Competition and Teacher Pay

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Abstract

Competition-based school reform could have a significant impact on teacher earnings. If school districts behave as typical oligopsonists, then increased competition could lead to higher salaries. However, if teachers receive a share of the economic rents generated by the monopoly power of school districts, then increased competition could lower teacher salaries.

This study uses panel data from 670 Texas school districts and more than 335,000 teachers to examine the relationship between school competition and teacher pay. The analysis suggests that an increase in competition results in higher wages for most teachers, but lower wages for teachers in relatively concentrated markets.

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I. INTRODUCTION

Recent policy initiatives are increasing the degree of competition in primary and secondary education. The number of charter schools is growing at an astounding rate. Vouchers are available to eligible children in Florida, Utah and a number of cities including Washington, D.C. (Kafer 2005). School districts throughout the country are experimenting with subcontracting campuses to private firms.

The increasing popularity of policies that foster school competition dovetails with a burgeoning literature on the effects of educational competition. Researchers have examined the effects of educational competition on academic outcomes, average public school spending, and school district efficiency.[1]

This research has largely ignored the effects of increased competition on teacher pay. Yet, reductions in the market power of school districts could have a significant impact on teacher earnings. If school districts behave as typical monopsonists or oligopsonists when hiring, then a lack of competition suppresses wages, and policies that are designed to enhance competition among schools could lead to higher teacher salaries, ceteris paribus.[2] Alternatively, if a school district's position as a monopolistic provider of education services generates economic rents, there are few parties to whom those rents might be dissipated besides school district personnel. If teachers have been able to appropriate a share of the economic rents generated by the market power of school districts, then increased competition could lower teacher salaries.[3]

This study uses individual data on more than 335,000 teachers from 670 Texas school districts to explore the empirical relationship between competition and teacher pay. The analysis suggests that a lack of competition in the public school system has led to significant market power for Texas school districts, resulting in lower wages for most Texas teachers, but higher wages for teachers in relatively concentrated markets.

II. THEORETICAL MODELS

The elementary and secondary education industry in the United States is often highly concentrated. More than 30% of U.S. educational markets are served by a single public school district, and there are many more that are served by only a handful of educational providers.[4]

The concentrated nature of education markets suggests two alternative models of teacher wage determination. The first is a classical oligopsony model: the second is a rent-sharing model.

An Oligopsony Model of Teacher Wage Determination

Boal and Ransom (1997) lay out a classic, Cournot model of oligopsony which can be directly applied to the educational labor market.[5] Firms (school districts) choose an employment level (L_i) to maximize the firm's revenue function (R_i), given the employment level of all other firms in the market (L_{i*}).

$$R_i(L_i) - w(L_i + L_{i*})L_i$$

The market wage is determined by the total employment of all firms. The first-order condition for each firm implies that it chooses L_i such that

$$(MRP_i - w)/w = (L_i / L)\varepsilon^{-1}$$

In this model, the gap between the marginal revenue product of firm i (MRP_i) and the wage rate (w) depends on the firm's employment share (L_i/L) and the inverse wage elasticity of labor supply for the market (ε^{-1}).[6] Assuming that firms have different marginal revenue products, the employment-weighted average across all of the firms in the market is:

$$\sum_i (MRP_i - w)L_i / wL = \varepsilon^{-1} \sum_i (L_i / L)^2$$

The last term on the right-hand-side is the familiar Herfindahl index of market concentration. Thus, wages are a function of market concentration (H), the (employment-weighted) average marginal revenue profit per worker ($\overline{MRP}$), and the elasticity of labor supply (ε).

$$w = \overline{MRP}\left(H\varepsilon^{-1} + 1\right)$$

Unfortunately, determining the effect on wages of changes in market concentration is not so simple as taking the partial derivative of (4) with respect to H. Market concentration can only change if the number of firms or the relative size of those firms changes. Average marginal revenue product also depends on the number of firms and the distribution of marginal revenue products across those firms. Thus, H cannot change without affecting $\overline{MRP}$. However, as discussed in Boal and Ransom (1997), if the market-level labor supply and marginal revenue product curves are held constant, then a negative correlation across markets between wages and market concentration implies oligopsony.[7]

A Rent-Sharing Model of Teacher Wage Determination

Oligopsony may not be the whole story, however. In communities where there is only one potential employer of teachers, there is also only one provider of educational services. Such territorial exclusivity could give rise to oligopoly power in the market for education services. Because public schools have no shareholders or capital owners per se, the parties most likely to capture any rents generated in the output market are the school district employees.

Blanchflower, Oswald and Sanfey (1996) develop a model in which employees share in any rents earned by the firm. Wages in an industry are a function of expected opportunity wages outside the industry, industry rents per worker (π / L) and the worker's relative bargaining power $(\phi / 1 - \phi)$ as follows:

$$w \cong \overline{w} + \left(\phi / \left(1 - \phi\right)\right) \cdot \left(\pi / L\right).$$

Assuming that capital is a quasi-fixed factor (at least in the short run) average profits per worker can be expressed as

$$\pi / L = \sum_i \left(MRP_i - w^o\right) \cdot \left(L_i / L\right) = \overline{MRP} - w^o$$

where w^o is the locally prevailing wage. Because the expected opportunity wages for educators $(\overline{w})$ should be a function of the prevailing wage (w^o) and the local unemployment rate (U), substituting equation (6) into equation (5) yields,

$$w \cong c\left(w^o, U\right) + \left(\phi / \left(1 - \phi\right)\right) \cdot \left(\overline{MRP} - w^o\right)$$

As equation (7) makes clear, rent-sharing implies that teacher wages are an increasing function of school district rents, if any.[8] In turn, a substantial literature indicates that school districts are inefficient, and that their inefficiency falls with competition (Taylor 2000, Grosskopf et al. 2001). Because such inefficiency is a potential source of economic rents, increased competition could reduce teacher pay.

III. EMPIRICAL MODEL

The classical oligopsony model and the rent-sharing model support very similar reduced-form equations for teacher pay. Let **Z** be a vector of factors other than market concentration that determine marginal revenue of school districts (primarily the determinants of local education demand and the educational technology). Assuming that market concentration is one of the sources of economic rents under the rent-sharing scenario, and that educator labor supply is a function of the determinants of expected opportunity wages, then

$$w_{oligopsony} = f\left(H, Z, w^o, U\right)$$

and

$$w_{rent-sharing} = g\left(H, Z, w^o, U, \phi\right)$$

The most obvious difference between the two reduced-form equations is that the rent-sharing equation has an extra term for the educator's relative bargaining power.[9] However, for the purposes of this analysis, the most important distinction between the two equations is the very different partial derivative with respect to market concentration (H). Holding constant the determinants of marginal revenue and labor supply ($\mathbf{Z}$, w^o, and U), the expected partial derivative of $w_{oligopsony}$ with respect to H is negative while the expected partial derivative of $w_{rent\text{-}sharing}$ with respect to H is positive. Thus, anticipating whether the trend toward greater competition in education will raise or lower teacher salaries turns on which model best explains teacher pay.

