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THE EFFECT OF UNIONIZATION ON FACULTY SALARIES 1978-1996: A TEST OF EMPIRICAL METHODS

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ABSTRACT

This article analyzes the union/nonunion differential for faculty compensation. Using a unique data set that includes information on public academic institutions over two time periods, the analysis examines whether the union premium changed over time. Various model specifications proposed by previous researchers are also tested. The estimated outcomes suggest: a) that the union/nonunion differential increased substantially in the 1990s compared to the late 1970s and early 1980s. Estimates suggest that unionized faculty experienced a 5 percent advantage in the 1970s-80s, and a 13 percent advantage in the 1990s; b) the model specification is somewhat sensitive to controls for heterogeneity and endogeneity, but less so for the unobserved individual effect typical with panel data. The growth in the union premium may be important as the legal structure underlying faculty unionization becomes more inclusive for private institutions and part-time faculty.

Since the 1970s, the unionization rate for higher education has grown dramatically. Despite setbacks in the private sector, the change in the underlying legal structure of collective bargaining in public universities and colleges permitted many faculty groups desiring better compensation or a higher level of governance to freely organize [1]. However, empirical research investigating unionization in higher education has found mixed evidence that collective bargaining increases compensation. This article examines the union effect on faculty compensation

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from 1978 to 1996 to explore why the evidence has been mixed and to investigate whether collective bargaining has an economic effect on faculty compensation.

The mixed empirical results may be due to several causes. First, it may be that the time period of study is correlated with external factors that either dampen or increase the union effect. Union wage theories focus on spillover and threat effects. While spillover effects tend to increase the union/nonunion salary gap (unions raise salaries and lower employment in the union sector, increasing employment and lowering salaries in the nonunion sector), the threat effect causes nonunion organizations to raise the salaries of their employees to avoid unionization. The overall unionization impact depends on the size of each effect and the interaction of labor supply and demand in both sectors. Union service models suggest that union power is dampened when the costs of unionization is high [2]. These theories suggest that the size of the union/nonunion compensation gap may be sensitive to business cycles, as well as when markets are "buyer" or "seller" dominated, because unions have relatively little power when markets are not tight.

A second problem is related to the changing methodology used to examine the faculty compensation differentials. Some researchers attribute their relatively smaller union effects to the advancement of statistical techniques that control for the heterogeneity of faculty salaries and for the endogeneity of the union variable [3, 4]. Compared to research outcomes from the 1970s, more recent studies may be finding small differentials because researchers have developed advanced techniques to control for a variety of statistical problems that lead to overestimation of the union effect.

This article examines the union-nonunion compensation differential over an 18-year time span to investigate whether the time factor and/or the statistical technique affects the size and direction of the union effect. The investigation uses selected years of panel data on institutions of higher learning from 1978 to 1996, collected by the National Center for Education Statistics (NCES) and compiled for the National Science Foundation [5, 6], union information from the *Directory of Faculty Contracts and Bargaining Agents in Institutions of Higher Education* [7, 8], and school classifications as provided by *Peterson's Register of Higher Education* [9]. Included in the analysis are cross-sectional and panel data estimation, as well as an investigation into the effect of controlling for heterogeneity, endogeneity, and unobserved institution effects.

FACULTY UNIONIZATION AND THE UNION-NONUNION COMPENSATION DIFFERENTIAL

As mentioned earlier, the unionization rate for higher education has grown since the 1970s. Hemmasi and Graf related the following statistics: Collective bargaining began in full force in the late 1960s, and between 1969 and 1979 the number of academic institutions with unionized faculty grew from 24 to 227 [10].

By 1990, the number of institutions under collective bargaining agreements grew to more than 1,000, which indicates a growth rate of more than 300 percent during the 11-year period. By 1997, nearly 37 percent of all faculty in the United States were unionized [11]. The rules that restricted collective bargaining in higher education in the 1960s were relaxed by many states over the past 30 years, particularly for public sector faculty. Prior to 1970, only five states provided a legal foundation for some form of collective bargaining for faculty (Missouri, New York, Oregon, Rhode Island, and Wisconsin). By 1997, legislation in 34 states and the District of Columbia permitted collective bargaining for higher education [8].

The above trend has occurred despite a setback to faculty collective bargaining due to the *Yeshiva* decision of 1980 [12]. In the *Yeshiva* case, the university successfully argued before the Supreme Court that its faculty members were managers, a category of workers not covered by the National Labor Relations Act [12]. This decision has been invoked by 38 schools wishing to decertify a faculty union or to prevent a union organizing effort [8] and has generally stifled the unionization of faculty at private institutions since 1980.

The growth of faculty unions is also counter to the declining unionization trend for the United States. Bacharach, Schmidle, and Bauer summarize the explanations given to account for the diametric trend unionization has taken in higher education [1], but one primary reason found empirically is economic—faculty groups join unions to increase their economic well-being.[1] Although economics may be a driving force for faculty unionization, contradictory evidence exists on whether unions actually increase faculty compensation. Economic theory indicates that, on average, collective bargaining increases wages for individuals who are unionized. As described by Freeman and Medoff, the collective voice of unions increases worker power in the marketplace [14]. The power of the collective voice translates into better wages and fringe benefits for unionized employees, not only because of the direct effect of union power, but also because unions reduce employment in their sectors, leading to negative spillover effects on wages and salaries in nonunion sectors.[2]

The union wage effect in the private sector has been estimated at anywhere from 10 percent to 15 percent [15, 16]. Early studies on faculty collective bargaining also found positive, albeit relatively smaller union effects on compensation [17-19]. However, some studies on faculty unionization suggest the union-nonunion compensation differential to be at most 1 percent to 4 percent or even

[1] Hemmasi and Graf found the desire for better salaries to be one of the primary reasons that faculties vote for collective bargaining. Williams and Zirkel reviewed the literature from the 1970s and concluded that compensation issues are a primary force in faculty unionization [13].

