td>
<td>2.4</td>
<td>4.4</td>
<td>5.4</td>
<td>5.5</td>
<td>27.6</td>
<td>32.6</td>
<td>33.9</td>
<td></td>
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<tr>
<td>Rich*Mean</td>
<td>–6.2**</td>
<td>–9.2**</td>
<td>–9.8**</td>
<td>–8.8*</td>
<td>–26.7**</td>
<td>–28.1**</td>
<td>–29.6*</td>
<td></td>
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<td></td>
<td></td>
</tr>
<tr>
<td>[βᵣᵢ]</td>
<td>2.4</td>
<td>4.1</td>
<td>4.7</td>
<td>4.9</td>
<td>11.9</td>
<td>13.5</td>
<td>15.9</td>
<td></td>
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</tbody>
</table>

Hypotheses tests

Hₐ : bₐ + bᵣᵢ < 0

<p>| | | | | | | | | | | | |</p>
<table>
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<tr>
<td></td>
<td>7.83</td>
<td>0.32</td>
<td>0.11</td>
<td>0.00</td>
<td>2.23</td>
<td>1.59</td>
<td>0.84</td>
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<td>[0.82]</td>
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Hₐ : bₐ + bᵣᵢ > 0

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</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>8.04</td>
<td>4.21</td>
<td>0.82</td>
<td>1.26</td>
<td>0.45</td>
<td>0.50</td>
<td>0.51</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>[0.00]</td>
<td>[0.02]</td>
<td>[0.18]</td>
<td>[0.13]</td>
<td>[0.25]</td>
<td>[0.24]</td>
<td>[0.24]</td>
<td></td>
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</tr>
</tbody>
</table>

Adj. R²          | 0.46       | 0.22       | 0.19       | 0.20       | 0.09       | 0.09       | 0.10       |            |            |            |            |

N. obs.         | 12865      | 12392      | 9156       | 8580       | 6798       | 5391       | 4995       |            |            |            |            |

Notes: The dependent variable is the local public spending per capita, expressed in dollars. Coefficients associated with the control variables (described in the text) are not reported. The school district mean household income and its standard deviation are expressed in thousand dollars. For data sources and summary statistics see Appendix B. Robust standard errors within parentheses. The hypotheses test sign restrictions derived from the theoretical model. The table reports the size of the tests and, within square brackets, the p-values at which the corresponding null hypotheses are rejected. * indicates significance at the 10 percent level, ** indicates significance at the 5 percent level, *** indicates significance at the 1 percent level.

is rejected at the 5% level in the unrestricted sample and even at the 10% level in the samples that retain fiscally unconstrained districts. Thus, we conclude that inequality decreases redistribution in poor districts and has ambiguous effects in the rich districts.

On the other hand, the effect of the tax base is positive in poor districts in all subsamples. The marginal effect in rich districts is negative but small and less precisely estimated. Therefore the net effect on redistribution in these districts stays positive but mostly statistically not different from zero.

While the estimated tax base effects corroborate to some extent the model’s predictions, data seems to suggest that different mechanisms are driving the response of spending per capita to higher inequality. In section 4, we have presented evidence that political participation is not independent of income at school district level. We have then shown that including this channel in the theoretical model generated tax rates which are increasing in average income (albeit lower in absolute terms compared to the benchmark) and decline monotonically with inequality. This scenario is consistent with the observed
positive tax base effects at all income levels as well as with the negative effect of inequality in poor districts. In these districts, higher dispersion in income leads to higher enrollments in public schools despite reduced funding per student. Other factors - such as tax progressivity or different preferences for education - might affect the behavior of the local tax revenues. Recall, for example, that case 2 of Proposition 2 established that for high values of $\gamma$, the tax rate increases with inequality. If rich districts are characterized by higher altruism toward children, beyond that captured in parental characteristics included in the regression, this may induce a stronger positive response of tax rates to inequality in these districts.

Taken together, the empirical exercises using U.S. school district data have enabled us to draw three main conclusions: First, we have found robust evidence across samples and estimation methods in favor of the hypothesis that changes in inequality have different effects on public education provision at different income levels. Second, while the literature usually focuses on redistribution measures such as spending per capita or spending per student, our empirical analysis of the effects on public school enrollment provided specific support for the theoretical framework proposed in the paper emphasizing fertility and opting out decisions. Finally, while in the theoretical framework we emphasize the role of income linked political power in education politics, our empirical analysis seems to suggest that this feature might be indeed relevant, particularly in explaining the response of local spending per capita to an increase in inequality.

