A continuous time panel study of the size and timing of negotiated wage changes: the effects of British incomes policy 1950-73

by

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Abstract

In this thesis, we propose an econometric estimator for a repeat-spell duration model involving fixed effects. Using data on nationally negotiated pay settlements in Britain between 1950 to 73, this estimator is applied to a model that simultaneously determines the timing as well as the magnitude of wage changes in a continuously varying economic environment. Our model is very robust in terms of the distributional assumptions and is particularly useful in analysing the effect on the exit rate of events occurring after the commencement of a spell, such as incomes policies. Consequently, this approach makes it possible to distinguish between two separate effects of such policies on the wage changes: delay and moderation.

We further analyse the impact of these policies by introducing a self contained policy sub-model in the system of renegotiation probability and realized wage change nexus, in order to find out the existence of consistent effect exerted by different degrees of allowance and enforcement.

We find a significant impact of incomes policy, enforced and non-enforced, on the size as well as the timing of negotiations, which is also confirmed by the results of dynamic simulation.

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Preface

I am grateful to Dr. Stephen Pudney who has provided me with invaluable ideas and advises, from whom I have learnt most of my econometrics throughout my years at the LSE. I am also grateful to professor Christopher Pissarides for his many valuable advises and consistent support. Finally, I would like to thank the ESRC Centre for Labour Economics for supporting this resarch.

To my parents

Chapter 1: The hazard

1.1 Introduction

In many occasions, we are interested in the spell of time, the duration of occupying a certain state under uncertainty. A state may represent a condition of the subject concerned, or it may simply be used to distinguish the time before and after the decision making. A termination of a spell is called failure and the spell is called the failure time. Provided that the timing of a failure occurs randomly, our objective is to study the underlying distribution of such failure times and investigate the effect of explanatory variables on such distribution and their behaviour overtime.

Some doctors may be interested in the time to death or recuperation of the several groups of patients who are treated with different drugs. A factory may wish to conduct a life testing of their products under different stress levels before putting them onto a market. Sociologist may wish to investigate peoples decision of when to get married or to have babies. Economists would like to know factors that influence a spell of unemployment, workers decision to retire, a length of strike or intervals of successive bargaining, if the termination of these spells are not predetermined but decided randomly by the factors under study.

These phenomena are all comprised of a sequence of choices that determine whether to continue or to exit the current state. Hence at any point in time during the spell, there exists a probability of making a transition into the other state. For example, the process of time to death of a patient is modeled as a sequence of probabilities, of death versus continue living, implicitly calculated at every moment up to his death. Time to re-employment is determined by a sequence of choice of whether to take an offer or continue searching over time. The length of a spell is then determined by repeating such choice problem sequentially. Hence, by investigating the effect of potentially interesting factors on the choice probability, we can study their influence on the length of time spent in a certain state, on a failure time.

Before embarking on the analysis, it is important to clarify 3 issues in defining the failure time properly: a definition of time origin, time scale, and a failure. First, it is essential that the starting point of the failure time be made explicit. It is usually the time when a subject enters the state concerned, although due to the nature of data sampling, the time origin may be set some point after the commencement of the spell. In either case, it is assumed that entrance into such a state is exogenous, so that the commencement of duration under study is non-stochastic. Our analysis is conditioned on the information that the subject has entered such a state, therefore, the failure time is always non-negative and the failure happens at most once for each duration. Secondly, in most cases, the clock time is used (e.g., hours, days weeks or months) for the measurement scale of time. In other cases, such as the experiment of stress level, some measure of cumulative fatigue may be used. As long as the non-negative units are assigned, the important thing is to have consistency in the scale measuring the passage of time throughout the analysis. Once such unit is decided, next step is to determine whether the failure time is better considered to be measured in discrete or in continuous time. This should depend on the nature of the failure time--the intervals of decision making sequences. In practice, it can be every day or month, but when the decision frequency is sufficiently short, the process is often thought of as in a continuous time. Finally, the definition of failure needs to be made explicitly. It is sometimes as clear as "death", although if the failure is the "recovery" from a disease, it is essential to make a precise definition of the term, "recover".

In this chapter, a basic concept of duration analysis is discussed while introducing the recent literature in the field. We confine our interest to the economic application of such analysis which have mostly been done in the field of labour economics.

First, a basic concept of the hazard function, in particular, its relation with the underlying duration distribution is discussed. A basic form of the proportionate hazard model and some useful distributions frequently used for the survival data analysis are introduced. Then, estimation methodology is discussed. Second section extends the basic model to incorporate some complications such as random heterogeneity terms, time varying explanatory variables, and multiple state/spells. Third section introduces actual economic applications of the duration technique. In particular, we will focus our interest on the past study of unemployment spells and strike durations. Fourth section introduces somewhat different approaches to the modeling of survival data, namely, non-parametric and Brownian motion approach.

1.2 The hazard

Let us denote δ to be a non negative random variable representing the length of a failure time, and t to be a time since the start of a spell. Since the entrance to the state concerned (i.e., the start of a failure time) is exogenous, all distributions we deal with are conditioned on these events. This is appropriate since the same conditioning often applies to the nature of sampling in practice. For example, a sample is extracted from the pool of unemployed when studying the duration of unemployment. Hence, the sampling is conditional on the event that those workers are already unemployed. Their spell is terminated by becoming employed or leaving the labour force altogether. From these practical point of view, working with a distribution which is conditioned on the initial state has its advantages.

1.2.1 The hazard

(i) Continuous time

Assume that the duration δ has a cumulative distribution function, F, and a density function, f, so that:

$$\Pr(\ \delta=t\ ) = f(\ \delta\,|\,x;\theta\ )$$
$$\Pr(\ \delta\leq\Delta\ ) = F(\ \Delta\,|\,x;\theta\ )$$

where x is a vector of explanatory variables and θ represents a vector of parameters. The survivor function is defined as:

$$S(\ \Delta,x\ ) = 1 - F(\ \Delta,x\ )$$

and it gives the probability of a spell lasting for at least as long as Δ. Hence, $S(0)=1$. Now, given the knowledge of the distribution function of durations, it is possible to derive a hazard function, which is a probability of exiting the current state at t provided the spell has lasted at least for t.

$$h(x,t) = \lim_{dt\to 0} \Pr(\delta\in(t,t+dt)\,|\,\delta\geq t) = \lim_{dt\to 0} \frac{\Pr(\ t\leq\delta\leq t+dt\,|\,t\leq\delta\ )}{dt}$$

$$= \lim_{dt\to 0} \frac{F(t+dt) - F(t)}{dt} \; \frac{1}{1 - F(t)}$$

$$= \frac{f(t\,|\,x)}{1 - F(t\,|\,x)} \qquad (1\text{-}1\text{-}1)$$

It is sometimes called the local exit rate, or the escape rate. In practice, the analysis is usually carried out by specifying the hazard function, from which the underlying conditional probability distribution, $f(t|x)$ or $F(t|x)$, are derived and they are then inserted to the likelihood function for estimation. A reason why the analysis tends to start with the specification of the hazard is because the hazard is more intuitive in terms of economic theory compared to the distribution function itself. Using the relation (1-1-1), it is possible to deduce the distribution functions in terms of h(.).

First, differentiate the log of survivor function with respect to t:

$$\frac{d \log S(t|x)}{dt} = \frac{d \log (1-F(t|x))}{dt}$$
$$= \frac{- f(t|x)}{1 - F(t|x)} = - h(t,x)$$

Integrate both sides over the range, Δ, say, gives:

$$\log (1 - F(\Delta|x)) = - \int_0^{\Delta} h(t,x)\, dt$$

Given that $F(0|x)=0$, by definition, it follows that the cumulative distribution function of a duration is expressed in terms of h(.) as follows:

$$F_{\delta} (\Delta|x) = 1 - \exp \left[- \int_0^{\Delta} h(t|x)\, dt \right]$$

and also:

$$f_{\delta} (t|x) = h(t,x) \left\{ \exp \left[- \int_0^{\Delta} h(u,x)\, du \right] \right\}$$
$$= h(t,x) \left[1 - F(t|x) \right]$$

Denoting integrated hazard, $\int_0^t h(u,x)\, du$ as $\Lambda(t,x)$, the survivor function can be written as:

$$S(t|x) = \exp (-\Lambda(t,x))$$

Non defectiveness is required for the duration to have a proper distribution so that the density integrates to 1. This is satisfied when $S(\infty)=0$, in other words:

$$\lim_{t\to\infty} \int_0^t h(u,x)\, du = \infty$$

This is called the admissibility condition. In practice, there are some cases when this condition are not necessarily met. Some people may stay being unemployed forever. In such cases, defective duration distribution arises since $F(\infty)<1$.

(ii) Discrete time

So far, δ had continuous distributions. However, there are cases where δ is better considered to be a discrete random variable, taking values $\tau_1<\tau_2<\ldots$ For example, a strike duration maybe recorded in the units of days if unions' decision is made daily whether to strike one more day. Then the underlying density function of duration is:

$$\Pr(\delta=\tau_j)= f_j$$

And its survivor function is:

$$\Pr(\delta<\tau_j) = S(\tau_j) = \sum_{x\geq\tau_j} f(x)$$

The hazard at τ_j is the probability of exit at $t=\tau_j$ given that the spell has not ended for the last τ_j periods, and is written as:

$$h_j = \Pr(\delta=\tau_j \mid \delta\geq\tau_j)$$

$$= \frac{f_j}{S(\tau_j)} \qquad \text{for } j = 1,2,\ldots.$$

And the underlying distributions in terms of the hazard is:

$$S(\tau_j) = \prod_{j \mid x_j<\tau} (1-h_j) \qquad (1\text{-}1\text{-}2)$$

and $$f_j = h_j \prod_{1}^{j-1} (1-h_j)$$

1.2.2 Forms of the Hazard

How the hazard is affected by the length of elapsed spell is called a duration dependence. When $dh(t)/dt=0$, the hazard is constant for any elapsed spell, t. This is the case of no duration dependence, and its underlying duration process is called stationary. The Markovian property is satisfied if the hazard is independent of its past history given the entrance to the current state, departure from which is under study. Out of which a stationary Markovian process is represented by the exponential duration distribution where its density function is $f(t)=\mu\exp(-\mu t)$. The corresponding hazard rate is $h(t)=\mu$ for all t. In this case, a probability to exit at any point in time during the spell does not change no matter how long the spell has been going on: the memory-less property. On the other hand, if the length of current duration affect the probability of exiting such state, the process is called semi-Markovian in a sense that the duration distribution is independent of the past durations spent in the other states but depends on the current elapsed duration. In this case, if $dh(t)/dt>0$, it is said to have a positive duration dependence and if $dh(t)/dt<0$, negative duration dependence. The existence as well as the form of such dependence often becomes a central issue in the duration problems in economics. For example, in the study of unemployment spell, the positive duration dependence will be observed if the longer the workers are unemployed, the lower the level of their reservation wage (i.e., the minimum level of wage they are willing to take on a job) Although if it is also assumed that the job offer rate decreases with a length of unemployment, this will have a counter effect on the hazard, consequently making the direction of duration dependence ambiguous. Note that in this particular problem, the elapsed duration affected the hazard through their impact on the reservation wages and job offer rate. The duration dependence can be generated from t directly, and also through the time path of explanatory variables, x(t), during the spell, when x are time varying. This creates some identification complications as we shall discuss later, and for the moment, we continue to assume that x is constant throughout the spell.

The most frequently used simplest formulation of the hazard is the proportionate form:

$$h(x,t;\theta)=h_1(x;\beta)\ h_0(t;\alpha)$$

$h_1(x;\beta)$ depends only on the explanatory variables and not on time, while $h_0(t;\alpha)$ depends on time and is known as the baseline hazard. This is the value of the hazard corresponding to $h_1(x)=1$. (β,α) constitute unknown parameter, θ. In this formulation, the duration dependence is determined entirely by $h_0(t)$, hence the explanatory variables can influence the scale of the hazard but not the form of its dependence on time. Therefore it is easy to discriminate the effects of time and other explanatory variables on the hazard. A form that $h_1(x)$ usually takes is $\exp(\beta'x)$, although linear polynomials or logistics can also be used. Exponential form is flexible and it naturally satisfies the basic requirement that the term $h_1(x)$, entering multiplicatively in this probability, is always non-negative. Hence:

$$h(t,x;\theta) = \exp(\beta'x)\, h_0(t)$$

Below, we list the examples of duration distributions often used in practice which underlies $h_0(t)$. They vary mainly in the forms of duration dependence they can accommodate. They also differ in the degree of simplicity of distributional forms.

(i) Exponential distribution

$$f(t) = \alpha\exp(-\alpha t)$$
$$h(t) = \alpha$$

When the underlying distribution of duration takes the exponential form, the transition probability is governed by the Poisson arrivals with rate α. This is to say, the probability of a transition from the current state within a small interval between t and t+dt is $\alpha dt+o(dt)$, where $\lim_{dt\to 0}\frac{o(dt)}{dt}=0$ indicates that a probability of making more than one transition is negligible. Hence, the probability of having n transitions (i.e., arrivals) in t periods is a familiar Poisson probability:

$$\frac{(\alpha t)^n}{n!}\exp(-\alpha t)$$

And obviously, α is the hazard rate which is constant throughout the spell reflecting the memory less property.

(ii)Weibull distribution

$$f(t) = \tau\alpha t^{\alpha-1}\exp(-\tau t^{\alpha})$$

$$h(t) = \tau\alpha t^{\alpha-1}$$

The Weibull distribution can be considered as an exponential distribution on a rescaled time axis, t^{α}, with parameter τ, where τ may depend on the explanatory variables, x. This distribution can bear either a negative or positive duration dependence according to the value of α. If α is smaller than 1, hazard is a monotonically increasing function of t. When $\alpha=1$, this distribution reduces to the exponential case with no duration dependence. Admissibility condition is satisfied as long as $\alpha>0$.

(iii)Gamma distribution

$$f(t) = \frac{\alpha(\alpha t)^{k-1}\exp(-\alpha t)}{\Gamma(k)} \qquad k>0$$

When $k=2$, this implies a distribution of time taken for the alternative Poisson arrival with rate α. It includes, as special cases, the exponential ($k=1$) and the log-normal (a limiting case as $k\to\infty$), and hence, it incorporates a variety of shapes. Gamma distribution is a representative distribution for any positive continuous variates, although a computation of the survivor function is difficult since it involves the incomplete gamma integral.

(iv)Gompertz-Makeham distribution

$$S(t) = \exp\left[-\gamma_0\,\delta + \frac{\gamma_1}{\gamma_2}[\,e^{\gamma_1} - 1\,]\right]$$
$$h(t) = \gamma_0 + \gamma_1\exp(\gamma_2 t)$$

When $\gamma_0=0$, it is called the Gompertz distribution. The hazard is an exponential function of the failure time, hence, changes more rapidly than, say, the Weibull distribution hazard. Positive duration dependence is seen for $\gamma_2>0$ and $\gamma_1>0$ or $\gamma_2<0$ and $\gamma_1<0$, and negative otherwise.

(v) Pareto distribution

This distribution can be interpreted as a distribution of exponential survival time whose rate of arrival differs between the individuals. Its density conditional on the arrival rate, γ, is:

$$f(t|\gamma) = \gamma \exp(-\gamma t)$$

Unconditional density of t is then:

$$f(t) = \int_0^\infty \gamma\exp(-\gamma t)\, f_\gamma(u)\, du$$

Convenient distribution for $f_\gamma(.)$ is the Gamma with mean α and index κ.

$$f_\gamma(u) = \frac{(\kappa/\alpha)(\kappa u/\alpha)^{\kappa\kappa\kappa-1} e^{-\kappa u/\alpha}}{\Gamma(\kappa)} \qquad \kappa>0$$

In such a case, unconditional density of duration becomes Pareto distribution with a density of a following form.

$$f(t) = \frac{\kappa\,(\kappa/\alpha)^{\kappa}}{(t + k/\alpha)^{\kappa+1}}$$

Hence, the corresponding hazard function is:

$$h(t) = \frac{\kappa}{(t + \kappa/\alpha)}$$

This distribution tends to the exponential as $\kappa\to\infty$, and has a very long tail for a small value of κ. It possesses a negative duration dependence.

(vi)Log-normal distribution

When log of duration is normally distributed with mean $-\log\mu$ and variance σ^2, duration, δ has a density:

$$f_\delta(t) = \frac{1}{\sqrt{2\pi}\sigma t} \exp\left[- \frac{(\log t\mu)^2}{2\sigma^2} \right]$$

and the survivor function is: $S(t)=1-\Phi((\log t\mu)/\sigma)$, where $\Phi(w)$ is the incomplete normal integral, over $-\infty$ to w. The hazard is f(t)/S(t), and it takes value 0 at t=0, and increases with t initially but becomes non-monotonic and tend to tail off to zero as t becomes large. The thickness of the tail and whether the maximum occurs depends on the value of σ. Application is difficult if censored observation are involved in the data. Also, this distribution is very sensitive to the small failure times.

(vii)Log logistic distribution

When w has a logistic distribution, so that:

$$f(w) = \frac{e^{w}}{(1 + e^{w})^{2}}$$

the density function of δ, where $\log\delta = \mu+\sigma w$, is:

$$f_{\delta}(t) = \frac{\lambda\gamma(\lambda t)^{\gamma-1}}{[1+(\lambda t)^{r}]^{2}}$$

where $\lambda=e^{-\mu}$ and $\gamma=1/\sigma$. This distribution provides a good approximation to the log-normal distribution, and the form of the hazard can be written more explicitly:

$$h(t) = \frac{\gamma t^{\gamma-1}\lambda^{\gamma}}{1 + (t\lambda)^{\gamma}}$$

For $\gamma>1$, the hazard has a single maximum, and for $\gamma<1$, it has a negative duration dependence.

1.2.3 Parametric Estimation

(i)Maximum Likelihood

Suppose that the underlying distribution of duration is known up to a vector of parameter, θ. Then the estimation of the unknown parameters are usually done by the maximum likelihood. We continue assuming that the entry to the current state is non-stochastic.

Depending on the nature of sampling, data may contain completed spells in addition to uncompleted observations. Uncompleted spells recorded may have ambiguous starting point (i.e., left censored) or ambiguous ending (i.e., right censored). Left censored observations have unknown starting points. This occurs for the spells that have already started at the time of sampling, hence, how long the spell had been going on prior to the sampling date is not known. This is common in a survey data where the information is based on several interviews conducted at arbitrary intervals. On the other hand, the right censored observations lack information on the termination points. We know as much as how long the spell has at least lasted, but the exact duration is again unknown. In a survey data, the right censoring occurs for those whose spell did not terminate at a time of the last interview.

In general, the likelihood function to be maximized is:

$$L = \prod_{i\in A} f_\delta(\delta_i) \prod_{j\in B} [1 - F_\delta(\delta_j)] \prod_{k\in C} \frac{f_\delta(\delta_k + s_k)}{1 - F_\delta(\delta_k)} \prod_{l\in D} \frac{1-F_\delta(s_l + \delta_l)}{1 - F_\delta(\delta_l)} \qquad \text{(1-2-1)}$$

Set A and B contain observations whose likelihood is the unconditional probabilities of δ. Set A contains observations on completed durations. Set B contains observations that are truncated from above, so that the only information available on such observation is that its spell has lasted at least as long as δ_j. Hence, the probability of observing such event is:

$$Pr(\delta > \delta_j) = 1 - F(\delta_j)$$

Set C and D both contain observations which are truncated from below, hence, they are the conditioned observations. Consider the interview survey which was conducted at 2 different times. The spell is known to have started δ_k (or δ_l in set D) prior to the time the first survey took place. Hence, this particular observation could never take a value less than δ_k. Hence, the likelihood for these observations have to be conditioned on the event, $t > \delta_k$, which occurs with probability, $1-F(\delta_k)$. Set C contains the observations on failure time which are truncated from below but whose complete length of spell is known. This is the case if the termination of event took place prior to the second survey. If the spell has lasted for s_k since the first survey, its failure time in total is $s_k + \delta_k$. Nonetheless, the probability of observing such event is conditional on $Pr(t > \delta_k)$ therefore, $f_\delta(\delta_k + s_k)/(1-F_\delta(\delta_k))$. Set D, on the other hand, contains observations which are truncated from below as well as from above. Conditioned on the event, $t > \delta_l$, all we know is that the spell has lasted at least for another s_l periods. In total, it has lasted at least as long as $s_l + \delta_l$. This is the case if the spell has not ended at a time of the second interview. The likelihood of observing such observation is, $(1-F_\delta(s_l + \delta_l))/(1-F_\delta(\delta_l))$.

Difficulties arise when there are observations on spells whose starting points are totally unknown. Flinn and Heckman (1982) have proposed a technique of deriving a likelihood contribution of the observed part of duration when the starting point is unknown. Suppose the observed part of a failure time is d. We consider, for such partly observed samples, the probability of observing δ ($\delta \geq d$) as an actual duration. A probability of sampling a duration, δ, for an individual with x is $f(\delta|x;\theta)$ at every point in time. Since this individual is in this state for a period δ, the total probability of sampling an individual with duration δ is $\delta f(\delta|x;\theta)$. In order to make this into a proper density so that it integrates to 1, we have to divide it by $E(\delta|x;\theta)$, hence it becomes: $\delta f(\delta|x;\theta)/E(\delta|x;\theta)$. On the other hand, given that the true duration is δ, the probability of observing d is uniform and its simply, $1/\delta$. From these two probabilities, form the joint density between δ and d, and integrate δ out over the range (d,∞) to yield a marginal probability for the observed d. This will be the contribution of the partially observed duration, d, in the likelihood function. As can be seen, it involves an integration of the distribution function, $f(\delta|x;\theta)$, hence, a numerical complication is unavoidable even for the distribution with a simple closed form.

In the straightforward case when all the observation are uncensored or right censored, full likelihood can be simply written in terms of the hazard as:

$$\log L = \sum_{i \in U} \ln h(t_i) - \sum_{j \in A} \Lambda(t_j)$$

where $\Lambda(t)$ is the integrated hazard over $[0,t]$ interval, U is the set of uncensored observations and A is the set of whole observations. The likelihood function is then maximized with respect to the unknown parameters which appear in the function in various ways according to the assumed functional form of the hazard.

Under the regularity conditions for f, $\hat{\theta}$ is consistent and $\sqrt{N}(\hat{\theta}-\theta)$ is asymptotically normally distributed with mean zero and variance:

$$-\left[E\left(\frac{1}{n} \frac{\partial^2 \log L}{\partial\theta\partial\theta'} \right)^{-1} \right]$$

which can be consistently estimated by:

$$-\left[\frac{1}{n} \frac{\partial^2 \log L}{\partial\theta\partial\theta'} \right]^{-1}$$

When the expression for the derivatives are not in closed form, numerical maximization should be adopted to obtain the maximum likelihood estimates. When the log likelihood is concave, the Newton's method works well. Cares must be taken to ensure that the global maximum is to be attained.

The maximum likelihood estimation can accommodate censored or truncated data as well as the time varying explanatory variables. But the entire functional form of the hazard, in particular, that of the baseline hazard is essential in constructing a likelihood function to be estimated parametrically. Moreover, when a random heterogeneity term is introduced, the density function for such term has to be also assumed so that the corresponding marginal density can be calculated. A possibility of correlation between the explanatory variables and the heterogeneity term cannot be easily handled. Problem often arises since the economic theory rarely provides any information on the form of such density function, and yet, such assumption is crucial to the estimation. The parameter estimates become inconsistent if they are wrong.

On the other hand, the semi-parametric estimation technique avoids such risks. One can build a likelihood function and derive consistent estimates without making assumptions over the parametric forms of the heterogeneity term or the baseline hazard. There, the baseline hazard and/or the heterogeneity distributions are treated as discrete functions over the elapsed spell, whose discrete mass points are estimated together with the parameters of explanatory variables. When the number of distinct duration spells is large, a number of parameters to be estimated should increase accordingly to avoid inconsistency, although this reduces efficiency. In practice, there are a few problems in obtaining a global maximum for such a semi-parametric likelihood. Another semi-parametric method is called a partial likelihood where the analysis is conditioned on the unknown form of the baseline hazard, hence, avoids making assumption over h_0. These methodologies are discussed in detail in the next section.

(ii) Partial Likelihood

Under this proportionate hazard formulation, it is possible to draw an inference about α without the knowledge of a functional form of the baseline hazard, h_0. Ranks rather than values of the observed failure times carry information which is enough to construct the likelihood. Hence, the parametric form of the baseline hazard doesn't have to be assumed.

Order the observed failure times as $t_1 < t_2 < t_3 < \ldots t_n$. If the j-th spell is of length t_i, it implies that the j-th spell lasted for t_i, which is the i-th shortest amongst all the observations. j-th observation faces explanatory variables x_j, so at the end of j-th spell, $x_j(t_i)$. Then the conditional probability that the j-th spell is of length t_i given the entire history is:

$$h(t_i,x_j(t_i)) \;/ \sum_{k \in R_i} h(t_i,x_k(t_i)) \qquad (1\text{-}2\text{-}2)$$

where R_i represents the risk set just prior to t_i. This set includes observations on durations which are longer than or equal to t_i and not yet terminated at spell duration t_i. (1-2-2) is a probability that the j-th spell ends at t_i given that one of the spells in the risk set R_i does end at t_i. In a sense, this probability is conditional on the entire history of all the failures and censoring before t_i and the values of all the explanatory variables up to and including time t_i, but without the knowledge that the j-th observation fails at t_i. The baseline hazard $h_0(t)$ cancels out because of the multiplicative form assumed by the proportionate hazard specification. Hence, a contribution to the likelihood of j-th duration observation which is the i-th shortest, is, $h_l(x_j)/\sum_{k \in R_i} h_l(x_k)$. Then the joint distribution of the observed set of failure times, $t_1,t_2,\ldots t_n$, can be obtained using the chain rules. The log of partial likelihood to be maximized is:

$$\log L = \sum_{i=1}^{n} \left[\ln h_l(x_j(t_i)) - \ln \sum_{k \in R_i} h_l(x_k(t_i)) \right]$$

Censored observation can easily be incorporated in the partial likelihood framework. If an observation is censored between duration t_i and t_{i+1}, its contribution to the likelihood is only towards the risk set, the denominator of individual's conditional probability for $t_1,\ldots t_{i+1}$. Ties (when there exists more than one duration observations with the same spell length) will reduce efficiency of the estimator, but can still be dealt with in a similar manner. Likelihood contributions of the tie observations will have the common denominators with the different numerators associated with each observations, possibly, with a different set of explanatory variables. A multiplicative random heterogeneity term can also be incorporated, although it will involve multiple integrations of order R_i, a dimension of the risk set, and such computation will be tremendously messy. A non-stationary duration distribution, which depends on a calendar time, is also easily accommodated. Denoting τ_j as the calendar time at which the j-th spell has started, $x_j(t)$ should be re-written as $x_j(t+\tau_j)$. Obviously, this estimation method is not useful if one is interested in the direction of duration dependence.

Rank information only is enough to make an inference about β can be seen by considering any one to one strictly increasing transformation. Under such a transformation, $h_0(t)$ will not be identified but the rank information on the observation on durations is not altered, which identifies the value of β. In this sense, the parameters of explanatory variables will not be identified if they are closely correlated with the elapsed duration. This method is sometimes called the "marginal likelihood" although it only means "marginal" with respect to the distribution function when the explanatory variables are not time varying.

(iii) Regression method

Under the proportionate hazard formulation, a log-linear regression model of the observed duration can usually be constructed because of the multiplicative explanatory factors in the hazard. Given the hazard function, the underlying density of duration is:

$$f_\delta(t|x) = h(t,x) \exp(-\Lambda)$$

where Λ denotes the integrated hazard, which is $\int_0^\delta h(u,x)du$. Transformation of variable from δ to Λ will give the density of Λ as:

$$f_\Lambda(i|x) = \exp(-\Lambda)$$

since the Jacobian of transformation is $|\partial\delta/\partial\Lambda| = h(t,x)$. This is an unit exponential distribution. Moreover, transforming Λ to $\varepsilon = -\log\Lambda$ would give:

$$f_\varepsilon(\varepsilon/x) \quad = e^{-\varepsilon}\exp(-e^{-\varepsilon}) \quad = \exp(-\varepsilon-e^{-\varepsilon})$$

This ε is the log of exponentially distributed random variable, and it follows a type I extreme value distribution. For the proportionate hazard form:

$$h(t,x;\theta) = \exp(\beta'x)\ h_0(t)$$

integrate this hazard over the range $(0,\delta)$ and take a logarithm to derive ε:

$$\varepsilon = -\log\Lambda = \beta' x + \log \int_0^{\delta} h_0(t)dt$$

Denoting the last term as $\ln \Lambda_0(\delta)$, this can be rewritten as:

$$-\ln\Lambda_0(\delta_i) = \beta' x_i + \varepsilon_i$$

for the i-th observation. This can be regarded as a regression equation with the error term ε_i which follows an extreme value distribution. This term has a fixed distribution and does not depend on β, although in practice, there is an additional regression residual from the least square estimation, that involves an error component because of the assumption we made over the form of the hazard. Variance of such least square residuals can be compared with the predicted value to test the existence of omitted regressors, for example. It is of an interest to draw inferences of parameters conditioned on this ancillary[1] statistics, but such attempt have not been made so far.

In the case of the constant hazard where $h(t,x)=he^{\beta' x}$, $\ln\Lambda_0(\delta_i)=\ln h+\ln\delta_i$, hence:

$$\log\delta_i = \varepsilon_i - \beta' x_i - \log h$$

Under the Weibull distribution, where $h(t,x)=\tau\alpha t^{\alpha-1}e^{\beta' x}$,

$$\log\delta_i = a_0 + a_1' x_i + \varepsilon_i$$

where $a_0=-\log\tau/\alpha$ and $a_1=-\beta/\alpha$ and again, ε_i follows extreme value distribution.

The advantage of this least square method is that it is simple and is distribution free, in a sense that no assumption is required for the distribution of the error term. And the correlation between the error term and the explanatory variables are easily incorporated. In addition to the examples above, log normal or log logistic distribution for the baseline hazard with time invariant explanatory variables can also be estimated by the simple regression of log of duration. When the explanatory variables include time-% varying series, dependent variable becomes more complex than merely the log of duration. Still, the regression model for the log of duration serves as a specification test as a preliminary analysis. If the model is correctly specified, the error term should possess the moments predicted by the theory. Drawbacks are that there is no simple way of handling severely censored or truncated data. Tobit type method can be applied based on the extreme value distribution of ε, though resulting non-linear computation will no longer be simple. Also, because of the explosive nature of logarithm near 0, it is not suitable for the data that contains observations on short durations which are contaminated with measurement errors, or if the study is aimed particularly at short term durations. Also, as can be seen from the example of Weibull distribution, it is not possible to identify β and α. Hence, if one assumes a certain distributional form for the baseline hazard, interesting parameters representing such distribution, particularly the duration dependence may not be estimated separately in this regression method. Cox and Oaks (1984) states that this ordinary least square regression's asymptotic efficiency relative to the maximum likelihood estimation depends on the mean values of ancillary statistic (i.e., standardized residuals, ε).

1.3 Generalization of a model

1.3.1 Random Heterogeneity

Consider a possibility that there are additional variations across individuals apart from those explained by the observable explanatory variables, or suppose that we have omitted some factors which influence the hazard. Up to now, we have been excluding these possibilities which is much too restrictive in practice. In general, unless we know a priori that individuals are homogeneous, it is necessary to take into account of the population variation apart from the variation caused by the observables. We denote an unobservable term representing a random heterogeneity as v. Its distribution function f_V is called a mixing distribution, and is defined over a range, R_V. The hazard and underlying duration distribution are then all conditioned on v so that they are now written as, $h(t|x,v;\theta)$ and $f_\delta(t|x,v)$, respectively. In order to obtain the unconditional distribution of duration, we need to integrate out for v over the range, R_V:

$$\begin{aligned} F(\delta|x) &= \int_{R_v} F(\delta|x,v)\, f_v(v)\, dv \\ &= 1 - \int_{R_v} \exp\left[- \int_0^\delta h(t|x,v)dt\right] f_v(v)\, dv \end{aligned}$$

$$f(t|x) = \int_{R_v} h(t|x,v) \left[1 - F(t|x,v)\right] f_v(v)\, dv$$

The unconditional hazard is derived as before:

$$h(t,x) = \frac{f(t|x)}{1 - F(t|x)}$$

In practice, a convenient pair of distributions are selected for the mixing distribution and the underlying duration distribution conditional on v, so as to come up with a simple closed form for the corresponding unconditional duration distribution. This is often called a reduced form method. Lancaster (1979) utilized Weibull for the conditional duration distribution and Gamma for v. Kennan (1985) used Logit for the hazard and Beta for v. In his study, given the assumed Logit distribution for the hazard, Beta specification for the mixing distribution was convenient for the following reasons: (1)its range lies in the unit interval, (2)its likelihood had a closed form, (3)it has only two parameters, and (4)the shape and location of the hazard could vary flexibly with the explanatory variables (can be uni-modal, U-shape, J-shape or uniform). The other commonly used combination of distributions for v and t are beta-logistic used by Heckman and Wills (1977).

What happens if we ignore such a heterogeneous term altogether? Lancaster (1979) found out that the spurious declining hazard is observed as a result of the omitted variables. If the individual characteristics are not wholly taken into account, this can falsely exacerbate the shortness or the lengthiness of a spell. An individual with a short spell will contribute to raise the hazard above its true value, and those with long spells will lower the hazard below its true level, both leading to an apparent negative duration dependence. Consider, for example, a study of a strike duration. Suppose v_n measures a degree of militancy of the n-th union. Then, it is natural to assume that the higher v_n, the longer the strike is likely to last (i.e., lower the hazard). If we fail to take into account of v_n, the estimated hazard will be higher than the true value for short strikes due to the groups with low level of v_n, and will be lower than the true value of the hazard for longer strikes due to high proportion of remained groups with high level of v_n. On the whole, the effect of omitting v_n alone would make it look as if the hazard declines with the length of strike.

In practice, it is often possible to deduce a form of the hazard from the economic theory while that of the mixing distribution is not. And yet, the estimates of the structural parameters are very sensitive to the specification of the hazard and/or the mixing distribution. Heckman and Singer (1984) has demonstrated that different directions of duration dependence was inferred from the different assumptions of underlying mixing distributions, and the parameter estimates were also found sensitive to the specifications of duration distributions. Although it does not necessarily lead to the inconsistency of the reduced form parameters (Lancaster (1985)), neglected multiplicative heterogeneity in the hazard gives inconsistent maximum likelihood estimates for the structural parameters. Also, these parameter estimates are sensitive to the specifications assumed for the distributions of the hazard and v_n. Hence the misspecification of the functional forms can lead to several errors, yet, there are rarely any theory which gives guidance to their functional forms. Is the making of such an *ad hoc* functional forms really necessary for the identification?

In general, introduction of the heterogeneity term raises an identification problem. Given only the data on durations, there exists more than one pair of specifications of $f_\delta(t|x,v)$ and $f_v(v)$ for which the same unconditional duration distribution, $F_\delta(t|x)$, applies. This is why these methods outlined above have specified the functional forms of both $f_\delta(t|x,v)$ and $f_v(v)$ in controlling the random heterogeneity. However, there is no need to specify the functional form of $f_v(v)$ in identifying both $h(t|x,v)$ and $f_v(v)$ as long as $f_\delta(t|x,v)$ is known and some conditions are satisfied. Given the functional form for $f_\delta(t|x,v)$, it is possible to uniquely solve for $f_v(v)$, hence, a standard practice of assuming the functional forms for both $f_\delta(t|x,v)$ and $f_v(v)$ over-parameterizes the duration model and may produce inaccurate estimates of the structural parameters. Specifically, in the proportional hazard models with a multiplicative random heterogeneity term:

$$h(t|x,v) = h_1(x;\beta)\, h_0(t)\, v$$

$$\Lambda_0(t) = \int_0^t h_0(u)\, du$$

the conditional hazards and the mixing distributions are identified (i.e., all h_1, Λ_0 and $f_v(v)$ are identified) if there exists at least one exogenous variable[2] taking values along the real line and $E(v)<\infty$ (Elbers and Ridder (1982)). Identification condition given by Heckman and Singer (1984) allows $E(v)=\infty$, which permits wider range of distributions to be a candidate for $f_v(v)$, but has heavier restriction on the form of the baseline hazard. It requires existence of a known constant, c, where $\Lambda_0(t')=c$ for a certain t' for all admissible Λ_0. For most widely used parametric hazard models (a class of Box-Cox hazard or a non-monotonic log logistic models), identification can be achieved without any regressors by specifying a functional form of the hazard up to a finite number of parameters and placing some restrictions on the moments of admissible mixing distribution. In other words, provided that there is a parameterization for $h(t|x,v)$, the non-parametric identification of $f_v(v)$ is achieved without any regressors. This directs us to consider the semi-parametric maximum likelihood estimation where consistent estimates of the structural parameters can be derived from parameterized $f_\delta(t|x,v)$ and non-parametric $f_v(v)$ (see section 1.5.2 (ii)).

1.3.2 Time varying explanatory variables

Observable explanatory variables may themselves vary overtime during the course of a spell. For example, amount of unemployment benefit depends on how long one has already been unemployed at the time of a claim. The hazard is then written as $h(t,x(t)|v)$. The duration distribution now depends not on the value of x at certain point in time but on the entire form of the time path x(t). Now, there is a possibility that the value of x(t) after the start of a spell can influence the hazard. Hence, continuous observations of x are required in the estimation. When observations are only available at finite discrete points, which is usually the case, or if one assumes that x is a discretely changing variable, an arbitrary time path for x has to be considered based on the given observation points. Accordingly, the numerical integration has to be done to derive the density function.

Another problem with the time varying regressors is best illustrated in the proportional hazard formulation, $h(t,x(t)|v) = h_0(t)h_1(v)h_2(x(t))$. The function h_2 (effect of time varying variables) is only identified from h_0 (baseline hazard) if there is sufficient independent variation in x(t), that is to say, if x(t) varies substantially across observations, so that $\ln(h_0)$ and $\ln(h_2)$ are linearly independent, or otherwise, multicollinearity arises. Also, a time path of x must be independent of the parameters of interest. Consider the comparison of the effect on survival of two alternative treatment, for instance. The blood pressure measured during the spell may exert a strong impact on the probability of failure. But the level of blood pressure may be endogenous to the treatment provided.

A method of estimation frequently used for the proportionate hazard model involving the time varying explanatory variables is the partial likelihood. The method was described in the estimation section 1.2.3 (ii). When the covariates don't depend on time, the product of conditional probability was interpreted as "marginal" likelihood of ranks. Now with the continuously varying covariates, the likelihood is no longer inferable of marginal nor conditional probabilities. This likelihood corresponds to the full likelihood without the term that reflects information on the gaps between successive failure times, and is called the partial likelihood. Cox and Oaks (1984) states that the estimator derived from this method will be consistent and asymptotically normally distributed. Consider the information matrix of the full and partial MLE. They converge to var(x(t)) and E(var(x(t)|t)), respectively, for groups defined by the failure time, t. Then, the relation:

$$\mathrm{Var}(x(t)) = E\,\{\mathrm{Var}(x(t)\,|\,t)\} + \mathrm{Var}\,\{E(x(t)\,|\,t)\}$$

shows that the asymptotic efficiency of the partial likelihood relative to the full likelihood will be high if the ratio of the between-spells component of var(x(t)) (i.e., var{E(x(t)|t)}) to the within-spells component (i.e., E{var(x(t)|t)}) is small. It depends on how useful the information provided by the gaps between successive failures would be in determining the coefficient on x. This condition applies unless the coefficient on x is far from 0, censoring depends strongly on x(t), or there are strong time trends in x. A loss in precision is greater under the partial likelihood for a finite sample. Conditional probability that the i-th spell terminated at t_j (j-th shortest of all observations) is now:

$$h(t_j,x_i) \,/ \sum_{k\in R} h(t_j,x_k) = h_2(x_i(t_j)) \,/ \sum_{k\in R_j} h_2(x_k(t_j))$$

where $x_k(t_j)$ denotes the value of explanatory variables facing the k-th spell at duration t_j. Here, the random heterogeneity term is not incorporated, although if multiplicative, it can be included in the partial likelihood framework. Still, it will involve multiple integration over the heterogeneity term of order equivalent to the number in the risk set, R_j.

1.3.3 Competing risks

When defining the failure time, it is important to make sure what the notation of the failure means. So far, this failure was defined so that a spell is terminated by a single cause. For example, unemployment spell ends when one becomes no longer unemployed. Nonetheless, in practice, unemployment spell can be terminated not only by a transition to the employment state, but also by a withdrawal from the labour force. In this case, the hazard has to take into account of the probabilities of a spell terminated by two different causes. As such, when a spell can end in several different ways, the competing risk settings occur.

Consider the case when there are m risks so that m different ways of failures to the current spell. Those risks are independent of each other. Let $h_i(t,x_i;\theta_i)$ denote the probability of exit caused by the i-th risk, with unique explanatory variables, x_i, and the corresponding parameter vector, θ_i. Termination of the current spell will be eventually caused by one of the m sources, hence its hazard rate is:

$$h(t,x;\theta) \quad = \quad \sum_{i=1}^{m} h_i(t,x_i;\theta_i)$$

where $x=\{x_1,x_2,..x_m\}$ and $\theta=\{\theta_1,\theta_2,...\theta_m\}$. That is, the hazard corresponding to the current spell is the sum of instantaneous exit probability due to each risk. In case of the unemployment spell, the hazard rate of exiting such state is the sum of two hazard rates, namely, into employment and out of labor force.

1.3.4 Multiple spells/multivariate failure times

So far, we have been dealing with the failure times of a single transition from one state to another. Under the competing risks framework introduced in the previous section, we allowed a spell to be terminated by several causes. That is to say, an individual, having exited from the current state, have a possibility of entering several different states. Nonetheless, all it mattered was the exit from the current state, and it was not relevant into which state the process continued onto. In this sense, there were only two possible states concerned: currently occupied state (entrance to which is exogenous) and all the other states. Here, we extend this two states with exogenous initial state setting to the one which can deal with the data that contain several spells over several states for each individual over time.

There are people who become unemployed, finds a job and go back to unemployment again sometime later in their life. Data that records employment history of each individuals may contain length of each unemployment spells they have experienced. Consider a study of open ended contract spells. Since they negotiate for the open ended contracts, there is a positive probability of re-contracting at anytime once the contract starts. Hence, there is no unique t where the contract ends with probability one. Data will contain a sequence of length between consecutive bargains for each group over time. And there probably are more than single duration data recorded for each group (i.e., multiple spells). In such cases, it is best to set the starting point of the "duration" as the actual start of each spell, so that there is no left censored observation. Note that the exogeneity of the entrance to each state concerned is still valid for every duration in the sample. In fact, if the sample period is long enough to allow for every individual to record at least one complete spell, there will not be a problem of right censoring since the sampling period can be adjusted accordingly for each individual. These are the case of multiple spells where the data contain several spells for each individual overtime.

In these cases, the vector of explanatory variables may include any function of past history or future variables of the individual. If the hazard of a certain individual depends not only on the current durations but also on its own history of durations, then the process is said to possess a lagged duration dependence. Then, even the semi-Markovian property fails to hold on the duration distribution because the hazard given only the current duration and the hazard conditional on the lagged duration are no longer equivalent. This is applicable, for example, to the case of unemployment history if we assume that people with a history of unemployment spell is more likely to become unemployed again. Or in the case of contract length, it may be that a group who had long contracts in the past tend to have longer contracts in the future. However, in both cases, it may be the underlying character (the fixed effect) of each individual (or group) that gives rise to what seems to be the lagged duration dependence. This difficulty in distinguishing the effect of individual or group heterogeneity from that of the lagged duration dependence makes it all the more important to control for all the idiosyncrasies of the individuals concerned. The lagged duration dependence is incorporated in the joint density by writing it as a product of individual densities each conditioned on its relevant history with the last term being the unconditional density for the initial condition. In order to test for the lagged duration dependence, at least one observation is required for each individual in a sample.

If the number of previous occurrence of event influence the hazard rate out of the current spell, then it is said to have an occurrence dependence. For example, a number of times one has become unemployed in the past may influence his chance of becoming unemployed in the future. In this case, the hazard should be allowed to vary as a number of past occurrence varies (c.f. Flinn and Heckman (1982), Heckman and Borjas (1980)).

On the other hand, the multiple state settings occur when there are more than one state to transfer at a time of exit from the current state. For example, at anytime, an individual can be in 3 different states, namely, employment, unemployment and withdrawal from the labour force. An exit from any one of these state involves a choice of 2 possible states to go into. If there are records of individuals who had experienced these transitions more than twice during the sample period, then this will be the case of multiple state as well as multiple spell (i.e., multivariate) duration data.

Consider a multiple state setting in which there are s possible states. From any particular state, there are potentially (s-1) different states to move into, and those (s-1) states act as competing risks to terminate the current spell. The distinct hazard needs to be specified according to a pair of states involved in the transition. That is to say, altogether, s(s-1) hazards have to be specified, possibly with parameters and explanatory variables unique to each transition. The hazard rate of transition from state i to j can be written as, $h_{ij}(t,x_{ij};\theta_{ij})$ for all $i,j = 1,2,\ldots s$ such that $i \neq j$. The form of possible dependence across the spells of each individual should be carefully studied in this setting.

1.3.5 Marked Failure

Apart from the data on the failure time, when there exists an additional measurement that is observed as a "result" of a failure, this further measurement is called a mark, and the failure time in conjunction with a mark is said to be of the marked failure type. In general, the mark is assumed to be a random variable, and is associated with the failure time, as well as a calendar time. With this kind of failure type, it is important to investigate the relation between the failure time and the mark, at the same time, study the time series structure of a mark.

In a study of contract length, there is a concomitant wage change whenever the contract ends, and in this sense, the new wage rate achieved is a "mark" to the contract duration process. It will be of interest, then, to investigate the relation between the agreed wage change and the length of the contract just terminated, as well as the time path of outside factors during the contract.

1.4 Economic application of duration analysis

Economic applications of duration analysis have mostly been confined to the field of labour economics. The transition between discrete labour market states, for example, from the state of unemployment to that of employment are analyzed by the duration model techniques while the choice of covariates or the specification of distribution were made according to the underlying economic theory. Here, we consider a study of job search, and briefly introduce how to model a sequence of employment choices by using the hazard/duration technique. We consider the discrete time as well as the continuous time setting. We also extend the case to time inhomogeneous environment. Followed by the example of job search model, we introduce some of the recent literature of unemployment spell and strike duration.

Assume that a person who is unemployed can exit its state only through obtaining an employment. And the process of finding a job consist of 2 steps. First of all, one has to receive a job offer. Given a certain job offer, a worker will accept it only if their offering wage is at least as high as their reservation wage (i.e., the minimum level of wage at which the worker is willing to work). Suppose that this worker has been unemployed for a period τ, and at that period, a job offer for such person arises at rate $\lambda(\tau)$. Let distribution of a wage offer be F(w) and a reservation wage $\bar{w}$ to be a function of ones current unemployment spell, τ. $\bar{w}$ maybe derived as a solution to the optimization problem of expected present value of future income stream. Then the probability of getting a job is a product of probabilities that an offer arises and the offered wage being larger than the reservation wage, $\bar{w}(\tau)$, that is:

Pr(accepting a job after τ periods of unemployment spell)

$$= \Pr(\text{offered wage} > \bar{w}(\tau))\ \lambda(\tau)$$

$$= (1 - F_W(\bar{w}(\tau))\ \lambda(\tau) \qquad (1\text{-}4\text{-}1)$$

This probability is conditioned on the fact that this worker has been unemployed at least for τ periods. This is the probability of getting out of unemployment after τ periods of unemployment spell given that he has been unemployed at least for period τ. Hence, it clearly is analogous to the hazard rate.

Consider this setting under the discrete time context, so that δ is a discrete random variable taking values $\tau_1 < \tau_2 < \ldots$ with the associated hazard h_j at the j-th shortest observation on duration, which is τ_j. The analysis based on such a discrete setting is appropriate if the data contains only the discrete failure times, for example, in days. In addition, it is plausible to assume the decision intervals to be defined according to such units, or that the data comes from the grouping of continuous data due to imprecise measurement (Kalbfleisch and Prentice (1973)). Consider the time homogeneous environment so that λ nor the reservation wage, $\overline{w}$, depends on the current duration. A probability of receiving i offers in τ periods is a probability in Bernoulli trial, and that is:

$$\Pr(\text{receive i offers in } \tau \text{ periods}) = \binom{\tau}{i} \lambda^{i}(1-\lambda)^{\tau-i}$$

A probability of staying unemployed at least for period τ is the sum over τ periods ($i=0$ to τ) of probabilities that there are i offers and none of them are accepted. Hence:

$$\begin{aligned}
\Pr(\delta > \tau) &= \sum_{i=0}^{\tau} \Pr(\text{receive i offers}) \\
&\qquad \times \Pr(\text{do not accept offers} \mid \text{receive i offers}) \\
&= \sum_{i=0}^{\tau} \binom{\tau}{i} \lambda^{i}(1-\lambda)^{\tau-i}(F_w(\overline{w}))^{i} \\
&= \left(1 - \lambda(1-F_w(\overline{w}))\right)^{\tau}
\end{aligned}$$

This is analogous to the survival function in discrete time. The corresponding density function is derived as a difference between consecutive discrete distribution function:

$$\begin{aligned}
\Pr(\delta=\tau) &= \Pr(\delta > \tau\text{-}1) - \Pr(\delta > \tau) \\
&= (1 - \lambda(1-F(\overline{w}))^{\tau-1} \; (1 - \lambda(1-F(\overline{w}))^{\tau} \\
&= \lambda(1 - F(\overline{w})) \left[1 - \lambda(1-F(\overline{w}))\right]^{\tau-1} \\
&= h \; (1 - h)^{\tau-1}
\end{aligned}$$

since $h(\tau)=h$ for any τ under the time homogeneous setting.

Let us now allow λ and/or $\overline{w}$ to depend on the current duration, that is, the time inhomogeneous setting. This allows the rate of incoming job offer and the reservation wages to vary according to how long one has already been unemployed. Again, the distribution function of duration is the product of conditional probability describing his state for each period from $t=0$ to $t=\tau$. For the first (τ-1) periods, it was either that the job offer did not come or the job offer did come but their offering wage was lower than his reservation wage at a time. Hence, the probability attached to it is, 1-h(i) for $i=1..(\tau-1)$. Given what had happened in the first $(\tau-1)$ periods, he then accepts an offer at τ-th period, and this occurs with probability $h(\tau)$, where $h(\tau)$ now depends on τ. Therefore, its unconditional probability is:

$$\Pr(\delta=\tau) = \prod_{i=0}^{\tau-1} [1 - h(i)]\, h(\tau) \qquad (1\text{-}4\text{-}2)$$

Unit that measures τ can be months, weeks, or days according to the sequence of decision intervals. But it is not clear if there exists any well defined natural time intervals according to which people make decisions. So, unless there is a clear time units for the their decision sequence, it usually makes more sense if these intervals are made sufficiently small. This leads us to operate the whole process under the continuous time setting.

