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Public sector unions and the free-rider problem

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Public sector unions and the free-rider problem

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To Mackenzie, Miles and Maxwell
Never be afraid to try

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CHAPTER 1 - PROBLEM STATEMENT AND LITERATURE REVIEW

1.1 Introduction

Many different approaches have been taken to try to identify models of union status/membership determination. The most successful course of action has to been to attempt to identify worker's preferences for union representation independent of the employment decision. If this is not done, it is difficult to discern if union status is due to preferences for union services or if it is simply a result of the employment process.

Many facets of collective bargaining have public good aspects. In the absence of a union shop, unions are required to represent all workers assigned to a given bargaining unit, regardless of whether a worker pays dues. Unions can influence which jobs are included in a bargaining unit, but are obliged to bargain for all workers regardless of membership status. Olsen (1965) claims that if a large group providing a collective good is to exist, it must be formed either with compulsory membership or with joint production of an excludable (club or private) good that it can tie to membership. Unions have used this argument to lobby for union security laws. The standard result is that if agents have incentives to free ride, there will be a sub-optimal provision of the public good. If the free rider problem is severe, the group will fail to form. However, unions have been able to exist in the absence of union security laws (Hundley 1993, Booth 1985). The question is how are unions able to overcome the free rider problem and arrive at a quasi-cooperative solution? Is it the reputation[1] or the "warm glow" effects that can only be obtained by being a union member or are unions able to

[1]In Booth (1985) reputation is a private good that can only be obtained joining the union. Thus, there is a disutility associated with being a free rider.

partially exclude non-members from wage gains, grievance procedures, and other union services that they are legally bound to provide to free riders?

The primary objective of this study was to identify a model of the membership/free rider decision conditional on current employment. Bargaining for wage increases and job security are the major services that a union provides to members and free riders. The ability to measure wage increases across individuals and across time is important in determining the relative importance of public versus private good aspects of wage bargaining. One would expect that an individual's membership decision would be more responsive to idiosyncratic wage increases than to wage increases that go to all employees. Grievance procedures are one way that unions can provide workers with a vehicle for job security. In times of lower demand, workers will perceive it more likely that they will be laid off or will be shifted to a different job. This would increase the probability that a individual worker would have a confrontation with management and thus make it more likely that the individual would want to use grievance procedures set up by the collective bargaining agreement. The rules and procedures for handing grievances are clearly a public good generated by the bargaining process. However, one can view an individual's use of the grievance procedures as a private good. This is particularly true if the union can and does pursue grievances more aggressively for union members.

A secondary objective is to model public sector quits. Labor supply to the public sector is of interest in its own right. However, there is an ulterior motive for looking at quits. Occupational choice and membership choice are not necessarily independent decisions. It is reasonable to assume that factors that affect union membership would also affect occupational choice. Those workers who have chosen not to remain in the public sector are not observed. Thus, there is a potential for non-random sample selection to bias the estimates. A model of quits can be used to develop controls for the potential selection bias.

I use a data set that is very rich in wage information which allows me to develop measures for the union's bargaining success that vary over time and across individuals. In addition, I can measure an individual's earnings relative to other employees in the individual's narrowly defined job in Iowa state government. In addition, I am able to measure employment changes within an individual's job or bargaining unit and within the state government as a whole. Thus, it is possible to determine the relative importance of private and public good aspects of union services in the union membership decision.

1.2 Previous Research

Many authors have examined various aspects of collective bargaining coverage and/or union membership. Studies of the allocation of workers between coverage and membership states have run headlong into the partial observability problem. That is, the employment decision and the representation decision are observed jointly. Farber (1983) points out that attempts to identify simple probit or logit models of union status when partial observability exists have not been very satisfactory. Instead, many authors have tried to model situations where the worker's preferences are directly or indirectly observable and/or the employment decision can be separated from the coverage/membership decision.

Farber and Saks (1980) developed a model of vote determination in National Labor Relations Board representation elections. The unique feature of this study is that workers' preferences for union representation are observed independent of the employment decision.

The authors use data gathered from interviews of workers who participated in NLRB representations elections. Interviewers observed how individuals voted in the election along with potential explanatory variables such as the individual's position on the intra-firm earnings distribution, seniority, race, sex, education, and age. Also observed were the individual's evaluation of the union's potential effect on relations with management, the probability of the worker being promoted, satisfaction with current job security, and the difficulty of finding an equivalent job.

The authors' main result is that workers at the lower end of the intra-firm earnings distribution tend be more likely to vote for union representation. This is presumably due to the empirical evidence that unions tend to raise average wage and decrease the variance of wage distributions within firms. However, this does not necessarily imply that they will join the union if it is certified (assuming they have the choice). Lower-paid workers may have the same or greater incentive to free ride as workers on the upper-end of the wage distribution.

The nonwage aspects of collective bargaining seem to have all the hypothesized effects. Workers were less likely to vote for the union if they felt that the labor/management relationship would deteriorate if a union were certified, if they felt that they had a good chance for promotion, or if they felt it would be difficult to find an equivalent job. Individuals that felt they were being unfairly treated or were dissatisfied with job security were more likely to vote for union certification. The authors found that sex, education and urban upbringing did not have a statistically significant impact on vote while age and race did have a significant impact on voting. Nonwhites and younger workers were more likely to vote for union representation. Farber and Saks suggest that older workers have fewer years of labor supply remaining to enjoy any benefits of unionization and could face more difficulties in finding alternative employment should they lose their current job. Nonwhites may perceive that they have a greater chance of experiencing discrimination by management and thus may benefit more by the establishment of union grievance procedures.

The importance of the Farber and Saks study is that it gets around the partial observability problem. The fact that it looks only at establishments that actually hold elections does limit the inferences that can be made about the population as a whole. Industries and firms with unions that have been established for a number of years may be substantially different from those currently holding elections. Thus, the results may not be representative of the population as a whole. In addition, the membership/free-rider decision is not addressed.

Farber (1983) develops a model of union status that incorporates two different decision-makers, workers and potential union employers, and allows for an excess supply of workers for union jobs. Using a 1977 cross section of data from the Quality of Employment Survey, Farber is able to identify individuals in nonunion jobs that desire union coverage. At the time of hire, workers in the sample are observed in one of three possible states: employed in a union job and preferring a union job, employed in a nonunion job and preferring a nonunion job, and employed in a nonunion job but preferring a union job. The author assumed that if a worker was employed in a union job he/she desired a union job. Thus, excess demand for union jobs is restricted to be non-negative. This restriction certainly seems to be reasonable and should not limit the usefulness of the results.

The author argues that modeling the union status determination in this fashion is much more enlightening than simple probit or logit analysis where both supply and demand factors are affecting the allocation of workers across coverage states. Farber points out that Poirier (1980) did develop an approach to identify and estimate models of union status when partial observability exists. However, the results required strict assumptions on functional forms that Farber claims have not proven useful in empirical applications.

The results from Farber's queuing model are quite interesting and shed a great deal of light on the determination of union status in the labor market. Farber's calculations indicate that nonwhites are significantly more likely to desire union jobs and to be working on a union job. The probability that a nonwhite is hired to a union job given that he/she desires a union job is not found significantly different than for whites. The low incidence of unionized workers in the south is found to be due both to lower demand for union jobs and lower supply of union jobs. The author also finds differences in unionization for his four major aggregate job classifications: professional and technical, service, clerical, and blue-collar workers. While blue-collar workers are more likely to be unionized than are other types of workers, the source of the variation is different for each job class. Clerical workers are less likely to be unionized because they are less likely to desire a union and less likely to be hired to a union job given that they desire one. Service workers are no less likely to desire a union job than blue-collar workers. However they are much less likely to be hired to a union job given they desire one. Professional and technical workers are less likely to be unionized because they are less likely to desire a union job and not because they are less likely to find a union job. Furthermore, older workers are less likely to be unionized predominantly because they are less likely to desire union representation.

Chaison and Dhavale (1992) use a method similar to that used here. They examine the choice between union membership and free riding using 1988 Current Population Survey data. Respondents to the CPS indicate if they are currently covered by a collective bargaining agreement and whether they are currently members of a union or employee association similar to a union. The analysis is limited to union workers in the twenty-one right-to-work states and pubic sector union employees where union security laws are illegal. Thus, each worker clearly has the choice between free riding or becoming a union member. However, the authors do reference studies that indicate that a fair percentage of residents of right-to-work states do not understand that they have this choice.

Chaison and Dhavale conduct probit analysis of the union member/free rider decision. Their results are not dramatically different from previous studies. They find that female, white, part-time, and more educated workers are more likely to free ride while older workers are more likely to be union members. Weekly earnings are negatively related to free riding. This result seems to run counter to Farber and Saks (1980). However, the results are not necessarily contradictory. It is not surprising that lower wage jobs are more likely to become unionized. This certainly does not imply that low-wage workers will have less incentive to free ride once the job in unionized. It is conceivable that they would have more. Income constraints may preclude them from becoming union members. Also, if wage gains and union membership are both endogenous, one would expect to find unions with the smallest degree of free riding at the upper end of the wage distribution.

The richness of wage information is a significant advantage that my study has over Chaison and Dhavales'. In addition to a current wage rate, I also know the position the individual occupies on the wage distribution relative to other workers in the same job. Also, I can relate his pay gains to the pay of other workers in similar jobs within and outside of the state government. Furthermore, I have intertemporal variation in the union's ability to secure wage gains for each particular job and for all jobs within the state government. This quality of data has not been found in other studies. In addition, the data compiled in this study includes pricing information for the unions. This allows me to estimate the effects of the dues rates and the substitutability of other workers' contributions for their own contributions.

Hundley (1993) develops a multinomial logit model of union status of public sector workers. The author uses 1985 Current Population Survey data to identify public sector workers occupying one of three states: not covered by collective bargaining, covered members and covered nonmembers. This distinction is important since the incidence of covered non-membership is much greater in the public sector than in the private sector. Using interstate data, the author examines the effects of various bargaining laws on union coverage and membership. The author also includes a host of personal and job characteristics in the model. However, the author provides little justification for their inclusion and does not report the parameter estimates for these variables. Thus, the results are not extremely useful for my study since preferences for collective bargaining are examined within a set of bargaining laws rather than across legal frameworks.

Booth (1985) is one of the few theoretical papers on union membership. Booth develops a social customs model of union membership in a theoretical context. Agent's utility functions are assumed to be defined over reputation and exogenously determined wages. The agents' reputations are determined by their choice to join the union or to free ride and how other agents in the model view their actions. Thus, the ability of others in the community to impose sanctions on free riders greatly influences the ability of the union to overcome the free rider problem and survive. Unions that have more homogenous members and that have members who work in close proximity would seem to have an advantage in imposing sanctions on free-riders. The author demonstrates that under certain conditions even large unions can overcome the free rider.

1.3 Overview of the Dissertation

This study develops a model of union membership. This model will allow for workers covered by collective bargain to make the choice between union membership and free-riding on the contribution of others. The ability to completely free ride is a feature found in the public sector and private sector where "Right to Work" laws have been enacted. Data for Iowa state employees is used to estimate the model. In addition a model of public sector quits is also developed. The estimates from this model are used to develop statistical controls for the potential bias from non-random selection of public sector workers.

Chapter 2 develops a theoretical model of union membership with a discrete membership choice. Chapter 3 provides a description of the data used. Chapter 4 develops the empirical model of union membership with the Iowa state employee data. Public sector quits are analyzed in Chapter 5. This chapter includes estimates of the impact of comparable worth wage gains on quits. The union demand model is reestimated with selection controls in Chapter 6. The estimates are used to calculate the impact of comparable worth on union membership. Chapter 7 summarizes the results and provides direction for further study.

CHAPTER 2 - THEORETICAL MODEL OF UNION MEMBERSHIP

2.1 Introduction

Unions are required to provide bargaining services to all covered workers without regard to workers' membership in the union. unless state laws allow union security clauses to be included in collective bargaining contracts. Union security clauses allow unions to force all workers covered by the collective bargaining contract to pay union dues. Unions cannot force workers to become members of the union because of freedom of association. They can require workers to pay fees to cover the costs of bargaining. The fees are typically less than the dues required to be a full member of the union. When unions obtain a collective bargaining contract, the contract covers all workers regardless of their union membership. In cases where workers cannot be forced to pay dues, it is quite plausible that unions would still have a mechanism for excluding non-payers for a portion of the services that they are legally bound to provide. This provides the motivation for describing union membership in the context of a joint product model similar to that used in Sandler and Murdoch (1990). Paying union dues will be modeled as producing public and private services for union workers. The ability of the union to withhold services from nonmembers would then depend on the degree to which membership produces public versus private goods and how each is valued by the worker.

2.2 Joint Product Model of the Demand for Union Services

Assume that an individual i chooses voluntary contributions to the union, x_i. Using x_i, the union produces a pure public good, z_i, and a private good, q_i, in the following manner:

$$q_i = f(x_i) \tag{2.1}$$
$$z_i = g(x_i) \tag{2.2}$$

f' and g' are strictly greater than zero and f'' and g'' are less than or equal to zero. Since z_i is a public good, agent i consumes the total amount of z that the union produces. That is, he or she consumes

$$Z = Z_i^* + z_i . \tag{2.3}$$

Z_i^* is the total amount of the public good produced by agents other than agent i. Assume that all individuals have identical production functions for the public good that exhibit constant returns to scale. That is, contributions made by others will be perfect substitutes for their own contributions in the production of the pubic good Z. Making this assumption allows us to express the total quantity of the public good as

$$Z_i = g(X_i^* + x_i) \tag{2.4}$$

Where

$$X_i^* = \sum_{i \neq j} x_j \tag{2.5}$$

Agents derive utility from consumption of the jointly produced public and private goods and from a composite good, y, which is purchased in a perfectly competitive market. Utility for agent i can thus be expressed as

$$u_i = u_i(y_i , q_i , Z_i^* + z_i). \tag{2.6}$$

u_i is maximized with respect to the budget constraint $I_i = p_y y_i + p_x x_i$. I_i is the nominal income of agent i and the p's are the respective prices. The utility maximization problem for agent i is

$$\underset{x_i, y_i}{Maximize} \; u_i(y_i , f(x_i), g(X_i^* + x_i)) \tag{2.7}$$

$$Subject\ to: I_i = p_y y_i + p_x x_i .$$

The model above is similar to that used by Sandler and Murdoch (1990) to determine Nash-Cournot versus Lindahl behavior for military alliances. They derive demand equations for the aggregate level of x. By including spill-ins in the income term[2] a nested test of the joint product model can be obtained by testing for the existence of an X^{*} term in the demand function. When the value of spill-ins is included in the income term, the pure public good model implies that X^{*} should not be a separate argument in the demand function. X^{*} contributing significantly implies that that this is an impure public good. The joint product model may then be appropriate. I express the demand functions in term of the individuals contributions. Clearly, the two methods are equivalent. Actually observing the level of public and private goods an agent receives would provide a direct test of the publicness of the good.

The first order conditions from the maximization problem in *(2.7)* are

$$\frac{\partial u_i}{\partial y_i} - \lambda p_y \leq 0, \tag{2.8A}$$

$$\frac{\partial u_i}{\partial q_i} f'(x_i) + \frac{\partial u_i}{\partial z_i} g'(x_i + X_i^{*}) - \lambda p_x \leq 0, \text{ and} \tag{2.8B}$$

$$I_i - p_y y_i - p_x x_i. \tag{2.8C}$$

Assuming an interior solution, the system of equations in *(2.8)* can be solved for demand functions and indirect utility functions. The demand for union services (the supply of contributions to the union) would be

[2]This is done by deriving the demand equations in terms of aggregate contributions. Thus the value of spill-ins $p_x X^{*}_i$, is added to the income term in the demand equation for X. The joint product model implies that the reduced form demand for X will still have an additional X^{*} as an argument. This specification can be used to test the joint product versus the pure public good model.

$$x_i = x_i(I_i, p_y, p_x, X_i^*). \tag{2.9}$$

This demand function is the optimal contribution to the union by an individual. Comparative static exercises can be performed to identify how optimal contributions change when prices, income, or contributions of others change. A comparative static that will be of particular interest will be how the optimal contribution changes as X^* changes. In general $\partial x/\partial X^*$ will be ambiguous. An increase in X^* causes an increase in spill-ins. The income effect implies that the agent would want to consume more of all goods. Because the agents are now given more of the public good, they would tend to substitute towards the private goods and away from contributions to the public good. A further ambiguity is introduced through the jointness in production of q and z. The only way an individual can consume more of the private good q is to also produce more z. As will be demonstrated, the results are ambiguous.

With regularity conditions on preferences and technology and a few additional assumptions, most of the comparative static results, including $\partial x/\partial X^*$ have unambiguous signs. To this end, assume that

$$\frac{\partial^2 u_i}{\partial y_i \partial z_i} \geq 0, \tag{2.10A}$$

$$\frac{\partial^2 u_i}{\partial y_i \partial q_i} \geq 0, \text{ and} \tag{2.10B}$$

$$\frac{\partial^2 u_i}{\partial q_i \partial z_i} \leq 0. \tag{2.10C}$$

The conditions above are needed to provide unambiguous signs for most of the comparative statics. The conditions imply that y and q, and y and z are net substitutes in consumption, and that q and z are net complements. There is no overwhelming reason to think that these conditions will hold. However, the conditions in *(2.10)* are certainly reasonable, especially for the problem at hand. It is plausible that q and z are identical goods except for the union's ability to exclude non-payers from consumption. That is the goods produced by the union are partially excludable. Thus, q and z would be additive in the utility functions. The following example will illustrate this point. Assume that union services are completely represented by wage increases. It is conceivable that the union can reserve part of the wage increase for members only with the remaining increase shared by all members. In this case, the workers' utility would depend on only the sum of the public and private wage increases. The conditions in *(2.10)* will hold in general for all strongly additive utility functions. Strongly additive utility functions have cross partial derivatives that are equal to zero.

The derivation of the comparative static results is in appendix A1. Table 2.1 summarizes the directions of change in x and y when income, prices and contributions of others change. All of the results in Table 2.1 assume that the conditions in *(2.10)* hold. Without these conditions, all of the comparative statics in Table 2.1 would have ambiguous signs. Especially interesting is the result that individual contributions decrease as contributions of others increase. This implies that individuals substitute away from own production of q and z when X^* increases.

Table 2.1 - Summary of Comparative Static

Partial of y_i	Sign	Partial of x_i	Sign
$\partial y_i/\partial X^{*}$	(+)	$\partial x_i/X^{*}$	(-)
$\partial y_i/\partial I_i$	(+)	$\partial x_i/\partial I_i$	(+)
$\partial y_i/\partial p_y$	(-)	$\partial x_i/\partial p_y$	(?)
$\partial y_i/\partial p_x$	(?)	$\partial x_i/\partial p_x$	(-)

2.3 Economic Efficiency of Union Contributions

x_i is the individuals' contribution in a competitive environment. This in general will not be a Pareto efficient level of contributions. Agents do not take into account the spill-ins that they create for other agents when they produce more q and z. Thus, they tend to stop short of socially optimal levels of contributions. Pareto efficient allocations could be obtained if all individuals were required to contribute to the collective good up to the point where the social marginal rate of substitution between x and y equals the relative price. To see this, consider an economy with n agents numbered $1,2,\ldots,n$. Each agent has optimization problem of the form in *(2.7)*. The conditions for Pareto optimal allocation can be derived from the following optimization:

$$\begin{aligned}
&\underset{x_1\ldots x_n, y_1\ldots y_n}{Maximize} \quad u_1(y_1, f(x_1), g(x_1 + X_1^{*})) \\
&\textit{Subject to:} \quad u_i(y_i, f(x_i), g(x_i + X_i^{*})) \geq U_i^0 \quad \forall i = 2,3,\ldots,n \\
&\qquad \sum_{i=1}^{n} x_i \leq X^0 \\
&\qquad \sum_{i=1}^{n} y_i \leq Y^0
\end{aligned} \qquad (2.11)$$

X^0 and Y^0 are the total amount of x and y available to the economy. U_i^0 is a fixed level of utility for each individual i equal to 2 through n. The first order conditions for a Pareto optimal allocation can be expressed as

$$
\begin{aligned}
&\frac{\partial u_1()}{\partial q_1} f'(x_1) + \frac{\partial u_1()}{\partial z_1} g'(x_1 + X_1^*) + \sum_{i=2}^{n} \gamma_i \frac{\partial u_i()}{\partial z_i} g'(x_i + X_i^*) - \lambda_x \le 0 \\
&\gamma_i \frac{\partial u_i()}{\partial q_i} f'(x_i) - \gamma_i \frac{\partial u_i()}{\partial z_i} g'(x_i + X_i^*) + \sum_{\substack{k=1 \\ k \ne i}}^{n} \gamma_k \frac{\partial u_k}{\partial z_k} g'(x_k + X_k^*) - \lambda_x \le 0 \quad i = 2,3,\ldots,n; \quad \gamma_1 \equiv 1 \\
&\frac{\partial u_1}{\partial y_1} - \lambda_y \le 0 \\
&\gamma_i \frac{\partial u_i}{\partial y_i} - \le 0, \quad i = 2,3,\ldots,n \\
&u_i(y_i, f(x_i), g(x_i + X_i^*), T_i) - U_i^0 \le 0 \quad \mathit{for\,all}\ i = 2,3,\ldots,n \\
&X^0 - \sum_{i=1}^{n} x_i \le 0 \\
&Y^0 - \sum_{i=1}^{n} y_i \le 0
\end{aligned}
\tag{2.12}
$$

where γ_i's are Lagrange multipliers that represent to the shadow price of the other agents marginal utility.