5.5 Robustness

Besides the controls described in the previous sections, we performed a number of additional robustness tests. We replicated the estimates using different income thresholds to separate rich and poor districts. We also computed clustered standard errors at state level in order to address any systematic bias in the response of the dependent variables beyond that captured by the set of fixed effects. We expanded the set of geographical control variables to include a measure of remoteness (fringe/distant/remote), in addition to urbanization degree and size as well as an alternative set of indicators capturing the position of the SD relative to the closest metropolitan area. In addition to the set of size categorical variables, population was also included as a control. We also included the share of immigrants in a district to capture peer-effects heterogeneity that could affect the quality of public education. As strategic behavior of the SDs, such as fiscal competition, could potentially alter our results, we have also controlled for the spending per student/capita in the SD closest in space, in terms of linear distance. A number of communities are served by overlapping SDs, dealing separately with elementary and secondary education. To avoid any double counting biases, we also estimated our equations excluding "secondary only" school districts. The main empirical results survive all these
6 Conclusion

The paper investigates the role of inequality in the determination of public education spending, using a probabilistic voting model of public education provision with a private schooling option and endogenous fertility. We show that modelling household income heterogeneity to be consistent with the skewness of empirical income distributions has important consequences for the qualitative properties of the political equilibrium.

Generalizing results in the previous literature, we find a non-monotonic relationship between inequality and per student public spending, depending on 1) the preference for fertility relative to children quality and 2) the average per capita income (the tax base) in the economy. For moderate fertility preferences, we show that a mean preserving spread decreases public spending per student but increases tax rates and public school enrollments when the average income per capita is low, while it has opposite effects in richer economies. A marginal increase in the tax base, holding income dispersion constant, also yields non-monotonic effects.

In the benchmark framework the households enjoy equal influence in local education politics. We show that in the U.S., participation in local education politics varies with socioeconomic indicators. We then extend the basic model to include income dependent political power and study its properties.

Finally, the empirical analysis of U.S. school district data lends strong support to predictions derived from the theoretical model, providing a way to reconcile previous contradictory results in the political economy literature on redistribution and inequality.

While the paper focuses on the effects of inequality on education spending, investigating the dynamic effects of education in this setup, endogenizing sorting across districts and exploring the resulting policy implications are all interesting topics for future research.

References


21 Results are available upon request from the authors.


Corcoran, S. and W. N. Evans (2009). Income Inequality, the Median Voter, and the Support for Public Education. *Department of Economics, University of Notre Dame*.


7 Appendix A

Proof of Proposition 1. The LHS of equation (19) is continuous and increasing in \( \bar{y} \), while the RHS is continuous and decreasing in \( \bar{y} \). Moreover, \( \lim_{\bar{y} \to \infty} \text{LHS}(\bar{y}) = \infty > \lim_{\bar{y} \to \infty} \text{RHS}(\bar{y}) = \mu/ [\delta(1 + \gamma \theta)] \). Next, \( \text{RHS}(y_l) = \mu/\delta = \alpha y_l/ [\delta(\alpha - 1)] > \text{LHS}(y_l) = y_l \).

By the Intermediate Value Theorem, the solution of equation (19) is interior and unique.

Proof of Proposition 2. Recall \( f(\mu, \sigma) = [y_l(\mu, \sigma)/\bar{y}(\mu, \sigma)]^{\alpha(\mu, \sigma)} \) and

\[
q^* = \frac{\phi \theta \mu}{1 + \gamma \theta (1 - f)}.
\] (A.1)

Next, we get \( \partial q^*/\partial \mu \):

\[
\frac{\partial q^*}{\partial \mu} = \frac{(q^*)^2}{\phi \mu} \left\{ \frac{\phi}{q^*} + \gamma \frac{\partial f(\mu, \sigma)}{\partial \mu} \right\}.
\] (A.2)

\[
\frac{\partial f(\mu, \sigma)}{\partial \mu} = f(\mu, \sigma) \left[ \frac{\partial \alpha}{\partial \mu} \ln \left( \frac{y_l}{\bar{y}^*} \right) + \alpha \bar{y}^* \frac{\partial \mu}{\partial \mu} \frac{\partial \bar{y}^*}{\partial \mu} - y_l \frac{\partial \bar{y}^*}{\partial \mu} \right]
\] (A.3)

\[
= f \left[ \frac{\partial \alpha}{\partial \mu} \ln \left( \frac{y_l}{\bar{y}^*} \right) + \alpha \frac{\partial y_l}{\bar{y}^*} \frac{\partial \mu}{\partial \mu} \right] - f \alpha \frac{\partial \bar{y}^*}{\partial \mu}.
\] (A.4)