[2] Unionization also can indirectly increase nonunion wages because employers institute union-type benefits to deter their workers from voting for unionization. This effect is usually small compared to the combined outcome of the direct union effect and the spillover effect.

negative [3, 4, 20]. In particular, the work of Rees [4] and Kesselring [3] suggests that a regression model of average compensation must account for the issues of observation-level effects or the endogeneity of the unionization variable. Earlier models that examined higher education compensation failed to do so.

In using institution data from 1970 through 1988 from the American Association of University of Professors (AAUP), Rees assumed that an unobserved school-specific predilection for unionization biases the union coefficient upward [4]. He employed a fixed-effects model to control for both time and school-specific effects and found a negative, but statistically insignificant, relationship between an institution's union status and its average total compensation. A later study by Rees, Kumar, and Fisher on Canadian universities found that the union-nonunion compensation differential was about 3 percent [21].

Kesselring followed a popular approach in estimating the union-nonunion compensation differential by assuming that union status is endogenous [3]. Using 1981 data on 471 doctoral-granting universities, he followed Lee [22] and estimated a switching regression model of faculty compensation. His results indicated that, *ceteris paribus,* schools under collective bargaining agreements pay about 16 percent less than their nonunion counterparts.

Counter to these results are those presented by Monks [23]. Using data on 8,198 individual faculty from the *1993 National Survey of Postsecondary Faculty* (NSOPF), Monks found a 7 percent to 14 percent salary differential that is robust to various model specifications, including a control for self-selectivity. Monks' result was larger than the 2 percent to 4 percent differential found with individual-level data for 1977 [24], which may be due to better sampling related to his 1993 data. To explain why his results differed from those of Kesselring and Rees, Monks argued that results may vary when using institution-level versus individual-level data. This point may be valid, as most studies that use faculty-level observations, even recent ones, suggest a positive, if small, union effect. The Kesselring and Rees studies used institutional data and found a zero or negative effect.

ESTIMATION OF THE UNION-NONUNION COMPENSATION DIFFERENTIAL FOR FACULTY

Description of the Data Used in the Analysis

This analysis uses data for public institutions from two sets of time periods, 1978-1985 and 1989-1996. Focus on the public sector is appropriate for these time periods because unionization of private sector institutions was severely curtailed by the 1980 *Yeshiva* decision, as mentioned earlier. The data include information on faculty compensation (salaries plus benefits), school-level faculty composition, tuition, financial statistics, and enrollment.

The reasoning behind two panel sets is to allow for time-related comparisons. First, as noted earlier, union effects may be time-varying. In finding small or negative union effects, Kesselring [3] and Rees [4] used data from the 1980s; in fact, Kesselring used only 1981 cross-sectional data. However, the 1978-1985 period reflects a major recession in the middle of the period and a decline in demand for faculty. The 1989-1996 period reflects an economic decline in the early part of the data, but a boom period in the latter part. Using the two panels may indicate whether business cycles are related to union premiums and allows for a time gap between the two panels (1986-1988) for comparative purposes. The two time periods also allow for comparisons to the other studies from the same time periods, particularly Kesselring [3], Rees [4], and Monks [23].

The National Center for Education Statistics (NCES) collects information on institutions of higher education over many years. The Computer-Aided Science Policy and Research (CASPAR) Database System, developed for the National Science Foundation, provides data on institutions of higher education from 1978 through 1988. The NCES Web site provides annual data after 1988. Note that during the 1980s the NCES adopted a policy of collecting some of the faculty data biennially, leaving data gaps in 1984 and 1987. Panel 1 uses data from 1978 through 1985, excluding the year 1984. The exclusion of one year in the panel does not affect the consistency of the parameter estimates, and while adding more observations would likely enhance efficiency, the large size of the final panel is sufficient to render this issue negligible.

Panel 1 (1978 to 1985) includes observations with full data for at least one year of analysis, leading to 3,214 observations for 582 institutions. The decision to use an uneven panel was made to ensure that the final sample captured the general population of public universities. An even panel would eliminate too many institutions from the analysis, particularly two-year schools.[3]

The second panel is from the NCES Web site. Omitting observations with missing data on key variables, the sample in this panel is 5,072 observations for 634 institutions. In this case an even panel was employed because there appeared to be some mismeasurement of the total compensation outlays. An even panel ensured that institutional salary data from one year to the next were reasonably related.[4] In addition, I wanted to test for unobserved school-specific effects, and a balanced panel makes for a better sample in this regard.

[3] The analysis was also conducted on the even panel. Despite the fact that this panel has a lower percentage of two-year schools than is representative of the population, the estimation did not substantially affect the final outcome in regard to the differential.

[4] In compiling the data, I noticed that there were a number of observations where the average compensation for an institution fell or rose by more than 50 percent. Therefore, to ensure that my data reflected actual average compensation for a school in a particular year, I eliminated any school where the average compensation fell or rose by more than 50 percent in any one year. I also ran the analysis on an uneven panel (where I included schools with at least two "clean" years of data), but some average compensation values were still suspect, and I preferred the current balanced panel.

The NCES data were matched to union information contained in the *Directory of Faculty Contracts and Bargaining Agents in Institutes of Higher Education* [7, 8] and for the first time period, to the institutional classification of two- or four-year status as presented in *Peterson's Register of Higher Education* [9]. Summary statistics of the sample are provided in Tables 1a and 1b.

Several differences emerged from a review of Tables 1a and 1b. The unionization rate increased from 33.7 percent to 36.0 percent. This growth rate is consistent with factual evidence that indicates a slow but steady growth in the 1980s and 1990s [24]. Unionized schools pay a higher unconditional average compensation for both time periods, and the union/nonunion compensation gap rose from an average 7.5 percent for 1978-85 period to 13 percent for the 1989-96 period. My first concern was that the increasing gap was due to an oversampling of two-year schools in the second time period, and these institutions may possess greater union effects. However, Figure l presents evidence that the compensation gap (in 1982-84 constant dollars) closed between 1989 and 1991, but then grew substantially from 1992 forward. A tightening of the differential also appears during the 1978-1979 and 1981-1982 school years. This latter trend has been noted in numerous studies about faculty salaries [1]. Real compensation for both groups has remained relatively flat over the time period, with little fluctuation. While not definitive, it appears as though the union-nonunion compensation gap closed slightly during recessionary periods and jumped substantially only between 1992 and 1993. Generally, union gains are tempered by slow growth in faculty compensation overall.