In general, the competitive allocation will result in less than the efficient level of x being contributed. This occurs because there is a positive increment to j's utility when i increases contributions to the union. The marginal benefit of spill-ins is represented by

$$\sum_{\substack{k=1 \\ k \ne i}}^{n} \gamma_k \frac{\partial u_k}{\partial z_k} g'(x_k + X_k^*) \ . \tag{2.13}$$

This is the benefit that others receive when agent i contributes to the union. This is not captured in a competitive pricing system. Cooperative solutions could reach a Pareto superior allocation if agents could be subsidized for these spill-ins. Agents could be taxed the value of the marginal benefit they receive in spill-ins. However, without some preference revelation mechanism this is generally not possible. Agents usually have an incentive to understate their true benefit from spill-ins. Thus, voluntary contributions will not reach optimal levels of provision by Pareto criteria.

2.4 Joint Product Model with Discrete Contributions

Imposing further restrictions on the maximization in *(2.7)* will make the analysis more realistic. Assume that individuals are faced with the decision to contribute voluntarily to the union (become members of the union) but they have to contribute an exogenously determined[3] dues level x_i^u. That is, they have to purchase union services in discrete increments. In addition, assume that agents have identical utility functions except for a vector of taste parameters T. Individuals will choose either to pay dues, or to contribute nothing and completely free-ride on other members' contributions. The individual's problem in *(2.7)* becomes a discrete choice of whether to contribute to the union with the additional constraint that x_i must be equal to constant amount $x^{\cdot}$. This can be expressed as

$$\underset{i}{Maximize}\, u_i(y_i, f(x_i), g(X_i^{*}+x')) \qquad (2.14)$$
$$Subject\ to:\ I_i = p_y y_i + p_x x'$$

This maximization yields an indirect utility function of the form

$$V(x', p_y, p_x, I_i, X_i^{*}). \qquad (2.15)$$

The agent chooses between two different allocations: one with $x'=0$ and one with $x'=x^u$. Thus, an agent will join the union if and only if

[3]Exogenous to the worker. x^u could be set endogenously by unions.

$$V(x' = x^u, p_y, p_x, X_i^*, I_i, T_i) - V(x' = 0, p_y, p_x, X_i^*, I_i, T_i) \geq 0. \tag{2.16}$$

Otherwise, the individual will choose to free ride on the production of the public good provided by other union members. Furthermore there is a critical value of x^u, call it x^c, beyond which it is no longer optimal for the individual to contribute to the union. By definition x^c solves

$$V(x' = x^c\ p_x, p_y, X_i^*, I_i, T_i) - V(x' = 0\ p_x, p_y, X_i^*, I_i, T_i) = 0. \tag{2.17}$$

Provided that the union can set dues such that x^u does not exceed x^c, the individual will be a member of the union. Otherwise, the individual will choose to free ride on the provision of the public good by other agents.

Figure 2.1 graphs $V(x'=x^u,...)$ and $V(x'=0,...)$ as x^u varies holding other arguments constant. This illustrates that as long as the required contribution is less than the critical value the individual will choose to be a union member. Once x^u exceeds x^c, the agent no longer joins the union and will only purchase the composite good y. $V(x'=x^u,...)$ has a unique

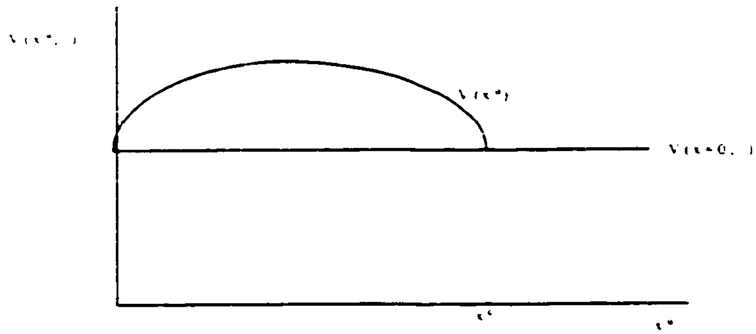

Figure 2.1 - Member and Non-member Utility

maximum that corresponds the optimal x_i in *(2.7)*. In other words, this is what the agent would do if she were not constrained to contribute to the union in discrete amounts.

It is common to assume in the context of the unconstrained problem *(2.7)* that an interior solution will be reached and that all agents would contribute. That is equivalent to saying that for equilibrium prices and income, $V()$ is positively sloped for all agents at $x'=0$. Clearly in the constrained model there may be complete free-riders for whom $x^u > x^c$. The interior solution assumption is usually made when modeling collective behavior of large groups such as cities, states or countries where a representative agent's utility is being maximized. In this case we want to model the decisions of heterogeneous individuals. Thus for *(2.14)*, there is not a compelling reason to assume that x^c will always be positive. Given prices, tastes, income and X^*, it may well be that x^c is zero for some individuals.

Figure 2.2 represents the utility maximization problem in *(2.14)*. Point A represents the consumption bundle of a free-rider. Point B is the consumption bundle of a union member. Given the constraints in the optimization problem, the only feasible budget set is bounded by line $(I/p_y)AX^*$ but also includes the point B. Thus, the efficient consumption set contains only point A and point B. The slope of line AB is equal to the negative of the price ratio. The line AB is the part of the budget constraint from *(2.7)*. Unless agents are identical, most agents optimization will result in a corner solution. The marginal rate of substitution will not be set equal to the price ratio. Given that the prices in the model reflect the true social cost of x and y, then a competitive solution has an additional source of inefficiency. How it compares to the allocation resulting from *(2.7)* is not clear. It depends upon how close the required contribution is set to each individual's optimal contribution. A union that could

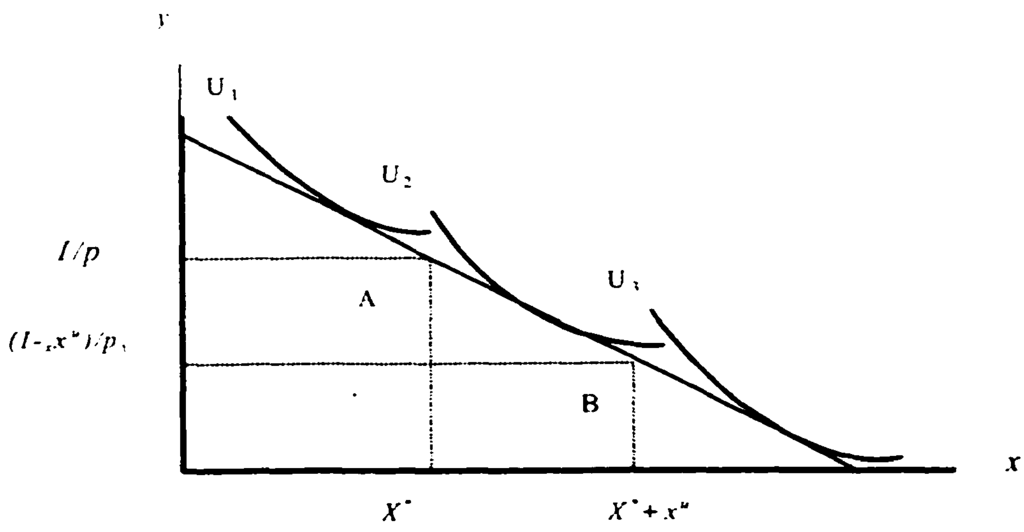

Figure 2.2 - Utility Maximization with Discrete Union Contributions

perfectly discriminate could set prices such that each agent contributes the socially optimal amount[4]. In this case, there would be no complete free-riders. However, there could be members who are not required to contribute. It is likely that x^u is set in such a manner that the aggregate contributions would be further from the socially optimal solution than the allocation resulting from *(2.7)*. Furthermore, because most agents will not be setting their own marginal rate of substitution equal to relative price, an additional source of inefficiency is added[5]. However, making efficiency comparisons is in general not possible. Cases exist such that *(2.7)* and *(2.14)* would result in Pareto non-comparable allocations.

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[4]This would mean setting the price of union services equal to marginal social benefit of x.
[5]unless x^u is such that each agent is contributing the socially optimal amount.

Sufficient but not necessary conditions for membership and for complete free-riding can also derived by examining the slope of the indifference curve at point A and point B. A sufficient condition for free riding would be if at point A

$$\frac{\partial u_i / \partial x_i}{\partial u_i / \partial y_i} \leq \frac{p_x}{p_y}. \tag{2.18}$$

In this case the indirect utility function is negatively sloped at $x^u = 0$. This corresponds to indifference curve U_1 in figure (2.2). No matter how small the required contribution is the agent will never choose to join the union (unless x^u was zero). Likewise, a sufficient condition for union membership would be if at point B

$$\frac{\partial u_i / \partial x_i}{\partial u_i / \partial y_i} \geq \frac{p_x}{p_y}. \tag{2.19}$$

In this case the optimal value of x is greater than the required contribution. This corresponds to the indifference curve U_2 in figure (2.2). The agent would produce more of q and z if contributions were not constrained at x^u. When neither of these conditions hold, knowing the slope the indifference curves is not enough. One must look at the level of utility at the two points A and B. This corresponds to indifference curve U_3 in figure (2). This is the case where the indifference map reaches a tangent on the line segment AB

An individual will only pay dues if the $x^u < x^c$. In other words, x^c is the maximum amount that an individual would be willing to contribute to insure that the level of public good is incremented above Z^*. The probability that an individual joins the union is the probability that $x^u < x^c$. That is

$$P(x_i = x^u) = P(x_i^c > x^u) = P(V(x_i = x^u) > V(x_i = 0)). \tag{2.20}$$

The question is, how does the willingness to pay for union services, x^c, vary as the demographic and job characteristics vary? In addition, what are the relative importance of the public and private good aspects of union services? Existence of private good aspects of bargaining services implies that unions are able to exclude, at least imperfectly, non-members from receiving the benefits that unions are legally bound to provide to all covered workers. Paying union dues reveals individuals' willingness to pay for union services, especially if the public and private components can be dichotomized. It is the willingness to pay that will be empirically modeled.

2.5 Supply of Union Services

Union management must decide how to set the dues charged for its services. Assume that union must provide equal services to all covered workers without regard to membership. Therefore, costs of providing union representation does not vary with union membership or the required contribution x^u. Costs could vary as the total number of covered workers varies. This seems to be reasonable in the short run since the size of the bargaining unit is deter-mined before the bargaining process. In addition, a considerable portion of bargaining costs could be fixed. Given this, unions can be viewed as maximizing total dues revenue by choosing the price for union membership p_x. Total union membership is simply the number of workers whose critical value, x^c, equals or exceeds union dues. If the union could set a p_x for each individual, they would simply set p_x such that $x^c = x^u$. In this case all workers would choose to join the union. A more realistic scenario is that the union must set dues based on uniformly based on occupational characteristics.

Define M as total membership of the union. M can be expressed as

$$M = \sum_{i=1}^{n} m_i(x^u, p_y, p_x, I_i, X_i^*, T_i) \tag{2.21}$$

Where

$$m_i(x^u, p_y, p_x, I_i, X_i^*, T_i) = \begin{cases} 0 \text{ } if \text{ } x^u > x^c \\ 1 \text{ } if \text{ } x^u \leq x^c \end{cases} \tag{2.22}$$

To maximize resources the union would set p_x such that

$$\varepsilon_x = \frac{\%\Delta M}{\%\Delta p_x} = -1. \tag{2.23}$$

Thus, unions would set dues at the point where the price elasticity of membership equals 1.

The union can maximize collections by setting dues such that an agent's critical value is as close to x^u possible. Thus, a union that is able to set dues over a more homogenous set of workers will have an advantage in collecting dues.

CHAPTER 3 - DESCRIPTION OF THE DATA

3.1 Introduction

One's ability to identify individual demand for union services is heavily dependent on the ability to observe workers' preference for union representation and member services. This can be difficult when union membership is not voluntary. In addition, measuring one of union primary duties, to bargain over wages, is sometimes difficult. This is especially acute in public sector labor markets. Wage gains in the public sector have been traditionally across the board increases with little change in relative wages. This makes it difficult to separate the effects of wages from other effects.

The primary source of data for this study is payroll records for state employees in Iowa. This provides information on state wages, employment status and demographic information. Wage information is also derived from the Current Population Survey data. This information is used to measure external wages.

3.2 Description of the Data

The data consists of payroll records for employees in the Iowa state government. Each individual's payroll record is observed every December from 1980-1992. Employment in the Iowa state government runs in the neighborhood of 18-20,000 employees each year. Approximately 13,000 of these employees are in unionized jobs. Thus, there is an ample amount of data to work with.

The variables used to model union membership and quits are defined in the next section (3.3.) The data includes employees in jobs that are covered by a collective bargaining contract and those workers whose job is not covered by collective bargaining.[6] The membership status is also observed for those workers covered by collective bargaining.

Data on the dues required for membership was compiled and merged with the payroll data. For those workers who pay dues this is straightforward. The payroll data indicates for what local the dues are being withheld. For non-members it is more difficult to determine to which local they would pay. For most situations, all employees in same job and division[7] paid to the same local. However, there were cases where no employees in a particular job and division paid union dues. In these cases, the divisions were combined by successively truncating the division code. With each truncation, the dues structure for the union/local associated with most employees was merged into the data where dues had not yet been determined. This was repeated until all covered employees had been assigned a dues structure. Once this information had been merged into the data, measures for the required dues contributions and the contributions of others in the bargaining unit were derived.

The average dues rate each year in real and nominal terms is displayed in Figure 3.1. There is not a great deal of change in real dues over time. Most of the dues rates work out to about 1.2% of the individuals biweekly salary, in other words, about one hour of the standard 80 hour pay period. However, there are differences between the locals, and most of the locals changed their pricing structure at least one time during the sample period.

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[6] Some jobs are not eligible for collective bargaining while other jobs are eligible but are not unionized. The majority of traditionally union jobs in the Iowa state government are organized. The allocations of individuals across coverage states will not be modeled. In other words, I model the union membership decision given that the worker is in a union job.

[7] These are the five digit job classification codes and ten digit division codes in the payroll data.

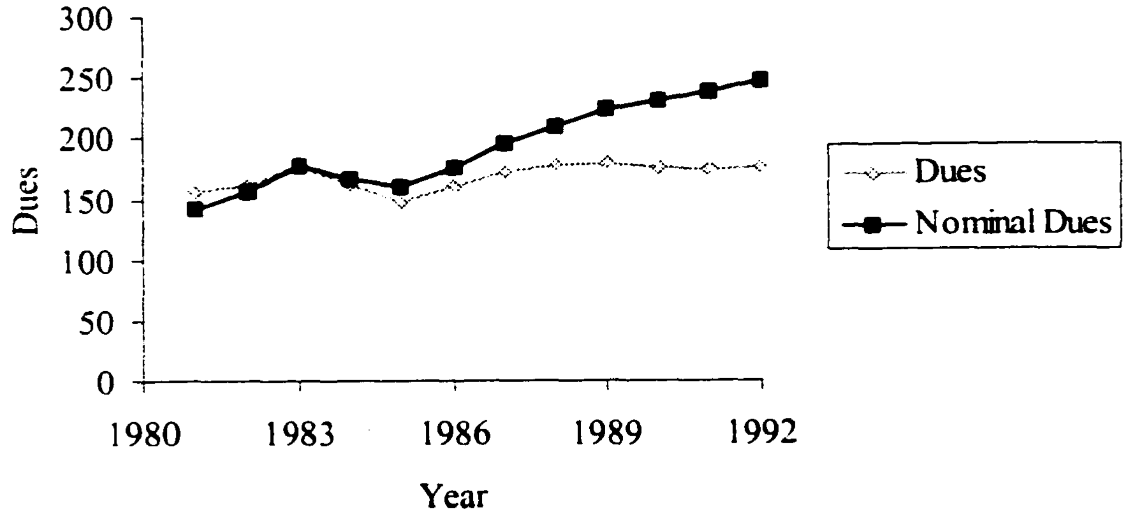

Figure 3.1 - Average Annual Dues Rates in Real and Nominal Terms, 1981-1992

Measures for external wages are taken from Current Population Survey data. Wages for detailed jobs are taken from the March Current Population Survey from 1980-1993[8]. The March Survey includes data on hourly earnings and occupation for a portion of the respondents. An average hourly wage is computed for each CPS detailed job. Two-year averages are then taken for each detailed occupation. These detailed jobs are mapped to a corresponding job or jobs within the Iowa state government payroll data and the CPS wages were merged with the payroll data based on the occupational mapping.

[8] The hourly wage data was not available for 1980-1982. This data was imputed using the predicted values from a regression of the log of hourly earning on the log of annual earnings, hours, sex , age and education and the Consumer price index and a trend. At the individual level, this regression explained 62% of the variability in log hourly earnings. The predicted values were used to compute an average hourly wage for each occupation.

Averages for individual and job characteristics are reported in Table 3.1 for union members, covered nonmembers and employees not covered by collective bargaining. Consistent with previous studies, wages for jobs that are not covered by collective bargaining are on average higher than unionized positions. This is consistent with previous studies that suggest lower wage jobs are more likely to become unionized. Also, free riders seem to be

Table 3.1 - Descriptive Statistics for Iowa State Government Employees by Collective Bargaining Status, 1981-1992

	Union Members		Covered Non-Members		Not Covered	
	Mean	**Std. Dev.**	**Mean**	**Std. Dev.**	**Mean**	**Std. Dev.**
Exit[a]	0.067	0.250	0.082	0.274	0.094	0.292
New Entrant	0.029	0.167	0.105	0.307	0.085	0.279
dln(Minimum Wage)	0.053	0.047	0.049	0.041	0.045	0.042
dln(Wage)	0.074	0.100	0.074	0.126	0.075	0.128
dln(CPS Wage)	0.043	0.040	0.042	0.049	0.057	0.067
ln(Wage)t-1	6.575	0.288	6.502	0.308	6.741	0.474
ln(CPS Wage)$_{t-1}$	6.433	0.302	6.422	0.303	6.554	0.375
Pay Step	1.164	0.093	1.139	0.096	1.165	0.117
dln(FTE)	-0.006	0.033	-0.002	0.034	0.003	0.033
Over time Indicator	0.523	0.499	0.427	0.495	0.087	0.281
Dues	167.911	37.654	169.580	31.530	0.000	0.000
Total Dues /1,000	156.668	96.087	138.202	90.162	0.000	0.000
State Tenure	10.500	7.491	9.766	8.042	11.676	9.021
Prior Experience	13.035	9.319	14.052	10.073	13.436	9.777
Part Time	0.012	0.109	0.035	0.184	0.055	0.228
Non-White	0.042	0.201	0.043	0.203	0.036	0.185
Female	0.429	0.495	0.492	0.500	0.509	0.500
Married	0.661	0.473	0.645	0.479	0.697	0.459
Manager	0.000	0.000	0.000	0.000	0.145	0.352
Professional	0.137	0.344	0.183	0.387	0.412	0.492
Technical	0.116	0.320	0.175	0.380	0.056	0.230
Clerical	0.153	0.360	0.309	0.462	0.290	0.454
Service/Blue Collar	0.594	0.491	0.333	0.471	0.097	0.296
n	55,721		103,380		76,995	

[a] The variable Exit only has values for employees in 1981 to 1991 since 1992 is the final year of data.

paid less and receive smaller wage increases on average than union members.

Statewide membership rates for males and females are displayed in Figure 3.2. Union membership trended upward in the late 1980's and early 1990's. Presumably, a large part of this upswing is the result of comparable worth wage gains. A portion of the increase in membership could be related to a standoff between the state governor and the unions over

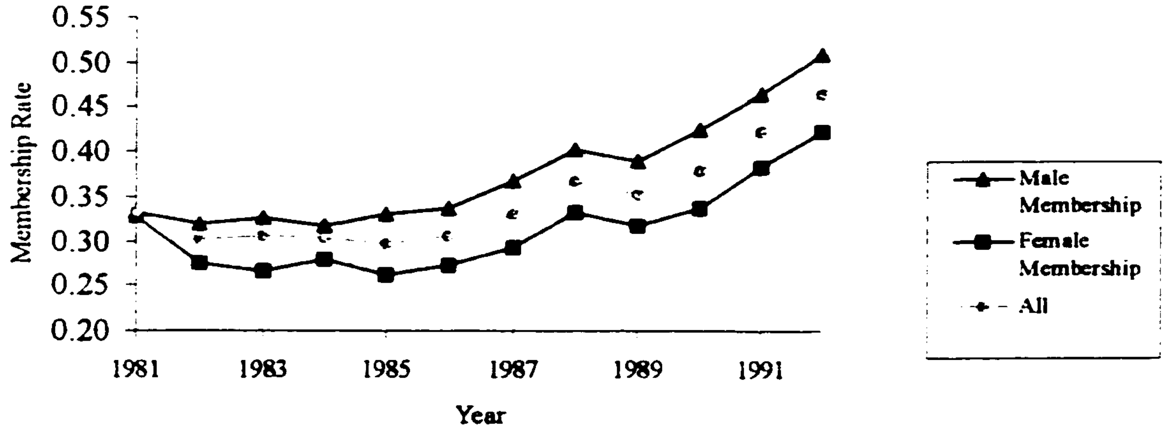

Figure 3.2 - Union Membership Rates for Males, Females and Overall, 1981-1992

wage increases the state tried to hold back during budget shortfalls that occurred at the end of the 1980s. Whatever the cause, the models estimated for both union membership and for quits will include annual dummy variables to control for macro effects. Males have a uniformly higher incidence of union membership than do females. This is not to say that females are more likely to be free riders than otherwise equivalent males. Some might suggest that this would be due to females being less attached to the labor market. Farber and Saks (1980) find evidence to the contrary. Females do tend to inhabit the lower portion of the wage distribution. It may be that otherwise equivalent females are more likely to be union members. The results of this study will help to resolve this question.

