PROBATION LENGTH AND TEACHER SALARIES: DOES WAITING PAY OFF?

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Tenure policies for elementary and secondary school teachers is a controversial issue in many states, but there is virtually no empirical evidence on how tenure affects teacher labor markets. This paper begins to fill this research void by using cross-state variation in tenure policies to identify the effects, if any, of the length of the probationary period on entry-level teacher salaries. Using data from the Schools and Staffing Survey, the authors investigate whether districts in states with longer probationary periods offer higher wages to teachers as a way to compensate for the extended evaluative period. Results suggest that they do, although effects are concentrated in districts that are most likely to be competing for teachers with districts in neighboring states with shorter probation periods. The authors also find that the relationship between probation length and wages is stronger for experienced teachers and in districts that engage in collective bargaining.

Tenure for elementary and secondary school teachers is a contentious issue in many states. Originally intended to protect teachers from arbitrary or unfair firings, tenure[1] today is often seen as a barrier to improving teacher quality, making it impossible for principals to remove bad teachers from the classroom. Because of this perception, many states have considered reforms to their tenure policies, focusing largely on increasing the length of the probationary period and streamlining the dismissal or appeals process. One noteworthy example is California's Proposition 74, rejected by the voters in 2005, which would have increased the probationary period for teachers from two to five years. Supporters of Proposition 74 argued that a longer probationary period would give principals more time to assess teachers as well as to mentor struggling teachers; opponents argued that a longer probationary period was unnecessary and would only deter teachers from entering the profession.

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A data appendix with additional results, and copies of the computer programs used to generate the results presented in the paper, are available from Eric Brunner, Department of Economics, Quinnipiac University, 275 Mount Carmel Ave, Hamden, CT 06518; eric.brunner@quinnipiac.edu.

[1] Although some states do not explicitly use the word "tenure" in their statutes, every state has laws that govern the terms of employment and "due process" for teachers. These laws cover the length of the probationary period that a teacher must serve before receiving tenure, the allowable reasons for dismissal of a tenured teacher, and the process for dismissal and appeals.

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As the debate over Proposition 74 suggests, opponents of reforms to teacher tenure tend to focus on the reaction of *teachers* to policy changes. The question of how *districts* are likely to respond, however, is generally ignored. If districts fear that reforms to the tenure process will affect the supply of teachers, they, like teachers, may certainly change their behavior. For example, if a longer probationary period deters individuals from becoming teachers or from teaching in particular states because of the increased uncertainty of gaining tenure, districts may respond by offering higher wages to newly hired teachers. Indeed, there exists substantial evidence that districts *do* respond to market conditions by adjusting teacher salaries. Numerous studies have documented the existence of compensating wage differentials that are required to attract teachers to districts with particular characteristics, such as those with a large concentration of minority students or being located in a market with a high cost of living (e.g., Antos and Rosen 1975; Kenny and Denslow 1980; Levinson 1988; Stoddard 2005). Similarly, Angrist and Guryan (2008) found that teacher salaries are positively related to the stringency of testing requirements imposed upon teachers for initial certification.

Even though researchers have explored many factors that affect teacher salaries, no study to date has examined the relationship between the length of the probationary period and teacher salaries. The purpose of this paper is to fill that gap in the literature. Specifically, we use data from the Schools and Staffing Survey (SASS), a nationally representative sample of schools, districts, and teachers, to examine whether districts respond to longer probationary periods by offering higher salaries to newly hired teachers. To identify the impact of probation length on teacher salaries, we focus on districts located in metropolitan areas that cross state boundaries and estimate models that include metropolitan fixed effects. Our identification strategy therefore utilizes only within-metropolitan area variation in the length of the probationary period, which allows us to control for unobservable characteristics of teacher labor markets that may be correlated with the length of a state's probationary period. In addition to cross-sectional models, we estimate longitudinal models that exploit the fact that six states increased the length of their probation periods during the 1990s. Because the number of states that changed the length of their probationary period is small, the results from the longitudinal models should be viewed primarily as an important specification check of our cross-sectional results.

Background and Context

There already exists a large literature on the determinants of teacher salaries. Many researchers have focused on the compensating differentials that teachers may require in order to work in schools and districts with particular characteristics. For example, Antos and Rosen (1975) and Levinson (1988) both found that white teachers demand higher wages for teaching nonwhite students, whereas Kenny and Denslow (1980) and Stoddard (2005) emphasized the role of cost of living and area amenities. A number of studies have suggested that unionization is positively correlated with higher salaries (e.g., Lankford and Wykoff 1997; Babcock and Engberg 1999), and that communities with a high demand for education may offer higher teacher salaries in order to attract better teachers (Loeb and Paige 2000). Finally, communities with constraints such as tax and expenditure limits generally have lower salaries (Figlio 1997).

Despite this relatively large literature on the determinants of teacher salaries, no study to date has examined the relationship between the length of the probationary period and teacher salaries. What evidence does exist on the topic comes from the more general labor economics literature on the use of probationary periods by private sector firms. For example, Wang and Weiss (1998) developed a model in which firms monitor new hires during a probationary period. Their model suggests that firms requiring a longer probationary period (and therefore a longer period of monitoring) must offer higher wages in the post-probationary period to attract workers of the same quality; thus, wages in the post-probationary period are an increasing function of probation length. Consistent with that hypothesis, Groshen and Loh (1994) found that the post-probationary wages of workers are positively correlated with the length of the probationary period.

In terms of the teacher labor market, probation length may affect wages if teachers view longer probationary periods as increasing uncertainty and thus raising the costs of becoming a teacher. This is similar to the entry–barrier effect on wages analyzed by Angrist and Guryan (2008) in the context of teacher testing and teacher wages. Specifically, these researchers argued that certification requirements establish barriers to entry into the teaching profession. Consequently, more stringent teacher testing could manifest itself into higher teacher salaries. Based on a panel of school districts drawn from the 1987–1988, 1990–1991, 1993–1994 and 1999–2000 waves of the SASS, they found evidence consistent with that hypothesis.[2] The notion that certification requirements act as a barrier to entry was also supported by Hanushek and Pace (1995), who found that state licensing requirements (both courses and tests) significantly reduce the probability that a prospective teacher will graduate from college with a teaching degree.