The percentage of faculty that was tenured declined between the two time periods (65 percent to 48 percent), likely due to the large number of senior faculty retirements in the 1990s. The percentage tenured for unionized faculty remained relatively higher in both time periods. The time periods also reflect the growing female faculty population; however, nonunion schools have a higher average percentage of females by the 1990s. The other noticeable difference between union and nonunion institutions is that nonunionized schools tend to be located in right-to-work states, which may reflect geographic differences in attitudes regarding the unionization of a professional group.

The Basic Regression Model of Faculty Compensation

This investigation began by estimating a basic salary regression with ordinary least squares (OLS) for each time period, then proceeding with models on each of the two panels that: a) account for heteroskedasticity, and b) assume that union status and compensation are simultaneously determined. A separate analysis to account for school-specific error was performed on the 1989-96 period.

To estimate salaries, the natural log of an institution's average total compensation for an academic year was used as the dependent variable. This value includes the total dollar value of all salaries and fringe benefits for full-time instructional faculty.[5] This variable, as well as all dollar values in the analysis, is adjusted by the 1982-84 CPI.

The independent variables include controls for school characteristics, financial characteristics, regional effects, and year effects.

School characteristics are as follows:

% Tenured: The percentage of faculty who are tenured. Earlier research [4, 25] found that average faculty compensation is positively related to a higher percentage of senior faculty. Senior tenured faculty are typically paid more than untenured faculty because of the positive effects of progress through the merit and promotion process.

% Female: The percentage of female faculty. This variable is included because generally female faculty are paid less on average than their male counterparts [20, 26]. Many factors can account for this disparity, including the fact that seniority is lower for female faculty and because discrimination may exist.

Enrollment: The number of full-time-equivalent students (in thousands) enrolled in the fall semester of an academic year. This variable captures the size of the school. Larger employers tend to pay higher salaries than smaller ones [28], and recent work on academic institutions finds a similar relationship in the higher education market [4].

Student-Faculty Ratio: Enrollment divided by the number of full-time faculty. The student-faculty ratio proxies for the workload of the average professor. The theory of wages suggests that increased work effort is positively correlated with a higher wage.

Union: Union is equal to 1 if school I is represented by a collective agent in time t and 0 if the school is not represented.[6]

School Type: School type is a binary variable that is equal to 1 if the college or university is a two-year institution and 0 if the institution is a four-year or graduate school. Many two-year and community colleges are included in this analysis. They typically pay much lower salaries, in part because faculty research quality is relatively lower on average than in four-year or graduate institutions.

Three financial variables are also included:

Tuition: The fall tuition fee (in thousands of dollars) represents the price that can be charged for an education. The direction of the relationship is

[5] A separate analysis of salaries and fringe benefits would answer some very interesting questions, such as, “Do unions provide better fringes?” Unfortunately, many of the schools in the NCES data set lack a breakdown of compensation; therefore, the analysis focuses only on total compensation.

[6] As with this study, other studies [3, 4, 20] used different editions of the *Directory of Faculty Contracts,* but assigned union status to schools once faculty had voted successfully for a bargaining agent.

Table 1a. Descriptive Statistics of Schools 1978-85

Variable	Description	All schools	Union	Nonunion
Average salary	Total salary expenditure/ No faculty	25,228.80 (4,656.28)	26,450.90 (5,089.64)	24,607.70 (4,289.56)
Ln average salary	Ln of average salary	10.120 (0.179)	10.166 (0.184)	10.096 (0.172)
Enrollment (1000s)	Enrollment (1000s)	7.588 (7.812)	7.695 (6.962)	7.433 (8.210)
Student/faculty ratio	Enrollment to faculty ratio	36.944 (18.517)	38.870 (18.484)	35.965 (18.4610)
Percent tenured	Percent of tenured faculty	65.462 (19.040)	71.390 (16.966)	62.450 (11.660)
Percent female	Percent of female faculty	30.417 (11.233)	30.414 (10.348)	30.419 (11.619)
Tuition (1000s)	Fall tuition (1000s)	0.811 (0.409)	0.919 (0.3760)	0.756 (0.4140)
Net revenues	Revenues-(Expenditure-Inst. Exp) in millions	14.449 (27.948)	11.129 (18.7780)	16.136 (31.4740)
Percent revenue	(Revenues-Inst. Exp)/ Revenues*100	77.299 (6.249)	76.376 (6.737)	77.768 (5.932)
Difference salary	School versus State average salary	–2.013 (14.665)	–1.529 (15.801)	–2.259 (14.049)
RTW	1 if right-to work state	0.384 (0.486)	0.106 (0.308)	0.525 (0.499)
Type	1 if a 2-year school	0.507 (0.500)	0.540 (0.499)	0.491 (0.500)
Union	1 if a unionized school	0.337 (0.473)		
Union*Type	Union*School type	0.182 (0.386)		
Union*PctFemale	Union*Percent female	10.248 (15.581)		
Sample size		3214	1083	2131

Data Source: The National Center for Education Statistics. Note that the data for 1978 to 1985 represents an unbalanced panel. Standard deviation in parentheses.

