COLLECTIVE BARGAINING AND STAFF SALARIES IN AMERICAN COLLEGES AND UNIVERSITIES

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Previous studies of union wage effects in higher education have examined faculty salaries, but not staff salaries. This study, using data from a 1997-98 survey conducted by the Association of Higher Education Facilities Officers and other sources, investigates how union coverage affected staff salaries at 163 U.S. colleges and universities. The authors estimate a union salary premium of 9-11%, with variation from near zero for some of the 47 occupations in their sample to 13-16% for others, such as the skilled building trades. The union/nonunion differential appears to be larger in 2-year than in 4-year institutions, but does not vary between the public and private sectors. Where faculty members are covered by a collective bargaining agreement, unionized staff members appear to enjoy an additional salary gain of 2-3%.

In 2001, a twenty-day sit-in at Harvard University brought the living-wage debate to the forefront of American consciousness. After a six-month study, the Harvard Committee on Employment and Contracting Policies, a 19-member committee of faculty, staff, administrators, and students that had been appointed by Harvard's president as a result of the discussions to end the sit-in, recommended giving raises to the university's lowest-paid employees and relying more on collective bargaining in the future to assure that the wages paid by subcontractors did not undercut local union wage scales (*Chronicle of Higher Education,* January 11, 2002). A three-day sit-in at the University of Connecticut that related to the living wage issue also yielded a substantive victory for campus workers. The protesters there generated an almost two-dollar increase in wages, as well as substantial improvement in benefits for many of the university's workers (*Chronicle of Higher Education,* May 25, 2001). Collectively, such

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The information on staff salaries and collective bargaining coverage used in this paper was provided conditional on the authors maintaining the confidentiality both of the data and of the identities of institutions included in the sample. However, if other researchers are granted access to these data by the Association of Higher Education Facilities Officers, the authors will be happy to provide them with the other variables used in this analysis so that they can replicate the study.

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Table 1. Collective Bargaining Coverage of College and University Staff in 1994.

Occupation Category	*Total Employees*	*Estimated Employees in Bargaining Units*	*Percent Represented*
White-Collar	1,070,142	250,573	23.4
Blue-Collar	306,335	131,232	42.8
Total	1,376,477	381,805	27.7

Sources: Digest of Education Statistics 1994 (Washington, D.C.: National Center for Education Statistics, 1994), pp. 228–29 (Total Employees); *Directory of Staff Bargaining Agents in Institutions of Higher Education* (New York: National Center for the Study of Collective Bargaining in Higher Education and the Professions, 1995) (Employees in Bargaining Units).

struggles represent a new battleground in American higher education.

The growth of living wage movements on almost one hundred campuses stems, in part, from increasing awareness among staff at those institutions that their pay compares poorly with that of similar staff at other institutions (Van Der Werf 2001). There are many potential explanations for these salary differences, including differences in local cost of living and in the resources that the academic institutions have available to pay faculty and staff salaries. One other possible explanation is the influence of staff unions. Previous studies of the impact of unions on salaries in academia have focused on faculty unions and have concluded that these unions have achieved at best a small percentage gain in salary for their members relative to faculty salaries at academic institutions in which faculty are without union coverage.[1] There have been no studies, however, of the impact of collective bargaining on staff salaries in higher education.

Our paper addresses this issue. We first provide some background data on the number of blue-collar and white-collar employees covered by collective bargaining agreements at American higher education institutions. We then use data from a 1997–98 study on the costs of staffing in higher education conducted by the Association of Higher Education Facilities Officers (APPA) and other sources to estimate models that explain the variation in academic institutions' salaries for a number of narrowly defined blue-collar and white-collar occupational groups that are employed by the academic institutions' facilities divisions.[2] Of primary interest to us is the extent to which the salaries of academic staff covered by collective bargaining agreements exceed the salaries of otherwise comparable academic staff who are not covered by such agreements.

Background Data

Table 1 presents data on the employment levels of blue-collar and white-collar staff members employed in American higher education in the mid-1990s, as well as the percentage of each group that was covered by a collective bargaining agreement. A much higher percentage of blue-collar employees (42.8%) than of white-collar employees (23.4%) were represented by staff unions. Because there are many more white-collar employees, in the aggregate about 27.7% of staff at American colleges and universities was covered by union contracts in the mid-1990s.

The salary and collective bargaining coverage data used in our study come from the APPA's *1997–1998 Comparative Costs and*

[1]See, for example, Ashraf (2000), Barbezat (1989), Kessering (1991), and Rees (1993). James Monks (2000) estimated union effects of 7–14%, which exceed the estimates found in other studies, but his sample consisted largely of two-year institutions.