Instead of considering the arrival of a job offer as the Bernoulli trial, suppose that such event takes place according to the Poisson process. This presumption is appropriate since the Poisson process is the continuous Bernoulli trial with instantaneous probability of success being $\lambda^* = \lim_{n\to 0}\frac{\lambda}{n}$ ($\neq 0$), where λ being the probability of success within an interval, n. To start with, consider once more a time homogeneous environment where a probability of receiving one offer in one period is λ. So, the probability of receiving one offer in an interval $(\tau,\tau+d\tau)$ is $\lambda d\tau + o(d\tau)$, where o() denotes the term with order of magnitude. If we denote g(k,t) as a discrete density function that there are k offers in t period:

$$\begin{aligned} g(i,\tau+d\tau) &= [\Pr(1 \text{ offer in interval } (\tau,\tau+d\tau) \,|\, (i\text{-}1) \text{ offers during last } \tau \text{ periods} \\ &\quad + \Pr(\text{no offer in interval } (\tau,\tau+d\tau) \,|\, i \text{ offers in last } \tau \text{ periods}) \\ &= g(i-1,\tau)(\lambda d\tau + o(d\tau)) + g(i,\tau)(1-\lambda d\tau - o(d\tau)) \end{aligned}$$

$$\begin{aligned} \frac{df(i,\tau)}{d\tau} &= \lim_{\Delta\to 0} \frac{g(i,\tau+\Delta) - g(i,\tau)}{\Delta} \\ &= \lim_{\Delta\to 0} \frac{g(i-1,\tau)\lambda\Delta + g(i,\tau)(1-\lambda\Delta) - g(i,\tau)}{\Delta} \\ &= \lim_{\Delta\to 0} \Big[g(i-1,\tau) - g(i,\tau)\Big]\, \lambda \\ &= -\lambda\, (g(i,\tau) - g(i-1,\tau)) \end{aligned}$$

where all the terms with $o(\Delta)$ disappear since $\lim_{\Delta\to 0}\frac{o(\Delta)}{\Delta}=0$. Solve this differential equation with the condition that $f(i,\tau)=0$ for $i<0$, then:

$$
\begin{aligned}
g(i,\tau) &= \frac{(\lambda\tau)^i}{i!} \exp(-\lambda\tau) \\
&= \text{Prob (i job offers in } \tau \text{ periods)}
\end{aligned}
$$

Hence, a probability of not accepting a job up to period τ is the sum over $i=1$ to infinity of a probability that all of i offers given in τ periods had offered wages lower than $\bar{w}$ (which, at the moment is time invariant). This is equal to the probability of unemployment spell being at least as large as τ, which is intuitively equivalent to the survivor function.

$$
\begin{aligned}
\Pr(\delta>\tau) &= \sum_{j=0}^{\infty}\left[\frac{(\lambda\tau)^j}{j!} \exp(-\lambda\tau)\, F(\bar{w})^j\right] \\
&= \exp(-\lambda\tau) \left[\sum_{j=0}^{\infty}\frac{[\lambda\tau F(\bar{w})]^j}{j!}\right] \\
&= \exp(-\lambda\tau) \exp[\lambda\tau F(\bar{w})] \\
&= \exp\,[-\,(1-F(\bar{w}))\lambda\tau] \\
&= \exp\,(-h\tau)
\end{aligned}
$$

Where the hazard function is the instantaneous probability to exit after τ period, and that is:

$$
h = \lim_{\Delta\to 0} \frac{\lambda\Delta(1-F(\bar{w}))}{\Delta} + \frac{o(\Delta)}{\Delta} = \lambda(1-F(\bar{w}))
$$

as before. Hence, the density of duration is:

$$
\Pr(\delta=\tau) = h \exp(-h\tau)
$$

In the time non-homogeneous setting, as we have dealt with in the discrete case, λ may depend on the current duration. In this case, the arrival of job offer comes from the non-homogeneous Poisson process, where the probability of i offers in the interval $(\tau,\tau+\Delta)$ is:

$$
\Pr(\text{1 offer in } (\tau,\tau+\Delta)) = \lambda(\tau)\Delta + o(\Delta)
$$

$$
\Pr(\text{i offers in } (\tau,\tau+\Delta)) = \frac{\left(\int_0^{\tau}\lambda(t)dt\right)^i}{i!} \exp\left(-\int_0^{\tau}\lambda(t)dt\right)
$$

Hence, the corresponding underlying distributions are:

$$\Pr(\delta > \tau) = \exp\left\{ - \left(\int_0^\tau \lambda(t)dt\right) (1 - F(\overline{w})) \right\}$$
$$= \exp\left(- \int_0^\tau h(t)\, dt\right)$$
$$\Pr(\delta = \tau) = h(\tau) \exp\left(- \int_0^\tau h(t)dt \right)$$

where the hazard also depends on the current duration, τ, hence:

$$h(\tau) = \lambda(t)(1-F(\bar{w}))$$

This can also be seen directly from the discrete duration distribution (1-4-2). As an interval becomes smaller, each h(i) becomes small. Hence, we can approximate $\exp(-h_i) \cong 1-h_i$ so that equation (1-4-2) becomes:

$$\Pr(\text{exit employment after } \tau \text{ periods}) = \exp\left(-\sum_{i=0}^{\tau-1} h(i)\right) h(\tau)$$

In order to transform this into the continuous formulation, replace summation by integration, and narrow the unit of interval, then, the limit of summation, τ-1, becomes close to τ, hence:

$$\Pr(\delta = \tau) = h(\tau) \exp\left(-\int_0^\tau h(t)\, dt\right)$$

This is the density function of the duration under continuous time. In practice, a problem lies in the specification of the hazard. Since economic theory does not imply the exact functional form of these components, they are often left unknown. Usually, the best the economic theory can do is to help list up a set of potentially influential variables to the process. Nonetheless, since such list can not be complete, a need for the heterogeneity term arises. This makes a matter worse since its functional form is utterly unknown. Most of the past literature have hence chosen to assume the functional form of the duration model rather arbitrarily, and resumed the estimation on the basis of such assumption. The non-robustness of this reduced form approach is stated in section (1.3.1).

One of the pioneering application of the duration analysis is the study of unemployment spells by Lancaster (1979). He assumed the proportionate hazard with a multiplicative heterogeneity and the Weibull duration distribution.

$$h(t,v) = v \exp(x'\beta)\alpha t^{\alpha-1}$$

where v is a random heterogeneity term and x is a vector of explanatory variables. He also assumed the distribution of v to follow the Gamma distribution. These assumptions are arbitrary since it was chosen purely on the ground of their simplicity while allowing for a monotonic duration dependence. The relation (1-4-1) suggests to include any variables that may influence (1-4-1) into x in the hazard. But the economic theory of unemployment cannot predict whether a negative or positive duration dependence should prevail. If we suppose that the reservation wage declines as a spell of unemployment persists, this alone will increase the hazard overtime. However, the opposite effect on the hazard will result if we assume the rate of job offer to decline with the current unemployment spell.

His particular concern was with the effect of heterogeneity term on the duration dependence when such term is omitted. He used a survey data on once unskilled, unemployed workers. Time invariant explanatory variables included age at a time of first interview, individual's unemployment history, and replacement ratio which is the ratio between income during unemployment to that of the last job. He formed a likelihood function and estimated the parameters excluding the heterogeneity term. Estimate of α is 0.77 which indicates a decreasing hazard. However, he found that this value increased as he included other explanatory variables which proved significant to the model. This is to say that the negative duration dependence implied by $\alpha=0.77$ from the first estimation was partly due to the omitted variables rather than the true negative duration dependence. This was further confirmed by the value of α estimates, $\hat{\alpha}=0.9$, when the heterogeneity term is included. When the heterogeneity was neglected, it gave rise to the spurious decline in the hazard over time.

Flinn and Heckman (1982) generalized the conditional hazard specification further to include the heterogeneity which is correlated across spells, and in addition, allowed for the time varying exogenous variables. Their hazard is called the Box-Cox conditional hazard since it involves Box-Cox transformation:

$$h(t) = \exp\left(\beta' x_n(t) + cv_n + \mu_1 \frac{t^{\gamma_1}-1}{\gamma_1} + \mu_2 \frac{t^{\gamma_2}-1}{\gamma_2}\right)$$

for $\gamma_1 \neq \gamma_2$, where x(t) is the vector value of time varying covariates at current duration, t, and v is the heterogeneous term. The last two terms approach logt as $\gamma_1 \to 0$ or $\gamma_2 \to 0$, respectively. This model converges to a Weibull duration distribution if $\gamma_1 = 0$ and $\mu_2 = 0$, to a Gompertz if $\mu_2 = 0$ and $\gamma_1 = 1$, and finally, to an exponential if $\mu_1 = \mu_2 = 0$.

Lancaster (1985) tried to determine the joint probability distribution function of unemployment durations and reservation wages by paying attention to the simultaneous relation between the two. The data he used is based on an interview survey conducted in two different years on unemployed people. The recorded wages and unemployment duration had multiple contemporaneous nature in a sense that they contain, for each individual, the length of unemployment spell they have experienced together with the wages in the following job. Also, at the time of interview, workers were asked what level of wages they were willing to start working and the elapsed unemployment spell up to the time of interview. Hence, this data contain reservation wage $\overline{w}(t)$, elapsed unemployment duration at each time of the survey, and the accepted wage. The elapsed spell t is a stochastic function of $\overline{w}$, in the ex-post sense, since longer the employment spell, higher the reservation wage he must have been using. On the other hand, $\overline{w}$ is a decreasing function of the elapsed duration since one would lower his reservation wage the longer he is unemployed. Hence, there are 2 causal relations between $\overline{w}$ and t. At the same time, since the accepted wage in a new job is an increasing function of $\overline{w}$, there must be a similar causal relation between accepted wages and the completed durations. As a result, the estimation was done by deducing the joint distribution of: (i)completed unemployment durations and accepted wages, and (ii)elapsed unemployment durations and reported asking wages. The hazard rate took a form of equation (1-4-1). He assumed the Pareto or Log normal distribution for $(1-F(\overline{w}))$ and posited a monotonic decreasing function of t for the reservation wage function. It was constructed so that it decreases with unemployment spell only until it reaches a certain lower bound and stays flat thereafter. Given this assumption on the functional form of the hazard, he approximated the moment generating function of log(w) and log(t), from which he derived a log linear simultaneous equations to be estimated. This formulation was able to accommodate the specification error or heterogeneity terms into the structural form error terms as long as the errors are uncorrelated with the regressors. He estimated the effects of factors such as age, years of schooling, marital status, disability and locations on this simultaneous relation. The results were consistent with the job search theory. To ensure identification, at least one explanatory variables which were not present in both equations were required. But such variables are hard to come by. Also, this regression method can neither cope with censored observations nor time dependent explanatory variables.

Other applications include those on strike durations. Kennan (1985), under discrete time setting, considered a continuation probability for any individual who is on strike. This is a probability of a strike lasting for one more day given it has already lasted for certain days. He introduced a multiplicative heterogeneous factor, v, into the individual continuation probability:

$$\Pr(\tau|x,v) = v\, p_1(\tau|x)$$

Its survival function is the probability of a strike lasting at least for τ days:

$$\Pr(\tau|x,v) = v^{\tau+1} \prod_{t=0}^{\tau} p_1(t|x)$$

The heterogeneity term was assumed to follow Beta distribution which was then integrated out to give the "aggregate" survival probability:

$$Q(\tau|x) = \int_{R_v} v^{\tau+1} \prod_{t=0}^{\tau} P_1(t|x)\, dF_v(v|x)$$

He found that the strike tend to resolve after 3 months. From the empirical hazard which is based only on the data on duration (see section 1.5.2 (i)), he found that the hazard has a U shape with respect to the strike duration. Hence, he adopt a logit distribution, that can accommodate U-shape, for $P_1(t|x)$. His main objective was to find out the impact of business cycle on the strike activities where the former was proxied by the level of industrial production. By adopting this duration technique, he was able to distinguish the effect of business cycle on the duration of strikes, which he found counter cyclical, and on the frequency of strikes, which was found pro-cyclical. The Beta-logit specification was convenient since its likelihood has a closed form and, in addition, the shape and location of the corresponding hazard specification is flexible enough to vary with the explanatory variables. It was not possible, however, to detect the precise effect of the unobserved heterogeneity on the aggregate hazard since the heterogeneity was not identified from the duration dependence.

Harrison and Stewart (1987) also analysed the strike activities using the duration technique. They adopted the discrete proportional hazards model with a multiplicative heterogeneity term. Their assumed conditional survivor function takes a form:

$$\Lambda(\tau,x|v)=\prod_{t=0}^{d}(1-\lambda_t)^{\exp(x'\beta+v)}$$

where λ_t is a discrete time analogue of the baseline hazard at current duration t, and v is the heterogeneity term. The hazard in a discrete setting is:

$$h(t,x|v) = 1 - \frac{\Lambda(t,x|v)}{\Lambda(t-1,x|v)}$$

Hence their hazard is:

$$h(t,x,v) = 1 - (1-\lambda_t)^{\exp(x\beta+v)}$$

They assumed the entire process to be in discrete time since a strike decision, naturally, is taken on the day to day basis. They assumed exp(v) to follow Gamma distribution and the function of the baseline hazard, $\log(-\log(1-\lambda_t))$, to be represented by a third order polynomial in t to allow for flexibility in the form of duration dependence. They found the strike duration to be counter-cyclical.

Given the proportionate hazard specification with a constant hazard, Stephen Jones (1988) has derived the log-linear relation between the expected value of the elapsed duration and the explanatory variables in his study of unemployment spells. He assumed an offer probability to follow the Pareto distribution, from which, he derived the expected value of elapsed unemployment spell to be log linear in reservation wage and random error term. His data consisted mainly of the censored unemployment spells. Nonetheless, the reservation wage measures were available at the time of each censoring (i.e., at a time of interviewing), which enabled him to study the relation between the *elapsed* unemployment spell and the reservation wages. A possibility of endogeneity of the reservation wage, in his simple linear regression model, was easily dealt with by instrumenting. The instruments were chosen so that they affect the reservation wage but not the exit rate (i.e., search cost) such as the level of benefits. They found that the results from this IV estimation were more significant and also economically sensible than the simple OLS estimates.

A problem inherent in the empirical application of duration analysis is how to specify the functional form of the hazard and their dependence on the observed and unobserved variables in accordance with the economic theory. It has been shown in section 1.3.1 that the results of estimation are very sensitive to the functional form assumed for the hazard as well as the mixing distribution. The identification problems are much easier to deal with if we have better knowledge about the functional form of the economic parameters. Despite this, it is generally difficult to deduce the functional form from the economic theory. As a result, most of the empirical studies have estimated the reduced form model whose underlying distributions of the duration as well as the random heterogeneity is assumed to follow some arbitrary distributions without the justification of the theory.

In order to simplify this inference problem, Lancaster and Chesher (1983) resorted to the design of data. They argued that as long as there exist a survey data that reports the elapsed duration and a mark at such duration, as well as the completed length and a mark at the termination of a spell, we can deduce the interesting parameters without arbitrary assumptions about their parametric forms. Their paper, based on search theory, succeeded in deducing rather than inducing (or estimating) the parameter values (mainly elasticities) from two surveys of unemployed people.

The analysis of job search involves a distribution of offer wage as well as that of arrival rate (see equation (1-4-1)). They used data that contained answers to two survey questions: the reservation wage and the expected wage conditioned on accepting a job. Under the stationary environment with a constant unemployment benefit, a random offering wage, and a random rate of incoming job offer, they maximized the expected present value of future income stream over an infinite horizon to yield the reservation wage equation in terms of a job offer distribution. In addition, there are data on conditional expected wage which is mathematically a function of the job offer distribution. Thus, a function of the unknown offer distribution can be substituted into the reservation wage equation to yield the relation between: (i)reservation wage and benefit level, and (ii)reservation wage and offer probability, without making any assumptions over the underlying distributions. These elasticities were calculated at their mean values, therefore should be robust against measurement errors. They have also deduced the elasticities between: (i)re-employment probability and benefit level, and (ii)re-employment probability and offer probability. For the computation of the latter elasticity, however, it was necessary to assume the form of offer wage distribution above the benefit level. They considered this to follow a member of the Pareto family. In the case of marked failure, one way to resolve the problem with the *ad hoc* reduced form method is to exploit the available data which are potentially informative about the transition process, in particular, the interrelation between the elapsed duration and the marks, as much as possible so as to infer the underlying parameter values from the statistical relations derived from the economic theory.

1.5. Brownian motion and Non-parametric method

1.5.1 Brownian Motion

An alternative approach to the inference problem may be to consider the determination of duration spells to follow some stochastic Markovian process. Then, a task is to derive a first passage time to the absorbing barrier of such process. If a transition of state is regarded as a continuous probabilistic process, a duration of occupying a certain state is a continuous random variable. We can consider this transition probability to be determined by the accumulation of a pressure measure, where a transition takes place when it reaches a certain threshold (i.e., the absorbing barrier) for the first time. This duration since the start of a spell until the transition is called the first passage time.

Lancaster (1972), in his study of strike duration, postulated x(t) to be a barrier representing the difference between wages demanded by the workers and that offered by the management, where t is the elapsed duration of the current strike. Strike duration is determined by bargaining and concessions such that it is terminated whenever the magnitude of x(t) becomes large enough to reach a certain threshold. Beyond the threshold, both sides will compromise to reach an agreement. He assumed this barrier process, $\{x(t);t>0\}$, to follow a simple Brownian motion with drift. This assumption implies that the process, $\{x(t);t>0\}$ satisfies the following properties of stationarity and independent increments: (1)for any interval (t_1,t_2), $E(x(t_2)-x(t_1))=\mu(t_2-t_1)$, (2)$[x(t_1)-x(t_0)]$, $[x(t_2)-x(t_1)]$,... of non-overlapping intervals for any $t_0<t_1<t_2<\ldots$ are independent of each other. Hence, the movement in x(t) occurred on one day is independent of its movement occurred in any of the previous days. The former is the stationarity property which implies that the average level of how close the parties are towards settling a strike (i.e., $x(t_2)-x(t_1)$) in any interval is independent of either t_1 or t_2, or their corresponding calendar time, but is proportionate to the length of interval, t_2-t_1. And this factor of proportionality, μ, is a "drift" indicating the mean proportionate rate per day of a drift towards the end of a strike. where t is the elapsed duration of the current strike. When these two assumptions are satisfied, the process $\{x(t);t>0\}$ is called a Markov process.

In addition, for any interval (t_1,t_2), if $[x(t_2)-x(t_1)]$ is distributed normally with variance $\sigma^2(t_2-t_1)$ and $E(x(t))=0$ for all t bigger than 0, then $\{x(t);t>0\}$ follows a Weiner process (Brownian motion). The process can be rescaled and the origin shifted so that at a start of a strike, $x(0)=0$. And the strike terminates when the absorbing barrier is reached, that is, when $x(t)=1$ for the first time. Given that the process $\{x(t);t>0\}$ follows a Brownian motion with drift, the first passage time of x(t) to the absorbing barrier, $x(t)=1$, is shown to follow Inverse Gaussian distribution. He estimated the influence of explanatory variables over the pressure measure, x(t), by using the maximum likelihood and found the fit of such distribution to be very good, although a large number of short strikes observed in the data failed to be explained. It is likely that these short strikes occurred as predetermined rituals by the union to exert threat, and thus, their spells were determined in a way entirely different from that embodied in the pressure factors, x(t). Jovanovic (1979) used a similar technique in his study of job search.

Consider formulating the labour contract durations in this way. Here, the the length of open-ended contracts are the spell duration to be studied, which is a random variable with a probability distribution characterized by the several economic variables that affect the occurrence of a wage negotiation. We consider a stochastic process of a certain pressure measure in continuous time and space, and a time to negotiation is determined whenever this process reaches a threshold point for the first time since the last negotiation. In other words, whenever the pressure builds up high enough so that the gain from re-negotiation is higher than the cost of doing so. Then, as long as we can consider such pressure factor to follow a Markovian process, it is possible to model the duration between consecutive wage changes as a first passage time of a stochastic process to an absorbing barrier. Specifically, this pressure measure represents a difference between a post tax real wage plus transaction cost, and a target wage, where the latter is derived as an optimal solution to the bargaining problem given the economic environment at a time. Let us write the pressure factor as follows:

$$p(t) = -\left(\frac{W(t)R(t)}{\Pi(t)}\right) - c(t) + x(t)$$

where W(t) : current nominal wage

R(t) : retention ratio

$\Pi(t)$: rate of inflation

c(t) : transaction cost

$x(t)$: target wage

Note that $W(t)$ will take a form of step function over time with a discrete jump at every negotiation. Given this formulation for the barrier, the hazard rate can be derived by specifying the stochastic process which governs $\{p(t);t>0\}$. The hazard, in this case, is the probability of $p(t)$ crossing a barrier in an interval $(t,t+dt)$ provided that there is no previous crossing for the last t periods. Since $R(t)$, $\pi(t)$ and $c(t)$ are exogenous and $W(t)$ being a predetermined variable, they evolve independently of the actions of agents. As a final step, we need to derive a density function of the first passage time and build a likelihood function to estimate the interesting parameters.

Specifically, we assume the pressure variable, $\{p(t);t>0\}$, to follow a simple Brownian motion with drift. Let such drift parameter be μ. One can rescale and shift the origin so as to make $p(t)$ to be 0 when a spell starts and end when the barrier is reached from below to $p(t)=1$ for the first time. For a stochastic process $\{p(t);t>0\}$ to be a Brownian motion, it has to comply with the stationarity, independent increment and normality assumption. This means, for any interval (t_2-t_1), $p(t_2)-p(t_1)$ is normally distributed with mean $\mu(t_2-t_1)$ and variance $\sigma^2(t_2-t_1)$. We also assume $\mu>0$ so that the average target wage increases proportionately with the length of the successive negotiations. In other words, μ is the mean proportionate rate of drift. The normality assumption implies a symmetry about the mean, $\mu(t_2-t_1)$, so that there is as much possibility for $p(t_2)-p(t_1)$ to be negative as it is to be positive. Also, for any interval (t_1,t_2) and (t_3,t_2), $p(t_3)-p(t_1)=[p(t_3)-p(t_2)]-[p(t_2)-p(t_1)]$, hence the left hand side, which is a sum of two independent normal variates, also follows a normal distribution with mean $\mu(t_3-t_1)$ and variance $(t_3-t_1)\sigma^2$. Finally, a joint density function of $x(t_0),\ldots x(t_n)$ can be written as a product of conditional densities with an initial condition that is:

$$\begin{aligned} f(x_0,x_1,..x_n) &= Pr(p(t_n)=x_n|p(t_{n-1})=x_{n-1})\ldots \\ &\qquad \ldots\ldots Pr(p(t_2)=x_2|p(t_0)=x_0)Pr(p(t_0)=x_0) \\ &= f_{t_n-t_{n-1}}(x_n-x_{n-1})\ldots\ldots f_{t_1-t_0}(x_1-x_0).\ f_{t_0}(x_0) \end{aligned}$$

since $Pr(p(t_n)=x_n|p(t_{n-1})=x_{n-1})=Pr(p(t_n)-p(t_{n-1})=x_n-x_{n-1})$ due to the stationary increment property. The estimates of μ and σ^2 can be computed by the maximum likelihood.

A problem with this method essentially is its inflexibility and its strict distributional assumptions. The density for the first passage time is only explicitly known for a very simple form of barrier, such as a constant or a linear function of time. Moreover, as we have seen in the case of contract length, the assumptions of independent and stationary increments are too strong in practice.

1.5.2 Non-parametric and semi-parametric approach

Another solution in avoiding the inference problem is to embark on a non-parametric or a semi-parametric method. Misspecifications of distributional forms not only lead to inefficiency but also to inconsistency of the estimators. By adopting a non-parametric estimator at least in some components of the hazard, we can be less dependent on the correctness of underlying distributional assumptions that are too often arbitrary.

(i) Empirical hazard

The actuarial approach, which is to construct the empirical hazard rate or the empirical survivor function from the data on duration alone, is useful in conjunction with the plotted graphs of survivor function or hazard to assess the goodness of fit, or as a preliminary step to surmise the potential functional forms before embarking on to the parametric method. An example of this is the work of Kennan (1985), in which his decision to use the logit duration distribution was based on the finding that the shape of his empirical hazard is a U-shaped function of the elapsed duration.

Sample survivor function, S(t) is the probability of a spell lasting at least as long as t. This can be estimated by a share of sample points exceeding t out of total number of observations. Suppose that we have a set of n observations in duration among which k ($\leq$n) observations are the completed durations. We can order observations on distinct completed durations so that $\tau_1 < \tau_2 < \tau_3 < \ldots \tau_k$. Then the number of spells neither completed nor censored before τ_j is:

$$n_j = \sum_{i=j}^{k} (m_i + g_i)$$

where, m_i : number of observation censored between τ_i and τ_{i+1}.

g_i : number of spells completed at duration τ_i

In other words, n_j is the number of observations in the risk set at τ_j whose duration has lasted at least as long as τ_j. The corresponding hazard rate at τ_j is a probability of completing a spell at τ_j conditioned on the event that a spell is at least as long as τ_j. In this case, there are n_j observations in a set which satisfies this condition. Therefore, the hazard rate can be estimated as: $\hat{h}(\tau_j)=g_j/n_j$. This is a maximum likelihood estimator of h_j where the log likelihood of this discrete duration observation is:

$$\log L = \Sigma\ [g_j \log(S(\tau_j)-S(\tau_j+0)) + m_j \log S(\tau_j+0)]$$

since a probability of failure at t_j is $S(\tau_j)-S(\tau_j+0)$ where $S(\tau_j+0)=\lim_{dt\to 0} S(\tau_j+dt)$. Recalling the relation between the hazard and the survivor function under the discrete setting:

$$S(\tau_j) = \prod_{i=1}^{j-1}(1-h(\tau_i)), \qquad S(\tau_j+0) = \prod_{i=1}^{j}(1-h(\tau_i))$$

After substitution, the likelihood to be maximized becomes:

$$\begin{aligned}\log L &= \sum_{j=1}^{k} \left\{ g_j \left[\log h_j + \sum_{l=1}^{j-1} \log(1-h_l)\right] + (n_j-g_j) \sum_{l=1}^{j} \log(1-h_l) \right\} \\ &\propto \sum_{j=1}^{k} [g_j \log h_j + (n_j-g_j) \log(1-h_j)]\end{aligned}$$

which yields $\hat{h}(\tau_j) = g_j/n_j$, and the corresponding estimated survivor function is a step function:

$$\hat{S}(\tau_j) = \prod_{i=1}^{j-1} \frac{(n_j - g_j)}{n_j}$$

This is called the Kaplan-Meier or product limit estimator. The corresponding integrated hazard is:

$$\hat{\Lambda}(\tau_j) = \sum_{i<j} \lambda(\tau_i)$$

If the number of distinct failure times $\tau_1,\tau_2,\ldots\ \tau_k$ are fixed and the number of failure at each τ_j $(j=1..k)$ increases as total sample size increases, then the standard asymptotic theory of inference on the maximum likelihood estimator applies. An asymptotic variance estimator of $\log(\hat{S}(\tau))$ is:

$$Var(\log\hat{S}(\tau)) = \sum_{j|\tau_j<\tau} (1-\hat{h}_j)^{-2}\, Var(1-\hat{h}_j)$$

$$= \sum_{j|\tau_j<\tau} \frac{g_j}{n_j(n_j-g_j)}$$

Inductively[3], therefore:

$$Var(\hat{S}(\tau)) = \hat{S}^2(\tau) \sum_{\tau_j<\tau} \frac{g_j}{n_j(n_j-g_j)} \qquad (1\text{-}5\text{-}1)$$

which is known as the Greenwood's formula.

A plot of the estimated integrated hazard or the survivor function against the elapsed duration helps determine the parametric form of the hazard. For example, exponential duration distribution should observe a constant hazard and a linear in duration integrated hazard. From a practical point of view, the plots of integrated hazard are usually smoother, hence, easier to interpret. The estimator is more accurate for the shorter durations where the number of observation in the risk set is high.

Another method of deriving the estimates of empirical hazard/survivor function is the use of a life table. Consider again the ordered observations on completed spells. Divide $[\tau_1,\tau_k]$ into some intervals (not necessarily equal). Then build a life table where the number of censored and completed observations are listed for each interval. Let those intervals be $I_1, I_2, \ldots I_k$. As before, let m_j be a number of censored data within the interval I_j, and g_j be a number of completed duration in the interval I_j. A number of observation in the risk set at t_j is $n_j=\sum_{l\geq j}(g_l+m_l)$. Then the life table estimator of the hazard, h_j, gives conditional probability of a failure during the j-th interval, I_j:

$$\hat{h}_j = \frac{g_j}{n_j-m_j/2}$$

for all n_j except for $n_j=0$, in which case, $\hat{h}_j=1$. m_j in the denominator is divided by 2 by assuming that not all but about half of the n_j observations are at risk throughout the interval, I_j. The corresponding life table estimator of the survivor function at the end of the interval I_j is:

$$\hat{S}(I_j) = \prod_{i=1}^{j}(1-\hat{h}_i)$$

Variances for this estimator is given by replacing n_j by $n_j-m_j/2$ in the

Greenwood's formula, (1-5-1).

This life table method is often used when the actual censoring times are unavailable but g_j's and m_j's are still known for each j-th interval.

(ii) Semi-parametric maximum likelihood

A problem of identifying the conditional duration distribution, $f_\delta(t|v,x)$, and the mixing distribution, $f_v(v)$, out of the observed data, $f_\delta(t|x)$, led the conventional study to specify the parametric distributions of both $f(t|v,x)$ and $f_v(v)$ even though their specification, particularly of $f_v(v)$, was rarely justified by the theory. Yet, different functional forms led to very different estimates, casting importance on the correctness of the parametric specifications to be adopted.

Semi-parametric method came up with a solution to combat these problems. They assume the parametric form of either f_v or $f(t|v,x)$ and use the non-parametric method to infer the nature of the other distribution which is left unknown[4].

The conditions discussed in section 1.3.1 under which $h(t|x,v)$ and $f_v(v)$ are identified secured the identification of $f_v(v)$ non-parametrically. Heckman and Singer (1984) have proved that it is possible to use the observed duration data to consistently estimate both $f_v(v)$ and the structural parameters of the conditional duration distribution, provided that we know the specification of $f(t|v,x)$. Specifically, for the proportionate hazard model in the presence of censoring and time-varying covariates, the non-parametric maximum likelihood estimation (NPMLE) yields consistent estimates when $h_1(x)$ and $h_0(t)$ were specified up to a finite number of parameters and $f_v(v)$ to have a certain behaviour in its tail distribution. They verified the conditions stated by Kiefer and Wolfowitz (1956) that ensure the existence of a consistent estimator of the mixing distribution and the structural parameters. If so, it is possible to derive the estimates of interesting structural parameters by imposing fewer arbitrary a priori identifying assumptions. Then, the conventional model which specifies both the form of $f(t|v,x)$ and $f_v(v)$ is over-parameterized and bears unnecessary risk of making inaccurate estimates.

The maximum likelihood estimator for the structural parameter, θ, and the mixing distribution $f_v(v)$ are derived by solving a problem:

$$\sup_{f_v \in V,\, \theta \in \Theta} \sum_{i=1}^{N} \ln \left(\int_{R_v} f_\delta(t_i|x,v)\, df_v(v) \right)$$

where Θ is the parameter space for θ, and $V=\{f_v:f_v(v)\geq 0$ is non decreasing, right continuous and $\int_v df_v(v)=1\}$. And N is the total number of observations. For a fixed θ, non-parametric MLE of $f_v(v)$ in an identified model is a finite mixture with at most N* points of increase, where N* is the number of distinct values of (t_i,x_i). Hence, the problem reduces to:

$$\sup_{\theta,p_1..p_I,v_1..v_{N^*},N^*\leq N} \sum_{i=1}^{N} \ln \left[\sum_{j=1}^{N^*} f(t_i|x_i,\theta,v_j)\ p_j\right]$$

subject to $\sum_{j=1}^{N^*} p_j=1$ and $p_j\geq 0$ for all j. In a case of general proportionate hazard model with a multiplicative heterogeneity, $h(t|x,v)=h_1(x)h_0(t)v$, transforming a variable from t to t*, say, where $t^*=\int h_1(x)h_0(u)du$ simplifies the likelihood. The density function of t* conditional on v is exponential, hence, $f(t^*|v) = ve^{-t^*v}$. Optimization problem is now:

$$\sup_{\theta,p_1..p_I,v_1..v_{N^*},N^*\leq N} \sum_{i=1}^{N} \ln \left[\sum_{j=1}^{N^*} v_j^d \exp(-t_i^* v_j)P_j\right]$$

where d=1 for the uncensored observation, and subject to $\sum_{j=1}^{N^*} p_j=1$ and $p_j\geq 0$ for all j. Asymptotic standard error cannot be computed from the Hessian of the likelihood since a dimension of the parameter space varies with N. They suggested the use of EM algorithm (Dempster, Laird and Rubin (1977)) to achieve the convergence to a stationary point. For a fixed θ that determines $h_1(x)$ and $h_0(t)$, estimate $f_v(v)$ using above NPMLE. Then, for each estimate of $f_v(v)$, estimates of θ is derived by the parametric maximum likelihood, which will then yield new values of t*. Given the new values of t*, new estimates of will be derived $f_v(v)$. In this way, the process is iterated until the convergence is achieved. There usually are multiple local maxima, hence, the estimation inevitably becomes sensitive to the starting point. Only the global maximum ensures consistency of the estimator. Results of the Monte Carlo experiments suggest that the NPMLE succeeded in estimating the parameters of structural model (i.e., θ) well despite the unreliable estimates for the mixing distribution. The NPMLE could not estimate more than four points of increase for $f_v(v)$, and they were poorly estimated. This suggests a possibility that as long as the mixing distribution is allowed to have a very flexible parametric form (with more than 2 parameters), corresponding MLE estimates of the structural parameters will not be heavily biased. They suggested the use of NPMLE as a test in determining a plausibility of the parametric MLE, although there is no formal test statistics to conclude its plausibility.

In their paper, Heckman and Singer (1984) estimated only the random heterogeneity term non-parametrically while the baseline hazard remained heavily parameterized. In a view that the parametric form of the baseline hazard is often difficult to determine except for their overall shape, it would be convenient if we could also estimate it non-parametrically. The baseline hazard can be considered as a discrete step function. Ideally, a separate parameter should be assigned for every distinct duration observation in a sample. But in practice, a number of steps have to be consulted for the sake of efficiency. Meyer (1986) discusses efficiency comparisons between the semi-parametric hazard and the parametric maximum likelihood estimates. He states that in a situation when the explanatory variables differ more across observations than over time, or when the explanatory variables include a time trend, the semi-parametric method loses little efficiency but insures consistency. The results, he states, are very similar to using the Cox's partial likelihood.

Having seen that those two components of the hazard are possibly estimated non-parametrically, Han and Hausman (1986) estimated both he baseline hazard and the heterogeneity term non-parametrically. They provided the asymptotic normality property of such semi-parametric maximum likelihood by requiring at least one pre-determined variables to be partly continuous so that the model is identified. Nonetheless, in practice, the estimation involving a discrete distribution with many points are difficult, particularly with respect to the mixing distribution estimates, as Meyer (1990) found in his recent paper on unemployment spells.

Meyer (1990) has investigated the effects of the level and the length of unemployment benefits on unemployment durations. The data contained both completed and right censored observations. Since they are all recorded in term of weeks, observations could be regarded as sampled out of discrete time duration distribution. But instead, he regarded them as incomplete observations in continuous time, hence, a contribution of a spell reported to have lasted for t weeks, say, in the likelihood is, $\Pr(t \leq \text{spell} < t+1)$, which is:

$$\{1-\Pr(\text{ spell} \geq t+1 \mid \text{spell} \geq t\ )\} \Pr(\text{ spell} \geq t) \qquad (1\text{-}5\text{-}2)$$

where the latter probability can be written as a product of another conditional distribution:

$$\prod_{k=0}^{t-1} \Pr(\text{spell} \geq k+1 \mid \text{spell} \geq k) \qquad (1\text{-}5\text{-}3)$$

Also, a contribution of the censored duration in the likelihood is merely Pr(spell≥t). Hence, the entire likelihood can be written in terms of the conditional probability of the form such as (1-5-3). This can be written in terms of the hazard as:

$$\exp\left(- \int_{k}^{k+1} \lambda_i(u)\, du \right) \qquad (1\text{-}5\text{-}4)$$

He posited a proportionate hazard form and included the time varying explanatory variables (namely, length until benefits lapse), which, he plausibly assumed to be constant between any unit interval, t and t+1 (i.e., week). The weekly observations on those variables are required for each i from the start till the end of the spell. But by taking the conditional probability form (1-5-3), he managed to avoid the numerical integration of those time varying explanatory variables altogether.

He then assigned different parameter values to every distinct value of duration observation in a sample and estimated together with the parameters of the explanatory variables. This way, the baseline hazard is estimated non-parametrically. Heterogeneity term could be estimated non-parametrically, although in practice, Meyer found it very difficult, hence, resorted to estimating with a gamma heterogeneity. Non parametric estimation of the baseline hazard may sacrifice efficiency, but insures consistency. He found that his estimates are more plausible than that of the totally parametric studies. Note, however, that he was able to allow as many parameters as there are distinct duration observations since the number of such observations was not too large. The effect of ignoring the heterogeneity or imposing a certain distribution function for such term is not known. As Meyer states, estimating the baseline hazard non-parametrically may make the existence of heterogeneity term unimportant. Even though they are theoretically identified, computationally, it seems much too difficult to estimate the heterogeneity as well as the baseline hazard non-parametrically. Another interesting finding is that the peaks found in the empirical hazard were also found in the non-parametrically estimated baseline hazard. These peaks, however, disappeared once he introduced a new explanatory variable that indicates the expected time until the benefits lapse. In this way, the non-parametrically estimated baseline hazard can also serve as a useful diagnostic device.

Footnotes to chapter 1

1. Ancillary means that this is a stochastic process that is exogenous to the individual under study and its marginal probability distribution is independent of the parameters of the hazard (Kalbfleish and Prentice. Chapter 5, 3.1). Nonetheless, the ancillarity here has been obtained from a relationship between the hazard and the duration's density, which presumes the validity of the hazard specification. In this sense, this transformed variable is an ancillary statistics whose distribution depend on the correctness of the parametric form assumed for the hazard.
2. This condition requires the existence of a mean for the mixing distribution.
3. For $\sqrt{n}\,(\hat{\theta}-\theta)\xrightarrow{d}N(0,\sigma^2)$, and $\sqrt{n}(g(\hat{\theta})-g(\theta))\xrightarrow{d}N(0,(\partial g/\partial\theta)^2\,|_{\theta=\hat{\theta}}\,\sigma^2)$, in this case, $g(\hat{s})=\log\hat{s}$, hence, $var(\log(\hat{s}))=var(\hat{s})$, and $(\partial g/\partial\theta)^2=(1/\hat{s})^2 var(\hat{s})$
4. This is not proved theoretically, but it is possible to specify conditions under which $h(t|v)$ is estimable non-parametrically once $f_v(v)$ is parameterized (Heckman and Singer (1984)).

Chapter 2: The microeconomic theory of bargaining

2.1 Introduction

There exists a vast literature that deals with the microeconomic theory of the labour market. Collective bargaining or some quasi collective bargaining such as wages council was the primary mode at least in Britain (80% in manual sector according to New Earning Survey 1978) during the period of our interest, between 1950-75. This enables us to work in the context of the union-firm bargaining when modeling the wage determination process. Note that for their wage determination process to be considered in the context of the firm-union bargaining, it is not necessary for the firms to be completely unionized. This setting is also valid if there exist a competitive sector with an infinitely elastic supply of labour in addition to the unionized sector within the firm. In those cases, size of the union sector becomes relevant in the determination of employment and wages.

First of all, we need to formulate the behaviour of the parties involved in the bargaining. Consider the utility function of a trade union and maximizing function of a firm. Then consider the constraint confronted by each party. Generally, the employees aim for higher pay as well as more jobs while the firm seeks for higher profit which is also a function of the level of employment and pay. Finally, the objective functions of both parties are together or separately maximized with respect to their choice variables subject to their constraints to derive an optimal solution to this bargaining problem.

There exist several competing models for this process, and these can roughly be categorized into four groups: the union monopoly model, the efficient bargain model, the right to manage model and the seniority model. They mainly differ in their assumption as to what constraints the union faces.

2.2 Microeconomic theory of bargaining

Before embarking on the detailed description of these models, let us assume for a moment the following. First, each individual has a decreasing marginal utility of income. Second, we principally consider each firm as a price-taker in its product market, although we also briefly go through the case of monopolistic competition. Under perfect competition, real revenue function is $R=pf(n)/c$, where c is a level of price index and p is a product price, both being given constants. While under monopolistic competition, the firm faces a downward sloping demand curve for its product that is decreasing in competitors (i.e., substitute's) price (s), and increasing in real demand shock (e). Retail price index has a role of deflating the product price towards consumers, but its effect on demand is likely to be ambiguous since c include prices of substitutes as well as complements. Hence, the product demand is a function of p/c, c, s and e. At equilibrium, supply (f(n)) equals demand, and the induced price equation is:

$$p = p(n,s,c,e) \qquad (2\text{-}2\text{-}1)$$

where n is number of workers. The revenue function is also a function of the same arguments, thus will affect the bargaining outcome. As we shall see later, the price taking firm makes a decision based on the product wage while the monopolistic firm bases their decision on the consumption wage. In practice, it may be too restrictive to assume perfect competition which predicts the level of negotiated pay and employment that is independent of the external influences on the product market.

As far as the specification of the utility function of the union is concerned, majority of studies have utilized either the general quasi-concave utility function or the expected utility, utilitarian preference approach. The example to the first type of utility function is the Stone-Geary function which represents a specific form of quasi-concave utility of union in terms of the wages and employment over and above their standard levels, with a weight attached reflecting the relative importance of two objectives. The function is represented as:

$$u = \left(\frac{w}{c} - \frac{\bar{w}}{c}\right)^{\theta}(n-\bar{n})^{1-\theta} \qquad (2\text{-}2\text{-}2)$$

where w denotes wage rate, n denotes employment level, and $\bar{w}$ and $\bar{n}$ denote minimum levels of wage rate and employment, respectively. θ is a weight reflecting the relative importance of wage and employment to the union, hence, $0 \le \theta \le 1$. The minimum level of wage, $\bar{w}$, usually reflects the level of outside wages; thus this formulation is able to capture the relative wage effect on the union preferences.

The latter formulation uses the theory of representative individuals where the utility of union members as a whole is the sum of individual utilities, each of which is identical. Moreover, layoffs take place randomly. If we denote a number of union members as m, actual employment as n, and $\bar{w}$ to be the wage in alternative employment, then the union's utility function is written as:

$$U = nu(\frac{w}{c}) + (m\text{-}n)u(\frac{\bar{w}}{c}) \qquad (2\text{-}2\text{-}3)$$

This is called the utilitarian utility function. The expected utility approach, on the other hand, assumes the union to be concerned with expected utility of its members rather than its sum, hence:

$$UU = [n/m]u(\frac{w}{c}) + [(m\text{-}n)/m]u(\frac{\bar{w}}{c}) \qquad (2\text{-}2\text{-}4)$$

U=UU when the membership is fixed. This specification of the utility function, unlike the Stone-Geary, explicitly states individual worker's preferences and it also incorporates the membership consideration into the utility function. These specifications, however, become invalid as soon as the number of employment exceeds the membership, that is often the case. We can rewrite (2-2-3) and (2-2-4) so that:

$$U = mu(\frac{w}{c}) + (u(\frac{\bar{w}}{c}) - u(\frac{w}{c})) * \max(\ 0,(m\text{-}n)\ ) \qquad (2\text{-}2\text{-}3)'$$

$$UU = u(\frac{w}{c}) + (u(\frac{\bar{w}}{c}) - u(\frac{w}{c})) * \max(\ 0,\frac{m-n}{m}\ ) \qquad (2\text{-}2\text{-}4)'$$

For the moment, we assume that the membership is exogenously determined unless otherwise stated. In special cases, these utilities reduce to maximizing the wage bill or the rent:

$$U = n\,\frac{w}{c} \quad \text{and} \quad U = n\,(\frac{w - \bar{w}}{c})$$

Workers represented by these utility functions are risk neutral since $U_{ww} = 0$.

2.2.1 Monopoly Union Model

The simplest of all the models is the monopoly union model where an utility maximizing union, constrained by the labour demand curve, sets wages unilaterally. This is followed by a firm choosing the level of employment accordingly. Hence, the equilibrium outcome lies on the labour demand curve.

Let R(n) be revenue function of the firm where labour and capital are the only input of this production function. Isoprofit curve is concave. The firm's objective is to maximize its real profit with respect to the level of employment, that is:

$$\max_{n} \pi = \frac{R(n) - wn - k}{c} \qquad (2\text{-}2\text{-}5)$$

where p is product price, k is fixed capital cost, and c is consumption price index. At the optimum, the firm chooses n so as to make the value of marginal product of labour equal to the real wage: $R_n=w$. In other words, at the profit maximizing point, they choose a level of employment corresponding to the given wage level on the labour demand curve (figure 2.1, L^d). The union has a concave utilitarian utility function (note for a fixed membership, the utilitarian and the expected utility approaches yield the same results), hence indifference curve of the union takes the usual convex shape exhibiting the tradeoffs between the level of wages and employment. Its utility is increasing in both w and n. This indifference curve approaches to $w=\bar{w}$ asymptotically from above as n goes to infinity, since workers are indifferent between working and not working at such wage level. The union maximizes its utility with respect to wage level while taking into account of the labour demand curve. For a standard concave utility function, u(.), the problem facing the union is:

$$\max_{w} U = n\, u(\frac{w}{c}) + (m-n)\, u(\frac{\bar{w}}{c}) \qquad (2\text{-}2\text{-}6)$$

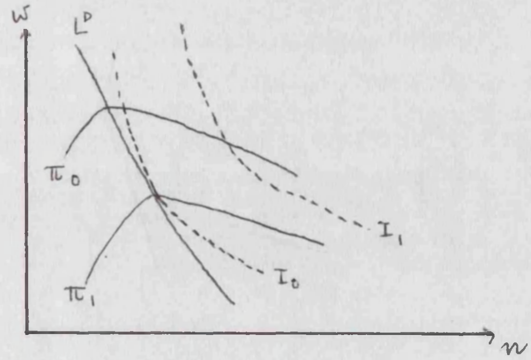
subject to $pf_n=w$, where f_n denotes a derivative of f with respect to the argument in the subscript. This is equivalent to writing:

$$\max_{w} U = n^*(\frac{w}{p})\, u(\frac{w}{c}) + [m-n^*(\frac{w}{p})]\, u(\frac{\bar{w}}{c}) \qquad (2\text{-}2\text{-}7)$$

where $n^*(\frac{w}{p})$..(2-2-8) is labour demand function. This is the level of employment that the firm would choose given the real product wage, $\frac{w}{p}$. On the other hand, under monopolistic product market, maximization of the utilitarian utility is subject to the labour demand function:

$$n^* = n^*(\frac{w}{c}, s, c, e) \qquad (2\text{-}2\text{-}9)$$

It depends on the real consumer wage and other factors affecting the product demand and not on the product price alone as in the perfect competitive case. In both cases, the first order condition which is set to zero at the maximum yields:



(Figure 2.1) Monopoly union

$$U_w = n_w \, [u(\frac{w}{c}) - u(\frac{\bar{w}}{c})] + nu_w$$

This implies that the solution is at a point of tangency between the union's indifference curve and the firm's demand curve, such as point A in figure 2.1. At this point,

$$-R_n/[nR_{nn}] = u_w \, w \, / \, [u(\frac{w}{c})-u(\frac{\bar{w}}{c})] \qquad (2\text{-}2\text{-}10)$$

Marginal cost in terms of reduced labour demand due to the wage rise (i.e., negative of wage elasticity of demand for labour) equals the union's marginal benefit from raising wages (i.e., elasticity of benefit from employment with respect to wages). In the long-run competitive equilibrium, where there is no union and $\bar{w}$ is the wage attainable elsewhere in the economy, the marginal product of labour equals the alternative wage, $\bar{w}$. In the current context, however, (2-2-10) implies $u(\bar{w}/c) < u(w/c)$, thus the marginal product of labour exceeds the alternative wage $\bar{w}$, indicating a job rationing ($n < \bar{n}$) created by the existence of the union which sets wages unilaterally.

Any exogenous shift in the union's utility that leads to a wage rise consequently reduces employment as the equilibrium moves up along the labour demand curve. For example, a rise in unemployment benefit (i.e., rise in $\bar{w}$) raises wages chosen by the union since $U_{w\bar{w}} = (-n_w * u_{\bar{w}})$ is larger than 0. The corresponding changes in the employment is determined by the firm according to the labour demand curve. In this way, changes in the variables that influence union's bargaining power or preferences, such as $\bar{w}$, affect the level of employment only in as much as they affect wages. As can be seen from the reduced form labour demand equation, (2-2-8) or (2-2-9), they should not be affected by variables such as $\bar{w}$ once they are controlled for wages. In other words, n* is affected by $\bar{w}$ only through the changes in w. Improvement of the product market can be captured by product price under perfect competitive framework, and level of GNP or competitor's price under monopolistic framework. If they have impact on the alternative wages through their effect on job opportunities, the movement of $\bar{w}$ will be pro-cyclical and so will the agreed wage. A rise in price of output, on the other hand, will have no effect on wages if the elasticity of labour demand is constant. This can be seen from the FOC:

$$u_w w \;/\; (u(w/c)-u(\bar{w}/c)) \;=\; -w^* \, n_w/n$$

where the right hand side is the elasticity of demand for labour. As long as the right hand side stays constant, the left hand side expression becomes also fixed, so does the wage level. In this case, the wage rate chosen by the union remains the same while the equilibrium employment may increase due to a shift in demand curve, creating the "wage stickiness". Moreover, $U_{wm}=0$, so that membership does not affect the union's desired wage in this model. Effect of changes in CPI has an ambiguous effect on the union's wages since the sign of the following expression is indeterminate:

$$U_{wc} \;=\; n'u'/pc^2\,(\bar{w}-w) \;-\; nwu''/c^3 \;-nu'/c^2$$

Here, for the purpose of clarity, u_w denotes the partial derivative with respect to w, while u' denotes its derivative with respect to its single argument, in this case, w/c.

With respect to the employment level, $pf_n=w$ under perfect competition, hence, n* is independent of c. However, under monopolistic market, demand for labour (2-2-9) is decreasing in w/c and increasing in s and e but ambiguously determined with respect to the movement of c. This is because CPI includes the prices of substitutes as well as complements to the particular product this firm deals with. Nonetheless, if the effect of a fall in real income induced by a rise in CPI outweighs the substitution effect towards the firm's product, then the effect of c would be negative on the number of employment.

This model is consistent with the empirical finding that the firms usually set employment unilaterally. However, in practice, unions do not usually set wages without negotiations. What also lacks in the current setting is the membership consideration, that m does not affect the negotiated wages ($U_{wm}=0$).