Table 1b. Descriptive Statistics of Schools 1989-96

Variable	Description	All schools	Union	Nonunion
Average salary	Total salary expenditure/ No faculty	26,984.87 (6,069.97)	29,182.48 (6,002.05)	25,746.52 (5,479.92)
Ln average salary	Ln of average salary	10.178 (0.226)	10.260 (0.210)	10.132 (0.221)
Enrollment	Enrollment (1000s)	7.387 (7.461)	7.999 (6.756)	7.042 (7.810)
Student/faculty ratio	Enrollment to faculty ratio	44.908 (24.428)	51.654 (27.839)	41.107 (21.362)
Percent tenured	Percent of tenured faculty	48.064 (35.097)	62.912 (32.365)	39.696 (33.788)
Percent female	Percent of female faculty	40.271 (11.420)	38.670 (9.480)	41.172 (12.289)
Tuition	Fall tuition (1000s)	1.105 (0.669)	1.355 (0.713)	0.964 (0.599)
Net revenues	Revenues-(Expenditure-Inst. Exp) in millions	21.782 (46.995)	17.399 (31.364)	24.252 (53.687)
Percent revenue	(Revenues-Inst. Exp)/ Revenues	60.706 (8.431)	61.209 (8.470)	59.815 (8.288)
Difference salary	School versus State average salary	–0.033 (0.204)	–0.022 (0.171)	–0.040 (0.221)
RTW	1 if right-to work state	0.492 (0.500)	0.153 (0.360)	0.683 (0.465)
Type	1 if a 2-year school	0.739 (0.439)	0.778 (0.415)	0.717 (0.450)
Union	1 if a unionized school	0.360 (0.480)		
Union*Type	Union*School type	0.281 (0.449)		
Union*PctFemale	Union*Percent female	13.937 (19.420)		
Sample size		5072	1828	3244

Data Source: The National Center for Education Statistics. Standard deviation in parentheses.

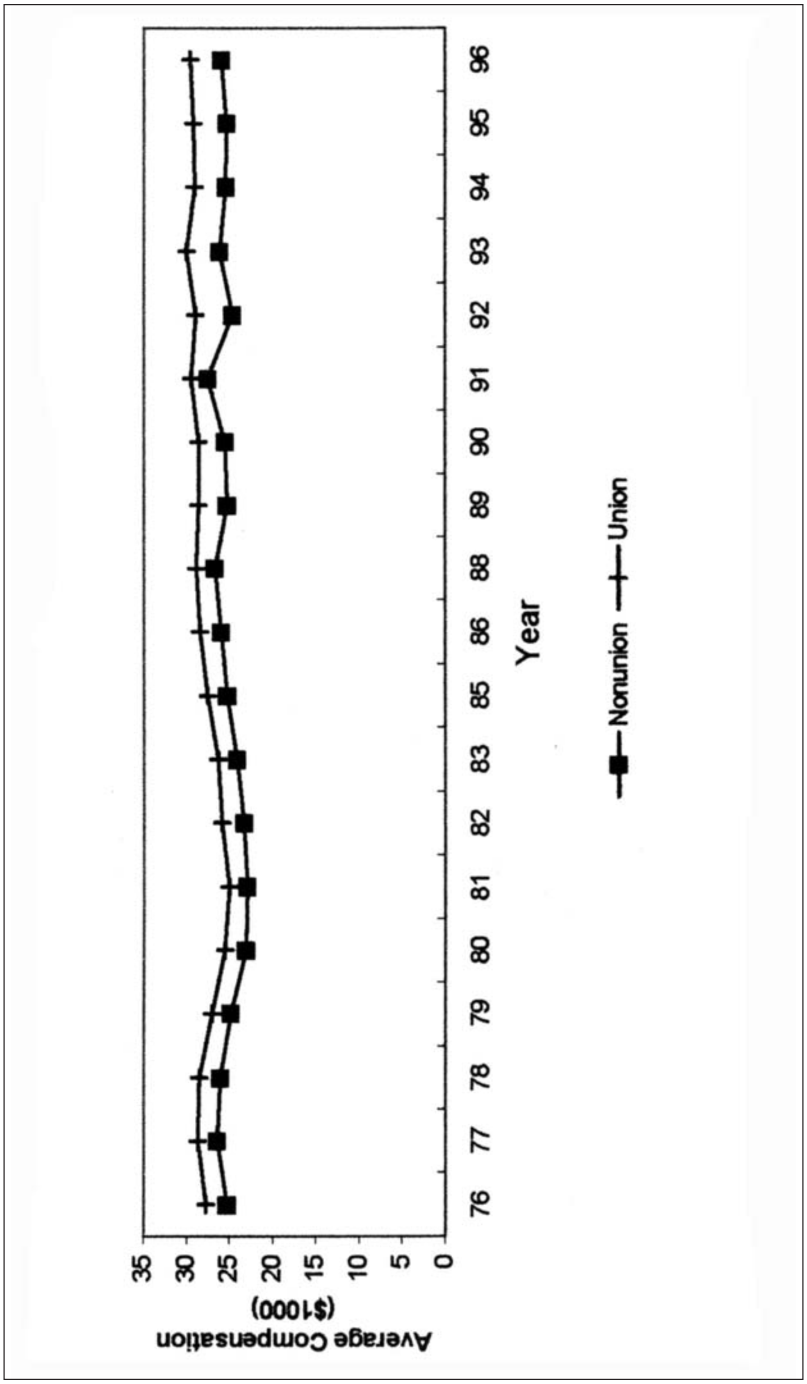

Figure 1. Average compensation by union status NCES data: 1976-1996.

ambiguous. High tuition may reflect school quality (and indirectly faculty quality, suggesting a positive relationship). Tuition may also be high (low) because other funding sources are low (high) (and therefore salaries just reflect the state's ability to pay at a particular time).

Net Revenues: Current fund revenues (total tuition and fees, government subsidies, and private gifts) less current fund expenditures, excluding instructional expenditures, are closely tied to the dependent variable. Net revenues indicate the size of the institution's excess funds once it pays for the daily upkeep of the institution sans instruction. The larger this value is, the more there is to share with faculty.

% Revenues: The percentage of revenues available after instructional expenditures are covered captures the financial viability of the school once it has covered its main annual expenditures on instruction.

It is expected that all three of the financial variables will be positively related to faculty salaries.

Regional dummy variables are included to control for cost-of-living differences across geographic locations. Finally, year dummy variables are included due to the above-mentioned changes in faculty average compensation over the time frame being investigated. Additionally, the union variable is interacted with School Type and percent female because evidence exists that collective bargaining appears more effective in making compensation gains for two-year schools than for four-year and graduate schools [20] and for senior female faculty [26].