We use (1) to write \( y_l \) and \( \alpha \) as functions of the first two moments, \( \mu \) and \( \sigma \):

\[
y_l(\mu, \sigma) = \frac{\alpha(\mu, \sigma) - 1}{\alpha(\mu, \sigma)} \mu, \text{ and } \alpha(\mu, \sigma) = 1 + \sqrt{1 + \frac{\mu^2}{\sigma^2}}.
\] (A.5)

We use (A.5) to find \( \partial y_l/\partial \mu \):

\[
\frac{\partial y_l}{\partial \mu} = \frac{\alpha - 1}{\alpha} + \mu \frac{\partial \alpha}{\partial \mu}.
\] (A.6)

where

\[
\frac{\partial \alpha}{\partial \mu} = \left( 1 + \frac{\mu^2}{\sigma^2} \right)^{-1/2} \frac{\mu}{\sigma^2} > 0.
\] (A.7)

Using (A.6) and

\[
\frac{1}{\bar{y}^*} \frac{\partial \bar{y}^*}{\partial \mu} = \frac{\partial q^* 1}{\partial \mu q^*}
\] (A.8)
in (A.4), we obtain:
\[
\frac{\partial f(\mu, \sigma)}{\partial \mu} = f \left[ \frac{\partial \alpha}{\partial \mu} \ln \left( \frac{y_i}{y^*} \right) + \frac{\alpha - 1}{y_i} + \frac{\mu}{\alpha y_i} \frac{\partial \alpha}{\partial \mu} \right] - f \frac{\alpha}{q^*} \frac{\partial q^*}{\partial \mu}. \tag{A.9}
\]

We use (A.9) in (A.2) and set \( y_i = (\alpha - 1)\mu/\alpha \). Rearranging terms, we get:
\[
\frac{\partial q^*}{\partial \mu} \left( 1 + f \frac{\alpha \gamma q^*}{\phi} \right) = q^* + \frac{\alpha \gamma (q^*)^2}{\phi} + \frac{(q^*)^2 f}{\phi} \frac{\partial \alpha}{\partial \mu} \left[ \ln \left( \frac{y_i}{y^*} \right) + \frac{\mu}{\alpha y_i} \right]. \tag{A.10}
\]

From equation (23) we see that \( \text{sign} \left( \frac{\partial N^*/\partial \mu}{} \right) = \text{sign} \left( q^* - \mu (\partial q^*/\partial \mu) \right) \). We use (A.10) and set \( y_i = (\alpha - 1)\mu/\alpha \) to compute \( q^* - \mu (\partial q^*/\partial \mu) \). We obtain:
\[
q^* - \mu \frac{\partial q^*}{\partial \mu} = -\frac{\gamma f \frac{\partial q^*}{\partial \mu}}{1 + f \frac{\alpha \gamma q^*}{\phi}} \left[ \ln \left( \frac{y_i}{y^*} \right) + \frac{\mu}{\alpha y_i} \right]. \tag{A.11}
\]

Denote by \( \omega(\mu, \sigma) = \ln \left( \frac{y_i}{y^*} \right) + \mu/(\alpha y_i) \). As \( \partial \alpha/\partial \mu > 0 \), \( \text{sign}(q^* - \mu (\partial q^*/\partial \mu)) = -\text{sign}(\omega(\mu, \sigma)) \implies \text{sign}(\partial N^*/\partial \mu) = -\text{sign}(\omega(\mu, \sigma)) \).

Next, we study the \( \text{sign}(\omega(\mu, \sigma)) \). From the expression of \( \omega(\mu, \sigma) \) we see that \( \omega(\mu, \sigma) \geq 0 \iff \mu/(\alpha y_i) \geq \ln \left( \frac{y_i}{y^*} \right) \iff \tilde{y}^* \leq \tilde{y} \), where \( \tilde{y} = y_i e^{\mu/(\alpha y_i)} \).