[2]The acronym APPA is derived from the earlier name of the organization, the Association of Physical Plant Administrators of Universities and Colleges.

Table 2. Distribution of Academic Institutions by Carnegie Category and Control in the APPA Sample.

	Funding		
Carnegie Category	*Private*	*Public*	*Total*
Associate	1 (.01/.18)	13 (.08/.34)	14
Baccalaureate	23 (.14/.19)	3 (.02/.03)	26
Doctoral	4 (.02/.02)	16 (.10/.02)	20
Masters	12 (.07/.09)	42 (.26/.10)	54
Research	7 (.04/.01)	42 (.26/.03)	49
Total	47	116	163

Note: Numbers in parentheses are (share of institutions in the sample in the category) / (share of the nation's 2,873 higher education institutions in these categories that fall in the category).

Staffing Report for College and University Faculties.[3] This data set provided information on salary levels and collective bargaining coverage for 47 narrowly defined occupations at 193 U.S. and Canadian colleges, universities, and elementary and secondary schools. We restrict our attention to U.S. higher education institutions that were classified as Research, Doctoral, Masters, Baccalaureate, or Associate (two-year) institutions by the Carnegie Foundation for the Advancement of Teaching (1994).[4] The sample that we used ultimately consisted of 163 institutions.

Table 2 presents a breakdown of the institutions in our sample by Carnegie classification and by form of control. Public institutions constitute the majority of the institutions in each Carnegie category in our sample, except for the Baccalaureate category. As the table indicates, our sample is not representative of American higher education as a whole. In particular, two-year institutions and private baccalaureate institutions are under-represented in our sample, while public masters and doctoral and all research universities are over-represented.

We restrict our attention to the 22 occupations in the survey that are not managerial or executive, and we include only those sample observations for which both an occupational salary level and whether the employees in the occupation were covered by a collective bargaining agreement are reported. Table 3 shows the difference in mean annual salaries between unionized and non-unionized employees for each occupation, the ratio of the mean salary in an occupation for employees who were covered by union contracts to the mean salary in an occupation for employees who were not covered by a union contract, and the difference in mean salaries between covered and noncovered institutions, as well as the standard deviation of the difference in the means.

In each occupation, employees covered by a union contract earned considerably more than employees not covered by a contract, with the raw differentials in the mean salaries varying across occupations from 17% to 42%. In each occupation the difference in mean salaries between covered and noncovered employees is more than twice the standard error of the difference in the means, allowing us to reject the

[3]We are grateful to Joseph Lally, Director of Business Operations for Cornell's Facilities Services Division, for granting us access to these data, under the condition that we keep the data confidential and not identify the specific institutions that participated in the survey.

[4]In addition to excluding Canadian and elementary and secondary institutions, we also excluded specialized U.S. institutions such as seminaries and conservatories. Institutions in the Research and Doctoral categories have annual research volumes and annual numbers of doctoral degrees granted that exceed specified minimum levels (the levels are higher for the Research category).

Table 3. Mean Occupational Salaries in 1997–98 for Employees Covered by Collective Bargaining Agreements and Not Covered by Collective Bargaining Agreements in the APPA Sample.

Occupation (n)	*(1) Mean Salary Noncovered*	*(2) Mean Salary Covered*	*(3) Difference in Means (Se)*[a]	*(4) Ratio of the Means (2)/(1)*
Secretary (143)	$21,953	$26,987	$5,025 (225)	1.23
Custodian (142)	16,993	22,850	5,857 (105)	1.34
Groundskeeper (150)	18,838	26,138	7,300 (108)	1.39
Carpenter (143)	26,206	35,962	9,756 (144)	1.37
Electrician (145)	27,701	38,629	10,928 (162)	1.39
Locksmith (119)	27,243	33,463	6,220 (765)	1.23
AC/Refrig (120)	26,576	37,600	11,024 (177)	1.41
Painter (131)	24,468	34,645	10,177 (149)	1.42
Plumber (139)	26,852	37,575	10,723 (173)	1.40
Programmer (82)	37,311	43,509	6,198 (683)	1.17
HVACTech (104)	30,866	37,357	6,491 (245)	1.21
UtilitiesOp (105)	24,758	36,307	11,549 (209)	1.47
GeneralMaint (110)	24,121	31,746	7,625 (200)	1.32
ElevMechanic (30)	31,633	44,053	12,420 (690)	1.39
VehMechanic (108)	25,914	32,424	6,581 (157)	1.25
Storekeeper (102)	23,750	28,689	4,939 (164)	1.21
GeneralLabor (84)	19,097	27,071	7,975 (189)	1.42
Security (57)	26,849	35,665	8,816 (770)	1.33
Machinist (71)	29,065	36,249	7,184 (256)	1.25
Mason (63)	27,392	35,717	8,325 (284)	1.30
Roofer (57)	26,354	35,623	9,270 (325)	1.35
SheetMtlWrkr (51)	28,286	36,530	8,244 (391)	1.29

[a]Standard error of the absolute difference in the means.