EMPIRICAL ESTIMATION OF THE UNION-NONUNION COMPENSATION DIFFERENTIAL

Cross-Section Analysis

Following Rees [4], OLS estimation of the log of average compensation was conducted for each year of the analysis. Table 2 presents the union-related coefficient estimates for each of the 15 years.[7] Estimated effects for two- and four-year schools are presented.[8] The union compensation differential for two-year schools demonstrates some annual fluctuation, rising to 7.9 percent in 1980, but falling to almost zero by 1985, suggesting some business-cycle effect. The differential is 3.7 percent in 1989, falls during 1991 recession and remains at about 4.5 percent thereafter. The union effect for four-year schools hovers between 4 percent and 5 percent from 1978 to 1985, rising slightly by 1985 to 4.9 percent. Unionized faculty in four-year schools have a 14 percent premium in 1989, and

[7] The OLS results are available from the author.

[8] The estimated effect is $\beta_{union} + \beta_{ufem}$*Ave. Percent Female + β_{utype}*Utype. β_{utype} falls out for four-year schools.

Table 2. Union Effect: Cross-Sectional Analysis

	β	s_β	β	s_β
	1978		**1989**	
Union	0.051	(0.034)	0.137**	(0.059)
Union*School type	0.009	(0.027)	–0.103**	(0.040)
Union*Percent Female	–0.0001	(0.0003)	0.0001**	(0.002)
Union Effect-2Yr.	0.053		0.037	
Union Effect-4 Yr.	0.045		0.140	
	1979		**1990**	
Union	0.074*	(0.046)	0.107**	(0.056)
Union*School type	–0.010	(0.030)	–0.068***	(0.038)
Union*Percent Female	–0.001	(0.001)	–0.0003	(0.002)
Union Effect-2Yr.	0.036		0.030	
Union Effect-4 Yr.	0.047		0.098	
	1980		**1991**	
Union	0.095**	(0.036)	0.063	(0.053)
Union*School type	0.031	(0.029)	–0.144***	(0.038)
Union*Percent Female	–0.002	(0.001)	0.003*	(0.001)
Union Effect-2Yr.	0.079		0.022	
Union Effect-4 Yr.	0.048		0.166	
	1981		**1992**	
Union	0.086	(0.038)	0.132**	(0.059)
Union*School type	0.009	(0.029)	–0.056	(0.038)
Union*Percent Female	–0.001	(0.001)	–0.001	(0.002)
Union Effect-2Yr.	0.054		0.052	
Union Effect-4 Yr.	0.044		0.107	
	1982		**1993**	
Union	0.072*	(0.036)	0.063***	(0.045)
Union*School type	0.004	(0.003)	–0.029	(0.029)
Union*Percent Female	–0.001*	(0.001)	0.0004	(0.001)
Union Effect-2Yr.	0.044		0.050	
Union Effect-4 Yr.	0.040		0.080	
	1983		**1994**	
Union	–0.073***	(0.036)	0.122**	(0.064)
Union*School type	–0.010***	(0.027)	–0.086***	(0.038)
Union*Percent Female	–0.001***	(0.001)	0.0003	(0.002)
Union Effect-2Yr.	0.032		0.046	
Union Effect-4 Yr.	0.042		0.132	

Table 2. (Cont'd.)

	β	s_β	β	s_β
	1985		**1995**	
Union	0.085***	(0.035)	0.194***	(0.065)
Union*School type	–0.042	(0.026)	–0.099***	(0.043)
Union*Percent Female	–0.001	(0.001)	0.0003**	(0.002)
Union Effect-2Yr.	0.007		0.046	
Union Effect-4 Yr.	0.049		0.132	
			1996	
Union			0.095**	(0.073)
Union*School type			–0.082**	(0.047)
Union*Percent Female			0.001	(0.002)
Union Effect-2Yr.			0.044	
Union Effect-4 Yr.			0.144	

Data Source: The NCES. Note that data on 1984 were not available. Union effects are estimated by β_{union} + β_{ufem}Ave. Percent Female + β_{utype}*Utype. Level of significance (α) is as follows: * = .10; ** = .05; *** = .01.

even with some fluctuation in the 1990s, the 14 percent union premium remains in 1996. Interestingly, a 16.6 percent union differential for four-year schools in 1991 came at a time when average faculty salaries experienced a sharp decline in real value [27]. These cross-sectional estimates suggest that unions at two-year schools have less ability to weather economic changes, but unions at four-year schools have some ability to keep compensation relatively steady, regardless of the economic situation.

These results are somewhat consistent with other research. Rees [4] notes that earlier work supports an annual 5 percent union effect during the late 1970s; he also finds about a 5 percent to 6 percent differential that grows during 1985 to 1987; however, his sample includes private schools, the inclusion of which would reduce the size of the union premium.[9]

Panel Data Analysis

To test for specification differences, Table 3 presents the coefficient estimates for several models related to the pooled data sets. Columns (1) and (3) present OLS estimates for comparison purposes. Columns (2) and (4) address the problem of heterogeneity. Schools with many instructors will likely have a broader range

[9] Rees also does not estimate separate union effects for two- and four-year schools. It may be that his 5 percent effect has variation by school type.