Descriptive statistics for covered workers are reported by major bargaining unit in Table 3.2. Employees are organized by bargaining unit. There are some non-bargaining unit employees in each bargaining unit. These would be typically managers and certain other staff that are not eligible for collective bargaining. The clerical unit was not organized until 1985.

Table 3.2 - Averages by Major Bargaining Unit for Iowa State Government Employees Covered by Collective Bargaining, 1981-1992

	Clerical	Technical	Blue Collar	Fiscal and Staff	Social Services	Security	Public Safety
Member	0.159	0.347	0.347	0.154	0.382	0.531	0.933
Exit*	0.078	0.078	0.074	0.077	0.083	0.085	0.035
New Entrant	0.092	0.076	0.066	0.079	0.074	0.112	0.041
dln(Minimum Wage)	0.046	0.053	0.049	0.044	0.055	0.049	0.057
dln(Wage)	0.082	0.076	0.065	0.079	0.075	0.072	0.066
dln(CPS Wage)	0.038	0.046	0.035	0.053	0.037	0.045	0.049
$\ln(\text{Wage})_{t-1}$	6.384	6.493	6.397	6.837	6.653	6.575	6.813
$\ln(\text{CPS Wage})_{t-1}$	6.384	6.333	6.307	6.849	6.362	6.626	6.689
Pay Step	1.163	1.151	1.130	1.156	1.141	1.131	1.221
dln(FTE)	-0.007	-0.005	-0.006	0.008	0.005	-0.006	-0.004
Over time Indicator	0.356	0.548	0.720	0.067	0.030	0.797	0.130
Dues	161.353	176.697	168.823	199.411	140.471	188.423	114.827
Total Dues /1,000	69.381	247.095	167.539	40.724	53.220	137.883	56.507
State Tenure	8.955	10.586	10.751	10.054	9.398	7.529	13.820
Prior Experience	15.687	12.081	15.304	14.217	13.239	14.774	7.544
Part Time	0.072	0.020	0.028	0.010	0.028	0.006	0.000
Non-White	0.069	0.038	0.037	0.040	0.044	0.044	0.018
Female	0.963	0.541	0.203	0.367	0.718	0.096	0.040
Married	0.570	0.610	0.720	0.615	0.632	0.724	0.801
Professional	0.001	0.011	0.000	0.916	0.620	0.005	0.157
Technical	0.000	0.327	0.177	0.054	0.000	0.134	0.038
Clerical	0.998	0.186	0.053	0.030	0.372	0.004	0.000
Service/Blue Collar	0.001	0.476	0.770	0.000	0.007	0.857	0.804
n	21,983	46,830	34,350	13,726	20,050	15,844	6,279

* The variable Exit only has values for employees in 1981 to 1991 since 1992 is the final year of data.

Also, the social services bargaining unit was not organized in 1983. Prior to that year, they were represented by AFSCME. The AFSCME union was de-certified and the United Professional union was certified in 1984.

3.3 Definition of Variables

The definition of the variables used to model union membership and quits are listed below:

Member - Equals 1 if union member and zero otherwise

Exit - Equals 1 if worker exits the state labor force at time $t+1$ and zero otherwise.

New Entrant - Equals 1 if the employee has less than one year of tenure and zero otherwise.

Collective Bargaining - Equals 1 if the employee is covered by a collective bargaining agreement and zero otherwise.

dln(Minimum Wage) - Log change in minimum biweekly wage for the employee's job (five-digit job code) between time t and $t-1$. If Minimum Wage is missing the statewide average of ln(Minimum Wage) for the aggregate job classification is used.

dln(Wage) - Log change in the employees biweekly wage between time t and $t-1$. If the Wage is missing the statewide average of dln(Wage) for the aggregate job classification is used.

dln(CPS Wage) - Log change in the biweekly wage for the relevant job in the CPS data. If the CPS Wage is missing the statewide average of dln(CPS Wage)for the aggregate job classification is used.

dln(Wage) - dln(Min Wage) - dln(Wage) - dln(Minimum Wage)

dln(Min Wage) - dln(CPS Wage) - dln(Minimum Wage) - dln(CPS Wage)

dln(Relative Wage) - dln(Wage) - dln(CPS Wage)

ln(Relative Wage) - ln(wage) - ln(CPS Wage)

ln(Wage)$_{t-1}$ - Log of the biweekly wage at $t-1$. If the Wage$_{t-1}$ is missing the statewide average of ln(Wage)$_{t-1}$ for the aggregate job classification is used.

ln(CPS Wage)$_{t-1}$ - Log of biweekly wage for the relevant occupation in the CPS at $t-1$. If the CPS wage is missing the statewide average of ln(CPS Wage) for the aggregate job classification is used.

Pay Step - Equals the ratio of the Wage to the Minimum Wage. If either Wage or the Minimum Wage is missing, the statewide average of Pay Step for the aggregate job classification is used.

Dues - Annual dues contribution required for union membership, deflated by the Consumer Price Index (CPI-U).

Total Dues /1,000 - Total contribution, in thousands of dollars, of all others in the bargaining unit.

dln(FTE) -Log change in the number of FTE's in the aggregate occupation category.

Over time Indicator - Equals 1 if the worker had overtime hours and zero otherwise.

Prior Experience - Number of years between age 18 and date of employment with the state

State Tenure - Number of years of employment with the state

Part Time - Equals 1 if the employee is part-time or seasonal. Part time is defined to include permanent and exempt part-time, intermittent, seasonal and temporary unauthorized positions.

Non-White - Equals 1 if the employee is a minority and zero otherwise. Minority includes Black, Asian, American Indian, and Hispanic individuals.

Female - Equals 1 if the employee is a female and zero otherwise.

Married - Equals 1 if the employee is married and zero otherwise.

Manager - Equals 1 if the employee's aggregate job classification is a manager and zero otherwise.

Professional - Equals 1 if the employee's aggregate job classification is a professional occupation and zero otherwise.

Technical - Equals 1 if the employee's aggregate job classification is a technical occupation and zero otherwise.

Clerical - Equals 1 if the employee's aggregate job classification is a clerical occupation and zero otherwise.

Service/Blue Collar - Equals 1 if the employee's aggregate job classification is a service or blue collar occupation and zero otherwise.

D82 - Equals 1 if 1982 and zero otherwise.

D83 - Equals 1 if 1983 and zero otherwise.

D84 - Equals 1 if 1984 and zero otherwise.

D85 - Equals 1 if 1985 and zero otherwise.

D86 - Equals 1 if 1986 and zero otherwise.

D87 - Equals 1 if 1987 and zero otherwise.

D88 - Equals 1 if 1988 and zero otherwise.

D89 - Equals 1 if 1989 and zero otherwise.

D90 - Equals 1 if 1990 and zero otherwise.

D91 - Equals 1 if 1991 and zero otherwise.

D92 - Equals 1 if 1992 and zero otherwise.

Chapter 4 defines which variables are used to model union membership. Chapter 5 defines the subset that is used to model quits. In both instances, descriptive statistics are reported.

CHAPTER 4 - EMPIRICAL MODEL OF UNIION MEMBERSHIP

4.1 Introduction

Observing preferences for unions is complicated by the fact that the choice of union versus non-union status is often made simultaneously with the choice of occupation or job. However, workers covered by collective bargaining contracts in the public sector and in "Right to Work" states are usually not required to pay union dues. The payroll data for Iowa state employees provides an opportunity to observe workers' preferences for union services via voluntary dues contributions. Workers clearly have the opportunity to (and do) change union status at any time without affecting their current job status. Thus, we can treat the workers choice of occupation at time t as exogenous to the union membership decision at time $t+n$.

It is possible that an individual will join the public sector labor force because it is organized. However, once employed in the public sector, the individual can make the choice of whether or not to be a union member each subsequent period. The only cost in change from union to non-union status, other than the dues, is filling out a form to start or stop the withholding of dues from the individual's biweekly check. However, changing sectors of employment has much greater cost of search, lost seniority, and lost specific human capital. Thus, this temporal separation of occupational and union choices provides the necessary identifying assumption for the estimation of the demand for union services.

Union membership is examined conditional on the worker being covered by a collective bargaining contract. The process of allocating workers into covered and non-covered jobs is thus ignored. Later, the selection process into state employment will be explicitly modeled. This will be done by modeling the decision to remain in a public sector job or to exit for employment elsewhere.

4.2 Empirical Model

Assume that each agent is maximizing utility, V, at time t by choosing membership status, m, given a set x_t of exogenous and predetermined factors and a random disturbance $:_t$. Each agent is free to choose the membership status that maximizes his utility. $:_t$ is observed by the agent, but is not observable to the econometrician. Employment at time t is not conditional on membership in the union. However, membership could conceivably affect future employment and earnings. Thus, we can identify the probability that a given individual chooses to pay dues and become a union member in the following manner:

Let $V_{i1} = V(m=1,x_i,u_i)$ be the utility derived by the ith individual if he pays dues and becomes a union member. Let $V_{i0} = V(m=0,x_i,u_i)$ be the utility derived by individual i if he does not join the union. Then we can define the difference in utility between the two different states as

$$M^* = V(m=1,x_i,u_i) - V(m=0,x_i,u_i) = x_i 2 + :_i \qquad (4.1)$$

where 2 is a vector of associated parameters.

A rational agent will choose $m=1$ if and only if $M^* > 0$. Otherwise, the agent will choose $m=0$ and free ride on the union services provided. Given $:_i \sim N(0,1)$, the probability that individual i is a union member is

$$P(m=1) = \int_{-\infty}^{x\beta+\mu} \phi(t)dt = \Phi(x\beta+\mu) \qquad (4.2)$$

where ϕ is the normal density function and Φ is the normal distribution function. The parameters. β_j have a direct relationship to the change in the expected value of membership. The calculation of the marginal effect of variable x_j on the expected value of membership. P is

$$\frac{\partial P}{\partial x_j} = \phi(x\beta)\beta_j . \tag{4.3}$$

The elasticity is calculated as

$$e_j = \frac{\phi(x\beta)\beta_j x_j}{\Phi(x\beta)} \tag{4.4A}$$

for nontransformed variables. The elasticity for a variable in log form is

$$e_j = \frac{\phi(x\beta)\beta_j}{\Phi(x\beta)} . \tag{4.4B}$$

The union may provide many services, but two of the most important are to negotiate for wage increases and job security. We can measure unions' services in terms of wage and employment changes[9]. These changes may be common across all jobs, so that the union services are public goods and consumed by all workers covered by the agreement. Conversely, a portion of the union services can be private goods, so that the wage and employment changes benefit only a subset of the covered workers.

Now define the following components of x as

Δw_{ik} - Percentage wage increase due to collective bargain for individual i in reference group k.

[9] Other public goods such as grievance procedures, working conditions, seniority rules are also provided by the union. However, these have spillover effects to non-bargaining unit employees also. These components are not specifically modeled but should affect workers' marginal valuation of union services.

Δw_k - Average percentage wage increase due to collective bargaining for individuals in reference group k.

Δw_{ip} - Percentage increase in individual i's opportunity wage.

Δe_{ik} - Percentage change in employment for individual i in group k.

Δe_k - Percentage change in total employment in group k.

D_{ik} - Annual dues required for the ith individual to be a union member of group k.

D_{-ik} - Annual dues contributions of other individuals in reference group k.

The public gains from collective bargaining can be measured by Δw_k - Δw_p and Δe_k. The gain in pay relative to private sector pay yields distinct benefits to workers. The union has no direct effect on employment growth[10] but does provide benefits if jobs are threatened. Therefore, there will be a greater demand for union services when $\Delta e_k < 0$. The private gains from collective bargaining for individual i can be expressed as Δw_{ik} - Δw_k and Δe_{ik} - Δe_k. These factors measure changes in wages and employment for the kth job relative to state employees as a whole. D_{ik} is the amount of dues or the price the union charges that individual to be a member. Using a revealed preference argument, those who are dues paying members are "willing to pay" at least D_{ik} for the union services. They are "voting" for an expansion of union services and are getting at least D_{ik} in benefit from the increase in resources to the union.

Now define the remaining elements of x as x_0. Then $x\beta$ in equation (4.2) can be written as

[10] The unions bargain over wage levels and not employment levels.

$$x_i\beta = \beta_1(\Delta w_{ik} - \Delta w_k) + \beta_2(\Delta w_k - \Delta w_{ip}) + \beta_3(\Delta e_{ik} - \Delta e_k) + \beta_4 \Delta e_k \\ + \beta_5 D_{ik} + \gamma\beta_5 D_{-ik} + x_{0i}\beta_0 \tag{4.5}$$

The parameter values in this specification reflect the degree of publicness of union services and specifically whether the public good and/or private good aspects of union services are driving membership. The total instantaneous wage service effect is related to β_1 and β_2 and the total employment service effect is related to β_3 and β_4. β_1 and β_3 reflect the private good effects of union services on membership. Likewise, β_2 and β_4 reflect the public good effects of union services. If $\beta_1 = \beta_3 = 0$, $\beta_2 > 0$ and $\beta_4 \neq 0$ then union wage and employment services would be treated only as public goods. At the other extreme, if $\beta_1 = \beta_2$ and $\beta_3 = \beta_4$ then union service would be valued only as private goods. The more likely case is that the union's wage and employment services have some public and some private good aspects.

The effect of own contributions and others' contributions are modeled as a linear technology. β_5 is related to average willingness to pay at the margin. As union dues increase or the contributions of other increase, fewer individuals will choose to contribute to the union. The parameter γ reflects the substitutability of others' contributions for their own contributions. In the case of a purely private good, γ is zero. When union services are a pure public good γ is equal to one. The joint product model implies γ will lie between zero and one.

The parameter β_5 provides some information about how the union might be behaving. Revenue maximization would suggest that the observed equilibrium should be such that demand is unitary price elastic. This could arise if the union was providing a pure public good[11]. This implies that union dues, in equilibrium, would be set such that membership is unitary elastic. We would expect to see a profit maximizing union setting dues such that the equilibrium occurs on the elastic portion of the union demand curve. Given that only one union represents any particular group of workers, unions have a monopoly on the provision of services, at least in the short run.

Median voter models have been used to model union determination and provision of services. In cases where dues paying is required, it's reasonable that the dues rate might reflect the marginal valuation of the median member. However, when membership is not required, the situation is different. The median worker has the ability to free ride as long as some portions of the services provided are public goods. Thus, the median worker may vote for union representation but not for the expansion of union services. This of course assumes that the services the union is able to provide are a function of the collective resources contributed. If in fact the ability to bargain is affected by the membership rate, the median voter may want the union to maximize membership or at least go beyond the point where total dues contributions are maximized. At any rate, the model being developed assumes that it is resources rather than solidarity that determines the services that the union can provide.
To the extent that contributions of others are substitutes for their own contributions, the likelihood of becoming a dues paying member should decrease as the contributions of others increase. The pure public good scenario implies that raising the contribution of others would, ignoring budget effects, have the same impact on membership as an increase in the dues rate. As union services become more rival in consumption, workers will value the contribution of

[11] The marginal costs of allowing additional covered workers to join would be zero if the services were a pure public good.

others less. For the limiting case of a pure private good, the contributions of others would not affect the decision to pay union dues.

4.3 Data

Payroll data from the Iowa state government is used to identify the public and private effects of union services. This data contains annual observations on Iowa state employees from 1980 to 1992. Included in the data is information about the union status of an individual. Thus, we can determine if the individual's job is covered by a collective bargaining contract and if dues are being voluntarily withheld from individual's biweekly paycheck. Table 4.1 provides descriptive statistics and definitions for employees covered by collective bargaining during the sample period 1982 to 1992. The first two years are not used because of an inconsistency in how dues paying was coded in the data.

Information is also provided on the amount of dues that are withheld. Some of the unions, such as the one representing the troopers, had dues that were constant across all jobs for a given year. However, for the AFSCME union, each local was allowed to have its own dues structure. Some locals had constant dollar amounts for all members. Others had dues that are set proportional to salary. Each local had an opportunity to adjust its pricing structure on an annual basis. Many of the locals changed their dues structure frequently during the 11 year period. The only constraint on local dues setting is the amount that must be collected to cover dues for the national organization.

In addition to each worker's actual biweekly pay, the minimum and maximum salaries for the job title are included. Public good wage increases are identified using year to year changes in the log of the minimum wage for the job title. The wage increase for an individual

Table 4.1 - Covered Iowa State Employees: Variable Definitions and Descriptive Statistics, 1982-1992

Variable	Definition	Mean	S.D.
Member	1 if union member	0.352	0.477
dln(Minimum Wage)	Log change in minimum biweekly wage for the employee's job	0.046	0.041
dln(Wage)	Log change in the employee's biweekly wage	0.072	0.118
dln(CPS Wage)	Log change in the biweekly wage for the relevant job in the CPS data	0.042	0.044
dln(Wage) - dln(Min Wage)	dln(Wage) - dln(Minimum Wage)	0.026	0.113
dln(Min Wage) - dln(CPS Wage)	dln(Minimum Wage) - dln(CPS Wage)	0.003	0.060
dln(Relative Wage)	dln(Wage) - dln(CPS Wage)	0.622	0.037
ln(Relative Wage)	ln(wage) - ln(CPS Wage)	0.770	0.019
ln(Wage)t-1	Log of the biweekly wage at t-1	6.550	0.294
ln(CPS Wage)$_{t-1}$	Log of biweekly wage for the relevant occupation in the CPS at t-1	6.445	0.299
Pay Step	Ratio of the actual biweekly wage to the minimum biweekly wage	1.150	0.094
Dues	Dues contribution required for union membership	169.962	34.746
Total Dues /1,000	Total contribution, in thousands of dollars, of all others in the bargaining unit	146.780	94.657
dln(FTE)	Log change in the number of FTEs in the aggregate occupation category	-0.002	0.035
Over time Indicator	1 if the worker had overtime hours	0.462	0.499
Prior Experience	Number of years between age 18 and employment with the state	13.686	9.780
State Tenure	Number of years of employment with the state	10.112	7.865
Part Time	1 if the employee is part-time or seasonal	0.028	0.164
Non-White	1 if the employee is a minority	0.044	0.205
Female	1 if the employee is a female	0.476	0.499
Married	1 if the employee is married	0.649	0.477
Professional	1 if the employee has a professional occupation	0.166	0.372
Technical	1 if the employee has a technical occupation	0.152	0.359
Clerical	1 if the employee has a clerical occupation	0.263	0.440
Service/Blue Collar	1 if the employee has a service or blue collar occupation	0.419	0.493
D82	1 if 1982	0.075	0.264
D83	1 if 1983	0.063	0.244
D84	1 if 1984	0.076	0.265
D85	1 if 1985	0.096	0.295
D86	1 if 1986	0.092	0.289
D87	1 if 1987	0.096	0.295
D88	1 if 1988	0.097	0.295
D89	1 if 1989	0.104	0.305
D90	1 if 1990	0.105	0.306
D91	1 if 1991	0.099	0.299
D92	1 if 1992	0.096	0.294
n =148,009			

is measured by the year to year change in the log of the actual biweekly wage. This includes changes in the individual salary that occur because of promotions or other occupational changes. The difference between the actual wage increase and the increase in the minimum salary measures the private good wage increase. The increase to the salary scale and the total wage increase that an individual receives are not necessarily equal. As is typical in public sector wage schedules, there is a series of pay steps within each pay plan that the workers receive based how long they have been in a specific job. Workers wages may also be changing due being promoted to higher paying jobs.

The level and rate of change of wages outside the state government labor market is derived from wage information contained in the March Current Population Survey. The March Survey contains wage information for detailed occupations. The log change in CPS wage for the occupation closest to the defined job for a public sector employee is used as an estimate of the employee's opportunity wage increase in the private labor market.

There is evidence that the wage effects can be identified. A considerable amount of intertemporal variation exists in the union's ability to secure wage increases. The data includes years of zero, moderate and large overall wage increases. During years where wage freezes are imposed, any change in an individual's pay is a private good (the public wage increase is zero.) In addition, comparable worth adjustments allow cross sectional effects of public wage increases to be identified.[12] Comparable worth adjustment provided exogenous shocks to relative wages in two separate years (1985 and 1987). These changes in relative wages not only occurred across the broad occupational categories but also created substantial

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[12] Mattila, Orazem and Turk (1999) use the comparable worth wage shocks to estimate an input demand system for state government.

changes in relative wages within each occupational category. Some jobs received as many as seven pay grade increases as a result of comparable worth, while other jobs received no increases. In addition, private wage gains also varied substantially.

4.4 Model Estimates

Three different specifications of the model are estimated. The specifications differ in the ways that wage increases influence the perceived benefits from being a union member.

The first specification allows absolute wage changes to drive membership. That is, the log change in actual wage, minimum wage and the private sector wage each enter the membership function. This specification allows for changes in wage differentials that come from external (private sector) forces to have a different effect than internal (state) movement in wages.

The second specification is based on equation 4.6. The private gain is measured by the difference between the minimum wage gain and the actual wage gain for each individual. The public wage gain is the amount of salary increase in the salary scale over and above the private sector wage increase for the specific job. This specification isolates the public and private wage increases.

The third specification further constrains the wage effect such that only the change in the relative wage of the state versus the private sector is important in the membership decision.