Source: Authors' computations from the APPA data. Only institutions that reported both a salary figure for an occupation and whether it was covered by a union contract are included. All of the differences of the means are statistically significantly different from zero at the .05 level of significance.

hypothesis that covered and noncovered workers' salaries are equal. The differences are largest in the skilled trades. Salaries for custodial workers, the group of employees that has been the focus of the living wage debate on many campuses, were the lowest in the group, and the unionized custodial workers in the sample earned about 34% more on average than custodial workers at academic institutions that were not covered by a collective bargaining agreement.

Estimating the Union/ Nonunion Salary Advantage of Unionized Academic Staff

The estimated differences in the salaries of academic staff covered by and not covered by union contracts reported in Table 3 are raw differences that do not control for characteristics of the institutions, or of the areas in which the institutions are located, that might be expected to influence staff salaries independent of unionization. For example, if academic institutions whose employees were organized also had greater financial resources, or were located in higher cost-of-living areas, than institutions whose employees were not organized, one would expect to observe the former paying higher salaries than the latter even if unionization per se had no effect on the salaries of staff at academic institutions. To estimate whether staff unions influence salaries, it is necessary to control for the other characteristics of the institutions that might be expected to influence salaries.

To accomplish this, we initially pool the data across occupations and institutions and estimate staff salary equations of the form

$$(1)\qquad \text{Log } (W_{ij}) = a_0 + a_1U_{ij} + a_2Y_j + a_3Z_j + a_4F_j + a_5d_i + e_{ij}.$$

In equation (1), W_{ij} is the annual salary paid to a staff member in occupation i at academic institution j, U_{ij} is a dichotomous variable indicating whether the employees in occupation i are unionized in institution j, Y_j is a vector of dichotomous variables indicating the Carnegie classification of institution j (two-year colleges are the omitted category), Z_j is a vector of other variables that vary across institutions and are expected to influence staff salaries, F_j is a dichotomous variable indicating whether the faculty at academic institution j are unionized, d_i is a vector of occupational dichotomous variables, and the e_i are random error terms. Because the dependent variable is the logarithm of salaries, a_1 can be interpreted as the estimated percentage by which the salaries of staff in institutions with collective bargaining for the occupation exceed the salaries of staff at institutions without collective bargaining for the occupation, after controlling for the other factors expected to influence salaries.[5]

We include in the Z_j a set of variables that influence the resources academic institutions have at their command out of which to pay the salaries of staff. These include the logarithm of the institution's endowment per student (lEndowm/Stdnt), the logarithm of its average undergraduate tuition (lTuition), and, for public institutions, the logarithm of its state and local government appropriations per student (lGovApprops/Stdnt).[6] Also included in this vector, to control for differences in cost of living or wage levels across areas, is the logarithm of the mean salary of custodians in the city in which the academic institution is located (lMeanCustodSal). When an institution was not located in a city for which we had mean custodian salary data, the mean custodian wage in the state was substituted. The occupational dichotomous (d_i) variables are included to control for differences in salaries across occupations, and the faculty union variable (F_j) is included to see if salary gains won by faculty unions spill over to staff unions' compensation. Finally, included in this vector is the logarithm of the average math and verbal SAT 75[th] percentile score for entering freshmen at the institution (lSAT). This variable, as well as the Carnegie category variables, were included to see if the "selectivity" of an academic institution, or its institutional type, influences the salary of its staff, once we have controlled for its financial resources.

The first column of Table 4 presents the estimates of equation (1). Staff members covered by a collective bargaining contract in our sample are paid about 9% more than staff members who are not covered by a contract, other factors held constant, and staff members appear to earn close to another 3% more if the faculty at the institution are also covered by a union contract. Staff members' salaries are clearly strongly related to the proxy for the cost of living or alternative wages in the area (lMean CustodSal), and we cannot reject the hypothesis that a 1% increase in the average wage of custodians in the area is associated with a 1% increase in academic institutions' staff members' salaries.[7]

We find little evidence that institutions that are better off financially, as measured by endowment per student, average tuition level, or, for public institutions, per student state and local government appro-

[5]More precisely, the estimated union/nonunion salary differential is given by $(e^{a_1} - 1)(100)$.