The existing literature suggests that either outcome is possible. As discussed in Medcalfe and Thornton (2006), most of the research since the 1970s has found either that teacher wages fall with market concentration or that there is no relationship between market concentration and teacher compensation. However, Hoxby (1994b) found that teacher wages rise with market concentration, a pattern she attributed to rent sharing.

Clearly, researchers can only find evidence of rent sharing in markets where there are rents to share. Taylor's (2000) survey of the literature concluded that the relationship between competition and school district efficiency is non-linear, and that small increases in competition have little or no impact on efficiency in markets that are already highly competitive. As such, rent-sharing implies that small changes in competition may have greater impact on wages in relatively concentrated market than in relatively competitive ones. Previous researchers may have been unable to detect rent sharing because they relied upon linear specifications.

It is also important to recognize the potential influence of teacher unions and teacher political action committees. Unions can mitigate the oligopsony power of school districts and increase the share of economic rents captured by teachers. Markets where unions are particularly strong or teachers are politically influential are more likely to exhibit signs of rent-sharing, all other things being equal. School districts are more likely to exploit their market power in markets where unions are weak and teachers lack political influence.

IV. THE DATA

The Texas public school system represents an attractive laboratory for exploring the relationship between competition and compensation. The state of Texas contains 67 Core Based Statistical Areas that are considered distinct labor markets by the U.S. Office of Management and Budget (OMB). Those 26 metropolitan areas and 41 micropolitan areas contain 670 traditional independent school districts, another 186 charter school districts, and more than 400 accredited private school districts.[10] As a result, there is a wide range of potential market structures located within a single state's policy environment.

Texas is also a right to work state that does not have a collective bargaining law for teachers (Krueger 2002). As such, analyses of Texas labor markets are not complicated by the need to measure and control for variations in the extent of teacher union influence.

Finally, the quality of the available data is another reason why Texas is particularly well suited to this analysis. Much of the variation in teacher compensation reflects the premia paid for individual-specific characteristics like experience or educational attainment. Because the distribution of teacher characteristics is likely to vary across labor markets, it is important to control for such characteristics when estimating the effect of competition on teacher salaries. Texas' administrative records provide the data necessary for such controls.

The Texas Education Agency (TEA) collects detailed information on the earnings, demographics and job characteristics of all school district personnel. The individual personnel files indicate years of experience, educational attainment, gender, ethnicity, effective days worked (percent day times number of days employed), school assignment and employing district. In addition, the teacher records include indicators for job assignments, and for the percentage of time each individual spends teaching in a field for which he or she holds a Texas state teaching certificate.[11]

The confidential TEA records yield an unbalanced panel covering 335,527 teachers who worked at least half time for a traditional Texas school district in a CBSA during at least one of the five school years from 1999-2000 through 2003-04.[12] Analysis using such a rich panel of payroll data will provide an excellent test of the relationship between competition and teacher pay.

A key element of such analysis is the definition of market area because it is critical to measuring H, **Z**, w^o and U from equation (8). As is common in the literature on school district competition (e.g. Chambers 1995 and 1997; Hoxby 2000; and Taylor et al. 2002), this analysis equates education markets with the labor market areas defined by the OMB. Thus, each CBSA is treated as a distinct market in this analysis.

As one might expect, market concentration in education varies substantially across the CBSAs in Texas. The Herfindahl index is the sum of squared enrollment shares for all public and private school districts in a labor market. In 2003-04, it ranged from below 0.10 in five metropolitan areas (Dallas, Fort Worth, Houston, Longview and San Antonio) to above 0.90 in four micropolitan areas and one metropolitan area. Metropolitan labor markets are generally less concentrated than micropolitan ones, but each market type included at least one highly concentrated market each year from 1999-2000 through 2003-04 (Table 1).

The **Z** vector contains market characteristics other than concentration that determine the marginal revenue product of school districts. Likely candidates include factors describing the education technology and the local demand for education.

Cost function analyses indicate that scale is the primary determinant of the educational technology. (For example, see Gronberg et al. 2004). Researchers generally find a non-linear relationship between cost and school district size. Therefore, I include in the Z vector not only the average school district enrollment in the labor market, but also its square.[13]

Education demand is frequently modeled as a function of the tax price of education and voter demographics. Following Hoxby (2001), I define the tax price of education as the additional tax revenue a district must generate to be able to spend an additional dollar on education. It is calculated from the statutory provisions of the Texas school finance formula as a function of school district wealth, student demographics and Texas' cost of education index.[14] The average tax price in each CBSA is a weighted average of the tax prices in the constituent school districts, weighted by each district's share of teacher employment in that market. Data on voter demographics in each CBSA—median income, the percent of the adult population with a high school diploma but no college degree, the percent of the adult population with at least a bachelors' degree, the percentage of residents that are school age children, and the percentage of residents over 65—come from the 2000 Census and subsequent population estimates.

The expected wage is a function of the local unemployment rate and the prevailing wage in a labor market (w^o). Data on unemployment rates come from the Bureau of Labor Statistics. I use the NCES Comparable Wage Index to measure the prevailing wage for college graduates in each CBSA (Taylor and Fowler 2006).[15]

Finally, the individual wages received by teachers could reflect district-specific compensating differentials as well as the market wage for teachers, and the premia paid for individual teacher characteristics. Districts with a student body that is perceived as unusually challenging to teach may have to pay a premium to staff their classrooms. On the other hand, districts with low enrollments are likely to have small class sizes as well, allowing them to hire at a modest discount. Similarly, given commuting costs and typical rent gradients, districts near the center of a metropolitan area may have to pay a premium to attract teachers while districts on the urban fringe may be able to hire at a modest discount. To control for such effects, the model includes student demographics (the percentage of students who are economically disadvantaged, limited English proficient, black and Hispanic), school district size (indicator variables for small and mid-size school districts using Texas Education Agency definitions) and the distance from the center of the closest metropolitan area.

Formally, the specification can be expressed as:

$$\ln(W_{idmt}) = \beta_H H_{mt} + \beta_{H2}(H_{mt})^2 + Z_{mt}\beta_Z + \beta_w w_{mt}^O + \beta_U U_{mt} + D_{dt}\delta + T_{it}\gamma + \varepsilon_{idmt}$$

where the subscripts i,d,m and t stand for individuals, districts, labor markets and time, respectively, $\mathbf{D}_{dt}$ is a vector of district-specific characteristics that could give rise to compensating differentials, and $\mathbf{T}_{it}$ is a vector of individual-specific characteristics. Because previous research has found a nonlinear relationship between market concentration and school district inefficiency, and such inefficiency is the hypothesized source of any economic rents, the specification also includes the square of the Herfindahl index. Following the literature on teacher compensation, the dependent variable ($\ln(W_{idmt})$) is the logarithm of full-time-equivalent monthly salary.