Later on, we will discuss the extension of this model that maximizes the Nash formula; a formula that takes into account of the membership consideration into the notion of bargaining power. Last point of this model to be noted is the inefficiency of the equilibrium outcome. As is clear from figure 2.2, there exist points above the labour demand curve (shaded area) that are Pareto superior to the equilibrium point, A. And there is no reason why either parties shouldn't move to any of these points where they can both be better off.

2.2.2 Efficient Bargain Model

The efficient bargain model, where the idea was first introduced by Leontief (1946), was put forward by McDonald and Solow (1981). They focused on the inefficiency of the monopoly union model, and derived an efficient outcome by assuming that the firm and union bargain over both wages and employment. The problem with the monopoly union model is that it leaves a ground for both bargaining agents to be better off. If the union also has a power to negotiate over the level of employment, they will accept lower wages for higher employment which will then increase the firm's profit. Hence, on the whole, the outcome will be Pareto superior. This model can also explain the phenomena of over manning (value of marginal product of labour$<$w), wage rigidity and pro-cyclical fluctuations of employment, all of which are the features commonly observed in the recent labour market.

A firm maximizes its profit with respect to n and w for a given level of union's utility, while the union maximizes its utility with respect to n and w for a certain level of profit. This is identical to maximizing the Lagrangean of a form:

$$L = \frac{R - wn - k}{c} + \lambda(\bar{u} - nu(\frac{w}{c}) - (m-n)u(\frac{\bar{w}}{c})] \qquad (2\text{-}2\text{-}11)$$

with respect to both n and w, for some arbitrary level of union's utility, $\bar{u}$. The first order condition is:

$$\frac{w - R_n}{n} = \frac{u_n}{u_w}$$

which, under the utilitarian union, is:

$$[R_n - w]/c = [u(\frac{\bar{w}}{c}) - u(\frac{w}{c})] / u' \qquad (2\text{-}2\text{-}12)$$

where u' denotes the derivative of u with respect to w/c. In case of perfect competition, $R_n = pf_n$ and the demand function for labour is:

$$n = n(\frac{w}{p}, \frac{w}{c}, \frac{\bar{w}}{c}, \frac{p}{c}) \qquad (2\text{-}2\text{-}13)$$

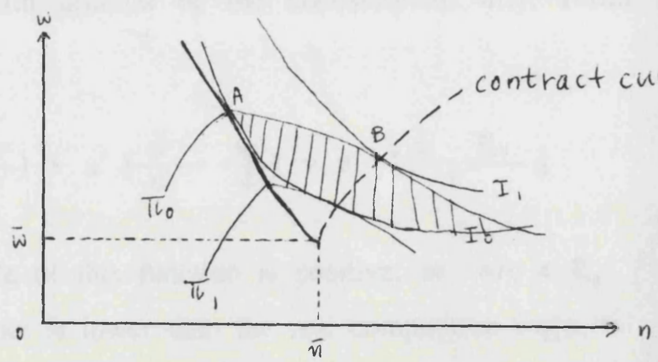
where the arguments follow from the specification of f_n. Under monopolistic competition, on the other hand, the first order condition is:

$$R_n = \frac{w}{c} - [u(\frac{w}{c}) - u(\frac{\bar{w}}{c})] / u' \qquad (2\text{-}2\text{-}14)$$

where R = R(n,s,c,e). The corresponding labour demand function is:

$$n = n(\frac{w}{c}, \frac{\bar{w}}{c}, c, s, e) \qquad (2\text{-}2\text{-}15)$$

In both cases, at the equilibrium, the slope of isoprofit curve for the firm is equal to that of indifference curve for the union. Hence, the equilibrium lies on the contract curve (a locus of such points of tangency) and is Pareto efficient; it is impossible to make either party better off without making the other party worse off by moving away from those points (point B in figure 2.2).



(Figure 2.2) Efficient bargain

Consider a point such as A. This point is on the demand curve, but any point in the shaded area is Pareto superior to this. Such superior points will continue to exist until it reaches point B; a point on the contract curve.

In addition to this revealing characteristic, following points are also observed. $R_n=w$ on the labour demand curve, and it intersects with the contract curve when $u(w)=u(\bar{w})$ (making RHS of (2-2-12) equal to 0). Then, the point $(\bar{w},\bar{n})$ represents the competitive outcome with no union and the outside utility, $u(\bar{w})$. A slope of the contract curve is positive for the risk averse union, because dw/dn is positive except at $(\bar{n},\bar{w})$, where the contract curve is vertical:

$$dw/dn = (R_{nn}u') \ / \ \{[w - R_n] \ u''\} \qquad (2\text{-}2\text{-}16)$$

The contract curve is also vertical for $u''=0$, that is, when the union is risk neutral. In this extreme case, the number of employment is not affected by the wages determined but only by the alternative wage, $\bar{w}$. For instance, this applies to the rent maximizing union, $u=(w/c-\gamma(\bar{w}/c))n$, where γ is a fixed parameter. In particular, when $\gamma=0$, it is strongly efficient and is socially Pareto optimal since the marginal product of labour is equal to the alternative wage throughout the economy. Bargaining outcome is also efficient for the risk neutral union with the expected or utilitarian utility function. With a risk loving unions, the contract curve slopes down, but this is very unlikely. More realistically, if the union is not interested in the level of employment (see section 2.2.5), the labour demand curve coincides with the contract curve.

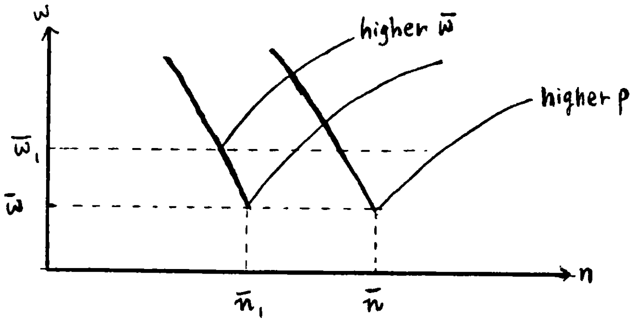
For $w>\bar{w}$, the first order condition implies wages in excess of the marginal product of labour along the contract curve, hence the level of employment is higher than that determined solely by the firm. Even stronger result is true when we take into account of the concavity of u(.). From the first order condition:

$$u(\frac{w}{c}) - u(\frac{\bar{w}}{c}) + u' \ [\frac{\bar{w}}{c} - \frac{w}{c}] = u' \ [\frac{\bar{w} - R_n}{c}] \qquad (2\text{-}2\text{-}17)$$

The left hand side of this function is positive, so, $\bar{w}/c \geq R_n$. Hence, the value of marginal product is lower than the real competitive wage, $\bar{w}/c$, implying over employment under the efficient bargain. Wage rate in excess of $\bar{w}$ invites excess supply of work force, thus, it needs some mechanism to ration the level of employment. Moreover, any exogenous shock that has the effect of raising wages, such as a rise in the union power, also has a positive effect on the level of employment. This result is distinctively different from that of the monopoly union model where the tradeoff between wages and employment always exists to the extent that the shock makes them move along the labour demand curve. For example, a rise in the unemployment benefit (i.e., rise in $\bar{w}$ to $\bar{w}_1$ in figure 2.3) shifts the contract curve up to the left making the starting point of the curve at $(\bar{n}_1,\bar{w}_1)$. Consequently, higher wage will be determined for a given level of employment.

$$dw/d\bar{w} = \frac{u'c}{u''(R_n - w)} > 0$$

On the other hand, a rise in output price shifts the labour demand curve hence will shift the contract curve to the right giving higher employment level at any given wage.



(Figure 2.3) Shift in the contract curve

$$dw/dp = - \frac{R_{np} u'c}{u''(R_n - w)} < 0$$

where $R_{np} = f_n$ under perfect competition. The wage stickiness under the efficient bargain is accentuated when the union seeks for a "fair share" deal, so that the wage bill is always a fixed proportion of the total revenue: $wn=kR(n,p)$..(2-2-18) (McDonald and Solow (1981)). This, together with a contract curve equation (2-2-17), yields:

$$\begin{bmatrix} \frac{u - \bar{u}}{(u')^2} & -R_{nn} \\ n & w-kR_n \end{bmatrix} \begin{bmatrix} \frac{dw}{dp} \\ \frac{dn}{dp} \end{bmatrix} = \begin{bmatrix} R_{np} \\ kR_p \end{bmatrix}$$

The numerator of dw/dp is equal to $-w\{R_{np}(1-nR_n/R)+nR_{nn}R_p/R\}$ whose sign is indeterminate. Under the fair share rule, the negotiated outcome will be at the intersection of the contract curve (eq.2-2-17) and a locus of the fair share rule (equity locus, eq.(2-2-18)), which is negatively sloping for the positive profit. In a recession, decline in the product demand (i.e., decline in p) shifts the contract curve to the left and the equity locus down to the left. Hence, the negotiated level of employment will decrease unambiguously while the wage can either go up or down. McDonald and Solow has given two examples where the labour demand or the inverse labour demand is not affected by p for a given level of wages or employment. When dw/dp=0, all the burden comes out in the form of reduced employment. In other cases, a sign of dw/dp becomes the opposite of $d/dn(NR_n/R)$. This implies that during a recession, if the revenue elasticity decreases, due to sales constraint for instance, the negotiated wage rate will increase.

Obviously, the most appealing aspects of this model is the Pareto optimality of the equilibrium outcome. According to their setting, however, the union and firm have to negotiate over the level of employment in the same way as they do over pay. The stylized facts suggest that the level of employment is largely adjusted unilaterally by the employers and very rarely do we see a contract specifying the level of employment or even a man-machine ratio. Firms usually adjust the size of their labour force without any negotiation with the workers, except when the adjustment involves a large number of compulsory redundancies. This maybe because that the firm may find it costly to negotiate over employment as well as wages. Also, the unions may be more interested in the condition of work rather than the number of employment. Once the wage is fixed, the firm always has an incentive to renege by jumping over to the labour demand curve and derives a higher profit while employing fewer number of workers than it would had it kept to the negotiation with the union. In order to prevent such asymmetry, the contract has to be complex enough to ensure the union's involvement in the determination of the employment level.

2.2.3 Formal Bargaining

So far, we have been using the Lagrangean method to solve the bargaining problem between two parties. The welfare level of either party was maximized subject to the constraints which usually involved the welfare level of the other party. This method was suitable in identifying the equilibrium level of wages and employment in the case of the monopoly union model, since the constraints and the choice variables each party confronted with enabled us to solve for a unique equilibrium point on the labour demand curve. On the other hand, under the efficient bargain setting, it was the contract curve equation, a locus of feasible equilibria, that was derived from the first order condition of the

Lagrangean optimization problem. Varying the the value of Lagrangean multiplier identified different points on the contract curve. We need additional information in order to determine a unique point on the contract curve that is to be a solution to this bargaining problem.

Consider, instead, an axiomatic rule that can accommodate several forms of bargaining. This theory operates in the context of the bargaining set and its efficient frontier instead of the contract curve itself, although the solution will be on the contract curve. This theory is justified axiomatically as well as strategically (Binmore, Rubinstein and Wolinsky (1984)).

First, define a bargaining set as a set of possible outcomes of payoffs. Given such a set, a formal solution to this bargaining problem is derived by maximizing the Nash formula, $(z-\bar{z})(\pi-\bar{\pi})^{\beta}$, subject to the constraint, $f(z,\pi)=0$, where z and π are the utilities of each party. This constraint is the frontier of the efficient payoff possibility curve that is decreasing and concave in (z,π) plane, for the axiomatic bargaining theory requires the bargaining set to be convex. The Nash formula is a product of each party's gains over and above the non-contract (fall-back, status quo) outcomes, $\bar{z}$ and $\bar{\pi}$. This product form is derived from the formal bargaining setup where the sequences of offers are made by each party and the compromise to achieve a settlement is induced by the lower welfare level (i.e., fall back utility) that has to be tolerated until the agreement is reached. To the union, it is the loss of forgone wages, to the firm, it is the loss of output and profits that prompt them to reach an agreement quickly. This assumption, that the bargaining outcome should never fall below the status quo utilities, is crucial or otherwise the union and firm would not bargain at all (strong individual rationality). In fact, this assumption is sufficient in ensuring the existence of a unique Nash outcome instead of assuming the agents' behaviour to be Pareto Optimal (Svejnar (1986)).

Under this setting, each party's optimal acceptance level is chosen by comparing the present value of immediately accepting whatever is offered by the other party and the present value of waiting for its desired offer to be accepted. In the latter case, one has to bear with the status quo level until its desired offer is accepted by the other party. By equating these two present values, we can derive the maximum tolerable stoppage time for each party. Whichever with the shorter tolerable time will have to compromise its offer. This comparison of the maximum stoppage time is equivalent to comparing the product of each party's welfare over and above their fall-back outcome, that is equivalent to maximizing the form of Nash maximand: $(z-\bar{z})(\pi-\bar{\pi})^{\beta}$. β here is interpreted as the bargaining power, which originally stems from the relative discount rates of each party. A party with the higher discount rate is less patient or less able to afford the long stoppages, the cost of disagreement is higher for such party, thus, has a lower bargaining power. The fall-back outcomes for the two parties are their welfare level during the stoppage. If the fall-back profit for the firm, $\bar{\pi}$, is the level of fixed cost, -k, then:

$$\pi - \bar{\pi} = [R(n) - wn]/c$$

The union is assumed to maximize its utilitarian utility, hence:

$$z = n[u(\frac{w}{c}) - u(\frac{\bar{w}}{c})] + mu(\frac{\bar{w}}{c})$$

where m is number of members. Their fall-back welfare level is $\bar{z}=mu(\bar{w})$, as long as the workers can derive outside income, $\bar{w}$, even during the strike.

$$z - \bar{z} = n\ (u(\frac{w}{c}) - u(\frac{\bar{w}}{c}))$$

Therefore the Nash formula to be maximized is:

$$[\ n\ (u(\frac{w}{c}) - u(\frac{\bar{w}}{c}))\ ][\ \frac{R - wn}{c}\ ]^{\beta}$$

subject to the constraint, $f(z,\pi)=0$. If the solution indicates utility level that is less than their outside option, the outcome of the agreement will not be on the frontier, $f(z,\pi)=0$, but at a corner solution.

With the additional information regarding the relative bargaining power between the two parties, this framework is general enough to embody both the monopoly union setting discussed in section 2.2.1, as well as the efficiency bargain problem of section 2.2.2. In particular, it derives a unique solution to the efficiency bargain model.

In the monopoly union, the union determines the level of wages unilaterally subject to the labour demand curve. Hence, $\beta=0$ and the problem now reduces to:

$$\max_{w}\ n^*(w)\ (u(\frac{w}{c}) - u(\frac{\bar{w}}{c}))$$

where n*(w) is the labour demand curve; a solution to the profit maximization of the firm, $R_n = w$.

Under the efficient bargain model, the union and firm bargain over both w and n so that the problem is now:

$$\max_{n,w} \; n\left(U\left(\frac{w}{c}\right) - U\left(\frac{\bar{w}}{c}\right)\right)\left(\frac{R - wn}{c}\right)^{\beta}$$

A ratio of the first order condition to this problem gives the contract curve equation (2-2-12). Furthermore, each first order condition gives a relation between w, n, β and their fall-back welfare levels. Note that these were not known under the analysis in the former section. This additional relationship identifies the equilibrium wage level, $w=w(p,c,\bar{w},k,\beta)$. We saw from equation (2-2-16) that the risk averseness (u'') of the union determined a location (slope) of the contract curve. This time, the bargaining power, β, determines the location of a solution on the contract curve.

To make this model even more appealing intuitively, Svejnar (1986) has let the bargaining power, β, to depend on the exogenous variables to provide a channel through which variables such as union membership or government policies have impact on the bargained level of wages and employment.

2.2.4 Right to Manage Model

The right to manage model as discussed in Nickell and Andrews (1981) and the monopoly union model together fall into a family of the labour demand (LD) model since its solution lies on the labour demand curve. Indeed, their comparative static results are very similar to each other. Yet, the right to manage model captures the stylized institutional facts that are not seen in the monopoly union model; the firm and the union bargain over the wage level, while the firm still holds a right to determine the level of employment unilaterally. By letting both parties bargain over wages, bargaining power (i.e., relative bargaining strength) can be incorporated to affect the outcome. This factor is ignored in the monopoly union model.

The firm maximizes its profit with respect to the level of employment, therefore:

$$\max_{n} \pi = \frac{R - wn}{c}$$

which gives the labour demand, n*(w), as a solution to the first order condition. The firm and union together bargain over w to maximize their utility, $\pi(w,n^*(w))$ and $U(w,n^*(w))$, given a certain level of π and u, with respect to w (In the efficient bargaining, it was with respect to both w and n). This is equivalent to solving a Nash formula:

$$\max_{n,w} n \left(u\left(\frac{w}{c}\right) - u\left(\frac{\bar{w}}{c}\right) \right)\left(\frac{R - wn}{c} \right)^{\beta}$$

subject to $R_n = w$. This can also be solved by:

$$\max_{w} n^*(.)\ (u(w) - u(\bar{w}))(R(n^*(.)) - wn^*(.))^{\beta}$$

where $n^*(.)$ is the labour demand function under the appropriate product market situation. First order condition is:

$$\frac{1}{nR_{nn}} - \frac{\beta\ n}{R - wn} + u' \ / \left(u\left(\frac{w}{c}\right) - u\left(\frac{\bar{w}}{c}\right) \right) = 0$$

where $R_n = w$. The comparative static results are the following. Variables that increase opportunities for the workers outside the firm raise the bargaining power of the union, hence, lead to higher wages. If the membership or union density affect the bargaining power parameter, β, dw/dm will no longer be zero, as was predicted by the monopoly union model. Similarly, an increase in $\bar{w}$ raises w. On the firm's side, higher product price or higher productivity enables the firm to have larger rent to be exploited by the unions. While if the firm adopt higher employment, higher wages become more expensive to the firm at the margin. If we expect the income effect to dominate, higher product price will result in higher wages. Higher fall-back profit increases the firm's bargaining power, hence lowers wages.

The shadow value of labour does not depend on alternative wage, and this is a feature of the labour demand model that also applies here. This property is often used to differentiate the labour demand model from the efficient contract model. Nevertheless, even if a firm stays along the labour demand curve, introduction of efficiency wage consideration (e.g., discouragement of shirking) can bring about the relevance of the outside opportunities in the employment equation. Extension of the LDM in this way makes a discrimination of the LD model from the EC model very difficult one to test empirically.

Layard and Nickell (1988) showed that under a *general equilibrium* framework with Cobb-Douglas technology and constant elasticity utility, the Nash outcome yields the same level of aggregate employment whether unions bargain over employment as well as wages or wages only. Moreover, if the elasticity of substitution between capital and labour exceeds unity (hence, there is greater substitutability between the two), employment level will be lower under the efficient bargain than under the right to manage model. Also, moving from the competitive labour market to the efficient bargaining reduces employment. If the workers know the relation of the labour demand curve, highly elastic demand for labour makes union more cautious when bargaining only over wages. This result is different under the partial equilibrium analysis, which assumes the same external opportunities regardless of the system of wage determination in each representative firm, where the employment is higher under the efficient contract. When the union can bargain over n as well as over w, they are given more power. It seems counter intuitive to find higher unemployment in the equilibrium when the union has more power, although their higher power is used principally to raise wages. The gain from higher wages more than compensates the loss from lower employment.

Under this right to manage model, the union is aware of the demand curve of labour, and yet, the union and employers do actually bargain over the wage level. This aspect is distinctly different from the monopoly union model. Also, they do not bargain over both the level of wage and employment as they did under the efficient bargain model. The outcome, therefore, is not Pareto optimal. Nonetheless, as Nickell and Andrews argued, the firm might find it too costly to negotiate over employment. Given such a cost consideration, this theory very well explains the stylized fact that employers continuously adjust their level of employment without bargaining with the union. The rationale as to why this form of bargaining takes place, however, is not given in these literature. Special case, in fact, is considered in the following section.

2.2.5 Seniority Model

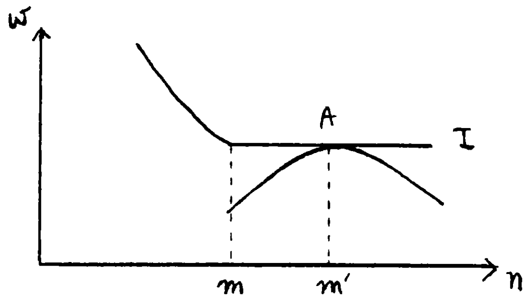
Here, we introduce a model which is a half-way-house between the efficient model and the labour demand model. It gives an efficient outcome (in a sense that it lies on the contract curve), while at the same time, it lies on the labour demand curve, hence the value of marginal product of labour is equal to the wage rate. This can be considered as a special version of the efficient bargain model which complies with the stylized fact that characterizes the right to manage model, yet, avoiding the inefficiency underlying the labour demand model. It justifies the form of bargaining seen in the right to manage model where it is only the wages which is negotiated and specified in the contract while the level of employment is chosen solely by the firm. There are two possible rationales which yield such setup for the bargaining and they differ in their specifications of the union's preference.

The first case assumes the strong discriminating power of the "insiders" (i.e., members) as opposed to outsiders and that it is only the insiders' preference that matters. Hence, as long as all the members are employed (i.e.,n≥m), the union is indifferent to the level of employment. It is only when the risk of its members becoming unemployed arises that the union starts caring about the level of employment. We assume that all the employees are paid the same wages. The utilitarian and expected utility function are re-written as:

$$U = u(w) + (u(\bar{w}) - u(w))^* \max [0 , (m-n)/n]$$

$$U = mu(w) + [u(\bar{w}) - u(w)]^* \max [0 , (m-n)]$$

In this case, union's indifference curve becomes horizontal for n≥m as in figure 2.4. For example, in an industry whose employment is expanding, the equilibrium occurs at point such as A, which is at tangency between the union's indifference curve and the isoprofit line; the firm's profit maximizing point (i.e., on the labour demand curve). When new employees are allowed to become insiders, however, the membership increases to m' and its effect becomes ambiguous. In this way, the transition of equilibria overtime for this model involves a complex problem of endogeneity because the union's preference may change as membership changes. If larger membership is considered to expose insiders with higher risk of being laid off in a recession, wage demand will decline. On the other hand, the larger the membership, the larger the marginal utility of pay for the union as a whole, inducing a stronger incentive for the union to demand for higher wages. This will lead to a decline in employment (therefore, fall in the number of insiders) when faced with a negative demand shock, since smaller number of insiders will try to convert all the rent into higher pay rather than

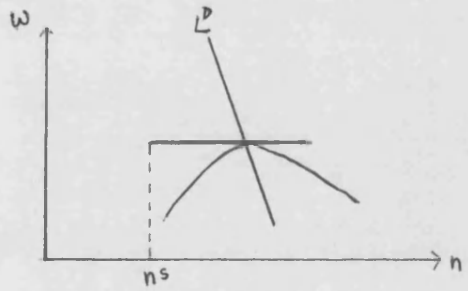


(Figure 3.4) Flat indifference curve: membership model to expand employment for the outsiders. Consequently, unemployment persists despite the fluctuations of the aggregate demand. This is called hysteresis (Blanchard and Summers (1987)) and it can explain the persistently high wage rises accompanied by equally high level of unemployment seen in the current U.K. The size and speed of adjustment of this effect also depends on how the entry and exit to the membership is determined.

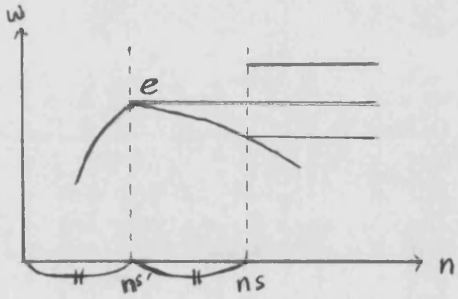
The second setup produces a horizontal segment in the union indifference curve by introducing a certain order to the laying off procedure such as, "last in first out" principle. Layoffs are not random; it assumes these to be done by seniority and the majority voting scheme within the trade union. Then, the behaviour of the union is characterized by the maximization of the median worker's utility. So that as long as the median seniority worker is insulated from the risk of being laid off, even though they can bargain over the level of employment, they are happy to leave its decision to the employers and merely interested in demanding higher wages. Only when a large number of members' jobs are threatened, does the union start exercising its power to bargain over the level of employment (concession bargaining). This argument is applicable whenever the lay off procedure takes place with some known sequence. For instance, it can be the order of tenure or age.

The equilibrium in this model can be considered as follows. As long as the median workers are insulated from the risk of redundancy, that is, if the level of employment, n, exceeds the median worker's seniority position, n^s, the union's indifference curve (represented by the n^s-th (median) worker) is horizontal. It is because for $n>n^s$, the union simply seeks for higher wages without caring about its consequence on employment. Outcome is on the isoprofit line touching the indifference curve at its peak, (see figure 2.5) where $w=R_n$ (i.e., on the labour demand curve). Moreover, at this point, the outcome is "efficient" since it lies on the contract curve. Since it is derived by maximizing the individual's (i.e., median worker's) utility, it is only Pareto efficient in a partial way. On the other hand, when there is a huge decline in demand that makes median voter to face a risk of redundancy, there will be a corner solution. This is no longer efficient, as depicted in figure 2.6.

Let us consider more formally what happens over the period. As long as the demand movement leaves a median seniority worker safely insulated from the risk of redundancy, it is still rational for the median worker to vote for the highest wage achievable without caring about the level of employment. This will lead to a higher wage with no growth in the level of employment for a positive demand shock. On the other hand, when there is a huge negative demand shock, huge enough to make the median worker to face a redundancy risk, there will be a corner solution as depicted in figure 2.6. There is a "concession bargaining" where the union vetoes upon the firm's optimal employment choice. At this point, it is neither efficient nor a steady state equilibrium. Next period, the union will be represented by a new median seniority worker at position $n^{s'}$ who will opt for fewer employment, and a new steady state equilibrium will be reached at point such as e in figure 2.6, given that the isoprofit line remains unchanged. But under this mechanism, even if there is no risk of unemployment over the current contract, as long as there is a negative shock, one time a median worker will never stays as a median worker since the level of employment falls. There will be a new median seniority worker representing the union's utility who continues to bid the wage up irrespective of the declined level of employment. In this way, the former median worker will be laid off eventually, which makes his strategy, that continues voting for the maximum possible wages at each settlement, irrational.



(Figure 2.5) Flat indifference curve: seniority model



(Figure 2.6) Concession bargaining

Layard (1989) argues that under the world of no demand uncertainty and long but finite life with a certain turnover each period, the median voter in the seniority model should act to achieve the highest possible wage in each period such that it will keep him employed up to his retirement. The only way to make sure that he will never get laid off is to remain being a median voter throughout his working life. Under the partial equilibrium framework, he analyzes how the convergence towards a steady state level of employment, n*, is reached from below and above its level. When there is a drastic fall in demand, as long as there are s most senior workers who leave every period, the median worker will vote for the wage level that reduces employment by s per period until the level of employment reaches a little more than 2n* but less than 2n*+s. So that the equilibrium employment level is reached from above without the median worker facing a risk of redundancy. One period later, he will vote for the level of wages corresponding to n* on the labour demand, so as to restore the steady state equilibrium.

Here is a formal model that illustrate both of the rationales discussed above:

$$\max_{w,n} u(w/c) \quad \text{such that, } R - wn \geq \bar{\pi} \text{ and } n - n^s \geq 0$$

where $n^s = m$ for the membership model. For the range $n \geq n^s$ (or $n \geq m$), this reduces to maximizing the union utility with respect to wages subject to the labour demand curve. If we exclude the possibility of a corner solution at $n = n^s$ (or $n = m$), this gives a unique solution (w*,n*). Let us solve for the Nash solution to this problem in order to reveal a unique solution. The maximand is:

$$\max_{w} L = [\, u(w/c) - u(\bar{w}/c) \,][\, R(n^*(.) - wn^*(.) - \bar{\pi} \,]$$

where n*(.) is a solution to $\max_{w}$ (R-wn), $\bar{\pi}$ and $u(\bar{w}/c)$ are the fall back utility of the firm and the union, respectively. First order condition is:

$$L_w = u'/c(u-\bar{u}) \;\; - n/[\, R(n^*)-wn^*-\bar{\pi} \,] = 0$$

with second order condition:

$$\Delta = \frac{u''}{c^3(u-\bar{u})} - \frac{u'^2}{c^2(u-\bar{u})^2} - \frac{n_w}{(\pi^*-\bar{\pi})} - \frac{n^2}{(\pi^*-\bar{\pi})^2} < 0$$

Comparative static is:

$$L_{w\bar{\pi}} = - n \,/\, (\pi^* - \bar{\pi})^{\,2} \quad < 0$$
$$L_{w\bar{w}} = u'^2 \,/\, c^3\, (u-\bar{u})^2 \quad > 0$$

Hence a rise in the firm's fall back profit decreases the negotiated wage while a rise in the union's fall back wage increases their bargaining power thus raises the negotiated wage.

$$dw/dp = - \frac{nf/(\pi^*-\bar{\pi})^{\,2}}{\Delta} > 0$$

The denominator is negative from the second order condition, thus, dw/dp is positive. However, an extreme wage rigidity can be observed when the production function is of constant elasticity form and $\bar{\pi}=0$. Then the problem reduces to :

$$\max_{w} \ (u-\bar{u}) \ \pi(n^*(.))$$

whose first order condition is:

$$(u-\bar{u})/u' \ = \ - \ \pi/\pi_w$$

Letting $f(n)=n^{\alpha}$ where $\alpha\in(0,1)$, the right hand side is equal to:

$$pf/n \ - \ w \ = \ w(1-\alpha)/\alpha$$

using the profit maximization condition for this particular production function. This is independent of p, hence, the wage is determined independent of product price (Oswald (1987)). This wage rigidity is confined to the case where $n>m$ ($n>n^s$). For $n<m$, the union cares about the level of employment, hence the bargaining setup becomes that of the efficiency bargain under which we also observe wage rigidity (see section 2). It is indeterminate for the case $n=m$, when the equilibrium is reached at the kink of union indifference curve where the contract curve is vertical. This is likely to happen under the membership context for the open union where all the employees automatically become members. Carruth and Oswald (1986) has simulated a estimates derived for the constant elasticity utility and production function with $\bar{u}=mu(\bar{w})$ and $\bar{\pi}=0$. They find a strong wage rigidity associated with an increase in the employment towards the number of membership. Once n reaches m, the employment level stays rigid while all the gains are transformed into higher wages. Further increase in the product price increases the level of employment above that of the membership indicating that by this time, it is also of insiders interest, to let outsiders have the job. Such rise in employment increases the firm's profit which eventually benefits the existing members. This refutes the conventional critique of this model predicting the union that forever claims for higher wages. They also find that the lower the risk aversion and production elasticities, the larger the price range that keeps employment level rigid at around the level of membership.

The feature of "last in first out" of this model is shown to be consistent with the evidence according to Medoff and Abraham (1981). Their paper, based on the newly collected US survey data of non-agricultural non-construction companies in the private sector, strongly supports the claims that the senior workers enjoy over-payment of their wages together with greater protection against redundancies and the latter point is particularly true amongst the unionized employees. Also appealing is the Pareto efficiency of the outcome which lies on the labour demand curve. At equilibrium, workers are not interested in the elasticity of labour demand except at a time of sudden sliding recession, in such circumstances, the concession bargaining may take place.

2.3 Empirical work on the union-firm bargaining theory

Since the introduction of the union-firm bargaining theory discussed in the former section, a vast number of studies were carried out to test their empirical significance. Very few used aggregate time series data where they estimated reduced form models of the wages and employment (Nickell and Andrews (1983)). Majority of them have concentrated on using more disaggregated data. Different industries have different characteristics, and face different economic climate over time. In particular, with respect to the union-firm relation, considerable institutional differences must exists in a way the bargaining takes place. In this sense, even the industry based data may not be adequate enough to take into account of such diversity. Hence, it is sensible for such empirical studies of the bargaining to be based on the data which is as disaggregated as possible in order to avoid the aggregation problems. Most of the empirical studies introduced in this section are based on the particular industries where their own institutional characteristics, such as their relation to the firms, are taken into consideration in forming the structure of bargaining.

These empirical work all assume that the union and firm act as though maximizing (or minimizing) well defined objective functions. It is generally agreed that the firm's objective function is something related to the level of profit (although it is questionable particularly under nationalized industries). However, when it comes to that of the union's, the issue is rather more complex. Some of the studies have focused on the nature of union's objective functions while the form of bargaining is assumed a priori as a maintained hypothesis. The others test the efficiency of contracts (that is to say to test the form of bargaining) given certain functional specifications for the objective functions of the parties involved. In particular, some concentrate on testing between the two polar models of employment determination: (1)the labour demand model where the union optimizes with respect to the wage while constrained by the labour demand curve, (2)the strong efficiency model with a risk neutral union and the firm bargains over both wage and employment. In this case, the problem is the choice of a criteria in discriminating those two polar models without assuming too strict the functional forms of the agents' objective functions.

However, as we have seen in the former section, there are cases when the efficient contract lies on the labour demand curve. Assuming that the members' utility is represented by a leader and there is a distinction between selfish members (insiders) and non-members (outsiders) or because of LIFO, it is plausible to expect that the horizontal segment emerges on the union's indifference curve. As long as all the members (or the representative members) are employed, union becomes locally indifferent about expanding the level of employment. With such an assumption, the simplified model predicts the negotiated wage to be a function of the firm's profits and the external pressures. Under this assumption, discriminating criteria mostly chosen for these studies becomes unidentified. Moreover, the rejection of either of these polar models, even if the union cares about the employment level, does not necessarily lead to the acceptance of the other. In the end, the actual bargaining seems best represented as weakly efficient with a good case for the efficiency wages consideration.

We start off with the first type of empirical literature which mainly concentrate on the test of the union objective function, followed by which the work involving a test of efficiency of a contract. And lastly, we review briefly those literature based on the model of insider-outsiders. Their assumed structure of the model, the method of estimation and their major findings are discussed.

2.3.1 Union's objective function

Farber (1978) is the first to carry out an empirical analysis on the union-firm bargaining problem. He formulates the union behaviour of the United Mine Workers (UMW) in the USA such that the union maximizes expected utility of a median aged member subject to a labour demand curve. Institutional background of the UMW provided a suitable background for his formulation of the monopoly union bargaining: the coal industry is labour intensive and UMW had a dominant force in an industry because of a growing concern over the OPEC oil embargo which made the union to be a monopolistic supplier of labour to this industry. UMW negotiated for three types of compensations, two types of benefits and a direct payment. He assumes all members to be identical in their preferences over the total compensations. It is represented by the constant absolute risk aversion utility function that permits risk neutrality as a special case. Since workers vary in their age, he has allowed their preference over the income composition to vary. Then, a leader of the UMW acts as if to satisfy the utility of the median aged worker so as to minimize his risk of losing a job. Hence the expected utility of the union is represented as:

$$\frac{n}{m} U^m + (1 - \frac{n}{m}) U^n \qquad (2\text{-}3\text{-}1)$$

where $U^m$is the utility of employed members represented by that of the median worker (i.e., leader) and U^n is the utility of unemployed members. Farber's aim is to investigate the nature of the union's objective, in particular, the parameters that reflect risk averseness, relative importance of direct wage payment versus fringe benefits and the rate of time preference. Using annual data from 1947 to 73, the first order condition derived from the utility maximization together with 7 other reduced form equations regarding the labour demand and other product market conditions are estimated by FIML where additive error terms are assumed to be multivariate normal. He uses the average hourly earnings in durable goods manufacturing sector as a measure of alternative wage. As a result, the union is shown to be highly risk averse (relative risk aversion coefficient is larger than 3.0) and the UMW's concern over employment is higher than that predicted by wage bill maximization. Also, he finds that the workers value a dollar received as the fixed benefits more than they do as the discretionary income. Price of oil is influential over the bargaining outcome, but on the whole, the union's policy is not very responsive to any shift in the demand for coal. Also, he argues that the omission of strike cost might have been responsible for the relatively high value of risk averse parameter, if the strike activity and the union demand were positively correlated.

Carruth and Oswald (1985) has formulated a union behaviour of the National Union of Mine Workers (NUM) in Great Britain, under the monopoly union setting. The NUM is the single unusually militant trade union in the British coal industry which is nationalized. Their study focuses on the objective of the union, in particular, the effect of the shift in labour demand, alternative wage and technological change on wages. They start with the utilitarian utility function where each identical worker is represented by a constant relative risk aversion utility with an additive "comparative" wage component. The reduced form labour demand is formed to include lagged value of exogenous variables such as price of oil to capture slow inter-fuel substitution. Technological progress is also incorporated, which they found best represented by a time trend rather than productivity per se. They estimate two models one with the instantaneous and another with the slower adjustment process between employment and wages. In the latter, the union determines a time path rather than a unique pair of wages and employment by solving an intertemporal optimization problem where the partial adjustment equation for the employment is: $\Delta N_t = \lambda(\hat{N} - N_{t-1})$, where $\hat{N}$ is the target level of employment and parameter λ reflects the speed of adjustment.

The first order condition derived by maximizing the appropriate present value of the utilitarian utility and the reduced form labour demand functions are together estimated by FIML. Alternative wage measure in their case is the average earnings in the manufacturing sector. As a result, they find the partial adjustment framework better in explaining the NMW's behaviour. Relative risk aversion coefficient is very significant at around 0.8 (1.1 under intertemporal model). Also, they found the significant employment adjustment parameter, λ, suggesting that a gap between the target and actual employment level in any one year is closed by as much as 50%.

Dertouzos and Pencavel (1981) has adopted Stone-Geary utility function for the International Typographers Union (ITU) under the framework of the labour demand model in the USA. Stone-Geary utility has wage bill or rent maximand as its special case and can explicitly measure the relative weight between employment and wage in the union utility.

$$U(w/p,n) = (w/p-\gamma)^{\theta}(n-\delta)^{1-\theta} \qquad (2\text{-}3\text{-}2)$$

This utility specification is also useful in finding out whether the absolute real wage, w/p, or the level above some "norm", γ, matters to the union. For instance, if γ is the lagged real wage, then the growth of real wage is what matters to the union. In this way, Stone-Geary functional form enables us to test for interesting structural parameters representing the individual's utility. Institutional background of the ITU, the homogeneity of its members and its renowned democracy, justifies this single utility function to be the representative of all its members. Also, the consensus is that the union is aware of its industry's competitiveness, hence, remain cooperative so as not to set wages unduly high. This is consistent with the assumption that the union chooses the value of wage subject to the labour demand curve.

The reduced form labour demand function and the wage equation derived from the first order condition are together estimated simultaneously by FIML. Although the variation of $\hat{\theta}$ across different union locals are notable, the wage bill and the rent maximand are both rejected. $\hat{\theta}$ lies somewhere between 0 and 0.5, implying the union who places relative importance on the level of employment rather than supernumerary wage, $(w/p-\gamma)$. This $\hat{\gamma}$ is significantly larger than 0, hence that the union makes a decision over wages with reference to some "norm". However, the overall preferences seem to vary substantially across the locals. In some cases, the parameter estimates are pretty implausible. This may be due to the strictness of the functional form assumed for the utility function posited above.

Pencavel (1984) has introduced the union utility function that is even more flexible than that of Stone-Geary to investigate the union behaviour still constrained by the labour demand curve. He also uses the ITU's annual data in 10 different cities. The addilog function represents the union's utility which can accommodate the rent or wage bill maximand in addition to the constant elasticity of substitution utility. This is expressed as:

$$U(w/p,n) = \mu \frac{(w/p-\gamma)^{1+\lambda} - 1}{1+\lambda} + (1-\mu) \frac{n^{1+\iota} - 1}{1+\iota} \quad 0 \leq \mu \leq 1 \qquad (2\text{-}3\text{-}3)$$

Because of the complexity of this utility function, instead of estimating the reduced form equation for the real wage, he directly estimates the marginal rate of substitution (MRS) which has a simpler expression. The MRS is then equated to the slope of a particular labour demand function, namely, $\exp(\alpha x)/r$. They are estimated jointly by non-linear two stage least squares since the level of employment is treated as endogenous. He finds that the utility function is generally quasi-concave in w and n. And the elasticity of substitution between w and n is somewhere between 0 and unity, reflecting little substitutability between the two. He has also found significant variations amongst different locals. Hence, he repeats the estimation by splitting the sample into large and small locals. The rent maximand hypothesis failed to be rejected by the larger locals, and in general, they tend to weigh the "supernumerary" wages more heavily than the employment (γ here is set proportionate to the wage of retail trade workers). This may be due to greater outside employment opportunities in bigger cities where the larger locals tend to be located.

On the basis of assumption that the union and firm bargain on the labour demand curve, what these empirical studies have shown in common are the following: the unions are concerned over the impact of employment on wages more than they do under the risk neutrality. Exogenous shocks that affect the product demand are shown to have varying degree of influence over the union. In fact, Pencavel, Dertouzos and Pencavel have found substantial variations in the union's objective function in different locations even within a single union (ITU). Pencavel (1985), in his survey, has computed the elasticity of substitution between wages and employment for the number of empirical work, but they widely vary between 0.18 to 2.1 and some of them are estimated imprecisely. Altogether, these results accentuate the importance of the institutional and regional factors that are unique to each union and firm concerned, hence, reminds us the necessity of using the disaggregated data in analysing such union-firm bargaining behaviour.

All the studies introduced so far focused on the form of the union's objective functions while the number of employment was always left to a firm to decide unilaterally. Although the main concern was in the union's objective function, it was still necessary to posit a certain functional form for the labour demand function. In this way, these analysis involved a joint test of the functional forms of the union objectives as well as the labour demand function. Hence, any results derived hitherto depends on the correctness of the alternative hypothesis, that is, the specification of the labour demand function.

2.3.2 Form of bargaining

In this section, we introduce the studies that test the efficiency of contract while the objective function of the union as well as that of the firm are specified as the maintained hypothesis. Bear in mind that such tests still depend crucially on the correctness of the maintained hypothesis.

Comparison of the efficiency of contract is analogous to testing the form of bargaining between two parties. In the labour demand model, the union maximizes its utility function with respect to wages while they face the trade off relationship between w and n of the labour demand curve. The result from such bargaining lies on the labour demand curve where the marginal revenue product of labour is equal to contract wage. However, this is not a Pareto efficient point. Both parties can gain by (efficient only in relation to the agents involved) moving away from the labour demand curve towards the contract curve. Pareto optimality is achieved on the contract curve, that is, when both parties bargain over wages and employment. Testing between these models have important policy implications, since, for instance, if a contract is efficient, the trade off relation between wages and employment would be severed.

Before discussing the studies which test the form of bargaining, let us discuss the work by Svejnar (1986) which tests the degree of efficiency within the paradigm of union and firm that bargain over both employment level and wages by solving Nash maximization problem. The literature introduced so far may be called "institutional literature", in which they contained variables that were likely to affect the bargaining power as regressors in the reduced form wage equation. However, they do not generate any explicit concept of the bargaining power nor do they incorporate the fear or the cost of disagreement as influential to the wage-employment determination process. On the other hand, the literature on axiomatic formal bargaining theory recognize these factors but lacks flexibility in accommodating other factors that is required for the empirical purposes. Svejnar has bridged them together and introduced the institutional aspect of the bargaining by allowing such factors to take effect via the bargaining power. As a result, the bargaining outcome is determined by such institutional factors, disagreement utility (status quo utility) and the fear of disagreement, all within the context of Nash bargaining. The disagreement utility reflects a threat point, which is analogous to the alternative wage. The fear of disagreement is a measure of risk averseness and it determines the location of the contract curve. Furthermore, additional consideration of the bargaining power determines the location of a unique solution on the contract curve. The Nash maximand is:

$$(U(n,w) - \hat{U}(\bar{w}))^{\gamma}(\Pi(n) - \hat{\Pi})^{1-\gamma} \qquad (2\text{-}3\text{-}4)$$

which is to be maximized with respect to w and n. U(n,w) is an expected utility function where the individual worker is represented by a constant risk aversion function; $U(w)=w^{\delta}/\delta$ ($\delta<0$ suggests risk averse agent). Under such setting, as long as the membership is more than employment, the outcome is efficient for any values of risk aversion parameter. The first order condition from the maximization yields two nonlinear equations for wages and employment which are estimated by iterative three stage least squares. Data used contains changes in wage levels of 12 major unionized companies between 1950 and 70. For the wage equation, he uses two data that differs in the way wages were recorded. One records wage changes at each bargain, while another reported every wage change including those brought about by the cost of living adjustment provisions (COLA), thus the latter has inflated sample size. Nonetheless, the results are very similar. A test for a strong efficiency hypothesis, where the workers are risk neutral, had failed to be rejected for the most companies although $\hat{\gamma}$ varied substantially across firms. In view of this, he estimates the unrestricted model under Cobb-Douglas production technology with the bargaining power γ that is allowed to vary with exogenous factors. He finds that the union's bargaining power rises with inflation and COLA, while negatively correlated with unemployment and wage/price control policies. These results supports the view that the exogenous factors influence the outcome above the threat point via their effect on the union's bargaining power. Under the context of bargaining over both w and n, this model did not provide a solution to be on the labour demand curve. But this does not mean a rejection of the labour demand model if it is supplemented by the additional assumption that the union is indifferent over the certain level of employment.

The following literature test the efficiency of the contract across different forms of bargaining. Union and firm may bargain over both employment and wages, or may bargain only over wages, in which case the outcome lies on the labour demand curve.

Let us consider how we can differentiate the labour demand model (LDM) from the efficient contract model (ECM) when there is no restriction over the form of bargaining. For the union utility function, $u=u(w,n)$, which is quasi-concave and continuous, and the firm's objective function, which is represented as profits, $\pi=R-wn$ the LDEM predicts that the employment should be set where the value of marginal product of labour is equal to the contract wage;

$$R_n = w \qquad (2\text{-}3\text{-}5)$$

On the other hand, ECM produces a contract curve equation which is a locus of possible outcome pairs, that is;

$$\begin{aligned} R_n &= n\,(U_n / U_w) + w \\ &= (e_{nw} + 1)\,w \qquad (2\text{-}3\text{-}6) \end{aligned}$$

where e_{nw} is the elasticity of employment with respect to wages. Comparing equation (2-3-5) and (2-3-6), it is the marginal rate of substitution that is present under ECM but not under LDEM. This additional term is non-zero only if $U_n>0$ and $U_w>0$. As has been discussed in the former section, when the union is indifferent over the level of employment so that $U_n=0$, efficient contract lies on the labour demand curve. In such a case, the LD and EC are indistinguishable[2]. For the moment, exclude such possibilities so that $U_n>0$ and $U_w>0$. Here, we introduce two representative papers that has tried discriminating these two models. They both used data of ITU in the USA but differs in the choice of criteria with which to differentiate the two. Brown and Ashenfelter (1986) is more restrictive than MaCurdy and Pencavel (1986) in their functional form specification, and they meet the conclusion from the opposite extreme forms of bargaining. Later, we represent other work which also tackled this problem.

Brown and Ashenfelter (1986) first of all, assume the elasticity term e_{nw} to depend on alternative wage, $\bar{w}$, which does not appear in the equation (2-3-5), so that equation (2-3-5) and (2-3-6) are identified. On the basis of this assumption, their test becomes an exclusion test of $\bar{w}$ in the employment equation. They assume reduced form of workers' marginal revenue product to take the following form:

$$\ln (R_n) = \alpha_0 + \alpha_1 X - \alpha_2 \ln(n) \qquad (2\text{-}3\text{-}7)$$

where X is the vector of exogenous variables. Assume either expected utility or Stone-Geary utility function to derive the marginal rate of substitution equation, which, substituted into (2-3-7) forms an employment equation. Then, the ECM and LDEM is differentiated by the existence of $\overline{w}$ in the employment equation. In particular, they have focused on the strong efficiency, an extreme case of the efficient contract, where a level of employment is determined only by the alternative wage. This holds if the workers are risk neutral, that is, if they are characterized as having a rent or some monotonic transformation of rent as the utility function. At this point, it is socially efficient. Even though this is very restrictive on the individual utility, they argue that, collectively, there is an incentive to act as if they are risk neutral by setting up an efficient employment benefit system which completely insures them from a risk of disemployment, even when an individual worker is risk averse. Then, under the world of uncertainty and imperfect information, workers as a group tend to act as risk neutral, and consequently makes otherwise weakly efficient[3] contract strongly efficient. This is so unless there exist some human or technical obstacles which hinder such insurance provision to be made. This argument makes their focusing on the test of strong efficiency, as opposed to just weak efficiency, somewhat less restrictive, since the individual worker does not have to be risk neutral to achieve the strong efficiency as a group. Nonetheless, for the efficiency of a contract to prevail, the weak efficiency has to be at least satisfied.

Their estimation reveals that the level of employment was affected by the contract wage even after controlling for $\bar{w}$, leading to the rejection of strong efficiency. This means the rejection of the ECM if the risk neutrality of the union or the existence of the complete insurance provision were true. On the other hand, if there were no such complete insurance provisions, the rejection of efficiency is probably due to the risk neutrality assumption which is too restrictive in a real world. Moreover, their entire analysis bases crucially on the assumption that the marginal rate of substitution between labour and wage depends on the alternative wage. If this is not the case, employment can be independent of $\bar{w}$ even under the efficient contract[4]. Also, in the course of estimation, they use lagged contract wages as instruments for the current wage. This is only valid if these instruments do not appear in the union's objective function or if there is no lagged partial adjustment mechanism in the formation of union policy.

MaCurdy and Pencavel's (1986) paper does not test the exclusion of alternative wages in the employment equation since their marginal rate of substitution does not necessarily depend on $\bar{w}$. Instead, they test the existence of the marginal substitution term itself in the the stochastic marginal value of labour equation. When there are input factors other than labour, at equilibrium, the ratio of marginal products of the two inputs is equal to the ratio of their prices. And this relation is equivalent under both regimes except for the marginal rate of substitution term, which is only present under the ECM. This term reduces to the elasticity term of equation (2-3-6) in the single input case. They estimate the stochastic marginal value of labour for a very general form of the marginal rate of substitution, namely, a non-linear function of employment and a set of union size and time dummies. They consider the Stone-Geary or rent maximizing union utility with the Cobb-Douglas or translog production function. They derive the marginal product for each input, whose which are substituted in the place of marginal value of labour in the final equation to be estimated. They find this additional term important, which is an evidence against the LDEM. However, this does not mean the acceptance of the ECM. For that, they required further assumption over the form of union maximand.