Table 3. Regression-Pooled Samples

	1978-1985		1989-1996		
Variable	OLS (1)	WLS (2)	OLS (3)	WLS (4)	Random effects (5) (weighted)
Constant	10.314*** (0.035)	10.402*** (0.035)	10.471*** (0.027)	10.450*** (0.027)	10.467** (0.027)
Percent tenured	0.001*** (0.0001)	0.001*** (0.0001)	0.001*** (0.0001)	0.001*** (0.0001)	0.001*** (0.0008)
Percent female	–0.001*** (0.0002)	–0.001*** (0.0001)	–0.004*** (0.0003)	–0.005*** (0.0003)	–0.005*** (0.0003)
Enrollment (1000s)	0.006*** (0.0001)	0.006*** (0.001)	0.004*** (0.001)	0.005*** (0.001)	0.005*** (0.001)
Student/Faculty	0.001*** (0.0004)	0.001*** (0.0001)	0.002*** (0.0001)	0.002*** (0.0001)	0.002*** (0.0001)
Tuition (1000s)	–0.020*** (0.008)	–0.024*** (0.008)	0.041*** (0.006)	0.038*** (0.006)	0.038*** (0.006)
Net revenues	0.002*** (0.0001)	0.003*** (0.0002)	0.001*** (0.0001)***	0.002*** (0.0001)	0.002*** (0.0001)
Percent revenues	–0.002*** (0.0004)	–0.003*** (0.0004)	–0.006*** (0.0003)	–0.006*** (0.0003)	–0.006*** (0.002)
School type	–0.121*** (0.007)	–0.111*** (0.006)	–0.013 (0.010)	0.022** (0.011)	0.024** (0.016)
Union	0.086*** (0.014)	0.090*** (0.014)	0.074*** (0.021)	0.105*** (0.022)	0.099*** (0.022)
Ufem	–0.001 (0.0003)	–0.001 (0.001)	0.002*** (0.001)	0.001** (0.001)	0.001*** (0.001)
Utype	0.002 (0.011)	–0.005* (0.011)	–0.095*** (0.014)	–0.115*** (0.016)	–0.117*** (0.017)
Estimated Union Effect:					
2-Year school	0.058	0.032	0.059	0.030	0.022
4-Year school	0.056	0.050	0.134	0.135	0.129
Adjusted R^2	0.563	0.534	0.398	0.353	0.347
F-Statistic	207.87***	184.88***	160.81***	132.91***	193.56**

Data Source: National Center for Education Statistics. All regressions include regional controls. All but Columns 7 include year controls.
Significance levels: * = 0.10; ** = 0.05; *** = 0.001.

of types of professors and therefore a larger variance for average compensation. Because tests for heterogeneity indicate the problem,[10] weighted-least squares (WLS), using the number of faculty as the weight, was employed.

Table 3 (Columns 1 and 3, bottom) indicates that four-year schools encounter a larger union premium; however, during the 1989-96 period, the average union premium is more than double that experienced by faculty at two-year schools. Columns (2) and (4) indicate that controlling for heteroskedasticity reduces the union differential for two-year schools. During the 1978-85 period, OLS estimates indicate a 5.8 percent compensation premium for two-year schools and a 5.6 percent premium for four-year schools. These values fall to 3.2 percent and 3 percent, respectively, with the correction. The estimate remains almost unchanged for four-year schools. These results suggest that the heterogeneity of compensation is a much larger problem for two-year schools.

Also note that the 3 percent to 5 percent differential for the 1978-85 period is consistent with the values found earlier [3, 20], even though the earlier studies use cross-sectional data. The relatively high values for the second panel are similar to the 7 percent to 14 percent differential estimated by Monks [23] using 1990s data.

Although most of the other coefficients remain consistent throughout all models, one other factor worth mentioning is the effect of tuition on compensation.[11] The estimates suggest that during the 1978-85 period, a $1,000 increase in tuition led to a 2.4 percent decrease in average compensation. However, in the second panel, a tuition increase leads to about a 3.8 percent increase in faculty compensation. The tuition effects may be reflective of the revenue-generating ability of the university at a particular time.

A second issue deals with the assumption of union status as an endogenous variable. Beginning with the empirical work of Ashenfelter and Johnson [29], many studies have estimated the union-nonunion wage or salary differential under the assumption that compensation and unionization are simultaneously determined. If this assumption is true, then the OLS and WLS estimation of the compensation model leads to biased and inconsistent parameter estimates.

The methodological approach to account for the simultaneity of union status and compensation was developed by Lee [22] and followed by many other researchers. Kesselring suggested that the endogeneity problem is more likely associated with institutional-level data because the union choice is made by the faculty group for the entire institution [3].

The system of equations proposed by Lee [22] as discussed in Maddala [30] is adapted as follows. Assume the following system of equations:

[10] The Park and Glejser tests were conducted. Following other research, the error variance is assumed proportional to the number of faculty.

[11] The coefficient on School Type also changes between the two panels; however, as mentioned earlier, the second period has a substantial percentage of two-year schools, and it may be that the change is due to the increase in two-year schools from the first to the second period.

$$\ln W_{uit} = \alpha_u + X_{uit}\alpha_{2u} + \epsilon_{uit} \tag{1}$$

$$\ln W_{nit} = \alpha_n + X_{nit}\alpha_{2n} + \epsilon_{nit} \tag{2}$$

$$U_{it}^* = \beta_1 + \beta_2(\ln W_{uit} - \ln W_{nit}) + Z_{it}\beta_3 - \epsilon_{it} \tag{3}$$

where the $\epsilon_j - N(0,\sigma_j)$.

Equation (1) is the compensation equation for the unionized schools, Equation (2) is the same for nonunionized schools, and Equation (3) determines the probability of whether or not a school is unionized at time t. The αs are the parameters of interest in the compensation equation, and X is a matrix of factors affecting compensation; β and Z represent the parameters to be estimated in the union probability equation and factors that affect the probability of unionization, respectively.

Under this scenario, W_{uit} is observed only if $U_{it}^* > 0$; otherwise, W_{nit} is observed. OLS of (1) and (2) gives inconsistent estimates because:

$$E(\epsilon_u|U^* > 0) \neq 0 \text{ and } E(\epsilon_n|U^* \leq 0) \neq 0$$

Lee proposed a two-stage least squares method to provide consistent parameter estimates when accounting for union endogeneity. This procedure requires that Equations (1) and (2) be substituted into Equation (3) and a probit regression is performed on the reduced form model of (3). Then, conditional on union status, the wage equations are:

$$\ln W^{uit} = \alpha_u + X_{uit}\alpha_{2u} - \sigma_{\epsilon u}\,[\phi(\Psi)/\Phi(\varphi)] + \eta_{uit} \tag{4}$$

$$\ln W^{nit} = \alpha_n + X_{nit}\alpha_{2n} - \sigma_{\epsilon n}\,[\phi(\Psi)/(1 - \Phi(\varphi))] + \eta_{nit} \tag{5}$$

where $\phi(\Psi)$ is the density function and $\Phi(\varphi)$ is the distribution function of a standard normal variable. The expression in the brackets is commonly called the inverse Mills ratio. φ is the estimated reduced form of Equation (3).