V. ESTIMATION AND RESULTS

The first column of table 2 presents coefficient estimates and robust standard errors from the baseline estimation of equation (9). Because the Herfindahl index, comparable wages, unemployment rates and the variables in **Z** do not vary within labor markets, the observations within any given labor market may not be independent. Therefore, the standard errors have been adjusted to reflect clustering by CBSA.[16]

As the table illustrates there is a significant–and nonlinear– relationship between market concentration and teacher salaries. Wages fall as market concentration rises—but only up to a point. Once the Herfindahl index exceeds 0.54, wages begin to rise with concentration.

The nonlinear pattern of teacher salaries suggests that both oligopsony power and rent sharing are at play in the market for Texas teachers. In relatively competitive markets, the oligopsony effect dominates, and wages fall with market concentration. However, in relatively concentrated markets, rent sharing best explains teacher pay. Rising concentration generates economic rents that lead to rising teacher salaries. Thus, the analysis suggests that policy changes designed to foster increased competition among school districts may have differing effects on teacher salaries, depending on the degree of competition already present in the market.

While the above analysis is a good first cut at the relationship between market concentration and teacher pay, it does not address the possible endogeneity of market competition. Caroline Hoxby illustrated the importance of treating competition among school districts as endogenous, using topographic information on rivers and streams as the exogenous source of variation (Hoxby 1994, 2000). Although her methodology has recently been the subject of heated debate (Rothstein 2005, Hoxby 2005), her basic observation remains indisputable: the amount of competition facing a school district may not be exogenous.

Unfortunately, Hoxby's preferred instrument–the physical location of rivers and streams–is not well suited to this analysis. The boundaries of Texas school districts are more a byproduct of political history than of natural barriers. Furthermore, charter schools and private schools have considerable influence on the amount of competition in an educational labor market, and their geographic distribution is unlikely to reflect regional topography. Finally, topography by its very nature does not vary over time. Taking full advantage of the five-year panel used in this analysis requires instruments that have not only cross-section but also time-series variation.

Accepting that the pattern of district boundaries in Texas is largely given by history, any endogenous variation in measured competition arises from the location decisions of private and charter schools, and from differences in the rate of enrollment growth among traditional public schools. The determinants of these two factors would be reasonable instruments for market competition.

Arguably, private and charter schools should locate where their expected revenues exceed their expected costs. Grosskopf, Hayes and Taylor (2004) demonstrate that the pattern of charter competition is well explained by the profit potential of school district neighborhoods. Their determinants of profit potential (which are derived from the Texas school finance formula and the local demand for schooling alternatives) include the school tax rate, student demographics, the rate at which students are failing standardized exams and the size of the educational market. I use the latter factors–the rates at which minority (black and Hispanic) and non-minority students are failing standardized exams and the size of the educational market (measured by the total land area of the CBSA and the total public and private enrollment per square mile)–as instruments for the public and private Herfindahl index.

Differences in enrollment growth among traditional public schools should reflect differences in population density and the availability and distribution of undeveloped land in the market. Therefore, I also include as instruments the population density of the core county in the CBSA, the average share of undeveloped land in the CBSA and the share of undeveloped land in the largest school district in the CBSA. Markets with a greater share of undeveloped land, holding constant the population density and the share of undeveloped land in the largest district, should be more competitive because they have more of a residential fringe.

To ensure that the seven instruments are at least nominally exogenous, I measure all of them with a five-year lag.[17] Thus, the instruments for market competition in 1999-2000 are average market characteristics from 1994-1995. Relying on a five-year lag ensures that the salary panel does not overlap with the panel of instruments. The most recent values used as instruments come from 1998-99, one year before the first salary observations.

Using the lagged determinants of school location and uneven enrollment growth (and their squares) as instruments for the Herfindahl index (and its square) yields the model presented in the second column of Table 2. A Durbin-Wu-Hausman test easily rejects ordinary least squares in favor of this instrumental variables specification.[18] Furthermore, analysis confirms that the excluded instruments have the necessary properties of good instruments, being both well correlated with the potentially endogenous variables and uncorrelated with the errors from the primary equation.

The most common indicators of the correlation between the endogenous variables and excluded instruments are an F-test of the joint significance of the excluded instruments in the first-stage regression, and the squared partial correlation between the excluded instruments and the endogenous variable, or partial R^2. Here, the relevant F-statistics are 16.63 for the Herfindahl index and 6.84 for its square, and in both cases the probability of a greater F-statistic is 0.000. Meanwhile, the partial R^2 is 0.7717 for the Herfindahl index and 0.6465 for its square.

With multiple endogenous variables, however, intercorrelations among the instruments may render one or more of the instruments irrelevant, and leave the model unidentified despite a statistically significant F-test and a high partial R^2 (Baum, Schaffer and Stillman 2003). Shea (1997) developed a partial R^2 statistic that is a more appropriate measure of instrument relevance for multivariate models. The Shea's partial R^2s are 0.4871 for the Herfindahl index and 0.4082 for its square, providing further evidence that the lagged determinants of school location and uneven enrollment growth are well correlated with the Herfindahl index and its square.[19, 20]

In addition to being well correlated with the potentially endogenous variable, good instruments must also be uncorrelated with the errors from the primary equation. Because the number of instruments excluded from the primary equation exceeds the number of endogenous variables, the IV specification in Table 2 is overidentified and a Hansen's *J* statistic can be used to test for whether or not the excluded instruments are uncorrelated with the errors from the primary equation. The Hansen's *J* statistic for the orthogonality of the excluded instruments is 14.61, and the probability of a greater *J* statistic is 0.263.[21] The null hypothesis is that the instruments are valid, so this failure to reject the null indicates that the instruments used in this analysis are also uncorrelated with the errors from the primary equation, and are therefore good instruments for the Herfindahl index and its square.

As the table illustrates, instrumenting for market concentration strengthens the previously indicated relationship. Highly competitive labor markets are best characterized by oligopsony while highly concentrated labor markets are best characterized by rent sharing. Intriguingly, the estimation also suggests that wages are as high in perfectly concentrated markets (those with a

Herfindahl index of 1.00) as they are in perfectly competitive markets (those with a Herfindahl index of 0.00), all other things being equal.[22]

While the 2SLS model is a clear improvement over the OLS model, it is possible to further improve the modeling by using a more efficient estimator. When heteroskedasticity is present (as is typically the case with panel data on earnings) the Generalized Method of Moments (GMM) estimator is more efficient than the 2SLS estimator (Baum, Schaffer and Stillman 2003). Therefore, the third column of table 2 presents GMM estimates of equation (9) and the corresponding cluster-adjusted robust standard errors.

As the table illustrates, the GMM model also indicates a non-linear relationship between market concentration and teacher pay. As with the OLS and 2SLS models, the GMM model indicates that the relationship between salaries and concentration is U-shaped, and that wages in perfectly concentrated markets are not significantly different from wages in perfectly competitive markets.