Oswald and Christofides (1987) has also tested the validity of the LDEM against the ECM in the context of Canadian private unionized sector in manufacturing and trade & services. Given the utilitarian union utility, perfect competitive product market, the reduced form employment is: $n=n(w/p,w/c,\bar{w}/c,p/c)$ under the efficient bargain, where c is consumer price index, and $n=n(w/p)$ under the LDEM. This assumption of the perfect competitive environment in which the producer competes is commonly assumed throughout the empirical studies discussed so far. However, this assumption may be too restrictive, and on this ground, they set up the case where the firm faces the monopolistic product market. Corresponding employment function is: $n=n(w/c,p,c,z)$ for the LDEM and $n=n(w/c,p,c,z,\bar{w}/c)$ for the ECM where z denotes a vector of demand shock. In both cases, a rejection of the LDEM follows if the alternative wage, $\bar{w}$, is found to be significant in the employment equation. Under the ECM, employment equation is a solution from the Nash optimization problem. Here, the bargaining power is assumed to depend on the government wages, incomes policy, negotiating stage of a contract and the outside rate of unemployment. Their data details in wage levels and number of employees at the time of each settlement, hence it is possible to measure the real wages at the beginning of the current contract and those at the end of the previous contract. They introduce a lag structure to the static representation of the employment equation to take into account of the adjustment cost. These lagged variables in the equation are their levels at each contract and not their quarterized or annualized averages. As a proxy to the alternative wage, they use 3 measures; regional wage rate, regional unemployment benefit and regional benefit duration. Consequently, the differenced employment equation that eliminates the fixed effect is estimated. They find weakly significant negative effect of the alternative wage variables in either product market conditions. Moreover, the strong efficiency fails to be rejected. On the other hand, employment equation estimated solely with the lagged product wage performs well with highly significant negative effect which is consistent with the LDEM. On the whole, these results failed to reach any consensus.

Bean and Turnbull (1987) and Card (1986) have also tested the efficiency of the contract by focusing on the significance of variables that affect the alternative wage (i.e., outside variables) in the reduced form employment equation. They also assume that the term e_{nw} in equation (2-3-6) depends on $\bar{w}$. Bean and Turnbull use annual data on British Coal industry and Card uses quarterly data on 7 major airline union in the USA. Under the ECM, contract wage is endogenous, therefore, appropriate selection of the instruments are necessary in order to identify the contract curve. However, Card treats contract wage exogenous on the ground that the lagged employment failed to Granger cause current wages in the autoregression of wages. As Bean and Turnbull points out, Granger non-causality is not sufficient nor necessary for the weak exogeneity required for the estimation of employment equation. Brown and Ashenfelter, on the other hand, has used the lagged contract wages while MaCurdy and Pencavel relied on the location dummies and time trends. Identification is ensured only if the instruments do not come into the union's utility function, while at the same time, affect the bargaining power. Therefore, Bean and Turnbull have chosen the variables that affect the status quo utilities of both parties as instruments. For a given production technology and utility function, they estimate reduced form labour demand equation using the appropriate instruments and find the outside variables such as benefits or manufacturing wage significant. However, they argue that if the labour market tightness (i.e., $\bar{w}$) predicts future contract wage and employment depends on future as well as current wages, then the significance of alternative wages in the static employment equation may not necessarily mean a rejection of the LDEM. When such intertemporal relation exists, alternative wage can affect the level of employment not only directly but also indirectly via its effect on the future contract wage.

Hence, they introduces a possibility of such intertemporal relation by imposing a decision making on the firm whether to mine now or later, on the basis of the exhaustiveness of coal. The union is assumed to remain having a myopic view while the firm maximizes the expected present value of profits subject to non-negative unmined reserve of coal at any time. They then derive and estimate the Euler equation. The instrument are any variables belonging to the information at (t-1) by rational expectation assumption. This enables us to see if the outside variables are correlated with employment such that they help determine the future contract wages. Results are rather consistent with the ECM prediction with the significant alternative wage variables.

Card has also introduced such dynamics by considering adjustment cost in the firm's objective function that include hiring, firing cost and cost of rearranging schedules. The firm minimizes the expected present value of total cost subject to the expected utility requirement of workers, and derive a reduced form employment equation where such adjustment cost yields a serial persistence term. If it is the labour demand model, so that the alternative wage has only indirect effect on n via future contract wage, then $\bar{w}$ and the contract wage should both have positive signs in the employment equation provided that the future contract wage and alternative wage is positively correlated. On the other hand, if the model is strongly efficient, so that there is only a direct effect of alternative wages on employment, it's effect on n should depend on the role that product price plays in forecasting the future contract wage. However, $\bar{w}$ and w are found to have opposite signs in the employment equation. And $\bar{w}$ reversed its sign when the product price was also introduced in the wage forecast equation. Hence, these two polar models are both rejected against a general unrestricted model where the employment depends on w as well as $\bar{w}$. Their point estimates suggest negative effect of alternative wage and positive effect of contract wage on n, although neither of them are significant. On the whole, the reduced form employment equation derived from the formal bargaining model has not performed better relative to the unrestricted model of auto regression.

The significance of alternative wage found in these models, however, does not yet mean an acceptance of the ECM. It may well be due to misspecifications of the production technology or the union's utility function. These results are consistent with the bargaining over man-machine ratios, in which case, the solution will be off the contract curve. Moreover, other theory of wage-employment relation such as the efficiency wage predicts that the discrepancy of w from $\overline{w}$ is necessary to prevent shirking, to reduce turnover or to induce effort, thus, leading to productivity growth. Under this theory, improvements in the outside opportunities will reduce productivity and increases the number of quits. Employment in such case will depend on the alternative wage even if bargaining does not take place over the level of employment. Nickell and Whadwani (1987) gives a similar explanation for their finding of the positive alternative wage effect in the employment equation. Their employment equation uses 219 U.K. firms over the period 1974-82, is particularly rich in financial factors by considering that the employers should sensitive to the costly bankruptcy. They argue that their finding is explained by the LDE model modified to allow for the efficiency wage considerations. As $\overline{w}$ increases, the employer raise w as long as the gain from increased productivity outweighs the labour cost. As can be seen, these considerations cast further doubt to the rejection of the LDEM merely on the basis of significant alternative wages in the employment equation. This is making it more and more difficult to empirically test between the LDEM and the ECM.

So far, we have been discussing the empirical studies that test between the labour demand and the efficient contract model of the union-firm bargaining. Discrimination between such models is possible for the utility function that has $u_w > 0$ and $u_n > 0$. As has been pointed out previously, when the union is indifferent to the level of employment (i.e., $u_n = 0$), this discrimination becomes impossible since the labour demand curve coincides with the contract curve. Let us call such model where the union is indifferent to the level of employment as the Flat Indifference Curve model (FIC) (Carruth, Oswald and Findlay (1985)). Some of the rationales for the FIC model are as follows: (1)layoffs take place by the order of inverse seniority (or of any known order) in the majority voting union, (2)there is a distinction between members and non-members and that the members are only interested in their own employment status (see section 5). This model provides many predictions that are not given by the previous models, in particular, the existence of wage differentials between workers in different plants despite controlling for their non-pecuniary characteristics. Such differentials are, instead, explained by the financial performances and the market conditions.

Carruth, Oswald and Findlay has tested the behaviour of the union at the flat segment of the indifference curve against the LDEM by estimating the reduced form wage and employment equation. Direct estimation of the reduced form equation avoids making any *ad hoc* specifications of the agents' objective functions. At such point, the variables that affect alternative wage should not affect the union's indifference curve since the employed members are indifferent to the welfare of the unemployed. They test the exclusion hypothesis of variables such as unemployment benefit and unemployment rate in the reduced form wage and employment equation. But using the annual data on British steel and coal industries for the post war era (1950-80), these variables are found significant, leading to a rejection of the FIC model. One of the explanation to this finding is the outside wage that was used to capture the comparison effect in the union utility. It is very likely that the average wages and unemployment benefit are collinear, hence, making it difficult to disentangle their effect in the wage equation. In other words, union's indifference over employment does not necessarily mean the absence of alternative wage effect in the wage equation since they may affect the union's fall back wage. Moreover, the unions they have chosen may not have been the most appropriate candidate for the FIC model, since the steel industry was experiencing massive layoffs at a time and the coal industry is regarded as having one of the most altruistic union.

Similar finding can be seen in Blanchflower, Oswald and Garrett (1988) and Beckerman and Jenkinson (1989) who have constructed a simple model of FIC for the union which is indifferent to the level of employment. Assume a risk neutral union so that their utility is linear in wage, and the bargaining outcome maximizes a Nash formula:

$$\max_{w} \quad (w-\bar{w})^{\sigma}(\pi-\bar{\pi})$$

where $\pi = \max_{n} R\text{-}wn\text{-}k$. The first order condition gives:

$$(\pi - \bar{\pi})*\sigma + (w - \bar{w})\,\pi_{w} = 0$$

while $\pi_w = -n$ by duality, hence:

$$w = \bar{w} + \frac{\sigma(\pi - \bar{\pi})}{n}$$

Thus, the equilibrium wage is a weighted average of fall back wage and profit per employee. As this simplified case suggests, one can view the wage determination as a kind of rent sharing where the workers' bargaining power is influenced by the external labour market conditions. Hence, pay depends not only on the external pressures that affect $\bar{w}$, but also on the internal pressures, that are, profitability of the firm including the market power and the market condition of the firm. And the relative importance of these two components depends on the power of insiders as opposed to that of outsiders. Note that this argument holds also for the non-union environment as long as insiders are not so close substitutes of outsiders.

Their regression estimates of the wage equation on internal and external variables have found both of these factors influential. In particular, Blanchflower, Oswald and Garrett have shown that internal factors do not have strong impact amongst the unskilled non-union sector; the sector with easily replaceable workers. As representing outsider pressures, they have found the regional, global unemployment, or the outside wages significant in determining the contract wages. On the other hand, since this theory assumes the insiders who only care about the size of their pay so as to set the expected employment to equal the union's membership, any increase in demand tends to translate into a wage gain rather than to an employment gain. This, in the empirical wage equation, should give rise to a positive industry or firm specific unemployment effect. This causes persistent unemployment despite the fluctuations in the aggregate demand; the hysteresis effect in the economy. Its pure insider view predicts their wage setting to be completely independent of the outside labour market conditions (Blanchard and Summers (1987)). In the work of Nickell and Whadwani (1989), they use the disaggregated firm level data and also found the role of insider and outsider pressures significant in their wage determination. In particular, they have found the hysteresis effect, as insider variables, that strengthened for the firms with decentralized bargaining. As for the outsider variables, they have found variables representing the state of labour market, particularly, the level of aggregate unemployment and the proportion of long term unemployed playing a significant role in the determination of contract wages at the firm level, therefore, rejecting the pure insider view. In

Beckerman and Jenkinson, they have found the significant negative aggregate unemployment and positive industry specific unemployment effect on the industry wages; a finding also consistent with the insider-outsider hysteresis. However, Nickell and Whadwani (1987), also using the firm level data, have found both aggregate and industry specific unemployment to depress wages. They resort rises in average wages to the extent that higher unemployment level boosts productivity, rationalised within an efficiency wage framework, and the rise in the long term unemployed.

2.3.3 Conclusion

It seems that the unions do care about wages as well as employment, but their exact form is still unknown. Based on the certain form of agents' objective functions, both the strong efficiency and the pure labour demand model are often rejected, although this does not necessarily mean the acceptance of the weaker efficient contract. Their test is mostly confined to the existence of alternative wages in the employment equation which has to be derived from the particular utility and production structure. Hence, it is based on the maintained hypothesis that these objective functions are correctly specified, but that may not be the case. On the other hand, the pure insider-outsider or the seniority theory implies the observationally equivalent employment function under the LDEM and ECM. And it seems to be able to explain the stylized facts that the union and employer both bargain over wages but only the employer retains the power to determine employment except in the case of massive layoffs (i.e., bargaining takes place on the labour demand curve). Still, empirically, the pure insider model where the outside labour market conditions becomes irrelevant, seems too simple. This may invalidates their assumption that workers do not care about employment, but introducing the possibility of comparability (relative wage) effect in the union's utility function can help avert the rejection of the model. On the whole, the more sophisticated version of the labour demand model which takes into account of considerations such as the efficiency wage theory seems better able to explain the wage-employment nexus. Last point: unions located in different cities have shown substantial variations in their objective functions, hence, empirical studies of bargaining seem to crucially depend on the union's institutional and local labour market characteristics.

Footnotes to chapter 2

1. Under general equilibrium, contract wage is equal to outside wage. (pp.16)
2. However, even if one believes that such situation may arise due to insider-outsider consideration or LIFO, there still is a ground for testing such framework against the LDEM. For as long as the level of employment is less than the level of membership, outcome should lie on the contract curve and not on the labour demand curve. This means that testing between LDEM and ECM by exclusion condition may be more applicable for the industry which is declining than that of expanding. (pp.31)
3. Weak efficiency means that the level of employment is affected both by contract wage and alternative wage.(pp.32)
4. Example of such quasi-concave utility function is: $u=(w/\bar{w})^{\alpha}\phi(n)$ where $\phi(n)$ is some increasing function of n. In this case, value of marginal product is independent of n, so as the marginal rate of substitution. (pp.32)

Chapter 3: The Data

3.1 Aberdeen database

The primary data to be used in our study is the "Aberdeen wages rate database" which was constructed at the University of Aberdeen by Elude, Steele and Bell (1977). Figures are collated originally from the Department of Employment publications, "Time rates of wages and hours of work" and "Changes in rates of wages and hours of work", where the exact level of wages were mainly collected from the former source. The data contain the details of all the wage settlements of the 191 largest national negotiating groups of manual workers between 1950 to 1975 inclusive. Those groups had to be covering at least 5000 workers some point in the sample period. They are divided into 2 files.

3.1.1 The agreement data

The agreement data file details information on all of the 191 agreement titles where the agreement title being the name of each bargaining group. They include the bargaining system employed, the trade unions involved, the geographical area covered and the number of workers covered. In general, each agreement title involves more than one trade union and the same trade union often appears in more than one agreement title. The file lists the outcome of centrally determined arrangement only and the changes negotiated at district, establishment or shop floor level are not reported. The bargaining systems are categorized into wages council, Whitley Council, local authority, nationalized industries, or private sector employers and others. This is important since each can face a very different bargaining environment. For example, public sector, in general, incorporates more sophisticated bargaining process and tends to be influenced more frequently by the government controls or arbitrations.

3.1.2 The settlement data

The settlements file contains information regarding each settlement for all bargaining groups over time. The data include settlements made up to 1975, where applicable, although the data start at different dates for different groups owing to the nature of each agreement. Note also that a number of settlements is not equal for each group which gives rise to the unbalanced nature to our panel data. They contain, for each title of the agreement, the date of implementation and settlement. These dates sometimes differ, particularly for the staged implementations. Any set of implementation dates that constitute stages have the same settlement dates but the converse is not necessarily true. While the former is always reported, the latter date is mostly not available. The file lists rates of basic wages or minimum entitlement and normal weekly hours negotiated at each settlement. Series of hourly rates were calculated as a weekly figure divided by normal weekly hours. Wage rates recorded are those of the highest paying district outside London. Wages reported of manual workers are categorized into several types, namely, top male, semi-skilled male, bottom male and female. They also contain several codings. First coding gives a detailed sources of government intervention or arbitration whenever they interfere with the negotiations. There are 8 categories of interventions and those are namely, National Arbitration Tribunal & Industrial Disputes Tribunal, Industrial Court or Industrial Arbitration Board, Court of Inquiry, Ad hoc Arbitration Boards & Single Arbitrators, Public Sector Arbitration Board, Committees of Inquiry & Investigation, National Board for Prices & Incomes and lastly, other forms of intervention. There is also a binary series which indicates the presence of indexation clause. When applicable, a wage change takes place in accordance to some formula based on the movement of retail price index rather than as a result of negotiation. There are very few groups who incorporates COLA clause as principal means in determining wages. We can identify these groups since the majority of their wage changes are coded as COLA induced. Basic wages are recorded in the data set in most cases, although the data identifies when only the minimum earning levels are available. The data also identifies any wage changes implemented as a part of long-term staged settlement. A staged settlement is part of the multiple wage changes that stems from one negotiation. The data records the dates of all wage changes as long as they arise from the different negotiations. With respect to the staged settlements, therefore, only the last implementation is recorded together with the wage changes that are aggregated to this last date. However, when the stages are implemented over more than one calendar year, they are recorded each year as separate settlements. In such cases, the coding distinguishes whether the stages were implemented across different calendar years or they all took place within one calendar year.

There are few information that we would like to know but were not available in this data set. The details of every implementations, when and by how much the wages have changed are crucial in our study of wage determination process. Although a distinction is made for each individual bargaining whether it is a part of staged settlements, it is not often possible to identify the date of the original settlement. Sometimes unplanned bargain takes place between the staged settlements. Different sets of long-term staged settlements may overlap with each other. Each consecutive staged wage changes are not necessarily coming from the same original settlement. It is then difficult for us to clarify the starting and the ending of a certain staged settlement. This is because of the nature of this particular coding employed in this data, and also, the settlement dates that are not available most of the time.

Secondly, the data set does not record every stages if they take place within the same calendar year. Hence, when the multiple stages were implemented within one calendar year, we need to unearth the details of every stage which has occurred in that same year prior to the last implementation date. Aggregation of wage changes to the last implementation date clearly understates the number of actual settlements occurred to this bargaining group over time, while, overstates the magnitude of wage changes at this last date. Same thing applies to the wage changes triggered by the indexation clause. For instance, under the threshold arrangement governed by the Counter-inflation order, which took place mainly from May 1974 till the end of the year, same amount of special payments, determined every month according to the movement of price index, were made to all the groups which joined the policy. In the data set, only the last implementation date in November 74 is recorded together with the wage rates that include the sum of the threshold payments summed up to this last date. In the process of retrieval, we assumed such special payments to have been incorporated into the basic pay on the date of the payments.

Thirdly, from the early 70's, a movement which aimed to achieve the equal pay for male and female engaged in equivalent jobs have prevailed throughout the industries. This was initiated by the bill presented in 1970 to the Parliament in which it recommended such objectives to be achieved by the end of 1975. During these period, there often were stages aimed only for female workers or sometimes, female workers and juvenile workers. It was considered, therefore, useful to distinguish such staged settlements aimed for particular workers from the rest of staged settlements.

Lastly, any planned wage changes known to be deferred due to the standstill arrangement need to be identified.

For these reasons, it was necessary to go back to the monthly Labour Gazette from which these data were originally collated. The additional information extracted are then summarized into the following new seven series, created for each wage change.

3.2 The new series

ista: 0:not staged, 1:staged, 2:staged according to the data but the gazette does not specify, 3:gazette specifies as staged while the data does not.

It indicates if a change in hourly wages is a part of a long run staged settlement. This distinguishes the case when it was recognized as staged only in the gazette but not in the original data, and vice versa.

insta: 0:not applicable 1:1st stage, 2:2nd stage, 3:......

Given that this wage change is a part of staged settlement, this variable states the stage of this particular implementation.

idfix: 0:not applicable, 1:not known, 2:staged but the date is not planned, 3:staged and it took place at planned date, 4:staged and had a planned date but altered.

This variable indicates whether the date (up to a month) of this particular implementation were planned when this settlement was first agreed, and if they were, whether they in fact took effect as planned. When the agreement is amended, the succeeding stages are treated as stemming from the separate settlement. However, since the first stage of amended settlement was planned at the beginning of the original settlement, it is coded as 3. Rest of the time, idfix=1 for the first stage of any staged settlement. When the plan was altered due to the standstill arrangement, idfix=4, and it is further indicated in another series called ifrez.

ifix: 0:not applicable, 1:not known, 2:staged but the size is not planned, 3: staged and the size was realized as planned, 4:staged but the size was not realized as planned

This variable states if the magnitude of wage change was planned at the initial agreement and if so, whether they actually took effect as planned. I assumed all of the following cases to have the planned wage changes:(1)wage changes planned in terms of a percentage of previously determined wage, (2)wage changes planned so as to achieve equal pay with the compatible male workers, (3)wage changes planned so as to equate the current minimum earnings level to the basic minimum wage, (4)hourly wage changes to take place as a result of planned changes in hours worked without a loss of pay, and this change in weekly hours were planned.

ifem: 1:staged implementation for female workers, 0:no

This is a binary variable indicating whether this implementation is a part of the staged settlement for female workers as a movement towards equal pay. This includes the case when the wage changes take effect both for male and female but the gazette states that only the female wages were planned. In most cases, staged wage changes for the female workers are also designated towards the juvenile workers who are mostly male workers.

ifrez: 1:deferred due to freeze policy, 0:no

This is a dummy variable indicating that the planned implementation date was deferred in accordance with the standstill arrangement.

ich: 1:working hours change, 0:no

This dummy variable states whether this wage change took effect as a result of planned change in hours worked without a loss of pay.

3.3 Series of exogenous variables

(1)Retention Ratio: Retention ratio was calculated as the ratio of disposable income to the income before tax, both derived before providing for the depreciation and stock appreciation. They are seasonally adjusted figures. The definition of personal disposable income is the total personal income before tax less tax on income, less national insurance and health contribution, less remittances abroad. Both series were taken from the appropriate series in Economic Trends (ET) which gives annual figures from 1946 to 1954 and quarterly figures from 1955 onwards up to 1975.

(2)RPI: Monthly, seasonally unadjusted figures for the general index of retail price on all items (as compared to that of foods or wholesale goods) are available from January 1949 to 1975 inclusive (1962 Q3=100). The monthly figures from mid 1947 to 1949 were derived from the series of interim index of retail prices from the Monthly Digest of Statistics (MDS) which was the only price index available for the pre 48 era. Figures for the later years are collected from ET. They were made consistent to the series of 1949 onwards through the figures in the overlapping periods.

(3)Unemployment: Quarterly figures for national unemployment levels in GB are extracted from British Labour Statistics (BLS) between 1948,Q2 and 1975 inclusive. They are the actual numbers in thousands who are registered as wholly unemployed excluding school leavers in GB and are seasonally unadjusted.

(4)Productivity: Annual index from 1950 to 59 and quarterly index from 1960 to 75 of output per head are extracted from ET (1960=100). They are the gross domestic product at constant prices (output-based) divided by total number of labour force which includes not only employees but also employers and self-employed persons. This series, "output-per-head" was first introduced in 1968, at the time, the series was re-calculated from 1950. The figures prior to 1950 were computed from the existing series of total labour force and index of gross domestic product according to the definition of this series. There was a problem since the series of total employed labour force changed its definition and began including private indoor domestic services since October 1948. From the overlapping periods, we deduced the numbers engaged in such service to be 2850 and computed the corresponding figures for pre 50 periods.

(5)Employment: Total number of employed labour force in thousand defined as total civil employment plus H.M. forces in GB. This is seasonally unadjusted and collected from BLS.

(6)Earnings: The level of average weekly earnings in pence are collected from MDS (originally from Labour Gazette). They are general averages covering all classes of manual workers, including unskilled, general labourers, and operatives in skilled occupation. They represent the actual earnings per week, including payments for overtime, night work, and those provided by payment by results. They also include non-contractual gifts and bonuses paid that are averaged out to give weekly figures.

(7)Wholesale input and output prices: Quarterly, seasonally unadjusted wholesale input price index (on materials and fuel purchased by manufacturing industry) and wholesale output price index (on all manufacturing products) are available from BLS since 1954 (1975=100).

(8)Unemployment per industry: Industry specific unemployment figures are available only for the manufacturing industries. They are "numbers of persons registered as wholly unemployed, by industry order" which were available quarterly from BLS.

(9)Profits per industry: Trading profits per industry is available only for the manufacturing industries from National Income and Expenditure. There are few difficulties in treating this series as consistent throughout our sample period. First, figures are only available according to the 1948 SIC coding for the earlier period. But their categories do not exactly match with the recent SIC. Second, figures are derived based on the tax assessment. Problem occurs since the particular tax system, on which the trading profits measure depends, has changed in 1965 from the profit and income tax to the corporate tax. The corporate tax is levied on total net income including non-trading income after deducting for capital allowances. But prior to 65, income and profit tax was levied on the gross trading profit. Hence, figures for later years are subject to wider margins of errors to the earlier figures. Nonetheless, there was no

variable	mean	std dev	min	max
rpi	108.35	40.04	55.3	264.3
retention ratio	0.842	0.029	0.761	0.876
earning	15.65	10.04	4.95	49.79
productivity	105.20	21.46	63.1	144.3
unemployment	419.46	168.20	162.5	928.6
employment	24385.1	801.83	22965	25358

(Table 3.1.A) Descriptive statistics

other series which reflects industry specific economic situation throughout our sampleperiod, hence, we had no choice but to resort to this series. In particular, we use the "gross" profits, that is the value before deducting for depreciation allowances.

These series are only available discretely, although they are supposed to be continuously changing over time. In particular, corresponding values of these series at the week of every negotiation are necessary for our analysis. We therefore interpolate the available data points linearly, constructing the time path made out of linear segments. This enables us to assign the weighted average of adjacent available values

industry	variable	mean	std dev	min	max
Food, drink	profits	413.34	131.52	195.0	641.5
	unemp	17.26	5.50	7.8	33.6
Chemical	profits	338.96	147.26	128.0	879.3
	unemp	7.18	3.11	3.2	16.9
Metal	profits	208.65	62.18	79.9	311.5
	unemp	9.78	7.98	2.7	48.4
Engineer	profits	571.90	228.84	250.0	1141.2
	unemp	26.59	15.81	9.7	83.4
Ship bldg	profits	22.93	8.58	10.1	42.0
	unemp	10.22	3.53	4.9	23.4
Vehicles	profits	193.27	64.0	107.0	437.2
	unemp	10.29	7.38	3.7	5.0
else Metal	profits	133.89	44.78	64.0	258.4
	unemp	9.95	6.48	3.3	34.1
Textile	profits	199.64	48.14	127.3	376.0
	unemp	14.67	6.28	5.9	34.0
Leather	profits	12.45	4.59	7.0	26.9
	unemp	1.10	0.41	0.5	2.4
Cloth,foot	profits	62.57	22.49	30.9	127.3
	unemp	7.87	2.68	3.8	18.1
Bricks,pottery	profits	129.91	79.40	46.0	340.9
	unemp	6.17	2.75	2.2	15.3
Timber,furniture	profits	37.93	24.70	20.1	128.3
	unemp	5.65	2.15	2.7	11.7
Paper,printing	profits	191.28	74.17	89.3	425.8
	unemp	5.77	3.28	1.8	16.7
Others	profits	77.55	35.98	31.5	159.5
	unemp	5.70	2.89	2.2	16.8
Construction	profits	187.21	119.33	46.0	496.1
	unemp	72.95	37.53	21.9	173.7

(Table 3.1.B) Descriptive statistics: profits and unemployment per industry

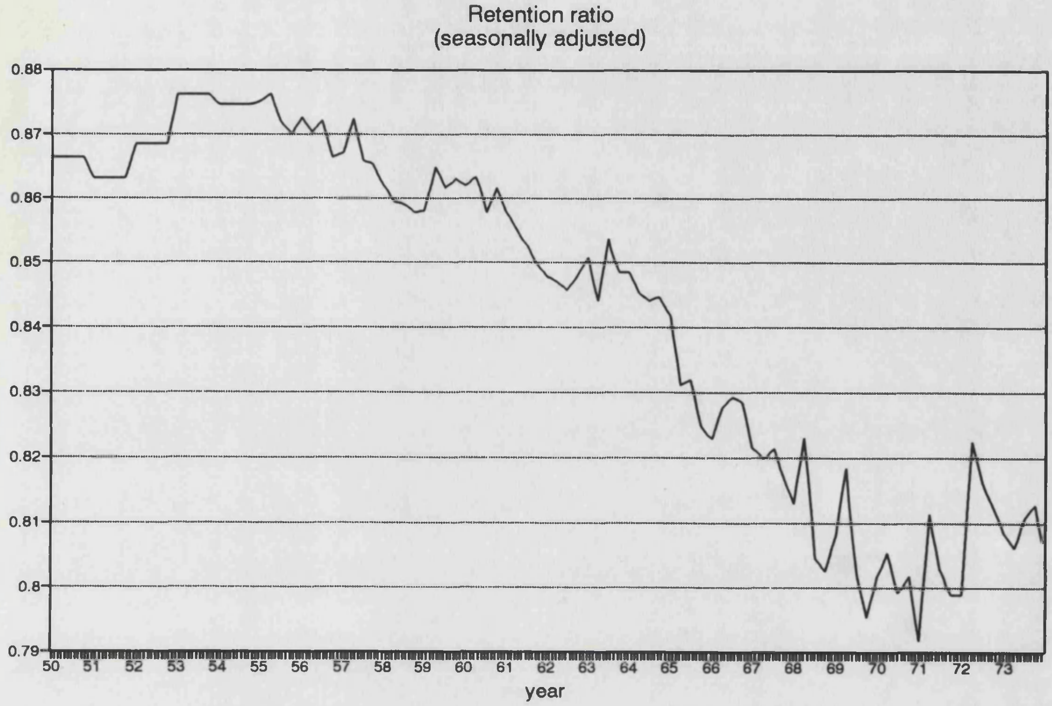


Figure 3.1

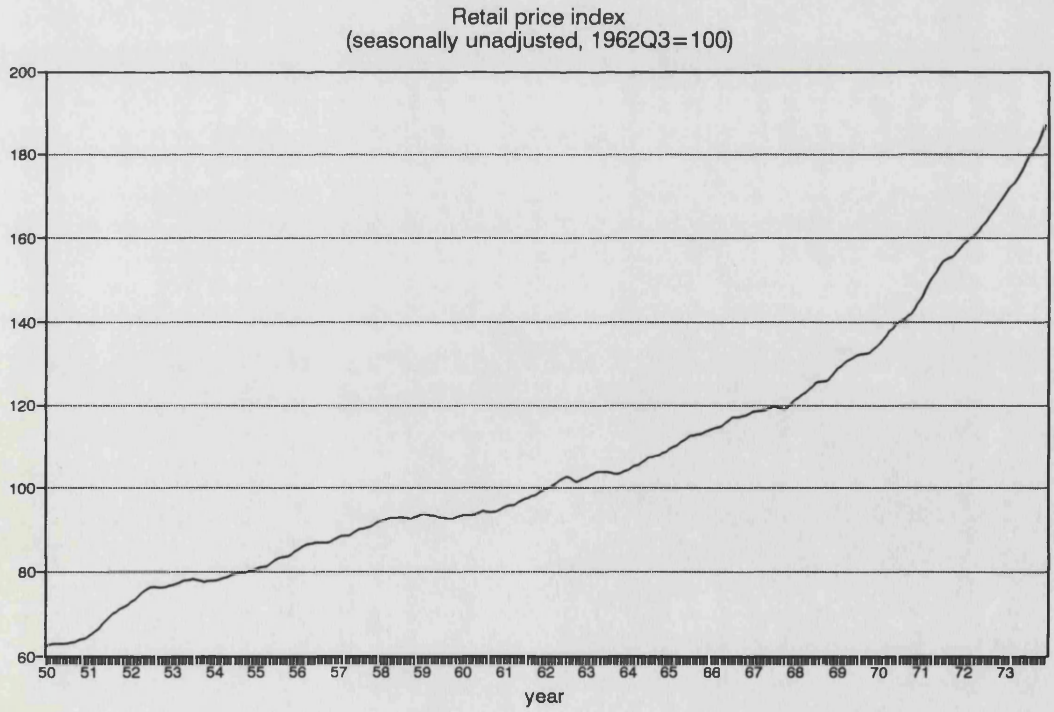


Figure 3.2

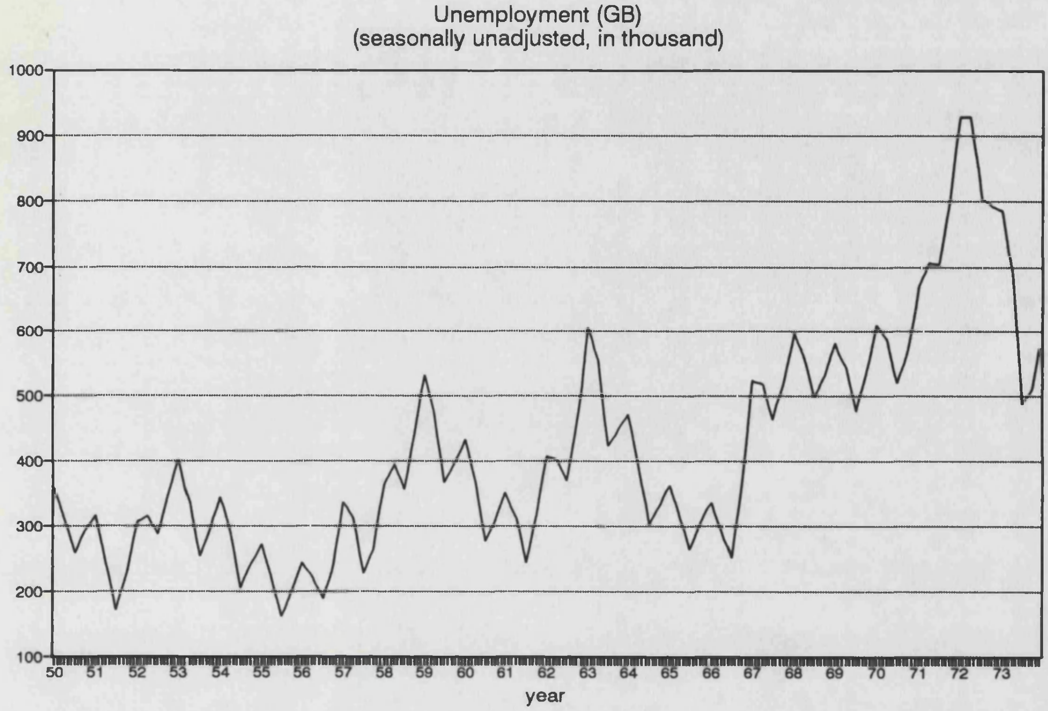


Figure 3.3

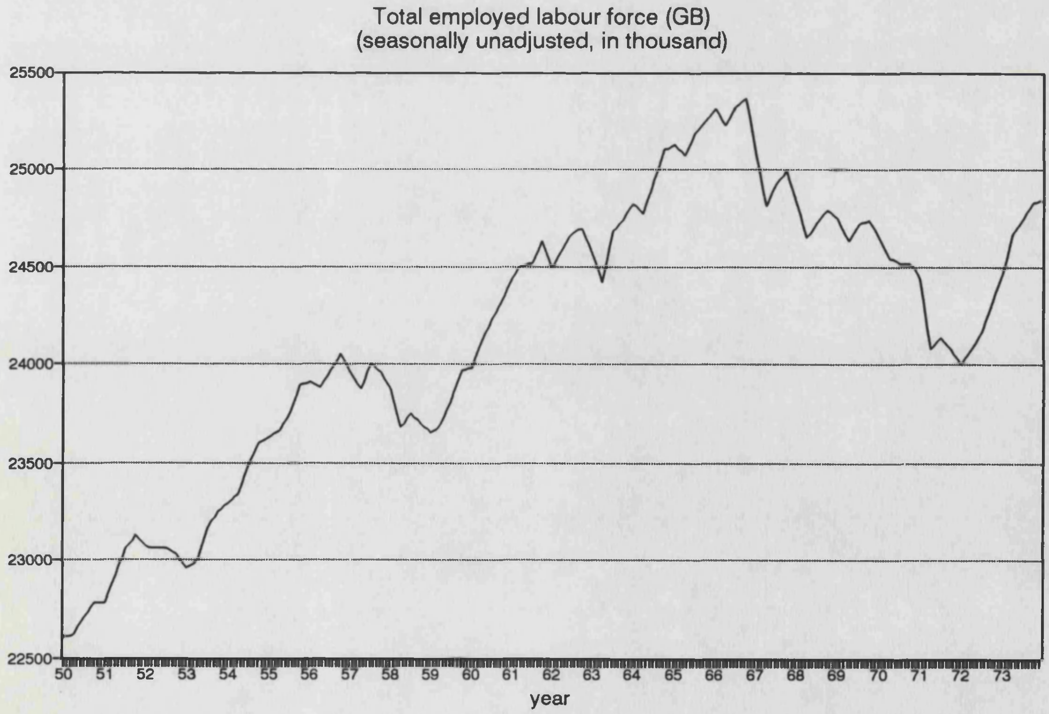


Figure 3.4

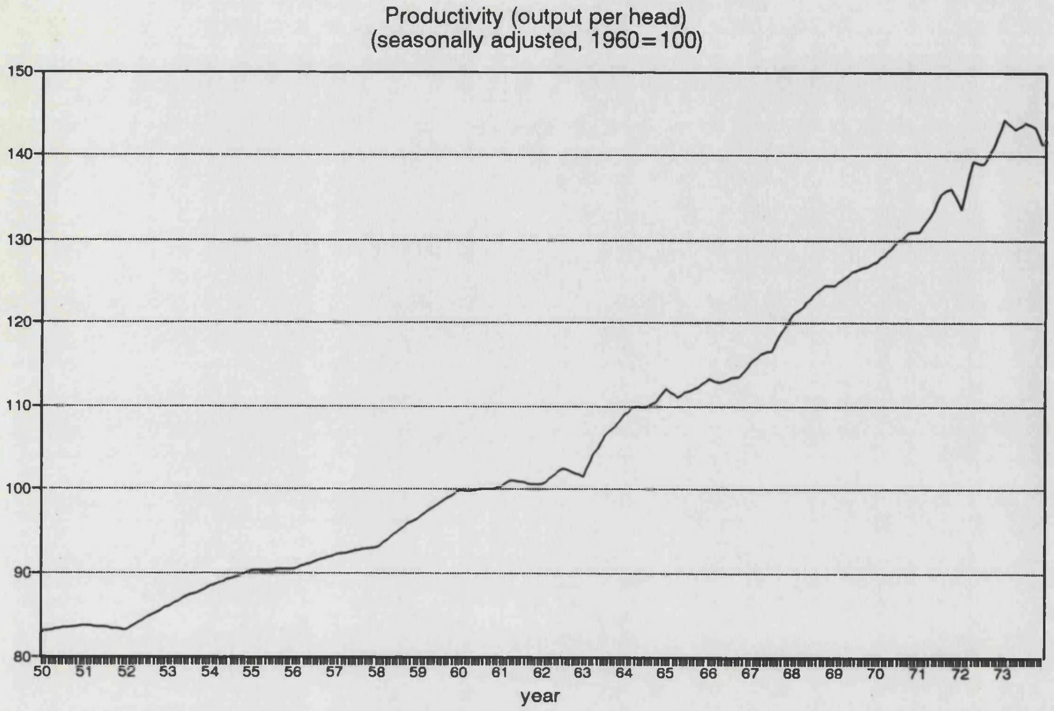


Figure 3.5

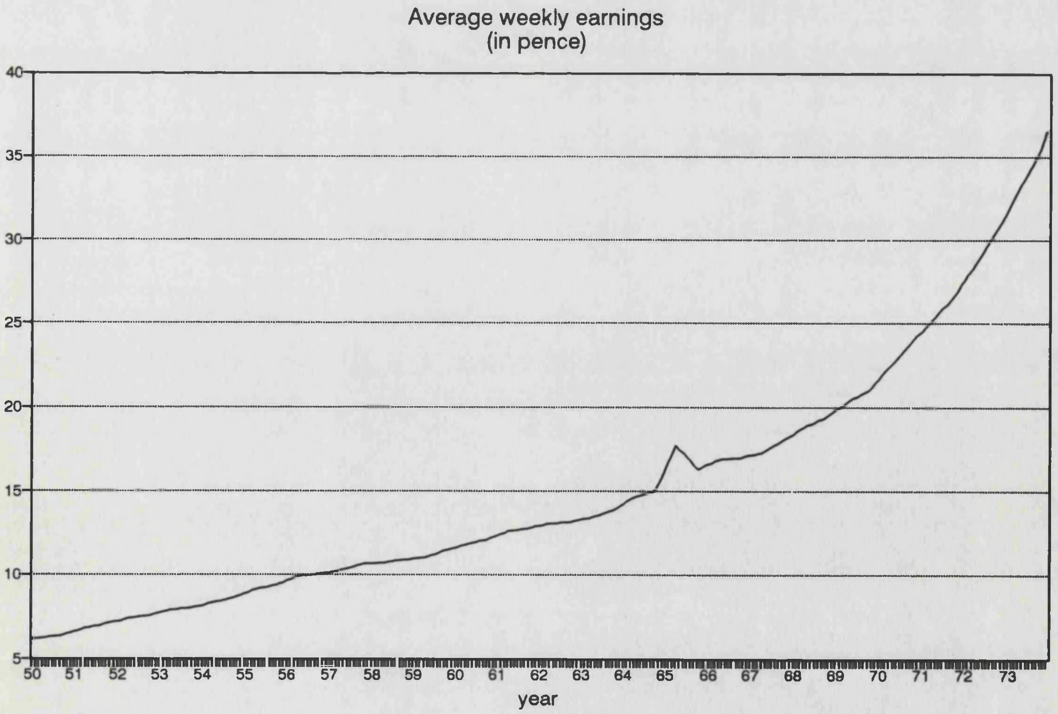


Figure 3.6

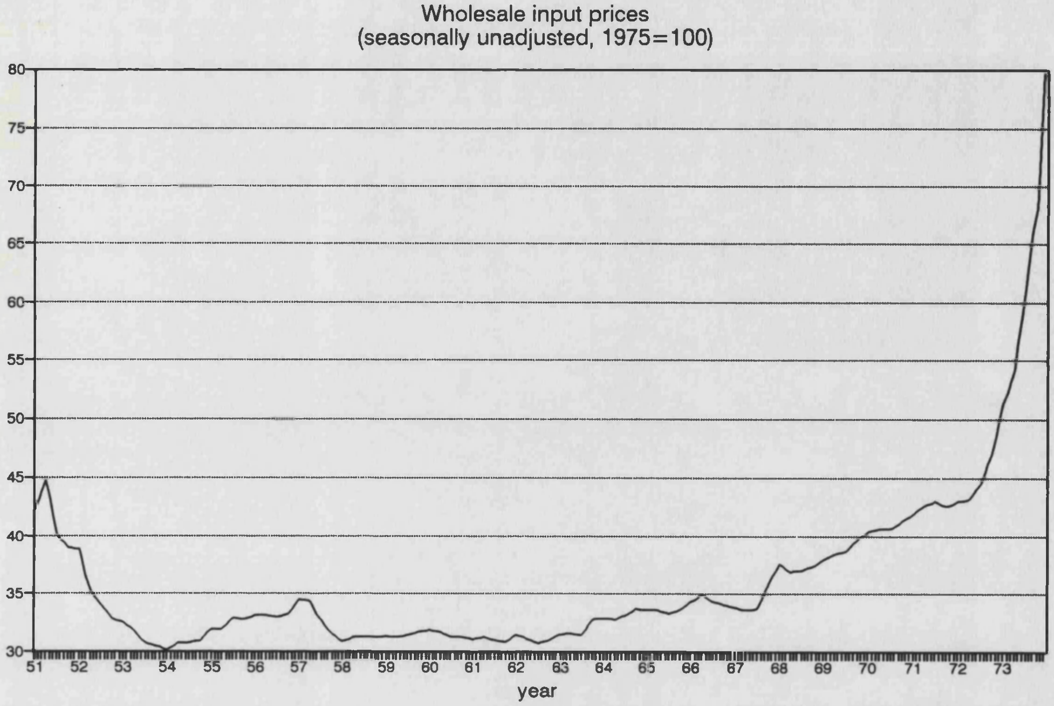


Figure 3.7

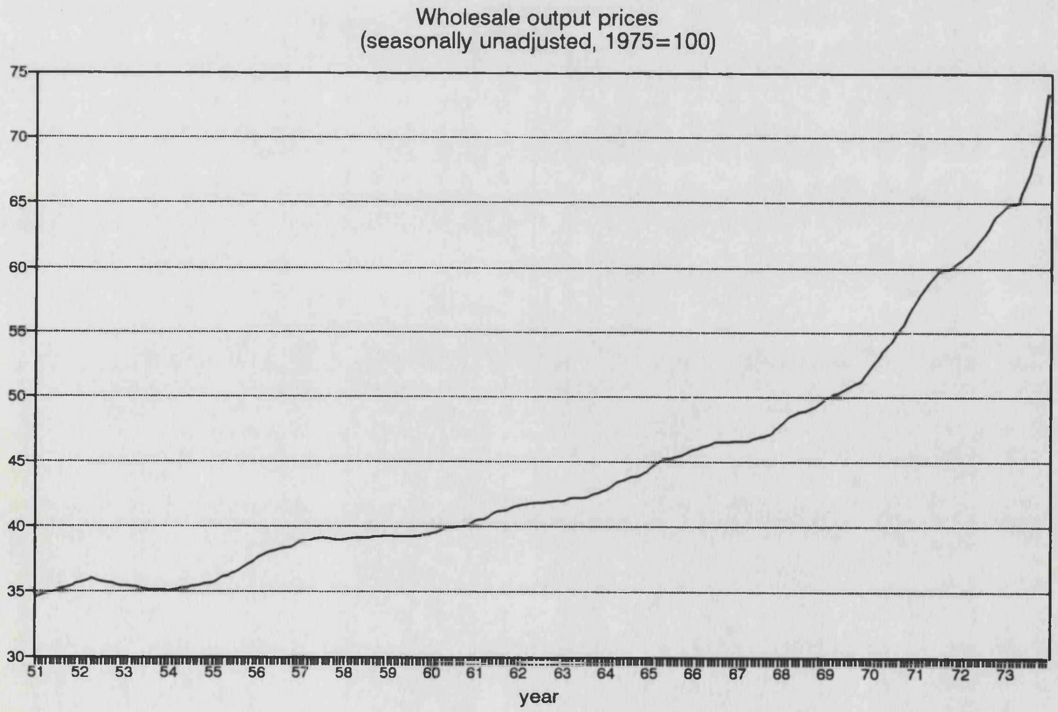


Figure 3.8

to be the corresponding value for the each implementation date recorded in the data set.Their movements overtime are plotted in figures (3.1)-(3.8) and their descriptive statistics are listed in table 3.1.A and 3.1.B.

3.4 Importance of National Agreement during 1947-1975

Increase in the size of earnings drift, the difference between the actual earnings and the basic wage rates, since the war drew concern of the Donovan Commission who blamed the cause of such gap to be the growing popularity of informal shop-floor bargaining. They argued that this phenomena has led to inefficiency by providing a ground for the shop stewards to exploit the plant-level institutions with their own interests which are not necessarily justified in the economic ground (Elliot, (1976)). They recommended to institutionalize the plant/company level bargaining so that the wage setting can keep stronger link with the level of productivity, therefore, leads to higher efficiency. In 1973 and 78, New Earnings Survey published the details of a number of workers affected by the various types of collective agreements, namely: national agreement only, national plus supplementary company-district-local agreement, company-district-local agreement only and no agreement. The actual question which was asked in the survey was, 'please indicate the type of negotiated collective agreement, if any, which affects the pay and condition of employment of this employee either directly or indirectly'.

Comparing the figures between 73 and 78, there indeed is a small but clear trends towards the disaggregated level of agreements (i.e., company, district, local agreements), particularly in manual groups. However, 72% and 64.6% of manual male workers were still affected by some form of national agreement in 73 and 78 respectively,(that is to say, they belong to either first or second categories above) indicating quite a substantial coverage.

Another source of information is the data from 1973 Department of Employment publication on their estimates of number of manual workers covered by national agreements and wages council boards. The coverage by such agreements has increased substantially from 48% to 65% over the period of 1950 to 1972. The percentage figures were computed by Elliot and Steele (1976) where the number of manual workers were taken from the seasonal population census. The proportions of coverage were then computed within each Main Order Heading (MOH). Although figures were adjusted to 1968 SIC basis, during the time of rapid employment decline, the number of coverage are likely to be overstated due to lags in adjusting the figures from the census, resulting in more than 100% as a proportion of coverage for these times. Provided that we take this into account, these percentage figures still provide useful information.

These two data sets are obviously not comparable to each other, basically due to difference in the method of deriving their information. Also, NES's figure excludes part time workers whereas the latter includes them. The former records the numbers *affected* while the latter records the numbers *covered* by the national agreements. Number of those merely *affected* can well be higher than those *covered*. And lastly, the first and the second category in NES can together cover wider agreements than the category of the "national agreements" in the latter data. Overall effect of these differences are not clear, although the chances are that the coverage in the NES data be biased towards overstatement.

Given these evidence, it is safe for us to say, that as far as our sample periods are concerned, the coverage of national agreements is significant.

Consider, now, the importance of basic rates out of the level of earnings. Donovan Commission blamed the increasing popularity of the informal shop-floor bargaining to the growth of earnings drift. Earnings' biggest component is the basic rates and the rest is mainly divided into shift payments and payment made under the payment by result scheme (PBR). New Earnings survey in 1968, 70 and 73 (with adjustments) has published the details of the earnings make-up for the full time males covered by major national collective agreement and wages council order. This enables us to decompose the earnings so as to derive the percentage of basic pay out of standard weekly earnings. According to them, although there are different tendencies for the public and private sectors over time, the basic rates were never less than 80% of the earnings for those periods, and was the growing factor of earnings for most of the industries. Payment by result scheme is essentially connected with productivity so that we should find the appropriate effect of productivity change on the earnings drift. The evidence shows that the average size of payment made under PBR has increased in most cases while the number of workers who received such payment have decreased. Elliot and

Steele argued that the latter effect outweighed the former, hence, reducing the overall size of wage gap overtime.

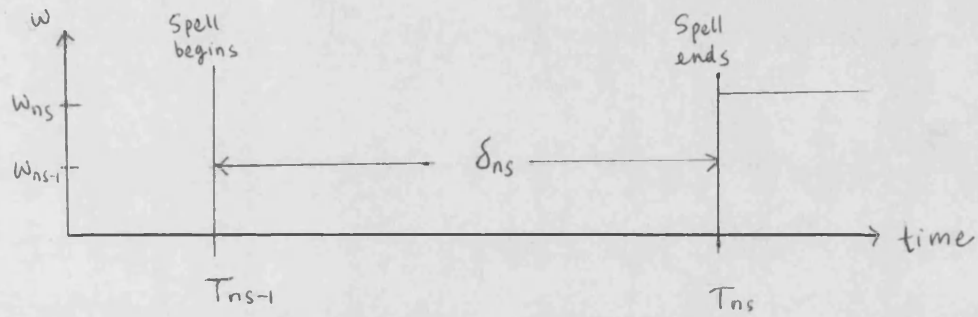
We have shown the followings: 1)the large proportion of earnings accounted for the basic wage rates negotiated at the national level, and 2)the coverage of national agreement/wages council has increased from 1950 to 1972 and had a significant absolute coverage level at 1978. From these evidence, we can conclude the importance of a role that national agreement played in the wage determination process in Britain, therefore justifies the use of data on national bargaining in analysing the wage determination process between 1950-75.

Few remarks can be made regarding the recent tendencies. The recent micro survey, CBI data bank, reveals the depressing role played by the national agreements. Their data were collected from the annual survey on each settlement group, the actual survey being answered by the management representatives of these groups. It was confined to the manufacturing sector only, and the "settlement group" can refer to the even more disaggregated level than the establishment itself. On average, 1.7 settlement groups per establishment were observed. According to the survey, most of the agreements were made at the disaggregated level and national agreements, if any, were popular only among smaller sized establishments and manual workers. Also, only around 49% of the basic rates were determined at national or national together with supplementary bargaining. CBI Data also shows that the 90-93% of the contracts have the duration of exactly 12 months (Gregory, Lobban and Thomson (1985)), which obviously is not the case for our data set throughout our sample periods. Since the late 70's, the tendencies towards more disaggregated level of bargaining grew stronger than what we observed in the early 70's. Disaggregated bargaining enables the negotiating group to have more frequent negotiations. These days, the annual contract seems to be a norm.