[6]For public institutions, this is a weighted average of in-state and out-of-state tuitions, with the weights depending on the fraction of the institution's students who come from each category.

[7]Substitution of a census housing price index in an area for the average custodian wage measure led to poorer fits and less statistically significant coefficients. It thus appears that the average custodian wage reflects area wage level differences more than it reflects cost of living differences.

Table 4. Logarithm of 1997–98 Salary Equations: Pooled Regressions.[a]
(Absolute Value of t Statistics in Parentheses)

	(1)	*(2)*	*(3)*	*(4)*
Bachelors	–.147 (5.5)	–.146 (3.5)	–.136 (7.0)	–.133 (4.2)
Doctoral	–.118 (4.5)	–.120 (4.6)	–.146 (4.5)	–.146 (4.6)
Masters	–.098 (4.5)	–.100 (4.6)	–.117 (4.4)	–.117 (4.4)
Research	–.077 (3.2)	–.078 (3.2)	–.082 (2.8)	–.083 (2.7)
lGovApprops/Stdnt	–.025 (6.2)	–.024 (6.0)	–.021 (4.2)	–.021 (4.2)
lEndowm/Stdnt	.000 (0.0)	.000 (0.0)	–.003 (0.9)	–.003 (0.9)
lTuition	–.024 (3.3)	–.023 (3.3)	–.018 (2.0)	–.018 (2.0)
lSAT	.359 (5.0)	.349 (4.8)	.320 (3.5)	.303 (3.3)
lMeanCustodSal	.997 (22.7)	.991 (22.6)	1.000 (17.7)	.994 (17.6)
UnionContract	.093 (7.9)		.108 (7.0)	
FacUnionized	.028 (2.2)	.028 (2.3)	.022 (1.4)	.024 (1.5)
uCustodian		.116 (3.5)		.120 (3.7)
uSecretary		.024 (0.7)		.023 (0.6)
uGroundskeeper		.044 (1.2)		.048 (1.4)
uCarpenter		.124 (3.7)		.127 (3.8)
uElectrician		.131 (4.0)		.135 (4.1)
uLocksmith		.046 (1.2)		.047 (1.3)
uAC/Refrig		.150 (4.2)		.154 (4.4)
uPainter		.161 (4.6)		.163 (4.8)
uPlumber		.136 (3.9)		.139 (4.1)
uProgrammer		.011 (0.2)		
uHVACTech		.034 (0.8)		
uUtilitiesOp		.156 (3.8)		
uGeneralMaint		.118 (2.9)		
uElevMechanic		.129 (1.7)		
uVehMechanic		.042 (1.1)		
uStorekeeper		.026 (0.6)		
uGeneralLabor		.152 (3.4)		
uSecurity		.068 (1.0)		
uMachinist		.044 (0.9)		
uMason		.086 (0.7)		
uRoofer		.088 (1.6)		
uSheetMtlWrkr		.063 (1.0)		
R^2/n	.602/2115	.604/2115	.638/1207	.641/1207

[a]Also included in the model were intercept terms for each occupation (one was excluded to avoid perfect collinearity) and dichotomous variables for non-reporting of endowment per student, government appropriations per student, tuition, and average SAT scores.

Variable explanations. Bachelors = 1 if the Carnegie Category of the institution was Baccalaureate, 0 otherwise; Doctoral = 1 if the Carnegie Category of the institution was Doctoral, 0 otherwise; Masters = 1 if the Carnegie Category of the institution was Masters, 0 otherwise; Research = 1 if the Carnegie Category of the institution was Research, 0 otherwise (the omitted category is 2-year college institution). lApprops/Stdnt = interaction between a (0, 1) dichotomous variable for being a public institution and the logarithm of state and local government appropriations per student; lEndowment/Stdnt = logarithm of endowment per student at the institution; lTuition = logarithm of the average undergraduate tuition at the institution; lSAT = logarithm of the average 75[th] percentile scores on entering students' verbal and mathematics SAT scores; lMeanCustodSal = logarithm of the mean salary for custodians in the city or state of the institution; UnionContract = 1 if the occupation was covered by a union contract at the institution, 0 otherwise; FacUnionized = 1 if the institution's faculty members were covered by a union contract, 0 otherwise. uCat: subsequent variables with the prefix u are equal to 1 if the indicated occupation was covered by a union contract at the institution, 0 otherwise (models in which union impact varies by occupation). See the Appendix for the list of occupation categories.

