VOTING IN A LOCAL SCHOOL ELECTION: A MICRO ANALYSIS

Daniel L. Rubinfeld*

IN recent years empirical studies of local school finance have relied to a large extent on the median voter and related models, tested with data aggregated to the precinct, school district, or local level.[1] While there are advantages to using aggregated data, the limited availability of data on the distribution of income, property tax payments, and other variables, as well as the possibility of bias associated with the grouping of households into aggregated units, suggest some important disadvantages.[2]

This paper attempts to analyze the demand for local public education using individual household data obtained through a survey of voters in two local school elections in a Detroit suburb.[3] In section I a model of voting in a school election is presented. The model assumes that individual voters determine their desired level of educational expenditures per pupil by maximizing a utility function subject to a budget constraint. Individuals decide whether to vote for or against a given millage request by comparing their desired expenditure level with the actual and proposed levels. On the basis of some assumptions concerning the stochastic nature of the individual utility functions, our analysis suggests that the probability of a yes or no vote can be estimated using a binary logit form.[4] In section II the model variables and estimates of the model parameters are presented and discussed. The estimation results are interpreted in the context of the voting model presented in section I and are compared to the results of several educational expenditure studies. In section III the outcome of the two local elections is analyzed, with an attempt made to explain the passage of the second election, in light of the failure of the first. In particular, the model is used to test the reaction of voters to the state circuit-breaker legislation which was enacted after the first election. Some concluding remarks are presented in the final section. Finally, some tests for the presence of bias in the survey responses are described in the appendix.

I. The Model

Assume that each individual within a school district maximizes a utility function

$$U(c, h, E; X, \epsilon) \tag{1}$$

where c represents units of non-housing consumption, h represents units of housing services consumed, E represents dollars of educational

Received for publication September 3, 1975. Revision accepted for publication February 26, 1976.

* The election survey was stimulated by, and planned jointly with, George Peterson of the Urban Institute, and was carried out with the financial support of the Urban Institute and the Institute of Public Policy Studies. The Institute of Public Policy Studies and the National Bureau of Economic Research provided research support for the analysis of the voting data. I am indebted to Sue Neal for her valuable research assistance, which was financed by the Michigan Research Seminar on Quantitative Economics. Paul G. Liddicoat, the Superintendent of Schools in Troy, Michigan was extremely cooperative in providing the lists of registered voters. In addition, Sue Neal, Edward Porter and Kemp Harshman assisted with the telephone follow-up of the mail questionnaire. Finally, William Apgar, Robin Barlow, Harvey Brazer, Paul Courant, Martin Feldstein, James Morgan, and A. Mitchell Polinsky provided helpful comments at various stages of the research.

[1] See, for example, Barlow (1970), Bass and Alexander (1974), Bradford and Oates (1974), Feldstein (1975), and Peterson (1973). For the use of aggregated data to analyze the demand for non-educational public goods, see, for example, Bergstrom and Goodman (1973), Birdsall (1965), Borcherding (1972), Deacon and Shapiro (1975), and Shapiro (1973).

[2] Even if the appropriate unit of observation is the individual household, the use of grouped data does not necessarily bias the estimated parameters. See, for example, the references in Theil (1971, p. 249) or Blalock (1961, p. 97).

[3] During the fall of 1973 a mail questionnaire (with a telephone follow-up) was sent to a sample of registered voters of Troy, Michigan. The questionnaire focused on the voting behavior of individuals in two local school millage elections, the first early in May 1973 and the second in June 1973. In May a proposal to renew 3.80 mills of school taxes for two years was rejected, but in June the same proposal was approved. Two samples were chosen; the first a sample of 833 individuals who voted in the June election, selected randomly within each precinct, and the second a sample of 487 individuals who did not vote in the June election, selected from the voter registration lists. Further details about the sampling process as well as the results of both elections are available from the author upon request.

[4] The model developed in the text focuses only on the yes-no voting choice for homeowners. Thus, the paper does not deal with the vote-no option or the voting choice of renters. Research relating to both renters and non-voters is empirically more difficult because of the low response rates involved.

expenditures per pupil, X is a vector of non-income attributes of individuals and their households, and ϵ is a random unobserved error.[5] The random error term accounts for the fact that the vector X does not include all relevant taste factors, as well as the fact that those included attributes may not serve as adequate proxies for taste. As a result, it is useful to view the value of the utility function as a random variable whose realization is dependent upon the specific individual sampled from the population of all individuals with attribute vector X.

Individuals are subject to the following household budget constraint:

$$Y = c + ph + t(phD) = c + p(1 + tD)h \quad (2)$$

where p is an annual rental price per unit of housing services (assumed constant across all housing units in the school district), t is the ad valorem effective school tax rate, D is a constant defined so that phD equals the capitalized value of the house, and Y is household income after federal and state taxes and including the imputed rental income of housing services (the price of the non-housing commodity is assumed to be 1). Finally, we assume that the budget of the local school board is in balance, i.e.,

$$E = tV \quad (3)$$

where V is the value of taxable property per pupil in the school district.[6]

By maximizing the utility function (1) subject to the budget constraint (2) and substituting the budget balance equation (3) we can solve to obtain the demand functions for the non-housing commodity and education, each as a function of income and prices. For our purposes it will be instructive to write the demand for educational expenditures per pupil as a function of income and the price (per dollar) of public schooling, conditional on the individual attributes and a random component:

$$E = g(Y, phD/V; X, \epsilon). \quad (4)$$

The price of public schooling, phD/V, represents the amount of local taxes which the household must pay if the school budget per pupil is raised \$1.00. Of particular interest is the fact that the price of schooling is independent of the school tax rate under a proportional property tax system. To see this, note that the incremental community tax rate necessary to finance an additional \$1 of spending per pupil is equal to $1/V$, the reciprocal of the value of taxable property per pupil. Then the price of schooling equals the incremental taxes paid by the household, which in turn equal the tax rate $(1/V)$ times the value of the house (phD). Thus, the price of schooling varies among households because of the variation in house value. In our sample, for example, the price of schooling ranges from 19¢ to \$2.38, and has a mean of \$1.00.