Unlike teacher testing and other certification requirements, longer probationary periods do not necessarily impose direct costs on teachers. Nevertheless, all else being equal, longer probationary periods are associated with a longer monitoring period and prospective teachers may view these as increasing the amount of uncertainty associated with becoming a teacher. In that sense, longer probationary periods may impose costs on new teachers and act as barriers to entry in much the same way as state licensing requirements. To understand this more clearly, consider how individuals decide to teach in a particular district, j. When deciding to become a teacher, they incur some entry costs, C_{Ej}, such as taking specific classes, passing a licensing exam, or other costs in money or time and effort associated with acquiring a teaching credential. After they begin teaching and, after a set amount of time, they either acquire tenure, with probability P_j and earn the expected income stream, $\tilde{W}_{Tj}$, or they leave teaching and earn the expected income stream, $\tilde{W}_{Aj}$. We can also allow for the possibility of additional costs associated with the probation period, C_{Pj}. C_{Pj} may be thought of as the costs associated with the uncertainty of probation, such as the additional stress of being evaluated or the extra effort that probationary teachers may exert in order to increase their chances of gaining tenure. Following Ehrenberg, Pieper and Willis (1998) and Angrist and Guryan (2003), we assume workers choose teaching jobs to maximize expected utility, V_j thus:

$$(1) \qquad P_j^* U(\tilde{W}_{Tj} - C_{Ej} - C_{Pj}) + (1 - P_j)^* U(\tilde{W}_{Aj} - C_{Ej} - C_{Pj}) = V_j$$

With this simple representation, and assuming that utility is increasing in wages and decreasing in entry or probation costs, we can see that any factor that increases P_j, $\tilde{W}_{Tj}$, or $\tilde{W}_{Aj}$,[3] *ceteris paribus,* will increase V_j whereas any policy that increases C_{Ej} or C_{Pj} will decrease V_j. In equilibrium, we expect V_j to be equal across all districts for teachers of equal quality. For an individual who has chosen to enter teaching, V_j must also exceed the expected utility from alternative occupations.

In this framework, there are two avenues through which longer probations may affect expected utility. One possibility is that longer probations reduce the probability (or the perceived probability) of tenure, P_j; another is that longer probations equate to more time when teachers are being evaluated and may be feeling stressed by uncertainty or are required to put in extra effort, thus increasing probation costs, C_{Pj}. An individual district can offset either of

[2] Kleiner and Petree (1988) also examined whether state licensing requirements affect average teacher salaries. Their analysis reveals no clear relationship between the two variables.

[3] It should be noted that the desirability of tenure is implicitly represented by assuming > .

these possible effects by offering higher wages, in which case we would expect districts in states with longer probations to have higher salaries.

Although the theory outlined above suggests a positive relationship between the length of the probationary period and teacher salaries, there are several reasons why teacher salaries may be unrelated to or even negatively related to the length of the probationary period. First, Gordon, Kane and Staiger (2006) have argued that even though states allow districts to dismiss probationary teachers for almost any reason, they seldom do so.[4] Based on teacher-level data from the 1999–2000 SASS, they found that among teachers who left teaching or transferred to another district, less than one percent cited being laid off or involuntarily transferred as the reason. Among teachers with three years of teaching experience or less (probationary teachers), less than two percent reported that they left teaching or moved to another district due to a layoff or an involuntary transfer. If most teachers are aware that the probability of being dismissed is extremely low, they may not view a longer probationary period as increasing uncertainty or imposing any additional costs. Second, it could also be that in districts with shorter probation periods, teachers may be evaluated more often or with greater scrutiny, which could imply costs are actually higher for teachers facing *shorter* probationary periods. In that case, probation length may have little (or even a negative) effect on teacher salaries. Given these findings, whether or not longer probationary periods lead to higher teacher salaries remains an empirical question.

Data

We use data from the 1999–2000 Schools and Staffing Survey (SASS), a national sample of schools and teachers that contains information on individual teachers, schools, and districts.[5] Since teacher salary schedules are typically set at the district level, our primary unit of analysis is the school district. District-level data on teacher salaries comes from the survey of school district administrators, a subcomponent of the SASS. In the empirical work that follows, we focus on beginning teacher salaries, which is measured as the salary of a teacher with a bachelor's degree and no teaching experience.[6]

Information on each state's probation length was taken from Loeb and Miller (2007) and verified from state statutes.[7] Figure 1 shows how probation length varies across the country. The majority of states have a 3-year probationary period; however, eight states (plus the District of Columbia) have 2-year probation periods, five states require 4 years, and two states require 5 years.[8] Also, technically, Wisconsin allows the probation length to be determined by