The parameter estimates for each model are estimated via the method of maximum likelihood. Table 4.2 contains the parameter estimates for the three different specifications.

The calculated elasticities are reported in Table 4.3. The marginal effects and the elasticities are evaluated at the overall averages of each independent variable.

The estimates are quite similar in the three specifications. Similar results exist for the effects of public and private wage changes. While the estimated parameter *for dln(CPS Wage)* is not statistically significant, imposing the restrictions of the of the Public/Private Wage Gain Model is rejected (Likelihood ration = 547.) The positive coefficients on dln(Minimum Salary) and dln(Wage) suggest that, in the short run, both individual and overall wage gains increase membership. Both parameters are significantly different from zero and thus not consistent with a wage services being provided a purely public or purely private goods. Membership seems to be more responsive to wage services that are more individual specific. This is supported by the absolute wage gain specification and the public private wage gain specifications. Both specifications suggest a more elastic membership response to individual specific wage gains than to across the board wage changes. This suggests that an exclusion mechanism for wage services will increase membership. This result seems to run counter to the tendency for unions to reduce the variation in wages and other differences between occupations. The difference in the magnitude of the estimated parameters for dln(Wage) and dln(Minimum Salary) suggests that membership will be increased more with individual wage gains as opposed to across the board increases.

The wage growth (dln(CPS Wage)) in the private sector has a positive but insignificant influence on membership. This suggests that any differences in external wage across occupations are not important in the membership decision. The annual dummies are

Table 4.2 - Union Demand Model Estimates

Variable	Absolute Wage Gain Model		Public/Private Wage Gain Model		Relative Wage Gain Model	
	β	$\partial p/\partial x$	β	$\partial p/\partial x$	β	$\partial p/\partial x$
dln(Minimum Wage)	1.3960	0.5083				
	(0.1197)					
dln(Wage)	1.8232	0.6639				
	(0.0469)					
dln(CPS Wage)	0.0751*	0.0274				
	(0.0874)					
dln(Wage) - dln(Min Wage)			1.7402	0.6338		
			(0.0465)			
dln(Min Wage) - dln(CPS Wage)			1.0648	0.3878		
			(0.0736)			
dln(Relative Wage)					0.6219	0.2265
					(0.0369)	
ln(Relative Wage)					0.7701	0.2805
					(0.0187)	
ln(Wage)$_{t-1}$	1.9321	0.7036	1.9035	0.6932		
	(0.0308)		(0.0307)			
ln(CPS Wage)$_{t-1}$	-0.4906	-0.1786	-0.5186	-0.1889		
	(0.0196)		(0.0196)			
Pay Step	0.3283	0.1196	0.3167	0.1153	1.3604	0.4956
	(0.0596)		(0.0595)		(0.0548)	
Dues	-0.0046	-0.0017	-0.0043	-0.0016	-0.0030	-0.0011
	(0.0001)		(0.0001)		(0.0001)	
Total Dues/ 1000	-0.0011	-0.0004	-0.0011	-0.0004	-0.0015	-0.0006
	(0.0001)		(0.0001)		(0.0001)	
γ	0.0002		0.0003		0.0005	
dln(FTE)	0.1831*	0.0667	-0.1658*	-0.0604	-0.2066*	-0.0753
	(0.3050)		(0.3040)		(0.3020)	
Overtime Indicator	0.0857	0.0312	0.0823	0.0300	0.0713	0.0260
	(0.0085)		(0.0085)		(0.0084)	
Prior Experience	-0.0026	-0.0009	-0.0029	-0.0011	-0.0045	-0.0016
	(0.0004)		(0.0004)		(0.0004)	
State Tenure	-0.0103	-0.0038	-0.0104	-0.0038	-0.0087	-0.0032
	(0.0006)		(0.0006)		(0.0006)	
Log Likelihood	-85176.4		-85449.9		-86562.3	

* Indicates not significant at the .05 level. Standard errors are listed in parentheses. Marginal effects are computed using the overall average for each x.

Table 4.2(cont.)

Variable	Absolute Wage Gain Model		Public and Private Wage Gain Model		Relative Wage Gain Model	
	β	$\partial p/\partial x$	β	$\partial p/\partial x$	β	$\partial p/\partial x$
Part Time	-0.4823	-0.1756	-0.4892	-0.1782	-0.5755	-0.2096
	(0.0267)		(0.0267)		(0.0264)	
Non-White	0.1026	0.0373	0.0995	0.0362	0.0878	0.0320
	(0.0175)		(0.0174)		(0.0173)	
Female	0.1592	0.0580	0.1635	0.0595	0.0251	0.0091
	(0.0088)		(0.0088)		(0.0083)	
Married	-0.0327	-0.0119	-0.0315	-0.0115	-0.0158	-0.0058
	(0.0076)		(0.0076)		(0.0075)	
Technical	0.5309	0.1933	0.5157	0.1878	0.3261	0.1188
	(0.0180)		(0.0180)		(0.0175)	
Clerical	0.1778	0.0647	0.1468	0.0535	-0.2574	-0.0938
	(0.0162)		(0.0162)		(0.0137)	
Service/Blue Collar	1.2152	0.4425	1.1972	0.4360	0.7951	0.2897
	(0.0173)		(0.0173)		(0.0147)	
Intercept	-10.1604		-9.5470		-1.1825	
	(0.1922)		(0.1899)		(0.0653)	
D82	-0.0113*	-0.0041	-0.0738	-0.0269	-0.6610	-0.2408
	(0.0256)		(0.0254)		(0.0221)	
D83	0.1534	0.0559	-0.0289*	-0.0105	-0.5911	-0.2153
	(0.0270)		(0.0258)		(0.0217)	
D84	0.0041*	0.0015	-0.0935	-0.0340	-0.5932	-0.2161
	(0.0267)		(0.0263)		(0.0238)	
D85	-0.0343*	-0.0125	-0.0907	-0.0330	-0.5145	-0.1874
	(0.0226)		(0.0224)		(0.0204)	
D86	0.0596	0.0217	-0.1398	-0.0509	-0.5081	-0.1851
	(0.0212)		(0.0194)		(0.0171)	
D87	-0.0319*	-0.0116	-0.0717	-0.0261	-0.3879	-0.1413
	(0.0277)		(0.0277)		(0.0267)	
D88	0.0574	0.0209	-0.0281*	-0.0102	-0.2735	-0.0996
	(0.0218)		(0.0214)		(0.0206)	
D89	-0.0357*	-0.0130	-0.1066	-0.0388	-0.2913	-0.1061
	(0.0307)		(0.0304)		(0.0298)	
D90	-0.0526	-0.0192	-0.1287	-0.0469	-0.2319	-0.0845
	(0.0201)		(0.0198)		(0.0195)	
D91	0.1494	0.0544	-0.0536	-0.0195	-0.0982	-0.0358
	(0.0196)		(0.0175)		(0.0168)	

Table 4.3 - Calculated Elasticities

	Absolute Wage Gain Model	Public and Private Wage Gain Model	Relative Wage Gain Model
dln(Minimum Wage)	1.52		
dln(Wage)	1.98		
dln(CPS Wage)	0.08		
dln(Wage) - dln(Min Wage)		1.89	
dln(Min Wage) - dln(CPS Wage)		1.16	
dln(Relative Wage)			0.68
ln(Relative Wage)			0.84
$\ln(\text{Wage})_{t-1}$	2.10	2.07	
$\ln(\text{CPS Wage})_{t-1}$	-0.53	-0.56	
Pay Step	0.41	0.40	1.70
Dues	-0.84	-0.80	-0.56
Total Dues x 1000	-0.17	-0.18	-0.24
dln(FTE)	0.20	-0.18	-0.22

capturing the general rate of wage inflation in the private sector. Thus, the absolute wage gain model may be over-parameterized.

dln(FTE) has an insignificant influence on union membership and the sign changes depending on the specification. Thus, we can't reject the hypothesis that changes in aggregate employment levels for the workers' occupational classification have no effect on union membership. This suggests that differences in private job security across job types are not important. That is not to say that overall job security is not a factor in the union membership decision. Because I have included annual dummy variables in each of the models, I have controlled for overall changes in state employment. It may well be that *dln(FTE)* does not accurately capture private job security effects or is dominated by statewide public job security effects.

Wage levels have a strong positive influence on membership, which is consistent across the all the models estimated. $\ln(\text{Actual Salary})_{t-1}$ is positively related to union membership. The absolute wage model and the public/private wage gain models both have an estimated elasticity of greater than 2. In all specifications, membership is positively related to relative wages. This relationship appears to be elastic in the absolute wage gain and the public/private wage gain models. Membership response is positive but inelastic in the relative wage gain model. The elasticity associated with *dln(Relative Wage)* is approximately equal to the difference between the public and private good wage elasticities.

The opportunity wage for the jobs outside state government ($ln(CPS\ Wage)_{t-1}$) has a negative influence on membership. The higher the opportunity wage the more likely the individual is to be a free-rider. This is consistent with theory. If a worker feels he is underpaid relative to what he could make outside state government, he will be less likely to be a union member. Workers in jobs that enjoy a smaller wage differential over private sector counterparts may not have benefited as much from past union bargaining and may be pessimistic about the ability of the union to bargain for wage increases. More importantly, they would be more inclined to exit state government employment. All these factors should contribute to this result.

Pay Step measures an individual's current pay relative to the minimum salary for their respective job. The estimated parameter associated with Pay Step indicates that those workers at the upper end of the wage distributions for their individual job are more likely to be union members. These individuals are more likely to have exhausted step increases. Thus, the only mechanism for wage increases would be from general pay increases that could come from union bargaining. More recently promoted workers tend to be at the lower pay steps. When workers advance to a job with a higher pay grade, they usually enter that grade at a lower step. However, this effect may be exacerbated by the higher turnover rates for less senior workers.

The level of dues or cost of membership has a negative impact on membership in all cases.[13] The price effect in Public/Private Wage Gain and Absolute Wage Gain specification are almost identical. The estimates suggest that a one-dollar increase in dues rate will cause a likelihood of membership to decrease by about .0016. The estimated coefficient for Dues in the Relative Wage Gain Model is similar to the other two specifications. In this case, an increase in dues would correspond to a .0011 decrease in likelihood of membership. This suggests that the estimated effect of dues is robust across the three specifications.

The calculated price elasticities (Table 4.3) indicate that overall union membership is price inelastic. Revenue maximization would suggest that, in equilibrium, demand would be unitary elastic. Based on this, the union may also be considering membership rate when setting dues. An alternative scenario might be where there is an uncoordinated supply of union services and myopic locals set dues given the demand and cost structure facing them. However, this is not consistent with the short run monopoly power unions have in providing collective bargain services.

The estimated γ is simply the ratio of the estimated parameter for *Total dues/1000* and *Dues*. This value is negative and statistically different from zero. Recall in the joint product model developed in Chapter 2 that their own contributions as well as the contributions of others enter the utility function. Thus, γ measures the degree to which contributions of others

[13] Models with interaction terms between the occupational dummies and the dues rate were also estimated. When the price response is allowed to vary across aggregate job types, only technical occupations had a positive coefficient. The coefficients for the other three occupation categories are all negative and close in magnitude.

can substitute for own contributions. In the case of a pure public good they would be perfect substitutes. The estimates of γ reported in Table 4.2 support the joint product model. The share parameter, γ, ranges from .0002 to .0005. The null hypothesis that $\gamma = 0$ is equivalent to testing that the coefficient for *Total dues/1000* is equal to zero. The calculated Wald statistic is 372 in the Absolute Wage Gain Model, 390 for the public private wages gain model and 720 for the Relative Wage Gain model estimates. These values are significant at any reasonable level. Whiles the values for γ are statistically significant, they are fairly small. This suggest that union membership is predominately a private good. If this is in fact the case, union security clauses aren't necessary to deal with free-riders.

While these estimates may be the first to quantify the relative importance of public and private aspects of union membership, this is not the first attempt to model the demand for public good with an impure public good. Many researchers have attempted to quantify the publicness of services provided by local governments[14]. In most cases, the research models demand in the context of the median voter. As such, the median voter's demand is modeled as a function of the per unit price of the public good, tN^{γ}. In this model t is the tax rate, N is the population size and γ is the congestion parameter. In this case, γ equal 1 implies that it is a public good and γ equal to zero implies that is a private good. In many cases, the empirical estimates derived from data have yielded estimates of congestion parameters that are greater than one and thus not consistent with the model.

One issue that has been thus far ignored is the influence of selection bias. Individuals are self-selecting in and out of state employment. It is quite likely that those individuals that

[14] See for example Gramlich and Rubinfeld (1982), Edwards (1986), Bercherding and Deacon (1972), and Bergstrom and Goodman (1973)

are most likely to leave state employment would also be less likely to be union members. They may be less likely to be union members not because they have less of a demand for union services, but because they have less of a chance of remaining in their current position. This would tend to bias downward the estimates of the marginal willingness to pay for union services. This issue will be addressed in the next chapter.

State tenure is negatively related to union membership in each of the specifications. Tenure before entering state employment (Prior Experience) is also negatively related to union membership. This could be due to an overall age effect. Farber and Saks (1980) found older workers were less likely to vote for union representation. Presumably, older workers have fewer years remaining to benefit from investing in the union. However, a likelihood ratio test with the null hypothesis that the coefficients on State Tenure and Prior Experience are equal is rejected. This suggests that tenure with state government has a different effect on union membership. Individuals' prior experience may come from all of the three following source: years spent in formal education, years spent in the labor force outside of state government, and time spent out of the labor force. This confounds the measurement slightly. Individuals with more tenure prior to entering state government would be older than others in their cohort and are likely to have more non-specific experience. Tenure with state employment is a cleaner measure and thus has a more straightforward interpretation. The magnitude of *State Tenure* is greater (more negative) than that of *Prior Experience* suggesting that firm specific experience has a larger negative impact on union demand.

The negative relationship seems at first surprising given that new and younger workers would receive fewer benefits from the union. Most of the exits from the public sector come from workers with very little tenure. Therefore, the presumption would be that the group of low-tenure workers would disproportionately include workers with little future interest in public sector pay growth and working conditions. Thus, selection bias would tend to bias the coefficient on *Prior Experience* and *State Tenure* toward the positive.

However, it is possible that as tenure increases incentives to invest in union membership decline. Tenure plays an important role in the services that the unions provide. As a worker gains state-specific experience, seniority rules tend to insulate the workers from the adverse effects of potential employment shocks. These rules apply to all covered workers regardless of membership, and so free riders cannot be excluded. These services would be viewed as more public, the more seniority the worker has.

Full time, non-white, female and single employees are each more likely to be union members. Previous research has generated mixed results of the impact of gender on the demand for union services. Some studies have argued that females tend to be less likely to be union members. Many have argued that it is due to less attachment to the labor force. Farber and Saks (1980) found positive but insignificant effects of gender. Chaison and Dhavale (1992) found that females were less likely to be union members. Freeman and Medoff (1984) find that while women are less likely to be union members, they are more likely to vote for unionization in elections. I would argue that limited wage information in other studies has precluded researchers from disentangling wage effects from other effects and this has led to mixed results. The results here suggest that females have higher marginal valuation of union membership and that wages are driving the empirical result that on average, women have lower rates of membership. Granted, one cannot make inferences about unions as a whole based on data in Iowa. However, the large relative wage changes induced by comparable worth should allow us to unravel wage effects from other fixed effects.

The occupational dummy variables indicate that Service/Blue Collar workers have the highest marginal valuation of union services and Professional employees have the lowest. Technical workers have greater interest in membership than do Clerical. Other studies suggest that unions tend to benefit low wage and lower skilled workers the most and thus narrow the size of the skilled/unskilled wage differentials. For instance, Farber and Saks (1980) found that workers on the lower end of the wage distribution were more likely to vote for union representation. The service/blue collar jobs in general tend to have higher union rates in the private market. Professional type jobs are more likely to have professional associations and not have collective bargaining. Clerical workers tend to have very low rates of unionization. They also tend to be at the lower end of the wage distribution, but this is controlled for in the model.

Non-white employees are significantly more likely to pay union dues. Presumably, minority workers place a higher value on union services, because they feel that collective bargaining and grievance procedures will help to protect them from race discrimination. This result is consistent with the results of numerous other studies[15].

The annual dummies are included to control for overall employment effects and any other macro economic effects. The dummies reflect the fact that there was a dramatic upswing in membership in the last half of the eighties and the beginning of the nineties. Part of this could be due to an unmeasured effect of comparable worth. Part of this could also be

[15] See for example, Farber and Saks (1980), Farber (1983), Chaison and Dhavale (1992), Davis and Huston (1995), and Sobel (1995).

due to a standoff between unions and the governor over wage increases that the unions had negotiated but were not given because of unexpected budget shortfalls. The union ultimately won the wage increases; however, this induced layoffs. Even though the total reduction in employment was small, a large number of workers were shifted to different jobs due to seniority rules and bumping rights. One would expect that this would increase an individual's desire for union representation, because it is more likely that the worker would be using grievance procedures.

4.5 Conclusions

Wages play an important role in determining union membership, especially the ability of unions to secure wage increases and historical wage differentials between the public and private sectors. Current wage gains by state employees and historical wage differentials have a strong positive influence on the perceived benefit from union membership. A wage increase that is less widely enjoyed elicits a stronger membership response.

Membership also responds significantly to the price of membership or the dues. The elasticity suggests that the union is pricing on the inelastic portion of the demand curve. It would seem that union revenue could be increased by increasing the dues rate. However, I only observe membership conditional on the choice to remain in state government. Factors that effect the marginal evaluation of union service also affect tenure. Thus, there is the potential for selection bias to influence the results. An empirical model of quits is developed in the next chapter. This model is used to create statistical controls for the potential selection bias. Once the selection bias has been accounted for, the model estimates will be examined. I will then quantify the impact of comparable worth on our ability to identify the model parameter. In addition, the model will be used to explore the impact of comparable on both quits and on union membership.

The contributions of others also have a negative effect on membership, suggesting a joint production of a public and private good. The magnitude of the congestion parameter suggests that union services are for the most part a private good.

CHAPTER 5 - PUBLIC SECTOR QUITS

5.1 Introduction

Both employer and employees bear costs when employment relationships are severed. Both invest in the creation of the job match through search costs and investment in firm specific human capital. These investments are foregone when the worker quits or is laid off. Asymmetric information or other information imperfections may be partly to blame. Some information about job characteristics and potential employees is not revealed fully to all parties until after employment. Once worker productivity becomes known, the employer and/or worker may find that their welfare can be improved if they terminate the relationship. Alternatively, new information may reveal that the worker is more productive than anticipated. In that case, employers may be able to reduce turnover by offering higher wages to reduce quits. Higher wages may also allow them to be more selective in the hiring process.

Labor market tenure is determined by many factors such as opportunity wages, job specific human capital and other non-wage benefits, search costs as well as other individual characteristics. Many researchers have explored the relationship between wages and tenure. Topel (1991) provides convincing evidence that wages rise with seniority. Others have taken wages as exogenous and focused on exit propensities and the influence of wages. Koch and Ragan (1986) argue that in unionized and public sectors, it is reasonable to assume wages cause quits, and not the reverse. Wage scales in the public sector are typically very structured and influenced by collective bargaining. Changes tend to be dominated by across the board increases in pay scales. Thus, at the micro level, it is plausible that workers take state wages as exogenous. This is the strategy that I adopt here.

Much of the research on quits has utilized the National Longitudinal Survey. However, the public sector is another area where turnover can be studied with micro data. States and the federal governments are usually very large employers and micro level employment data is more readily available. Unfortunately, relative wages in the public sector are very stable and typically aren't as responsive to demand and supply shocks as are wages in the private sector. However, the implementation of comparable worth pay plans in Iowa state government provides an exogenous shock to relative wages that can be used to help identify the effects of public sector wages on quit propensity.

This chapter develops an empirical model of quits for state employees in Iowa using payroll data for state employees from 1980-1992. The purpose is two fold. One is to examine the role of tenure and wages in turnover. The second is to develop statistical controls for selection bias for the union membership model developed previously. Section 5.2 develops the empirical model. Section 5.3 provides a descriptive analysis of the payroll data. Section 5.4 reports the parameter estimates for three specifications. Section 5.5 explores the impact of comparable worth on parameter identification. Section 5.6 employs the empirical model to estimate the impact of comparable worth quits.

5.2 Empirical Model of Quits

Each worker chooses at time period t to remain employed in the public sector or to exit to the private sector. Employees compare the net benefit between the two employment opportunities based on a vector of exogenous factors, x_{it-1} and a stochastic shock ε_{it}. Let $e_t=1$ represent the choice of remaining in state government employment at time t and let $e_t=0$

represent exiting state government for employment elsewhere. An individual will choose to remain in the public sector labor market if

$$E_t^* = U(e_t=1, x_{it-1}, \varepsilon_{it}) - U(e_t=0, x_{it-1}, \varepsilon_{it}) > 0. \tag{5.1}$$

Unfortunately, the value of E_t* is not revealed to us. The sign of E_t* is revealed by observing if the individual is in the public sector labor market at time t. Assume that the net benefit from changing jobs can be represented as

$$E_t^* = x_{it-1}\alpha + \varepsilon_{it}. \tag{5.2}$$

Assuming that ε_{it} is distributed normally, the decision to remain employed with the state government can then be represented as

$$P(E_t^* > 0) = \int_{-\infty}^{x_{t-1}\alpha} \phi(z)dz = \Phi(x_{t-1}\alpha) \tag{5.3}$$

where ϕ is the normal density function and Φ is the normal distribution function. The marginal effect of an individual variable[16] x_j on the probability of an exit, P, is

$$\frac{\partial P}{\partial x_j} = \phi(x\alpha)\alpha_j. \tag{5.4A}$$

However, when x_j also has a quadratic term, the marginal effect is

$$\frac{\partial P}{\partial x_j} = \phi(x\alpha)(\alpha_j + 2\alpha_{j+1}x_j) \tag{5.4B}$$

when $x_{j+1} = x_j^2$.