3.5 Duration of contract

Duration, δ_{ns} of the s-th contract for the n-th group is computed as a difference in weeks between the s-th and (s-1)-th implementation dates, $(T_{ns}-T_{ns-1})$. In the data, they are listed together with the prevailed wage, w_{ns-1}, that was implemented at the (s-1)-th bargain (See figure 3.9).

Of all the implementation of wage changes recorded, some of them are designated only for a particular category of workers. Also, not all the bargaining group have the usual 4 categories of workforce: top-male, semi-male, bottom-male and female. In this sense, a problem arises in deciding which category's wage change should be appropriated for each implementation, It may be possible to take some weighted average of all the wage changes that took place at the same bargaining as a measure of w_{ns} corresponding



(Figure 3.9) Timing of settlement

to the s-th contract which lasted for δ_{ns}. Although a decrease in sample size is inevitable, a less arbitrary way is to create 4 different files according to the categories of workers, and this is the method adopted in our study. In this case, not the length between any successive bargains but the length between the consecutive bargains that determine wages for a certain type of workers is recorded as the contract length to be studied.

Of all the 147 negotiating groups, 84 groups included all 3 grades for male workers, namely, top, semi-skilled and bottom. 92 groups included both top and semi grade while 87 included both bottom and semi grade. Only 36 groups did not have female grade.

3.5.1 Indexed contracts: wages council groups

Before start analysing the data, we had to drop some observations in order to satisfy certain criteria. First of all, we exclude any groups that has missing implementation dates in the midst of its time dimension. Any missing observation at the start or at the end of the sample periods are just discarded. This should not affect our study which requires the stream of multiple bargaining history that is continuous. In this process, we have to drop 16 groups from our analysis.

Secondly, we exclude those groups who utilize the cost of living allowance clause (COLA) as a main means of wage determination. We are interested in the duration between the settlements that are properly negotiated via bargaining rather than determined automatically by some function of retail price index. When a group has COLA as a primary mode that triggers wage changes, it can be thought of as having one long term settlement with a protection measure against inflation. It is different from the staged settlement where workers know a priori when and by how much the wage will change in the future, since under COLA, neither size nor timing of such payments are known a priori. This kind of indexation provision in the contract is common in North America. With such contracts, the elasticity of indexation becomes one of the key choice variable in addition to the contract length. At the moment, we are applying our analysis to the open-ended contracts that was started with the negotiated bargaining. Any fixed length contract with the indexation clause should be distinguished from the others because of the differences in their nature of wage determination process. The groups belonging to the building, construction, and printing industries apply to this category, and altogether, there are 30 such groups out of 191 groups in our total sample (this 30 groups includes 3 groups which were deleted from the sample due to the missing observation).

There have been a few wage changes which were made by COLA even amongst those groups who don't utilize COLA as principal wage determining mode. They are usually due to temporarily imposed government policy in 1974-1975. The threshold arrangement was introduced in May 1974 and continued till the end of the year (or on to 75 for some groups). There, a special payments of 40 pence per week for every percentage rise in RPI above 7% was made as a compensation for a rise in inflation for all groups who joined the policy. Amount of payment was revised every month, as can be seen in figure 3.10 where the mode is around 4 weeks. For these groups, COLA is not the primary mode of wage change determination, but was due to the government policy which is completely exogenous to the bargaining. Hence, it is not necessary to exclude the whole observations belonging to such groups, but important to bear in mind that these few wage changes were determined entirely by different nature than by bargaining. Fortunately, these policy took effect during the last few years of our sample period, hence, we can exclude those COLA induced wage changes by discarding all the observations made after the second quarter of 1973 without distorting the rest of the sample.

Few cases (4 cases) remain where the indexation was adopted to determine the wage changes. But their involvement of COLA itself are likely to have been determined during the negotiation without any intervention from outsiders. Hence, it is different from those groups who utilize COLA as a main tool of wage determination, and we regard the 4 cases as ordinary wage determinations.

The durations which are affected by the indexation clause are either started off by the policy or ended due to the commencement of such policy. Excluding the groups with missing observations and groups with principally indexed wages, we divide the sample of contracts into COLA and non-COLA categories. Using 3944 observations on spells between wage changes that apply to any one type of workers, a mean of contract length is 47.79 weeks, which is just under one year, with a standard deviation of 28.9 weeks, indicating significant variations in the duration. Once we exclude the contracts affected by the indexation clause (mainly due to 1974's threshold policy), a mean contract length increases to 54.11 weeks with still a significant variation of 25.77 weeks. 12.6% of contracts are affected by COLA in a way stated above, exhibiting a mean of 9.73 weeks with a standard deviation of 13.5 weeks. The same analysis is given in table 3.2 below for the data on wage changes for the top-male, semi-male, bottom-male and female categories (for bottom male, see figure 3.10 and 3.11).

In the analysis which follows, we exclude all the observations after second quarter of 1973. This eliminates most of the contracts which were affected by COLA without

		non-COLA	COLA
top-male	nob	2499	415
	$E(\delta)$	54.39(25.1)	16.97(20.9)
semi-male	nob	1977	329
	$E(\delta)$	55.32(25.1)	15.91(23.0)
bottom-male	nob	2500	460
	$E(\delta)$	56.26(26.1)	17.0(21.4)
female	nob	2353	377
	$E(\delta)$	56.1(27.1)	16.60(21.7)

($E(\delta)$=mean duration in weeks, standard deviations in parenthesis.)

(Table 3.2) Number of wage implementations and
the mean analysis of contract length induced by COLA

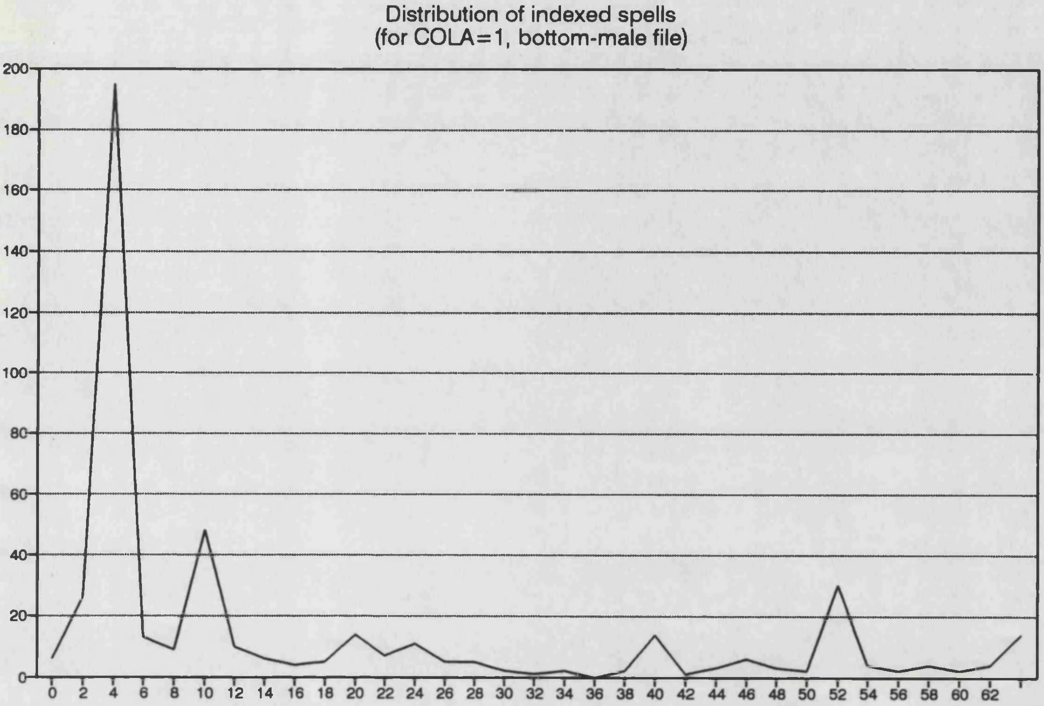


Figure 3.10

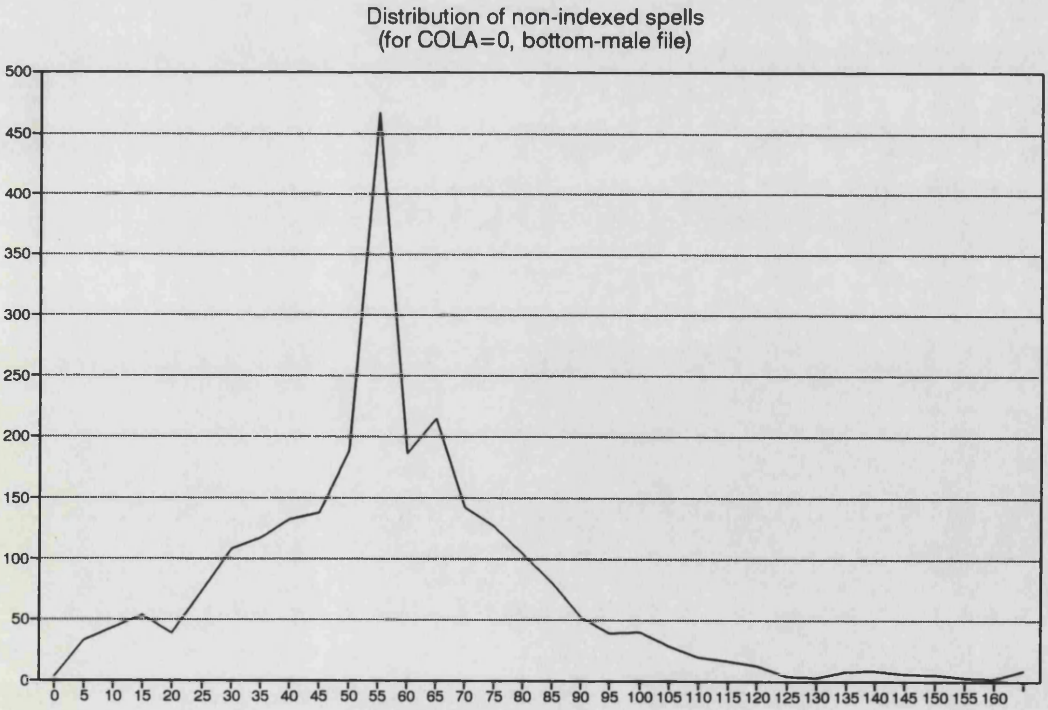


Figure 3.11

destructing the continuous time dimension of our panel.

Consequently, there are 108 groups altogether in a bottom-male file. Amongst those, 61 groups belong to manufacturing industries. For the top-male file, there are 106 groups altogether with 52 groups belonging to manufacturing. 83 groups in the semi-male file, amongst which there are 45 manufacturing. Lastly, female file contains data on 98 groups, where 61 belongs to manufacturing.

With respect to the bargaining system employed, see the table below. The bargaining groups that typically employs Whitley Council system are railway and hospital staffs, and those belonging to "others" category includes police, prison, government establishments, and atomic energy. Hence, if a group neither employs wages council nor private bargaining system, the group is identified as being in a public sector (i.e., they employ either (2), (3), (4) or (6) bargaining systems in table 3.3). Wages council is the statutory body that fix and monitor the rates of pay of many low paid occupational groups. They function in a similar way to the other negotiating bodies, with either side having the right to initiate a negotiation at any time. They often sets "Wages Regulation Orders" which is to be followed by the members of Wages Councils. These are usually in accordance with the social target aimed by the government, such as increase in the minimum level of wages to achieve equal pay for female and juvenile workers or decrease in the working hours. This distinctive, rather bureaucratic system, might have played influential role in the process of bargaining compared to the others with the independent systems.

Bargaining system	Bottom	Top	Semi	Female
(1) Wages Council	43	36	30	50
(2) Whitley Council	2	2	2	1
(3) Public/Nationalized	7	6	6	2
(4) Local Authority	3	6	5	0
(5) Private employer	51	53	38	43
(6) others	2	3	2	2
Public sector ((2),(3),(4),(6))	14	17	15	5
Total	108	106	83	98

(Table 3.3) Number of groups by the bargaining system employed

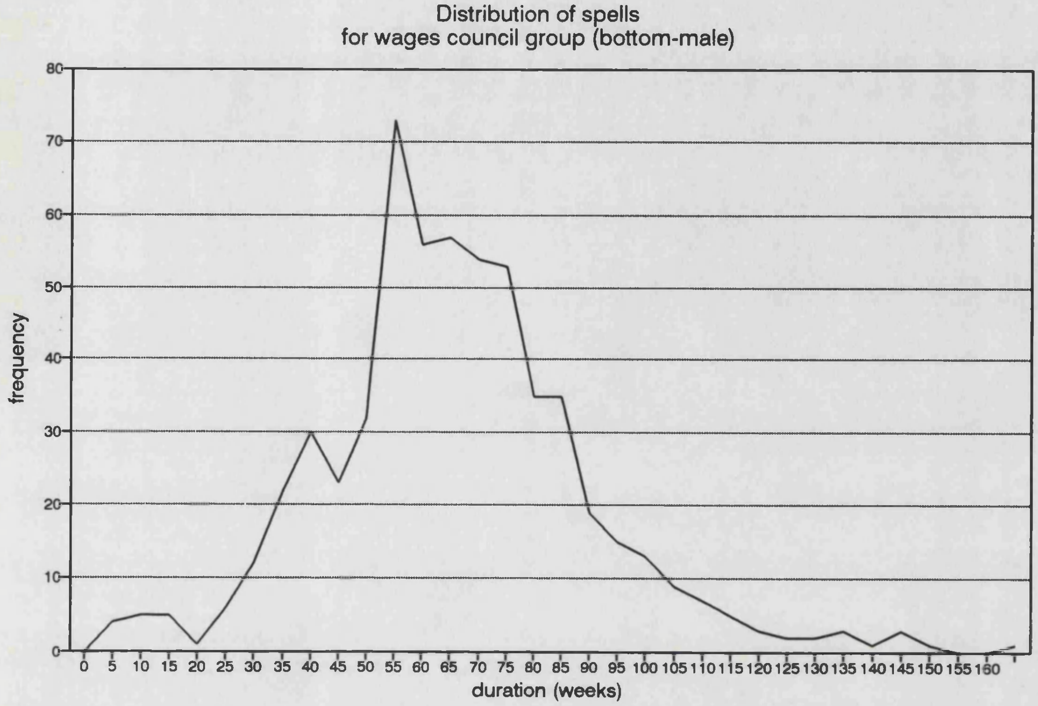


Figure 3.12

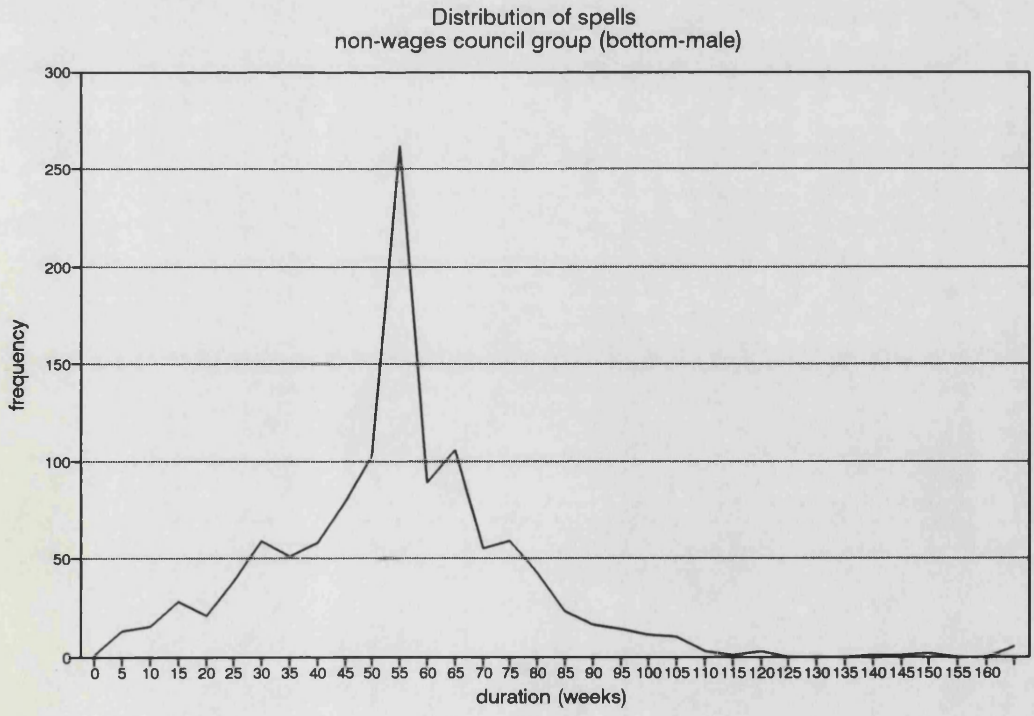


Figure 3.13

Nonetheless, as far as the frequency of wage implementation is concerned, we found them broadly similar, although the wages council groups tend to have a longer contract and larger variation (mean=64.78 and standard deviation=25.91) compared to the others (mean 52.35 and standard deviation 20.95 weeks). See figure 3.12 and 3.13.

The data also lists if any wage change was associated with the government's arbitration. In particular, table 3.4 lists the number of wage changes in a sample of the bottom-male file which was affected by 8 different forms of government interventions. The average length of spells terminated by the settlement affected by the arbitration and

	sample size	variable	mean	standard dev
National arbitration Tribunal/Industrial Disputes Tribunal	34	δ	59.08	10.36
		%w	0.048	0.017
Industrial arbitration board	24	δ	52.63	25.50
		%w	0.046	0.017
Court of Inquiry	4	δ	57.50	6.35
		%w	0.133	0.102
Ad hoc arbitration boards	10	δ	54.70	10.07
		%w	0.060	0.027
Public sector Arbitrator	3	δ	52.0	12.0
		%w	0.044	0.006
Committees of Inquiry and Investigation	1	δ	18.0	-
		%w	0.102	-
Prices and Income	6	δ	57.83	30.08
		%w	0.101	0.055
Others	1	δ	39.0	-
		%w	0.022	-
No Arbitration	1673	δ	56.58	23.73
		%w	0.067	0.046
Total	1756	δ	56.53	23.48
		%w	0.067	0.046

(Table 3.4) Number of settlements affected by the government's arbitration - bottom male the mean of corresponding wage changes are also listed in the table. Even though it would be of an interest to examine the influence of the arbitration on the timing or the degree of wage changes, unfortunately, the sample size of those affected by such intervention is too small to draw much inference.

In the light of recent finding that about 90 percent of the settlement takes place annually, (Gregory, Lobban and Thomson (1985)), we have examined the sample distributions of spells that occur before and after 1965, in figure 3.14 and 3.15. Despite such evidence, there still seems to exist substantial variations in the contract length during the period 1960-1970. Although the mean spell length falls somewhat from 57.90 to 54.65 weeks, there is no corresponding decrease in a dispersion about this mean. On the contrary, the standard deviation increases from 19.79 to 27.63 weeks.

3.5.2 Details of staged settlement

14.6% of the wage changes recorded in the data are staged a priori (excluding those brought about by the indexation clause). When a particular wage change is staged, such change is brought about as a consequence of the agreement that plans to yield wage changes in some stages. Although 2.4% of such settlements were not found as staged in the gazette, we treat all the settlements recorded as staged either in our original data or in the gazette to be the 'staged' settlements. 38.5% of staged settlements are known to have magnitude of future wage changes fixed at the time when the first stage was negotiated. Only 0.9% of those were altered when implemented (this figure increases to around 2.4% when including data for post 73 Q2 periods). Out of all the staged wage changes, 72.3% were known to have planned dates (up to a month) at the time of first stage, out of which, 15.02% failed to take place as planned. Moreover, 20.8% of these were in fact deferred due to wage freeze. Incidence of such a deferment took effect mostly from 1966 to 67, during the time of 6 months stand still, a statutory order imposed by the government.

When we look at the length of spell which is associated with the staged settlement, they have significantly shorter duration exhibiting a mean of 45.24 weeks compared to non-staged counterparts with a mean of 59.24 weeks (bottom-male file). Associated duration of a spell here implies either its start or its end is the staged bargaining.

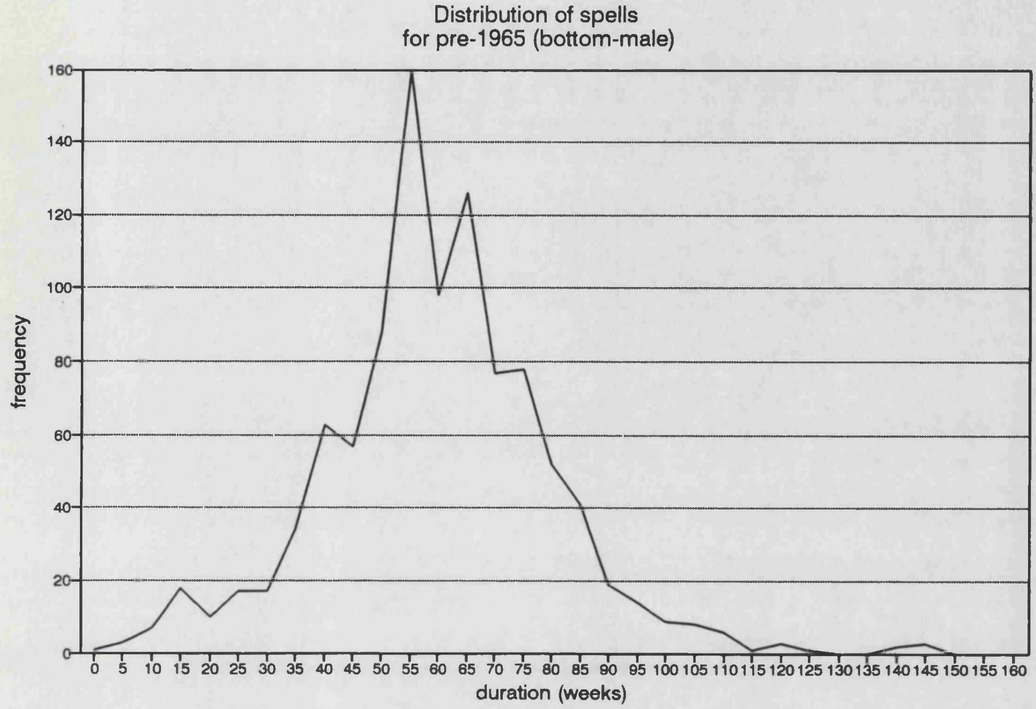


Figure 3.14

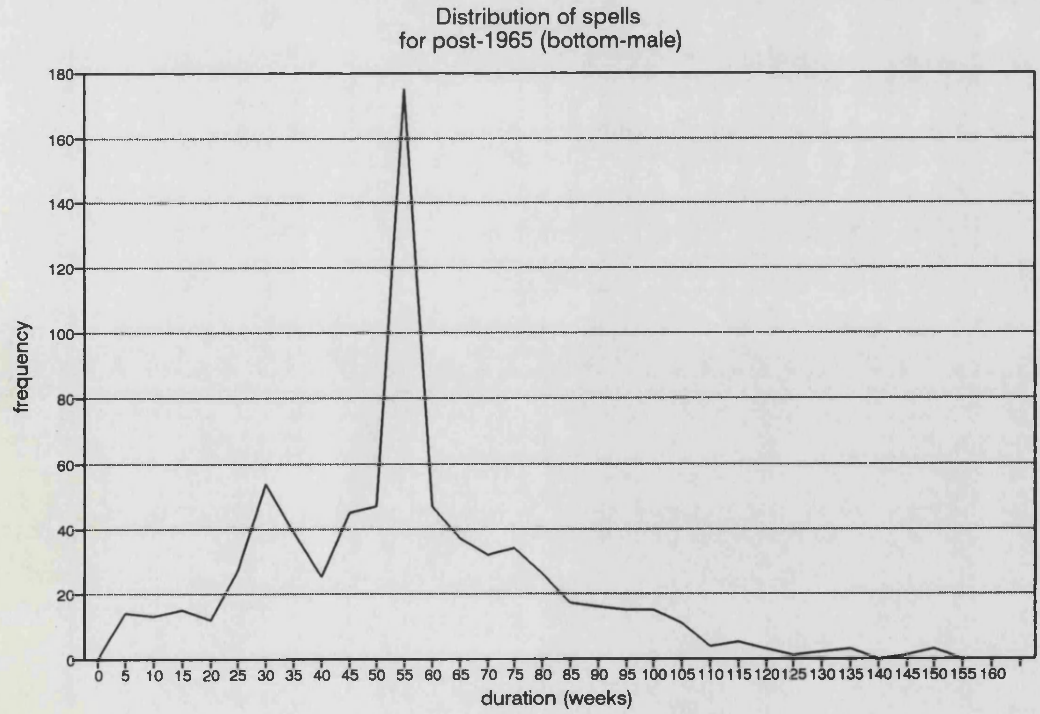


Figure 3.15

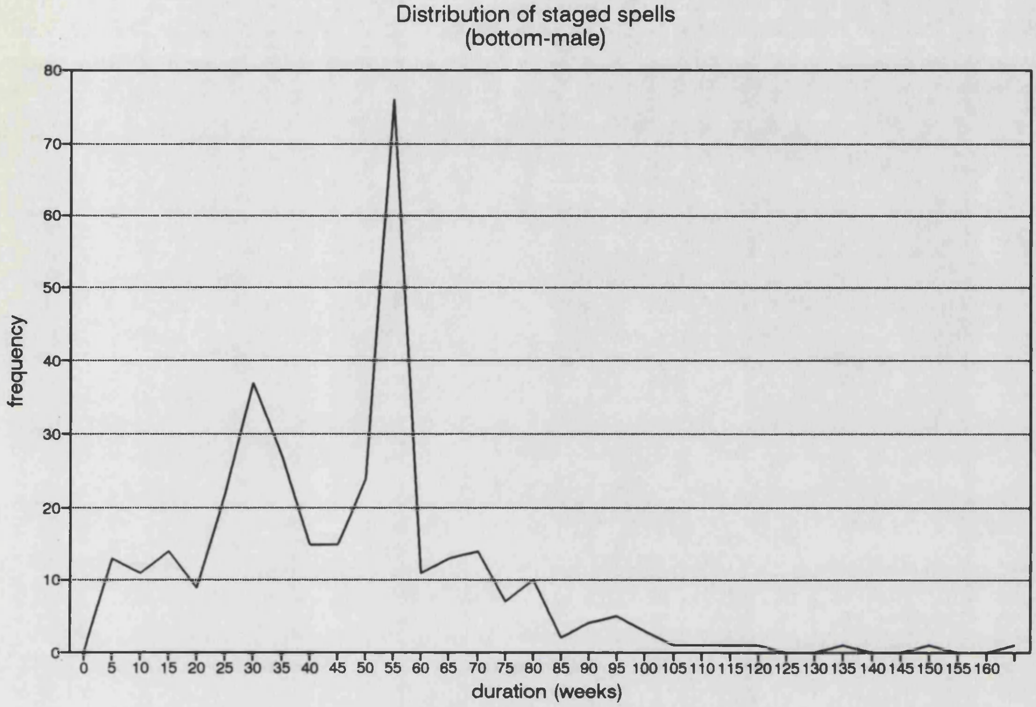


Figure 3.16

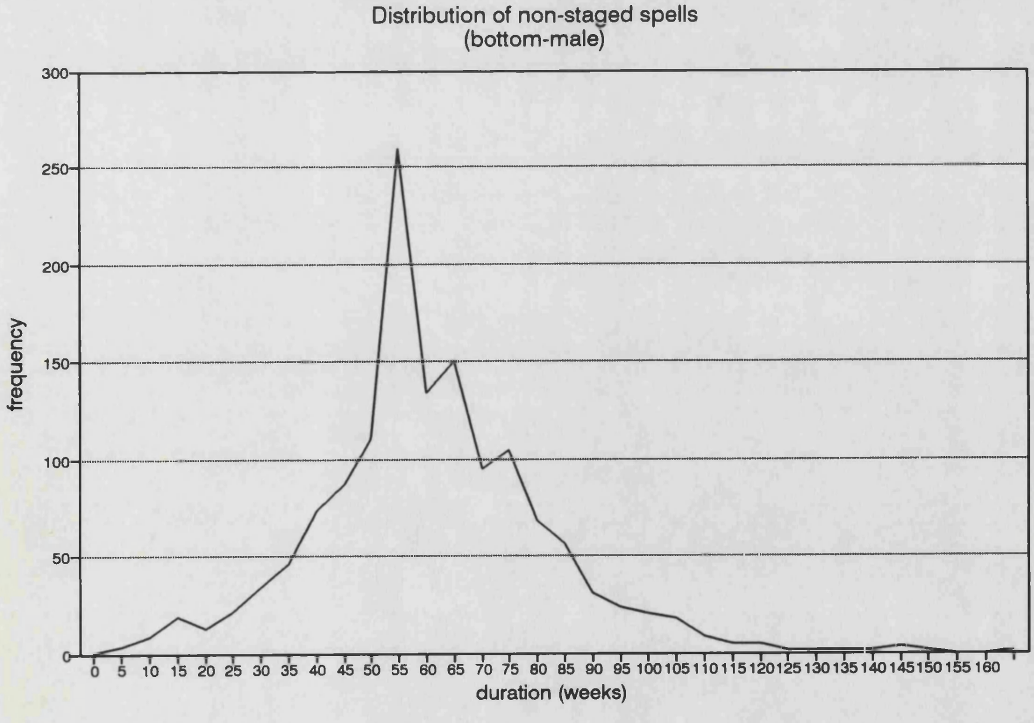


Figure 3.17

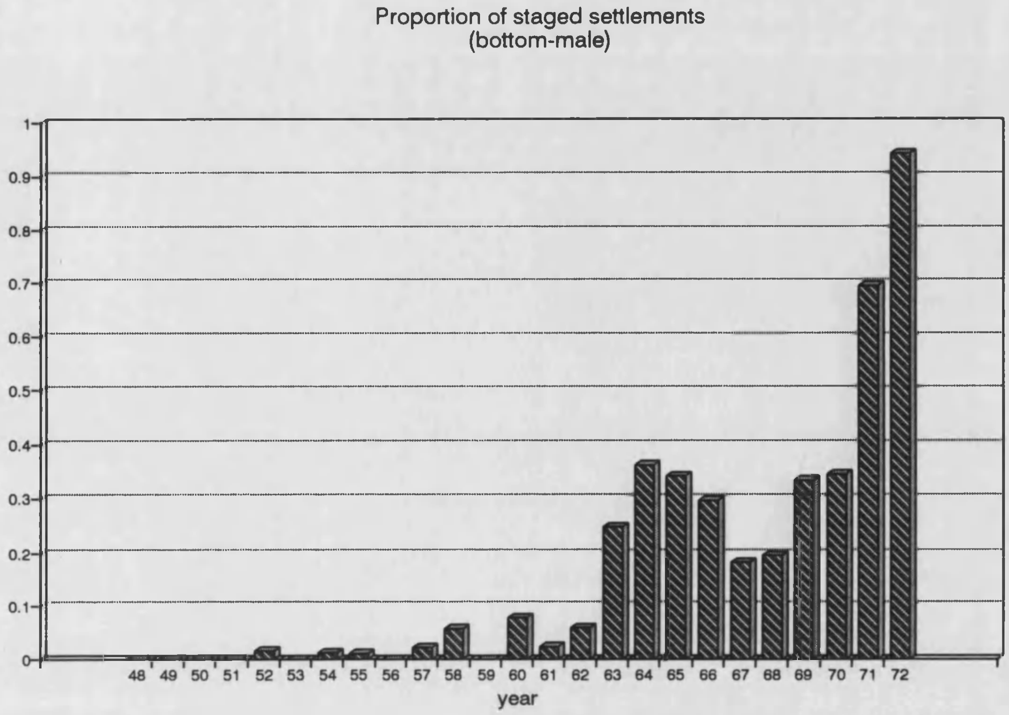


Figure 3.18

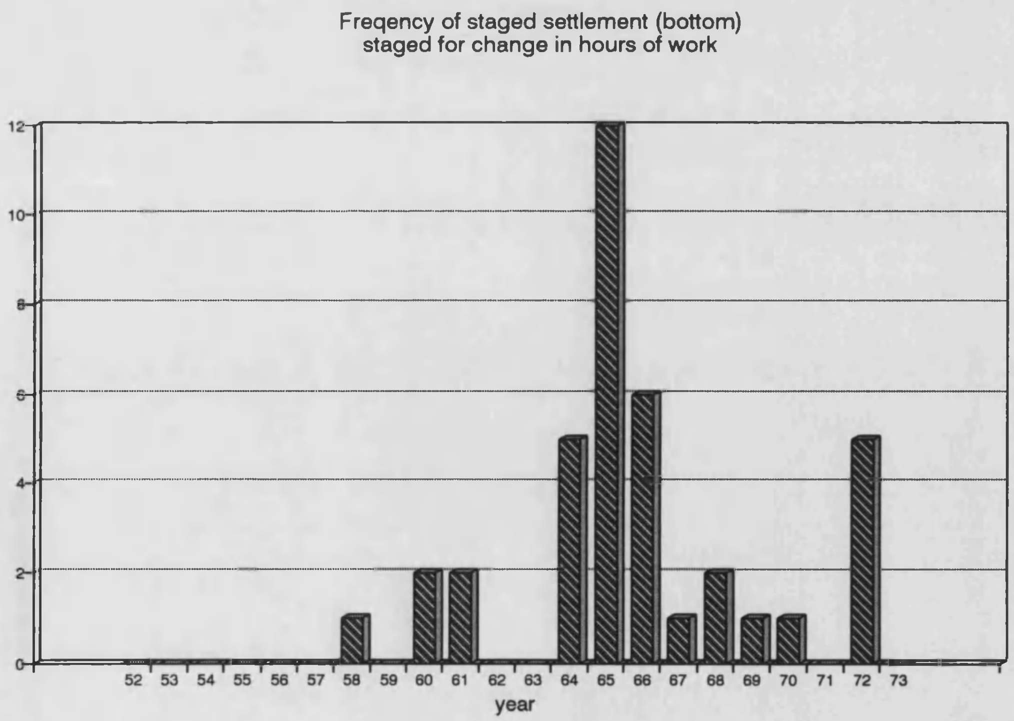


Figure 3.19

		non-staged	staged
Top	nob	18770	341
	E(δ)	56.48(22.75)	42.87(21.16)
	E(%w)	0.067(0.046)	0.07(0.059)
Semi-male	nob	1569	267
	E(δ)	58.83(25.23)	43.79(21.83)
	E(%W)	0.015(0.207)	0.060(0.09)
Bottom male	nob	2009	325
	E(δ)	60.00(24.6)	44.00(23.7)
	E(%w)	0.067(0.047)	0.066(0.053)
Female	nob	2343	387
	E(δ)	52.74(30.03)	38.22(24.57)
	E(%w)	0.082*0.08)	0.08(0.071)

(Table 3.5) Mean analysis of contract length and rate of wage changes for the staged and non-staged settlements

Nonetheless, there remains a substantial degree of dispersion around these means with standard deviation 25.80 and 22.05 weeks, for staged and non-staged, respectively (see figure 3.16 and 3.17).

Look at the figure 3.18 which plots proportion of the staged bargains out of entire settlements that occurred each year during the sample period. Clearly, peaks are found in 65, 66, and then again in 71 and 72. In 65 and in 66, there was major movement towards less working hours (mainly from 42 to 40 hours per week) without the loss of weekly pay. As can be seen from the figure 3.19 which shows the frequency of staged settlements that specifically reduces hours of work, such stages are most frequently seen in 65 and 66. There was another major movement towards fewer working hours seen in 60 and 61 (mainly from 44 to 42 hours a week), but this time, they didn't take form of stages. Another peak in the early 70's for the staged settlements are to make women and juvenile workers equal with respect to their pay. Specifically speaking, equal pay act was put into operation in 70 and 72 which aim to achieve equal pay by the end of 1975. Mean analysis of duration (δ) and rate of change in wages are given in table 3.5 for staged and non-staged implementations. It is clear from the table that the bargaining involving female wages are more often staged than the others, and also, the rate of change of wages brought about at those staged implementations are on average, higher. This is explained by the frequent staged settlements for women that took place in order to achieve the equal pay, since the level of female wages were so much lower than the male equivalents until the early 70's. The women's pay just about caught up with the male counterparts by 1975.

3.5.3 Mean analysis for each industry

Significant variations can be seen in the frequency of bargaining across different industries. Focusing on the bottom-male category, the manufacturing and construction industries have slightly shorter contract length on average (see figure 3.20 and 3.21) than they do otherwise (manufacturing mean=55.5 weeks, others mean=58.7 weeks), although both have similar standard errors at around 23 weeks. In particular, the Dock industry, fur industry and hat, cap millineries have very infrequent negotiations, lasting at least for 1 and a half years on average, as a result, there are as few as 10 negotiations during our sample period. The longest contract length of 285 weeks observed is in the fur industry. On the other hand, industries such as saw milling, food manufacturing, and post offices derive wage changes as often as every 6-7 months on average, reporting more than 40 negotiations over the sample period. For the top-male category, 13 bargains were observed on average during the entire sample period (excluding Q3,1973 onwards) with standard deviation of 7 bargains. The Aerated water, licensed residential and non-residential establishments, merchant navy and police have relatively infrequent negotiations with mean duration of above 76 weeks. Plastic, heat installment and forestry, on the other hand, have bargains on average every 40 weeks or less. Even within an individual agreement group, there is a significant variation in the contract length with average standard deviation of around 20 weeks.

What is interesting, however, is that those groups with infrequent negotiation does not necessarily come up with larger percentage change in wages at each settlement. Look at table 3.6 which shows the mean and standard deviation of the duration and the percentage wage changes realized during the sample period for the groups with particularly infrequent negotiations (Top-male wage changes). Compared to the

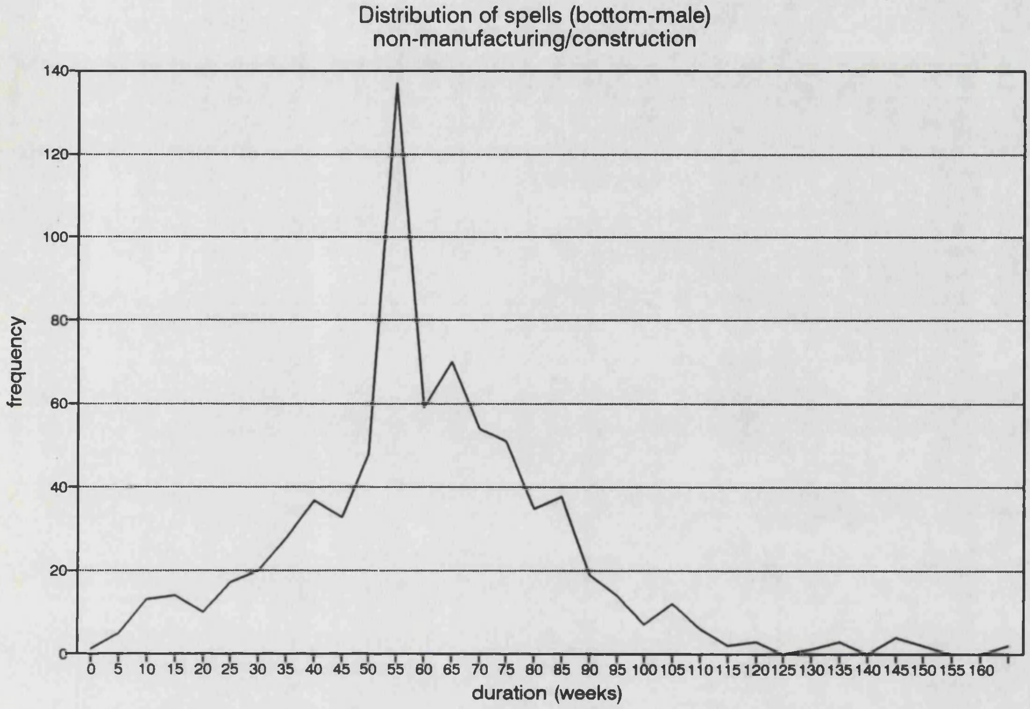


Figure 3.20

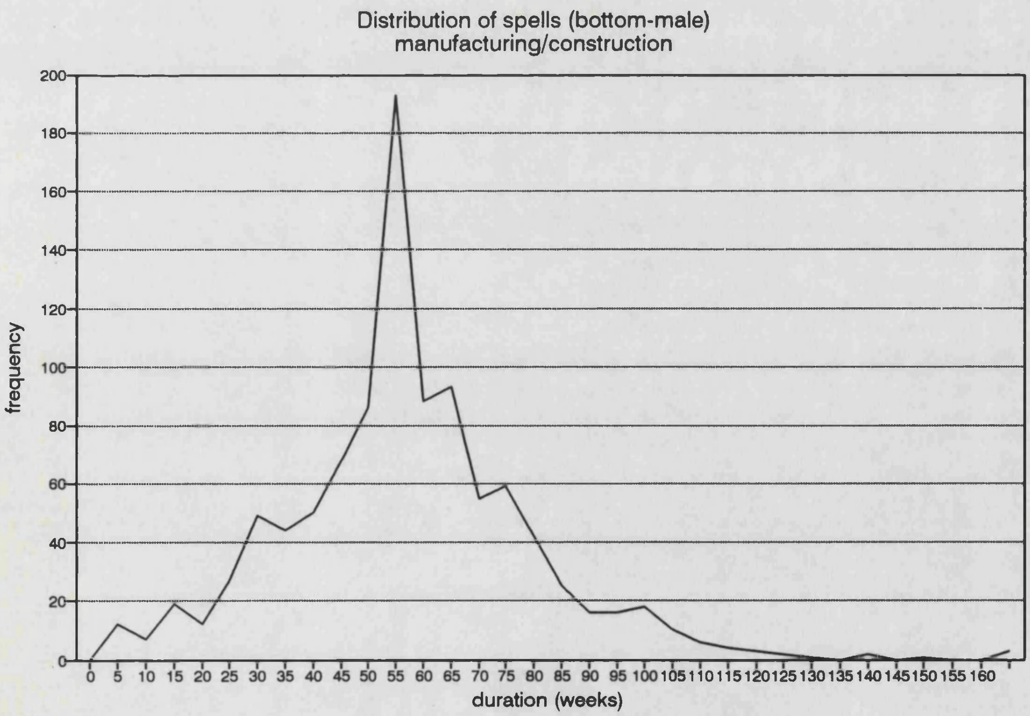


Figure 3.21

merchant navy and police groups, the licensed residential/non-residential groups exhibit much lower rate of wage changes despite of their having longer contract length on average. This is probably due to difference in their size of wage drift between private and public sectors. During the early 1960's, rates of pay in the public sectors were considered heavily controlled by the incomes policies, while the earnings in private sector rose rapidly. A sharp rise in the wage rate of public sector manual workers took place towards the late 60's in order to narrow the gap between two sectors.

	no. of settlements	$E(\delta)$	$E(\%w)$
Licencenonresident	12	92.3 (37.17)	0.09 (0.045)
Licenceresident	12	84.9 (22.3)	0.029 (0.067)
MerchantNavy	14	75.9 (33.4)	0.1 (0.08)
Police	15	73.0 (44.3)	0.11 (0.089)
HeatInstall	31	38.7 (18.82)	0.05 (0.039)
Forestry	28	40.7 (20.18)	0.06 (0.05)
whole group	2211	54.38 (23.03)	0.068 (0.049)

(Table 3.6) Mean analysis of contract length and rate of wage changes for some groups (Top male category)

3.6 Incomes Policy

Since the mid 1960's, the incomes policy emerged as an additional instrument of macroeconomic policy by the government in reducing the rate of change of nominal wages by intervening the mechanism of wage setting.

The policy takes variety of forms. It may be a wage freeze that demand a complete nominal wage standstill, or a twelve months policy that restricts no more than a single settlement per year, or a ceiling that specifies a certain target level in the rate of wage change. The coverage of these policies vary so that they may be aimed only towards the public sector or also to the private sector. Furthermore, the degree of enforcement

period		norm	comment
1948 Feb-51 June (2501) (2677)	wage freeze	0	not clear whether applied on the earnings or basic wage
1956 Mar-56 Dec (2921) (2964)	wage plateau		voluntary for the nationalized sector not to increase prices
1961 Jul-62 Mar (3202) (3237) fd_{n1}	wage freeze	0	enforced in the public sector, voluntary for the private sector
1962 Apr-63 Mar (3238) (3288) cd_{n1}		2-2.5%	in line with national output, enforced in the public sector only, voluntary for the private
1963 Apr-65 Apr (3289) (3392) cd_{n2}		3-3.5%	enforced in the public sector only
1965 Apr-66 Jul (3393) (3458) cd_{n3}		3-3.5%	enforced in the public sector only
1966 Jul-66 Dec (3459) (3484) fd_2	stand still	0	statutory in all sectors
1967 Jan-67 June (3485) (3510) fd_2(contd')	severe restraint	0	statutory in all sectors
1967 Jul-68 Mar (3511) (3544) cd_4	relaxed severe restraint	3-3.5%	voluntary restraint
1967 Jul-70 Dec (3511) (3692) d12	Twelve month rule		minimum of 12 months to separate the settlements

(Table 3.7) History of incomes policy

period		norm	comment
1968 Mar-69 Dec (3545) (3640) cd_5		3.5%	statutory limit to all sectors
1970 Jan-70 Jun (3641) (3666) cd_6		2.5-4.5%	statutory limit to all sectors
1971 Aug-72 Sep (3723) (3779)	N-1		1% less than last wage change
1972 Nov-73 Mar (3788) (3809) fd_3	wage freeze	0	statutory to all sectors
1973 Apr-73 Nov (3810) (3840)		£1+4% (≈6.7%)	statutory limit to all sectors
1973 Nov-74 Feb (3841) (3853)		8.5%	statutory limit to all sectors
1974 Feb-74 Jul (3854) (3874)		approx 13%	statutory limit same as before plus cost-of-living payment
1974 Jul-75 Jul (3875) (3926)			voluntary pay increase to cover the cost of living
1975 Jul-76 Jul (3927) (3978)		£6/week (≈10.4%)	zero for those earning more than £8500 per annum

(Table 3.7) (contd')

also varies from voluntary to statutory.

Table above lists the brief history of the incomes policy since 1948. (The figures in the parenthesis are the absolute weeks since 1900.) First major freeze policy was in effect between July 1961 and March 1963, but was only enforced to the government and wages council employees and was left voluntary for the private sectors. Second freeze that lasted for 5 months between July and December 1966 was a statutory policy to all sectors and any wage implementation dates planned to take place during this period had to be deferred. This freeze policy was in effect until June 1967, although the deferred wage changes during the first half of the policy were honored during this period. The third statutory standstill policy took effect between November 72 and March 73 for 4 months.