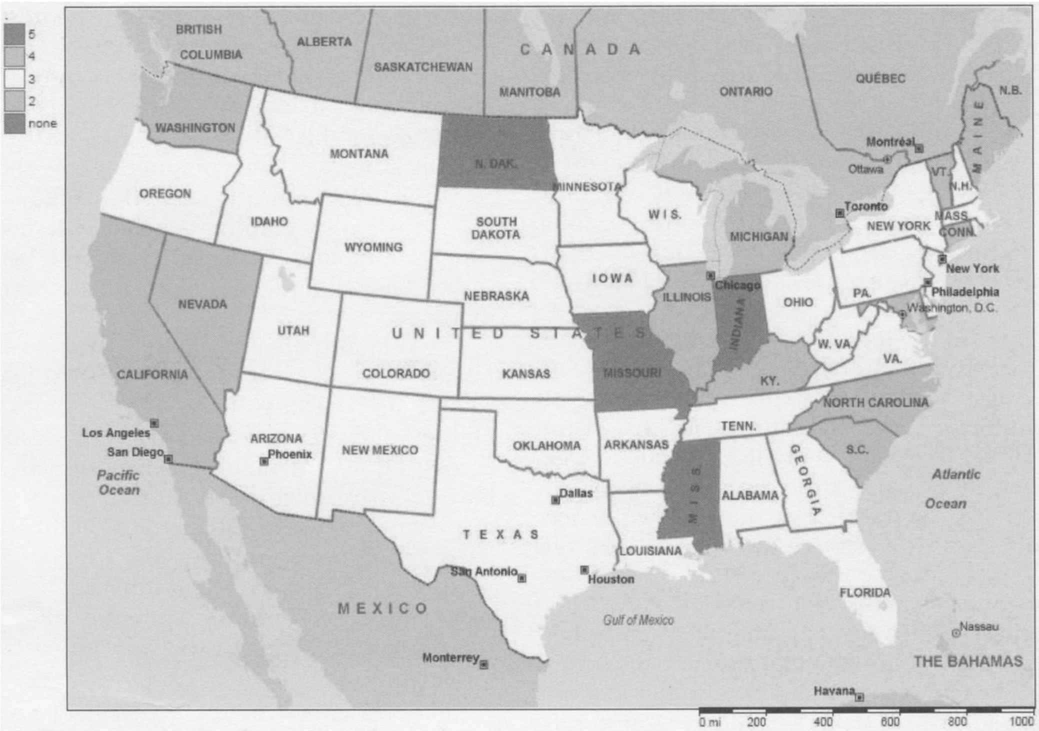
[4] Survey evidence supports this argument. Specifically, based on a 2007 survey of 1,010 K–12 public school teachers, Duffett et al. (2008: 3) found that 69% of teachers "say that when they hear a teacher at their school has been awarded tenure, they think that it's just a formality—it has very little to do with whether a teacher is good or bad."

[5] We focus on the 1999–2000 SASS because the timing corresponds perfectly with the district demographic characteristics measured by the 2000 Census. However, we also estimated all cross-sectional models with data from the 2003–2004 wave of the SASS, as well as a pooled sample with both the 1999–2000 and 2003–2004 waves. The results are generally consistent with those based on the 1999–2000 sample and are available upon request.

[6] We also conducted the analysis with experienced teacher salaries; however, if higher wages are offered to offset the uncertainty and costs associated with longer probations, we would expect the effect to be stronger for teachers nearer the beginning of their careers. Consistent with that expectation, the results for experienced teacher salaries are weaker than for beginning teacher salaries. Furthermore, given that all the districts in our sample are operating on a step-and-column salary schedule, there is most likely a "mechanical" relationship between salaries for beginning and experienced teachers (i.e., if beginning salaries are relatively high in a given district, experienced salaries are likely to be relatively high as well). Separating out the effects of probation on the level of salaries versus the return to experience is beyond the scope of this paper and thus, we focus solely on beginning salaries. Results for experienced salaries are available upon request.

[7] See Loeb and Miller (2007) for a full listing of the statutes. Since the policies in Loeb and Miller are for 2005 and the SASS data is from 1999-2000, we checked each state for any changes in the last several years.

[8] Indiana awards teachers "semi-permanent" status after two years but they are not considered "permanent" until they have completed five years of teaching.



each district, through collective bargaining. Because in practice almost all districts negotiate probation periods of 3 years, we treat Wisconsin as a 3-year state in our analysis. North Dakota and Mississippi do not specify a specific probation length and are thus excluded from our analysis.

We include a number of additional state-level characteristics in our empirical work to mitigate any correlation between the length of the probationary period and other state-level characteristics. The first three variables are indicator variables that take the value of unity if a state required (1) a basic skills test, (2) a subject skills test, and (3) either a teacher knowledge exam or a teaching assessment exam for initial teacher certification in 1999. We include these variables for two reasons. First, as noted previously, Angrist and Guryan (2008) found that teacher salaries are positively related to the stringency of teacher testing. Second, the testing requirements imposed upon teachers for initial certification may be correlated with the length of a state's probationary period. For example, states may respond to shorter probationary periods by requiring more stringent testing for initial teacher certification. Data on state-level teacher testing requirements in 1999 were obtained from the National Center for Education Statistics and the National Association of State Directors of Teacher Education and Certification (NASDTEC) Manual on Certification & Preparation of Educational Personnel (1999).

Using the 1999 SASS teacher sample, which identifies the undergraduate institution each teacher attended, we are also able to calculate the fraction of teachers in a state who graduated from a selective undergraduate institution.[9] We include this variable to control for systematic variation across states in teacher quality that might

[9] Following Reback (2004), we define selective institutions as those receiving one of Barron's top three selectivity ratings (very competitive, highly competitive and most competitive). We thank Randall Reback for sharing these data.

also be correlated with the length of the probation period.[10] For example, in states where residents have a strong preference for teacher quality, state policy may favor longer probationary periods even as districts, independently of this policy, offer higher wages.[11]

Following the literature on the determinants of teacher salaries, we include two additional state-level variables in our analysis: the log of median household income (obtained from the 2000 Census), and an indicator variable for whether a state had a potentially binding tax and expenditure limitation measure (TEL) in 1999.[12] We code a state as having a potentially binding TEL if any of the following apply: (1) it imposed a property tax rate limit *and* a limit on property tax assessment increases, (2) it imposed a property tax revenue limit, or (3) it imposed a limit on general revenue and expenditures. Information on states with potentially binding TELs is taken from Figlio (1997) and updated to 1999 using information on local tax and assessment limitation measures prepared by Mikhailov (1998) and Mullins (2004).

We also use information from the 1999–2000 SASS and the National Center for Education Statistics' Census 2000 School District Tabulation Data to create a number of district-level control variables. Those variables include (1) the fraction of students below the poverty level, (2) the fraction of students that are non-white, (3) an indicator variable for whether the district engages in collective bargaining, (4) the log of district enrollment, (5) the log of median family income, (6) the fraction of the population age 25 or older with a college degree, (7) an indicator for whether the district is located in a rural area, (8) an indicator for whether the district is located in the South, (9) an indicator for whether the district serves only elementary students, and (10) an indicator for whether the district serves only high school students. All of these variables are designed to capture either amenity and cost characteristics of districts that may affect the supply of teachers, or community characteristics that may affect the demand for teachers, either of which may influence teacher salaries. Finally, to control for systematic variation in the wage a teacher could earn in an alternative profession, we include the comparable wage index (CWI) prepared by the National Center for Education Statistics. The CWI measures regional variation in the salaries of college graduates who are not educators. The NCES constructs a CWI for four different geographic areas: school districts, labor markets, states, and combined regional areas. We use the 1999 district-level CWI in the empirical work that follows.