Implicit in our model is the assumption that the price of schooling remains unchanged (or changes an insignificant amount) regardless of the outcome of the local referendum. This is equivalent to assuming that (a) individual housing consumption remains fixed, i.e., individuals do not move or rehabilitate their existing house; (b) the price of housing remains unchanged, i.e., housing units are not reassessed;[7] and (c) the per pupil value of taxable property remains unchanged.[8] To the extent that any changes that occur in the variables p, h, and V are small or unforeseen by voters, the model is likely to provide a reasonable approximation to reality.

To reiterate our assumption that the demand for educational expenditures per pupil is a random variable conditional upon the vector of household and individual attributes, we express each individual's desired level of expenditures per pupil as

[5] We are assuming that all individuals have identical utility functions, but that the value of utility obtained by the maximization process is conditional on the attributes of the individual households. To simplify the analysis (without loss of generality) we have imbedded the choice among private goods and non-educational public services into the non-housing commodity c.

[6] V may include nonresidential as well as residential property. In addition, equation (3) can be generalized to include the effects of non-matching and matching state and federal aid to education. For details, see Edelson (1974) and Peterson (1973).

[7] Thus, the price of housing does not appear explicitly in equation (4) because the price is constant for all families in the school district and assumed independent of the outcome of the election.

[8] As a result of these assumptions our model rules out the real possibility that individuals will vote for a millage proposal solely on the grounds that it is likely to raise the value of their own property, or to lower the cost of the public services by raising the value of taxable property in the community.

$$E_i^* = E^*(Z) + u_i \quad (5)$$

where $Z = (X, Y, phD/V)$ is the vector of individual attributes, household income, and the price of schooling, and u_i is a random variable.

Within a single community, all households consume the same level of expenditures per pupil. In any given millage election, they are limited to a choice between one of two levels of educational expenditures per pupil. We will denote the original level by E_o and the proposed level by E_p.[9] Under the presumption that each individual must vote yes or no, we expect the voting choice to involve a comparison of the desired level of expenditures per pupil with the actual and proposed levels. When the desired level is higher (lower) than both E_o and E_p, we expect the individual to vote yes (no). When the desired level lies between E_o and E_p, the vote may be either yes or no.

To make the voting decision explicit, and to clarify our analysis when the desired level lies between E_o and E_p, we assume that for each individual there exists a cut-off level of expenditures per pupil (E_i') that determines the point at which the difference between the desired and actual expenditure levels is sufficient to generate a yes vote. We will assume that the cut-off level varies randomly across individuals, so that we may write $E_i' = E' + v_i$, where v_i is a random variable and E' is between E_o and E_p.[10] Then, the specific voting rule is

vote yes if $E_i^* \geqslant E_i'$

vote no if $E_i^* < E_i'$.

Under these assumptions we may represent the probability that an individual will vote yes as follows:

$$\begin{aligned} \text{prob (yes)} &= \text{prob}(E_i^* \geqslant E_i') \\ &= \text{prob}(v_i - u_i \leqslant E^*(Z) - E'). \end{aligned} \quad (6)$$

Notice that for each individual $v_i - u_i$ is a random variable, while $E^*(Z) - E'$ is fixed. The expression in (6) makes it clear that we may express the probability of an individual voting yes in terms of the cumulative distribution function associated with the random variable $v_i - u_i$. We will assume that the probability density function is logistic,[11] from which it follows that

$$\text{prob (yes)} = 1/(1 + e^{-(E^*(Z)-E')}) \quad (7)$$

or

$$\ln\left[\frac{\text{prob (yes)}}{1 - \text{prob (yes)}}\right] = E^*(Z) - E'. \quad (8)$$

Under the assumption that $E^*(Z)$ is a linear function of the vector Z (for example), our model suggests the use of logit estimation to obtain parameter estimates in the following equation:[12]

$$\ln\left[\frac{\text{prob (yes)}}{1 - \text{prob (yes)}}\right] = Z\beta. \quad (9)$$

We will use equation (9) to estimate the parameters associated with the voting model in the following section.

II. Model Estimation

The voting model was estimated using data collected from a 1973 survey of eligible voters of Troy, Michigan. The responses to the survey provided us with a list of attributes of voters as well as estimates of household income and the price of public schooling. The variables used in the estimation process are listed in table 1.

The estimated logit equation associated with the May election appears below, with asymptot-

[9] There will, of course, be a unique tax rate associated with each expenditure choice. We will implicitly assume that the tax rate differential is sufficiently small so that the passage of the millage would have an insignificant effect on the number of children sent to private school. If this were not the case, the price of schooling might not be independent of the tax rate.

[10] The analysis still holds if v has zero variance, but is substantially complicated if v is correlated with the individual attributes.

[11] The logistic function is chosen primarily because the properties of the estimation procedure are more desirable than those associated with the choice of a uniform distribution (leading to the specification of a linear probability model) or a normal probability distribution (resulting in a probit model). For details concerning the logistic specification, see Cox (1970) and Nerlove and Press (1973). The logistic specification follows directly from the assumption that u and v are statistically independent and follow a Weibull distribution. See McFadden (1973) for details.

[12] The constant term will pick up the effect of E', while the variance associated with the logistic specification will be incorporated in the estimates of β. For a given election there is no way of estimating this variance, so that the logit parameter estimates can only be determined up to a scalar multiple.