The elasticity is calculated as

$$e_j = \frac{\phi(x\alpha)\alpha_j x_j}{\Phi(x\alpha)} \tag{5.5A}$$

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[16] The time subscript is dropped here for notional ease

for nontransformed variables. The elasticity for variables in log form is

$$e_j = \frac{\phi(x\alpha)\alpha_j}{\Phi(x\alpha)}. \qquad (5.5B)$$

5.3 Data

Payroll data from the Iowa state government is used to develop the model estimates. The data are observations of employment in the last payroll period for each calendar year from 1980 to 1992. The data is essentially the same as that used to model union membership in the previous section. However, when modeling quits, the data is not limited to employees covered by collective bargaining. Because this data has employee information that spans 13 years, it is especially appealing for modeling quits. Over 13 years there will potentially be considerable variation in relative pay between the public and the private sectors.

Table 5.1 lists descriptive statistics for the variables used to model the propensity to quit. One reason for modeling quits is to develop statistical controls for selection bias that may exist in the union membership model estimates developed in chapter 4. Management is not covered by collective bargaining and therefore does not really have the opportunity to join the union.[17] In developing the selection bias controls, it is desirable to have a model that is specific to the population. The estimates change little since management makes up such a small proportion of employees. Thus, I don't exclude management employees when estimating the exit models used to develop selection bias controls for the union membership model.

An average of eight percent of workers exit from the state labor force each year. Most of the exits observed are permanent. Only about six percent of the workers observed

[17] A few cases exist where managers will pay union dues. However, they are not covered by the collective bargaining agreement and their contributions are most likely motivated by things other than the union services provided.

Table 5.1 – Iowa State Government Employees, Variable Definitions, Means and Standard Deviations, 1981-1991

Variable	Definition	Mean	SD
Exit	1 if workers exits the state labor force, zero otherwise	0.082	0.275
dln(Wage)	Log change in the employees biweekly wage	0.071	0.123
dln(CPS Wage)	Log change in the biweekly wage for the relevant job in the CPS data	0.048	0.056
dln(Wage) - dln(CPS Wage)	Log change biweekly wage minus the log change in the biweekly wage for the relevant job in the CPS data	0.024	0.135
ln(Wage)$_{t-1}$	Log of the biweekly wage at t-1	6.573	0.377
Prior Experience	Number of years between age 18 and employment with the state	13.627	9.855
State Tenure	Number of years of employment with the state	10.385	8.266
State Tenure2**	The square of the number of years of employment with the state	176.175	261.562
New Entrant	1 if the employee has less than one year of tenure	0.085	0.279
Collective Bargaining	1 if the employee is covered by a collective bargaining agreement	0.668	0.471
Part Time	1 if the employee is part-time or seasonal	0.036	0.186
Non-White	1 if the employee is a minority	0.039	0.194
Female	1 if the employee is a female	0.482	0.500
Married	1 if the employee is married	0.665	0.472
Manager	1 if the employee is a manager	0.047	0.213
Professional	1 if the employee has a professional occupation	0.245	0.430
Technical	1 if the employee has a technical occupation	0.122	0.328
Clerical	1 if the employee has a clerical occupation	0.266	0.442
Service/Blue Collar	1 if the employee has a service or blue collar occupation	0.319	0.466
D82	1 if 1982	0.089	0.284
D83	1 if 1983	0.089	0.285
D84	1 if 1984	0.090	0.287
D85	1 if 1985	0.091	0.288
D86	1 if 1986	0.087	0.282
D87	1 if 1987	0.090	0.287
D88	1 if 1988	0.091	0.288
D89	1 if 1989	0.096	0.294
D90	1 if 1990	0.097	0.295
D91	1 if 1991	0.091	0.288
n=217,056			

leaving the state government return before the end of the sample period. Almost 24% of the state labor force for the sample period is comprised of workers with three or less years of state tenure. Around 11% have more than 25 years with the state. Average tenure is slightly more than ten years. Figure 5.1 displays the distribution of tenure in the Iowa state government. The figure suggests that a substantial proportion of exits come from workers

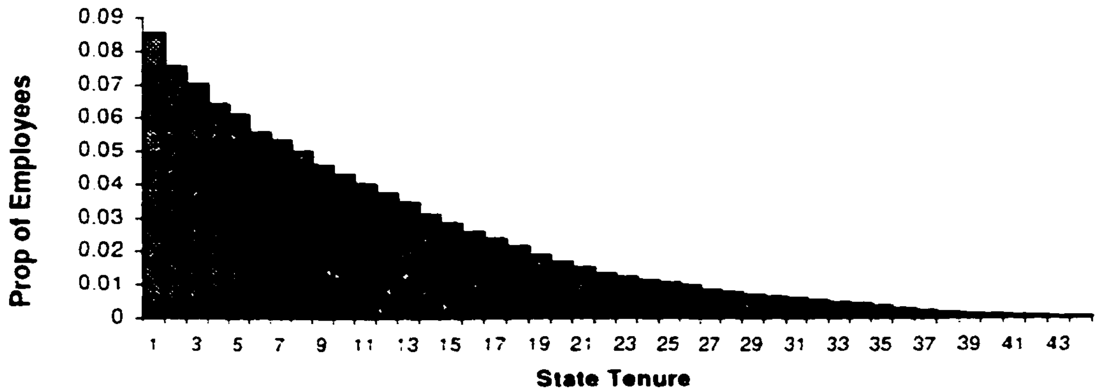

Figure 5.1 – Proportion of Employees by Years of State Tenure, 1981-1991

with very little tenure. Figure 5.2 reports the exit rates during the sample period by the years of state tenure when the exit occurs. The graph plots the actual quit or exit rate versus tenure. The data confirms the notion that relatively new workers have a higher tendency to quit. Calculated exit rates are high in the early years and then decline rapidly. Eventually the incidence of exits increases as workers approach retirement. The high rate of turnover in the early years of employment has been found in other empirical studies. This fact has been

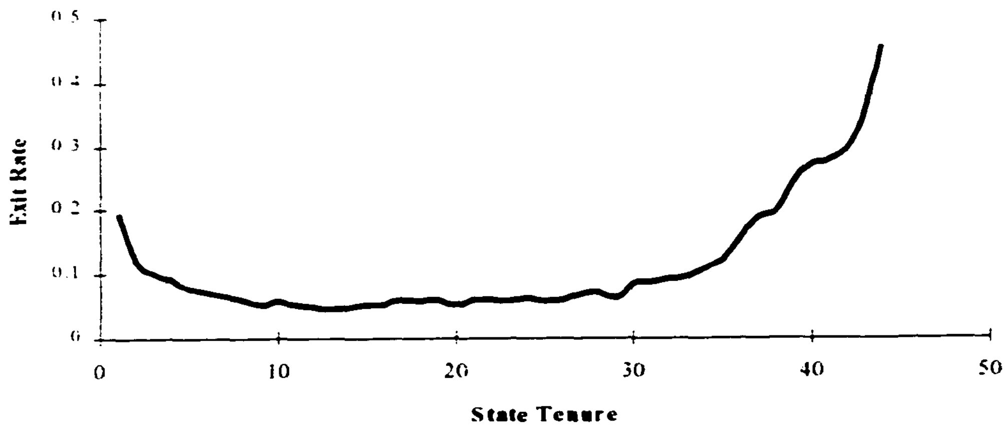

Figure 5.2 – Exit Rates by Years of State Tenure, 1981-1991

stylized into models of "matching." These models assert that asymmetric information exists between employee and employer. After a particular employment relationship is established, information is revealed to both employer and employee regarding the "match" between the skills needed in the occupations and the skills embodied in the individual. Good matches tend to survive. At any rate, it appears that a quadratic relationship may exist between state tenure and propensity to quit.

Figure 5.3 graphs the exit rates for each year in the sample. The year is defined as the year the employee is last employed. The exit rates decline noticeably after 1985. 1985 is particularly of interest, because it is when the comparable worth play plans were first implemented. A second implementation occurred in 1987. Without considering any other factors, it appears that exit rates declined after the comparable worth pay plans were implemented. Thus the increases in wage differentials, or at least the notion of comparable worth, may have decreased incentives to leave the state labor market.

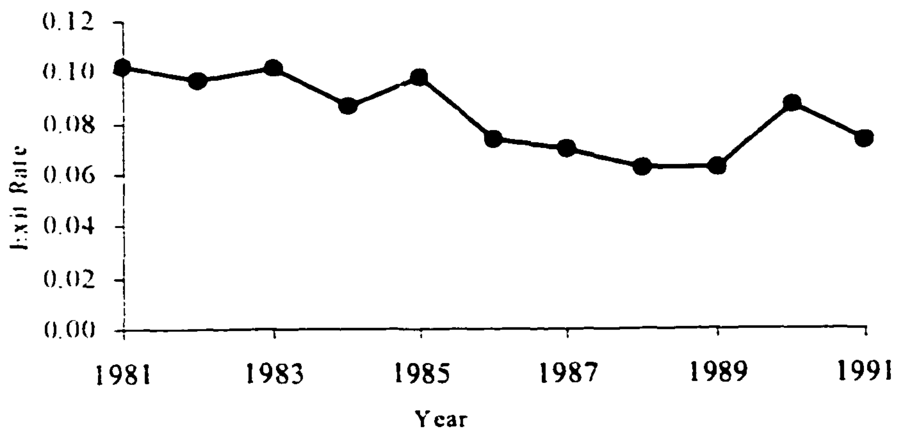

Figure 5.3 - Exit Rate by Year, 1981-1991

As demonstrated previously, the implementation of comparable worth plans induced changes in relative wages that are rare in public sector labor markets[18]. In the case at hand, intertemporal changes in an individual's wages and in the wage scales for each job classification are observed throughout the sample period. Two separate quasi-exogenous shocks to relative wages occurred as a result of the implementation of comparable worth pay plans. This phenomenon not only provides changes in relative wages between aggregate occupations, but also generated changes in relative wages within each occupation. Figure 5.4 shows the average changes in starting pay for the five aggregate job types. The implementation of comparable worth in 1985 and 1987 clearly created inter-job variability in wage gains. Furthermore, these changes are not homogenous within aggregate job types. Figure 5.5 shows within each aggregate occupational classification the standard deviation of

[18] Kim (1989) found that, in the California state government, relative pay in the 1980's was almost perfectly explained by relative pay in the 1930's.

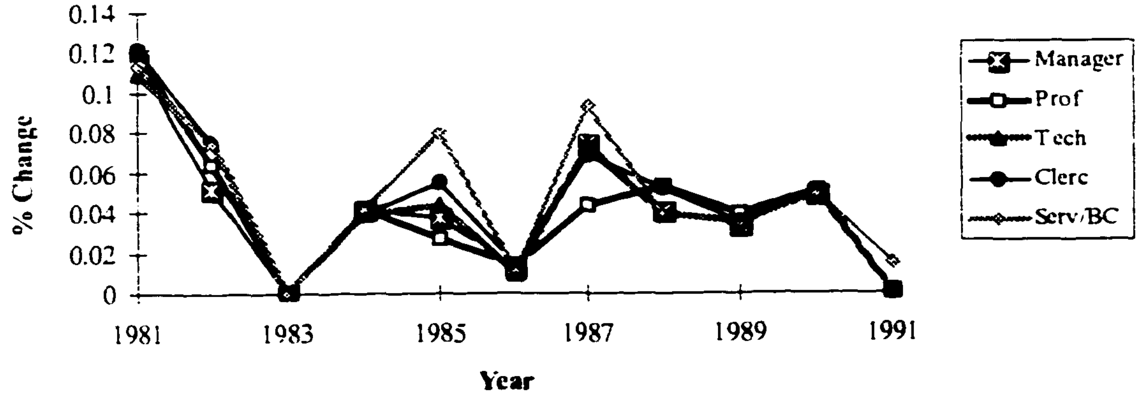

Figure 5.4 – Average Percentage Change in Starting Pay, 1981-1991

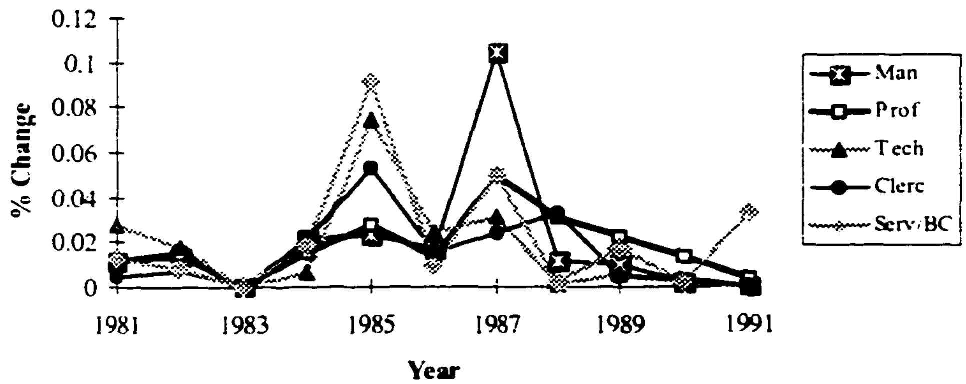

Figure 5.5 - Individual Standard Deviations of Starting Pay, 1981-1991

the percentage changes in starting pay. These shocks help identify the effects of wages on quits.

5.4 Empirical Exit Model Estimates

Three specifications were estimated, each differing in the assumed effect of contemporaneous wage gains on quits. The first specification allows internal and external wage gains to have asymmetric effects on the quit propensity. In this case, the current wage received in the state government and the opportunity wage outside of state employment enter separately into the model. The second specification assumes that workers condition on the relative wage between the public and private sector. The third specification does not allow changes in external wages, other than at a macro level, to affect quit propensities.

Table 5.1 lists the variables used to model quits. $\ln(\text{Wage})_{t-1}$ measures the individuals' relative position on the intra-firm wage distribution. Also included are measures to control for wage levels, tenure, and non-specific experience. Also include are demographic variables to capture effects of gender, marital status, minority, broad occupational categories and part time employees. Similar variables have been included in many other studies. Thus, comparisons can be made to the previous research results.

The models are estimated using annual dummy variables to capture time specific macro effects. Another alternative would be to include macro variables such as overall price indices, unemployment rates and other cyclical factors. I choose the dummy variable approach since many of the macro factors, such as grievance activities and management/employee relations could not be separately controlled for. Once one includes the dummy variables, time specific macro variables are redundant. In addition, transforming any of the independent variables by a time specific deflator will have absolutely no effect on any of the coefficients estimated, expect for the time specific constants.

The model estimates are reported in Table 5.2 and Table 5.3. Table 5.3 contains the annual dummy estimates and Table 5.2 contains continuous and categorical variable estimates. Included in the table are the parameter estimate and the estimated marginal effect per equations 5.4A and 5.4B.

Table 5.2 – Exit Model Full Sample Estimates

	Internal/External Wage Model		Relative Wage Model		Internal Wage Model	
Variable	α	$\partial p/\partial x$	α	$\partial p/\partial x$	α	$\partial p/\partial x$
dln(Wage)	-0.3902	-0.0445			-0.3716	-0.0424
	(0.0386)				(0.0383)	
dln(CPS Wage)	0.2008	0.0114				
	(0.0758)					
dln(Wage) - dln(CPS Wage)			-0.3533	-0.0403		
			(0.0348)			
ln(Wage)$_{t-1}$	-0.2798	-0.0319	-0.2663	-0.0304	-0.2697	-0.0308
	(0.0243)		(0.0236)		(0.0242)	
Prior Experience	0.0115	0.0013	0.0115	0.0013	0.0114	0.0013
	(0.0004)		(0.0004)		(0.0004)	
State Tenure[b]	-0.0527	-0.0017	-0.0528	-0.0018	-0.0528	-0.0018
	(0.0017)		(0.0017)		(0.0017)	
State Tenure2**	0.0018		0.0018		0.0018	
	(0.0000)		(0.0000)		(0.0000)	
New Entrant	0.2106	0.0240	0.2116	0.0241	0.2115	0.0241
	(0.0143)		(0.0143)		(0.0143)	
Collective Bargaining	-0.0872	-0.0100	-0.0840	-0.0096	-0.0881	-0.0101
	(0.0104)		(0.0103)		(0.0104)	
Part Time	0.5586	0.0638	0.5625	0.0642	0.5613	0.0641
	(0.0176)		(0.0175)		(0.0176)	
Non-White	0.1792	0.0205	0.1791	0.0204	0.1783	0.0204
	(0.0187)		(0.0187)		(0.0187)	
Female	0.0627	0.0072	0.0637	0.0073	0.0640	0.0073
	(0.0096)		(0.0096)		(0.0096)	
Married	-0.1002	-0.0114	-0.1005	-0.0115	-0.1002	-0.0114
	(0.0086)		(0.0086)		(0.0086)	
Manager	-0.0614	-0.0070	-0.0642	-0.0073	-0.0627	-0.0072
	(0.0225)		(0.0225)		(0.0225)	
Technical	-0.2043	-0.0233	-0.2012	-0.0230	-0.2020	-0.0231
	(0.0168)		(0.0167)		(0.0167)	
Clerical	-0.2334	-0.0266	-0.2274	-0.0259	-0.2302	-0.0263
	(0.0155)		(0.0153)		(0.0155)	
Service/Blue Collar	-0.1023	-0.0117	-0.0958	-0.0109	-0.1010	-0.0115
	(0.0147)		(0.0144)		(0.0147)	
Intercept	0.6937	0.0792	0.5936	0.0677	0.6363	0.0726
	(0.1613)		(0.1551)		(0.1606)	
Log Likelihood	-58,023.4		-58,026.0		-58,053.9	

* Indicates not significant at the .05 level. Standard Errors are listed in parentheses. Marginal effects are computed using the overall average for each x. [b] Marginal effect includes the impact of the quadratic term.

Table 5.3 – Exit Model Full Sample Estimates – Time Dummy Variables

	Internal/External Wage Model		Relative Wage Model		Internal Wage Model	
Variable	a	$\partial p/\partial x$	a	$\partial p/\partial x$	a	$\partial p/\partial x$
D82	-0.0136*	-0.0015	-0.0128*	-0.0015	-0.0154*	-0.0018
	(0.0181)		(0.0181)		(0.0180)	
D83	-0.0066*	-0.0008	-0.0137*	-0.0016	0.0048*	0.0005
	(0.0186)		(0.0183)		(0.0181)	
D84	-0.0780	-0.0089	-0.0824	-0.0094	-0.0736	-0.0084
	(0.0186)		(0.0184)		(0.0185)	
D85	0.0500	0.0057	0.0476	0.0054	0.0482	0.0055
	(0.0185)		(0.0185)		(0.0185)	
D86	-0.0927	-0.0106	-0.0932	-0.0106	-0.0966	-0.0110
	(0.0199)		(0.0199)		(0.0199)	
D87	-0.1084	-0.0124	-0.1119	-0.0128	-0.1133	-0.0129
	(0.0206)		(0.0205)		(0.0206)	
D88	-0.1533	-0.0175	-0.1585	-0.0181	-0.1557	-0.0178
	(0.0214)		(0.0213)		(0.0214)	
D89	-0.1633	-0.0186	-0.1717	-0.0196	-0.1627	-0.0186
	(0.0218)		(0.0215)		(0.0218)	
D90	0.0472	0.0054	0.0406	0.0046	0.0432	0.0049
	(0.0216)		(0.0214)		(0.0215)	
D91	-0.0245*	-0.0028	-0.0274*	-0.0031	-0.0296*	-0.0034
	(0.0224)		(0.0224)		(0.0224)	

* Indicates not significant at the .05 level. Standard Errors are listed in parentheses. Marginal effects are computed using the overall average for each x.

The estimates indicate an inelastic relationship between quits and wages. The estimates for dln(wage) suggest that a one-percent increase in state wages induces an instantaneous .0004 percentage point reduction in quits. This estimate is common across all three specifications used. The corresponding elasticity of quits with respect to public sector wages is approximately -0.8. The permanent effect of wages, as measured by $ln(wage)_{t-1}$, is somewhat smaller but similar in magnitude to the instantaneous effect. When job-specific external wages, $dln(CPS\ Wage)_{t-1}$, are allowed to enter the model explicitly, the effect of external wage changes is less pronounced than the effect of state wages. Theory would suggest that occupational choice would be conditioned on relative wages. However, based on the likelihood ratio test, one must reject the hypothesis that these two coefficients have equal but opposite values. It is conceivable that quits are less sensitive to external wages due to the non-wage cost of changing employers. Thus, there may be compensating differentials that cause the public sector employment to be relatively less responsive to external wages. The difference could also be due to error in the measurement of the true opportunity wage.