We restrict our sample in a number of ways. First, we limit the sample to local elementary, high school, and unified districts and drop charter schools, state-operated institutions, and other non-traditional districts. Second, we drop districts that do not utilize a teacher salary schedule and therefore do not report information on the salary of teachers with a bachelor's degree and no teaching experience. Finally, we exclude a small number of districts with missing Census data on fraction poverty, fraction college-educated and median household income. Table 1 provides the means of the variables used in our analysis, for the full sample and by length of the probationary period.

Empirical Framework

To examine the relationship between the length of the probationary period and teacher salaries, we begin by simply regressing the log of teacher salaries on a set of dummies that correspond to the various probation lengths, a set of state-level controls, and a set of district-level controls.

[10] We calculate our measure of teacher quality at the state level rather than the district level due to the survey design of the SASS. Specifically, the SASS teacher sample is designed to be representative at the state level but not at the district level.

[11] However, we note that with the example of teacher quality, the bias could presumably go either direction; supporters of longer probation periods argue that longer probations will increase teacher quality whereas opponents argue that longer probations decrease teacher quality. Since there is no empirical evidence for either argument, it remains unclear whether voters who desire higher teacher quality would support longer or shorter probation periods.

[12] Figlio (1997) found that binding tax and expenditure limitation measures are associated with lower cost-of-living adjusted salaries.

Table 1. Sample Means, by Length of Probationary Period

	Full Sample	*Two Years*	*Three Years*	*Four Years*	*Five Years*
Teacher Salaries					
BA Salary, No Experience	25,987	26,919	25,659	27,198	25,236
State Characteristics					
Length of Probationary Period	3.04				
Median H.H. Income	41,331	43,699	40,556	43,610	39,751
Binding TEL	0.67	0.43	0.72	0.60	1.00
State Basic Skills Tests	0.76	0.57	0.75	1.00	1.00
State Subject Skills Tests	0.63	0.57	0.59	0.80	1.00
State Other Tests	0.57	0.57	0.59	0.40	0.50
Selective College	0.21	0.24	0.20	0.19	0.18
District Characteristics					
Alternative Wage Index	0.90	0.94	0.88	0.94	0.86
Fraction Poverty	0.16	0.16	0.16	0.14	0.13
Fraction Non-White	0.25	0.35	0.25	0.21	0.13
Bargain	0.63	0.88	0.59	0.67	0.49
Enrollment	6,983	10,440	6,420	7,183	4,892
Elementary District	0.08	0.21	0.05	0.08	0.01
High School District	0.03	0.07	0.02	0.04	0.00
Rural	0.30	0.26	0.33	0.24	0.27
South	0.32	0.29	0.34	0.40	0.00
Median Family Income	47,481	49,663	46,371	51,956	46,606
Fraction College Educated	0.20	0.22	0.19	0.20	0.17
Number of States	46	7	32	5	2
Number of Districts	4,145	575	2,870	476	224

Specifically, we estimate a model of the following form:

$$(2) \quad lsalary_{js} = \beta_0 + \beta_1 Three_s + \beta_2 Four_s + \beta_3 Five_s + X_{js}'\delta + Z_s'\gamma + \alpha_m + \varepsilon_{js},$$

where $lsalary_{js}$ denotes log of teacher salaries in district j located in state s, $Three_s$, $Four_s$, and $Five_s$ are indicator variables that take the value of unity if the length of the probationary period in state s is 3, 4, or 5 years respectively (the omitted category is states with a 2-year probationary period); X_{js} is a vector of district-level control variables; Z_s is a vector of state-level control variables; and ε_{js} is a random disturbance term.

Note that in equation (2) the impact of probation length on teacher salaries is identified using all cross-sectional variation in length of probationary period. By estimating equation (2) using all districts, we are essentially asking whether, all else being equal, salaries in a state such as Colorado, which has a 3-year probationary period, differ from salaries in a state such as Maine, which has a 2-year probationary period. An obvious concern with this identification strategy is that it fails to take into account the regional nature of teacher labor markets. For example, Boyd et. al. (2005) and Reininger (2007) found that teachers restrict their job searches to relatively small geographic areas, which implies that teacher labor markets are geographically small in size. That leads to the concern that our parameter estimates may be biased by unobservable teacher labor market conditions that are correlated with teacher salaries and with the length of the probationary period.

To better account for the regional nature of teacher labor markets, we developed an alternative identification strategy that exploits variation *within* metropolitan areas (i.e., core based statistical areas (CBSAs)) to identify the impact of probation length on teacher wages.[13] Specifically, we estimate the following model:

$$(3)\quad lsalary_{jms} = \beta_0 + \beta_1 Three_s + \beta_2 Four_s + \beta_3 Five_s + X_{jms}'\delta + Z_s'\gamma + \alpha_m + \varepsilon_{jms},$$

where $lsalary_{jms}$ denotes the log salary of teachers in district j, located in metropolitan area m and state s, and α_m is a set of metropolitan-area fixed effects. The inclusion of metropolitan-area fixed effects implies that we are now identifying the impact of probation length on teacher salaries based solely on those metropolitan areas that cross state boundaries *and* contain states with different probationary periods. Restricting our attention to variation within metropolitan areas allows us more accurately to model the localized nature of teacher labor markets. In addition, the inclusion of metropolitan-area fixed effects allows us to control for any metropolitan-level unobservables that may be correlated with teacher wages or the length of the probationary period.[14]