Labor Market Effects

Although the estimated relationship between market concentration and teacher pay is highly significant, one might still wonder if there are important market characteristics that have been omitted from the analysis. For example, the supply of potential teachers might be greater in college towns than in other markets. The Herfindahl index could be picking up the influence of unobserved market characteristics rather than the effect of market concentration.

The final three columns of Table 2 address this concern. Each of these models (OLS, IV and GMM) includes indicator variables for each labor market. Labor market characteristics that do not vary over time (such as the percentage of adults with a college degree) have been dropped from the estimation, so it is no longer necessary (or credible) to cluster the data by CBSA.

However, other variables such as the Comparable Wage Index still do not vary within labor markets in any given year, so the observations within a labor market in any given year may not be independent. Therefore the data have been clustered by market-years. There are 355 market-year clusters (67 markets X 5 years).

The IV and GMM models treat the Herfindahl index and its square as endogenous, and the two excluded instruments that do not vary over time (land area and land area, squared) have naturally been dropped from the analysis. Because they are not effective instruments in the presence of market fixed effects, the share of undeveloped land and its square have also been dropped. However, the five remaining excluded instruments (the passing rates, enrollment density, population density of the core county in the CBSA, and the share of undeveloped land in the largest school district in the CBSA) continue to be good instruments for the Herfindahl index and its square. The Shea's R^2s for the Herfindahl index and its square are 0.388 and 0.229, respectively, and the probability of a higher Hansen's J statistic is 0.090. A Durbin-Wu-Hausman test easily rejects ordinary least squares in favor of an IV specification for the model with labor market fixed effects.[23]

Adding labor market fixed effects to the analysis does not alter the basic results. Teacher salaries have a significant and non-linear relationship with market concentration. Salaries fall with concentration in relatively competitive markets, but rise with concentration in relatively concentrated markets. Furthermore, wages in perfectly competitive markets remain indistinguishable from wages in perfectly concentrated ones.

Adding labor market fixed effects greatly increases the magnitude of the estimated marginal effect of market concentration, however. For example, the Herfindahl index for Dallas fell from 0.086 in 2000 to 0.072 in 2004. The fixed effects model suggests that this 1.4

percentage point decline in market concentration increased teacher salaries by 3.5%, ceteris paribus. In contrast, the model without fixed effects indicates that this change in concentration would have raised salaries by no more than 0.7%.

An increase in the coefficient estimates is not unexpected. The model with labor market effects estimates the influences of changes in market concentration within a given labor market. The within-market analysis can be thought of as a model of the short-run elasticity of wages with respect to market concentration, while the pooled, cross-section analysis can be thought of as a model of as modeling a longer-run elasticity. Although their analysis focused on nurses rather than teachers, Hirsch and Schumacher (2005) also found that the short run elasticity of wages with respect to market concentration greatly exceeds the longer-run elasticity.

The increase in the estimated elasticity of wages with respect to market concentration also suggests that market concentration in education is correlated with unobserved market characteristics, and consequentially, that cross-sectional estimates of the impact of concentration may understate the influence of concentration on teacher pay, at least in the short run.

The within-market analysis indicates that small changes in market concentration can have relatively large effects on teacher pay. A one percentage point increase in market concentration can reduce teacher wages by up to 3.0%, or increase teacher wages by up to 3.7%, depending on the initial level of concentration. At one extreme of the actual Texas experience, market concentration rose by 3.8 percentage points (from 0.538 in 1999-2000 to 0.576 in 2003-04) in Plainview, Texas, a change the model predicts would have increased salaries by 2.7%. At the other extreme, market concentration in Beeville, Texas fell by 6.5 percentage points (from 0.612 in 1999-2000 to 0.548 in 2003-04), a change the model predicts would have decreased teacher wages by 5.3%. Changes in market concentration that exceed 6.5 percentage points are outside the scope of the data, so great caution should be used in estimating the effects of such changes.

The first derivative of equation (9) with respect to the Herfindahl index indicates whether salaries increase or decreases with market concentration. Based on the GMM model with labor market fixed effects, I generated 95% confidence intervals for the marginal effect of market concentration and used them to classify labor markets into oligopsonistic markets, rent-sharing markets, and competitively neutral markets.[24] I classified as oligopsonistic any CBSA where the confidence interval for the marginal effect contained only negative values when evaluated at the CBSA's Herfindahl index, and classified as rent-sharing any CBSA where the confidence interval for the marginal effect contained only positive values. By this criterion, CBSAs where the Herfindahl index was below 0.38 were oligopsonistic markets, while CBSAs where the Herfindahl index was above 0.58 were rent-sharing markets.

In 2004, 29 Texas CBSAs were best characterized as oligopsonistic, 20 CBSAs were neutral, and 18 CBSAs exhibited signs of rent sharing. Nineteen of the state's 26 metropolitan areas were oligopsonistic, but three metropolitan areas (Midland, Odessa and San Angelo) were rent-sharing markets. Meanwhile, there were five more rent-sharing micropolitan markets than there were oligopsonistic micropolitan markets.

During the 2003-04 school year, the vast majority of Texas teachers worked in markets where wages would rise with competition. Eighty-nine percent of the Texas teachers in the sample worked in one of the 29 oligopsonistic labor markets. Only 3.7% of the teachers worked in rent-sharing markets. Intriguingly, the teachers who worked in rent-sharing markets were significantly more experienced and less well educated, on average than teachers who worked in oligopsonistic ones.

Sensitivity Analysis

Readers familiar with Texas education markets might wonder if these estimates of concentration effects are biased by the geographic distribution of Texas teachers. Nearly 24% of the teachers in the sample work in the Houston metropolitan area, and another 17% of the teachers work in the Dallas metropolitan area. Furthermore, Houston and Dallas also have the two lowest values for the Herfindahl index (0.063 and 0.072 in 2003-04, respectively). It is reasonable to suspect that these two markets have undue influence on the regression analysis.

The second column of Table 3 presents the fixed effects analysis excluding the Dallas and Houston metropolitan areas. As the model makes clear, excluding Dallas and Houston from the analysis reduces the magnitude of the estimated marginal effects, but has little qualitative impact on the estimation.

Given the differences in teacher demographics across market types, one might also wonder if the relationship between salaries and competition could vary with teacher experience or educational attainment. Arguably, additional years of experience in public education or an advanced degree could make a teacher both less mobile across occupations and better positioned to capture any school district rents. Either mechanism could have a perceptible influence the relationship between competition and teacher pay.

To explore this possibility, I divided teachers into three experience groups–beginning teachers, experienced teachers and highly experienced teachers–and divided non-beginning teachers into two educational attainment groups–teachers with a bachelor's degree or less and teachers with a master's degree or more–and re-estimated equation (9).[25] Following the definitions used by the National Center for Education Statistics, I consider teachers with three or fewer years of experience to be beginning teachers (U.S. Department of Education 2005). By this definition, roughly one quarter of the teachers in the sample are beginning teachers. Roughly one quarter of the teachers in the sample have at least 20 years of experience. I consider them highly experienced teachers. All remaining teachers are considered experienced teachers. Table 3 presents selected coefficient estimates.