Theory would also suggest that job-specific human capital would diminish the incentives to quit and that non-specific human capital would tend to make workers more mobile and hence more likely to change jobs. The model estimated here is consistent with those presumptions. Prior Experience, the measure for non-specific human capital, tends to increase the incidence of quits, albeit with a very small elasticity. Each additional year of experience prior to state employment increases the probability of a quit by .0013. The estimated elasticity is approximately .31. Conversely state specific human capital, as measured by state tenure, decreases the quit rate. State tenure may have a slightly larger impact than prior experience. An additional year of state experience decreases the likelihood of a worker quitting by .0017, which corresponds to an elasticity of about -.32. In addition, workers with no state specific experience have an exit rate 2.4 percentage points higher than those with one year of tenure. The estimates for job tenure are consistent with Light and Ureta (1992). However, the impact of prior experience is negative in their hazard model. Their result seems counter-intuitive. The data they used has a richer set of information on tenure. They were able to exclude years of education from their prior experience variable, which I am not able to do. However, they find that more education decreases the hazard rates, so that would not explain why they found prior experience to decrease quits. More importantly, they were able to also control for time between jobs and the cumulative amount of time spent in the labor market. This very well could explain the differences between our results. Since my data did not include information on past labor market activity, it is not possible to determine if non-specific human capital or time spent outside the market, or both, are driving my result of prior experience increasing quits. At any rate, it seems less plausible that non-specific experience would decrease quit propensity, unless it is simply identifying individuals that are more attached to the labor market and have therefore built up more years of labor market experience.

As expected, collective bargaining reduces turnover. Collective bargaining tends to improve wages and working conditions[19]. The collective voice provided by unions, in terms of contracts and grievance procedures, provides a vehicle to resolve issues so that workers do not have to vote with their feet. I find that workers covered by a collective bargaining agreement have a one-percent lower quit rate. This suggests that collective bargaining does reduce quits, which is consistent with Light and Ureta (1992). One could argue, however, that unions are more likely to try to win, and actually win, representation with groups of workers that have a lower quit rate. There are a few groups organized within the sample period, the largest being the clerical bargaining unit.

The occupational dummies indicate that Clerical and Technical worker are the least likely to quit, followed by Service and Managers. Professional are most likely to exit, all else equal. This certainly seems reasonable since Professionals tend to be more mobile and are more likely to be competing in a national labor market.

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[19] See Freeman and Medoff (1981) for empirical evidence.

The notion that unions are simply organizing workers that are already less likely to quit can be addressed with the data at hand. Interacting "Collective Bargaining" with the occupational dummy variables allows us to explore the idea that collective bargaining has different impacts on quits depending on occupation. Thus, we add 4 additional dummy variables to Internal Wage model. The calculated likelihood ratio statistic is approximately 12. This is marginally significant at the .025 level but not at the .01 level. Only the Service/Blue Collar interaction coefficient is significant. Thus, there is evidence to suggest that the impact of collective bargaining does not depend on the occupation of the employee. This suggests that collective bargaining is reducing incentives to quit.

The estimates for the individual characteristics are similar to the results of other studies. I find that non-white workers are 2% more likely to exit and part time workers are 6% more likely to exit. In addition, marriage tends to decrease exit rates by 1%. These results are consistent with findings by Mclaughlin (1991) and to some degree, with Light and Ureta (1992). Married individuals are more likely to have dependents on their income, and thus spells of unemployment are more difficult to ride out. However, when Light and Ureta (1992) estimate their model separately for males and females, they find that marriage decreases the hazard rate for males but increases it for females. As found in almost every empirical study, I find that females are more likely to exit. The estimates suggest that an average female worker has an exit rate that is one percentage point higher than an observationally equivalent male. Light and Ureta (1992) and Shorey (1983) find evidence that the structural relationships differ between male and females. However, Viscusi (1980) did not find gender differences in quit behavior.

Wages and tenure have strong impacts on the likelihood of quits for state employees in Iowa. As found in other studies, the incidence of quits is very high in the first few years of employment. Individual wages seem to be more important than opportunity wages. Clearly, turnover can be reduced by increasing relative wages. Unfortunately, it is difficult to measure the costs of quits on both employers and employees. Thus, it is difficult to make any statements about the economic efficiency of the observed quits.

5.5 Impact of Comparable Worth on Parameter Identification

Comparable worth clearly impacts wages in Iowa state government. However, what impact this has upon our ability to estimate the model parameters is not so obvious. Mattila *et al* (1999) show that comparable worth was crucial in estimating an input demand system for state government. It would seem logical that it would also be crucial here as well. Mattila *et al* modeled broad inputs at an aggregate division level, while I am looking at quits by modeling the individual workers. There will be more variability in relative wages at an individual level as opposed to the division and aggregate occupational level.

The impacts of the comparable worth wage adjustments are explored by dividing the data into three time periods: pre-comparable worth, comparable worth, and post-comparable worth. Pre-comparable worth years are 1981-1984, comparable worth runs from 1985 through 1987, and post-comparable worth years are from 1988 through 1991. Estimates for the models are derived from each of the three periods. The internal wage model is re-estimated for the three sample periods, and the estimated parameters can be compared. The main interest is how comparable worth affects our ability to estimate the impact of wages on incentives to quit

Table 5.4 reports the estimates of model 1 for each period. Table 5.5 reports the parameter estimates, α, for each period relative to the full sample estimates that were reported in Table 5.2.

Although the comparable worth period covers the fewest years, the estimates for

Table 5.4 Internal Wage Model - Time Specific Parameter Estimates

	Pre Comp. Worth 81-84		Comp Worth 85-87		Post Comp Worth 88-91	
	α	$\partial p/\partial x$	α	$\partial p/\partial x$	α	$\partial p/\partial x$
dln(Wage)	-0.3170	-0.0362	-0.3518	-0.0402	-0.4233	-0.0483
	(0.0565)		(0.0777)		(0.0764)	
$\ln(\text{Wage})_{t-1}$	-0.1959	-0.0224	-0.2327	-0.0266	-0.4070	-0.0465
	(0.0388)		(0.0494)		(0.0432)	
Prior Experience	0.0090	0.0010	0.0153	0.0017	0.0120	0.0014
	(0.0006)		(0.0008)		(0.0007)	
State Tenure[b]	-0.0640	-0.0026	-0.0465	-0.0011	-0.0448	-0.0013
	(0.0030)		(0.0033)		(0.0028)	
State Tenure2**	0.0020		0.0018		0.0016	
	(0.0001)		(0.0001)		(0.0001)	
New Entrant	0.1767	0.0202	0.1352	0.0154	0.2978	0.0340
	(0.0215)		(0.0293)		(0.0260)	
Collective Bargaining	-0.0415	-0.0047	-0.0740	-0.0085	-0.1862	-0.0213
	(0.0167)		(0.0226)		(0.0197)	
Part Time	0.5536	0.0632	0.5721	0.0653	0.5286	0.0603
	(0.0289)		(0.0349)		(0.0298)	
Non-White	0.1665	0.0190	0.1324	0.0151	0.2037	0.0233
	(0.0331)		(0.0399)		(0.0278)	
Female	0.0576	0.0066	0.0844	0.0096	0.0699	0.0080
	(0.0160)		(0.0186)		(0.0160)	
Married	-0.0944	-0.0108	-0.0985	-0.0112	-0.1005	-0.0115
	(0.0138)		(0.0168)		(0.0148)	
Manager	-0.1357	-0.0155	-0.0151*	-0.0017	-0.0357*	-0.0041
	(0.0372)		(0.0420)		(0.0384)	
Technical	-0.2399	-0.0274	-0.1818	-0.0207	-0.1738	-0.0198
	(0.0277)		(0.0318)		(0.0284)	
Clerical	-0.2334	-0.0266	-0.2178	-0.0249	-0.2215	-0.0253
	(0.0250)		(0.0301)		(0.0265)	
Service/Blue Collar	-0.1185	-0.0135	-0.0864	-0.0099	-0.0920	-0.0105
	(0.0252)		(0.0278)		(0.0243)	
Intercept	0.2653*		0.3076*		1.3924	
	(0.2563)		(0.3387)		(0.3035)	
Log Likelihood	-23,273.2		-15,361.3		-19,312.0	
n	77,415		58,249		81,392	

* Indicates not significant at the .05 level. Standard Errors are listed in parentheses. Marginal effects are computed using the overall average for each x. [b] Marginal effect includes the impact of the quadratic term.

dln(Wage) and $\ln(\text{Wage})_{t-1}$ derived from the comparable worth period are closest to the full sample estimates. However, the magnitude of the estimates is very similar. The comparable worth period contains approximately 30% fewer observations than the pre or post periods do.

All in all, the results are very consistent between each of the periods. Qualitatively, the results are identical for the three periods. With the exception of the manager dummy and the annual dummies, all variables are significant at the .025 level. While this is not surprising given the number of observations, it is surprising to get such consistency across the sample periods given the variability seen in the exit rates over the years studied. The stability of the parameters suggests that reasonable estimates can be estimated with data that covers a narrower window of time. Clearly, comparable worth impacts the parameter estimates. However, the impact won't be as dramatic when micro data is used. In general, the ability to estimate the impact of wages on quits can be enhanced by using study behavior over longer periods and for periods where significant relative wage changes occur.

Table 5.5 - Internal Wage Gain Model Relative Parameter Estimates

	Pre Comp. Worth 81-84	Comp Worth 85-87	Post Comp Worth 88-91
dln(Wage)	0.85	0.95	1.14
$\text{Ln}(\text{Wage})_{t-1}$	0.73	0.86	1.51
Prior Experience	0.79	1.34	1.04
State Tenure	1.21	0.88	0.85
State Tenure2**	1.11	0.98	0.89
New Entrant	0.84	0.64	1.41
Collective Bargaining	0.47	0.84	2.11
Part Time	0.99	1.02	0.94
Non-White	0.93	0.74	1.14
Female	0.90	1.32	1.09
Married	0.94	0.98	1.00
Manager	2.17	0.24	0.57
Technical	1.19	0.90	0.86
Clerical	1.01	0.95	0.96
Service/Blue Collar	1.17	0.86	0.91

5.6 Comparable Worth Impact on Quits

The estimates presented previously indicate that wage gains reduce quits. Since comparable worth caused permanent shifts in relative wages, it is reasonable to ask what impact the comparable worth wage increases had on quits. Unlike across the board wage adjustments, changes in relative wages can have permanent effects on the makeup of the labor force.

The comparable worth wage adjustments occurred in two specific years, 1985 and 1987. This makes it easier to explore a counterfactual state where no comparable worth wage increases occurred. The strategy used is to back out the wage increases and use the model to predict quits in the absence of the comparable worth wage adjustments. These predictions are compared to the model predictions with the comparable worth wage increases left intact.

The comparable worth wage increases are identified by looking at the change in minimum salary in 1985 and 1987. Since all changes in the pay plans in 1985 were the result of comparable worth, the adjustment in 1985 can be measured as the log change in the minimum salary for each detailed job. The impact in 1987 is muddled by the fact that wage increases included non-comparable worth changes to the pay scales. Fortunately, previous work has identified the number of pay steps each occupation received as a result of comparable worth. Analysis of the pay plans indicates that there is approximately a 4% difference between each pay scale. Let $d\ln(\text{Min Wage})_{1985,i}$ be the log change in the minimum biweekly salary for job i in 1985. Let $CWADJ_i$ be the increase in the pay grade for job i as a result of comparable worth. The wage impact of comparable for job i in 1987, $dCW_{1987,i}$ is then calculated as

$$dCW_{1987i} = MAX(0.04*(CWADJ_i) - dln(Min\ Wage)_{1985,i}, 0).$$

The results are measures for the impact of comparable worth in 1987 for each job within the state government.

The mechanics of creating the data for the counterfactual state are straightforward. The calculated 1985 comparable worth wage change for a given job is subtracted from $dln(Wage)_{1985}$ and from $ln(Wage)_{t-1}$ for each year between 1986 and 1991. The calculated 1987 comparable worth wage change for a given job is subtracted from $dln(Wage)_{1987}$ and from $ln(Wage)_{t-1}$ for each year between 1988 and 1991. The remaining data is left intact.

Aggregate estimates for the impact of the comparable worth wage gains in the post comparable worth period are reported in Table 5.6. Estimates are shown for the five aggregate job classes, for male and female employees, and overall. The results suggest that comparable worth reduced quits by 3.9% or about 59 employees per year or 410 over the seven-year period.[20] However, the comparable worth wage adjustments did not affect each occupational category equally. As seen in Figure 5.4 previously, the average wage increase is the largest in the Service/Blue Collar occupations. It follows that the impact on quits would thus be the largest for this occupation. Comparable worth reduced the number of quits by an estimated 186 employees, about 5.7%, during 1885-1991. The predicted quit rate is estimated to be .44 percentage points lower. As you can see from the table, this estimate is much larger than for the other four occupations. A somewhat distant second were Clerical workers. The estimates suggest that the number of clerical quits was reduced by 3.5% over

[20] The simulations were also run without including the annual dummies in the model. This tended to magnify the estimated impact of comparable worth by about 105 quits overall. This could be the result of the dummies capturing the across the board wage increase or the dummies measuring some latent effect of comparable worth.

Table 5.6 – Impact of Comparable on Quits Based on the Counterfactual State of No Comparable Worth Wage Gains, 1985-1991

	Average Number of Employees Per Year	Exit Rate	Predicted Actual Exit Rate	Predicted Counter-Factual Exit Rate	Change in the Exit Rate	Change in number of Employees per Year	Percentage change in Exits
Manager	935	6.26%	5.76%	5.90%	-0.140%	-1	-2.43%
Prof.	5067	7.82%	8.09%	8.33%	-0.237%	-12	-2.93%
Tech.	2422	6.12%	5.95%	6.12%	-0.175%	-4	-2.94%
Clerical	5248	8.00%	7.83%	8.10%	-0.273%	-14	-3.49%
Serv/BC	6276	7.44%	7.48%	7.91%	-0.425%	-27	-5.67%
Male	10248	6.69%	6.70%	6.91%	-0.202%	-21	-3.01%
Female	9701	8.28%	8.26%	8.65%	-0.390%	-38	-4.73%
Overall	19949	7.47%	7.46%	7.75%	-0.294%	-59	-3.93%

1985-1991. This corresponds to around 187 workers (27 per year) over the seven years following comparable worth. The impacts on Professional and Technical workers are almost identical, although they have dissimilar quit rates. The number of quits was reduced by 2.9% in both cases. Since there are more than twice as many Professionals as Technical employees, the decrease in the number of Professionals is much higher than the number of Technicians. The impact on management, as you might expect, is very small.

Comparable worth adjustments tended to be larger in jobs traditionally dominated by females. Thus, it is logical to ask to what degree this slowed female quits. The results in Table 5.6 support the notion that females enjoyed larger comparable worth wage increases, and thus incentives for them to quit were reduced. The state labor force is almost evenly split with roughly 48% female.[21] The estimates are that females quits were reduced by 4.7% (265 quits), while quits involving male employees were reduced by 3% (145 quits.) Given the

[21] See Table 5.1

asymmetric impact of comparable worth, one might expect that females could gain in terms of their employment share. Figure 5.6 shows the actual proportion of female workers for each year. After 1987, the proportion of female employees climbed noticeably. From 1985 to 1991, the number of female employees posted a net increase of 397. Some of this was due to

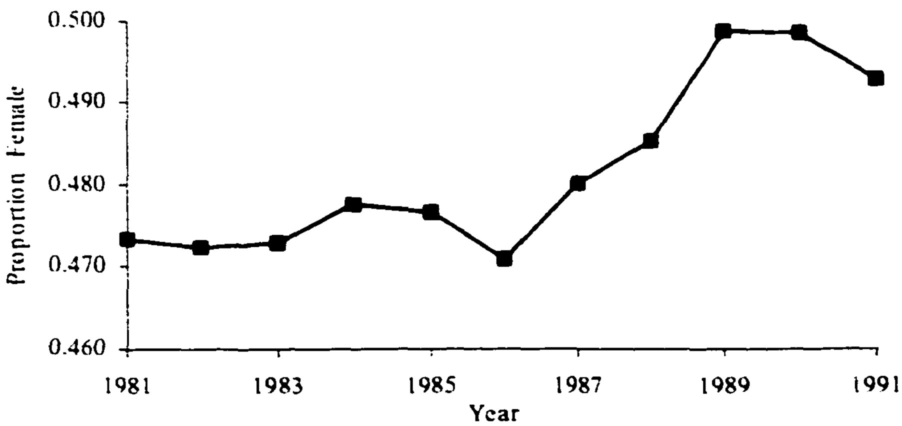

Figure 5.6 – Proportion of Female Workers 1981 to 1991

changes in overall employment levels. A corresponding decrease of 191 male employees occurred over the same period. The magnitude of the changes is not inconsistent with the simulation results. Furthermore, increases in relative wages most likely affected the relative number of female applicants. Orazem and Mattila (1998) found that comparable worth wage adjustments would induce more women applicants, especially in traditionally female dominated jobs. All this suggests that comparable worth had a somewhat larger impact on female employees, permanently increasing females' share of employment in the state labor force.

The next step is to look at the impact of the comparable worth adjustments for each year after comparable worth. Figure 5.7 graphs the predicted change in the number of quits

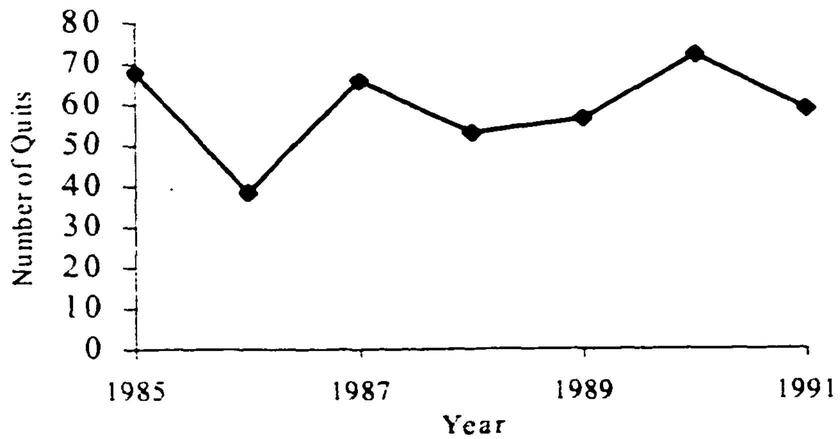

Figure 5.7 - Predicted Decrease in Quits As a Result of Comparable Worth

due to comparable worth. Figure 5.8 displays the corresponding predicted percentage change in quits each year. The largest impact in terms of number of employees appears to be in 1985 and in 1987, when the wage changes occurred. There was also a spike in 1990. This was most likely precipitated by the wage freeze that was imposed in that year. The accelerated exit rates in 1990, presumably because of the wage stagnation, magnified the estimated effect of comparable worth. The largest percentage changes in quits occurred in 1987. However, it does seem that comparable worth did have a persistent impact on quit rates. The impact seems to be very slowly declining after 1987.

5.7 Conclusions

Wage changes, especially relative wage changes, have a significant impact on quits. In addition to a contemporaneous effect, there appears to be a persistent impact that results from relative and absolute wage changes. Comparable worth did indeed induce a relative wage change. This seems to have reduced the number of quits by about 4%. This may have resulted in an increase in the proportion of females in the state labor force. The results of the

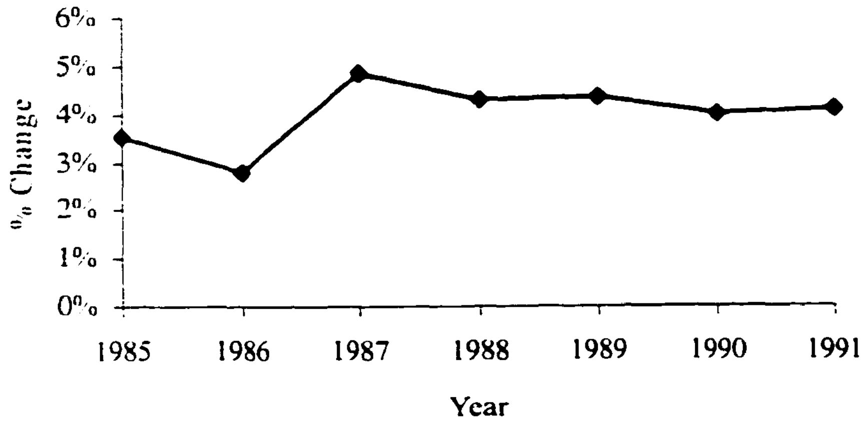

Figure 5.8 – Predicted Percentage Reduction in Quits Due to Comparable Worth

models estimated suggest that quits have an inelastic response to wages. Quits also are affected by, to a lesser degree, prevailing wages in the private sector. Unfortunately, I have no direct measure for the costs of separations to both employer and employees. Having this information would allow one to address the social costs and benefits of labor market turnover.

Other parameter values, for the most part, are consistent with previous work. Firm specific experience tends to decrease quits, while non-specific experience tends to increase this propensity. Collective bargaining has the expected effect of reducing quits. Female, single, non-white and part time workers are more likely to quit.

The next step is to use the model developed here to control for selection bias that may exist when we attempt to model union membership. This is the topic of the next chapter.

CHAPTER 6 - UNION MEMBERSHIP MODEL WITH NON-RANDOM SELECTION

6.1 Introduction

Observing voluntary payments of union dues provides a mechanism for the revelation of workers' preference for services unions provide. The model developed in chapter 4 assumed independence between the choice of union membership and occupation. Unfortunately, we only have the opportunity to observe an individual's contribution toward collective bargaining if he has already chosen to participate in the public sector labor market. This fact is not necessarily a problem as long as the stochastic disturbances for union membership and public sector employment are not correlated. However, the estimated parameters for the membership decision will be biased if correlation between the errors exists. Thus, the model developed in Chapter 5 is used to control for the potential correlation and provide unbiased estimates of the parameters of the dues models.