TABLE 1. — DEFINITION OF VARIABLES

SEX	= 1 if female, 0 if male
MAR	= 1 if married with spouse present, 0 otherwise
OTHER	= 1 if separated, divorced or widowed, 0 otherwise
*A*35–49	= 1 if aged 39 to 49, 0 otherwise
*A*50–64	= 1 if aged 50 to 64, 0 otherwise
*A*65+	= 1 if aged 65 or over, 0 otherwise
*PUB*1	= 1 if 1 child is in public school, 0 otherwise
*PUB*2	= 1 if 2 children are in public school, 0 otherwise
*PUB*3	= 1 if 3 children are in public school, 0 otherwise
*PUB*4	= 1 if 4 children are in public school, 0 otherwise
*PUB*5	= 1 if 5 or more children are in public school, 0 otherwise
PRIV	= 1 if the family has 1 or more children in private school, 0 otherwise
YEARS	= number of years living in Troy
SCHOOL	= 1 if individual is employed as a teacher (public or private), 0 otherwise
ln (*INC*)	= natural logarithm of annual household income in dollars
ln (*PRICE*)	= natural logarithm of price of public schooling in dollars

Note: *INC* was constructed by assigning to each individual category the midpoint of the range of possible response. The endpoints were set arbitrarily. Some experiments suggested that the results were not very sensitive to these assumptions. For this reason no attempt was made to obtain a continuous income variable by fitting the data to (for example) a Pareto income distribution. The variable *PRICE* was created from the survey response in which voters estimated their total property tax payments. *PRICE* was set equal to a continuous variable describing property taxes paid (determined in a manner analogous to the creation of the variable *INC*) divided by the product of the effective tax rate times school revenue per pupil. The logit model was also estimated using dummy variables to represent income and price, but the results are not included (details are available from the author).

We included the sex dummy to allow for the possibility that women and men might perceive the benefits and costs of educational expenditures differently. Our expectation was that because women tend to bear a large share of the responsibility for child care, they would value the potential benefits associated with the educational system more highly than men. Whether women or men are more cognizant of the tax costs associated with a favorable referendum outcome seemed less clear. The sex dummy coefficient was insignificant here, but was significant as expected in the June election (reported later). As will be discussed in the next section, the change in significance of *SEX* in the voting model helps to explain why the June election passed when the May election failed by a substantial margin.

We included marital status dummies on the premise that family structure would have an important effect on voting behavior. Relative to single individuals (without children) we expected that individuals from households with spouse present would be more directly aware of the potential benefits of education whether or not school age children were present, and therefore would be more likely to favor the referenda. The relationship between single individuals and those married without spouse present is more

ic standard errors in parentheses (* = significant at the 5% level):[13]

$$\begin{array}{lllllll}
-23.15^{*} & +\ .24(SEX) & +\ 1.13(MAR) & +\ 1.09(OTHER) & +\ .08(A35\text{–}49) & +\ .61(A50\text{–}64) & \\
(3.84) & (.24) & (1.13) & (1.47) & (.30) & (.41) & \\
+\ 1.04(A65) & +\ 1.44^{*}(PUB1) & +\ 1.39^{*}(PUB2) & +\ 1.30^{*}(PUB3) & +\ 2.00^{*}(PUB4) & +\ 2.16^{*}(PUB5) & \\
(.79) & (.34) & (.35) & (.42) & (.58) & (.79) & \\
-\ .56(PRIV) & -\ .02^{*}(YEARS) & +\ 3.07^{*}(SCHOOL) & +\ 2.14^{*}(\ln(INC)) & -\ 1.21^{*}(\ln(PRICE)) & & \\
(.42) & (.01) & (.84) & (.37) & (.44) & &
\end{array}$$

degrees of freedom = 408; chi-square = 156.2 *.

[13] The model was estimated using a maximum likelihood estimation procedure supplied by Charles Manski and a comparable procedure developed by Forrest Nelson for the National Bureau of Economic Research. Laxmi Rao of the NBER provided valuable research assistance with the NBER estimation package.

difficult to predict, and was unlikely to be statistically important in any case because of the limited number of respondents in the latter category. In fact, the marital status dummy coefficients were all insignificant (at the 5% level), although their signs were consistent with our a priori expectations. While insignificant, the numerical values of the parameters associated with the *OTHER* dummy suggest that those married without spouse present were almost as likely to vote favorably in the election as those married with spouse present.[14]

The effect of the age dummies is difficult to predict when income, number of children in school and years in the community are held constant. Since we did not control for the total number of children in the household, our expectation was that as age increases in the 18-49 range, the probability that an individual will have school age children in the future declines, and therefore the probability of voting yes should also decline. However, the results show that the age dummy coefficients were insignificant, generally with a positive sign.[15]

The number of children in public school is perhaps the best measure for the direct benefits that households perceive from public schooling. When children are of school age, households are most likely to be aware of the costs and benefits associated with a vote for higher school taxes. Thus, we expected the presence of at least one child in public school to have a substantial positive impact on the probability of a yes vote. The presence of additional school-age children should also increase the probability of a yes vote, although we expected that, beyond a certain point, the marginal gain of reallocating the household budget towards private expenditures would outweigh the gain from public expenditures and the probability of a yes vote would decline. The presence of children in private school should have a strong negative effect, however. Families sending their children to private school are likely to perceive little benefit from the public school system, while facing a substantial tax bill associated with the public school system.[16]

Each of the dummy variable coefficients representing public school children was highly significant. Our results suggest that the presence of the first public school child raises substantially the probability that the individual will vote yes in the election. After the first child, the probability of a yes vote remained roughly constant until the fourth and fifth school children were present, at which point it increased substantially. Also, the May election results suggest that in Troy the demand for educational expenditures per child actually levels off when the number of children grows to four or more, not an unreasonable result, given the large demands for private goods which arise with increased family size.

The private school dummy coefficient was negative but insignificant in the May election. This suggests that the presence of children in private school does increase the probability of a no vote, but the magnitude of this effect is small relative to the effect associated with the presence of at least one child in public school.

We have also included the number of years in residence as an explanatory variable in the logit model. Several potentially conflicting motives might be measured by the residence variable. On one hand, as time of residency increases, voters are more likely to feel a part of the community, and, other things equal, to vote for expenditures which are deemed to be socially beneficial. On the other hand, as time of residency increases, voters become more aware of the virtues and deficiencies of the schools as well as the overall burden of the tax system. Whether these motives will lead to a no vote in protest or a yes vote for improved quality is difficult to say a priori. Our results suggest that, to the extent that we have been able to keep other things (such as life-cycle stage) constant, as the time of residence increases voters tend to vote no, either in criticism of the educational

[14] Using a likelihood ratio test, we could not reject the null hypothesis that the *MAR* and *OTHER* coefficients are identical (at the 5% level).