In the empirical work that follows, we estimate the parameters of equations (2) and (3) using two separate samples: all relevant districts in the SASS sample and only those districts that engage in collective bargaining. We present separate estimates based on the sample of collective bargaining districts for several reasons. First, the literature on unionization and teacher salaries suggests that collective bargaining agreements may affect the structural determinants of teacher salaries. For example, Moore (1976) found that the elementary–secondary salary differential is smaller in districts that engage in collective bargaining, suggesting that teacher unions (like their private sector counterparts) tend to bargain for standardized wage policies. Similarly, Zwerling and Thomason (1995) and Ballou and Podgursky (2002) found that the returns to experience are higher in districts with collective bargaining agreements whereas Babcock and Engberg (1999) found that the returns to education and experience are positively related to the median level of teacher experience and the fraction of teachers with a master's degree in a collective bargaining unit. Second, and more fundamentally, Easton (1988) argued that when contracts are collectively bargained, salary and job characteristics comparisons across districts may play a more important role in negotiations. Consistent with that notion, Babcock, Engberg and Greenbaum (2005) discovered that in

[13] Our definition of metropolitan areas is based on the 2003 core-based statistical areas (CBSAs), developed by the Office of Management and Budget. A number of CBSAs are part of larger statistical areas known as Combined Statistical Areas (CSAs). CSAs comprise several CBSAs that have strong commuting and employment linkages. Whenever a CBSA is part of a larger CSA, we use the CSA as our measure of a metropolitan area. The one exception is the New York-Newark-Bridgeport, NY-NJ-CT-PA CSA, which is the largest CSA in the country and spans several hundred miles from northern Connecticut to central New Jersey and Pennsylvania. The vast size of this CSA makes it unlikely that it corresponds to local teacher labor markets. Consequently, for the New York-Newark-Bridgeport, NY-NJ-CT-PA CSA we continue to define "teacher labor markets" in terms of the CBSAs that comprise the larger CSA.

[14] In our metropolitan-area fixed-effects model, the effect of probation length is identified based on the 358 districts that are located within a metropolitan area that crosses state boundaries and contains states with different probation lengths. One might wonder whether the characteristics of these districts differ systematically from the characteristics of districts in metropolitan areas that do not cross state boundaries. To address that question, we calculated separate summary statistics for districts in metropolitan areas that cross state boundaries and for districts in metropolitan areas that do not cross state boundaries. In general, the two sets of districts look similar. However, teacher salaries in districts in metropolitan areas that cross state boundaries tend to be slightly higher on average and the comparable wage index also tends to be slightly higher. These districts also tend to have higher family incomes, as well as slightly lower fractions of non-white students and students in poverty.

districts with "high" union strength, the characteristics of comparison districts referred to by the union during the negotiation process affect negotiated outcomes. Thus, if teacher unions use probation length as a bargaining tool, the results of Babcock et al. (2005) suggest that the impact of length of probation on teacher salaries may be larger in districts that engage in collective bargaining.

Results

Results based on the estimation of Equation (2) are presented in Table 2. All of the models we estimate are weighted using the SASS district sampling weights. In addition, all standard errors are clustered at the state level to allow for within-state autocorrelation of the disturbance term. Column 1 reports results based on the full sample of SASS districts whereas column 2 reports results based on the sub-sample of districts that engage in collective bargaining. Overall, these results provide little evidence that length of probation affects teacher salaries: in no case are any of the estimated coefficients on the probation length dummies statistically significant.

It is worth noting, however, that many of the estimated coefficients on our control variables are statistically significant, and their signs are generally consistent with the previous literature on the determinants of teacher salaries. For example, consonant with Taylor (2010) and Rose (2007), both specifications suggest that teacher wages are positively related to the wages earned by workers in other professions (alternative wage index). Similarly, consistent with prior research, teacher wages are positively related to district enrollment, median household income and the fraction of students that are non-white.[15]

As noted in above, the OLS results fail to take into account the regional nature of teacher labor markets. In Table 3, therefore, we present the results for our metropolitan-area fixed-effects specifications. Our core results are reported in columns 1 and 2. Column 1 presents results based on the full sample of SASS districts that are located within a CBSA while column 2 presents results based on the subsample of districts that also engage in collective bargaining. In the interest of brevity, Table 3 contains only the coefficients for the three probation dummy variables; however, we note that all specifications include all the control variables included in Table 2.[16]

When compared to the OLS results, the metropolitan area fixed-effects results are relatively striking—columns 1 and 2 show that the estimated coefficients on all three probation dummies are positive and statistically significant. Furthermore, the estimated coefficients reported in there display a consistent pattern. Beginning teacher salaries tend to increase monotonically with the length of the probationary period. For example, the results in column 1 suggest that relative to states with 2-year probationary periods, beginning teacher salaries are 6.1%, 6.7%, and 7.8% higher in states with 3-, 4- and 5-year probationary periods, respectively. Finally, a comparison between column 1 and column 2 reveals that probation length has a much greater impact on teacher salaries in districts that engage in collective bargaining (i.e., relative to the estimates for the full sample, the coefficients in column 2 are all larger in magnitude).[17] That

[15] See for example Antos and Rosen (1975), Eberts and Stone (1986) and Zwerling and Thomason (1995).

[16] The coefficients on the control variables are qualitatively similar across specifications and generally consistent with the previous literature on teacher salaries.

[17] One may be concerned about how attributes of school districts differ in the sample that contains all CBSAs and the subsample that contains only those districts that engage in collective bargaining. In general, the two samples tend to be similar with one big exception: most school districts located in southern states do not engage in collective bargaining. Thus, the collective bargaining subsample contains substantially fewer Southern districts. To examine how this restriction affected our results, we also estimated models where we drop districts located in the South. For the full sample of CBSAs located outside the south, the estimated coefficients on the probation dummies increase in magnitude compared to the estimates reported in column 1. This is not surprising given that we have effectively dropped a large fraction of the districts that do not engage in collective bargaining, so the results tend to mirror the collective bargaining results reported in column 2. However, we still find that even in non-Southern districts, collective

Table 2. Ordinary Least Squares Estimates
(Dependent Variable: Log Beginning Teacher Salary)