The estimated parameters for equation 5.3 will be used to develop the statistical controls for selection bias. The internal wage gain specification reported in Tables 5.2 and 5.3 will be used to derive a term, λ, as described in Heckman (1979). The ideal control for selection bias would model occupation choice for the entire labor market. However, the data necessary to develop such a model is difficult to obtain. I thus rely on the rich information of the payroll data and the model developed in Chapter 5 to provide the selection bias controls. This means the model only incorporates workers who were employed with the state at some point. Those workers who have never chosen state employment cannot be explicitly modeled.

6.2 Empirical Model with Selectivity Control

The model developed in Chapter 4 is extended to incorporate the non-random selection that may exist in the payroll data for the Iowa state government. Again, we assume that net utility from membership is

$$M^* = V(m=1, X_{it}, \mu_{it}) - V(m=0, X_{it}, \mu_{it}). \tag{6.1}$$

As before, we do not observe the net benefit from membership, M^*, but may observe its sign. Thus, as developed in Chapter 4, the probability that $M^* > 0$ is given by

$$P(M^* > 0) = \int_{-\infty}^{x_t\beta+\mu_t} \phi(z)dz = \Phi(x_t\beta + \mu_t). \tag{6.2}$$

Unfortunately, we cannot observe the sign of M^* for those individuals that exit the public sector labor market. An individual would choose to remain employed with the state if their net benefit is higher by remaining in their state job versus exiting and being employed in the private sector. Let $e=1$ represent the choice of remaining in state government employment, and let $e=0$ represent exiting state government for employment elsewhere. An individual will choose to remain in the public sector labor market if

$$E^* = U(e=1, X_{it-1}, \varepsilon_{it-1}) - U(e=0, X_{it-1}\varepsilon_{it-1}) > 0. \tag{6.3}$$

The stochastic error processes μ_t and ε_{t-1} are distributed

$(\mu_t, \varepsilon_{t-1}) \sim bn(0, 0, 1, \sigma^2, \rho)$.

The sign of M^* is only observed if E^* is greater than zero. Unless ρ is equal to zero, estimating 6.2 directly with the Iowa payroll data would yield biased parameter estimates. The estimation procedure must control for this potential bias. The model developed in Chapter 5 can be used to derive controls for the self-selection out of state employment. The decision to remain employed with the state government is represented as

$$P(E^* > 0) = \int_{-\infty}^{x_{t-1}\alpha + \varepsilon_{t-1}} \phi(z)dz = \Phi(x_{t-1}\alpha + \varepsilon_{t-1}). \qquad (6.4)$$

The estimated parameters for equation 6.4 are used to calculate a selection correction factor, λ. λ is calculated as

$$\lambda_{it} = \frac{\phi(x_{it-1}\alpha)}{\Phi(x_{it-1}\alpha)}. \qquad (6.5)$$

Incorporating the selection bias term into equation 6.2 we have

$$P(M^* > 0 | E^* > 0) = \int_{-\infty}^{x_t\beta + d\lambda_{it} + \mu_t} \phi(z)dz = \Phi(x_t\beta + d\lambda_t + \mu_t) \qquad (6.6)$$

Estimating equation 6.6 provides unbiased estimates of β. We can then calculate the elasticities specified in equations 4.4A and 4.4B by setting λ equal to zero. These elasticities are thus for the uncensored population.

6.3 Data

The payroll data for the Iowa state employees is again used to estimate the membership model. The 13 year panel of data provides the variability needed to identify parameters of both the membership and the exit models. The data used in Chapter 5 utilized information for both union and non-unionized employees. In other words, the model for exits is estimated using a broader population than just the organized workers in the Iowa state government. This allows us to develop the selection correction factor for those workers who may not have been in a job covered by a collective bargaining agreement in *t-1*.

6.4 Model Estimates

The general form of equation 6.6 was estimated for the absolute wage, the public/private wage and the relative wage models developed in Chapter 4. An additional dummy variable was added for employees who are new entrants into the state labor pool. Since λ is a function of x_{t-1}, it cannot be calculated for those workers who were not in the state labor pool in t-1. λ is defined to be zero for these cases, and the effect is captured through a dummy variable.

Table 6.1 contains the parameter estimates of the three specifications of equation 6.1. The parameter estimates for the most part do not change drastically when the selection bias is accounted for. The estimates for the impact of wage services are very similar. The impacts of employment changes are now all positive in all the specifications, but not significantly different from zero. It does appear that most of the parameter estimates, with the exception of the dues parameter, get smaller in magnitude once selection is accounted for.

The estimates for Pay Step have changed, however. The estimates for the absolute wage gain model and the public/private model changed from positive and significant to negative and significant when selection bias is controlled for. The Relative Wage model estimate for Pay Step decreased in magnitude but is still positive. It appears that the estimated positive impact of pay step on union membership was an artifact of self-selection. Those workers who have lower tenure will be more heavily represented at the lower steps of the pay plan. These workers are also more likely to exit. Once selectivity is controlled for, the impact of relative position appears to be negative. This implies that as workers gain job specific tenure, the willingness to pay for union membership decreases. This seems to be counter intuitive. Unions typically bargain for seniority rights of their constituents. When employees have to be reallocated, preferences are usually given to more senior workers. However, the collective bargaining agreement typically spells out these arrangements and the rules apply to

Table 6.1 - Union Demand Estimates with Selectivity Controls

Variable	Absolute Wage Gain Model		Public/Private Wage Gain Model		Relative Wage Gain Model	
	β	$\partial p/\partial x$	β	$\partial p/\partial x$	β	$\partial p/\partial x$
dln(Minimum Wage)	1.3577	0.3788				
	(0.1205)					
dln(Wage)	1.7321	0.4833				
	(0.0471)					
dln(CPS Wage)	0.1650*	0.0460				
	(0.0878)					
dln(Wage) - dln(Min Wage)			1.6515	0.4669		
			(0.0468)			
dln(Min Wage) - dln(CPS Wage)			0.9565	0.2704		
			(0.0652)			
dln(Relative Wage)					0.5838	0.1041
					(0.0376)	
ln(Relative Wage)					0.6701	0.1195
					(0.0191)	
ln(Wage)$_{t-1}$	1.8250	0.5092	1.7982	0.5084		
	(0.0319)		(0.0318)			
ln(CPS Wage)$_{t-1}$	-0.4280	-0.1194	-0.4559	-0.1289		
	(0.0198)		(0.0198)			
Pay Step	-0.3909	-0.1091	-0.3981	-0.1125	0.3856	0.0688
	(0.0653)		(0.0652)		(0.0622)	
Dues	-0.0048	-0.0013	-0.0046	-0.0013	-0.0034	-0.0006
	(0.0001)		(0.0001)		(0.0001)	
Total Dues/ 1000	-0.0010	-0.0003	-0.0010	-0.0003	-0.0014	-0.0002
	(0.0001)		(0.0001)		(0.0001)	
γ	0.0002		0.0002		0.0004	
dln(FTE)	0.4068*	0.1135	0.0760*	0.0215	0.0149*	0.0027
	(0.3065)		(0.3055)		(0.3037)	
Overtime Indicator	0.0763	0.0213	0.0728	0.0206	0.0632	0.0113
	(0.0086)		(0.0086)		(0.0085)	
Prior Experience	-0.0006*	-0.0002	-0.0010	-0.0003	-0.0005*	-0.0001
	(0.0004)		(0.0004)		(0.0004)	
State Tenure	-0.0113	-0.0031	-0.0113	-0.0032	-0.0092	-0.0016
	(0.0006)		(0.0006)		(0.0006)	
λ	0.2392		0.2314		0.4596	
	(0.0265)		(0.0264)		(0.0259)	
λ Dummy	-0.6209		-0.6242		-0.6307	
	(0.0164)		(0.0164)		(0.0162)	
Log Likelihood	-84401.9		-84665.8		-85661.0	

* Indicates not significant at the .05 level. Standard errors are listed in parentheses. Marginal effects are computed using the overall average for each x

Table 6.1 (Cont.)

Variable	Absolute Wage Gain Model		Public and Private Wage Gain Model		Relative Wage Gain Model	
	β	$\partial p/\partial x$	β	$\partial p/\partial x$	β	$\partial p/\partial x$
Part Time	-0.3678	-0.1026	-0.3772	-0.1066	-0.3807	-0.0679
	(0.0286)		(0.0286)		(0.0282)	
Non-White	0.1503	0.0419	0.1461	0.0413	0.1689	0.0301
	(0.0180)		(0.0180)		(0.0179)	
Female	0.1720	0.0480	0.1758	0.0497	0.0577	0.0103
	(0.0090)		(0.0089)		(0.0085)	
Married	-0.0529	-0.0148	-0.0513	-0.0145	-0.0545	-0.0097
	(0.0079)		(0.0079)		(0.0078)	
Technical	0.4756	0.1327	0.4620	0.1306	0.2396	0.0427
	(0.0188)		(0.0187)		(0.0180)	
Clerical	0.1312	0.0366	0.1021	0.0289	-0.3175	-0.0566
	(0.0171)		(0.0170)		(0.0141)	
Service/Blue Collar	1.2063	0.3366	1.1892	0.3362	0.7907	0.1410
	(0.0176)		(0.0175)		(0.0148)	
Intercept	-9.3890		-8.7809		-0.8595	
	(0.1944)		(0.1921)		(0.0709)	
D82	-0.0321*	-0.0089	-0.0941	-0.0266	-0.6409	-0.1143
	(0.0257)		(0.0255)		(0.0223)	
D83	0.1126	0.0314	-0.0680	-0.0192	-0.5890	-0.1050
	(0.0271)		(0.0259)		(0.0219)	
D84	-0.0215*	-0.0060	-0.1187	-0.0336	-0.5786	-0.1032
	(0.0268)		(0.0264)		(0.0241)	
D85	-0.0803	-0.0224	-0.1365	-0.0386	-0.5416	-0.0966
	(0.0227)		(0.0226)		(0.0205)	
D86	0.0429*	0.0120	-0.1547	-0.0437	-0.4842	-0.0864
	(0.0213)		(0.0195)		(0.0174)	
D87	-0.0526*	-0.0147	-0.0926	-0.0262	-0.4001	-0.0714
	(0.0279)		(0.0278)		(0.0268)	
D88	0.0305*	0.0085	-0.0541	-0.0153	-0.2953	-0.0527
	(0.0220)		(0.0216)		(0.0207)	
D89	-0.0548*	-0.0153	-0.1258	-0.0356	-0.3137	-0.0559
	(0.0309)		(0.0306)		(0.0300)	
D90	-0.0816	-0.0228	-0.1566	-0.0443	-0.2717	-0.0485
	(0.0203)		(0.0200)		(0.0196)	
D91	0.1482	0.0413	-0.0530	-0.0150	-0.0788	-0.0141
	(0.0197)		(0.0176)		(0.0170)	

all covered workers. Hence, the rules are very much a public good. These rules do little in the way of protecting newer employees. It could be argued that as a worker gains tenure, the value of the seniority rules increases, since they are less likely to be adversely affected when employment changes. Less tenured workers are more likely to be adversely impacted and presumably then perceive a higher likelihood of filing a grievance via the collective bargaining grievance procedures. While unions may be obliged to represent all workers in formal grievance procedures, they most likely would not pursue a case involving a non-member as aggressively as they might dues paying members.

Table 6.2 contains the calculated elasticities for the models with and without the selection correction. For the most part, the magnitudes increase when selection bias is accounted for. Remember that it appeared that the parameter estimates were getting smaller when the selection correction was included. However, the predicted probability of paying dues when evaluated at sample means also became smaller (from .33 to .20), and so the resulting elasticities became larger when evaluated at sample means. Consequently, the elasticities are being evaluated at different points on the normal density and distribution functions even though the same values for x are used in both cases. Since the uncensored population has a lower membership rate, and both the censored and uncensored populations are in the lower half of the distribution, $\Phi(x\beta)$ decreases proportionally more than $\varphi(x\beta)$. Consequently the ratio, $\varphi(x\beta)$ $\Phi(x\beta)$ increases. This increase dominates the effect of the smaller βs estimated with the correction for sample selection.

The effect of relative wages and wage gains has essentially the same interpretation with and without the selection bias term. However, controlling for selection bias, the

Table 6.2 - Union Demand Elasticities and Dues Crowding Parameter

	Absolute Wage Gain Model		Public/Private Wage Gain Model		Relative Wage Gain Model	
	Without λ	With λ	Without λ	With λ	Without λ	With λ
dln(Minimum Salary)	1.52	1.90				
dln(Wage)	1.98	2.43				
dln(CPS Wage)	0.08	0.23				
dln(Act Sal) - dln(Min Sal)			1.89	2.30		
dln(Min Sal) - dln(CPS Wage)			1.16	1.33		
dln(Relative Wage)					0.68	1.02
ln(Relative Wage)					0.84	1.17
ln(Actual Salary)$_{t-1}$	2.10	2.56	2.07	2.50		
ln(CPS Wage)$_{t-1}$	-0.53	-0.60	-0.56	-0.63		
Pay Step	0.41	-0.63	0.40	-0.64	1.70	0.77
Dues	-0.84	-1.14	-0.80	-1.08	-0.56	-1.01
Total Dues /1000	-0.17	-0.21	-0.18	-0.21	-0.24	-0.35
γ	0.0002	0.0002	0.0003	0.0002	0.0005	0.0004
dln(FTE)	0.20	0.57	-0.18	0.11	-0.22	0.03

elasticity for Pay Step becomes negative, since the parameter estimate is negative, in the model with absolute wage and public/private wage gains. This is again likely due to the correlation between tenure and propensity to exit.

One of the more interesting results is the elasticity for the dues rate. In the absolute wage gain model and the public/private wage gain, the magnitudes are now slightly greater than one and the relative wage gain model essentially equal to one. These estimates are consistent with an equilibrium where union revenues are maximized. Unitary elasticity has some implications as to the "public" nature of union services. Public goods are necessarily non-rival in consumption. In this case, allowing more members to join, at least at the margin, would not increase the costs of providing services, given the definition of the bargaining units. However, membership rates could have some impact on the efficiency of the union in providing union services. Unions with more members may have more leverage with management during contract negotiations. However, this may call into question the exogeniety of the wage gains received by workers. At the individual level, wages are assumed exogenous to the employee.

A pure public good implies that individuals do not care who purchases additional units of the good. Once the good is produced, it is available for all to share in non-rival consumption. While the estimated impact of other's dues contributions does decrease the likelihood of paying dues, own contributions have a much larger impact. This suggests that others' contributions are not perfect substitutes for their own contributions.

The crowding parameter, γ, is again simply the ratio of the effect of others' contributions relative to the impact of own contributions. The estimates are almost identical across the three specifications. The estimates are very small, consistent with findings in Chapter 4. While the other contributions do have a negative and statistically significant impact on membership, this effect is much smaller than their own contributions. This suggests that union membership is much closer to a private good than to a pure public good.

6.5 Impact of Comparable Worth on Parameter Identification

Comparable worth provided permanent shocks to relative wages in the Iowa state government. Chapter 5 demonstrated that comparable worth was important to estimate the impact of wages on quits. It should be expected that comparable worth wage adjustments would also enhance our ability to estimate the impact of wages on union membership.

The impacts of the comparable worth wage adjustments are explored, as was done in Chapter 5, by dividing the data into three periods: pre-comparable worth, comparable worth, and post-comparable worth. Pre-comparable worth years are 1982-1984, comparable worth runs from 1985 to 1987, and post-comparable worth years are from 1988 to 1992. The Public/Private Wage Gain model is estimated for each of the periods. The results are reported in Table 6.3.

The difference in the estimates across the sub-periods is greatest for the wage variables. The pre-comparable worth estimates have smaller estimated impacts of public and private wage gains and of wage levels, and the estimated impacts of external wages and public wage gains are not significant. This supports the notion that relative wage changes caused by comparable worth are necessary to estimate the effects of wages. The pre-comparable worth period and the comparable worth period estimates are derived from the same number of years. Sample sizes differ as the clerical bargaining unit was not organized until 1985 and the professional bargaining unit was not organized for 1983. The wage estimates for the comparable worth period are similar to the post comparable worth period, which has 5 years of data, and the overall estimates reported earlier. The wage estimates derived from the comparable worth and the post-comparable worth periods are very similar. The impact is most dramatic on the wage variables. If one had only the pre-comparable worth sample, the inferences about the role of wages would be much different.

6.6 Comparable Worth Impact on Union Membership

The estimates presented previously have shown union membership to be very wage elastic. Thus, it is reasonable to expect that the comparable worth wage gains had a dramatic effect on union membership. This hypothesis is supported by the data: union membership did increase after the comparable worth pay plans were implemented. This section uses a similar

Table 6.3 - Union Demand -Public/Private Wage Gain - Time Specific Parameter Estimates

	Pre Comp. Worth 82-84		Comp. Worth 85-87		Post Comp. Worth 88-92	
Variable	β	∂p/∂x	β	∂p/∂x	β	∂p/∂x
dln(Wage) - dln(Min Wage)	0.2401	0.0679	1.2996	0.3674	2.0083	0.5677
	(0.1088)		(0.1015)		(0.0711)	
dln(Min Wage) - dln(CPS Wage)	0.0870*	0.0246	2.0269	0.5730	0.9741	0.2754
	(0.1409)		(0.1307)		(0.0991)	
ln(Wage)$_{t-1}$	0.2914	0.0824	1.6869	0.4769	2.1115	0.5969
	(0.0850)		(0.0620)		(0.0457)	
ln(CPS Wage)$_{t-1}$	0.0166*	0.0047	-0.2773	-0.0784	-0.6858	-0.1939
	(0.0488)		(0.0387)		(0.0274)	
Pay Step	0.6569	0.1857	-0.0590*	-0.0167	-0.9710	-0.2745
	(0.1409)		(0.1307)		(0.0991)	
Dues	-0.0032	-0.0009	-0.0042	-0.0012	-0.0045	-0.0013
	(0.0003)		(0.0002)		(0.0002)	
Total Dues/ 1000	-0.0065	-0.0018	-0.0016	-0.0005	-0.0009	-0.0002
	(0.0003)		(0.0001)		(0.0001)	
γ	0.0021		0.0004		0.0002	
dln(FTE)	-1.5922	-0.4501	0.4211*	0.1190	1.4301	0.4043
	(0.7544)		(0.5112)		(0.5127)	
Overtime Indicator	0.1467	0.0415	0.0495	0.0140	0.0742	0.0210
	(0.0192)		(0.0165)		(0.0120)	
Prior Experience	0.0003*	0.0001	-0.0002*	-0.0001	-0.0009*	-0.0003
	(0.0009)		(0.0008)		(0.0007)	
State Tenure	-0.0026*	-0.0007	-0.0133	-0.0038	-0.0120	-0.0034
	(0.0016)		(0.0013)		(0.0009)	
λ	0.4426		0.1988		0.2477	
	(0.0619)		(0.0505)		(0.0372)	
λ Dummy	-0.7040		-0.6636		-0.5895	
	(0.0318)		(0.0306)		(0.0254)	
Part Time	-0.4623	-0.1307	-0.4618	-0.1306	-0.3134	-0.0886
	(0.0781)		(0.0562)		(0.0374)	
Non-White	0.2035	0.0575	0.1668	0.0472	0.1328	0.0375
	(0.0447)		(0.0378)		(0.0232)	
Female	0.1073	0.0303	0.2100	0.0594	0.1706	0.0482
	(0.0201)		(0.0176)		(0.0125)	
Married	-0.0534	-0.0151	-0.0560	-0.0158	-0.0593	-0.0168
	(0.0177)		(0.0149)		(0.0110)	
Technical	0.6041	0.1708	0.5594	0.1582	0.5042	0.1426
	(0.0396)		(0.0380)		(0.0278)	
Clerical	0.1962	0.0555	0.1912	0.0541	0.1365	0.0386
	(0.0391)		(0.0365)		(0.0238)	
Service/Blue Collar	1.2037	0.3403	1.2552	0.3549	1.2435	0.3515
	(0.0375)		(0.0372)		(0.0247)	
	n=31.822		n=42.144		n=74.043	

* Indicates not significant at the .05 level Standard errors are listed in parentheses. Marginal effects are computed using the overall average for each x and the full parameter estimates to evaluate the normal density function.

methodology as that used in Chapter 5 to estimate the impact of the comparable worth wage gains on union membership.

As in Chapter 5, the comparable worth wage increases are identified by looking at the changes in minimum salary in 1985 and 1987. Since all changes in the pay plans in 1985 were the result of comparable worth; the adjustment in 1985 can be measured as the log change in the minimum salary for each detailed job. The impact in 1987 is obscured by the fact that wage increases included non-comparable worth changes to the pay scales. Analysis of the pay plans indicates that there is approximately a 4% difference between each pay scale. Let $dln(Min\ Wage)_{1985,i}$ be the log change in the minimum biweekly salary for job i in 1985. Let $CWADJ_i$ be the increase in the pay grade for job i as a result of comparable worth. The wage impact of comparable worth for job i in 1987, $dCW_{1987,i}$ is calculated as

$$dCW_{1987i} = MAX(0.04*(CWADJ_i) - dln(Min\ Wage)_{1985,i},\ 0).$$

The results are measures for the impact of comparable worth in 1987 for each job within the state government.

The mechanics of creating the data for the counterfactual state are straightforward. The calculated 1985 comparable worth wage change for a given job is subtracted from $dln(Wage)_{1985}$, $dln(Min\ Wage)_{1985}$, and from $ln(Wage)_{t-1}$ for each year between 1986 and 1992. The calculated 1987 comparable worth wage change for a given job is subtracted from $dln(Wage)_{1987}$, $dln(Min\ Wage)_{1987}$, and from $ln(Wage)_{t-1}$ for each year between 1988 and 1992. The remaining data is left intact. The estimates from the Public/Private wage gain model used are used to evaluate the change in membership resulting from comparable worth wage gains. Comparable worth wage adjustments are job specific, so they do not impact the measured private wage gains.