Using the expressions for \( y_i \) and \( \alpha \) from (A.5), we can express \( \tilde{y} \) as a function of the first two moments of the income distribution, \( \mu \) and \( \sigma \):
\[
\tilde{y}(\mu, \sigma) = \mu \frac{z}{z + 1} e^{1/z}, \tag{A.12}
\]

where \( z = \sqrt{1 + \mu^2/\sigma^2} \) and \( e \) is the Euler’s constant.

In order to see if \( \tilde{y}^* \leq \tilde{y} \) holds, we evaluate the \( \text{LHS} \) and \( \text{RHS} \) of equation (19) at \( \tilde{y} \). The \( \text{LHS} \) is increasing in \( \tilde{y} \), while the \( \text{RHS} \) is decreasing in \( \tilde{y} \). Thus, the inequality \( \tilde{y}^* \leq \tilde{y} \) holds if \( \text{LHS}(\tilde{y}(\mu, \sigma)) \geq \text{RHS}(\tilde{y}(\mu, \sigma)) \), or
\[
\frac{\partial q^*}{\partial \mu} \left( 1 + f \frac{\alpha \gamma q^*}{\phi} \right) \geq \frac{1}{1 + \gamma \theta \left[ 1 - e^{-(1+z)/z} \right]} \tag{A.13}
\]

Notice that the inequality implies a restriction in \( \mu \) and \( \sigma \). In the following, we study
the properties of functions $h(\mu, \sigma)$ and $v(\mu, \sigma)$.

\[
\frac{\partial h}{\partial \mu} = \left(1 + \frac{\mu^2}{\sigma^2}\right)^{-1/2} \frac{\mu}{\sigma^2} \left[ \delta \frac{e^{1/2}}{(z + 1)^{1/2}} - \frac{z^{1/2}}{z + 1^{1/2}} \right] \\
= - \left(1 + \frac{\mu^2}{\sigma^2}\right)^{-1/2} \frac{\mu \sigma}{\sigma^2} \delta \frac{e^{1/2}}{(z + 1)^{1/2}} < 0 \tag{A.14}
\]

\[
\frac{\partial v}{\partial \mu} = \frac{e^{-1/(z+1/z)}}{\{1 + \gamma_\theta [1 - e^{-1/(z+1/z)}]\}^2} \left[ 1 + \frac{\mu^2}{\sigma^2}\right]^{-1/2} \frac{\mu}{\sigma^2} > 0. \tag{A.15}
\]

Consequently, $h(\mu)$ is decreasing and $v(\mu)$ is increasing in $\mu \in (0, \infty)$. Both functions are continuous. In addition, $\lim h(\mu) = \delta e / 2$, $\lim v(\mu) = 1/[1 + \gamma_\theta (1 - e^{-2})]$, $\lim h(\mu) = \delta$, and $\lim v(\mu) = 1/[1 + \gamma_\theta [1 - (1/e^2)]]$.

We distinguish three cases:

1) $\lim v(\mu) \geq \lim h(\mu) \iff 1/[1 + \gamma_\theta (1 - e^{-2})] \geq \delta e / 2 \iff \gamma \leq \gamma = [2/(\delta e) - 1] / [\theta (1 - e^{-2})]$; In this case $h(\mu) < v(\mu)$ for any $\mu \in (0, \infty) \implies \hat{y}^* < \hat{y} \implies \omega(\mu) < 0 \implies \partial N^*/\partial \mu > 0$.

2) $\lim v(\mu) \leq \lim h(\mu) \iff \gamma \geq \gamma = [(1/\delta) - 1] / [\theta (1 - (1/e))]$; In this case $h(\mu) > v(\mu)$ for any $\mu \in (0, \infty) \implies \hat{y}^* > \hat{y} \implies \omega(\mu) > 0 \implies \partial N^*/\partial \mu < 0$.

3) $\lim v(\mu) \leq \lim h(\mu)$

   \[ \left\{ \begin{array}{ll}
   \lim v(\mu) & \leq \lim h(\mu) \iff \gamma > \gamma = [(2/(\delta e)) - 1] / [\theta (1 - e^{-2})] \\
   \lim v(\mu) & > \lim h(\mu) \iff \gamma < \gamma = [(1/\delta) - 1] / [\theta (1 - (1/e))] \\
   \end{array} \right. \]

In this case, by the Intermediate Value Theorem, the two functions intersect once in $\hat{\mu} \in (0, \infty)$. There are two subcases here:

3.1) $\mu \in (0, \hat{\mu}) \implies h(\mu) > v(\mu) \implies \hat{y}^* \leq \hat{y} \implies \omega(\mu) \geq 0 \implies \partial N^*/\partial \mu \leq 0$;

3.2) $\mu \in (\hat{\mu}, \infty) \implies h(\mu) < v(\mu) \implies \hat{y}^* > \hat{y} \implies \omega(\mu) < 0 \implies \partial N^*/\partial \mu > 0$.