As the table illustrates, all of the specifications find a significant and non-linear relationship between market concentration and teacher pay. In all cases, one can resoundingly reject the hypothesis that the Herfindahl index and its square are jointly zero. All but two sub-samples also find that wages under perfect competition are not significantly different from wages under perfect concentration. The exceptions are for highly experienced teachers and teachers with a Masters degree. For these teachers, rent sharing leads to wages that are significantly higher in perfectly concentrated markets than they are in perfectly competitive ones.

Notably, the rents extracted by teachers in relatively concentrated markets are not the result of direct union activity. Because Texas is a right to work state that does not have a collective bargaining law for teachers, teachers do not gain their share of the rents through collective bargaining at the school district level.

On the other hand, many work rules favored by teacher unions–such as maximum class sizes–have been written into the regulatory code by the Texas Legislature. One of those legislatively imposed regulations is a minimum salary scale that is increasing in experience but not in educational attainment, tops out at 20 years of experience, and does not vary from district to district. As table 4 illustrates, the steps in the salary scale are not even, and were unchanged from the 1999-2000 school year through the 2001-02 school year. For the 2002-03 school year, the Legislature mandated a $100 per month increase in pay for each teacher. However, state-level budget constraints led the Legislature to take back half of the increase in 2003-04, effectively lowering the salary scale by $50 per month.

The salary scale is binding with respect to salary for less than 2.3% of the teachers in the sample, but it is binding for more than half of the teachers in some of the smaller micropolitan areas. The minimum salary scale could alter the relationship between compensation and concentration by suppressing wage differences that might otherwise be observed in response to differences in market concentration. If so, then the GMM estimates of the relationship between market concentration and teacher pay could be biased.

Table 5 presents the selected coefficients from a Tobit re-estimation of the three teacher experience models, assuming that the wages are censored at each teacher's statutory minimum wage. To accommodate a Tobit specification, I subtracted the log of the individual's statutory minimum wage from both sides of equation (9).[26] Thus, the dependent variable has been redefined as the difference between the observed log wage and the log of the statutory minimum wage for the individual's level of experience. The Herfindahl index and its square are treated as endogenous, using Newey's (1987) two step estimator. The standard errors are not adjusted for clustering of the data, and are undoubtedly too small. That said, hypothesis tests based on the unadjusted standard errors indicate that in all three cases there is a significant and non-linear relationship between market concentration and teacher pay.

Each panel of figure 1 compares the percentage difference from perfect competition predicted by the GMM and Tobit models. Each line in the figure traces out the percentage difference between predicted salaries in perfectly competitive markets and predicted salaries in markets with the designated value for the Herfindahl index. The top panel shows the differences for beginning teachers, the middle panel shows the differences for experienced teachers and the bottom panel shows the differences for highly experienced teachers. In all three panels, the dashed lines are the confidence bands around the GMM estimates.

As the figure illustrates, there are few qualitative differences between the Tobit estimates and the GMM estimates. For both beginning teachers and highly experienced teachers, the Tobit estimates lie well within the confidence interval from the GMM estimates. For experienced teachers, the Tobit estimates are somewhat less sensitive than the GMM estimates to market concentration in the middle of the distribution, but well within the confidence interval in most of the markets identified as rent-sharing by the GMM estimation. Because Tobit point estimates can be sensitive to heteroskedastcity in the errors, and the standard errors from the Tobit models are not adjusted for clustering by labor market, the GMM models are arguably more conservative. Therefore, the GMM models by experience group are the preferred specification.

A Comparison with Previous Literature

The GMM models by experience group indicate that a one percentage point increase in market concentration can reduce teacher wages by up to 3.4%, or increase teacher wages by up to 4.7%, depending on teacher experience and the initial level of concentration. As such, these estimates of the marginal effect of market concentration are substantially larger than those found by previous researchers. Borland and Howsen's research predicts that a one hundred percentage point increase in concentration (from a Herfindahl index of zero to a Herfindahl index of one) lowers average teacher salaries by less than 3%.[27] Vedder and Hall (2000) found that salaries were 2% lower in monopsonistic counties than in counties with 10 public school districts (which, if they were equal sized, would be equivalent to a Herfindahl index of .10). Merrifield (1999) found that a one hundred percentage point increase in concentration would reduce salaries by less than 1%, while Medcalfe and Thornton (2006) found that it would have no effect.

The results are also strikingly different from those found by Hoxby (1994b). Hoxby's analysis of teacher pay in U.S. metropolitan areas found evidence of rent sharing, but no evidence of oligopsony power for school districts. After controlling for the endogeneity of market concentration, her analysis indicated that wages in perfectly concentrated markets were 18% higher than wages in perfectly competitive markets. This analysis suggests that oligopsony is more common than rent sharing, and that wages in perfectly concentrated markets are comparable to wages in perfectly competitive markets for all but the most experienced or highly educated teachers.

The non-linear specification used in this analysis is an important departure from the existing literature. Previous researchers reported only linear versions of their models, and a linear model obscures much of the relationship between market concentration and teacher pay.[28] The specification used here can be thought of as an unrestricted version of those previous models. As is clear from table 2, restricting the model to be linear with respect to market concentration is strongly rejected.

With the exception of Luizer and Thornton (1986), previous researchers also relied on average salaries to measure compensation, and included few, if any, controls for teacher characteristics. As figure 1 illustrates, the relationship between market concentration and teacher pay can be sensitive to differences in teacher experience.

Finally, previous researchers used cross-section but not time-series variations to estimate the relationship between concentration and compensation. As the analysis demonstrates, cross-sectional estimates may understate the influence of market concentration on teacher pay, at least in the short run.

Restricting the analysis to something comparable to the existing literature yields coefficient estimates that are similar to those in the existing literature on monopsony in the education market. As Table 6 illustrates, a linear, cross-sectional analysis that does not control for differences in the geographic distribution of teacher characteristics finds evidence of mild oligopsony, and no sign of rent-sharing. Thus, the differences from the previous literature appear driven primarily by methodological improvements rather than by unusual characteristics of the teacher labor market in Texas.

VI. IMPLICATIONS AND CONCLUSIONS

This analysis of individual data on more than 335,000 Texas teachers suggests that a lack of competition in the public school system has led to significant market power for school districts. In relatively competitive markets, school districts have oligopsony power and an increase in competition would tend to increase teacher pay. However, in relatively concentrated markets, teachers are able to share in some of the rents generated by the oligopoly power of school districts and an increase in competition would tend to reduce their expected earnings.