[15] The age coefficients remain insignificant even when the *YEARS* available is dropped from the model.

[16] There are positive benefits for homeowners without children in public school associated with increased public expenditures, if the expenditures are capitalized into property values. See Oates (1972) for some empirical evidence on this subject. For a more general discussion of the importance of allowing for the private school option when estimating educational income and price elasticities, see Barzel (1973) and Edelson (1973).

system, or possibly in opposition to the growing burden of local taxes.

The highly significant school dummy was included to account for the fact that the sample of respondents is overrepresented by school teachers and their spouses. As expected, school teachers are more likely to vote yes in the election relative to individuals with similar non-occupational attributes.[17]

The income variable serves as a measure of the capacity of households to consume both private and public goods. On the assumption that local school education is a normal good, we expected, other things equal, that income and the demand for public schools would be positively correlated. In the context of our voting model, this suggests a positive relationship between income and the probability of a yes vote. In the estimated equations the income variable was positive and significant, consistent with a positive income elasticity of demand for education.[18]

As the price of schooling rises, other things equal, we expected that the quantity of educational expenditures per pupil demanded would fall, as would the probability of voting yes in the election. Despite the fact that property tax payments are positively correlated with income, we found that the coefficient of the price of schooling variable was negative and significant. This result is consistent with a negative price elasticity of the demand for education.

It is natural to ask what the coefficients of the income and price variables tell us about the actual magnitudes of the income and price elasticities of the demand for educational expenditures. Unfortunately, however, there is not sufficient information in the existing data set for separate income and price elasticities to be calculated. Only if the study were to analyze situations in which voters from the same community faced two or more distinct expenditure choices would such a calculation be potentially feasible. To see why this is true, note that the elasticities calculated directly from the logit estimation are the elasticities of the logarithm of the odds of voting yes with respect to income and price, respectively (hereafter called the odds income and price elasticities). Since the variance of the logistic distribution cannot be determined, the estimated parameters of the logit voting model are known only up to a linear transformation. As a result, only the ratio of the two odds elasticities has a cardinal interpretation. Fortunately, however, the ratio of the odds income and price elasticities is equal to the ratio of the expenditures per pupil income and price elasticities.[19] For this reason our survey results can give some interesting insights into issues which relate to the relative magnitude of the educational income and price elasticities.

We will proceed by describing the odds income and price elasticities obtained using several definitions of income and price and a logarithmic form for both variables.[20] This makes it possible to examine the sensitivity of the magnitude of the income elasticity of demand for education to the particular definitions used and to reach some tentative conclusions concerning the relative magnitudes of the biases in the price and income elasticities of demand for education obtained from aggregated studies. Finally, we show how our estimates of the ratio of the income and price elasticities may be used to obtain information about the issue of whether local public education is efficiently supplied

[17] A detailed explanation of the response of school teachers is presented in the appendix.

[18] These results are also consistent with the analysis of perceived educational benefits given by Neenan (1972).

[19] To see why this is true, note that the application of the cumulative logistic distribution in equation (6) maps expenditure levels into probabilities of voting yes in the election. Since the probability of voting yes is monotonically related to the log of the odds of voting yes, we may represent the relationship between the log of the odds and expenditures per pupil as $R = \ln[\text{prob (yes)}/(1 - \text{prob (yes)})] = F(E)$, where F is a monotonic and continuously differentiable function. Then by substituting in equation (9) and applying the chain rule of differentiation we find that (Y = income, P = price):

$$\partial E/\partial Y = [\partial F^{-1}(R)/\partial Z\beta]\beta_Y$$

and

$$\partial E/\partial P = [\partial F^{-1}(R)/\partial Z\beta]\beta_P.$$

Then,

$$\frac{\partial E}{\partial Y}\frac{Y}{E}\Big/\frac{\partial E}{\partial P}\frac{P}{E} = \frac{\partial R}{\partial Y}\frac{Y}{R}\Big/\frac{\partial R}{\partial P}\frac{P}{R} = (\beta_Y/\beta_P)(Y/P).$$

With the logit model specified with the attributes in logarithmic form, the ratio of the elasticities equals (β_Y/β_P) and thus is independent of the attribute values.

[20] We also estimated the model using linear forms for the income and price variables, but do not present the results here for convenience. The inclusion of logarithmic income and price terms resulted in a better fit than the inclusion of linear forms of the variables.

TABLE 2. — RELATIVE INCOME AND PRICE ELASTICITIES

Equation	Income Variable	Income Elasticity	Price Variable	Price Elasticity	Ratio
1	ln (*INC*)	1.00	ln (*PRICE*)	—.56	—1.79
2	ln (*INC*)	.96	ln (*PADJ*)	—.56	—1.71
3	ln (*PINC*)	1.12	ln (*PRICE*)	—.69	—1.62
4	ln (*PINC*)	1.06	ln (*PADJ*)	—.65	—1.63

Notes: 1.) All elasticities are determined only up to a scalar multiple. 2.) The income elasticity associated with the first set of results is normalized to equal 1.00. 3.) *PINC* is the sum of household income and 6% of the estimated value of the individual's house. 4.) *PADJ* is price adjusted to account for the federal deductibility of local property taxes. 5.) When logit results were obtained using linear income and price variables, the estimated elasticities about the mean tended to be lower for both income and price. The ratio of the elasticities (income to price) tended to be somewhat higher than the results printed above.

when financed under our existing system of property taxation.