Data Sources. Bachelors, Doctoral, Masters, Research, Public/Private, UnionContract, and uCat: APPA Survey. lApprops/Stdnt, lEndowment/Stdnt, lTuition: Webcaspar. lSAT: *America's Best Colleges—1998* (Washington, D.C.: *U.S. News & World Report,* 1997). lMeanCustodSal: *Metropolitan Area Occupational Employment and Wage Estimates* (U.S Bureau of Labor Statistics, 2000; http://www.bls.gov/bls/blswage.htm). FacUnionized: Directory of Faculty Contracts and Bargaining Agents in Institutions of Higher Education, National Center for the Study of Collective Bargaining in Higher Education and the Professions (Baruch College, New York, 1997).

priation level, pay staff members higher salaries. Indeed, institutions with higher tuition levels and public institutions with higher state appropriations per student actually appear to pay their staff lower salaries, other factors held constant. That institutional financial variables do not influence staff salary levels is not completely surprising. Academic institutions draw employees in the occupations included in our sample primarily from local labor markets, and while institutions may talk about wanting to attract employees of the highest possible quality, they are less concerned about doing so for support staff than they are for faculty, upon whom presumably an institution's reputation is strongly based.[8]

We do find, however, that the selectivity of an institution's undergraduate students, as measured by their SAT scores, is positively related to the salaries of staff in these occupations. An explanation for this finding is that to attract high-quality students, institutions find they need to provide both academic *and* nonacademic services of high quality, and thus they offer higher staff salaries to attract, motivate, and retain higher-quality staff.[9]

The estimates above assume that the impact of staff unions on staff salaries is the same across all occupations. In column (2) we report the results of estimating a model in which the impact of staff unions is allowed to vary across occupations. The estimated collective bargaining coverage differentials vary from about 1% to 16%, but the differential is statistically significantly different from zero in only 10 of the 22 occupations. The differentials appear to be largest in many of the skilled trades, where unions historically have achieved substantial wage gains for their members. Given the relatively small number of observations for a number of the occupations, it is difficult to distinguish between the hypothesis that collective bargaining coverage has no effect on university staff members' salaries in these occupations and the hypothesis that larger sample sizes might yield a statistically significant effect.[10]

To get around the small sample size problem and to lay the foundation for the sensitivity analyses that we conduct in the next section, we restrict our attention to the nine occupations for which we have the greatest numbers of observations; in column (3) we provide the estimated coefficients when equation (1) is estimated using this restricted sample of data. The estimated collective bargaining coverage differential is around 11%, which is very close

[8]Indeed, when we estimated a similar equation to explain the logarithm of average full professors' salaries, we found that the logarithm of endowment per student was positively related to this salary measure, indicating that richer universities do attempt to attract higher-quality faculty. Neither the presence of a faculty union nor the share of staff occupations that were covered by collective bargaining agreement was significantly associated with the average faculty salaries. Finally, in contrast to the results for the staff salary equations, use of the housing price index rather than the custodian wage marginally improved the performance of the faculty salary equation. This lends credence to the assertion that universities hire faculty in a national market and thus variables that reflect area cost of living will influence their salaries.

[9]At first glance, one might be tempted to conclude also that, other factors held constant, staff members at 2-year institutions (the omitted Carnegie category) earn between 8% and 15% more than their counterparts employed at baccalaureate, masters, doctoral, and research institutions. However, as a referee has pointed out to us, this comparison would not be very useful, because other factors are not constant across Carnegie categories. For example, SAT scores are not reported for 2-year institutions, and hence our SAT variable is coded zero and the dichotomous variable that is included in the model to control for nonreporting of SAT scores is coded one for each 2-year college. When the mean values, by Carnegie category, are substituted for each of the variables in the model, including those not reported in the table (except for staff unionization, which we set equal to zero), the model predicts that in the absence of staff unions, average staff salaries are actually lowest at the 2-year institutions.

[10]For example, only 30 institutions provide data for elevator maintenance workers (Table 3), and larger sample sizes might indicate that the estimated union differential for this occupation of about 13 percentage points was statistically significant. A formal F test allows us, however, to reject at the .05 level of significance the hypothesis that the impact of staff unions on staff salaries is the same across all occupations.

to that found in the unrestricted sample. In column (4) we present estimates for the restricted data sample that allow the effects of union coverage to vary across occupations. The estimated differentials are virtually identical to those found in column (2). We find statistically significant union coverage differentials in the range of 12–16% for the skilled building trades crafts—carpenters, electricians, painters, and plumbers—and statistically insignificant differentials that are close to zero for administrative secretaries, groundskeepers, and locksmiths. In addition, custodians covered by union contracts earn about 12% more than custodians not covered by contracts, all other variables held constant.