**Proof of Corollary 1.** We use equation (A.10). As $\partial \alpha / \partial \mu > 0$, if $\omega(\mu, \sigma) > 0$ then $\partial q^*/\partial \mu > 0$. As established in Proposition 2, $\omega(\mu, \sigma) > 0$ when $\gamma \geq \gamma$ or when $\gamma \in (\gamma, \gamma)$. We establish the case when $\gamma \in (\gamma, \gamma)$. As the RHS of equation (A.10) contains some positive terms in addition to $\omega(\mu, \sigma) \implies$ there exists $\hat{\mu} > \mu$ such that $\partial q^*/\partial \mu > 0$ on the interval $\mu \in (0, \hat{\mu})$.

**Proof of Proposition 3.** First, we take the derivative of $f(\sigma) = [y_i(\sigma) / y(\sigma)]^{\alpha(\sigma)}$ with respect to $\sigma$:

\[
\frac{\partial f(\sigma)}{\partial \sigma} = f(\sigma) \left[ \frac{\partial \alpha}{\partial \sigma} \ln \left( \frac{y_i}{y} \right) + \alpha \frac{\partial y_i}{\partial \sigma} \frac{\partial y_i}{\partial \sigma} - \alpha \frac{\partial y_i}{\partial \sigma} \right] \tag{A.16}
\]

\[
= f(\sigma) \left[ \frac{\partial \alpha}{\partial \sigma} \ln \left( \frac{y_i}{y} \right) + \frac{\alpha \partial y_i}{y_i \partial \sigma} - f(\sigma) \frac{\alpha}{y} \frac{\partial y^*}{\partial \sigma} \right]. \tag{A.17}
\]

Next, we calculate $\partial y^*/\partial \sigma = (\partial q^*/\partial \sigma) / \phi \theta \delta$, $\partial \alpha / \partial \sigma = -(\mu^2 / \sigma^2) [1 + (\mu / \sigma)^2]^{-1/2} < 0$, and
\[ \frac{\partial y_t}{\partial \sigma} = (\mu/\alpha^2)(\partial \alpha/\partial \sigma) < 0. \] We use (A.17) in the expression of \((\partial q^*/\partial \sigma)\), (25), and group terms to obtain:

\[
\frac{\partial q^*}{\partial \sigma} \left\{ 1 + \frac{\mu f(\sigma)}{1 + \gamma \theta [1 - f(\sigma)]^2} \frac{\alpha}{\delta} \right\} = \frac{\phi \theta \mu f(\sigma)}{1 + \gamma \theta [1 - f(\sigma)]^2} \frac{\partial \alpha}{\partial \sigma} \left[ \ln \left( \frac{y_t}{y^*_t} \right) + \frac{\mu}{\alpha y_t} \right].
\]

From the expression above we can see that \(\text{sign}(\partial q^*/\partial \sigma) = -\text{sign}(\omega(\mu, \sigma))\). Also, \(\text{sign}(\partial N^*/\partial \sigma) = \text{sign}(\partial \tau^*/\partial \sigma) = \text{sign}(\omega(\mu, \sigma))\).

We studied the properties of the function \(\omega(\mu, \sigma)\) in the proof of Proposition 2. Thus, there are three cases:

1) \(\gamma \leq \gamma = [(2/(\delta e) - 1) / [\theta(1 - e^{-2})] \implies \omega(\mu) < 0 \implies \partial \tau^*/\partial \sigma < 0, \partial N^*/\partial \sigma < 0, \partial q^*/\partial \sigma > 0;\)

2) \(\gamma \geq \gamma = [(1/\delta) - 1] / [\theta (1 - (1/e))] \implies \omega(\mu) > 0 \implies \partial \tau^*/\partial \sigma > 0, \partial N^*/\partial \sigma > 0, \partial q^*/\partial \sigma < 0;\)

3) \(\gamma \in (\gamma, \gamma)\). There are two subcases here:

3.1) \(\mu \in (\bar{\mu}, \mu) \implies \omega(\mu) \geq 0 \implies \partial \tau^*/\partial \sigma \geq 0, \partial N^*/\partial \sigma \geq 0, \partial q^*/\partial \sigma \leq 0;\)

3.2) \(\mu \in (\bar{\mu}, \infty) \implies \omega(\mu) < 0 \implies \partial \tau^*/\partial \sigma < 0, \partial N^*/\partial \sigma < 0, \partial q^*/\partial \sigma > 0.\)

**Proof of Proposition 4.** Denote \(z = (y_t/y^*_t) \in (0, 1]\). Then, the equilibrium enrollment is determined by

\[
y_t = \frac{\mu}{\delta} \frac{1}{1 - z^{\alpha - \nu}} \frac{1 - z^{\alpha - \nu}}{1 + \gamma \theta (1 - z^{\alpha - \nu})}
\]

(A.19)

Denote the left and the right hand sides of (A.19) with LHS and RHS respectively. It is easy to verify that \(\lim_{z \to 0} LHS = +\infty\) and \(\lim_{z \to 1} LHS = y_t\), \(\lim_{z \to 0} RHS = \mu/((\delta(1 + \gamma \theta))\). Using l’Hospital rule, \(\lim_{z \to 1} RHS = \mu(\alpha - \nu)/(\delta \alpha)\). Clearly, \(LHS\) is monotonically decreasing in \(z\). The RHS can be first decreasing and then increasing in \(z\) since

\[
\frac{\partial RHS}{\partial z} > 0 \iff 1 - z^{\alpha - \nu} \left( \frac{1}{1 - z^{\alpha - \nu}} \right) > \frac{\alpha - \nu}{\alpha(1 + \gamma \theta)}
\]

and since \((1 - z^{\alpha - \nu})(1 - \gamma \theta / (1 + \gamma \theta) z^{\alpha - \nu}) / (1 - z^{\alpha}) > 1, \forall z \in (0, 1], a sufficient condition for \(\partial RHS/\partial z > 0\) is \(z > ((\alpha - \nu)/(\alpha(1 + \gamma \theta)))^{1/\nu}\). (i) Thus a sufficient condition for uniqueness is

\[ RHS_{z=0} < LHS_{z=1} \iff \mu/((\delta(1 + \gamma \theta)) < y_t \]

(A.20)

If furthermore \(RHS_{z=1} > LHS_{z=1} \iff \mu(\alpha - \nu)/(\delta \alpha) > y_t \iff \nu < \alpha - (\alpha - 1)\delta/\alpha\), the equilibrium enrollment is interior, otherwise \(z = 1(y^*_t = y_t)\). Using the definition of \(\delta\) and
(A.5) in (A.20) and solving for $\gamma$ results in $\gamma > (\alpha/(\alpha - 1)(1 - \theta)^{(1/\theta - 1)} - 1)/\theta > 0$. Thus, if household’s concern for children is high enough, there is a unique equilibrium threshold for private enrollment.

(ii) Intuitively, as $\nu$ goes to zero, the problem is reduced to the benchmark, which has a unique equilibrium. Since $\partial LHS/\partial z < 0$, imposing $\partial RHS/\partial z > 0$ guarantees uniqueness. This condition can be further rewritten as

$$(1 - z^{\alpha - \nu}) \left[ \alpha z^{\nu} (1 + \gamma (1 - z^{\alpha - \nu})) + \gamma (1 - z^\alpha) (\alpha - \nu) \right] > (\alpha - \nu)(1 - z^\alpha) (1 + \gamma (1 - z^{\alpha - \nu})).$$

The inequality holds for any $z < 1$ as $\nu \to 0$.

8 Appendix B

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<th>Table 5: Summary statistics</th>
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The data on local TELs has been taken from Yuan et al. (2007). Three criteria have been used to separate the states where local politics have more bearing on school finance. First, to construct sample No-TELs 1, we have retained states that i) do not have state mandated local tax rate limits and revenue/expenditure limitations or, ii) if they have any, they can be changed through a local vote. Second, to construct sample No-TELs 2, we also remove states that have annual property assessment increase limits under 5%. Relative to the benchmark sample, No-TELs 1 drops AR, CA, IN, IA, NM, NY, OH, PA, UT, WY. Sample No-TELs 2 also excludes CO, FL, GA and OR.