Table 2 lists the relative income and price elasticities obtained from the May election data.[21] To make the results most easily comparable we have arbitrarily normalized the elasticities by setting the income elasticity measured in the first equation equal to one. A comparison of equations 1 and 2 allows us to ascertain the changes which occur when the price variable is adjusted to account for the deductibility of property taxes under the federal income tax. Our results suggest that failure to account for property tax deductibility causes the estimated income elasticity of demand to be biased upward. A comparison of equations 1 and 3 allows us to study the effect of adding the imputed rental income of housing to household income. The results suggest that failure to account for imputed rental income when constructing an income variable tends to bias the income and price elasticities of the demand for educational expenditures downward. The income elasticity result is consistent with our expectation that permanent income elasticities are higher than elasticities calculated from annual income.[22]

It is interesting to compare the information contained in our relative estimates of income and price elasticity to estimates obtained from aggregated studies. While aggregated data allow for the explicit calculation of the magnitudes of the income and price elasticities, the nature of the aggregation process makes it likely that the estimates of elasticity will be biased. The bias arises because the aggregated educational expenditure equations are usually misspecified, either by omitting variables such as children in public and private school and marital status, or by including the incorrect price term. To consider the nature of this specification bias we might examine the results of some aggregated studies as summarized in table 3.[23]

TABLE 3. — INCOME AND PRICE ELASTICITIES FROM AGGREGATED STUDIES

		Income Elasticity	Price Elasticity	Ratio
I.	Barlow (Michigan)	.64	—.34	—1.91
II.	Bradford-Oates (New Jersey)	.65	—.36	—1.81
III.	Peterson (Michigan)	1.11	—.60	—1.85

To focus on what may be the most important source of specification bias, consider the correct price term under the assumption that each individual in a community lives in an identically valued house. In this case the correct price term in the aggregated model is simply $phD/V = (n_s/n)\ (R/V)$ where n_s is the number of public school students, n is the number of families and R is the aggregate residential tax base per public school pupil.[24] The price term using the study by Barlow is roughly equivalent to the term (R/V) and thus does not account for the variation across communities in the variable (n_s/n). Bradford-Oates, on the other hand, use (n_s/n) as a price variable, but do not account for the variation in (R/V). Peterson uses the correct price term, but is unable to account for other potential sources of specification bias due to data limitations. Peterson's result suggests that Barlow and Bradford-Oates underestimate price and income elasticities of demand for education in such a way that the

[21] We have chosen to focus on the May results because of difficulties associated with the price term in the June election. These difficulties are discussed in section III.

[22] This result follows because random measurement errors associated with the income variable will cause the income elasticity to be underestimated.

[23] Some of the specification problems associated with these elasticities are discussed in Peterson (1973). For an analysis of the relationship between specification error and estimation using grouped data in the context of housing demand, see Polinsky (forthcoming).

[24] To simplify our discussion we have not taken into account the fact that the price of schooling may vary across communities because the matching rate on state aid varies. This variation is used by Feldstein (1975) to estimate price elasticities of demand for Massachusetts communities. However, Feldstein's results are not directly comparable here because wealth (tax base) per pupil is used as an explanatory variable rather than income.

ratios of the elasticities remain approximately the same. However, our results in table 3 suggest that the aggregated studies tend to overestimate the magnitude of the ratio of the income and price elasticities.

Another useful application of our estimates of the ratio of income and price elasticities is the determination of whether the property tax leads to an efficient level of school expenditures. In his 1970 study of school expenditures in Michigan, Barlow calculates (for each of seven income groups) the marginal benefit/burden ratio-at-efficiency, which equals the ratio between voters' marginal benefit and marginal tax burden at the efficient level of output. He then argues that a tax structure will lead to the efficient level of output if the crucial voter (the one who is just above the median in the ranking of voters by benefit/burden ratios-at-efficiency) possesses a marginal benefit/burden ratio-at-efficiency equal to unity. After calculating a benefit/burden ratio of 0.5, Barlow concludes (tentatively) that on average the property tax leads to an inadequate level of school expenditures per pupil.

One of the crucial data requirements of Barlow's study is the ratio of income and price elasticities. To see how sensitive Barlow's calculations are to the specification problems associated with his estimated income and price elasticities, we recalculated his results utilizing the information contained in equation 4 of table 2. Since we view this exercise solely as an illustration of the usefulness of the Troy voting results, we have not attempted to adjust for other problems inherent in the Barlow study.[25] We find that the crucial voter has a benefit/burden ratio-at-efficiency of 0.57.[26] This does not alter Barlow's conclusion about the inefficiency of property taxation, but does lower somewhat the statistical confidence associated with his result.

III. The Passage of the June Millage Proposal

Despite the fact that voters faced identical referenda, the results of the May and June elections were markedly different. In May the proposal failed, receiving only 37.5% of the vote, while in June the proposal barely passed, receiving 51.9% of the vote.

TABLE 4. — ANALYSIS OF VOTING CHANGES

Mean Values of Variables for Selected Voting Strata	*INC*	*PUBLIC*	*SEX*	*YEARS*
Homeowners not voting in May, but voting yes in June ($N = 42$)	23,631	1.40	.81	4.81
Homeowners voting in both elections ($N = 395$)	20,605	1.42	.58	11.57
Homeowners who switched from no to yes ($N = 19$)	20,789	1.89	.79	11.74
All homeowners ($N = 540$)	20,572	1.69	.59	10.44

Note: *INC* = household income; *PUBLIC* = number of children in public school; *SEX* = 1 if female, 0 if male; *YEARS* = years of residence.

Perhaps the best overall picture of the voting change that occurred between the May and June elections can be seen by examining table 4. The table makes it clear that the passage of the June election was brought about largely by an influx of voters to the polls who had not voted in the earlier election. The increased turnout was clearly more important in the final determination than the switching of votes from no to yes. Those who did switch their votes were not markedly different from the entire sample in terms of income, number of school children and years lived in the community. However, those turning out for the June election (but not in May) were distinguishable from the average voter in the sample. New turnouts who voted yes in June tended to be women, with higher household income, but with fewer children in public schools. In addition, they were recent residents of Troy, having lived in the community for less than five years on average.

An additional important change which occurred between the May and June elections was the passage of circuit-breaker legislation by the state of Michigan. Essentially this legislation provided income tax relief to individuals whose property tax payments exceeded a specified fraction of their income.[27] Coupled with the fact

[25] See Edelson (1973) and Barzel (1973) for additional discussion of the Barlow paper.