Several extensions warrant brief mention here. We also estimated, for both of the samples used in Table 4, models that allowed the estimated collective bargaining coverage differentials to vary with the Carnegie category of an academic institution and whether the institution was public or private. The estimated differentials were virtually identical for public and private academic institutions. However, they did vary across Carnegie category. In particular, by far the largest estimated differential, varying from about 20% to 30% depending on the sample, was observed for the two-year colleges. This finding is consistent with James Monks's (2000) finding, cited earlier, that faculty unions' "impacts" on salaries, although smaller in magnitude than staff unions', are largest at the two-year institutions.

Testing for the Sensitivity of Our Findings to Alternative Specifications

Our primary concern is the effect of unionization on staff employees' salaries. Table 5 summarizes the results of additional econometric modeling we conducted to investigate the sensitivity of the estimated union coefficient to the variables included in the analyses and to the econometric methods used. To provide a baseline, column (1) reports again the estimated union coefficients that are reported in column (4) of Table 4.

The estimates in Table 4 come from a model in which all of the coefficients of the explanatory variables, save for the intercept term and collective bargaining coverage in the occupation, are assumed to be constant across occupations. A more general specification would allow the coefficients of all variables to vary across occupations by estimating separate equations for each occupation. The collective bargaining coverage coefficients we obtained when we did this are shown in column (2) of Table 5. In the main, these are very similar to the corresponding coefficients found in column (1), never varying by more than .02.

The estimates presented in columns (1) and (2) treat each occupational equation as independent. They ignore the fact that there may be some omitted institution-level variables that influence the salaries of staff in all occupations. For example, the union/nonunion wage advantage for an occupation at an institution may depend on the fraction of the other staff occupations at an institution that are covered by collective bargaining agreements. Hence the wages of any given staff occupation at an academic institution may depend on the unionization of all staff occupations at the institution.

We attempted to re-estimate the models underlying column (2), including as an additional explanatory variable the fraction of all nine occupations that were covered by collective bargaining agreements.[11] Unfortunately, when one of the nine occupations was covered by a contract, the vast majority of the other occupations also were covered by a contract. Hence the coverage by union contract variable for an occupation was very highly correlated with the fraction of the nine occupations at the institution that were covered by union contracts. The high degree of collinearity prevented us from estimating such a model.

[11] Ehrenberg and Goldstein (1975) followed a similar procedure in their study of the impact of public sector unions on the wages of different occupational categories of public employees.

Table 5. Logarithm of 1997–98 Occupational Salary Equations: Coefficients of Union Variables—Sensitivity Analyses. (Absolute Value of t Statistics in Parentheses)

Occupation	*(1)*	*(2)*	*(3)*	*(4)*
Administrative Secretary	.023 (0.6)	.031 (0.6)	.010 (0.2)	.004 (0.1)
Custodian	.120 (3.7)	.130 (3.4)	.067 (2.0)	.092 (2.1)
Groundskeeper	.048 (1.4)	.014 (0.2)	.027 (0.4)	–.046 (0.9)
Carpenter	.127 (3.8)	.120 (2.6)	.107 (1.6)	.102 (2.0)
Electrician	.135 (4.1)	.113 (2.3)	.123 (1.8)	.110 (2.1)
Locksmith	.047 (1.3)	.058 (1.3)	.047 (0.9)	.033 (0.7)
Air Conditioning and Refrigeration	.154 (4.4)	.168 (3.2)	.115 (1.8)	.142 (2.5)
Painter	.163 (4.8)	.143 (3.0)	.122 (2.0)	.131 (2.5)
Plumber	.139 (4.1)	.133 (2.7)	.157 (2.4)	.115 (2.1)

Column explanations:
(1) Union coefficients from Table 4, column (4);
(2) Union coefficients from separate occupational salary equations;
(3) Union coefficients from separate occupational salary equations—seemingly unrelated regressions for the sample of institutions that reported data for all nine occupations;
(4) Selectivity bias corrected estimates corresponding to the estimates in column (2).

A second way to get at this issue is simply to treat the nine occupational salary equations as a single system and to allow the error terms to be correlated across equations. Estimating this system using the method of seemingly unrelated regressions will increase the efficiency of our estimates; however, as long as none of the other statistical assumptions is violated, the estimates reported in column (2) will remain unbiased.[12]

The method of seemingly unrelated regressions will increase the efficiency of the estimated coefficients only if the identical explanatory variables do not appear in each equation. In our system, the only explanatory variable that varies across occupations is whether employees in an occupation are covered by a collective bargaining agreement at an institution. We have already indicated that the fraction of occupations organized at an institution is highly correlated with whether any one of the occupations is organized across institutions. Given this fact, it is not surprising that the estimated union coefficients we obtained when we re-estimated the model by seemingly unrelated regressions (column 3) were very similar to the coefficients found in column (2) of the table. Any differences are probably due to sampling error, since the seemingly unrelated regression model could only be estimated using data on the subset of institutions that reported occupational salary and unionization data for all nine occupations.