Based on the GMM estimates by experience group, I generated 95% confidence intervals for the marginal effect of market concentration, and used them to classify labor markets into oligopsonistic markets, rent-sharing markets, and competitively neutral markets. An intriguing pattern emerges (Table 7). For all teachers, markets where the Herfindahl index is below 0.26 are oligopsonistic, and wages fall with concentration. For markets where the Herfindahl index is between 0.26 and 0.34, wages fall with concentration for all but the highly experienced teachers. For highly experienced teachers, the marginal effect is insignificant. For the narrow window from a Herfindahl index of 0.34 and a Herfindahl index of 0.37, the marginal effect of concentration is significant (and negative) only for experienced teachers. For markets with a

Herfindahl index between 0.37 and 0.49, all three types of teachers are indifferent to marginal changes in market concentration. For a Herfindahl index between 0.49 and 0.54, the marginal effect of market concentration is significantly positive for highly experienced teachers, but insignificant for all other teachers. For Herfindahl index values above 0.54, wages rise with concentration for all experienced teachers, and for Herfindahl index values above 0.68, wages rise with concentration for all teachers.

The decomposition thus reveals that there are 15 Texas markets where all types of teachers would benefit from an increase in competition. At least some teachers would benefit and no teachers would lose in another 13 markets. There are 13 markets where all three models predict that wages would not change with competition, and 26 markets where increased competition would lead to lower salaries for at least some teachers.

The vast majority of Texas teachers work in oligopsonistic markets and would therefore benefit from an increase in educational competition. More than 88% of the teachers with less than 20 years of experience would benefit from increased competition. Seventy-nine percent of the highly experienced teachers would also benefit. Only 2% of beginning teachers, 5% of experienced teachers and 6% of highly experienced teachers could expect increased competition to lower their pay.

The prospects of significant labor market impacts raises the stakes associated with competition-based school reform. In particular, a district's loss of oligopsony power could lead to higher wages that at least partially offset the hoped for savings from efficiency gains. For example, consider the charter school movement in Texas. More than 70% of the charter school students and 54% of the charter schools in Texas are located in the Houston, Dallas or San Antonio metropolitan areas. Based on the degree of competition in those markets before the first charter schools opened their doors, the models predict that a one percentage point increase in competition in those markets would have increased teacher wages (in those markets) by 2.5%. Given that teacher payroll is more than 40% of current operating expenditures, a 2.5% increase in wages represents a significant–and largely unrecognized– downside risk to competition-based school reform.

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NOTES

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1.For surveys of the literature on educational competition, see Belfield and Levin (2002) and Taylor (2000). Recent contributions to this literature include Booker et al. (2004) and Grosskopf et al. (2004).

2. For example, see Boal and Ransom 1997, Luizer and Thornton 1986, Merrifield 1999, Vedder and Hall 2000, or Medcalf and Thornton 2006.

3. For example, see Blanchflower, Oswald and Sanfey 1996, Hildreth and Oswald 1997 or Black and Strahan 2001.

4. According to the National Center for Education Statistics' F-33 school finance files for the 2001-02 school year, 805 of the 2593 U.S. education markets are served by a single local education agency. Education markets are defined as the CMSA for all districts in a CMSA and as the county for all other districts.

5.This model has been referred to as the "classic" monopsony model (e.g. Hirsch and Schumacher 2005). The "new" monopsony literature emphasizes labor market frictions such as commuting costs or worker preferences regarding non-wage attributes of jobs rather than market concentration per se (e.g. Manning 2004; Manning 2003, Brueckner, Thisse and Zenou 2002; Bhaskar, Manning and To 2002).

6.Note that $L_i + L_{i*} = L$.

7.Provided, of course, that labor supply is not perfectly elastic.

8. School districts could maximize potential rents by raising or lowering expenditures on school quality, depending on the elasticity of student enrollment with respect to quality. As discussed in Bayer and McMillan (2005), the elasticity of demand with respect to school quality increases with competition. Therefore, potential school rents from a reduction in quality are smaller in more competitive markets.

9.If ϕ does not vary systematically across labor markets, the equations are observationally equivalent.

10.A private school district, such as the Catholic Archdiocese of San Antonio, can operate many schools.

11.On average Texas teachers spend more than 25% of their teaching time in a field for which they are not certified.

12. Charter school personnel were excluded from the wage analysis.

13.Analyses using the average log enrollment and its square yield very similar results.

14.The cost of education index, which is a component of the Texas school finance formula, indicates differences in the predicted cost of hiring teachers in 1989. For more on Texas' cost of education index, see Alexander et al. (2000).

15. The BLS had not yet adopted the new CBSA definitions, so I assigned to each school district the corresponding county and metropolitan area estimates of the unemployment rate, and then

calculated the CBSA unemployment rate as a weighted average of the unemployment rates for each district in the CBSA. Similarly, the average CWI for each CBSA is a weighted average of the CWI for each district in that CBSA. In both cases, the weights are each district's share of teacher employment in the CBSA.

16. Because school districts are contained within CBSAs, clustering by CBSA captures any correlation in the errors within school districts.

17. Of course, because the CBSA definitions are constant over time, the land area in a CBSA is constant over time, and lagging has no effect on this instrument.

18.The Durbin-Wu-Hausman test statistic is 9.32 with 2 and 66 degrees of freedom. The probability of a greater test statistic is 0.0003.

19. Shea's partial R^2 "may be interpreted like any R^2" (Baum, Schaffer and Stillman 2003, p. 15). Previous empirical work by Fedorov and Sahn (2005) refers to Shea's partial R^2s of 0.34 and 0.22 as "fairly high."

20.The Anderson-Rubin test for the joint significance of the endogenous regressors in main equation (which is robust to weak instruments and clustering) also indicates that the model is identified. The chi-squared test statistic is 212.33 with 14 degrees of freedom, and the probability of a greater test statistic is 0.000.

21.This test statistic is consistent in the presence of intra-cluster heteroskedasticity (Baum, Schaffer and Stillman 2003).

22.The chi-squared statistic for the hypothesis that the coefficients sum to zero is 0.12 with 1 degree of freedom. The probability of a higher test statistic is 0.7257.

23.The Durbin-Wu-Hausman test statistic is 6.59 with 2 and 334 degrees of freedom. The probability of a greater test statistic is 0.0016.

24. "The calculation of standard errors is based on the 'delta method'." (StataCorp 2003, 225).

25.Relatively few beginning teachers also have an advanced degree. The regressions for beginning teachers and experienced teachers use fixed effects for each year of experience rather than the years of experience and experience squared.

26. The log of the statutory minimum wage is perfectly collinear with the fixed effects for years of experience and time, and therefore drops out of the first two models. All teachers with 20 or more years of experience share the same statutory minimum wage in any given year, so the log of the statutory minimum wage also drops out of the third model.

27. Using data on average salaries in Kentucky for the 1989-90 school year and a variety of specifications, Borland and Howsen (1992, 1993, 1996) estimate that the difference in average salaries between perfectly competitive and perfectly monopsonistic markets was less than $700 per year. The average teacher salary in Kentucky at that time was $26,292.

28.Borland and Housen (1993) used a switching regimes technique to search for a competitive threshold, but because they did not allow concentration to influence wages on both sides of the switchpoint, they could only detect either monopsony or rent sharing, not both.