[26] Our calculations relate to table 1 of Barlow (1970, p. 1032). We have taken the information in columns (1)-(4) and (7) as given and have recalculated columns (5) and (6) using our price and income elasticity information.

[27] The Property Tax Relief Act of 1973 provides that homeowners and renters who are not eligible for special relief as senior citizens, eligible veterans or blind persons, are entitled to a credit equal to 60% of the amount by which their

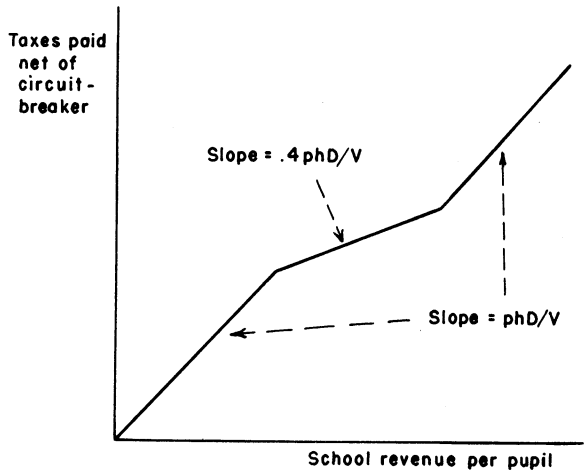
that the Superintendent of Schools placed a heavy emphasis on tax issues while campaigning for the June millage, this suggested that we might use the voting data to test the responsiveness of voters to the presence of the circuit-breaker.[28]

When the circuit-breaker is in effect the analysis increases in complexity, since the price of schooling is no longer independent of the tax rate. The correlation between the tax rate and the price of schooling can be seen by examining figure 1. The graph depicts the relationship between school revenues per pupil and property taxes paid net of circuit-breaker relief for a family that is eligible for a 60% credit on property taxes that exceed 3.5% of its income, subject to a maximum credit of $500. For those not eligible for property tax relief and those receiving the maximum credit of $500, the price of schooling remains equal to phD/V. However, for those eligible for the tax credit and not receiving the maximum benefit, the price of schooling falls to 0.4 phD/V on the margin. Those receiving a $500 credit face the original price (on the margin) of phD/V, since they treat the credit as a lump sum income transfer. It follows that the correctly measured price of schooling will depend upon the tax rate, which has the effect of introducing a nonlinear budget constraint into our analysis.

Since we see no convenient way of introducing the nonlinear constraint into our analysis, we have chosen to utilize two proxies for the correct price term to test whether the presence of the circuit-breaker legislation is able to explain the passage of the June millage proposal. First, the June logit model was reestimated using a "marginal price of schooling" (equal to 0.4 phD/V in figure 1).[29] This assumes that voters respond to the tax cost associated with an incremental change in the school budget. We expected this price term to overestimate the effect of the circuit-breaker. Second, the June model was estimated using an "average price of schooling," which assumes that voters respond to the average tax cost associated with the total proposed budget. We expected this price term to underestimate the effect of the circuit-breaker.[30]

The estimated logit equation for the June election, using the logarithm of the average price of schooling ($\ln(APRICE)$) as an explanatory variable, appears below, with asymptotic

FIGURE 1. — THE CIRCUIT-BREAKER AND THE PRICE OF SCHOOLING



Notes: School revenue per pupil $= tV$, where $t =$ school tax rate.

Taxes paid net of circuit breaker

$$= \begin{cases} tphD \text{ if } tphD \leqslant .035INC \\ tphD - .6(tphD - .035INC) \\ tphD - CRED \text{ if } tphD > .035INC \text{ and } CRED \leqslant \$500 \\ tphD - \$500 \text{ if } tphD > .035INC \text{ and } CRED > \$500. \end{cases}$$

property taxes exceed 3.5% of their household income, with a maximum credit of $500. For senior citizens, the percentage of income falls from 3.5% to 0% as income falls below $3,000, and the reimbursable rate is 100%. According to the Advisory Commission on Intergovernmental Relations (1975), the average benefit per capita in 1974 (of those receiving benefits) was $159.25, while their per capita cost was $14.26.

[28] The campaign made extensive use of local tax worksheets, which emphasized not only the availability of tax credits under the circuit-breaker legislation, but also the deductibility of property taxes under the federal income tax and the savings available to taxpayers under recent reforms in the state income tax. We have no way of knowing to what extent strategic considerations concerning the timing and size of the proposed millages may have affected the election outcomes.

[29] For senior citizens who are eligible, the marginal price of schooling is equal to zero. The mean of the marginal price variable was $.54.

[30] The "average price of schooling" was set equal to $(phD - CRED/t)/V$ where $CRED$ is the income tax credit. Its mean was equal to $.83.

standard errors in parentheses (* = significant at the 5% level):

$$
\begin{aligned}
&-\underset{(5.52)}{35.17^*} + \underset{(.26)}{.71^*}(SEX) + \underset{(1.25)}{1.52}(MAR) + \underset{(1.61)}{1.06}(OTHER) + \underset{(.30)}{.02}(A35\text{–}49) + \underset{(.41)}{.81}(A50\text{–}64)\\
&+ \underset{(.86)}{.67}(A65+) + \underset{(.35)}{1.92^*}(PUB1) + \underset{(.36)}{1.78^*}(PUB2) + \underset{(.47)}{2.18^*}(PUB3) + \underset{(.62)}{2.44^*}(PUB4) + \underset{(.72)}{1.82^*}(PUB5)\\
&- \underset{(.43)}{.63}(PRIV) - \underset{(.01)}{.03^*}(YEARS) + \underset{(1.24)}{3.99^*}(SCHOOL) + \underset{(.53)}{3.26^*}(\ln(INC)) - \underset{(.63)}{2.67^*}(\ln(APRICE))
\end{aligned}
$$

degrees of freedom = 430; chi-square = 205.6*.