	(1) All Districts	(2) Only Collective Bargaining Districts
Three-Year Probationary Period	-0.007	-0.007
	(0.035)	(0.039)
Four-Year Probationary Period	-0.011	0.011
	(0.053)	(0.061)
Five-Year Probationary Period	0.023	0.024
	(0.046)	(0.057)
Log Alternative Wage Index	0.386**	0.360**
	(0.060)	(0.073)
Fraction Poverty	0.116**	0.115
	(0.055)	(0.083)
Fraction Non-White	0.078**	0.111**
	(0.025)	(0.035)
Bargain	0.012	
	(0.013)	
Log District Enrollment	0.015**	0.017**
	(0.004)	(0.005)
Elementary District	-0.015	-0.021
	(0.014)	(0.015)
High School District	0.017	0.021
	(0.024)	(0.022)
Rural	-0.006	-0.004
	(0.005)	(0.006)
South	0.022	-0.021
	(0.032)	(0.046)
Log District Median Income	0.054**	0.068**
	(0.026)	(0.029)
Fraction College-Educated	0.041	0.016
	(0.056)	(0.057)
Log State Median H.H. Income	0.428**	0.462**
	(0.093)	(0.127)
Basic Skills Test	0.010	0.018
	(0.018)	(0.021)
Subject Matter Test	-0.011	-0.021
	(0.031)	(0.035)
State Other Tests	0.067**	0.073**
	(0.021)	(0.034)
Binding TEL	-0.042	-0.065*
	(0.029)	(0.035)
Selective College	-0.109	-0.036
	(0.115)	(0.185)
Observations	4145	2615
R^2	0.602	0.630

Notes: Robust, clustered standard errors are in parentheses.
*Statistically significant at the .10 level and **at the .05 level.

Table 3. Metropolitan Area Fixed Effects Estimates
(Dependent Variable: Log Beginning Teacher Salary)

	Districts Located in MSA		*Districts Located in MSA or Within 30 Miles of an MSA*	
	(1)	*(2)*	*(3)*	*(4)*
	All Districts	*Only Collective Bargaining Districts*	*All Districts*	*Only Collective Bargaining Districts*
3-Year Probationary Period	0.061**	0.097**	0.056**	0.095**
	(0.022)	(0.040)	(0.020)	(0.034
4-Year Probationary Period	0.067*	0.231**	0.057*	0.214**
	(0.036)	(0.048)	(0.034)	(0.046)
5-Year Probationary Period	0.078**	0.250**	0.066**	0.231**
	(0.030)	(0.057)	(0.028)	(0.054)
Observations	3135	2087	3308	2168
R^2	0.820	0.813	0.823	0.818

Notes: Robust, clustered standard errors are in parentheses. All specifications include the full set of control variables listed in Table 2.
*Statistically significant at the .10 level and **at the .05 level.

probation length has a much larger impact on teacher salaries in districts that engage in collective bargaining suggests that teachers in these districts are more successful at using the length of the probation period as a bargaining tool during negotiations over teacher salaries. This is generally consistent with Babcock et al. (2005), who found that union strength increases the importance of union-backed salary and job characteristic comparisons across districts to negotiated contracts.

The results reported in columns 1 and 2 consistently suggest that beginning teacher salaries are positively related to the length of the probation period. One concern with those results, however, is that the number of observations used to identify the impact of probation length on teacher salaries is relatively small. Recall that in our metropolitan-area fixed-effect specifications we are identifying the impact of probation length on teacher salaries based solely on those metropolitan areas that cross state boundaries *and* contain states with different probationary periods. For the full sample of districts located in a CBSA (column 1), the effect of probation length is identified, based on 32 observations in 2-year probation states and 150, 105, and 71 observations in 3-, 4-, and 5-year probation states, respectively. Similarly, for the subsample of districts that also engage in collective bargaining, the effect of probation length is identified, based on 25 observations in 2-year probation states and 113, 74, and 26 observations in 3-, 4-, and 5-year probation states, respectively.

To address concerns about the relatively small number of identifying observations and to provide a robustness check of the results reported in columns 1 and 2, we re-estimated our metropolitan-areas fixed-effects specifications using an expanded sample of districts. Specifically, of the 4,145 districts in the SASS sample, approximately 25 percent are located outside a CBSA. However, many of those districts are located in a county that borders a CBSA. To expand our sample size, we first used the longitude and latitude of all districts in the SASS sample to measure the distance from the centroid of a district to the centroid of the closest CBSA. We then defined districts as belonging to a CBSA if (1) the district is located in a county that is part of a CBSA

bargaining agreements tend to increase the strength of the relationship between teacher pay and probation length. Results are available upon request.

or (2) the district is located in a county that borders a CBSA and is within 30 miles of the centroid of the CBSA. Furthermore, for districts located outside a CBSA but within 30 miles of the centroid of a CBSA, we restrict the sample to include only those districts classified as non-rural.

Using this alternative definition of metropolitan-areas substantially increases the number of identifying observations. For the full sample of districts located in or near a CBSA, the effect of probation length is now identified based on 61 observations in 2-year probation states, and 191, 127, and 91 observations in 3-, 4-, and 5-year probation states, respectively. For the subsample of districts that also engage in collective bargaining, the effect of probation length is now identified based on 54 observations in 2-year probation states and 134, 82, and 32 observations in 3-, 4-, and 5-year probation states, respectively.

Results based on this alternative definition of metropolitan areas are reported in columns 3 (full sample) and 4 (collective bargaining subsample). Similar to the results in columns 1 and 2, the estimated coefficients on the probation dummies for the expanded sample are all positive and statistically significant, and the magnitude of the estimated coefficients once again consistently increases with the length of the probation period. Also note that the estimated coefficients on the probation length dummies in columns 3 and 4 tend to be slightly smaller in magnitude than the corresponding estimates in columns 1 and 2. This is perhaps not too surprising given the fact that teacher labor markets tend to be highly localized; as we expand the definition of a teacher labor market to incorporate a larger area, we are most likely straining the definition of the size of an actual teacher labor market. Nevertheless, the fact that the results for the expanded sample are quite similar to our core results is reassuring and provides additional evidence that beginning teacher salaries are positively related to the length of the probation period.[18]

State Fixed-Effects Specifications

Up to this point, all of our specifications attempt to identify the impact of probation length on teacher salaries by exploiting cross-sectional variation. One remaining concern is that even with metropolitan-area fixed effects there may still be state-level unobservable variables that are correlated with both probation length and teacher salaries, thus creating bias in our estimates. Our results may then simply reflect this correlation rather than a compensating differential for teachers. This concern is perhaps mitigated by the fact that we find a stronger relationship between probation length and teacher salaries in the fixed-effects specifications than in the OLS specification; that is, because probation length is a state-level policy, any bias created by state-level preferences would presumably affect all districts, not just those close to other states. Nevertheless, in this section, we exploit the fact that six states changed the length of their probationary period during the 1990s in order to estimate models that rely on both cross-sectional and temporal variation in teacher salaries and probation length.[19] Specifically, we use the 1990, 1993, 1999, and 2003 waves of the SASS to estimate pooled cross-sectional models that include both time and state fixed effects.