The probit regression includes the exogenous variables listed earlier and two additional variables that are assumed to affect the outcome of U*. RTW is equal to 1 if the school is located in a right-to-work state, and 0 otherwise. It is included to proxy for the general attitude toward unionism within a region. The percentage difference between the school's average compensation and that of the average compensation paid by all institutions within the state is included because faculty may be motivated to unionize if they perceive their compensation package to be lacking when compared to others in their locality.

Results from the probit estimation are in Appendix 1, but they are not discussed in detail here because the focus is on the salary equations. Note that the overall predictive ability for the probit was more than 80 percent for both time periods. Table 4 presents the union and nonunion wage equations without accounting for endogeneity (Columns 1, 3, 5, and 7) and with the inverse Mills ratio included

(Columns 2, 4, 6, and 8). All regressions account for heterogeneity with WLS. The separate estimation of the union and nonunion compensation equations indicate that this model form is appropriate in comparing union and nonunion compensation. Although many of the parameter estimates are similar, several key coefficients are different. After controlling for endogeneity, the coefficient on tuition increases, and the effect is larger for unionized schools. At first I considered whether unions enable faculty to persuade institutions to "share the pie" as tuition increases, and the two-step procedure better accounts for the relationship between unions and tuition. However, the marginal effects related to the revenue variables do not indicate that revenue sharing in general occurs. And, Appendix 1 indicates that unionization is inversely associated with tuition and net revenues. Perhaps unions view high tuition values as a sign to negotiate for better salaries. This may be true because the effect is larger in the 1978-85 period when faculty unions were relatively new. By the 1989-96 period, many of the unions had been in place for many years and negotiation tactics may have changed substantially.

The coefficient related to school type also indicates that the two-step procedure yields estimates consistent with factual evidence during the 1990s. The 2SLS (two-stage least squares) procedure indicates a 6.1 percent lower average faculty compensation for two- versus four-year schools in the union sector; the differential is 16.2 percent lower in the nonunion sector. The 2000-2001 salary report of the American Association of University Professors (AAUP) [27] indicates that one main factor increasing the faculty compensation distribution during the 1990s was the growing pay gap between faculty at public research institutions versus other types of institutions. It is not surprising that the coefficient associated with school type is larger for the nonunion sector, which includes the top echelon of public universities and many lower-tiered schools. Further, unions tend to reduce salary inequality within unionized private sectors [31], so it is not unreasonable to find a substantially smaller average compensation difference between two- and four-year schools in the union sector.

The most interesting outcome of this set of regressions is the estimated union-nonunion wage differential. The differential is found by estimating the average lnW for the union and nonunion groups and calculating $\exp(\ln W_u - \ln W_n)$ [3]. Using this method, the differential is approximated at about 5.3 percent for the 1978-85 period and 13.5 percent for the 1989-96 panel. This result is greater than those found for the earlier period [3, 4, 20] and within the range of estimates presented by Monks [23] for the second period. The fact that the 1978-85 estimate is outside of the 0 percent to 4 percent range from most of the earlier research is not troublesome. These earlier studies include private institutions in their analysis. While the inclusion of private schools may be tenable on data from earlier than 1980, it may be that including private schools after 1980 does not paint an accurate picture of the union influence on faculty compensation. As mentioned earlier, the organization of faculty in private institutions was substantially reduced by the

Table 4. Regression—By Union Status

	1978–1985				1989–1996			
	Union		Nonunion		Union		Nonunion	
Variable	WLS (1)	2SLS (2) (weighted)	WLS (3)	2SLS (4) (weighted)	WLS (5)	2SLS (6) (weighted)	WLS (7)	2SLS (8) (weighted)
Constant	10.595***	11.8172***	10.282***	10.278***	10.403***	10.553***	10.555***	10.648***
	(0.060)	(0.069)	(0.042)	(0.038)	(0.041)	(0.077)	(0.034)	(0.360)
Percent tenured	–0.006***	-0.006***	–0.0014***	–0.002***	0.015***	0.001**	0.0005	0.0004
	(0.0001)	(0.0004)	(0.0003)	(0.0002)	(0.001)	(0.0002)	(0.001)	(0.002)
Percent female	–0.002***	–0.003***	0.0007***	–0.001***	0.001***	–0.003***	0.002***	–0.005***
	(0.0004)	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.001)	(0.0002)	(0.0003)
Enrollment (1000s)	0.0054***	–0.0104***	0.007***	–0.0007	0.001***	0.011***	0.001***	0.002
	(0.001)	(0.001)	(0.001)	(0.001)	(0.0001)	(0.002)	(0.0001)	(0.002)
Student/Faculty ratio	0.002***	0.005***	0.001***	0.002***	–0.004***	0.001	–0.004***	0.002***
	(0.0003)	(0.0003)	(0.0002)	(0.0002)	(0.0004)	(0.001)	(0.0003)	(0.0002)
Tuition	0.0003	0.224***	–0.017*	0.113***	0.045***	0.049***	0.037***	0.026
	(0.020)	(0.018)	(0.009)	(0.010)	(0.008)	(0.008)	(0.009)	(0.058)