Finally, our estimates of the salary advantage that staff members working in unionized academic environments have over staff members working in nonunion academic environments treat staff coverage by a collective bargaining agreement as being exogenous. If, for example, the institutions in which we observe staff covered by a collective bargaining agreement were initially the institutions in which staff compensation was lowest, other factors held constant, our estimates will understate the extent to which academic staff unions have improved their members' compensation

[12]The seemingly unrelated regression model was developed by Zellner (1962).

relative to the compensation of academic staff at institutions not covered by collective bargaining agreements.

In the absence of a panel data set that would permit us to estimate how changes in staff salaries at academic institutions are related to changes in collective bargaining coverage, we use the sample selection bias correction method first developed by James Heckman (1979) and Lung Fei Lee (1978).[13] Collective bargaining coverage for an occupation at an institution is assumed to be a function of the percentage wage gain that workers in the occupation might be expected to receive if they voted for union coverage, as well as variables that likely influence the workers' "tastes" for collective bargaining coverage.

Academic staff members' tastes for collective bargaining are assumed to be related to the proportions of private and public employees in the institution's state covered by collected bargaining agreements, each interacted with a dichotomous variable indicating whether the institution was public or private. This specification allows the impact of the magnitude of collective bargaining coverage for private and public employees in the state to differentially influence the probability of union coverage for a staff occupation at an academic institution, depending on whether the academic institution is public or private.

The expected percentage wage gain is assumed to depend on all of the variables that enter the occupational wage equation. This specification allows the impact on staff salaries of each explanatory variable found in equation (1) to differ for workers covered by and not covered by collective bargaining agreements.[14]

A two-stage approach is then followed. In the first stage we estimate a probit reduced form probability of union coverage equation for workers in each occupation, which is specified to be a function of all of the variables found in equation (1), along with the statewide proportions of employees who are union members in the public and private sectors, interacted with whether the academic institution is public or private.[15] The estimates of this equation allow us to compute an estimate of the inverse mills ratio for each observation; this is added as an additional explanatory variable, and equation (1) is then re-estimated. Inclusion of this estimated inverse mills ratio in the model controls for the nonrandom nature of union coverage.[16]

The estimated union coefficients that we obtained when the sample selection bias correction method was used are shown in column (4) of Table 5. In most cases these estimates prove to be very similar to the OLS estimates reported in column (2). The estimated union coefficients for custodians, carpenters, electricians, heating and cooling technicians, painters, and plumb-

[13] A similar survey was undertaken by the APPA in 1999–2000. However, the number of institutions present in both survey years is considerably smaller than the sample size used in this study, and the number of institutions that were in the survey in both years and actually saw a change in union coverage for a staff occupation during the period was close to zero. Hence, the longitudinal data cannot be used to provide estimates of the relationships that are of interest to us.

[14] We have allowed our estimated collective bargaining coverage effects to vary with some of the variables that enter the salary equations and found, for example, that in most cases collective bargaining coverage effects were not larger in the public sector than in the private sector and that, within the private sector, higher endowment levels per student were not associated with higher collective bargaining effects. However, in a few occupations they were larger for staff employed in two-year colleges than for staff employed in other academic institutions.

[15] A table with the estimated coefficients of the union coverage equation for each occupation is available from the authors upon request. The major interesting finding from these probits is that for several occupations, the proportion of public sector employees who are organized statewide does have a positive effect on the probability of observing the occupation at a public academic institution covered by a collective bargaining agreement.

[16] The coefficients of the estimated inverse mills ratios were negative in all nine occupational staff salary equations but were statistically significantly different from zero at the .05 (.10) level of significance in only one (three) cases.

ers remain statistically significant, and each coefficient is close to its value in the OLS equations. In five of the six occupations the selectivity corrected coefficients are 1–3 percentage points smaller. The estimated union coefficients for secretaries, groundskeepers, and locksmiths are statistically insignificantly different from zero, as they were in the OLS estimation.