[18] We also estimated models that exploit variation within state borders to identify the impact of probation length on teacher salaries. Specifically, we restricted our sample to include only those districts within 30 miles of a state border and created a set of border fixed effects that take the value of unity for all districts on either side of a particular state border. We then estimated models nearly identical to equation (3) except we replaced the metropolitan-area fixed effects with border fixed effects. Results based on this specification were qualitatively similar to those reported in Table 3. Specifically, for the sample of all districts located within 30 miles of a state border, the estimated coefficients on all three probation length dummies were positive but only the estimated coefficient on the 3-year probation dummy was statistically significant. For the sub-sample of districts that also engaged in collective bargaining, the estimated coefficients on all three probation length dummies were positive and statistically significant and the magnitude of the estimated coefficients once again consistently increased with the length of the probation period. Results are available upon request.

[19] The six states are Connecticut (3 years to 4 years in 1996), Iowa (2 years to 4 years in 1998), Illinois (3 years to 4 years in 1998), Michigan (2 years to 4 years in 1993), North Carolina (3 years to 4 years in 1997) and Pennsylvania (2 years to 3 years in 1996).

The inclusion of state fixed effects implies that we are only using *within*-state variation to identify the impact of probation length on teacher salaries. Thus, the models we estimate in this section control for any state-specific time-invariant unobservables that may be correlated with the length of the probationary period.

The pooled cross-sectional models we estimate are similar in spirit to our metropolitan-area fixed-effects models. Specifically, our pooled cross-section model takes the following form:

$$(4) \quad lsalary_{jsmt} = \beta_0 + \beta_1 L_{st} + X_{jsmt}'\delta + Z_{st}'\gamma + \alpha_s + \eta_{mt} + \varepsilon_{jsmt},$$

where $lsalary_{jsmt}$ denotes the natural log of teacher salaries in district j, in state s, in metropolitan area m, in year t; $L_{s,t}$ denotes either the length of the probationary period in state s, in year t, or a set of probationary length dummies; α_s is a set of state fixed effects, and η_{mt} is a set of metropolitan-specific time effects. The inclusion of these time effects in equation (4) implies that we are now controlling for any within-metropolitan area arbitrary time trends in teacher labor market conditions that may be correlated with changes in the length of probationary periods. Furthermore, the inclusion of both state fixed effects and metropolitan-specific time effects implies we are now identifying the impact of probation length on teacher salaries using only those metropolitan areas that cross state boundaries *and* contain a state that changed its probationary period during the 1990s. For example, consider the Chicago metropolitan area which contains districts located in Illinois, Indiana, and Wisconsin. In 1998, Illinois increased the length of its probationary period from two years to four years while the probationary periods in Indiana and Wisconsin remained unchanged. Thus, districts in the 1999 and 2003 wave of the SASS that are located in the Illinois portion of the Chicago metropolitan area represent our "treatment" group whereas districts in the Indiana and Wisconsin portion of the Chicago metropolitan area represent our control group.

We estimate the parameters of equation (4) by pooling data from the 1990, 1993, 1999, and 2003 waves of the SASS. From these four waves, we obtained data on the salary of teachers with a bachelor's degree and no teaching experience, district enrollment, fraction of non-white students, and whether a district engaged in collective bargaining.[20] We then merged that data with the same state-by-year and district-by-year control variables listed in Table 1.[21]

Table 4 reports results based on the estimation of our pooled cross-sectional models. Once again, all specifications are weighted using the SASS district sampling weights and all standard errors are clustered at the state level to allow for within-state autocorrelation of the disturbance term.[22]

[20] The 1990 wave of the SASS does not provide information on whether a district engaged in collective bargaining. Therefore, we obtained data from the 1987 Census of Governments on whether or not school district employees' were covered by a collective bargaining agreement and matched the districts in the 1990 wave of the SASS to the 1987 Census of Governments data.

[21] We obtained state-by-year data on teacher testing requirements for initial certification from the National Center for Education Statistics and state-by-year data on states with potentially binding TELs from Figlio (1997), Mikhailov (1998) and Mullins (2004). For 1990 and 1993 we use state median household income from the 1990 census; for 1999 and 2003 we use state median household income from the 2000 census. To construct district-level estimates of median household income, fraction poverty, and fraction college-educated in 1990 and 1993, we use district-level data from the Special School District Tabulations of the 1990 Census. To construct estimates of those same variables for 1999 and 2003 we use district-level data from the Special School District Tabulations of the 2000 Census. The National Center for Education Statistics only provides data on comparable wages going back to 1997. Consequently, in 1990 and 1993 we use the 1997 comparable wage index to control for systematic variation in the wage a teacher could earn in an alternative profession. For 1999 and 2003 we use the actual comparable wage index for those years.