Abbreviations

BLS: Bureau of Labor Statistics
CBSA: Core Based Statistical Area
CWI: Comparable Wage Index
GMM: Generalized Method of Moments
IV: Instrumental Variables
NCES: National Center for Education Statistics
OLS: Ordinary Least Squares
OMB: U.S. Office of Management and Budget
TEA: Texas Education Agency
2SLS: Two Stage Least Squares

Table 1
The Distribution of the Herfindahl Index of School Competition in Texas

	Number of Observations	Mean	Standard Deviation	Minimum	Maximum
1999-2000					
All CBSAs	225,627	0.1811	0.1752	0.0691	1.0000
Metropolitan areas	207,052	0.1510	0.1353	0.0691	0.9846
Micropolitan areas	18,575	0.5168	0.2147	0.2160	1.0000
2000-2001					
All CBSAs	233,076	0.1788	0.1753	0.0668	1.0000
Metropolitan areas	214,270	0.1493	0.1362	0.0668	0.9843
Micropolitan areas	18,806	0.5154	0.2147	0.2130	1.0000
2001-2002					
All CBSAs	239,659	0.1750	0.1730	0.0657	1.0000
Metropolitan areas	220,898	0.1464	0.1349	0.0657	0.9843
Micropolitan areas	18,761	0.5117	0.2122	0.2180	1.0000
2002-2003					
All CBSAs	242,960	0.1706	0.1712	0.0645	1.0000
Metropolitan areas	224,918	0.1429	0.1330	0.0645	0.9841
Micropolitan areas	18,042	0.5150	0.2145	0.2196	1.0000
2003-2004					
All CBSAs	246,364	0.1689	0.1705	0.0629	1.0000
Metropolitan areas	228,055	0.1413	0.1324	0.0629	0.9749
Micropolitan areas	18,309	0.5123	0.2136	0.2200	1.0000

Table 2
The Effect of Competition on Teacher Salaries

	Pooled Cross-Section Models			**Labor Market Fixed Effects**		
	OLS	**IV**	**GMM**	**OLS**	**IV**	**GMM**
Herfindahl	-0.316	-0.530	-0.564	-1.680	-2.852	-3.038
	(0.061)**	(0.092)**	(0.069)**	(0.330)**	(0.494)**	(0.480)**
Herfindahl, squared	0.290	0.516	0.593	1.505	3.412	3.353
	(0.069)**	(0.101)**	(0.081)**	(0.321)**	(0.699)**	(0.682)**
Tax price	-0.013	-0.006	-0.001	-0.033	-0.041	-0.035
	(0.023)	(0.025)	(0.020)	(0.020)	(0.025)	(0.024)
Average district enrollment	0.002	0.002	0.002	0.007	0.008	0.009
	(0.001)	(0.001)*	(0.001)**	(0.002)**	(0.002)**	(0.002)**
Average district enrollment, squared	-0.000	-0.000	-0.000	-0.000	-0.000	-0.000
	(0.000)	(0.000)*	(0.000)**	(0.000)**	(0.000)**	(0.000)**
Average income (log)	0.005	-0.050	-0.122			
	(0.065)	(0.064)	(0.054)*			
Percent school-age	0.505	0.625	0.462	0.576	-0.022	0.586
	(0.222)*	(0.269)*	(0.217)*	(0.420)	(0.527)	(0.456)
Percent over 65	-0.015	-0.061	-0.048	2.510	2.455	2.401
	(0.166)	(0.178)	(0.166)	(0.496)**	(0.587)**	(0.572)**
Percent high school graduates	0.166	0.250	0.319			
	(0.125)	(0.119)*	(0.099)**			
Percent college graduates	0.053	0.202	0.214			
	(0.135)	(0.154)	(0.124)			
Unemployment Rate	-0.003	-0.004	-0.006	0.002	0.000	0.000
	(0.003)	(0.003)	(0.002)*	(0.002)	(0.002)	(0.002)
Comparable wage index	0.133	0.054	0.163	0.238	0.180	0.112
	(0.085)	(0.097)	(0.079)*	(0.054)**	(0.059)**	(0.054)*
Micropolitan area	0.078	0.083	0.073			
	(0.021)**	(0.021)**	(0.015)**			

Number of Obs.	1,187,693	1,187,693	1,187,693	1,187,693	1,187,693	1,187,693
R^2	0.8776	0.8759	0.8731	0.8860	0.8857	0.8857
Prob > F, β_H=0, β_{H2}=0	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
Prob > F, $\beta_H + \beta_{H2}$=0	0.2569	0.7257	0.3115	0.1854	0.1332	0.3637

Note: The dependent variable is log of full-time-equivalent monthly salary. All specifications also include indicator variables for each school year, district characteristics (distance from the center of the closest metro area, indicators for small and midsized districts and student demographics (percent low income, limited English proficient, black and Hispanic)) and an array of individual characteristics (gender and ethnicity, years of experience and its square, percent time teaching in a field for which the teacher was certified, and indicators for educational attainment (no college degree, MA and Ph.D.), classroom assignment (math, science and special education), assignment to a high school, and first year in the district). All but the OLS specifications treat the Herfindahl index and its square as endogenous, using the determinants of school location and uneven enrollment growth as instruments (see text). The robust standard errors (in parentheses) are clustered by labor market for the first three models, and by market-years for the three models with CBSA fixed effects. * significant at 5%; ** significant at 1%.

Table 3
Sensitivity Analysis, Selected Coefficients from GMM Models with Labor Market Fixed Effects

	Full Sample	Excluding Dallas and Houston	Beginning Teachers	Experienced Teachers	Highly Experienced Teachers	Experienced Teachers without Advanced Degrees	Experienced Teachers with Advanced Degrees
Herfindahl	-3.035	-1.908	-1.851	-3.042	-2.546	-2.753	-3.545
	(0.481)**	(0.467)**	(0.443)**	(0.483)**	(0.502)**	(0.577)**	(0.539)**
Herfindahl, squared	3.348	2.238	2.492	3.470	3.318	3.084	4.102
	(0.683)**	(0.656)**	(0.521)**	(0.683)**	(0.735)**	(0.851)**	(0.783)**
Model includes labor market characteristics	yes	yes	yes	yes	yes	yes	yes
Model includes teacher and district covariates?	yes	yes	yes	yes	yes	yes	yes
Model includes year indicators?	yes	yes	yes	yes	yes	yes	yes
Model includes CBSA fixed effects?	yes	yes	yes	yes	yes	yes	yes
Number of Observations	1,187,693	702,997	297,582	637,762	252,349	631,299	258,812
Number of Clusters	335	325	335	335	335	335	335
Shea's partial R^2, Herfindahl	0.3881	0.3705	0.4005	0.3827	0.3861	0.3803	0.3813
Shea's partial R^2, Herfindahl^2	0.2288	0.2026	0.2696	0.2204	0.2098	0.2211	0.2000
Prob > Hansen's *J* statistic	0.0901	0.1426	0.0639	0.0776	0.1052	0.0872	0.1115
Centered R^2	0.8857	0.8843	0.7366	0.7988	0.6783	0.8664	0.8471
Prob > F, $\beta_H=0$, $\beta_{H2}=0$	0.0000	0.0002	0.0000	0.0000	0.0000	0.0000	0.0000
Prob > F, $\beta_H + \beta_{H2}=0$	0.3672	0.2687	0.4983	0.1572	0.0373	0.2149	0.0434

Note: Selected coefficient estimates from GMM specifications identical to those in the last column of Table 2, but estimated on the designated sub-samples. Cluster-adjusted robust standard errors in parentheses. * significant at 5%; ** significant at 1%.