Aggregate estimates for the impact of the comparable worth wage gains in the post comparable worth period are reported in Table 6.4. Estimates are shown for the five

Table 6.4 - Impact of Comparable Worth on Union Membership Based on the Counterfactual State of No Comparable Worth Wage Gains, 1985-1992

	Actual		Predicted		Change Due to Comparable Worth		
	Number of Employees Per Year	Membership Rate	Actual Membership Rate	Counterfactual Membership Rate	Membership Rate	Number of Members per Year	Percentage change in Membership
Prof.	2448	30.51%	31.02%	28.37%	2.645%	65	8.53%
Tech	2054	28.67%	28.63%	25.68%	2.946%	60	10.29%
Clerical	4384	21.23%	21.34%	18.31%	3.032%	133	14.20%
Serv/BC	5637	53.78%	53.71%	46.79%	6.920%	390	12.88%
Male	7150	40.30%	40.70%	37.19%	3.511%	251	8.63%
Female	7373	32.78%	32.56%	27.17%	5.387%	397	16.55%
Overall	14523	36.48%	36.57%	32.10%	4.464%	648	12.21%

aggregate job classes, for male and female employees, and overall. The results suggest that comparable worth increased union membership 12.2% or about 648 members per year. This corresponds to an increase of about 4.5 percentage points in the membership rate. The magnitude is expected considering the estimated wage elasticities. Comparable worth wage adjustments did not affect each occupational category equally. As seen in Figure 5.4 previously, the average wage increase is the largest in the Service/Blue Collar occupations. It follows that the impact on membership would thus be the largest for this occupation. Comparable worth increased Service/Blue Collar membership by an estimated 390 members per year, about 12.9%, during 1885-1992. The predicted membership rate is estimated to be 6.9 percentage points higher, more than twice the rate increase in the other occupations. The number of members among Clerical workers increased by 14.2% per year over the period

1985-1992. This corresponds to 133 members per year. The impact on Professional and in Technical workers is similar. Comparable worth increased membership for Professionals by 65 members (8.5%) per year, while Technical membership increased by 60 members (10.3%) per year.

As stated before, the estimated elasticities for wages are large. Thus, it should be expected that the wage changes induced by comparable worth would have an estimated permanent effect on union membership. If no other relative wage shocks occur, the estimated models should provide predictions of the impact of comparable worth that converge to a constant percentage point increase. This in fact will not happen because other changes occur throughout the period.

Figure 6.1 graphs the estimated increase in members by year, and Figure 6.2 graphs the estimated change in the membership rate resulting from comparable worth wage changes. It seems clear that effects of comparable worth are not temporary. After the period of comparable worth wage gains, the impact seems to flatten out to slightly more than 600 members per year. One should keep in mind that the model has annual dummy variables that are capturing the overall effect of the comparable worth, among other things. Excluding the annual dummies from the model resulted in an increase in the estimated impact of comparable worth by about 20 members per year. The number of covered workers peaked in 1990, and thus this year had the largest estimated impact of comparable worth on the number of members. Since the actual membership rate continued to climb towards 50% throughout the post comparable worth period, the estimated difference in the membership rate also climbed[22].

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[22] The marginal effect is determined by the wage parameters and the value of the density function. The density function reaches a maximum when the cumulative distribution function is 0.5.

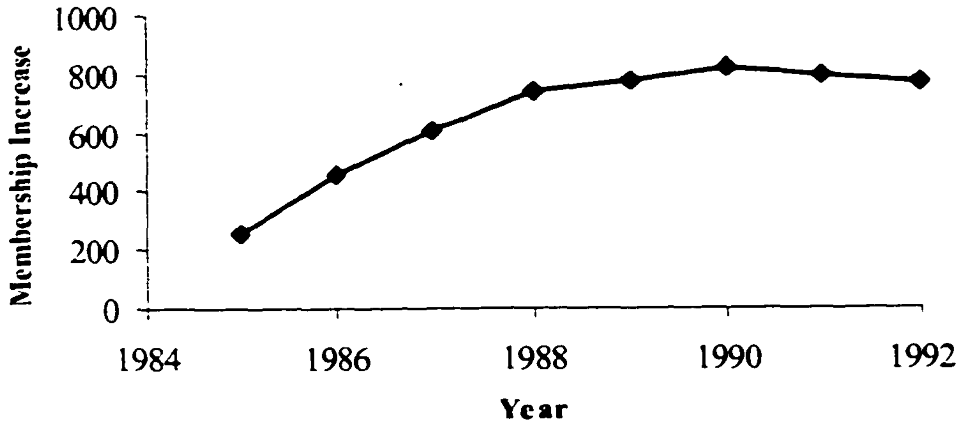

Figure 6.1 - Estimated Increase in Union Members Due to Comparable Worth

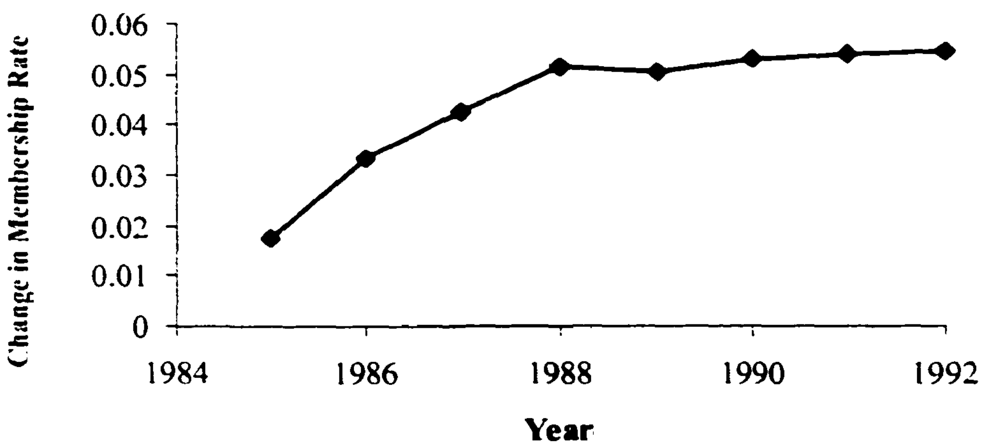

Figure 6.2 - Increase in the Membership Rate Due to Comparable Worth

Comparable worth adjustments tended to be larger in jobs traditionally dominated by females. Thus, we should expect to see a larger increase in female membership, which the estimates support. Female membership increased by an estimated 16.6% (397 members per year), while male membership was increased by 8.6% (251 members). Figure 6.3 displays the predicted[23] and the counterfactual membership estimates for each year. There is an apparent upswing in membership during the post-comparable worth years. This was true for both males and females. However, the impact on the female membership appears to be larger. This reinforces the idea that comparable worth had its largest effects on jobs that were traditionally dominated by females. Thus, the comparable worth wage caused a disproportionate increase in female membership.

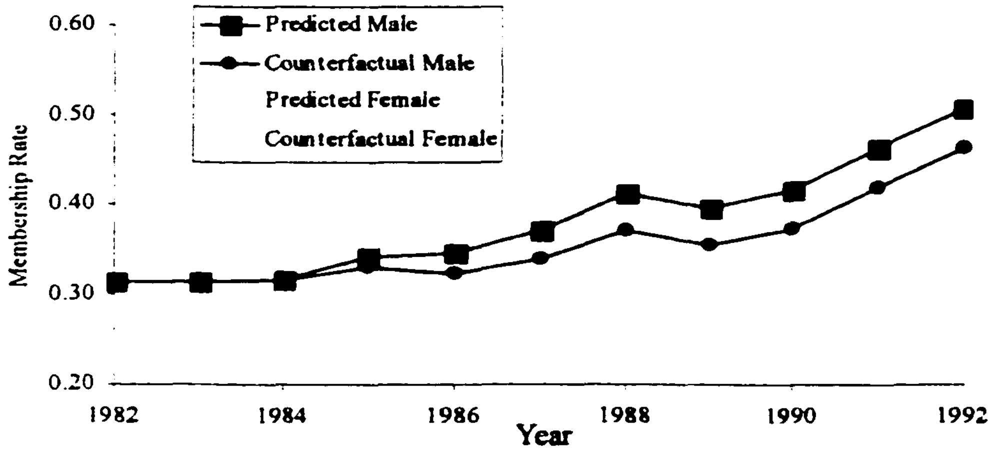

Figure 6.3 - Annual Predicted and Counterfactual Membership Rates for Males and females

6.5 Conclusions

The union demand model developed in chapter 4 was extended to control for non-random selection of workers in the state labor market. The control for non-random selection was developed from the model of quits developed in Chapter 5.

[23] Since annual dummies are included in the model, there is almost no difference between the actual and the aggregate predicted membership rate for each year.

Selection bias does appear to affect the model estimates. The most notable change is that the estimated elasticity for their own dues contributions becomes very close to unity. This suggests an equilibrium where union dues revenue is maximized. The relative magnitudes of own contributions and the contributions of others again suggest the joint product model is appropriate. While the contributions of others affect membership, they are far from perfect substitutes for their own contributions.

There appears to be a more elastic response to private wage gains than to non-employee specific rates of wage increase. This is consistent with the union providing members some benefits that are not provided to non-members. However, the ability to exclude the non-members would negate the union argument that free riders weaken the union's ability to provide collective bargaining at an optimal level. It is more likely that state employees who get atypical wage gains are more favorably disposed to the public sector union.

Comparable worth wage adjustments provided an exogenous shock to relative wages that is critical when estimating the demand for union services. The estimates suggest that comparable worth increased union membership by about 650 members per year. Since females were in jobs which got larger wage increases, large membership increases were observed among females. Similarly, Service/Blue Collar and Clerical workers had the largest increases in membership.

Selection does not affect the qualitative conclusions derived earlier. Full time, non-white, female, single, and Service/Blue Collar workers value union services more. This implies that one need not be overly concerned with non-random selection when trying to make qualitative statements about different sub-groups of the labor pool.

CHAPTER 7 – CONCLUSIONS

This study developed a model of union membership in a framework where union membership jointly produces a public and private goods. Included in the model are measures of the wage increases that go to individuals and wage increases that go to the group as a whole. In addition, the model includes measures to control for the price of union membership. Thus, willingness to pay at the margin is also addressed.

The data used in this study provided significant advantages in modeling union membership. Because public sector union membership is voluntary, it is possible to identify the demand for union services separately from the decision to accept employment in the public sector. The data contains detailed wage information on both individual wages and wage scales for jobs. Individual wages were subject to exogenous shocks from comparable worth wage adjustments that occurred during the sample period. Shocks to relative wages are rare in public sector labor markets. Most studies have not been able to control for union dues and, thus, have not been able to address willingness to pay for union services. Finally, the data allowed estimation of an empirical model of quits This model was used to provide statistical controls for non-random selection.

Union membership is very responsive to wage gains. As found in some other studies, union membership is positively related to earnings. Membership was found to be wage elastic. Also, membership appears to be more responsive to idiosyncratic wage gains than wage gains that accrue to all employees. The model was also used to estimate the impact that comparable worth wage increases had on membership. The wage gains from comparable worth increased membership by over 12%. This is consistent with the actual increase in membership that is observed in the post-comparable worth period.

Union membership appears to be price inelastic when selection bias is not controlled for. Once controls for the potential bias are included, it appears that union membership is unitary elastic. It does not appear that other's contributions are close substitutes for own contributions. This suggests that union contributions are, for the most part, private goods.

Most exits occur in the early years of the employment relationship. Thus, early in the employment relationship, the likelihood of observing an exit declines as the individual gains state-specific experience. Quits are negatively related to wage gains and positively related to private sector wage gains. As a result of this estimated relationship, comparable worth wage gains reduced incentives for female and traditionally lower wage employees to exit. This estimated decrease in female quit rate is corroborated by an observed increase in the proportion of females in the state labor force.

Further study on the demand for private good aspects union services could be enhanced if one could have access to data on grievance procedures and other measures of employee/management relations.

APPENDIX - COMPARATIVE STATICS OF JOINT PRODUCT MODEL

Comparative static results will be derived for (2.7). Throughout this appendix the i subscripts will be suppressed unless it is necessary for clarity. The first order conditions for this problem as listed in (2.8) are

$$\Psi_1 = u_y - \lambda p_y = 0\,0 \qquad A.1$$

$$\Psi_2 = u_q f'(x_i) + u_z g'(x_i + X_i^*) - \lambda p_x = 0 \qquad A.2$$

$$\Psi_3 = I - P_y y_i - P_x x_i = 0\,0 \qquad A.3$$

The Jacobian is

$$J = \begin{pmatrix} \frac{\partial \Psi_1}{\partial y_i} & \frac{\partial \Psi_1}{\partial x_i} & \frac{\partial \Psi_1}{\partial \lambda} \\ \frac{\partial \Psi_2}{\partial y_i} & \frac{\partial \Psi_2}{\partial x_i} & \frac{\partial \Psi_2}{\partial \lambda} \\ \frac{\partial \Psi_3}{\partial y_i} & \frac{\partial \Psi_3}{\partial x_i} & \frac{\partial \Psi_3}{\partial \lambda} \end{pmatrix} = \begin{pmatrix} u_{yy} & B & -p_y \\ B & A & -p_x \\ -p_y & -p_x & 0 \end{pmatrix}. \qquad A.4$$

where: $A \equiv U_{qq} f'(x)^2 + 2U_{qz} g'(x) f'(x) + U_{zz} g'(x)^2 + U_q f''(x) + U_z g''(x)$

$B \equiv U_{yq} f'(x) + U_{yz} g'(x)$

$|J| > 0$ for maximum

The comparative static for a change in others' contributions X^* can be expressed as

A.5

$$J\begin{pmatrix} \dfrac{\partial y_i}{\partial X_i^*} \\ \dfrac{\partial x_i}{\partial X_i^*} \\ \dfrac{\partial \lambda_i}{\partial X_i^*} \end{pmatrix} = \begin{pmatrix} -\dfrac{\partial \psi_1}{\partial X_i^*} \\ -\dfrac{\partial \psi_2}{\partial X_i^*} \\ -\dfrac{\partial \psi_3}{\partial X_i^*} \end{pmatrix} = \begin{pmatrix} -u_{yz}g'(x_i + X_i^*) \\ -u_{qz}g'(x_i + X_i^*)f'(x_i) - u_{zz}\left(g'(x_i + X_i^*)\right)^2 - u_z g''(x_i + X_i^*) \\ 0 \end{pmatrix}$$

Using Cramer's rule one can solve for $\partial x_i/\partial X^*$ as

A.6

$$\frac{\partial x_i}{\partial X_i^*} = \frac{\begin{vmatrix} u_{yy} & -u_{yz}g'(x_i + X_i^*) & -p_y \\ B & -u_{qz}g'(x_i + X_i^*)f'(x_i) - u_{zz}\left(g'(x_i + X_i^*)\right)^2 - u_z g''(x_i + X_i^*) & -p_x \\ -p_y & 0 & 0 \end{vmatrix}}{|J|} \quad 0$$

$$= \frac{P_y P_x U_{yz} g'(x_i + X_i^*) - \left[-P_y^2 \left(U_{qz} g'(x_i + X_i^*) f'(x) + U_{zz} g'(x_i + X_i^*)^2 + U_z g''(x_i + X_i^*) \right)\right]}{|J|}$$

0

The marginal change in agent's consumption of y when the contributions of others increase is

A.7

$$\frac{\partial y_i}{\partial X_i^*} = \frac{\begin{vmatrix} -u_{yz}g'(x_i + X_i^*) & u_{yq}f'(x_i) + u_{yz}g'(x_i + X_i^*) & -p_y \\ -u_{qz}g'(x_i + X_i^*)f'(x_i) - u_{zz}g'(x_i + X_i^*)^2 - u_z g''(x_i + X_i^*) & A & -p_x \\ 0 & -p_x & 0 \end{vmatrix}}{|J|}$$

$$= \frac{p_x p_y\left(-u_{qz} g'(x_i + X_i^{*}) f'(x_i) - u_{zz} g'(x_i + X_i^{*})^2 - u_z g''(x_i + X_i^{*})\right) + p_x^2 u_{yz} g'(x_i + X_i^{*})}{|J|}$$

The comparative static effects of changes in income can be represented as

A.8

$$J\begin{pmatrix} \frac{\partial y}{\partial I} \\ \frac{\partial x}{\partial I} \\ \frac{\partial \lambda}{\partial I} \end{pmatrix} = \begin{pmatrix} -\frac{\partial \Psi_1}{\partial I} \\ -\frac{\partial \Psi_2}{\partial I} \\ -\frac{\partial \Psi_3}{\partial I} \end{pmatrix} = \begin{pmatrix} 0 \\ 0 \\ -1 \end{pmatrix}$$

The marginal effect of changes in income on consumption of y can again be derived using Cramer's rule. The comparative static for y is

A.9

$$\frac{\partial y}{\partial I} = \frac{\begin{vmatrix} 0 & B & -p_y \\ 0 & A & -p_x \\ -1 & -p_x & 0 \end{vmatrix}}{|J|}$$

$$= p_x\left(u_{yq} f'(x) + u_{yz} g'(x + X^{*})\right) - p_y\left(u_{qq} f'(x)^2 + 2u_{qz} g'(x + X^{*}) f'(x) + u_{zz} g'(x + X^{*}) + u_q f''(x) + u_z g''(x + X^{*})\right) \Big/ |J|$$

The marginal effect of I on x_i is

A.10

$$\frac{\partial x_i}{\partial I} = \frac{\begin{vmatrix} u_{yy} & 0 & -p_y \\ B & 0 & -p_x \\ -p_y & -1 & 0 \end{vmatrix}}{|J|} = \frac{p_y\left(u_{qy} f'(x) + u_{zy} g'(x + X^{*})\right) - p_x u_{yy}}{|J|}$$

The comparative static results with respect to p_y are derived from

A.11

$$J\begin{pmatrix}\frac{\partial y}{\partial p_y}\\ \frac{\partial x}{\partial p_y}\\ \frac{\partial \lambda}{\partial p_y}\end{pmatrix}=\begin{pmatrix}-\frac{\partial \Psi_1}{\partial p_y}\\ -\frac{\partial \Psi_2}{\partial p_y}\\ -\frac{\partial \Psi_3}{\partial p_y}\end{pmatrix}=\begin{pmatrix}\lambda\\ 0\\ y\end{pmatrix}.$$

The change in the optimal consumption of y when p_y changes is

A.12

$$\frac{\partial y}{\partial p_y}=\frac{\begin{vmatrix}\lambda & B & -p_y\\ 0 & A & -p_x\\ y & -p_x & 0\end{vmatrix}}{|J|}$$

$$=\Big(-p_x y\left(u_{yq}f'(x)+u_{yz}g'(x+X^*)\right)+p_y y\Big(u_{qq}f'(x)^2+2u_{qz}g'(x+X^*)f'(x)$$

$$+u_{zz}g'(x+X^*)^2+u_q f''(x)+u_z g''(x+X^*)\Big)-p_x^2\lambda\Big)\Big/|J|\ .$$

The change in optimal contributions of x when p_y changes is

A.13

$$\frac{\partial x}{\partial p_y} = \frac{\begin{vmatrix} u_{yy} & \lambda & -p_y \\ B & 0 & -p_x \\ -p_y & y & 0 \end{vmatrix}}{|J|}$$

$$= \left(-p_y y\left(u_{yq} f'(x) + u_{yz} g'(x + X^*) \right) + p_y p_x \lambda + p_x y u_{yy} \right) \Big/ |J|$$

The comparative static for changes in p_x can be derived from the following.

A.14

$$J \begin{pmatrix} \dfrac{\partial y}{\partial p_x} \\ \dfrac{\partial x}{\partial p_x} \\ \dfrac{\partial \lambda}{\partial p_x} \end{pmatrix} = \begin{pmatrix} -\dfrac{\partial \Psi_1}{\partial p_x} \\ -\dfrac{\partial \Psi_2}{\partial p_x} \\ -\dfrac{\partial \Psi_3}{\partial p_x} \end{pmatrix} = \begin{pmatrix} 0 \\ \lambda \\ x \end{pmatrix}.$$

The change in optimal consumption of y when p_x changes is

A.15

$$\frac{\partial y}{\partial p_x} = \frac{\begin{vmatrix} 0 & B & -p_y \\ \lambda & A & -p_x \\ x & -p_x & 0 \end{vmatrix}}{|J|}$$

$$= \Big(\lambda p_x p_y - p_x x\left(u_{yq} f'(x) + u_{yz} g'(x + X^*) \right)$$

$$+ p_y x\left(u_{qq} f'(x)^2 + 2u_{qz} g'(x + X^*) f'(x) + u_{zz} g'(x + X^*)^2 + u_q f''(x) + u_z g''(x + X^*) \right)\Big) \Big/ |J|$$

The change in optimal x when p_x changes is

A.16

$$\frac{\partial x}{\partial p_x} = \frac{\begin{vmatrix} u_{yy} & 0 & -p_y \\ B & \lambda & -p_x \\ -p_y & x & 0 \end{vmatrix}}{|J|}$$

$$= \left(-p_y x \left(u_{yq} f'(x) + u_{yz} g'(x + X^*) \right) - p_x^2 \lambda + p_x x u_{yy} \right) \Big/ |J|$$

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