Concluding Remarks

This study is, to our knowledge, the first published effort to estimate the effect of collective bargaining coverage on the salaries of staff members at American higher education institutions. When we treated collective bargaining coverage as exogenous, we obtained an estimated union/nonunion salary differential in the range of 9–11% for the occupations in our sample. The magnitude of the differential appears to be larger in 2-year institutions than in 4-year institutions, but does not vary between the public and private sectors. When faculty members are covered by a collective bargaining agreement, staff members in these institutions appear to gain an additional increment in salary of about 2 to 3 percentage points. Thus, it appears that there are spillovers from faculty unions to staff unions.

The impact of collective bargaining coverage on staff salaries varies widely across the occupations in our sample. While some unions, in particular many of those in the skilled building trades, appear to have won wage gains for their members in the 13–16-percentage-point range, other staff unions appear to have won much smaller wage advantages or none at all for their members. Generalizing our approach to allow for separate estimating equations by occupation, to allow for interdependencies of the error terms across equations, and to allow for staff units covered by collective bargaining agreements to be a nonrandom sample of all staff unions does not substantially alter our findings, although the last of those three modified specifications yields somewhat smaller estimated effects in several occupations.

The limitations of our study should be kept in mind. First, the sample of 163 academic institutions used in our study is not fully representative of the population of over 3,000 two- and four-year colleges and universities in the United States. The 22 occupations whose salaries we analyze all relate to employees from the facilities division of America's colleges and universities, and the effects we estimate for them are not necessarily representative of the effects for staff unions that one might observe for a wider range of college and university staff employed in other areas (for example, housing and dining, athletics, academic support, student services, and external relations).

Second, the data to which we have access contain no information on employee characteristics, such as education and experience. Thus they do not permit us to test the hypothesis that the higher salaries of staff members found at academic institutions with staff unions are offset by compensating staff members' productivity gains. These gains may occur because the higher salaries permit the institutions to hire better-quality employees, reduce staff turnover, and increase staff job tenure, or because collective bargaining coverage directly leads to improved staff morale and productivity.[17] Third, we have no information on the extent to which staff union contracts restrict the ability of academic institutions to contract out work and thus the extent to which academic institutions respond to salary gains won by staff unions by contracting out more work to external subcontractors. Hence we cannot conclude from our findings that the academic institutions are worse off (in a cost sense) from having staff unions.

Nonetheless, our study suggests that collective bargaining coverage does influence staff salaries in higher education. The *National Labor Relations Act* governs collective

[17]Numerous studies have tested the Freeman and Medoff (1984) exit-voice hypothesis and found that private sector unions tend to reduce employees' turnover rates. Daniel Rees (1994) found that faculty unions in the United States reduce turnover rates of associate and assistant professors.

bargaining for staff of private academic institutions, while state public employee bargaining laws govern collective bargaining for staff at public academic institutions. Our findings suggest that a direct way to achieve better salaries for college and university employees employed in many staff occupations is to encourage them to organize and bargain collectively. Unlike private college and university faculty members, who are effectively precluded from collective bargaining at many institutions because of the Supreme Court's decision in the *Yeshiva* case, there is no such prohibition to prevent staff at these institutions from organizing.[18]

We also find that when other factors are held constant, including the proxy for area wage levels and collective bargaining coverage, there is no evidence that more financially well-off academic institutions pay their staff higher salaries. Whether public pressure should be brought to bear on a wider range of academic institutions that have the financial resources to improve their staff members' salaries if they choose is an open question. We say "should be" rather than "can be" because the welfare implications of higher salaries for staff at higher education institutions are not obvious: higher staff salaries will, after all, mean higher costs and, ultimately, higher tuition levels, unless these increased costs can be offset by productivity gains.

[18]See *NLRB v. Yeshiva University,* 944 U.S. 672 (1980).

Appendix
Occupational Descriptions

Occupation	*Description*	*Occupation*	*Description*
Secretary	Secretarial/Clerical	UtilitiesOp	Utilities Operator/Maintenance
Custodian	Custodial/Housekeeper	GeneralMaint	General Zone Maintenance Worker
Groundskeeper	Groundskeeper	ElevMechanic	Elevator Mechanic
Carpenter	Carpenter	VehMechanic	Vehicle/Equipment Mechanic
Electrician	Electrician	Storekeeper	Storekeeper/Expediter
Locksmith	Locksmith	GeneralLabor	General Labor Worker
AC/Refrig	Air Conditioning and Refrigeration	Security	Security Worker
Painter	Painter	Machinist	Machinist/Welder
Plumber	Plumber/Pipefitter	Mason	Mason
Programmer	Computer Program/Analyst	Roofer	Roofing Worker
HVACTech	HVAC/Controls Technician	SheetMtlWrkr	Sheet Metal Worker

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