[22] In the interest of brevity, we once again report only the estimated coefficients on the probation length variables but note that all models include the full set of control variables listed in Table 1, with one exception. Information on the selectivity of a teacher's undergraduate institution was not collected in all waves of the SASS. Consequently, we omitted the state-level teacher quality variable from the analysis. We note, however, that an examination of the teacher quality variable for other years reveals that the percentage of teachers who graduate from selective colleges within a state is relatively stable over time. Specifically, the

Table 4. Coefficient Estimates for Pooled 1990-03 SASS Samples with State Fixed Effects (Dependent variable: Log Beginning Teacher Salary)

		(1) *State Fixed Effects*	*(2)* *MSA-Year Fixed Effects and State Fixed Effects*
		All Districts	
A.	Length of Probationary Period	0.008	0.018**
		(0.008)	(0.007)
B.	3-Year Probationary Period	0.001	0.024*
		(0.020)	(0.012)
	4-Year Probationary Period	0.018	0.035**
		(0.016)	(0.015)
	Observations	16007	1712
	R^2	0.86	0.91
		Only Collective Bargaining Districts	
C.	Length of Probationary Period	0.010	0.020**
		(0.010)	(0.009)
D.	3-Year Probationary Period	−0.001	0.011
		(0.024)	(0.020)
	4-Year Probationary Period	0.018	0.043*
		(0.016)	(0.025)
	Observations	9430	1073
	R^2	0.86	0.9

Notes: Robust, clustered standard errors are in parentheses. Column 1 includes year and state fixed effects and column 2 includes metropolitan/year fixed effects and state fixed effects.
*Statistically significant at the .10 level;**at the .05 level.

The first column of Table 5 reports results based on a specification in which we simply include state fixed-effects and exclude the metropolitan-specific time trends; the second column reports results based on the metropolitan-area/year fixed-effects and state fixed-effects specification given by equation (4).

Panels A (all districts in sample) and C (sub-sample of districts that engage in collective bargaining) report results based on a specification in which we simply include a probation length variable that takes on the values of 2, 3, 4, or 5 in year *t*. In column 1 (state fixed effects only), the estimated coefficients on the probation length variable are positive but small in magnitude and never statistically significant. In contrast, in our preferred specification (column 2), the estimated coefficients on the probation length variable are positive, much larger in magnitude, and statistically significant at the 5-percent level in both Panels A and C. The fact that the estimated coefficients in column 1 are smaller in magnitude and statistically insignificant is perhaps not too surprising given that this is the specification that is most likely to suffer from omitted variable bias due to unobservable teacher labor market conditions that vary across time and market.

Panels B and D of Table 4 report results based on specifications where we replace the probation length variable with a set of dummy variables for whether a district was located in a state with either a 3-year or a 4-year probationary period in year *t*. Note that since no state changed its probationary period to 5 years, and no 5-year probationary

correlation between the teacher quality measure for states over the years 1993, 1999, and 2003 is 0.96. Thus, most of the variation across states in teacher quality will be captured by the state fixed effects in equation (4).

state changed its probationary period, we do not include a 5-year probation dummy (since it would be perfectly collinear with the state fixed effects).[23] Similar to the results reported in Panels A and C, all but one of the estimated coefficients on the probationary period dummies are positive. The one exception is the estimated coefficient on the 3-year probationary period dummy in Panel D, column 1, which is negative but small in magnitude and statistically insignificant. Furthermore, in our preferred specifications reported in column 2, three of the four estimated coefficients on the probation dummies are statistically significant, the one exception being the 3-year probation dummy in Panel D. Similar to Table 3, the magnitudes of the estimated coefficients on the probation length dummies reported in column 2 increase monotonically with the length of the probationary period. Thus, the results from these state fixed-effects models continue to suggest that longer probationary periods are associated with higher beginning teacher salaries.

The fact that many of the estimated coefficients on the probationary dummies remain statistically significant in the state fixed-effects specifications is relatively surprising given that only six states actually changed their probationary periods during the 1990s. Specifically, one might expect all of the estimated coefficients on the probationary dummies to lose significance given the relatively small amount of variation in probation length in this analysis. Thus, the fact that many of the estimated coefficients in Table 4 remain statistically significant provides surprisingly strong evidence to support the notion that longer probationary periods are associated with higher teacher salaries.

Conclusion

In discussions of teacher tenure reform, extending the length of the probationary period is often suggested as one way to increase teacher quality. Supporters of such a policy typically argue that longer probations allow principals to do a better job of screening teachers before awarding tenure. Opponents of longer probations counter that the added uncertainty may discourage qualified individuals from entering teaching in the first place. One way that districts may offset this uncertainty is by offering higher wages to new teachers. The debates on this issue, however, have rarely considered the financial costs of a change in probation length, and there is no empirical evidence on how districts respond to differences in probation length. In this paper, we offer the first evidence that districts may, indeed, react to differences in probation length by adjusting salaries.

Our core analysis of the relationship between probation length and teacher salaries is based on cross-sectional data on beginning teacher salaries from the 1999–2000 wave of the Schools and Staffing Survey. To control for the localized nature of teacher labor markets, we exploited the fact that numerous metropolitan areas cross state boundaries and a number of those areas contain states with different probationary periods. We exploited this within-metropolitan area/across-state variation in probation length and estimated models that include metropolitan-area fixed effects. Results based on this identification strategy suggest that salaries for beginning teachers are measurably higher in districts whose states require longer probationary periods. Furthermore, the effect of probation length on salaries is particularly strong in districts that engage in collective bargaining.

We examined the robustness of our results by exploiting the fact that six states changed the length of their probationary period during the 1990s and by estimating

[23] Recall that in the metropolitan-area/year fixed-effects model we are identifying the impact of probation length on teacher salaries using only those metropolitan areas that cross state boundaries and contain a state that changed its probationary period. For the full sample of districts located in a CBSA (top panel), the 3- and 4-year probation dummies are identified based on 109 and 146 observations respectively, whereas the 2-year probation period control group is identified based on 233 observations. Similarly, for the subsample of districts that also engage in collective bargaining (bottom panel), the 3- and 4-year probation dummies are identified based on 87 and 124 observations respectively, whereas the 2-year probation period control group is identified based on 178 observations.

pooled cross-sectional models using data from the 1990, 1993, 1999, and 2003 waves of the SASS. Once again, based on a variety of specifications that include both time and state fixed effects, we find that longer probationary periods are generally associated with higher teacher salaries.

Finally, our results highlight the importance of the local nature of teacher labor markets. State policymakers considering proposals to increase the length of teacher probationary periods should be aware that districts closer to neighboring states with shorter probations will likely bear costs that may not be felt as strongly by districts elsewhere in the state. This may be particularly true if those districts also engage in collective bargaining.

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