Net revenues	0.006***	0.009***	0.0017***	0.0034***	–0.00004	0.0001	0.002***	0.002***
	(0.0003)	(0.0003)	(0.0002)	(0.0002)	(0.0003)	(0.0001)	(0.0001)	(0.003)
Percent revenues	–0.006***	–0.013***	–0.0012**	–0.004***	–0.00373***	–0.006	–0.007***	–0.007***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.0005)	(0.006)	(0.0004)	(0.001)
School type	–0.102***	–0.027**	–0.108***	–0.067***	–0.034*	–0.061	–0.011	–0.162***
	(0.013)	(0.010)	(0.007)	(0.0063)	(0.015)	(0.074)	(0.01)	(0.051)
Adjusted R^2	0.541		0.529		0.355		0.331	
F-Statistic	75.48***		142.000***		56.96***		90.08***	
λ		–0.579***		–0.400***		–0.001		–0.001
		(0.024)		(0.020)		(0.024)		(0.02)
Log-likelihood		900.091***		1674.02***		726.549***		1023.04***
N	1083	1083	2131	2131	1828	1828	3244	3244

Data Source: National Center for Education Statistics. All regressions include year and regional controls. Significance levels: * = 0.10; ** = 0.05; *** = 0.001.

1980 *Yeshiva* decision. Therefore, any analysis of data after that time which includes schools that lack the legal opportunity to unionize will lead to an underestimation of the union coefficient.

A final test of model specification involves the unobserved error that may be associated within schools due to the panel nature of the data. In this case, the model becomes:

$$\ln W_{it} = \alpha_u + X_{it}\alpha_{2u} + c_i + \epsilon_{it} \tag{6}$$

where the variables are defined as before and c_i is the unobserved school effect not captured by the explanatory variables.

This study assumes that c_i is a random effect so that the several binary explanatory variables of interest remain in the model.[12] Column (5) of Table 3 presents the results of estimation of Equation (6), controlling for heteroskedasticity. The union premiums under this method are 2.2 percent for two-year schools and 12.9 percent for four-year schools. These estimated premiums are smaller than the differentials estimated with WLS, but the reduction is not as substantial as in Rees [4].

CONCLUSION

This article uses data that combines information on individual school compensation, faculty, financial status, and union status and examines the faculty compensation differential by union status over two time periods. The results indicate that time is a factor in the size of the differential. The union premium was between 4 percent and 5 percent during the 1978-85 period, even when controlling for heterogeneity and endogeneity. However, while two-year schools experienced a similar effect in the 1989-96 period, faculty at four-year institutions experienced a much larger union effect that was estimated between 12 percent and 15 percent. In reviewing this outcome in light of Figure 1, it may be that faculty unions helped average salaries rebound slightly during the economic boom of the 1990s. However, it appears that the main effect was that faculty unions kept average salaries inflation-proof, but the buying power of average salaries for nonunion schools fell. Future research should consider the time period of the union data in order to make generalizations about the union effect on faculty pay.

The analysis also suggests that the differential is somewhat sensitive to the econometric specification of the error term in faculty compensation models, but that the different model specifications made only a small impact on the overall

[12] With a fixed-effects assumption, the time invariant variables will be excluded from the estimation. That would eliminate the union-related and school type variables from the analysis. Underlying the random effects model is the assumption that Cov(X,c) = 0. While this may be too restrictive for the data at hand, any bias from possible correlation between the explanatory variables and the error term is likely to be small.

estimates. Employing 2SLS yielded almost identical estimates from the WLS model. The 2SLS estimate for union/nonunion differential is 5.3 percent for the 1978-85 panel and 13.5 percent for 1989-96 panel, compared to the WLS estimates of 5 percent and 13.5 percent, respectively. However, controlling for endogeneity of the union variable appears especially suitable for the 1978-85 panel, where the coefficient on the Mills ratio is statistically significant.[13]

The fact that the union premium has grown may also have important policy implications for the future. Recent court and NLRB decisions have diluted the 1980 *Yeshiva* decision. In 2001, the Supreme Court ruled that the burden of proof regarding the managerial status of employees falls on the employer [32]. Since the 1990s, NLRB decisions, regardless of whether the board finds faculty to be "managerial" employees or not, require that institutions clearly define the managerial duties of faculty who desire a collective bargaining unit [33, 34]. Additionally, a 2003 NLRB decision at the regional level supports the unionization of part-time faculty if they have worked for the institution for several years [35]. If these rulings are an indication of the future direction of private university collective bargaining, then the rather substantial 13 percent premium may draw many new institutions to unionization, particularly in small colleges where faculty pay is low.

[13] There seems to be some controversy on this topic. Lewis [16] and Barbezet [20] suggested that the 2SLS estimation, while innovative, is proving to be useless at the least, and perhaps more worrisome, leading to spurious conclusions about the impact of unions on compensation.

APPENDIX 1. Probit (Weighted)
1 = Unionized School[a]

	1978–1985		1989–1996	
	Coefficient	Marginal effect	Coefficient	Marginal effect
Constant	–3.173*** (0.471)	–0.934	0.474** (0.240)	0.114
Percent tenure	0.010*** (0.002)	0.003	0.011*** (0.001)	0.003
Percent female	0.135*** (0.003)	0.004	–0.008*** (0.002)	–0.001
Enrollment	0.049*** (0.007)	0.015	0.025*** (0.007)	0.007
Student faculty ratio	–0.012*** (0.002)	–0.004	0.004*** (0.001)	0.001
Tuition	–0.741*** (0.098)	–0.218	–0.349*** (0.051)	–0.104
Net revenues	–0.016*** (0.002)	–0.005	–0.005*** (0.001)	–0.001
Percent revenues	0.035*** (0.006)	0.010	–0.020*** (0.003)	–0.006
School type	–0.086 (0.079)	–0.026	0.247*** (0.091)	0.074
RTW state	–0.385*** (0.083)	–0.113	–0.775*** (0.058)	–0.231
Difference from state average	0.012*** (0.022)	0.004	0.496*** (0.127)	0.148
Log-likelihood function	–1317.875		–1850.839	
Chi-square statistic	1471.746***		2928.986***	
Percent correctly predicted	80.09%		82.7%	

[a]All regressions include controls for region and year of observation.
* = α at .10. ** = α at .05; ***α = .01.

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