The average level of rate of wage changes that took place during each quarter and the average length of contract which ended at each quarter are computed and plotted against time in figure 3.22 and 3.23, respectively. It is clear from figure 3.22 that there are sharp drops during the second and the third freeze policies and moderately stable level of rate of change in wages during the ceiling policies. The level of wage changes are high during the fd1 freeze policy, although the graph plots the average over whole groups while the fd1 policy was only enforced towards public and wages council groups. Notable point is the sharp increase in the level of wage changes starting in Q2 1970, around the time the cd6 was called off. According to figure 3.23, there is a sharp decline in the contract length that terminates during the first few quarters of the freeze policies, fd1 and fd2. Also look at figure 3.24 which plots the number of negotiations occurred during each quarter. There are sharp decline in the number of wage changes during all the freeze policies. In principal, there should not be any bargaining occurring during the freeze policy, but if they do, we observe that they have short contract length during such period. This is true even for fd1, during which the average wage changes did not decrease (first quarter has seen a sharp increase). This, again should be explained by the extent of coverage of this policy. Otherwise, before the policy comes into full effect, a number of groups may well have rushed to have

	policy	number of bargaining
fd_{n1}	Freeze (61 Jul-62 Mar)	21
cd_{n1}	Ceiling(62 Ap -63 Mar)	50
cd_{n2}	Ceiling(63 Ap -65 Ap)	88
cd_{n3}	Ceiling(65 Ap -66 Jul)	85
fd_2	Freeze (66 Jul-67 Jun)	49
cd_4	Ceiling(67 Jul-68 Mar)	71
cd_5	Ceiling(68 Mar-69 Dec)	165
cd_6	Ceiling(70 Jan-70 Jun)	54
fd_3	Freeze (72 Nov-73 Mar)	3
d12	Twelve (67 Jul-70 Dec)	358

(Table 3.8) Number of bargaining occurred during the policy

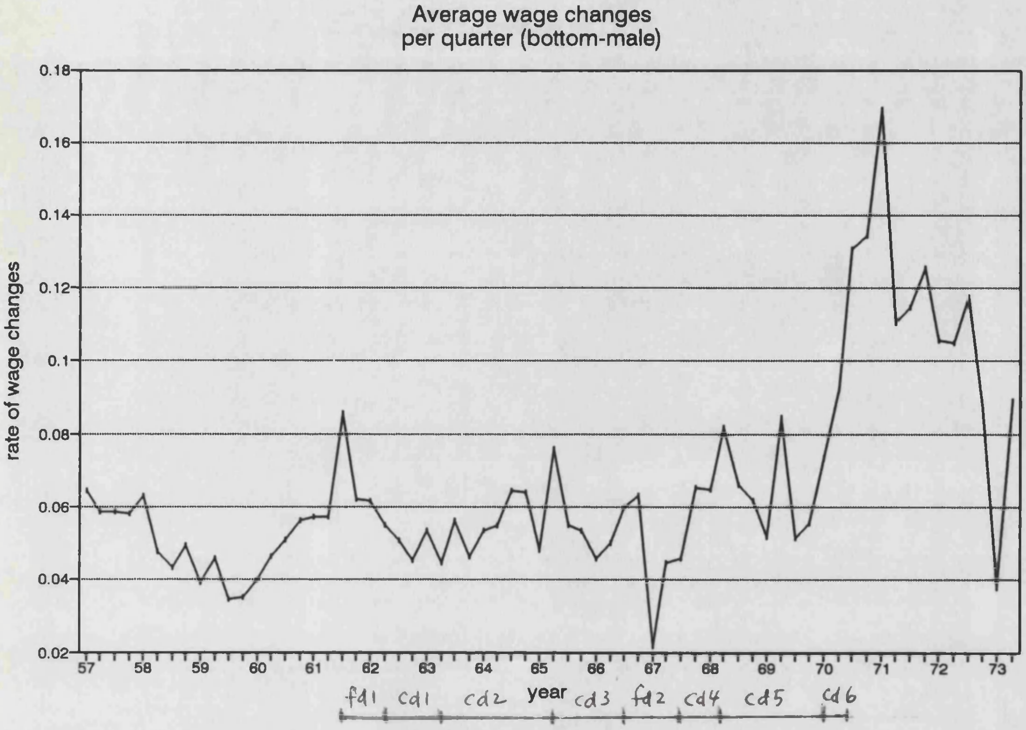


Figure 3.22

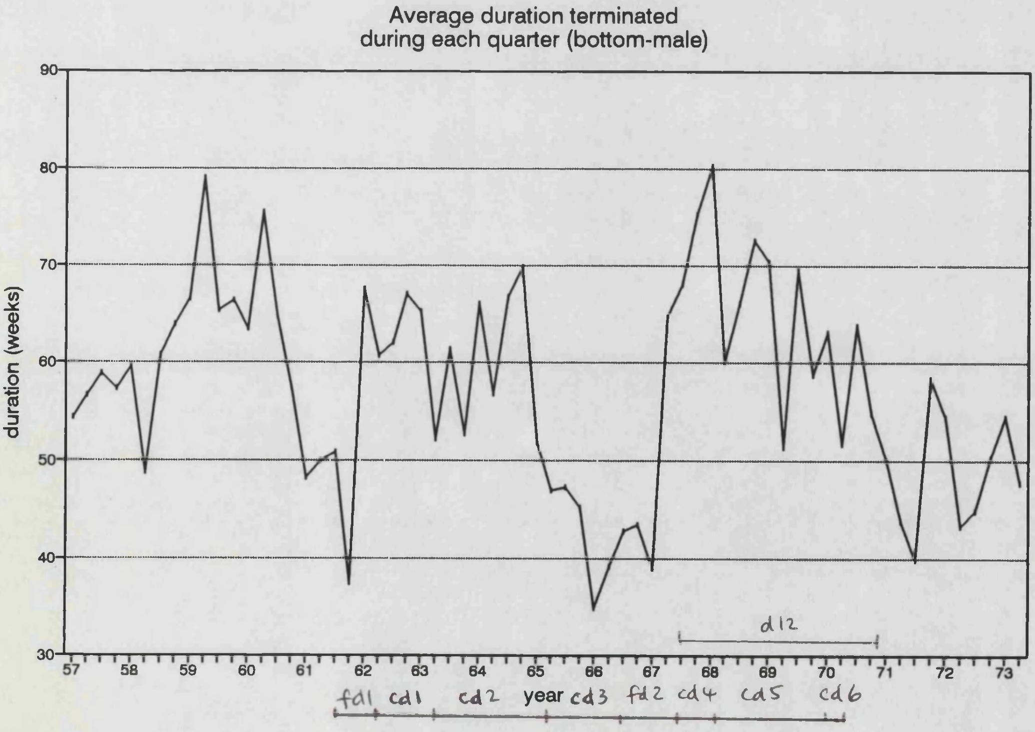


Figure 3.23

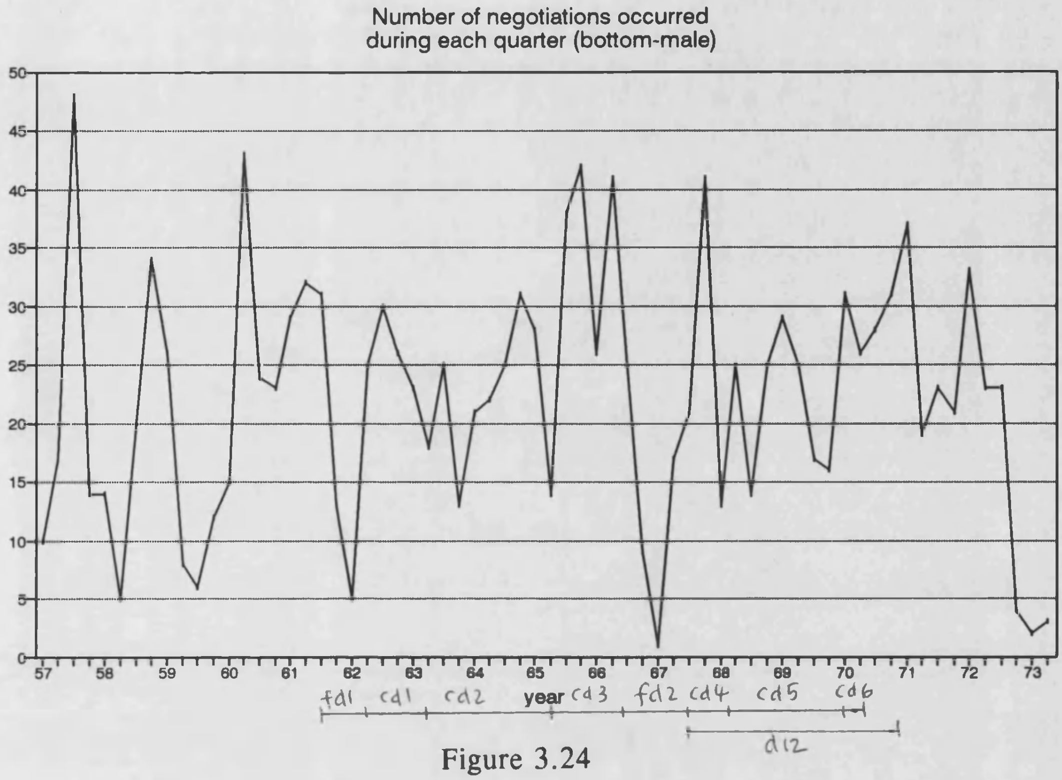


Figure 3.24

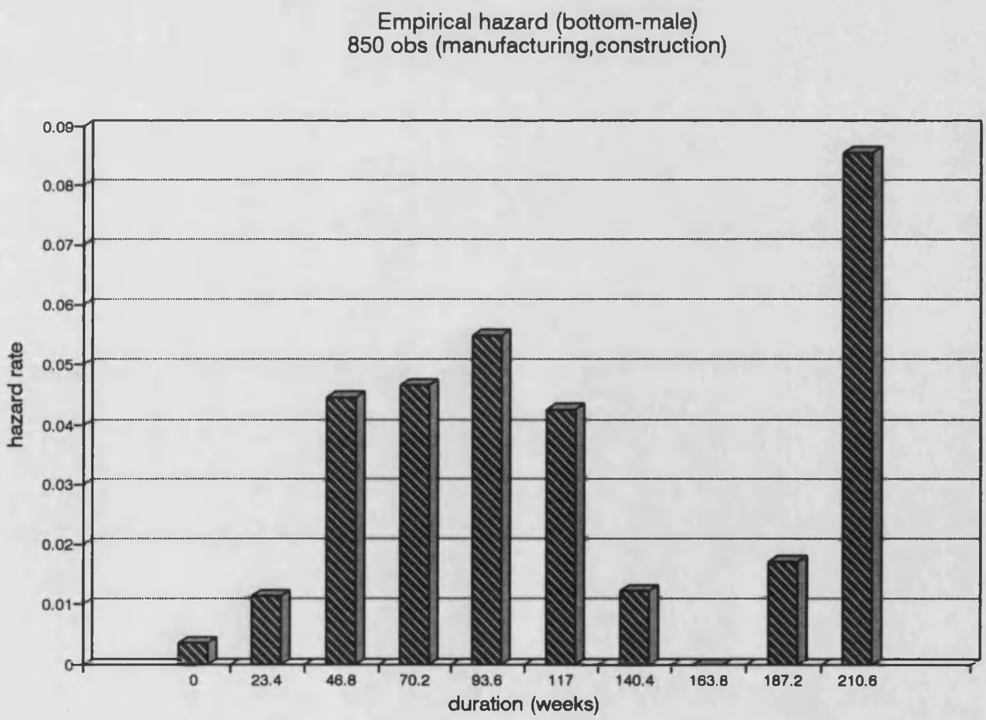


Figure 3.25

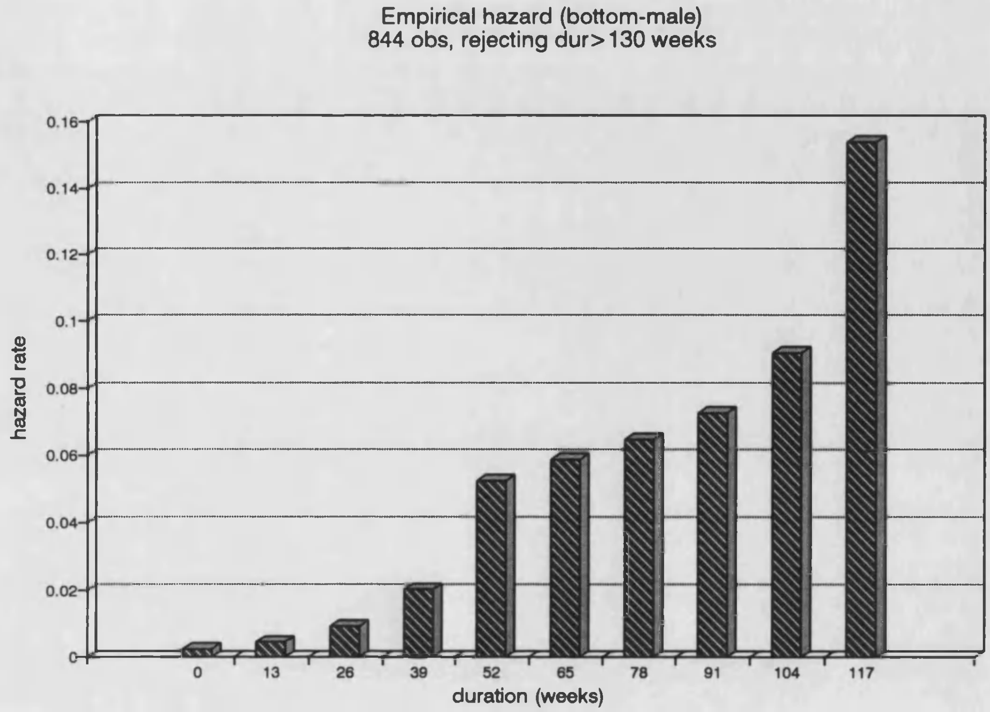


Figure 3.26

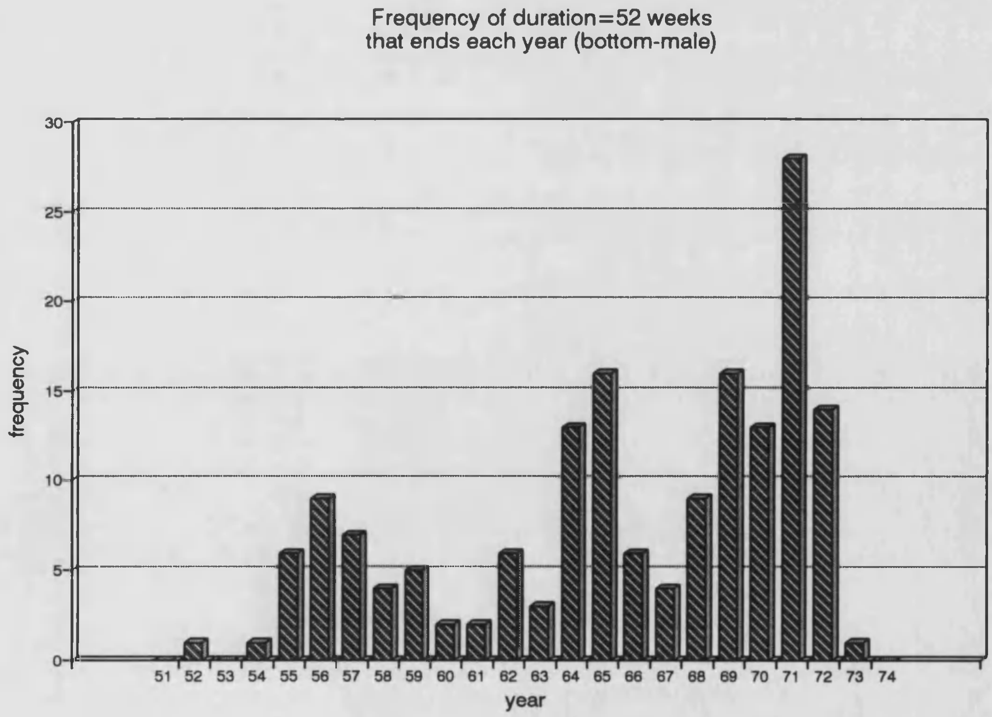


Figure 3.27

wage changes, resulting in small number of settlement with short durations. Number of bargainings occurred during each incomes policy period is listed in table 3.8 for the bottom-male category of our sample which is truncated at 73, Q2. fd_{n1}, cd_{n1}, cd_{n2}, and cd_{n3} are enforced only towards non-private groups, therefore, the table lists number of bargaining struck by non-private groups during the corresponding period.

3.7 The Empirical Hazard

Figure 3.25 shows the empirical hazard function, estimated as a step function over the 23.4-week intervals (sub-divide maximum duration observation length = 234; into 10 equal intervals) using 850 observations on durations for the manufacturing and construction groups, bottom-male category. It was calculated according to the life table method (see section 4.2), which computes the hazard rate in each interval as a proportion of number of spells terminating in that interval out of a size of the "risk set": set of duration observations that have not terminated at the interval's lower limit. This figure suggests a rising hazard rate, at least up to around 22 months, with a substantial jump at 52 weeks. Due to the nature of the life table method, the empirical hazard is high towards the maximum length of duration observations. Of course, such calculations take no account of the behaviour of the explanatory variables, and do not necessarily reflect accurately the behaviour of the baseline hazard. It is also sensitive to the width of the intervals assumed, and to the size of observations. When we exclude observations which recorded longer than 130 weeks, the shape of the empirical hazard becomes monotonically increasing (fig 3.26), although a jump at around 1 year still exists.

A notable point emerging from all this information is the importance of the peak around 52 weeks, corresponding to a regular pay round. These 52-week spells are distributed quite evenly across the bargaining groups apart from the concentration observed in spells terminating in 1971 (17% of annual contract length are observed in 1971, see figure 3.25). In the whole sample, however, there is still considerable dispersion around the annual peak. The major aim of our study is to explain this dispersion in terms of the economic influences, in particular, of the income policies.

Chapter 4: Economic framework

4.1 Introduction

Data on individual pay settlements are a valuable source of evidence on the nature of bargaining behaviour and the effects of government intervention in the labour market. By analyzing the results of individual wage negotiations, it is possible to avoid the severe aggregation problems inherent in a study of wage determination at the macro level[1] and take into account of the complex institutional features of the labour market such as varying contract length, differed increments, varying bargaining calendar and the degree of centralization in bargaining. However, a full dynamic analysis of micro bargaining data raises formidable methodological issues since it must explain both the magnitude of wage changes and their timing against the backdrop of continuously evolving economic environment.

The timing of wage changes is particularly difficult to analyse, and has received much less attention in applied econometrics literature than have the magnitude of those changes. This is so despite the fact that by analyzing both timing and magnitude of wage changes, it is possible to discriminate the effectiveness of a government control policy on the two, hence can test a widely-held view that past incomes policies have been effective only as a means of delaying, rather than moderating pay increase. Notable exception to this neglect of the timing issue include Pencavel (1982) and Lee (1987) using British data, and Christofides and Wilton (1983), Christofides (1985) and Cecchetti (1985) who use North American data. However, analytical methods used by those authors do not reflect the full dynamic nature of the wage determination process which is very important in the case of open ended contracts.

The purpose of chapter 4 and 5 is to develop and apply a method of modeling the occurrence of wage settlements and the magnitude of wage changes associated with those settlements. In our data we observe, for each discrete event, not only the time since its last occurrence (i.e., duration, δ_{ns}), but also the realized wage change at the end of each duration (i.e., w_{ns}). The task is therefore to build a model that simultaneously determines such pair of variables.

It is important, before embarking on the statistical model, to recognize the stylized fact on the nature and the degree of organization of the bargaining system so that the assumption about the way timing and wage is determined can be made appropriately for our data. Unlike in North America, agreement groups in

Great Britain during our sample period (1950-75) generally negotiated open ended agreements which fixed the nominal wage but left unspecified the date of the next negotiation[2]. Those negotiation were made mostly at national level which was the most prevalent at least in the manual sector at a time, hence, government control policy such as incomes policy was easy to implement. There were rarely indexed contracts, and even in few cases where a contract length was determined a priori (as a staged settlement), they were often violated. Hence, under this system of pay bargaining, there was a probability that a contract ends at any point in time. In other words, duration of a contract could be influenced by events occurring *after* a negotiation took place; such as an introduction of incomes policy or a sudden rise in the unexpected inflation. The mechanism, then, of a process towards a new negotiation can be plausibly considered as being triggered by the accumulation of such events in the economic environment. The accumulation will consequently put enough pressures towards unions so that a new negotiation will be called for when the potential gain exceeds, for the first time, the cost of having a new negotiation. Therefore, spell length will depend on the entire time path of all the relevant determinants observed from the previous bargain onwards.

This form of relation cannot be appropriately estimated by conventional regression methods but by a duration model. A probability of having a new negotiation at any point in time, once a contract starts, is analogous to the notion of the hazard rate, which can depend on the entire time path of the evolving environment up to a given point in time during the spell. Precisely speaking, the hazard rate at elapsed duration, δ, is defined as a probability of the spell ending at t conditional on all the currently observable values of explanatory variables and on the fact that the spell has not previously terminated since t periods ago. This hazard rate outlines the process up to a point of renegotiation. Given such point, the variables that have exerted influence on the hazard together with other forward-looking factors will then determine the magnitude of wage changes via bargaining. In this way, it is possible to consider separately the determinants of the hazard rate for renegotiation and the agreed wage changes conditional upon a new settlement taking place. This composes a joint density function for the pair of observations, $\{w_{ns}, \delta_{ns}\}$, therefore, completing a system of simultaneous relation between the timing of a bargain and a concomitant wage changes.

In this chapter, economic model is presented and specification of suitable variables is discussed on the bases of which an econometric model is built and applied to a British bargaining data in the following chapter 5.

4.2 Conceptual framework

Suppose that union and firm come to an agreement on the level of nominal wage and the timing of such a wage change via bargaining. They may decide some of the wage changes to be implemented in multiple stages whose dates are determined a priori. Nonetheless, such planned dates may be violated and eventually, a new negotiation will take place. Hence, under this system of bargaining, there is a possibility of a new negotiation occurring at any point in time. Once the nominal wage is fixed, it stays invariant or changes only as much as planned until new negotiation. Meanwhile, economic environment continues to vary, and the rigid nominal wage becomes more and more unfavorable to the workers. Workers will decide to step forward for a new negotiation only when the gains from the negotiation outweighs the cost of doing so. And here, we assume such criteria to be the difference between an optimal real wage, real take home pay and transaction cost.

Transaction costs involve resources necessary in conducting negotiations during the process of preparation as well as administration of having negotiations, implementing a wage change and a potential cost of industrial actions. In a latter case, there will be a further cost of industrial relation deterioration, cost involved in settling internal equity problem when wage change occurs. Their cost should largely differ for different bargaining machinery involved in negotiations (whether it takes place with private employers, with nationalized industry or with wages council). Different firms or industries may also have different likelihood of a strike occurrence, ways of implementing wage changes. Size of union members or number of unions involved in a bargaining unit can also be considered to play a role.

The optimal real wage is the outcome of a bargaining given the economic environment and the bargaining power of each group were they to bargain continuously. Consider the case of n-th bargaining group, s-th observed contract. Let us denote t as an elapsed duration since calendar time T_{ns-1}, hence, a duration t corresponds to $T_{ns-1}+t$ in calendar time. Denote the time path of such optimal wage for the n-th bargaining group as $w^*_n(t)$, retention ratio as rr(t), retail price index as p(t) and a cost of negotiation as c(t) where t corresponds to any point in time between T_{ns-1} and T_{ns}. Then, renegotiation takes place when,

$$w^*_n(t) - c_n(t) - \frac{w_n(T_{ns-1})rr(t)}{p(t)} \geq 0 \qquad (4.2.1)$$

for the first time since the last negotiation, T_{ns-1}. To be more general, (4.2.1) can be written as:

$$w^*_n(t) - c_n(t) - \frac{w_n(T_{ns-1})rr(t)}{p(t)} - e_n(t) \geq 0 \qquad (4.2.2)$$

where $e_n(t)$ is some random non-observable cost element idiosyncratic to the particular bargaining group, n. As long as above inequality is false for any t from the start of a contract, such a contract continues to survive. The hazard formulation takes into account precisely of this relationship continuously from the start of the contract up to the point of actual renegotiation.

$$h_{ns}(t) = \text{Prob}(\ w^*_n(t)-c_n(t)- \frac{w_n(T_{ns-1})rr(t)}{p(t)} = e_n(t) \text{ for the first time}$$
$$\text{since the start of (s-1)th contract} \mid \text{contract has already lasted at least for t period})$$
$$= h(\frac{rr}{p}(t),\ c_n(t),\ w^*_n(t)) \qquad (4.2.3)$$

where (rr/p)(t) and $w^*_n(t)$ are the entire time path of each variable since T_{ns-1} up to t where t can be any point in time between $[T_{ns-1}, T_{ns}]$. We consider the likelihood of a departure from the current contract defined between $t \in [0, \delta_{ns}]$, hence the corresponding duration density for this observation is a function of the entire time path of explanatory variables up to time T_{ns}. Note that the hazard at T_{ns} with elapsed duration δ_{ns} includes time path of explanatory variables up to T_{ns}, however, not the information that this contract actually terminates at T_{ns}. The hazard is greater (i.e., the likely occurrence of renegotiation is higher), larger the gain compared to the cost. Therefore, the hazard should be decreasing in (rr/p)(t) , $c_n(t)$ and increasing in w*(t).

On the other hand, the wage equation is the realization of w(t) at t where the inequality (4.2.2), for the first time since t=0, is satisfied. We observe such threshold to be reached for the n-th group's s-th contract at time T_{ns}. Then a potential wage level which triggers renegotiation for the n-th group's s-th contract is:

$$w_{ns} = (w^*_{ns} - c_{ns})\ (\frac{rr}{p})_{ns}$$
$$= w\ ((\frac{rr}{p})_{ns}\ c_{ns}\ w^*_{ns}) \qquad (4.2.4)$$

where $(rr/p)_{ns}$, c_{ns} and w^*_{ns} are the entire time path of those variables up to and including T_{ns}. Although this is the value which triggered renegotiation, once the contracts terminates, the level of w_{ns} that merely satisfies inequality (4.2.2) is not necessarily the wage to be claimed by the workers at the negotiation. This is because workers, knowing that the contract will last for some time, not only take care of their past uncompensated inflation, but also try to make as good ex-ante provisions as possible for the future uncertainty. Such provisions depend on workers expected length of next contract and general uncertainty over future, in particular, over price and income tax. In this sense, the agreed wage becomes:

$$w_{ns} = w\left(\left(\frac{rr}{p}\right)_{ns}, c_{ns}, w^*_{ns}, \Omega_{ns} \right) \qquad (4.2.5)$$

where, again, subscript ns refers to their entire time-path up to and including T_{ns}. Ω_{ns} is the forward-looking expectational element made at T_{ns} over the future uncertainty. It is decreasing in first and second arguments and increasing in w^*_{ns} and future uncertainty.

4.3 Bargaining

Let us now turn to the nature of optimal real wage variable, w*(t), which is the outcome of bargaining between firms and unions were they to bargain continuously overtime. Since bargaining takes into account both of demand and supply constraints, optimal wage will depend not only on the wage setting behaviour of both parties, but also on factors such as firms' profitability or their pricing policies. There are several theories that depicts union-firm behaviour in determining wage and employment (chapter 2). Essentially, firm and union each tries to optimize their profit and utility function subject to constraints. In the monopoly union model, union chooses the level of wage given the labour demand curve (i.e., which is a locus of firm's profit maximizing pair, (w,n)). On the other hand, the efficient bargain model has both parties bargain over employment and wages. Seniority or insider-outsider model provides a rationale for the union to be indifferent over employment above a certain level, hence a bargaining takes place only over wages. A large number of empirical studies have been carried out to test the validity of these models, however, these tests do not provide conclusive results. First of all, their tests rest on the specification of either a form of bargaining (i.e., the efficiency of the outcome) or the agents' objective function. Hence, a rejection of the hypothesis is often skeptically dependent on the correctness of the maintained hypothesis. In particular, these studies often test the presence of alternative wages in an employment function based on the assumption that it affects marginal rate of substitution of the union. But the presence of the alternative wage does not necessarily mean an acceptance of the efficient contracts. Labour demand model (LDM) in conjunction with efficiency wage consideration can explain the role of alternative wage in an employment function even if a firm and union bargain over wages only. On the other hand, insignificant alternative wage may not always lead to the LDM. Insider-outsider or seniority consideration gives rationale for a flat indifference curve (FIC) which yields observationally equivalent employment function for the LDM and efficient contract model. This model has an advantage since its fundamental assumption of layoffs by seniority is commonly observed in practice, and is also able to explain the recent phenomena of persistent inflation and concomitant high unemployment rate. The test of the FIC is usually based on the claims that the level of wages determined during the bargaining should not be affected by any indicators of outside opportunities such as unemployment or unemployment benefit, since workers do not care about the welfare of the unemployed members. This restriction is often violated in the empirical studies. However, violation of such claims does not necessarily lead to the rejection of the FIC if a comparison effect is found to exists in the union utility function, which is often the case.

As one can see, empirical evidence is far from conclusive with respect to which model best suits the reality. Nonetheless, we adopt a simple version of the FIC model in our process of generating the determinants for the reduced form optimal wage equation. This choice of model is not crucial to our analysis, however. Derived determinants for the optimal wage is very similar and also easily generalized to accord with other bargaining setup.

Consider a simple bargaining where workers and employers can bargain over the level of wages and employment. Suppose a lay off takes place by some known order, such as that of inverse seniority. Then, a firm and union bargain only over wages since union is indifferent over the level of employment (see chap 2.1.5) even though workers *can* bargain over employment if they want to. A firm sets the level of employment unilaterally for a given wage by maximizing its profit so that the outcome lies on a labour demand curve. That is, if a firm maximizes its profit:

$$\max_{w} \quad [\ R(n,p^p) - w\ n - k\ ]\ /\ p \qquad (4.3.1)$$

where R is a revenue function, k is an exogenously given level of capital, p is retail price index, p^p is wholesale price. Then the solution lies on the labour demand curve, n*(w/p), hence, a profit function is:

$$\pi = R\ (n^*(w/p),p^p) - w\ n^*(w/p) - k \qquad (4.3.2)$$

For a firm facing a monopolistic competition, demand function for its product is affected by its own product price as well as other prices and real demand shock (e). In that case, labour demand becomes:

$$n=n^*(\ w/p,\ p^p,\ p,\ e) \qquad (4.3.3)$$

Firm's fall back profit, $\bar{\pi}$, corresponds to its level during a strike, which would be -k, in this case.

On the other hand, union, who is only interested in demanding higher pay is simply assumed, without a loss of generality, to have utility level which equals real wage.

$$u = \frac{w}{p} \qquad (4.3.4)$$

Their fall-back utility is $\bar{w}/p$ where $\bar{w}$ denotes alternative wage or wage attainable elsewhere. The bargaining solution is then assumed to be derived by maximizing a Nash maximand:

$$(\pi - \bar{\pi})^{1-\sigma}(\ \frac{w}{p} - \frac{\bar{w}}{p}\ )^{\sigma} \qquad (4.3.5)$$

where σ is the relative bargaining power of the union. The first order condition of this maximization problem yields:

$$w = (\frac{\sigma}{\sigma-1})\ \frac{\pi-\bar{\pi}}{n} + \bar{w} \qquad (4.3.6)$$

since π_w=-n by duality. Therefore, in this very simple model with risk neutral unions, wage is a weighted average of alternative wage (fall back level of wages) and profit per employee[3]. Profit term is a function of the "internal variables" that affect firm's financial performance which in turn determines the size of rent to be shared between two parties. They include technological progress, input/output prices, demand shock and its product market competitiveness. In addition, if a firm is constrained in their borrowing , the liquidity and cash flow consideration may play a role since they can restrict firm's activities. Outside wage, on the other hand, is a measure of "external effect" reflecting outside opportunities that can be represented by unemployment benefit, average earnings elsewhere, or global and local unemployment rate. The weights, σ, reflects bargaining power, which can be influenced by virtually any variables that affect bargaining environment. They may include indicators of bargaining machinery, external market situation, number of members or union density within an industry or government wage policies. Some of these factors are also considered to affect the derived wage level through alternative wages and cost of negotiation.

4.4 Specification of variables

4.4.1 The duration component

In view of equation (4.2.3), we categorize the influence on the hazard as: (1)change in real take home pay, (2)outsider influences, (3)capacity-to-pay of the firm, and (4)negotiation cost. In particular, we categorize their effect on the hazard into those initially existing at the start of a spell (they may contain group's entire history of past failures and wage changes, as well as an indicator of secular trend- a calendar time - if it is considered to affect the level of hazard) and another which continues exerting effect during the spell through their values at each point in time. With respect to the within spell varying covariates, it is important to make sure such explanatory variables are deterministic or if not, that they carry no information about the parameter of interests. So that for example, for our particular interest lies in the effect of incomes policy on the timing of negotiation, we have to assume that the timing of incomes policy is not endogenous to the other environmental variables that are also considered to affect the hazard, such as inflation. Otherwise, we will have difficulties in interpreting the coefficients of incomes policies since the incidence of such policies themselves may well be affected by the inflation.

First criteria is attributed to the real purchasing power of the workers who wish to keep up their standard of living. Declining real take home pay makes a new negotiation more likely as contract proceeds. In our model, we have decomposed this factor into two. Ex-post uncompensated part known at T_{ns-1} and another part which have accumulated since the start of the contract. The former is represented by the percentage change in the real take home pay received at the last consecutive bargains (i.e., between T_{ns-2} and T_{ns-1}), denoted as DRW1, and the latter is its changes since the start of the current contract (i.e., between T_{ns-1} and t where $t\in[T_{ns-1},T_{ns})$), denoted as DRR(t).

DRW1 serves a role of the "price-catch up" variable in Christofides and Wilton (1980) paper, which is an indicator of uncompensated inflation carried over from the previous contract. This variable essentially states whether the level of real take home pay has declined or increased between the last consecutive bargains. If workers aim is to keep up their standard of living, failure to do so at the beginning of the current contract, so that the uncompensated inflation is carried over into the following contract, should induce workers to have another pay rise quickly irrespective of what happens to the economic environment after T_{ns-1}. On the other hand, a large positive DRW1 suggests overcompensated workers who should be content with the existing wage level may economic environment stays the same . In general, the latter case is more plausible since workers, knowing that the contract will last for sometime would take into account of the ex-ante provision for the inflation uncertainty (see specification for the wage equation). In general, a high value of DRW1 may suggest two things: workers expectation of high inflation over the next contracts or high cost of negotiation. Their perception of high cost should make them claim for high nominal wage at T_{ns-1} since they expect not to have another negotiation for a long time. Hence, this would lower the hazard. Effect of rising expected inflation depends on its realized values, which will be picked up by the following, DRR(t).

Given the initial condition, DRW1, it is then the accumulated changes in real take home pay that concerns workers, hence, its entire time path from the start of the spell becomes relevant. This is captured by the latter variable, DRR(t). This variable is measured as a difference in the net real pay between each month τ^{i}_{ns} $(i=1,2,..k_{ns})$ and the starting month of the contract (τ^{0}_{ns}), reflecting the buildup nature of the decline in real purchasing power. This variable measures how better or worse off the workers have become in terms of their real take home pay at a given nominal wage since the start of the contract. When this value declines, we would expect unions more likely to step forward to the new negotiation therefore 'increases' the hazard. However, as far as the average real pay is concerned during our sample period, it is found consistently increasing with much higher rate than did inflation. This fact is confirmed by looking at DRW1 whose mean is 0.01655 and standard deviation, 0.0386. Its positive average indicates inflation rate that is lower that of net nominal wage. This may suggests that workers, during our sample period, were almost always compensated above the inflation rate, therefore did not care very much about its movement. It makes sense then, that only when the decline in real take home pay becomes large enough to threaten their standard of living, does it start giving a push to the occurrence of new negotiation. Against the backdrop of such rising real take home pay over the period, we assume this threat to operate through expectations, so that we expect the unexpected rather than expected decline in the real take home pay to affect more strongly the probability of negotiation.

Original variable expressing a change in real take home pay is decomposed into expected and unexpected component. Hence:

$$w_{ns-1}[\frac{rr}{p}(\tau^i_{ns})-\frac{rr}{p}(\tau^0_{ns})] = w_{ns-1}[\frac{rr}{p}(\tau^i_{ns})-\frac{rr^e}{p}(\tau^i_{ns})] + w_{ns-1}[\frac{rr^e}{p}(\tau^i_{ns})-\frac{rr}{p}(\tau^0_{ns})]$$

$$= \text{unexpected} + \text{expected} \qquad (4.4.1)$$

We assume that the rr/p series can be explained by distributed lag of its previous values and assume also that the knowledge of such relationship is used to form future expectation of rr/p. We consider 12 months lag structure for the monthly rr/p series for our sample. Denoting rr/p(τ^i) as P_t, rr/p(τ^{i-1}) as P_{t-1}, and so on, the estimated equation is:

$$\begin{aligned} P_t = \; & \underset{(0.06)}{1.95\,P_{t-1}} - \underset{(0.13)}{0.97\,P_{t-2}} - \underset{(0.14)}{0.70\,P_{t-3}} + \underset{(0.14)}{1.40\,P_{t-4}} - \underset{(0.16)}{0.70\,P_{t-5}} \\ & - \underset{(0.16)}{0.34\,P_{t-6}} + \underset{(0.16)}{0.70\,P_{t-7}} - \underset{(0.16)}{0.38\,P_{t-8}} - \underset{(0.14)}{0.14\,P_{t-9}} + \underset{(0.13)}{0.34\,P_{t-10}} \\ & - \underset{(0.12)}{0.21\,P_{t-11}} + \underset{(0.06)}{0.04\,P_{t-12}} + \underset{(0.4*10^{-5})}{2.1*10^{-6}} \end{aligned}$$

R^2 (adjusted) = 0.99993

standard error is in parenthesis. Result of this estimate is then used to generate the sequence of i future monthly rr/p.

First, consider that the i-th future expectation is formed at the start of the spell, τ^0, without further updating of information. For a typical n,s-th spell observation:

$$\frac{rr^e}{p}(\tau^i_{ns}) = E\left[\frac{rr}{p}(\tau^i_{ns}) \mid \Omega(\tau^0_{ns})\right] \qquad (4.4.2)$$

where $\Omega(\tau^0_{ns})$ is the information set available at τ^0_{ns} and E denotes expectation operator that uses relevant information set to derive i-th month forecast. Specifically, this is:

$$\frac{rr^e}{p}(\tau^i_{ns}) = \begin{cases} \frac{rr}{p}(\tau^0_{ns}) & \text{for } i=0 \\ \sum_{j=1}^{i-1} b_j \frac{rr^e}{p}(\tau^{i-1-j}_{ns}) + \sum_{j=i}^{12} b_j \frac{rr^e}{p}(\tau^{i-1-j}_{ns}) + b_0 & \text{for } 0<i\leq 13 \\ \sum_{j=1}^{12} b_j \frac{rr^e}{p}(\tau^{i-j}_{ns}) + b_0 & \text{for } i>13 \end{cases}$$

Under this assumption, workers form their expectation over future uncertainty based on the information at the start of the spell.

Meanwhile, our second assumption that assumes workers to update their expectation with one month lag throughout throughout the contract spell gives:

$$\frac{rr^e}{p}(\tau^i_{ns}) = E\left[\frac{rr}{p}(\tau^i_{ns}) \mid \Omega(\tau^{i-1}_{ns},\tau^0_{ns})\right] \qquad (4.4.3)$$

Ω is the information set available at the month τ^{i-1}_{ns} or τ^0_{ns}, whichever comes later. Hence at the start of a contract, the workers know the level of rr/p, and they derive the monthly update of rr/p levels thereafter. Specifically:

$$\frac{rr^e}{p}(\tau^i_{ns}) = \begin{cases} \frac{rr}{p}(\tau^0_{ns}) & \text{for } i=0 \\ \sum_{j=1}^{12} b_j \frac{rr^e}{p}(\tau^{i-j}_{ns}) + b_0 & \text{for } i>0 \end{cases}$$

In practice, we use the rate of change of unexpected and expected terms in equation (4.4.1) and may also include log of real wage level at the start of a spell to measure any existence of secular trend effect on the hazard. That is, if there is any tendency for the average duration to shorten due to higher real take home pay.

Figure (4.1) and (4.3) plots the actual and expected real take home pay according to the expectational form (4.4.2) and (4.4.3) for typical observations taken from a sample and comparing it with the actual series. As can be seen from the figure, expectation with 1 month lag provides as good prediction as the actual series while the forecasts based on the first assumption is rather like a monotonic linear function of time. Third dotted line represents expectations made with 3 months rather than 1 month lags. This assumption is the halfway house between the other two extreme cases set forth. Interestingly, expectation with 3 months lag traces rather closely with that based on (4.4.2), even though update of information still takes place with additional two months lag. Figure (4.2) and (4.4) plot unexpected component corresponding to the expectational

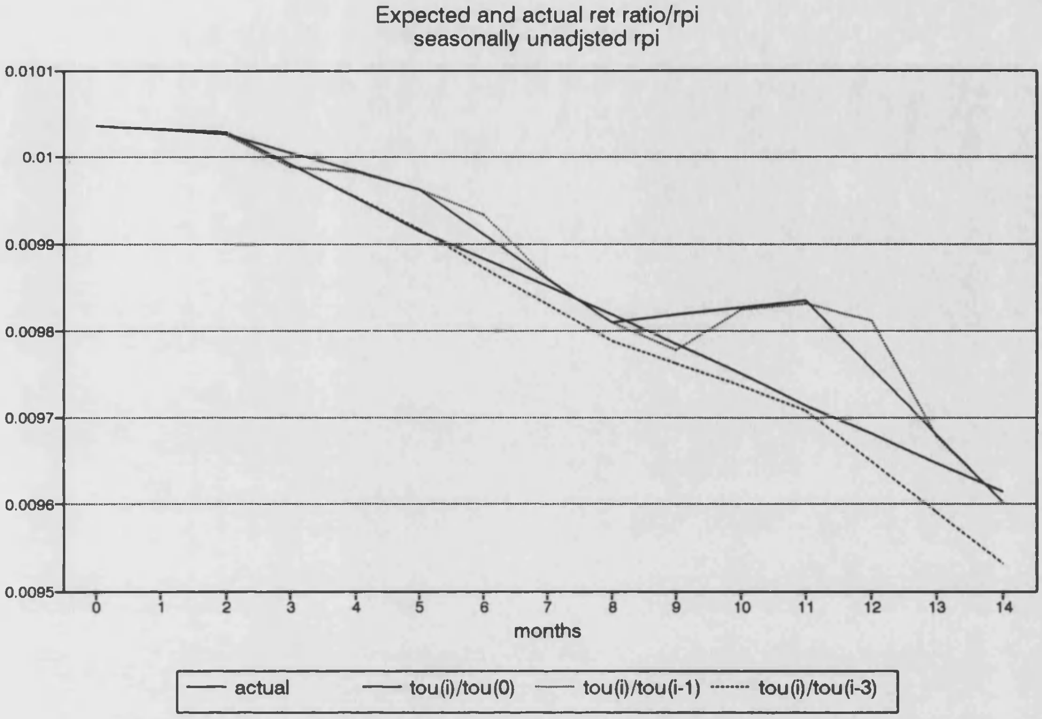


Figure 4.1

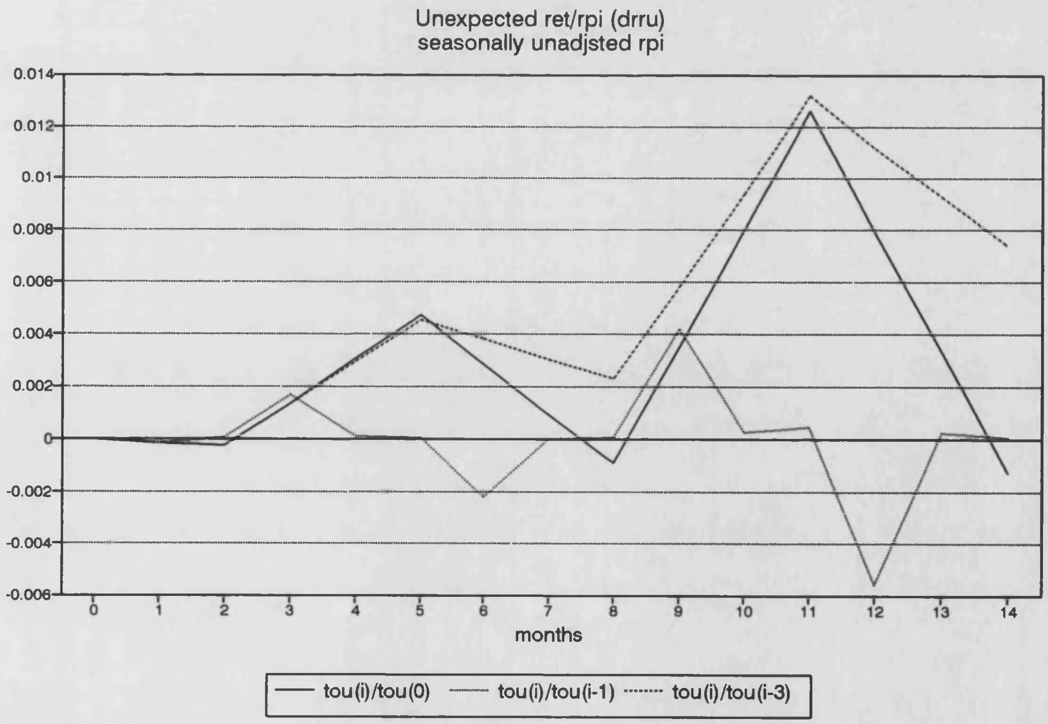


Figure 4.2

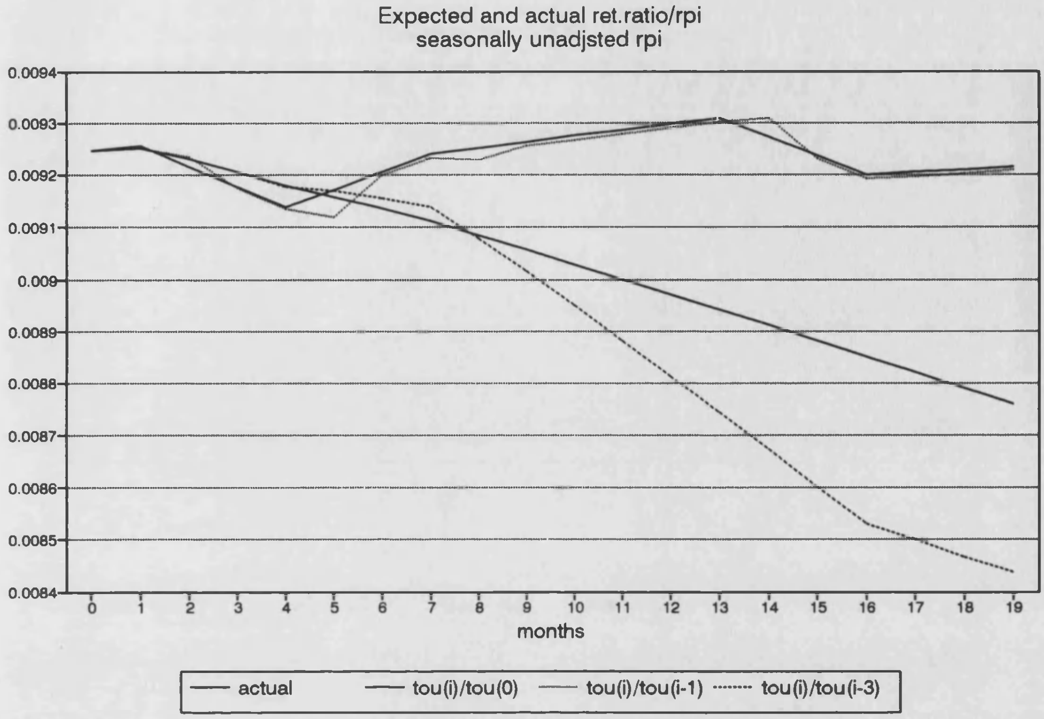


Figure 4.3

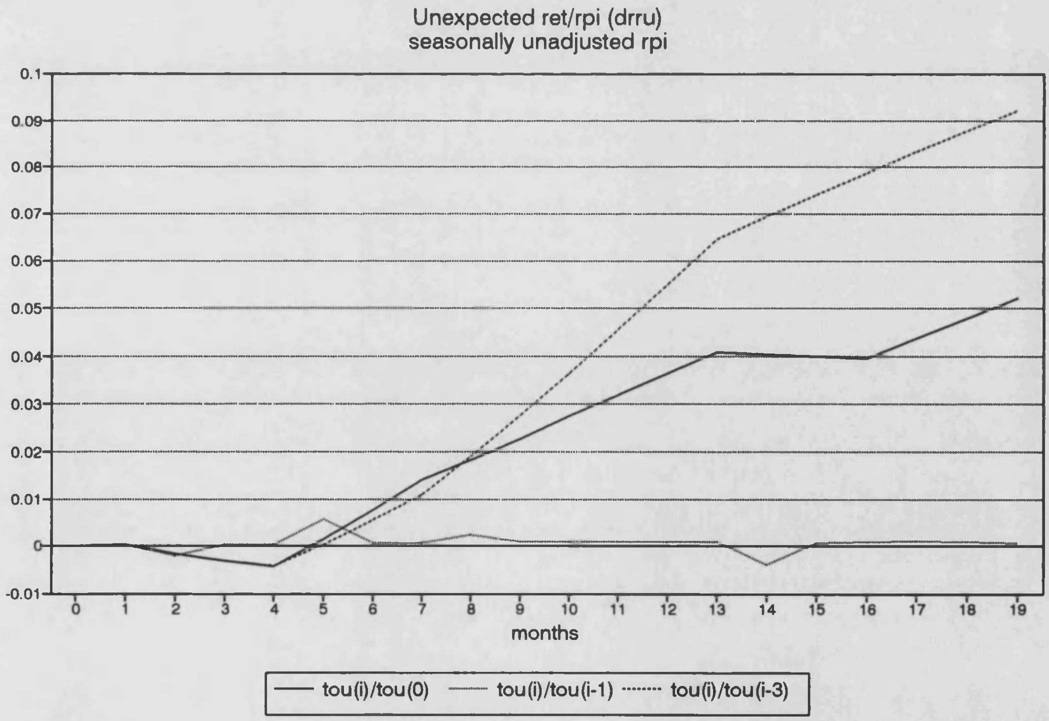


Figure 4.4

assumptions depicted in figure (4.1) and (4.3), respectively. These figures suggest that unexpected component largely differs according to the expectational forms assumed. The second assumption with as little as 1 month lag provides very small magnitude of unexpected component, which may seem too unrealistic, however, considering that the monthly inflation index is available to the public, it may not be so implausible after all. Typically, these forecasts tend to diverge away from the actual level as the spell lengthens even though this is not always the case. There also is a tendency for such divergence to be negative implying pessimistic attitude of workers who tend to expect for worse. Uniqueness of our model enables to measure the response of the hazard to such continuously varying factors throughout the contracts.

It is possible to a certain extent to insulate workers from this aspect of pressure by adopting a cost of living allowances which links nominal wages to the rate of inflation (Gray (1978)). However, this introduces a new variable, namely, an elasticity of indexation as a determining factor in the contract. In our current context, we excluded any bargaining groups which employed COLA system (Cost of Living Allowances) as a main device of wage determining process and our sample period do not include early 1974 when the majority of bargaining groups adopted threshold payment scheme introduced by the government. Nonetheless, Gray concludes that even with the optimal degree of indexation, contract length is affected inversely by uncertainty and positively by transaction cost. Empirically, Christofides (1985), Christofides and Wilton (1983) found that uncertainty induces noncontingent wage changes and such unexpected rise in inflation was found to play a significant role in a determination of contract length both for an indexed and non-indexed contracts.

Secondly, we consider external factors that affect $\bar{w}$. According to our formulation, they affect the hazard via the effect on the optimal wage, w^*. In view of this, we have considered two factors, namely, unemployment and relative wages. Global, as well as industry-specific measures of unemployment are considered and are both expected to exert negative effect on the hazard since their signaling for the excess demand for labour will lead to stronger bargaining power of the union, hence, push up w^*. On the other hand, relative wage measured as the ratio of current level of average earnings in the whole economy to the previously agreed basic weekly wage for each t, also serves as a surrogate to the excess demand in the labour market, hence, affects bargaining power. However, this is not the only way that relative wages can affect w^*. Comparative wage effect, fairness among workers is usually important in determining wages (Beckerman and Jenkinson (1989)). It reflects how better or worse off the workers belonging to this agreement group are compared to everybody else in the economy. The effort is not so much to narrow a gap with the others but rather to avoid the gap to widen. Hence, when this ratio becomes very high, even if the real take home pay has not changed, there is an incentive for the union to go ahead with a new bargaining since they now realize that the wages outside are getting higher faster than theirs.

Let us now consider the "internal" variables (Oswald and Blanchflower (1988)) which affects firm's financial performances. In a view that the wage determination via bargaining is sharing of a rent which is the surplus above the mere production cost, the size of this rent itself should play an important role. And this "ability to pay" serves as third criteria. Variables such as productivity, profit or sales per employee can be used as a direct measure of the size of rent. Competitiveness of the product market, changes in the number of employment as representing a growth or decline of a firm can also be of influence.

Rent sharing view also casts importance on the variables that affect bargaining power of the participants. Not only the size of the cake to be shared but also the relative power of two parties together determine the final size of each share. The existence of various forms of bargaining machinery or trade union, higher union membership or density should increase workers bargaining power in addition to the labour market variables discussed above (i.e., unemployment, outside wages). Note that even without a trade union, groups of workers with skills can have similar effects on their bargaining power due to their replacement cost both in terms of money and time. (Lindbeck and Snower (1986,1988)). Then, higher the proportion of skilled worker, higher the bargaining power of the workers. Also, given a certain size of rent, capital intensive firm maybe more willing to accept higher wages.

With respect to the actual data availability, unfortunately, it was not possible to find many disaggregated variables that reflects above measures. We have trading profit measure per industry, and the number of negotiators for each bargaining group. Number of workers covered at each agreement is known for each negotiating group at 1950, 55, 60, 65, 70 and 72. These figures are interpolated to give appropriate values at the time of each negotiations, and its value relative to total working population at a time is used as a measure of "membership". We also incorporated global productivity measure to take into account of the technological growth. When unions are informative about the rise in profit or productivity, they may expect a greater chance in exploiting the gains from increased productivity by claiming for higher wages and also, may wish to claim for it sooner than they would otherwise. Number of trade unions involved is indicated for all the bargaining groups. For the majority of them, the numbers involved does not change over the sample period.