Table 4
Texas Minimum Monthly Teacher Salary

Years of Experience	1999-2000 through 2001-02	2002-03	2003-04
0	$2,424	$2,524	$2,474
1	2,481	2,581	2,531
2	2,539	2,639	2,589
3	2,596	2,696	2,646
4	2,717	2,817	2,767
5	2,838	2,938	2,888
6	2,959	3,059	3,009
7	3,072	3,172	3,122
8	3,178	3,278	3,228
9	3,279	3,379	3,329
10	3,373	3,473	3,423
11	3,464	3,564	3,514
12	3,549	3,649	3,599
13	3,628	3,728	3,678
14	3,705	3,805	3,755
15	3,776	3,876	3,826
16	3,844	3,944	3,894
17	3,908	4,008	3,958
18	3,968	4,068	4,018
19	4,026	4,126	4,076
20	4,080	4,180	4,130

Table 5
Selected Coefficients from Tobit Models with Market Fixed Effects

	Beginning Teachers	Experienced Teachers	Highly Experienced Teachers
Herfindahl	-2.233	-2.860	-1.502
	(0.106)**	(0.065)**	(0.101)**
Herfindahl, squared	2.781	3.638	2.088
	(0.150)**	(0.098)**	(0.155)**
Model includes teacher covariates?	yes	yes	yes
Model includes district covariates?	yes	yes	yes
Model includes year indicators?	yes	yes	yes
Model includes labor market fixed effects?	yes	yes	yes
Standard errors adjusted for clustering?	no	no	no
Number of Observations	297,582	611,719	278,392

Note: Selected coefficient estimates from Tobit specifications identical to those in Table 4.

Table 6
Selected Coefficients from Linear, Cross-Sectional Analyses of Aggregate Data
2003-04

	CBSA Averages	**District Averages**
Herfindahl	-0.089	-0.033
	0.022**	0.020
Model includes teacher covariates?	no	no
Model includes district covariates?	yes	yes
Model includes labor market covariates?	yes	yes
Number of Observations	67	669
Shea's partial R^2, Herfindahl	0.8881	0.8492
Prob > Hansen's *J* statistic	0.2182	0.0907
Centered R^2	0.9066	0.6067

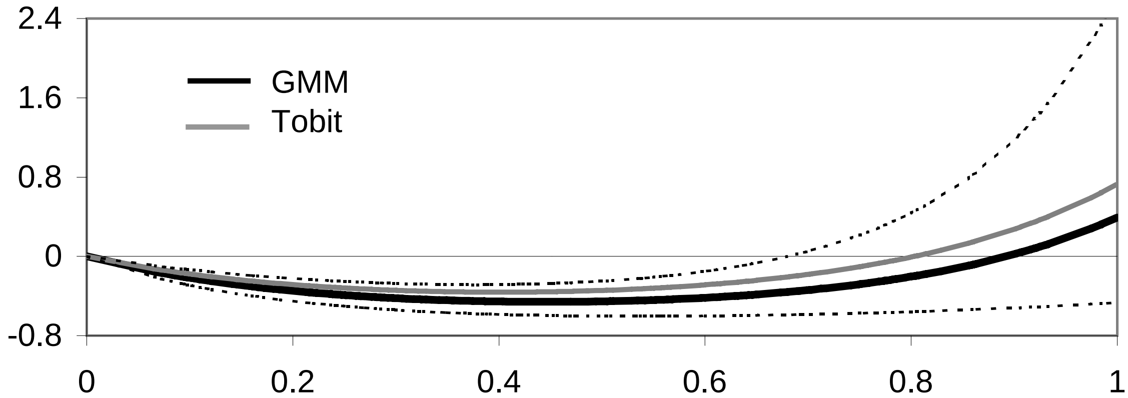
Note: The dependent variable is log of average, full-time-equivalent monthly salary. Both specifications also include the full set of district and labor market covariates from Table2. Both specifications treat the Herfindahl index as endogenous, using the same instruments as table 2. Both regressions are GMM models weighted by the number of teachers. The robust standard errors (in parentheses) are clustered by labor market. * significant at 5%; ** significant at 1%.

Table 7
The Pattern of Oligopsony and Rent Sharing

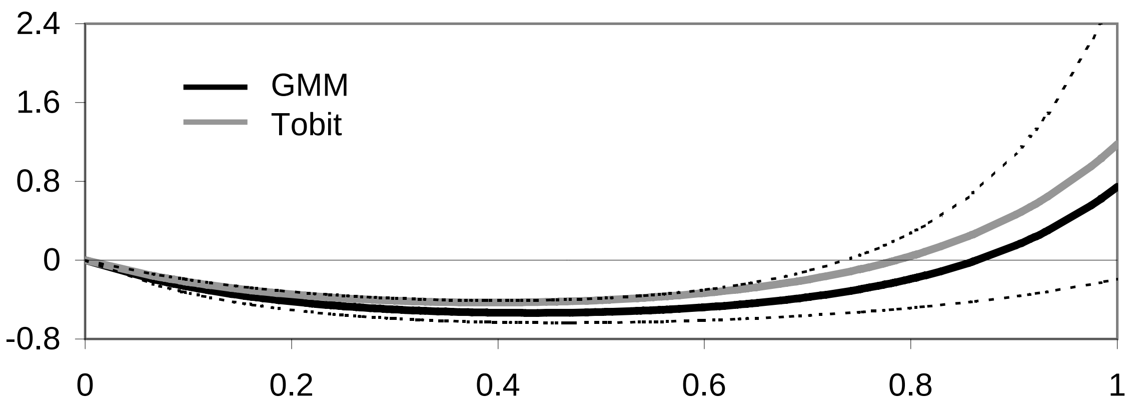
	Herfindahl Index						
	0.00-0.26	0.26-0.34	0.34-0.37	0.37-0.49	0.49-0.54	0.54-0.69	0.69-1.00
Beginning Teachers	O	O					R
Experienced Teachers	O	O	O			R	R
Highly Experienced Teachers	O				R	R	R

O– Oligopsonistic Market
R– Rent-Sharing Market

Beginning Teachers



Experienced Teachers



Highly Experienced Teachers