A comparison of these results to the results of the May election presented earlier shows that the *SEX* variable, which was insignificant in May, is now strongly significant. In addition, the significance of the *YEARS* variable has increased from the previous election. Most of the other parameters have increased in magnitude as well as in significance, which can be explained in part by the increased sample used to estimate the June logit model.[31] This statistical evidence is supported by a careful examination of the responses to several open-ended questions by those who switched votes and those who voted yes in June, but did not vote in May. Of those switching votes, a majority expressed a concern for the reduction in school quality which might result if the millage were not passed. A minority of those responding cited a reduced fear of interdistrict bussing as their reason for supporting the referendum.[32] None cited tax reasons as the primary cause of a voting switch. Finally, of those voting in June, but not in May, a substantial majority expressed a concern for the quality of schools, which was strengthened as a result of the failure of the first election, and by the advertising of the school system.[33]

When the logit model was reestimated using the marginal price of schooling variable (unlogged), the price term was negative, but insignificant. In addition, the explanatory power of the logit model fell in both cases in comparison to the model which was estimated using an unadjusted price term. To pursue the matter further we examined the tax credit eligibility associated with various subsamples of June voters. According to our admittedly crude calculations, 83% of those voting in June who owned homes and paid taxes were eligible for some tax credit (the mean credit was $192). Of those who voted yes in June, but did not vote in May, 78% were eligible by our calculations to receive some form of credit (the mean was $202). Finally, of those switching their votes from no in May to yes in June, only 63% were eligible for the credit (the mean was $137).

[31] The fact that the estimated parameter vector for the June election is not a scalar multiple of the May estimated parameter vector suggests that we ought to consider whether the voting relationship has remained stable between the two elections. The issue of stability is an important one because we are analyzing the survey responses of actual voters only. Since the turnout differed somewhat between the two elections, and our selection of the sample to be analyzed was contingent upon the vote-no vote distinction, our estimation may produce biased estimates of the true voting parameters. A test for the presence of bias might involve a test of the null hypothesis that the estimated parameter vectors are scalar multiples of each other. However, we have not attempted such a test because of the difficulties associated with the nonlinear budget constraint in the June election.

[32] Of the 19 who switched votes, 3 listed a reduced fear of bussing as the primary cause.

[33] Of the few that did not mention quality of school as a decision variable, one focused on the bussing issue, while several (school teachers) argued for job security.

These results suggest that the presence of the circuit-breaker does not provide much help in explaining the voting changes that occurred between the May and June elections.

To this point we have used the Troy voting data primarily as a means of validating the voting model. However, the data and the model can also be used for forecasting and policy analysis. As an example, we used the estimated voting model to predict the likely effect of the Michigan circuit-breaker under the assumption that voters become sufficiently aware of the circuit-breaker so as to respond to the change in the price of schooling in accordance with the assumptions of our model. To do so, we calculated the predicted values of the May logit model, replacing the original price variable with the "marginal price of schooling" and the "average price of schooling," respectively. Our calculations suggest that accurate knowledge of the circuit-breaker will increase the probability of voting yes in the school election by as little as 3.7% and as much as 9.0%.

IV. Summary and Conclusions

We have developed a model of individual voting behavior in local school millage referenda, which has been estimated using a binary logit model and individual survey responses associated with the first of two elections in Troy, Michigan. The data are generally consistent with the notion that individual voters act in a manner which is coincident with their own self-interest, since the number of children in public school, income, and the price of schooling are three of the most important explanatory variables in the model. By reestimating our model using alternative definitions of income and price we were able to reach some conclusions concerning the bias in estimates of income and price elasticity that result from the failure to account for the deductibility of local property taxes and the inclusion of imputed rental income as income. In addition, our estimates of the relative magnitude of the income and price elasticities provide some information about specification biases involved with aggregated expenditure studies.

We reestimated the logit model using responses associated with the second Troy election to analyze its passage. We found that the change in school price associated with the introduction of the Michigan circuit-breaker legislation does little to explain the outcome. Finally, our results suggest that to the extent that voters become aware of the import of the circuit-breaker, the likelihood that school millage elections pass may increase substantially.

Despite the difficulties associated with the analysis of survey data, we find that the use of such data to elicit information about the demand for public education provides a useful complement to those studies which rely on precinct and school district data. We stress, however, that further research along both theoretical and empirical lines is needed. The most important theoretical improvement would involve a generalization of the model to encompass the voter turnout decision in local school elections. Empirical improvements would involve the creation of data sets which make possible the calculation of separate income and price elasticities of the demand for public schooling (e.g., by considering voter responses to different millage options).

APPENDIX

Analysis of Survey Bias

Of those who actually voted in the May election in Troy, 37.5% voted for the millage renewal; the corresponding figure for the June election was 51.9%. However, of the sample of voters who actually responded to the questionnaire, 55.1% said they voted yes in May, and 63.1% said they voted yes in June. The fact that in both elections the sample contains a higher (relative to the actual election) percentage of yes voters is likely to be the result of two factors. First, yes voters, through self-selection, may have a higher probability of response than no voters. Second, there may be a tendency for some no voters (intentionally or unintentionally) to record their vote as a favorable one. Because the second millage did pass, we would expect this effect to outweigh the tendency of yes voters to record their vote in the negative.

In an attempt to distinguish between these two causes of survey bias, the following procedure was utilized. Let p equal the probability that an individual who actually voted yes in the election will respond to the questionnaire, and q equal the probability that a no voter will respond. Assuming that p and q are constant among precincts and utilizing the fact that voting tallies varied substantially among precincts, we may obtain least squares parameter estimates for

p and q. Within each precinct i, we expect the following relationship to hold:

$$R_i = q + (p - q)Y_i \qquad \text{(A1)}$$

where R is the actual response rate to the survey of those who voted and Y is the proportion of voters in the population who favored the millage proposal.

Equation (A1) was estimated using a cross-section of 18 precincts and June election data. Because the number of survey respondents varied substantially among precincts, weighted least squares estimation was used to adjust for heteroscedasticity. The estimated variances were calculated under the assumption that respondents in each precinct were drawn randomly from a binomial distribution with mean response rate R_i. The regression yields an estimated value of p equal to 0.586 and an estimated value of q equal to 0.503. However, we were unable to reject (at the 10% level) the null hypothesis that p equals q.