Lastly, we consider factors that influence transaction cost or bargaining power, including factors that affect bargaining environment. The values of past duration are incorporated in order to test the possibility of lagged duration dependence, that is, whether exiting rate from the current contract depends on the history of past contract length. If some group who negotiate infrequently tend to do so due to high cost inherent to this group irrespective of the economic environment, such attribute should be captured by the group specific factors. However, there may be a case for such cost factors or the way it affects the hazard to change overtime. Other exogenous factors such as incomes policy dummies, existence of government's arbitration, if the industry involved is nationalized or it belongs to Wage Council, manufacturing or non-manufacturing, geographical coding, number of unions involved within the bargaining group and number of workers involved in the negotiation (i.e., membership) are also incorporated. Higher such proportion, higher the bargaining power, higher the cost of industrial action, hence stronger the tendency for new negotiation. With respect to the incomes policy, there are several forms such as wage freeze (or stand still) which put severe restraint on the frequency of bargaining and the ceiling policy or the policy with the specified target level for the wage increase which put restraint both on the magnitude and the frequency of bargaining. And the policy can be statutory or voluntary as well as enforced only on public sector or on all the sectors. In this sense, it is not enough to capture the effect of different forms of policies by simple dummies, however, we adopt this for the moment as a preliminary analysis. These dummies are at least capable of distinguishing the effect of these different policies on the magnitude and on the frequency of wage changes.

4.4.2 The wage change component

On the basis of the equation (4.2.5), we now discuss the exact specifications of the explanatory variables which, we have assumed, comes in linearly to the rate of nominal basic wage equation. According to the equation (4.2.5), basic influences on the wage changes are : price/retention ratio changes, labour market condition, firm's internal situation, cost of negotiation and future uncertainty. Note that this equation is conditioned on the fact that the contract has just terminated at T_{ns}. This is the major deviation from the hazard counterpart. Hence, workers, when determining their wage claim, are aware of the total decline in their real wage during the last contract between T_{ns-1} and T_{ns}, as well as the average earnings level and the unemployment level at T_{ns}, and most of all, the fact that the contract terminated at T_{ns}. Knowledge of these information are used basically to make up for the losses workers have experienced throughout the previous contract during which nominal wage was constant. In addition, workers, knowing that the new contract will again fix the level of nominal wage for some time, will demand much more than a mere compensation for the past. They will try incorporating the ex-ante expectation for the future, in particular, of negotiation cost and price/retention ratio uncertainty. In this way, we can consider the determinants of current wage claim in two components over time: compensation for the past and the ex-ante provisions for the future.

First of all, we consider the ex-post compensation against the lost real take home pay at T_{ns} in a comparable form found in the hazard equation. Obvious variables that describes such loss is a change in real take home pay between T_{ns-1} and T_{ns} while nominal wage were fixed at w_{ns-1} (DRR). This variable depicts a rate of change in infla/retention ratio during the previous contract. Additionally, we may consider a variable, DRW1, a change in real pay between T_{ns-2} and T_{ns-1} with nominal wages w_{ns-2} and w_{ns-1}, respectively. This represents a sort of "initial condition" for the ns-th spell. In a world of continuous contracting between two parties, there may be a case when losses experienced during a certain contract is not totally compensated at the end of such contract, but is carried over to be compensated in the successive wage bargain. Hence, a negative DRW1 can be incorporated at the following negotiation into the w_{ns}, although, our sample shows such case is very rare (i.e., DRW1>0 mostly). Still, larger the real wage increase at the last negotiation, smaller the increase we would expect in the following negotiation if everything stays constant. However, as have been discussed in the duration component section, this will not be true if the DRW1 signals as the latest trend in the negotiation cost and the future uncertainty, hence will help predict the future contract length. If that is the case, this explanation may suggest an opposite positive effect of DRW1 on w_{ns}, in which case, we tend to observe groups succeeded in obtaining large real wage change tends to do the same in the future.

Labour market condition, we assume, as considered in the duration equation to be represented by the relative wage and unemployment variables. Relative wage not only captures excess demand but also "fairness" criteria which must be important in this competitive world. Unemployment, global and industry specific can be both included. Hysteresis effect of the outsider-insider consideration can be informally tested by observing their effect on the wage claim. If hysteresis exists, industry specific unemployment, unlike global unemployment, should exert positive effect on the wage claim. However, this test is not very precise since bargaining group which our data is based on contain several unions of several industries, hence industry specific unemployment does not necessarily reflect any particular union's membership consideration. Other variables which may explain excess demand, such as vacancy rate, unemployment/vacancy ratio are also experimented but found insignificantly superior to the unemployment variables.

Firm's internal situation is again mainly reflected by industry specific trading profit since no other disaggregated data, particularly those concerning sales, product market competitiveness and financial situations were available. Global productivity measure represents technical progress. We have tried to seek an effect of input prices but they were found insignificant. Under the profit sharing framework, more profitable the firms is, more rent there is for the workers to exploit in terms of higher wages.

Consider the future uncertainty and negotiation cost. Workers, expecting a rise in inflation during the future contract period, will demand and obtain a larger current wage settlement to compensate them for the future expected rise in prices. If workers know the timing of the next negotiation, their wage claim will be such that it just covers their expected rise in inflation up to the planned date. However, without such knowledge, it all depends on the ex-ante expectation of how long the contract will last. Cost consideration is an important determinant of future uncertainty. If workers feel that such cost is high, they will try to derive the best out of the current negotiation, therefore bargain harder for they know it will be a long time before they can negotiate again. On the other hand, if workers know such cost is low, they may not claim high wages since they feel they can negotiate for higher wages whenever they like. Such negotiation cost is generally determined by the time invariant group-specific factors such as bargaining system employed, number of trade unions involved. However, these factors miss out the elements that fluctuate over time. Probably, the best indicator that can be used to help workers gauge expected future contract length and inflation are simply the value of wage changes over the last consecutive negotiations (i.e., DRW1) or/and the lagged duration.

Obviously, if a wage change is determined a priori as a stage implementation, nature of wage determination process may be completely different. ISTA depicts those determined a priori and amongst the staged wage changes, if the dates as well as the magnitudes are planned, they are coded in the ISFIX. Wage change, when staged, is usually smaller in magnitude than otherwise bargained.

Incomes policies also affect the nature of wage determination. Ceiling policy is made to suppress rate of wage rise henceforth, if effective, should exert a negative influence. The effect of wage freeze policy, on the other hand, is not too clear. During such a policy, there should not be any wage changes unless it is completely backed with a rise in productivity. Twelve month policy, which is aimed to restrict wage rise at most once a year, may raise the size of wage rise because of the lessened frequency of wage changes. Although each policy differs in their degree of enforcement and their extent (some ceiling allowed up to 3.5%, some up to 6.7%), we have resumed using separate dummies for each policies. In this way, we can at least separate out policies with different ceilings and voluntary from statutory policies. With respect to the extent of the coverage, policies enforced only to the public sector have corresponding dummies only for the public sector bargaining groups.

4.5 Summary

What matters to the unions in determining when and by how much to negotiate involves past, present and future consideration given the fact that these two parties, namely unions and employees, renegotiate continuously overtime. How much of uncompensated pressure factors have been carried over into this new contract, what is the prevailed economic situation since the last bargain and how is such a loss compared to the cost of renegotiation? And when the bargain is struck, apart from the ex-post loss compensation, workers need to incorporate the prospect for the future. These are the problems facing the unions. Our formulation of separating the effect into the ex-ante factors, within the spell factors and the future expectation factors should be able to cover all of above accumulating effect. By considering the determinants for the timing of a bargain and wage change given that bargain takes place, we can deal with them separately, and yet, be able to embody full simultaneous relation between the two. The important fact is that these explanatory variables are continuously varying overtime and their entire time path during the course of contract is naturally, considered to affect when the next negotiation occurs. The great advantage of our model as compared to the conventional studies of wage determination process using a simple regression analysis is that our model can take into account of these dynamic effect of time varying regressors on hazard.

Footnotes to chapter 4

1. See, for instance, Black and Kelejian (1972) who formulated wage variable as one-quarter percentage change while assuming uniform quarterly bargaining pattern with one-year contracts, Ashenfelter and Pencavel (1975) who took into account of the proportion of workers who experienced change in wages in a given quarter and Smith and Wilton (1978) who extended to accommodate multi-year variable length contracts and deferred increments by using weights
2. In North America, they typically agree fixed term contracts often with indexation provisions. Length of contract as well as the size of wage increase are determined at the time of bargain, hence, events occurring after the bargain is struck play no role in determining when the next bargain will take place (though, there are cases where contingent wage changes occur). Also, in Britain, between 1979 and 84, Gregory, Lobban and Thomson (1985) found that over 90% of settlements recorded in the CBI survey of pay settlements were made on a strict annual basis. This may be due to an increasing number of settlements which were bargained at a very disaggregated level (phenomena started in late 70's), or the absence of formal incomes policy during the 80's. This rigidity has not always existed in the past (as our data show), and may cease to exists in the future if circumstances change.
3. For a more general utility function $u(w,\bar{w},p,t,n)$, where union may care about the level of employment, the comparative statics are very similar. First order condition of Nash maximization yields:

$$\frac{s}{1-s}\,\frac{p-\bar{p}}{n}=\frac{u-\bar{u}}{u_W}$$

which defines the bargained utility level:

$$u=\left[\frac{s}{1-s}\,\frac{p-\bar{p}}{n}\right]u_W+\bar{u}=\frac{s}{1-s}\,p^*\,u_W+\bar{u}$$

Differentiating this equation gives comparative statics: $dw/d\bar{u}>0$ and $dw/dp^*>0$, where p^* denotes profit per employee above its fall back level. Tax (t), retail price index and input prices have ambiguous effect on wages.

Chapter 5: The Model and Estimation

5.1 Introduction

In this chapter, we build a model that simultaneously determines the occurrence and the magnitude of wage changes based on the economic framework discussed in the preceding chapter.

During our sample period, bargaining groups in U.K. generally negotiated open ended agreements which fixed the nominal wages but left the date of the next negotiation unspecified. There were cases where a contract length was determined a priori as a staged settlement. However, even in such circumstances, there was no legal compulsion to adhere to the agreement, and as a result, deviations from agreed negotiation dates were common. Hence, under such system, re-negotiation was a possibility at any time once a contract began. In other words, there always was a non-zero probability of re-negotiation that was subject to the influence of events that took place after the commencement of the contract, such as a sudden rise in inflation rate or changes in the incomes policy regime. This probability at any point in time is easily associated with the hazard rate, which, in this case is analogous to the probability of negotiation conditional on the elapsed length of the current contract.

For each discrete wage change, we observe the elapsed time since the last settlement and the associated wage variable which comes about as a result of the termination of such contract. This setting, in a duration model, is best represented as the marked failure type (see Cox and Oaks (1984)) where there is a measure of a certain variable, "mark", (i.e., magnitude of wage change) corresponding to each failure (i.e., occurrence of wage change). Generally, there is a strong relationship between the timing of a failure and its mark, and our case is certainly not an exception. This necessitates a simultaneous equation system in order to understand the possibly dynamic interdependence between the two. It is surprising, however, that they are very often treated as exogenous and merely thrown into each others regression equation as an independent variable.

Here, we propose a duration model of the marked failure type that explains the occurrence of wage settlement and the magnitude of concomitant wage changes. In this way, it is possible to distinguish the two separate effects of any external influences, in particular, of incomes policy, on this system of wage determination process. Namely, on delay and moderation of wage changes.

The relation will be based on the joint distribution function of the duration and the wage changes where the former is represented by the hazard rate. Modeling the hazard rate directly enables us to analyze the effect of the accumulating explanatory variables that trigger wage changes more naturally and intuitively than modeling the duration distribution itself. Furthermore, because these pressure factors are continuously evolving and exerting influence on the negotiation probability throughout the spell, it is crucial for the analysis of the observed spell length to depend on the entire time path of these variables rather than their values at any one point in time. This sort of duration relation is a *functional* rather than a function, and is not appropriately estimated by the conventional regression method.

In addition to the introduction of the time varying explanatory variables in the duration component, our model faces some other complications. First is a clear existence of the group-specific effect. Some bargaining groups seem to have tendencies towards relatively long or short duration spells, often in a way that is not clearly related to the observable characteristics such as degree of unionization or the bargaining machinery (e.g., private, wages council, nationalized industry). Thus our model must allow for the presence of unobservable group-specific effects. Secondly, we have very little a priori information about the nature of the unobservable influences in the model. In order to make our model as free as possible from the dependence on *ad hoc* assumptions about the random disturbance terms and their lack of correlation with the explanatory variables, we opt for statistical techniques that are robust in this sense. This method bears *a* further advantage in dealing with the multiple spells nature of the data. Since our data contains multiple observations recorded by a single group over time, the hazard associated with one duration observation depends on the entire bargaining history of its group, in particular, its lagged durations. Therefore, a joint density function for the duration observation is a product of distributions, each conditioned on its history and the unobservable factors. Then, the marginal distribution can only be derived by going through a multiple integrations which will be terribly messy for any functional forms assumed for the unobservables. Inclusion of the time varying explanatory variables can further complicate such integration. Our model is totally devoid of such multiple integrations as well as any *ad hoc* distributional assumptions, yet, capable of deriving the estimates of interesting parameters.

Final note on the applicability of the model developed here. Although we focus on the particular issue of the wage determination process, this model can be applied to many other phenomena which involve a study of spell durations whose termination probability may depend on the events occurring during the spell. In particular, if there were an endogenous "response" variable associated with each duration spell, their dynamic relation can be determined simultaneously. For instance, the strike duration can be studied together with the wage level achieved at the end of each conflict.

In this chapter, we will discuss our statistical model separately for the wage change and the duration component, followed by a discussion of the estimation methodology. Last section is devoted to the results.

5.2 The statistical model

5.2.1 Notation

We have data on N bargaining groups, indexed by a suffix n. We observe S_n bargains for the n-th group over a historical period, and these are indexed by a suffix s. For each bargain, we observe the following:

δ_{ns} the s-th spell duration for the n-th bargaining group

w_{ns} the rate of change in nominal wage brought about at the s-th bargain, at the termination of the s-th spell.

z_{ns}, ξ_{ns} vectors of explanatory variables (used respectively in the duration and wage components of the model) which are spell specific: they may vary across the spells, but invariant within the spells: they may contain lagged δ_{ns} and w_{ns}.

c_n, d_n vectors of explanatory variables that vary across bargaining groups but are constant over the whole observation period: they represent time invariant idiosyncrasies of each group and tend to include qualitative variables.

$x_n(t)$ vectors of exogenous variables that vary continuously over time, both within and between contracts.

u_n an unobservable, invariant effect specific to the n-th group.

v_{ns}, ν_{ns} spell-specific random errors relating to the duration and wage components of the model, respectively.

5.2.2 The joint distribution of duration and wage increase

We assume the existence of a random group specific effect u_n. Conditional on all unobservable variates and past history, δ_{ns} and w_{ns} have a joint distribution that can be decomposed as follows into the duration and the wage change component.

$$f(\delta_{ns},w_{ns}|c_n,d_n,z_{ns},\xi_{ns},X_{ns},u_n,v_{ns},\nu_{ns}) = f(\delta_{ns}|c_n,z_{ns},X_{ns},u_n,v_{ns}) \\ * f_w(w_{ns}|d_n,\xi_{ns},\delta_{ns},u_n,\nu_{ns}) \qquad (5\text{-}2\text{-}1)$$

where X_{ns} represents the time path of $x_n(t)$ between the (s-1)-th wage bargain onwards. f_w is conditioned on the current duration, δ_{ns}, hence the density of current wage change is conditional on the timing of the bargain. Note that with z_{ns} and ξ_{ns} defined appropriately, the joint distribution is conditional on all the past bargains for this group, n.

By parameterizing the distribution in the form (5-2-1), we are assuming that the timing of negotiation is determined prior to the size of the resulting wage increase. In other words, negotiators first decide when to increase and then consider the size of increase that is appropriate. This is reasonable considering that, in practice, the amount of wage change is determined at the time of bargaining which can be different from what negotiators wished to achieve prior to the negotiation (section 4.4.2). In this way, the length of the current contract may affect the size of wage changes achieved at the end of such contract, although the current contract length is not affected by the amount of realized wage changes agreed when the contract is over. We consider the two components of (5-2-1) separately.

5.2.3 The duration component

The distribution of observed duration is conditioned on the information set at the time of the last negotiation and the time path of the explanatory variables $x_n(t)$ since then. We start indirectly with the hazard function rather than the distribution of duration itself, since the hazard is the most natural way of representing the accumulating pressures on negotiators, hence makes more economic sense in interpreting the effects of these pressure variables. Here, the hazard rate at duration t is the probability of a renegotiation occurring at the elapsed duration conditional on the present and past circumstances and the fact that the spell has already lasted for t. In other words, it is the instantaneous exit rate from the current contract which has lasted for t duration periods. To be more precise:

$$\Pr(\delta_{ns} \in (t,t+dt) \mid \delta_{ns} \geq t,c_n,z_{ns},x_n(t),u_n) = h(t|c_n,z_{ns},x_n(t),u_n)dt \quad (5\text{-}2\text{-}2)$$

where $h(t|.)$ is the hazard function. Adopting the conventional proportional hazard specification:

$$h(t|c_n,z_{ns},x_n(t),u_n) = \exp\{\beta' x_n(t)+\gamma_0+\gamma_1' c_n+\gamma_2' z_{ns}+u_n\}\ h_0(t;\alpha) \qquad (5\text{-}2\text{-}3)$$

where α, β, γ_0, γ_1 and γ_2 are fixed parameters requiring estimation. The function $h_0(t;\alpha)$ denotes the baseline hazard which represents the pressure on negotiators engendered purely by the passage of time. When this baseline hazard equals 1, the conditional probability of renegotiation depends on time only through the time varying covariates $x_n(t)$ and not on time itself.

Using standard results (see, for example, Pudney (1989, chapter 6)) the probability density of duration can be written as:

$$f_\delta(\delta_{ns}|c_n,z_{ns},X_{ns},u_n,v_{ns}) = h(\delta_{ns}|c_n,z_{ns},x_n(t),u_n,v_{ns})\ \exp\{-I_{ns}\} \qquad (5\text{-}2\text{-}4)$$

where I_{ns} is the integrated hazard:

$$\begin{aligned} I_{ns} &= \int_0^{\delta_{ns}} h(t|c_n,z_{ns},X_{ns},u_n,v_{ns})dt \\ &= \exp(\gamma_0+\gamma_1' c_n+\gamma_2' z_{ns}+u_n+v_{ns}) \int_0^{\delta_{ns}} \exp\{\beta' x_n(t)\}h_0(t;\alpha)dt \qquad (5\text{-}2\text{-}5) \end{aligned}$$

The theory of the hazard function implies that the integrated hazard has a standard exponential distribution, thus the variable $\iota_{ns}=-\ln(I_{ns})$ is distributed unconditionally as type I extreme value distribution. To see this, the Jacobian of the transformation from δ_{ns} to ι_{ns} is:

$$\left| \frac{\partial\, \delta_{ns}}{\partial\, \varepsilon_{ns}} \right| = \left| \frac{I_{ns}}{h\,(\delta_{ns}|z_{ns},X_{ns},u_n,v)_{ns}} \right| \qquad (5\text{-}2\text{-}6)$$

and thus the density function of ι_{ns} is the following:

$$f\,(\iota) = \exp(\,-\iota - e^{-\iota}) \qquad (5\text{-}2\text{-}7)$$

This is the density of a type I extreme value distribution with mean $\psi(1)$ and variance $\psi'(1)$, where $\psi(.)$ is the digamma function, $d\ln\Gamma(\xi)/d\xi$.

Now define a variable:

$$\eta_{ns} = - \ln \left[\int_0^{\delta_{ns}} \exp \{\beta' x_n(t)\}\, h_0(t;\alpha)\, dt \right] \tag{5-2-8}$$

Equations (5-2-5) and (5-2-8) imply a random effects regression structure with η_{ns} as its dependent variable:

$$\eta_{ns} = \gamma_0^* + \gamma_1' c_n + \gamma_2' z_{ns} + u_n + \varepsilon_{ns} \tag{5-2-9}$$

where the intercept is $\gamma_0^* = \gamma_0 + \psi(1)$ and the spell-specific error term is $\varepsilon_{ns} = \iota_{ns} + v_{ns} - \psi(1)$ which is constructed to have a zero conditional mean.

Note that, although ι_{ns} must be distributed as type I extreme value, much can be said about ε_{ns} unless we assume a particular distributional form for v_{ns}. Even if such assumption is to be made, conventional maximum likelihood estimation will not be straightforward. Recall that the joint density function is the product of conditional distributions, each conditioned on the past contract history and the unobservables that include u_n and v_{ns}. The maximum likelihood estimation then requires the distributional assumptions for both u_n and v_{ns} to conduct a formidable multiple integration that removes these unobservables to derive a marginal density for the observed δ_{ns}. Such distributional assumption tends to be very *ad hoc* and the likely correlations between the error term and the explanatory variables, in particular, the lagged durations, creates further difficulties. Moreover, when the hazard involves the time varying explanatory variables, the integrand becomes a very complicated function of t, which leads to a very complicated and expensive numerical integration. For this reason, we choose to work with the regression formulation, (5-2-9), making only relatively weak assumption on the structure of serial correlation across settlements[1] and that ε_{ns} has a conditional mean of zero. In this sense, our methods will be semi-parametric.

We now turn to the dependent variable, η_{ns} which is not directly observable because of its dependence on the parameters of the model. The main problem in constructing η_{ns} concerns the vector $x_n(t)$. These are potentially continuous variables, but are only observed at discrete monthly or quarterly intervals. There are two options open to us. We can either interpolate these discrete observations with some form of spline function (as is done by Diamond and Hausman (1984)), or we can treat the time path of x_n as a step function with discrete jumps at each of the observation points. The latter alternative is much simpler, and not obviously more arbitrary, and is the approach adopted here.

Assume that, for the s-th bargain struck by group n, the timing of observation is as illustrated in figure 5.1. Within the interval between the (s-1)-th and s-th bargains, δ_{ns}, there are k_{ns} monthly observations on $x_n(t)$; these occur at dates $\tau_{ns}[0]$, $\tau_{ns}[1]$, $\tau_{ns}[2]$ ....$\tau_{ns}[k_{ns}]$. The dates of the month of the observation immediately prior to the start of the spell is $\tau_{ns}[0]$, and the date of the first observation recorded after the end of the spell is $\tau_{ns}[k_{ns}+1]$.

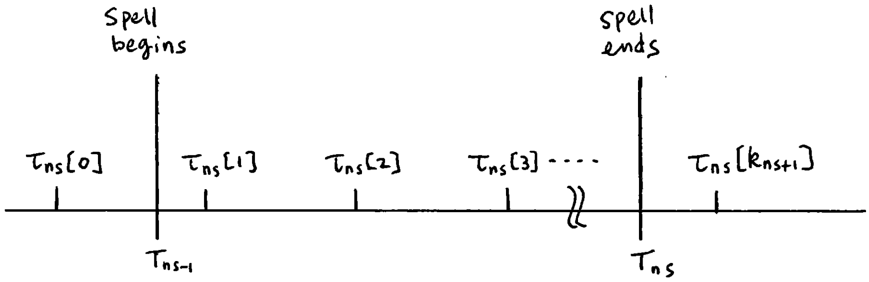


Figure 5.1 The timing of observations

If $x_n(t)$ remains fixed between each monthly interval during the duration, δ_{ns}, η_{ns} can be written as:

$$\eta_{ns} = -\ln \{ \sum_{i=0}^{k_{ns}} \exp\left[\beta' x_n(\tau_{ns}[i])\right] \int_{a_{ns}[i]}^{b_{ns}[i]} h_0(t;\alpha)dt \} \qquad (5\text{-}2\text{-}10)$$

where:

$$a_{ns}[i] = \max \{ T_{ns-1}, \tau_{ns}[i]\} - T_{ns-1} \qquad (5\text{-}2\text{-}11)$$

$$b_{ns}[i] = \min \{ T_{ns}, \tau_{ns}[i+1]\} - T_{ns-1} \qquad (5\text{-}2\text{-}12)$$

There is no convincing economic theory available to tell us the nature of the baseline hazard function, $h_0(t;\alpha)$, beyond a presumed tendency towards an annual pay round. Approximating $h_0(.)$ directly requires the computation of the integrals in (5-2-10). We have therefore approximated the integrated baseline hazard,

$\Psi(t;\alpha)=\int h_0(t;\alpha)\, dt$.

Generally, our dependent variable is:

$$\eta_{ns} = -\ln \left\{ \sum_{i=0}^{k_{ns}} \exp\left[\beta' x_n(\tau_{ns}[i])\right] \left[\Psi(b_{ns}[i]) - \Psi(a_{ns}[i])\right] \right\} \qquad (5\text{-}2\text{-}15)$$

Note that neither the linear form $\beta'x_n$ nor the polynomial in (5-2-13) has an intercept, since this can be absorbed into γ^*_n in (5-2-9).

For given parameter values, and any of our approximations, $\Psi(t;\alpha)$, η_{ns} is easily computed. However, the specification of $\Psi(t;\alpha)$ has to be very carefully chosen, for otherwise, the model becomes quickly unidentified. As can be seen from the definition (5-2-15), if there exists a parameter vector, α^* say, for which $\Psi(t;\alpha^*)$ is constant for any values of t, there is an identification problem. If we set β to zero and allow α to approach α^*, the η_{ns} will be constant given that $\Psi(0;\alpha)=0$ by definition. The dependent variable in a differenced equation, $\Delta\eta_{ns}$, becomes zero, hence, choosing the γ parameters to be zero will give a zero residual: a perfect fit for the regression equation (5-2-8). The following is examples of such approximations that are inappropriate for this reason:

$$\Psi(t;\alpha) = \exp(\alpha t) - 1$$
$$\Psi(t;\alpha) = t^{\alpha}$$

In both cases, $\alpha^*=0$ leads to a constant dependent variable for any value of t. The integrated hazard corresponding to the second example is the Weibull duration distribution, whose unidentification problem is well known in the context of static regression analysis. There, the dependent variable is log(t) and the structural parameter, β cannot be separately estimated from α. In practice, these identification problems can become more intractable by the insensitivity of numerical computation process. Nonetheless, we have experimented with a wide variety of forms for $\Psi(t;\alpha)$ with appropriate normalizations in order to avoid above identification problems. These include a flexible polynomial:

$$\Psi(t;\alpha) \approx \exp\{t + \alpha_2 t^2 + ... + \alpha_m t^m\} - 1 \qquad (5\text{-}2\text{-}13)$$

which has to have constraints on the α's so that a negative value for $\Psi(t;\alpha)$ is avoided. Another approximation is the constant baseline hazard with a superimposed annual jump:

$$\Psi(t;\alpha) \approx t+\alpha_2\lambda(t) \qquad (5\text{-}2\text{-}14)$$

with a restriction that $\alpha_2 \geq 0$ and:

$$\lambda(t) = \begin{cases} 0 & \text{for } t<48 \text{ weeks} \\ t-48 & \text{for } 48 \le t \le 72 \text{ weeks} \\ 12 & \text{for } t>72 \text{ weeks} \end{cases}$$

We have also experimented with a step function with jumps at fixed 13-week intervals.

5.2.4 The wage change component

The rate of wage change, conditional on the occurrence and timing of a bargain, can be modeled by means of a linear regression:

$$w_{ns} = \phi_0 + \phi_1 d_n + \phi_2' \xi_{ns} + \phi_3 u_n + \nu_{ns} \qquad (5\text{-}2\text{-}16)$$

where d_n and ξ_{ns} are vectors of observable explanatory variables. Note that ξ_{ns} may include current and lagged duration and lagged wage changes. The random effect, u_n, is included to allow for the same group specific factors that influence both the duration and wage change decisions of the bargaining. The disturbance term, ν_{ns}, is assumed to have zero means conditional on the explanatory variables.

5.3 Estimation: fixed effects

We have already ruled out maximum likelihood for its complexity and lack of robustness and chosen instead to work with the regression format of equations (5-2-8) and (5-2-16). However, ordinary regression analysis is inappropriate for the following reasons.

(i) The dependent variable of the duration equation, η_{ns}, is a nonlinear function of endogenous variable, δ_{ns}, and some unknown parameters, which makes it inappropriate to use the standard nonlinear least squares format.
(ii) In practice, lagged durations and wage changes will be included in the spell-specific explanatory variables, z_{ns} and ξ_{ns}, that are not independent of the group-specific effect, u_n. This correlation leads to inconsistent least squares estimation.
(ii) Since we have little a priori information about the nature of u_n, it is best to avoid making strong assumptions about it. Therefore, elimination of u_n before estimation is appropriate.

The first and the second issues call for a general method of moment type estimator where the moment restriction is considered with the appropriate instrumental variables which, according to (iii), should be applied to the transformed model so that the group effects do not appear. These objections have already been discussed in the literature (see Nickell (1981) and Bhargava and Sargan (1983)), but the kind of transformation and the form of moment restrictions particularly, the set of appropriate instruments, have to be uniquely considered for this model so as to avoid any inconsistencies.

Consider the method of eliminating u_n. The usual within-group transformation (i.e., covariance model) is inappropriate for the familiar reason that the transformed error, $\varepsilon_{ns}-\bar{\varepsilon}_n$, is correlated with all the observed durations, wage changes and spell-related values of $x_n(t)$. Although the asymptotic bias created as a result of such correlations will approach zero as the time dimension of our sample approaches infinity, that is not the case here[2]. This makes it almost impossible to find a set of admissible instruments. Instead, we have chosen to use a simple first-difference transformation. Under this transformation, the admissible instruments can be found amongst suitably lagged variables as long as we assume the serial correlation structure between the settlements. Note that Arellano's (1989) estimator would be a worthwhile alternative in deriving a set of admissible instruments. Thus, the system transformed for the estimation is:

$$\iota_{ns} - \iota_{ns-1} = \gamma '(z_{ns} - z_{ns-1}) + \varepsilon_{ns} - \varepsilon_{ns-1} \qquad (5\text{-}3\text{-}1)$$
$$w_{ns} - w_{ns-1} = \phi_2'(\xi_{ns} - \xi_{ns-1}) + \nu_{ns} - \nu_{ns-1} \qquad (5\text{-}3\text{-}2)$$

Another difficulty, unique to this model, concerns the treatment of the exogeneity assumptions. Suppose for example that the variables in the vector $x_n(t)$ are exogenous in the sense that their entire trajectories are statistically independent of the unobserved part of the wage determination process: a very strong definition of exogeneity. This implies that the value of $x_n(t)$ at any *exogenously* determined time t is also independent of all random elements in the model. However, the value of $x_n(t)$ observed at an *endogenously* determined time t is a variable endogenous to the bargaining model. To see this more clearly, consider the likelihood function of the entire observed samples while assuming that each duration and wage change component depends only on their most recent lagged values:

$$f(\delta_1,\delta_2,..,\delta_m,w_1,w_2,..w_m | \underline{x}^*,\underline{\xi}^*,\underline{\nu},\underline{\varepsilon}) =$$

$$f_\delta(\delta_1|z_1,x_1^*,\varepsilon_1)f_w(w_1|\delta_1,\xi_1^*,\nu_1)\ldots$$
$$\ldots f_\delta(\delta_m|\delta_1..\delta_{m-1},w_1..w_{m-1},x_m^*,\varepsilon_m)f_w(w_m|\delta_1..\delta_m,w_1..w_{m-1},\xi_m^*,\nu_m)$$

where $\underline{x}^*$ and $\underline{\xi}^*$ are the time-path of exogenous variables excluding the lagged endogenous variables that affect the duration and wage distributions. Subscript, n, has been dropped for simplicity. More specifically, for each s-th observation on duration, x_s^* includes spell specific exogenous variables known at the beginning of the s-th spell (i.e., T_{s-1}) and the time-path of exogenous variables from then onwards. On the other hand, for the s-th wage change observation, ξ_s^* includes variables observed up to and including the end of the s-th spell (i.e., T_s). The difficulty in defining the exogeneity of variables occurs particularly with respect to the duration component. For a typical s-th observation of duration, recall:

$$\eta_s = \gamma_0^* + \gamma_1'c + \gamma_2'z_s + u + \varepsilon_s$$

Where c, z and X_{ns} in η_s, excluding the lagged endogenous variables, constitutes x_s^* in above notation. Then:

$$E(u+\varepsilon_s|\delta_1..\delta_{s-1},w_1..w_{s-1},x_s) = 0$$
$$E(u+\nu_s|\delta_1..\delta_s,w_1..w_{s-1},\xi_s) = 0$$

Any exogenous variables observed at times related in any way to the duration of the current spell, δ_s, are not exogenous to the s-th duration equation since they are determined in part by the length of the current spell, which in turn depends on ε_s. While for the wage equation, since the wage is determined after the decision to terminate the contract is made, the current duration, δ_s, is strictly exogenous with respect to ν_s, and so as the other variables as long as they are observed prior to or at the end of the s-th bargain (i.e., up to and including T_s).

The presence of lagged endogenous variables in z_s or ξ_s makes this problem even more complex for the differenced equation (5-3-1) and (5-3-2). Now, the two compound error terms, $\{\varepsilon_s-\varepsilon_{s-1}\}$ and $\{\nu_s-\nu_{s-1}\}$ are correlated with $\{z_s-z_{s-1}\}$ and $\{\xi_s-\xi_{s-1}\}$, respectively. Assuming no serial correlation in $\{\varepsilon_s\}$:

$$E(\varepsilon_s-\varepsilon_{s-1}|\delta_1\ldots\delta_{s-2},w_1\ldots w_{s-2},x_{s-1}^*) = 0$$

However,

$$E(\varepsilon_s - \varepsilon_{s-1} | \delta_1 \ldots \delta_{s-1}, w_1 \ldots w_{s-1}, x^*_s) \neq 0$$

Thus, the exogenous variables x^*_{s-1} and the lagged endogenous variables up to (s-2)th bargain are appropriate instruments for the duration equation (5-3-1). In other words, the variables observed at the end of (s-2)-th spell (i.e., T_{s-2}) or earlier are all admissible instruments. Note that, in addition, exogenous variables observed after the end of the (s-2)-th spell are still valid instruments as long as their observation points are not related to the time of bargaining, T_{s-1} or T_s. Consequently, in practice, admissible instruments are all the explanatory variables observed at the (s-2)-th bargaining (i.e., T_{s-2}) or earlier, as well as the exogenous variables observed at fixed intervals of 12 and 24 months after the end of the (s-2)-th spell (i.e., T_{s-2}+12-months and T_{s-2}+24-months). This is because the spell length of around 12 months is the average in the data.

For the wage equation (5-3-2), ξ^*_s is correlated with ν_{s-1} since ν_{s-1} affects the timing of the following s-th bargain, the time that ξ^*_s is observed. Nonetheless, since the wage equation is conditioned on the timing of wage change, δ_{s-1} is strictly exogenous with respect to ν_{s-1}. Therefore, δ_{s-1} and any exogenous variables observed at the end of the (s-1)-th spell (i.e., T_{s-1}) are all admissible instruments.

$$E(\nu_s - \nu_{s-1} | \delta_1, \ldots \delta_{s-1}, w_1 \ldots w_{s-2}, \xi^*_{s-1}) = 0$$

Estimating equation forms (5-3-1) and (5-3-2) has the advantage that the unobservable group effects, u_n, are eliminated, but it also has some disadvantages. The intercepts and the observable group idiosyncrasies, c_n and d_n, are also eliminated and their coefficients are not directly estimable. Moreover, in performing this covariance transformation, we are eliminating a large amount of variation in the sample, therefore sacrificing efficiency. Nevertheless, this is the price one has to pay for computational convenience and statistical robustness.

5.3.1 Estimation of the wage equation

In writing our estimators, we use the following notation for sample moments. A matrix M_{ab} is the matrix of cross-moments between the differenced forms of two vector variables, a_{ns} and b_{ns}, while a vector m_{ab} is the vector of cross-moments between the differenced vector a_{ns} and the differenced scalar b_{ns}. Thus:

$$M_{ab} = N^{-1} \sum_{n}^{N} \sum_{s}^{Sn} (a_{ns} - a_{ns-1})(b_{ns} - b_{ns-1})' \qquad (5\text{-}3\text{-}3)$$

$$m_{ab} = N^{-1} \sum_{n}^{N} \sum_{s}^{Sn} (a_{ns}-a_{ns-1})(b_{ns}-b_{ns-1}) \tag{5-3-4}$$

where b_{ns} is a vector in (5-3-3) but a scalar in (5-3-4).

Assume a vector of variables, q_{ns}, to contain suitable instruments. Then, this vector will satisfy the following conditions[3]:

$$\operatorname*{plim}_{N\to\infty} m_{q\varepsilon} = 0 \tag{5-3-5}$$

$$\operatorname*{plim}_{N\to\infty} m_{q\nu} = 0 \tag{5-3-6}$$

Provided the matrix inverses exist, the usual instrument variable estimator can be defined for ϕ_2 in the linear rate of wage change equation, (5-3-2). This can be derived by considering the minimum of the criterion function with respect to ϕ_2:

$$m_{q\nu}' \, (W)^{-1} \, m_{q\nu} \tag{5-3-7}$$

in which, we can choose the weighting matrix to be $W = M_{qq}$, and derive:

$$\hat{\phi}_2 = (M_{\xi q} M_{qq}^{-1} M_{q\xi})^{-1} \, M_{\xi q} M_{qq}^{-1} m_{qw} \tag{5-3-8}$$

This is consistent for $N \to \infty$ under standard conditions. To see its asymptotic distribution, expand (5-3-8) to derive:

$$\sqrt{N} \, (\hat{\phi}_2 - \phi_2) = (\bar{M}_{\xi q} \bar{M}_{qq}^{-1} \, \bar{M}_{q\xi})^{-1} \, \bar{M}_{\xi q} \bar{M}_{qq}^{-1} \, (\sqrt{N} \, m_{q\nu}) + o_p(1) \tag{5-3-9}$$

where $\bar{M}_{\xi q}$ and $\bar{M}_{qq}$ are plims of $M_{\xi q}$ and M_{qq}. The last component of (5-3-9) can be written as:

$$N^{1/2} \, m_{q\nu} = N^{-1/2} \sum_{n} e_n \tag{5-3-10}$$

where:

$$e_n = \sum_{s=2}^{S_n} (q_{ns} - q_{ns-1})(\nu_{ns} - \nu_{ns-1}).$$

But e_n has a zero mean vector and a covariance matrix that can be consistently estimated by:

$$\hat{A} = \sum_{s=2}^{S_n} \sum_{t=2}^{S_n} \hat{e}_n \hat{e}_n \qquad (5\text{-}3\text{-}11)$$

where:

$$\hat{e} = \sum_{s=2}^{S_n} (q_{ns}-q_{ns-1})(w_{ns}-w_{ns-1}-\hat{\phi}_2'(\xi_{ns}-\xi_{ns-1}))$$

Specifically, expression of $\hat{A}$ reduces to the following if we assume ν_{ns} to be independent and homoscedastic both across groups and over time (i.e., $E(\nu^2_{ns})=\sigma^2_\nu$ for all n and s):

$$\hat{A} = \hat{\sigma}^2_\nu \sum_{s=2}^{S_n} \sum_{t=2}^{S_n} (2\Delta_{st}-\Delta_{s,t-1}-\Delta_{t,s-1})(q_{ns}-q_{ns-1})(q_{nt}-q_{nt-1})' \qquad (5\text{-}3\text{-}12)$$

where Δ_{st} is the Kronecker delta. And the variance of ν_{ns} conditional on $q_{ns}$and q_{ns-1} can be estimated consistently by:

$$\hat{\sigma}^2_\nu = \frac{\sum_{n=1}^{N} \sum_{s=2}^{S_n} (w_{ns} - w_{ns-1} - \hat{\phi}_2'(\xi_{ns}-\xi_{ns-1}))^2}{2 \sum_{n=1}^{N} (S_n - 1)} \qquad (5\text{-}3\text{-}13)$$

It then follows from the standard arguments that $\sqrt{N}(\hat{\phi}_2 - \phi_2)$ is asymptotically (in N) normal and that a valid asymptotic approximation to $\mathrm{cov}(\hat{\phi}_2)$ is:

$$\hat{V}_{\phi_2} = [M_{\xi q}M^{-1}_{qq}M_{q\xi}]^{-1}[M_{\xi q}M^{-1}_{qq}(\hat{A})\, M^{-1}_{qq}M_{q\xi}][M_{\xi q}M^{-1}_{qq}M_{q\xi}]^{-1}/N^2 \qquad (5\text{-}3\text{-}14)$$

In applying our estimators, it is important to check the specification of the model by means of some suitable misspecification test. The obvious test to apply in this context is a development of Sargan's (1985) Instrumental Variable χ^2 test for covariance between the errors and instruments. Since $\hat{\phi}_2$ we have just derived is not efficient, the form, $m_{q\hat{\nu}}'(M_{qq})^{-1}m_{q\hat{\nu}}/\sigma^2_\nu$, does not have the usual asymptotic χ^2 distribution even under the hypothesis of correct specification. Instead, expand the estimator of expression (5-3-10) about the true parameter value:

$$\begin{aligned}\sqrt{N}\, \hat{m}_{q\nu} &= [\, I - \bar{M}_{q\xi}(\bar{M}_{\xi q}\bar{M}^{-1}_{qq}\bar{M}_{q\xi})^{-1}\bar{M}_{\xi q}\bar{M}^{-1}_{qq}\,]\, \sqrt{N}\, m_{q\nu} + o_p(1) \qquad (5\text{-}3\text{-}15)\\ &= \bar{B}\, \sqrt{N}\, m_{q\nu} + o_p(1)\end{aligned}$$

Therefore:

$$N\ \hat{m}_{q\nu}'(\bar{B}\bar{A}\bar{B}')^{-}\ \hat{m}_{q\nu} \quad \xrightarrow[D]{} \chi^2\ (r) \qquad (5\text{-}3\text{-}16)$$

where $\bar{A}=\mathrm{plim}\hat{A}$, $[]^{-}$ denotes a generalized inverse and $r=\mathrm{rank}(\bar{B}\bar{A}\bar{B}')$ is the number of instruments minus the number of coefficients in ϕ_2. $\bar{A}$ can be consistently estimated by $\hat{A}$, by expression (5-3-11) which is based on the consistent estimator of ϕ_2, $\hat{\phi}_2$. Thus, an asymptotically valid specification test statistic which will follow $\chi^2(r)$ under the null hypothesis of correct specification is the following:

$$C = N\ \hat{m}_{q\nu}'[\ \hat{B}\hat{A}\hat{B}'\ ]^{-}\ \hat{m}_{q\nu} \qquad (5\text{-}3\text{-}17)$$

where:

$$\hat{B} = I - M_{q\xi}[M_{\xi q}M_{qq}^{-1}M_{q\xi}]^{-1}M_{\xi q}M_{qq}^{-1} \qquad (5\text{-}3\text{-}18)$$

However, estimator $\hat{\phi}_2$ described by (5-3-8) is consistent but not efficient. Efficient estimator of ϕ_2 can be derived as a second step, once having derived $\hat{\phi}_2$. Let:

$$\Delta\hat{\nu}_n = \Delta w_n - \Delta\xi_n\,\hat{\phi}_2\ ,$$
$$M_{q\Omega q} = 1/N \sum_n \{\Delta q_n'\Delta\hat{\nu}_n\ \Delta\hat{\nu}_n'\ \Delta q_n\}$$

And use $M_{q\Omega q}$ as a weighting matrix, W, for the minimizing problem (5-3-7):

$$\min_{\phi_2}\ M_{\nu q}\ M_{q\Omega q}^{-1}\ M_{q\nu} \qquad (5\text{-}3\text{-}19)$$

Hence the optimal estimator for ϕ_2 is:

$$\tilde{\phi}_2 = (M_{\xi q}\ M_{q\Omega q}^{-1}\ M_{q\xi})^{-1}\ M_{\xi q}\ M_{q\Omega q}^{-1}\ m_{qw} \qquad (5\text{-}3\text{-}20)$$

Its asymptotic variance is now estimated by:

$$\mathrm{avar}(\tilde{\phi}_2) = (M_{\xi q}\ M_{q\Omega q}^{-1}\ M_{q\xi})^{-1}\ /\ N \qquad (5\text{-}3\text{-}21)$$

Given this efficient second step estimator, the Sargan statistics for over identifying restrictions which is asymptotically equivalent to (5-3-17) is straightforwardly:

$$\begin{aligned} sar &= N\,(M_{\tilde{\nu} q}\, M^{-1}_{q\Omega q}\, M_{q\tilde{\nu}}) && (5\text{-}3\text{-}22)\\ &= (\sum_n \Delta\tilde{\nu}'_n \Delta q_n)\,(\sum_n \Delta q'_n \Delta\hat{\nu}_n \Delta\hat{\nu}'_n\, \Delta q_n)^{-1}\,(\sum_n \Delta q'_n\, \Delta\tilde{\nu}_n) \quad \sim \chi\,(r) \end{aligned}$$

where $\Delta\tilde{\nu}_n$ is the residual given $\tilde{\phi}_2$, $\Delta\hat{v}_n$ is the residual given $\hat{\phi}_2$ and r is the number of instruments minus the number of coefficients in ϕ_2. This can be seen by rewriting above statistic as:

$$sar = N\, m_{\nu q}\, C_N\, (I_p - G(G'G)^{-1}G')C_N'm_{q\nu} \qquad (5\text{-}3\text{-}23)$$

where: $C_N C_N' = M^{-1}_{q\Omega q}$, $G = C_N'\, M_{q\xi}$ and p is the number of instruments. Also, $\sqrt{N} m_{\nu q} C_N \overset{a}{\sim} N(0,I_p)$ and $\text{rank}(I_p\text{-}G(G'G)^{-1}G')=r$, hence (5-3-23) follows $\chi^2_{(r)}$.

The intercept, ϕ_0, and the coefficients of the group-specific explanatory variables, ϕ_1, cannot be estimated without some assumption about the nature of the unobservable group-specific effects, u_n. Assume that u_n has zero mean conditional on all exogenous variables, then a regression (with an intercept) of $w_{ns}\text{-}\hat{\phi}_2\xi_{ns}$ or $w_{ns}\text{-}\tilde{\phi}_2\xi_{ns}$ on d_n will generate consistent estimates of ϕ_0 and ϕ_1. Consider only the second step estimator, $\tilde{\phi}_2$ and suppose that the first element of vector d_n is one. And let ϕ_1 now has ϕ_0 as its first element. Then:

$$\tilde{\phi}_1 = (\sum_n S_n\, d_n\, d_n')^{-1}\, (\sum_n \sum_s d_n(w_{ns}\text{-}\tilde{\phi}_2'\xi_{ns})') \qquad (5\text{-}3\text{-}24)$$

Then, its asymptotic covariance matrix is consistently estimated by $\sum_n a_n a_n'$ where:

$$\begin{aligned} a_n = [\, \sum_n S_n(d_n d_n')\,]^{-1}[\, \sum_n\, (d_n S_n(\phi_3\tilde{}\, u_n) + d_n\sum_s \tilde{\nu}_{ns} \\ + d_n\sum_s \xi'_{ns}(M_{\xi q}\, M^{-1}_{q\Omega q}\, M_{q\xi})^{-1}\, M_{\xi q}\, M^{-1}_{q\Omega q}\, m_{q\tilde{\nu}})] \end{aligned}$$

where $(\phi_3\tilde{}\, u_n)=\sum_s c_{ns}/S_n$, $\tilde{\nu}_{ns}=c_{ns}\text{-}(\phi_3\tilde{}\, u_n)$ and $c_{ns}=w_{ns}\text{-}\tilde{\phi}_2'\xi_{ns}$. Note that:

$$\text{plim}\, \frac{\sum_n\sum_s (\, w_{ns} - \hat{\phi}_0 - \hat{\phi}_1' d_n - \hat{\phi}_2'\xi_{ns})}{\sum_n S_n} = \sigma^2_\nu + \phi^2_3\sigma^2_u \qquad (5\text{-}3\text{-}25)$$

and this residual variance can be used to estimate ϕ_3 when independent estimate of σ^2_u is available.

5.3.2 Estimation of the duration equation

Although the duration component of the model is highly nonlinear, only slight modifications to the estimation described in the previous section are necessary. First step instrumental variable estimator is calculated by minimizing the following criterion function which is analogous to (5-3-7). However, such minimum has to be located numerically since the dependent variable is a non-linear function of the unknown parameters.

Hansen (Econometrica?)

$$\psi(\vartheta) = m_{q\varepsilon}'(M_{qq})^{-1}m_{q\varepsilon} \qquad (5\text{-}3\text{-}26)$$

where $\vartheta' = (\beta', \alpha', \gamma_2')$ and the vector q_{ns} contains the admissible instrumental variables which are different from those used to estimate the wage equation.

Denote this estimator $\hat{\vartheta}$. Under the standard regularity conditions, this estimator is consistent and asymptotically normal. Our asymptotic approximation to the covariance matrix of $\hat{\vartheta}$ is:

$$\hat{V}_{\vartheta} = [M_{dq}M_{qq}^{-1}M_{qd}]^{-1}[M_{dq}M_{qq}^{-1}(\hat{A})\ M_{qq}^{-1}M_{qd}][M_{dq}M_{qq}^{-1}M_{qd}]^{-1}/N^2 \qquad (5\text{-}3\text{-}27)$$

where:

$$M_{dq} = \partial\ \hat{m}_{q\varepsilon}' \ / \ \partial\hat{\vartheta}$$

and $\hat{A}$ is now redefined as:

$$\hat{A} = \sum_{s=2}^{S_n} \sum_{t=2}^{S_n} \hat{e}_n\, \hat{e}_n \qquad (5\text{-}3\text{-}28)$$

where:

$$\hat{e} = \sum_{s=2}^{S_n} (q_{ns}-q_{ns-1})(\hat{\eta}_{ns}-\hat{\eta}_{ns-1}-\hat{\gamma}_2'(z_{ns}-z_{ns-1})) \qquad (5\text{-}3\text{-}29)$$

The Sargan test statistic is a similar variant of (5-3-17):

$$C = N\ \hat{m}_{q\varepsilon}'[\ \hat{B}\hat{A}\hat{B}'\ ]^{-}\ \hat{m}_{q\varepsilon} \qquad (5\text{-}3\text{-}30)$$

where:

$$\hat{B} = I - M_{qd}[M_{dq}M_{qq}^{-1}M_{qd}]^{-1}M_{dq}M_{qq}^{-1} \qquad (5\text{-}3\text{-}31)$$

Under the null hypothesis of correct specification, C is asymptotically $\chi^2(r)$, where r is the number of instruments minus the number of elements in ϑ.