If we were to use the estimated values of p and q to predict the voting behavior of those actually returning the questionnaire, we would predict that 55.6% of those returning the questionnaire would respond that they voted yes, and 44.4% would respond no (253 would say yes and 202 no). In the actual survey 63.1% of the respondents replied that they voted in the affirmative. This suggests that of the 11.2 percentage point difference between the actual population yes vote and the recorded survey vote in June, 3.7 points can be explained by bias resulting from self-selection of the respondents, and 7.5 points can be explained by incorrect responses of those returning the questionnaire.

These results can only be taken as suggestive, however, because of the limited number of observations (precincts) available. Further information about survey bias can be obtained by comparing the characteristics of the sample of respondents to the characteristics of residents of the city of Troy obtained from the 1970 census. The mean income of Troy residents was $16,110 in 1969, while the mean income of the sample was $20,208. Allowing for the presence of inflation, this suggests that any bias resulting from a positive correlation between income and probability of survey response will be small. A more likely source of bias appears when we note that in Troy 51.2% of the over 18 population is female, while 57% of those responding were female (66% of the females voting in June said that they favored the millage proposal). Finally, of the residents of Troy, 6.6% of those employed were involved in elementary and secondary teaching (many were out-commuters). However, of those responding to the questionnaire 13.6% of those voting (and 10.5% of the entire sample) reported an occupation directly related to primary and secondary school systems. Of those who voted and were directly involved with an educational occupation, 96% reported favoring the June millage proposal. This suggests that self-selection may have played a greater role in the resulting survey bias than was predicted on the basis of the regression model described previously. For example, if we do not count those with educational occupations among the respondents, our sample vote tally would be 59% for the proposal and 41% against.

To retest for the importance of self-selection bias, we reestimated the probabilities of response for yes and no voters (p and q) using response rates calculated when all responding teachers had been eliminated from the sample. The weighted least squares estimates of p and q were 58.4% and 45.4%, respectively. On this basis we estimated that of the 7.1 percentage point difference between the percentage responding yes in the sample (59.0%) and the percentage voting yes in the election (51.9%), at most 6.2 percentage points can be explained by a self-selection process on the part of school teachers. Thus, after allowing for teacher self-selection, as little as 0.9 percentage points may be explained by the incorrect responses of those returning the questionnaire.

REFERENCES

Advisory Commission on Intergovernmental Relations, "Property Tax Circuit-Breakers: Current Status and Policy Issues," Washington, D.C., Feb. 1975.

Barlow, R., "Efficiency Aspects of Local School Finance," *Journal of Political Economy* 78 (Sept./Oct. 1970), 1028-1040.

Barzel, Y., "Private Schools and Public School Finance," *Journal of Political Economy* (Jan./Feb. 1973), 174-186.

Bass, G. V., and A. J. Alexander, "Schools, Taxes, and Voter Behavior: An Analysis of School District Property Tax Elections," Rand Corporation, Santa Monica, April 1974.

Bergstrom, T. C., and R. P. Goodman, "Private Demands for Public Goods," *American Economic Review* 63 (June 1973), 280-296.

Birdsall, W. C., "A Study in the Demand for Public Goods," in Richard A. Musgrave (ed.), *Essays in Fiscal Federalism* (Washington, D.C.: The Brookings Institution, 1965), 235-294.

Blalock, H. M., *Causal Inference in Non-experimental Research* (Chapel Hill, University of North Carolina Press, 1961).

Borcherding, T. E., and R. T. Deacon, "The Demand for the Services of Non-Federal Governments," *American Economic Review* 62 (Dec. 1972), 891-901.

Bradford, D. F., and W. E. Oates, "Suburban Exploitation of Central Cities and Governmental Structure," in H. Hochman and G. Peterson (eds.), *Redistribution Through Public Choice* (New York: Columbia University Press, 1974), 43-92.

Cox, D. R., *The Analysis of Binary Data* (London: Methuen & Co., Ltd., 1970).

Deacon, R., and P. Shapiro, "Private Preference for Collective Goods Revealed Through Voting on Referenda," *American Economic Review* 65 (Dec. 1975), 943-955.

Edelson, N. W., "Efficiency Aspects of Local School Finance—Comments and Extension," *Journal of Political Economy* (Jan./Feb. 1973) 158-173.

———, "Budgetary Outcomes in a Referendum Setting," chapter 19 in R. W. Lindholm (ed.), *Property Taxation and the Finance of Education* (Madison, University of Wisconsin Press, 1974).

Feldstein, M. S., "Wealth Neutrality and Local Choice in Public Education," *American Economic Review* 65 (Mar. 1975), 75-89.

McFadden, D., "Conditional Logit Analysis of Qualitative Choice Behavior," in P. Zarembka (ed.), *Frontiers in Econometrics* (New York: Academic Press, 1973).

Neenan, W. B., *Political Economy of Urban Areas* (Chicago: Markham Publishing Co., 1972).

Nerlove, M., and S. J. Press, "Univariate and Multivariate Log-Linear and Logistic Models," Rand Corporation, Santa Monica, Dec. 1973.

Oates, W. E., "The Effects of Property Taxes and Local Public Spending on Property Values: An Empirical Study of the Capitalization and the Tiebout Hypothesis," *Journal of Political Economy* 77 (Nov./Dec. 1969), 957-971.

Peterson, G., "The Demand for Public Schooling," Urban Institute Working Paper no.1207-28, Washington, D.C., 1973.

Polinsky, A. M., "The Demand for Housing, A Study in Specification and Grouping," *Econometrica*, forthcoming.

Shapiro, P., "A Model of Voting and the Incidence of Environmental Quality," in H. W. Gottinger (ed.), *Systems Approaches and Environmental Problems* (Gottingen: Vandenhoeck and Ruprecht, 1975).

Theil, H., *Principles of Econometrics* (New York: John Wiley & Sons, Inc., 1971).

U.S. Department of the Treasury, Internal Revenue Service, "Statistics of Income: Individual Income Tax Returns," 1971.