Again, since $\hat{\vartheta}$ is a consistent but inefficient estimator, we can conduct another numerical optimization to derive the efficient second step estimator, $\tilde{\vartheta}$, of ϑ. The criterion function is now:

$$m_{q\varepsilon}'(\hat{V}_n)^{-1}m_{q\varepsilon} \qquad (5\text{-}3\text{-}32)$$

where:

$$\hat{V}_n = 1/N\ \Sigma\ \Delta q_n'\Delta\hat{\varepsilon}_n\Delta\hat{\varepsilon}_n'\Delta q_n \qquad (5\text{-}3\text{-}33)$$

where $\Delta\hat{\varepsilon}_n=\Delta\hat{\eta}_{ns}-\hat{\gamma}_2'\Delta z_{ns}$. Since we treat $\hat{V}_n$ as fixed, the gradient of (5-3-32) (i.e., value of objective function) during the iterative computation of the 2nd step estimator is given as:

$$\frac{\partial\ \text{vof}}{\partial\ \theta} = \begin{Bmatrix} 2\,\frac{\partial\ M_{q\varepsilon}}{\partial\beta}\,\hat{V}_n^{-1}\,M_{q\varepsilon} \\ 2\,\frac{\partial\ M_{q\varepsilon}}{\partial\alpha}\,\hat{V}_n^{-1}\,M_{q\varepsilon} \\ 2\,\frac{\partial\ M_{q\varepsilon}}{\partial\gamma_2}\,\hat{V}_n^{-1}\,M_{q\varepsilon} \end{Bmatrix} \qquad (5\text{-}3\text{-}34)$$

Covariance matrix of the 2nd-step estimator is:

$$\text{cov}(\tilde{\theta}) = (D\ \hat{V}_n^{-1}\ D)^{-1}/\ N \qquad (5\text{-}3\text{-}35)$$

where $D = 1/N \sum_n \frac{\partial\Delta\varepsilon_n'}{\partial\ \theta}\Big|_{\theta}\ \Delta q_n$. Corresponding Sargan statistic is:

$$C = (\Sigma\ \Delta\tilde{\varepsilon}_n'\Delta q_n)\ (\Sigma\ \Delta q_n'\Delta\hat{\varepsilon}_n\Delta\hat{\varepsilon}_n'\ \Delta q_n)^{-1}\ (\Sigma\ \Delta q_n'\ \Delta\tilde{\varepsilon}_n) \sim \chi\ (r) \qquad (5\text{-}3\text{-}36)$$

where $\Delta\tilde{\varepsilon}_n=\Delta\tilde{\eta}_{ns}-\tilde{\gamma}_2'\Delta z_{ns}$ and r is again the number of instruments minus the number of parameters.

Recovering of the character variables estimates can be done again by assuming u_n to have a conditional zero mean. Then regress $\tilde{e}_n=\tilde{\eta}_n - \tilde{\gamma}_2'Z_n$ on C_n', a vector of character variables whose first column is one. Let γ_0^* be the first element of γ_1, then the covariance matrix for their coefficient is derived as $\sum_n a_n a_n'$ where :

$$a_n=(\Sigma S_n C_n C_n')^{-1}\left[\Sigma S_n C_n\tilde{u}_n+\Sigma C_n(\sum_s^{S_n}\tilde{\varepsilon}_{ns})+\sum_n\left(C_{ns}\sum^{S_n}\frac{\partial\ e_n}{\partial\ \theta'}\Big|_{\theta=\tilde{\theta}}(D\hat{V}_n^{-1}D)^{-1}D\hat{V}_n^{-1}M_{q\tilde{\varepsilon}}\right)\right] \qquad (5\text{-}3\text{-}37)$$

and: $\tilde{u}_n = (\sum_s \iota_{ns})/S_n$, for $\iota_{ns} = \tilde{\eta}_{ns} - \tilde{\gamma}_2' z_{ns} - \tilde{\gamma}_1' C_n$

5.4. Estimation: random effects

The fixed effects IV estimation has the major advantage that it provides consistent estimates of α, β, γ_2 and ϕ_2 quite independently of any assumptions about the nature of the group effects, u_n. However, if we are prepared to assume that the u_n are stochastic and distributed independently of the instruments, q_{ns} for the duration equation and q^*_{ns}, for the wage equation, a more efficient IV estimation can be used.

Consider the n-th bargaining group. For this group, we can define the following residual vector:

$$\begin{aligned}\lambda_n(\theta)' &= [(\iota_{n1}-\gamma_0^*-\gamma_1'c_n-\gamma_2'z_{n1}) \ldots (\iota_{ns}-\gamma_0^*-\gamma_1'c_n-\gamma_2'z_{nS_n}),\\ &\qquad (w_{n1}-\phi_0-\phi_1 d_n-\phi_2'\xi_{n1}) \ldots (w_{ns}-\phi_0-\phi_1 d_n-\phi_2'\xi_{nS_n})]\\ &= [\, u_n+\varepsilon_{n1}, \ldots, u_n+\varepsilon_{nS_n}, \phi_3 u_n+\nu_{n1}, \ldots, \phi_3 u_n+\nu_{nS_n}] \qquad (5\text{-}4\text{-}1)\end{aligned}$$

where θ is now $(\alpha, \beta, \gamma_0^*, \gamma_1, \gamma_2, \phi_0, \phi_1, \phi_2)$. Conditional on the exogenous variables, $\lambda_n(\theta)$ has a zero mean and a covariance matrix:

$$\Omega_n = \left[\begin{array}{cccc|cccc} \sigma_\varepsilon^2+\sigma_u^2 & \ldots & \sigma_u^2 & \sigma_u^2 & \phi_3\sigma_u^2 & \ldots & & \phi_3\sigma_u^2 \\ \vdots & \ddots & & \vdots & \vdots & \ddots & & \vdots \\ \sigma_u^2 & \ldots & & \sigma_\varepsilon^2+\sigma_u^2 & \phi_3\sigma_u^2 & \ldots & & \phi_3\sigma_u^2 \\ \hline & & & & \sigma_v^2+\phi_3\sigma_u^2 & \ldots & & \phi_3\sigma_u^2 \\ & & & & \vdots & \ddots & & \vdots \\ & & & & \phi_3\sigma_u^2 & \ldots & & \sigma_v^2+\phi_3\sigma_u^2 \end{array}\right]$$

$$= A \otimes I_{(S_n)} + bb' \otimes ee' \qquad (5\text{-}4\text{-}2)$$

where e is the $S_n \times 1$ unit vector and:

$$A = \begin{bmatrix} \sigma_\varepsilon^2 & 0 \\ 0 & \sigma_v^2 \end{bmatrix} \qquad (5\text{-}4\text{-}3)$$

$$b' = [\, \sigma_u \quad \sigma_u \phi_3] \qquad (5\text{-}4\text{-}4)$$

The unknown components of (5-4-3) and (5-4-4) can be estimated from the fixed-effect residuals, which can then be used to construct an estimated covariance matrix, $\hat{\Omega}_n$. The optimal IV estimator is the value of θ that minimizes:

$$\left[\sum_{n=1}^{N} \lambda_n(\theta)'\hat{\Omega}_n^{-1} Q_n\right]\left[\sum_{n=1}^{N} Q_n'\hat{\Omega}_n^{-1}Q_n\right]^{-1}\left[\sum_{n=1}^{N} Q_n'\hat{\Omega}_n^{-1}\lambda_n(\theta)\right], \qquad (5\text{-}4\text{-}5)$$

where $Q_n' = [q_{n1} \ldots q_{nS_n}, \ldots q^*_{n1} \ldots q^*_{nS_n}]$. The inverse of $\hat{\Omega}_n$ can be written as:

$$\hat{\Omega}_n^{-1} = A^{-1} \otimes I - \hat{B} \otimes ee', \qquad (5\text{-}4\text{-}6)$$

where:

$$\hat{B} = \frac{\hat{\sigma}_u^2 \hat{\sigma}_\varepsilon^2 \hat{\sigma}_v^2}{\hat{\sigma}_\varepsilon^2 \hat{\sigma}_v^2 + S_n \hat{\sigma}_u^2(\hat{\sigma}_v^2 + \hat{\phi}_3^2\hat{\sigma}_\varepsilon^2)} \left[\begin{array}{c|c} \dfrac{1}{\hat{\sigma}_\varepsilon^4} & \dfrac{\hat{\phi}_3}{\hat{\sigma}_\varepsilon^2 \hat{\sigma}_v^2} \\ \hline \dfrac{\hat{\phi}_3}{\hat{\sigma}_\varepsilon^2 \hat{\sigma}_v^2} & \dfrac{\hat{\phi}_3^2}{\hat{\sigma}_\varepsilon^4} \end{array}\right], \qquad (5\text{-}4\text{-}7)$$

Thus the product sum in the objective function (5-4-5) is:

$$\sum_n Q_n' \hat{\Omega}_n^{-1} Q_n = \hat{A}^{-1} \otimes \sum_{s=1}^{S_n} q_{ns}q_{ns}' - S_n^2 \hat{B} \otimes \bar{q}_n\bar{q}_n' \qquad (5\text{-}4\text{-}8)$$

and:

$$\sum_n Q_n'\hat{\Omega}_n^{-1}\lambda_n(\theta) = \left[\begin{array}{c} \dfrac{1}{\hat{\sigma}_\varepsilon^2}\sum_{s=1}^{S_n} q_{ns}(\iota_{ns} - \gamma^*_0 - \gamma_1'c_n - \gamma_2'z_{ns}) \\ \dfrac{1}{\hat{\sigma}_v^2}\sum_{s=1}^{S_n} q_{ns}(w_{ns}\phi^*_0 - \phi_1 d_n - \phi_2'\xi_{ns}) \end{array}\right]$$

$$- S_n^2 \left[\hat{B}\left[\begin{array}{c} \bar{\iota}_n - \gamma^*_0 - \gamma_1'c_n - \gamma_2'\bar{z}_n \\ \bar{w}_n - \phi^*_0 - \phi_1'd_n - \phi_2'\bar{\xi}_n \end{array}\right]\right] \otimes \bar{q}_{n.} \qquad (5\text{-}4\text{-}9)$$

Under standard conditions, this nonlinear 3SLS estimator is consistent and has a limiting normal distribution with the approximate covariance matrix:

$$V_{\theta} = [(\sum_n \hat{G}_n \hat{\Omega}_n^{-1} Q_n)(\sum_n Q_n' \hat{\Omega}_n^{-1} Q_n)^{-1} (\sum_n Q_n' \hat{\Omega}_n^{-1} \hat{G}_n')], \tag{5-4-10}$$

where:

$$\hat{G}_n = \frac{\partial \lambda_n(\hat{\theta})'}{\partial \hat{\theta}} \tag{5-4-11}$$

5.5 Results

Based on the economic framework discussed in chapter 4, we present some results for one class of observations: wage agreements affecting unskilled males in the manufacturing and construction industries. Narrowing the scope of our study to these industries was inevitable due to restricted availability of the industry specific data availability. We also lost some observations through the construction of lagged values as instruments. Consequently, our sample consists of 61 bargaining groups with 850 wage agreements altogether.

Following are the definitions of the variables used in the equations to be estimated. W(t) denotes the negotiated wage at calendar time t, R(t) is the retention ratio and P(t) is the retail price index. The s-th spell started at T_{s-1} and ended at T_s in calendar time, where $T_s - T_{s-1} = \delta_s$. The continuously varying explanatory variables are typically available only on quarterly or monthly basis. Their spell specific values corresponding to each settlement date are interpolated appropriately. During the spell, they are assumed to take the forms of step functions with monthly increments.

With respect to the incomes policy, we use simple on/off dummies for each episode. Although we cannot derive any implications on the impact of each degree of norm and enforcement, it enables us to find out which policies were successful in delaying or moderating the pay rise. We have consistently failed to detect any influence of the pre-1961 or post 1970 incomes policies (only three settlements are observed in our sample during the 1972/73 policy), and they are consequently excluded form the analysis. The list of dummy variables we have used to generate the results are listed below.

Variables with argument, t, vary with time during the spell, while the others are either spell or group specific.

Real pay variables

lnrw(t) = $\ln(W(T_{s-1})R(t) \ / \ P(t))$

Drr(t) = $R(t)P(T_{s-1}) \ / \ R(T_{s-1})P(t) - 1$

$Drr^e(t) = E[R(t)/P(t)|\Omega]\ P(T_{s-1})/R(T_{s-1}) - 1$

$Drr^u(t) = E[P(t)/R(t)|\Omega]\ R(t)/P(t) - 1$

where Ω is the appropriate information set

Drw1 $= W(T_{s-1})R(T_{s-1})P(T_{s-2})\ /\ W(T_{s-2})R(T_{s-2})P(T_{s-1}) - 1$

Dw $= W(T_s)/W(T_{s-1}) - 1$: dependent variable in the wage equation

Capacity to pay variables

lnprin(t) = log of the industry's gross profit at time t;

prin(t) = rate of change of industry's gross profit at time t since T_{s-1};

prin11(t) = rate of change of industry's gross profit at time t since 11 months ago;

prod(t) = rate of change of global productivity since T_{s-1};

whip(t) = rate of change of whole sale input prices since T_{s-1};

whop(t) = rate of change of whole sale output prices since T_{s-1};

Outsider influences

rel(t) = the ratio of average UK earnings at time t to the group's current wage;

lnun(t) = log of the aggregate unemployment in 1000's (excluding school leavers);

lnunin(t) = log of the industry specific unemployment in 1000's;

Other variables

size(t) = changes in the number of workers covered by the bargaining group as a proportion of total UK employment at time t since T_{s-1}.

$staged_s$ = 1 if the s-th wage increase (made at T_s) is part of a staged settlement, = 0 otherwise;

$sfix_s$ = 1 if the s-th wage increase is fixed in advance as a part of a staged settlement, = 0 otherwise;

$dfix_s$ = 1 if the date of s-th wage increase is fixed in advance as a part of a staged settlement, = 0 otherwise;

WCouncil = 1 if the group belongs to the wages councils sector, = 0 otherwise;

Public = 1 if the groups belongs to the public sector, = 0 otherwise;

TU = number of Trade Unions involved in the negotiation group.

Incomes policy

Wage freeze

fd1(t) = 1 for non-private sector groups at time $t \in$(Jul 61-Mar 62), = 0 otherwise;

fd2(t) = 1 for $t \in$(Jul 66-Jun 67), = 0 otherwise;

Wage ceilings

cd1(t) = 1 for non-private sector groups at time $t \in$(Apr 62-Mar 63),

= 0 otherwise;

cd2(t) = 1 for non-private sector groups at time t∈(Apr 63-Apr 65),

= 0 otherwise;

cd3(t) = 1 for non-private sector groups at time t∈(Apr 65-Jul 66),

= 0 otherwise;

cd4(t) = 1 for t∈(Jul 67-Mar 68), = 0 otherwise;

cd5(t) = 1 for t∈(Mar 68-Dec 69), = 0 otherwise;

cd6(t) = 1 for t∈(Jan 70-Jun 70), = 0 otherwise;

Twelve months policy

d12(t) = 1 for t∈(Jul 67-Dec 70), = 0 otherwise.

5.5.1 Duration equation

Equation (5-3-1) is estimated by the iterative generalized method of moments, where the iteration is mostly done by DFP. The programming language used is GAUSS, in particular, a routine called OPTMUM. This routine increased its speed immensely when we provided analytical gradients. Nonetheless, in general, we have found a convergence of such a complicated nonlinear objective function with numerous instruments very hard to achieve. In particular, attainment of the global minimum seems to depend, to a certain extent, on starting values and the scaling of the explanatory variables. The best way, it seems, is to start the estimation only with a handful of covariates then gradually increase the number of explanatory variables.

In interpreting the estimates, note that a positive coefficient implies that an increase in the corresponding variable *reduces* the duration of the spell on average. This interpretation of the coefficient on the average ex-post duration is only strictly valid for the spell specific variables such as lagged durations, DRW1, Dfix and group idiosyncrasies. With respect to the time varying covariates, their immediate impact is exerted on the hazard, and their coefficients represent the comparative statics impact on the log of duration.

(i) Duration dependence (1)

As we have discussed in the section 5.2.3, it is a very tricky task to find a general parametric form for the baseline hazard that will not lead to an identification problem. Look at the expression of our dependent variable, η_{ns}:

$$\eta_{ns} = -\ln\left[\int_0^{\delta_{ns}} \exp\{\beta' x_n(t)\}\, h_0(t;\alpha)\, dt\right] \tag{5-5-1}$$

In the simplest of all cases where there is no time-varying covariates, x, η_{ns} reduces to $-\ln\int_0^{\delta_{ns}} h_0(u)du$. There is a well known identification problem with the baseline hazard distribution assumption such as Weibull ($h_0(t)=t^{\alpha}$), where our dependent variable becomes $\eta_{ns}=-\alpha\ln(\delta_{ns})$; the parameters are only estimable up to a scale factor. Our regression type model, in general, is prone to an identification problem of the slope parameter in the integrand, since the slope cannot be identified as long as the area under the curve (i.e., the integrated values) remains unchanged. Provided the appropriate normalization and/or sufficient nonlinearity to the functional form of $h_0(t;\alpha)$ should solve this problem. It is not certain, however, whether that is also the case empirically.

Moreover, even if the functional form is appropriate so that the duration dependence can be identified, the collinearity amongst the time path of x(t) and the elapsed duration during the spell will bring about another source of identification problem. Theoretically speaking, this issue is not a problem if there are enough within spell variations in x(t) across the observations. Nonetheless, in practice, variables such as inflation, productivity and profit move very closely with the time trend.

The first issue precludes a form of the integrated hazard, $\Psi(t;\alpha)$, that has a value of parameter α which makes $\Psi(t;\alpha)$ constant for any elapsed duration, t. If $\Psi(t;\alpha^*)=\Psi^*$, which is independent of t, then, as α approaches α^*, the dependent variable will be reduced to a simple linear form:

$$\eta_{ns} = - x(\tau^0_{ns})\beta - \ln \Psi^* \qquad (5\text{-}5\text{-}2)$$

making β unidentified. According to this restriction, following approximations for $\Psi(t;\alpha)$ also lead to the unidentification problem: t^{α} (Weibull), $\exp(\alpha t)-1$, or $\ln(1+(\lambda t)^{\alpha})$ (log logistic). Moreover, there are other specifications that do not fall into the restriction above but still result in the similar identification problem. Such examples are: $h_0(t)=\exp(\alpha t)$ or $h_0(t)=1+\alpha\lambda(t)$ where $\lambda(t)$ is a dummy for a certain interval of t. In these specifications, the coefficient estimates representing the duration dependence tend to converge to values that are independent of the duration observations.

Consider again the simplest case with no explanatory variables, and simply minimize the residual sum of squares (therefore ignoring the group random effect). Based on the hazard, $h(t)=h_0(t)e^{\gamma}$, our problem reduces to minimizing:

$$\sum_n \sum_s \varepsilon_{ns}^2 = \sum_n \sum_s \{ -\ln [\int_0^{\delta_{ns}} h_0(u)\, du] - \gamma \}^2$$
$$= \sum_n \sum_s \{ -\ln \Psi(\delta_{ns},\alpha) - \gamma \}^2 \qquad (5\text{-}5\text{-}3)$$

For example, if $h_0(t)=\exp(\alpha t)$, α has to be non-negative for the implied duration distribution to generate a proper distribution, so that $\Psi(\infty)e^{\gamma}=\infty$. Still, if α is allowed to be negative, the minimum is achieved at a sufficiently negative value of α such that $\exp(\alpha\delta_{ns})=0$ for any δ_{ns}, therefore making our dependent variable constant. On the other hand, if α is constrained to be non-negative, there is no longer a value of α that leads to a constancy of the dependent variable. Nevertheless, the convergence is reached at $\alpha=0$, implying the constant hazard. In this particular case, the corresponding integrated baseline hazard is, $\Psi(\delta_{ns})=(\exp(\alpha\delta_{ns})-1)/\alpha$, which is an increasing function of α for $\alpha\geq 0$. Nonlinear minimization of the difference between the log of $\Psi(\delta_{ns};\alpha)$ and a constant term has resulted in a corner solution at $\alpha=0$. In the case of $h_0(t)=(\exp(\alpha t)+1)$, non-negativity constraint on α is not necessary for the admissibility condition. The minimum is again derived at a sufficiently negative value of α that makes $\exp(\alpha\delta_{ns})\cong 0$, reducing our dependent variable to be $-\ln(\delta_{ns}-1/\alpha)$. Even for the specification with a normalization such as $h_0(t)=1+\alpha\lambda(t)$ or $h_0(t)=t+\alpha\lambda(t)$, where $\lambda(t)=1$ for a certain range of t, the optimum is again at $\alpha=0$ ($\alpha\geq 0$ constrained). Also, for $\Psi(t)=\exp(t+\alpha t^2)-1$, which directly approximates the integrated baseline hazard, the minimum is reached at 0, again failing to capture any form of duration dependence.

These simple models can be easily extended to include explanatory variables that are both spell specific and time-varying. The hazard is now written as: $h(t)=h_0(t)\exp(\beta' x(t)+\gamma' z_{ns})$. The inclusion of such variables, however, is irrelevant to the property of the resulting estimates of α. For example, table 5.1 lists the coefficient estimates of the hazard equation with a baseline hazard that has a discrete annual jump:

$$h_0(t;\alpha) \cong \begin{cases} 1 & t<48 \text{(weeks)} \\ 1+\alpha & 48 \leq t \leq 72 \\ 1 & t>72 \end{cases} \qquad (5\text{-}5\text{-}4)$$

which, in terms of the integrated baseline hazard, is:

$$\Psi(t;\alpha) \cong \begin{cases} t & t<48 \\ t+\alpha(t-48) & 48\leq t \leq 72 \\ t+24\alpha & t>72 \end{cases} \qquad (5\text{-}5\text{-}5)$$

where $\alpha>0$ is constrained by setting $\alpha=\exp(b)$ and the value of b is iterated during the estimation. If there exists a sudden increase in the hazard at around the 52nd week that is not explained by any other variables considered in the equation, we should find a significantly positive value for $\hat{\alpha}$. On the contrary, as can be seen, the estimates for b is largely negative, hence, $\hat{\alpha}\cong 0$. Table 5.2, on the other hand, lists the coefficient estimates when we approximate the continuous duration dependence to be discrete. We assume that the elapsed duration affects the hazard discretely at every month just as the other time varying covariates do, instead of letting them affect the hazard continuously as they have been so far. Since this eliminates the difference term in the integrated baseline hazard, α is allowed to become negative, although this means the admissibility condition may not be satisfied for some time path of x(t). We can write the baseline hazard as having its argument, $(\tau^i_{ns}-\tau^0_{ns})$; an approximation for δ_{ns}. Our dependent variable now has an expression:

$$\eta_{ns} = -\ln \sum_{i=0}^{k_{ns}} \left\{ e^{x(\tau^i_{ns})\beta} \; h_0(\tau^i_{ns}-\tau^0_{ns}) \; [b^i_{ns}-a^i_{ns}] \right\} \qquad (5\text{-}5\text{-}6)$$

$$\text{where} \qquad b^i_{ns} = \max \{ T_{ns-1}, \tau^i_{ns} \} - T_{ns-1}$$

$$a^i_{ns} = \min \{ \tau^{i+1}_{ns}, T_{ns} \} - T_{ns-1} \qquad \text{as before.}$$

In particular, consider the simplest monotonic duration dependence such as:

$$h_0(\tau^i_{ns}-\tau^0_{ns}) = \exp[\alpha \; (\tau^i_{ns}-\tau^0_{ns})] \qquad (5\text{-}5\text{-}7)$$

In our data, $\hat{\alpha}$ converged to a negative value exhibiting an acute negative duration dependence, which is pulling the hazard down to zero soon after the commencement of a spell (see the estimated average hazard[3] in fig 5.2). Recall the simple case with no explanatory variables where $h=\exp(\alpha\delta_{ns})\gamma$. The convergence was achieved for a sufficiently negative α, because it made the dependent variable, η_{ns}, constant for any δ_{ns}. Likewise, in this case, a sufficiently negative value of α deprives x(t) of any impact on the hazard soon after the start of a spell. In fact, large and negative α can reduce our nonlinear optimization problem to the simple linear regression with a dependent variable:

$$\eta_{ns} \cong -\beta' x(\tau^0_{ns}) - \ln(\tau^1_{ns} - T_{n,s-1}) \qquad (5\text{-}5\text{-}8)$$

where the parameters are only estimable up to a scale factor. As expected, estimated coefficients, particularly those of incomes policies are markedly different from those of table 5.2. For example, neither of the freeze policies exerts any significant impact, which is totally unreasonable considering the marked reductions in the occurrence of negotiation found in a sample during these policies. In spite of this, when we include a discrete annual jump in addition to the model (5-5-6), so that:

$$h_0(\tau^i_{ns}-\tau^0_{ns}) = \exp\left[\alpha_1 (\tau^i_{ns}-\tau^0_{ns}) + \alpha_2\, d(\tau^i_{ns})\right] \qquad (5\text{-}5\text{-}9)$$

where $d(\tau^i_{ns}) = \begin{cases} 1 & \text{for } 11 \le i \le 13 \\ 0 & \text{otherwise} \end{cases}$

Surprisingly, α_2 is found significant (table 5.3). It may be that due to numerical insensitivity, the noise, albeit small, around the 52nd week in the residuals is taken up by this additional variable[4]. Consequently, the average estimated hazard has a jump at around 1 year (fig 5.3) and its decline is still implausibly sharp. This acute decline in the hazard is not due to the spell specific terms. To see this, we can examine the movement of the estimated hazard over time conditional on the values of spell specific factors and the timing of starting point, by choosing an arbitrary observation point and plot $\hat{h}(t)$ over t. Fig 5.4 is such a plot for a private group and fig 5.5 is for a wages council group. They both represent the hazard that declines very fast over time. Obviously, the negative value of $\hat{\alpha}$ is exerting the dominant impact in pulling the hazard down to zero irrespective of the time path of x(t).

What happens, then, if we just keep $h_0(t)=1$ and assume that the hazard is only affected by the elapsed duration via x(t)? This is a reasonable assumption provided that the variables considered in x(t) sufficiently captures the accumulating pressure depicted by the theory in chapter 4. But then, the same identification problem with the duration dependence is replicated as long as there is at least one covariate which is collinear with the elapsed duration during the spell across observations. In table 5.4, coefficients on Drr^{e}[5] and the log of real wages are behaving in a way α did in table 5.2 or 5.3, exhibiting the wrong positive signs and pulling the hazard down to zero. Decline in the level of real take home pay should, *ceteris paribus*, increase the chance of new negotiation which implies a negative coefficient for both variables. However, our estimates show a strong tendency that the lower the expected real take home pay, the more unlikely the contract will terminate, which is hard to comprehend. These puzzling coefficients are reflected in the form of declining average estimated hazard, as can be seen in figure 5.6. Also, there is no sign of an annual jump in the residuals (fig 5.7), nor is such jump significant when actually allowed for in the equation. Again, the movement of the estimated hazard over time, for the same groups as those in fig 5.4 and 5.5 is shown to be declining implausibly fast (fig 5.8, 5.9: other groups showed similar figures).

Despite their possible shortcomings, these results so far do emphasize one important point: that the timing of wage bargains in this period was endogenous, and strongly influenced by the economic developments. For example, the incidence of 1966-67 freeze policy (fd2) increases the average duration by 84.5% and 1% point decline in relative wage variable, Rel, reduces the contract length by 2.08% on average.

(i) Duration dependence (2)[6]

We have found strong influences of the time varying covariates on the hazard despite the tendencies for the coefficient estimates of few monotonic variables to converge to values unrelated to the duration observations. As can be seen in table 5.2-5.4, such monotonic variables pull the hazard down to zero so fast that it is, in effect, nullifying any contribution of other x variables on the hazard, soon after the start of a spell. This is due to the multiplicative nature of the way x's affect the hazard. Consider, instead, specifying the hazard as:

$$h_{ns}(t) = \exp\{\gamma_0+\gamma_1'c_n+\gamma_2'z_{ns}+u_n\} \prod_{j=1}^{nb}\{\exp(\beta_j'x_j(t)) + 1\} \qquad (5\text{-}5\text{-}10)$$

where nb is the number of time-varying covariates. This allows variables to continue exerting impact on the hazard even when $\beta_j x_j \rightarrow -\infty$ for any j-th explanatory variable. In fact, when $\beta_j x_j \rightarrow -\infty$ for all j, h(t) reduces to a constant hazard. Corresponding expression for the dependent variable is:

$$\eta_{ns} = -\ln \{ \sum_{i=0}^{k_{ns}} (a^i_{ns} - b^i_{ns}) \prod_{j=1}^{nb}\{\exp(\beta_j'x_j(t)) + 1\} \qquad (5\text{-}5\text{-}11)$$

This is an arbitrary specification in as much the same way as the proportionate hazard specification. The coefficients, however, no longer represent elasticities, therefore are more difficult to interpret. Table 5.5 lists the first step estimates based on above specification. Negative coefficient still represents an inverse relation with the hazard, although the coefficient that makes $\beta_j x_j$ largely negative now simply provides quantitatively trifle effect on the hazard. Hence, a large negative coefficient estimates seen in table 5.5 for a1 suggests an unimportant and insignificant pure elapsed duration effect on the hazard. The first three incomes policies, aimed only at wages council and public sector, have been also quantitatively unimportant and statistically insignificant (because of their huge standard errors, it was not possible to proceed with the second round estimator). Inclusion of these variables has led to a large Sargan statistics. Other variables have shown signs that are consistent with the theory. The overall movement of the hazard implied by these estimates are in figure 5.10. It is still decreasing, at least up to 2 years of duration observation, though its slope is no longer sharp enough for the admissibility condition to fail (see fig 5.11). Since there is no dominant time effect that is significantly affecting the way the hazard moves over time, it is all the more misleading to gauge a duration dependence solely from this figure. Recall that it depicts the average estimated hazard computed at one point in time for each observation, namely, at the completed duration. Instead, look at figures 5.12-5.13 which plot estimated hazard over a 50-month period for an arbitrarily chosen observation, one from a private sector and another from the wages council. The level of the estimated hazard varies for each observation partly due to differences in their initial condition, and their movement is largely affected by the time path of x(t). For a full analysis, we need a dynamic simulation to measure the impact of each x, in particular, of the incomes policies.

Still within the framework of the original parametric specification, we have considered estimating the hazard excluding variables such as Drr^e and log of real wages[7] that has been causing the implausible decline in the hazard. In other words, we exclude any time varying factors whose coefficient estimates had the tendencies to converge to the corner solutions that are independent of the duration observations. Look at table 5.6. Here, all the signs of estimated coefficients are consistent with the theory, including profit and relative wage variables which tend to vary monotonically during the spell. The incidence of fd1 reduces the hazard amongst non-private groups by 70.8% which is as much as 123% increase in the duration on average. Second freeze increases the average duration by 73% in all groups. Most of the incomes policies are significant with around 43 to 75% impact on the hazard except for cd4 which is hardly effective in altering the timing of wage changes. Static elasticity of duration with respect to industry unemployment is 0.475 in a single spell and $0.457/(1+0.09)=0.436$ in the long run. Similarly, elasticity with respect to industry specific profit is -0.48. An interesting finding is the importance of public sector dummy which has a tendency to raise the hazard and therefore shortens the contract length on average (out of 850 observations on durations used in the analysis, 18 are of public sectors and its mean duration is 47.33 weeks with 23.22 standard deviation compared to the mean of 56.01 and the standard deviation of 24.68 of private and wages councils combined). The average estimated hazard is found to be non-monotonic with a jump at around 2 years of the completed duration. It still exhibits a decline for the first 2 years, as well as high residuals for the short duration (fig 5.14-5.16)[8]. Note that the relatively large mean residuals at the extremes of the duration range are based on very small numbers of observed spells, and that the negative correlation between the residuals and the observed duration is an inherent feature of this type of plot, even with a correctly specified model. Plots of arbitrary wages council and private sector groups are in figure 5.17 and 5.18. As can be seen, their duration dependence through x(t) again varies largely with x, and their dynamic impact have to be examined by dynamic simulation.

Look at figure 5.19 which plots average estimated hazard based on the maximum likelihood estimates where all the explanatory variables except for the elapsed spell are those known at the start of a spell (results listed in table 7.4). Our data suggest an increasing hazard given the assumption of the Weibull baseline hazard and no heterogeneity. Their estimated average pdf at completed duration observations suggests a plausible peak at around 6-15 months (fig 5.20). One would ideally like to argue that the positive duration dependence observed from the ML estimates is misspecified since once we take into account of the continuous varying economic environment, the underlying pure duration dependence is found to be actually negative. Nonetheless, the plot of the estimated hazard averaged at each completed duration observation is still declining at least for the first 2 years. Moreover, it is hard to imagine the introduction of a group and individual heterogeneity and their possible correlations with the other x's alone have caused the estimated hazard to decline. But can we draw inferences on the duration dependence from the shape of estimated hazard, or to conclude the validity of the hazard specification from the estimated pdf, both averaged at the completed duration observations? The answer to this has to resort to dynamic simulation which will be discussed in chapter 6.

But before concluding this section, let us consider the implication of the declining average estimated hazard on the duration distribution that is required in the dynamic simulation. In particular, the extent to which the omission of the pure duration dependence is responsible.

When only the static explanatory variables are incorporated in the hazard, the parameter estimates derived from the regression format are only estimable up to a scale factor, hence, they are the re-scaled values of the true structural parameters. To see this, consider when x(t) are constant at the start of a spell so that it is absorbed in z_{ns}, thus, $h(t)=\exp(\gamma' z_{ns}+u_n)h_0$. If the true duration distribution is monotonic, such as Weibull, the corresponding hazard can be written as:

$$h(t) = \exp(\gamma' z_{ns}+u_n)\alpha t^{\alpha-1} \qquad (5\text{-}5\text{-}12)$$

And the survivor function required for the dynamic simulation is:

$$S(t) = \exp(-\exp(\gamma' z_{ns}+u_n)t^{\alpha}).$$

The regression model derived on the basis of this hazard formulation is: $-\ln\delta_{ns}=\gamma^{*\prime} z_{ns}+u^*_{ns}$, where the coefficient estimates derived are the rescaled values of their true parameters, namely, $\gamma^*=\gamma/\alpha$ and $u^*_n=u_n/\alpha$. Substituting back the rescaled estimates into the estimated hazard gives: $h^*(t)=\exp\ \gamma^* z_{ns}+u^*_{ns}$. The true hazard h(t) can then be written in terms of h*(t) as:

$$h(t) = \left\{ h^*(t)/(\alpha\ t^{\alpha-1}) \right\}^{\alpha} \qquad (5\text{-}5\text{-}13)$$

And the true survivor function, S(t) is: $S(t)=(S^*(t))^{\alpha}$, (5-5-14), where $S^*(t)=\exp\{-\exp(\gamma' z_{ns}+u_n)t^{\alpha})^{1/\alpha}\}$. But S(t) and S*(t) are not the same. Even if our estimated hazard declines over time, the true underlying hazard can be increasing for a certain range of α. In this static case, one can identify the rescaling factor, α, so that the average estimated distribution implied by α best represents the actual sample distribution (for example, choose α so that the implied value of prob($5<t<20$months) is closest to its sample distribution).

Extending this argument to the case where x(t) varies during the spell, however, is not so straightforward. There is no reason why the time varying x(t) which has sufficiently non-collinear variation over t should not be identified. And indeed, we have seen their strong influences that are in accord with the economic theory in the previous analysis. In the presence of such x(t) that are not constant during the spell, a simple relation shown in equation (5-5-14) is no longer appropriate to gauge the true duration distribution to be used in the dynamic simulation.

Looking more closely into the issue, the timing of wage change is simulated by a conditional probability to exit within a monthly interval, which in terms of the survivor function, is: (S(t-1)-S(t))/S(t-1). Since x(t) are observed monthly, this probability can be written as:

$$1 - \exp\left\{ - \exp(\gamma' z_{ns} + \beta' x(\tau_{ns}[k_{ns}])) \int_{\tau_{ns}[k_{ns}]-T_{ns-1}}^{\delta_{ns}} h_0(u)\, du \right\}$$

where k_{ns} is the number of maximum monthly x's observed during this n,s-th spell, which started at T_{ns-1} and lasted for δ_{ns}. If the true underlying baseline duration distribution follows the Weibull, the above conditional probability is approximately:

$$1 - \exp\left\{ - [\ \exp(\gamma/\alpha' z_{ns} + \beta/\alpha' x(\tau_{ns}[k_{ns}]))4.3\ ]^{\alpha}\ ((k_{ns}+1)^{\alpha} - k_{ns}^{\alpha}) \right\}$$

where a month is approximated as 4.3 weeks, since δ_{ns} is measured in weeks. However, this is different from the same conditional probability based on h*(t):

$$1 - \exp\left\{ - [\ \exp(\gamma^{*\prime} z_{ns} + \beta^{*\prime} x(\tau_{ns}[k_{ns}]))4.3\ ] \right\}$$

even if the estimated comparative static effects of the explanatory variables are the same between them, that is, if $\gamma^*=\gamma/\alpha$ and $\beta^*=\beta/\alpha$. Look at table 5.7 that lists coefficient estimates when the baseline hazard is restricted to follow the Weibull with parameter $\alpha=2$. This value is chosen since the maximum likelihood estimates with the Weibull duration distribution with no time varying x has a value of $\hat{\alpha}=2.4$. Implied elasticities of duration (i.e., β/α and γ/α) are very similar except for the relative wage. Even though imposing $\alpha=2$ restricts the hazard to increase with time, as can be seen in the upward sloping average estimated hazard in fig 5.21, that of the survivor function (i.e., the area under the hazard) is hardly different (fig 5.22). Since the exit probability that determines the timing of wage changes in the dynamic simulation is a function of monthly changes in the survivor function, these two specifications, even though the underlying pure duration dependence is different, may turn out to produce similar results in the simulated wage levels.

We have found it difficult to identify the elapsed duration, as well as any other variables that are collinear with the passage of time during each observation spell. Fortunately, relative wage and industry specific profit series, even though their behaviour during the spell is almost monotonic, seem to have enough variations across observations for them to be identified.

Nonetheless, the question remains: if these x(t)'s we have taken into account in table 5.6 adequately represent the accumulated environmental pressure, or additional imposition of $\alpha=2$ is yet better. Dynamic simulation is the only way through which we can measure the appropriateness of our specification and compare them with the estimates derived from the static MLE that assumes the Weibull duration distribution.

5.5.2 The wage equation

The coefficient estimates of the conditional wage change equation are presented in Table 5.8. The majority of significant coefficients have the expected signs. Possible exceptions are Drw1, the rate of change of real take home pay over the previous contracts, and the log of global unemployment, although neither of them are precisely estimated. The positive DRW1 implies that a high wage settlement in real terms, has a slight but weekly significant tendency to be followed by further high settlements, which is an unexpected finding. One would expect that if a contract starts off with a high real pay that more than compensates the inflation rate over the previous contract, *ceteris paribus*, then workers will not wish nor will be in a position to claim another high wage. Interestingly, in addition to this finding, we observe insignificant current and lagged durations.

Recall the economic framework discussed in chapter 4. We have assumed that workers will step forward for a new negotiation when, for the first time since the last negotiation, the real wage becomes less than the target wage net of negotiation cost. Consider the role of elapsed durations. The timing of wage change will be affected by the elapsed duration if, for example, there exists a certain "norm" in the economy to have one negotiation per year. Then, deviation from such a "norm" becomes a costly behaviour that affects the renegotiation probability. Given the changes in the other explanatory variables over the spell, once the bargaining occurs, δ_{ns} does not affect the wage claimed at the end of such contract as long as δ_{ns} does not help to predict future uncertainty or negotiation cost. We consider workers as having a certain target level of potential remuneration that is a function of all the economic variables, such as $w^*_n(t)$, c(t) and p(t). They will renegotiate whenever the potential remuneration level reaches the target level (i.e., $w^*_n(t)-c(t)-w_{ns-1}p(t)\geq 0$). The actual wage level, w_{ns}, will then be determined by additionally taking into account of the future uncertainty. Since this target level is fixed a priori (to 0), w_{ns} is not affected by the contract spell, δ_{ns}, unless it affect the component of the target level. This interpretation may also be valid for DRW1. Even though DRW1 indicates the existence of any uncompensated inflation carried over to the current contract, its measure also embodies the length of the last contract, δ_{ns-1}, together with the movement of RPI and tax rates during the last contract. In this way, it can act to signal the latest trend in the negotiation cost and future uncertainty, hence, will help predict the future contract length. If this impact dominates the effect as a signal of uncompensated inflation, it should have a positive sign. And this is consistent with our observation that groups who succeed in obtaining large wage changes tend to do the same at the next pay round. Furthermore, in our data, we find the current and lagged durations insignificant when estimated with and without the DRW1. Our Conclusion, hence, is that DRW1 and neither δ_{ns} nor δ_{ns-1} helps predict the negotiation cost and the future uncertainty. High DRW1 suggests a high negotiation cost predicted at the last bargaining, which is likely to be inherited in the future cost. This makes workers to bargain harder for the current wages.

A possibility of simultaneity of δ_{ns} in the wage equation is not a problem since the wage equation is conditioned on the timing of wage changes.

Contrary to the hazard equation, the group's wage relative to the average level of earnings in the economy as a whole has a major influence on the level of wage claim in terms of the statistical significance in addition to the real retained wage, the number of workers negotiating, the change of global productivity, and of course, the incomes policies. Wage rises achieved on the basis of relative wage were the commonplace especially in the public sector. It is said that once a public sector group succeeded in deriving a favorable settlement by claiming their wage level were too low compared to the private sector counterparts, such move was quickly followed by other public sector groups. We have seen the public sector catching up with the private sector in this way mostly between 1969-70. A one percentage point increase of the global productivity increases $\dot{w}_{ns}$ by 0.4 percentage points. On the other hand, industry's trading profit exerts slightly negative influence. In view of the fact that only sustained profits affect wage claims (Carruth and Oswald (1989)), we have included the log of average profits over the last two consecutive negotiations, but it was insignificantly positive. The global unemployment has a positive effect while that of the industry specific is negative, and they are both only weakly significant. This is contrary to the prediction of the Hysteresis and the insider-outsider explanations. We have also included the lagged industry unemployment to incorporate the membership consideration (i.e., how quickly unemployed loses insider status), hence allowing for a possibility of increased industry unemployment to be associated with higher wage demands in the future. It showed an insignificant negative effect. Also, when the insignificant factors are taken away from the wage equation, as shown in table 5.9, unemployment variables become totally insignificant. The number of workers negotiating at a time varies from group to group, and is in the form of deviation since the start of the last contract. It is also divided by the total UK employment in order to make this variable net of aggregate fluctuations. Its significantly positive coefficient suggests the importance of membership that enables the group to achieve higher wages by raising their bargaining power.

Our result also shows that wage changes, when brought about as a part of staged settlement, are reduced by around 1.5 percentage points, although they are insignificant. Only about 38.5% of the staged wage changes are recorded to have their size planned in our sample, most of which were tied to the regulations that aim to achieve equal pay between male and female/juvenile workers. Considering how badly those juveniles were paid to start with, the effect of the stages implemented by such regulation could well have been positive. In fact, the mean rate of wage changes amongst the staged implementations is hardly different from that of the non-staged (0.066 versus 0.067). Significant group-specific influences are confined to the wage council sector, where a wage rise tend to be lower than that of other comparable groups by as much as 5 percentage points. This observation, together with the evidence from the hazard equation of table 6.4, indicates that the wages council groups derive lower average rate of wage changes per unit of time than the other groups. This is not an unexpected finding, since the wages council sector mainly covered a large group of low-paid workers in the private sector. Not only was it hard for them to breach the limits imposed by the policies but they also failed to derive some of the full benefit specified by the policies.

The incomes policy effects appear very strong for all the policies except for fd1. The estimated impacts of the late 1960s policies (cd5) is as much as 7 percentage points. It seems that not only the ceiling policy but also the freeze policy lowers the magnitude of wage changes. Combining the results of the wage and the hazard equations, we see very few wage changes that have taken place during either of the freeze policies. But if they did, their size was relatively large during fd1 while small during fd2. During the ceiling policies the size of wage rises were severely suppressed. Amongst them, Cd5 seems to be most successful in reducing $\dot{w}_{ns}$ although it did not appear to have any effect on the frequency of wage changes. On the other hand, the 12-month rule policy seems to have had an opposite positive effect, as we might expect, with an impact of about 1-2 percentage points. This policy was hardly significant in the hazard equation.

Finally, table 5.9 lists the estimates of the preferred equation. Overall effects are intact. Industry unemployment now has the effect of reducing the rate of wage changes by 0.02 percentage points for its 1% increase.

5.5.3 Summary

In the model that takes into account of the continuously evolving economic background, duration dependence (i.e., the way the hazard evolves with elapsed duration) arises from two sources: the baseline hazard and the time-varying explanatory variables. A problem inherent in our regression type model is the difficulty in distinguishing between the duration dependence and the effects of time-varying variables that are collinear with the elapsed duration during the spell. This problem will not arise if the economic factors supposed to affect the probability to negotiate have adequate variations across observations. In our model, we have been able to measure the impact of several factors, both time varying as well as spell or group specific, that we consider appropriate in determining the negotiation probability. It was not, however, possible to estimate a coefficient of pure duration dependence due to the identification problem. Hence, the question remains as to whether the true duration dependence has been adequately captured by the limited number of time varying covariates we considered thus far. Nonetheless, an important point of our model is the finding that the timing of the wage bargains in this sample period was endogenous; it was strongly influenced by the economic developments. Our overall results (Table 5.5, 5.6 and Table 5.9) show the importance of the real take home pay in determining the timing of negotiation while a concern over the relative wage becomes paramount in determining the size of wage claim. As expected, the overall impact of the incomes policies is strong on both the hazard and wage changes, with an exception of the ceiling policy, cd4, which has a particularly weak impact on the probability of negotiation. Note also that irrespective of the exclusion of the duration dependence, our coefficient estimates on x(t) bear the interpretation of comparative statics impact on the log of duration.

A full analysis of the impact of time-varying covariates has to be done using the dynamic simulation technique, which will be discussed in the next chapter. It would also be of interest to find out how the prediction based on the static model, where the exit probability is not affected by any event once the spell starts, differs from that based on our dynamic model. This will also be done by means of dynamic simulation in the following chapter.

We can put the duration and wage change equations together to yield a single combined incomes policy effect on the average rate of wage change per unit of time, namely, 1 year. If w denotes the rate of change of agreed wages, then w/δ is the average rate of wage change per unit of time. Its response to a new incomes policy is:

$$\Delta(w/\delta) \cong \delta^{-1} \Delta w - (w/\delta) \Delta \ln(\delta)$$

by differentiation. Consider the effect of fd2 (Jul 66 to Jun 67 freeze policy), for example. This has coefficients of -0.730 and -0.037 in the duration and wage equations. The average value of wage change that time was around 6%, with duration around a year. Thus:

$$\begin{aligned}\Delta(w/\delta) &\cong -0.037 - 0.06 \times 0.730 \\ &= -8.1\%\end{aligned}$$

Our estimates imply, therefore, that total incomes policy impact reduces the rate of increase of wages by up to 8 percentage points per year.

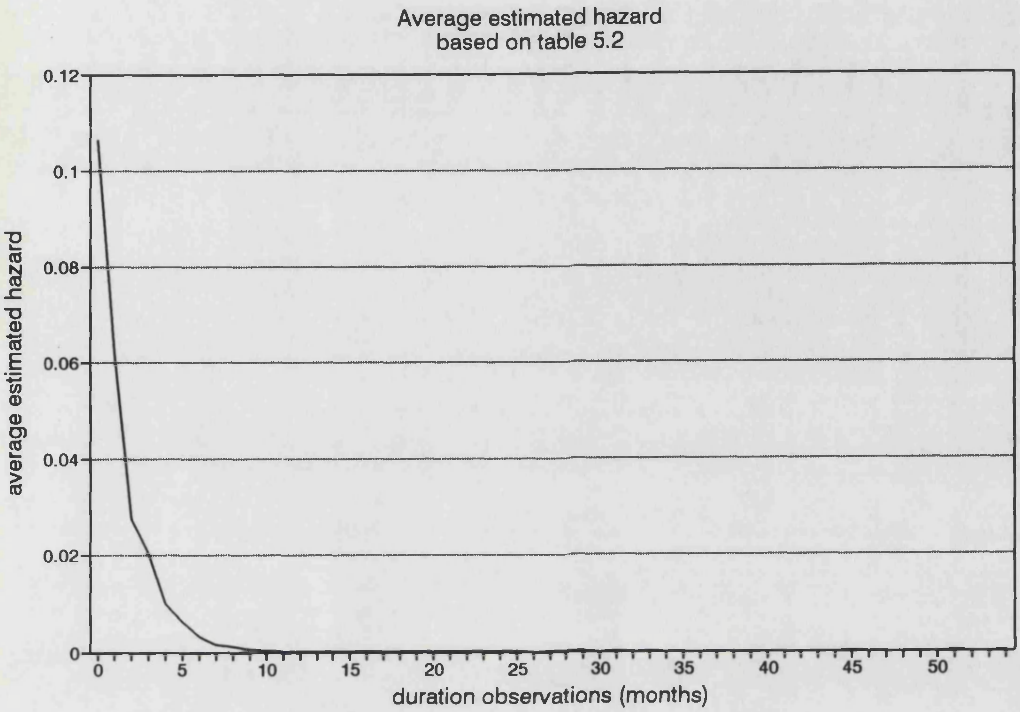


Figure 5.2

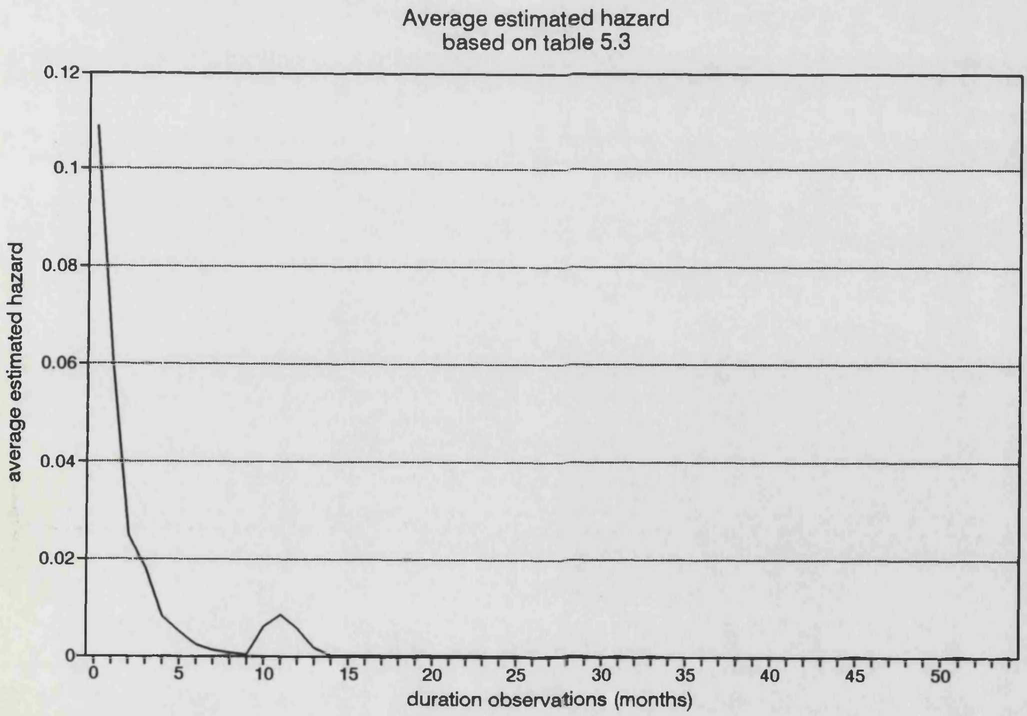


Figure 5.3

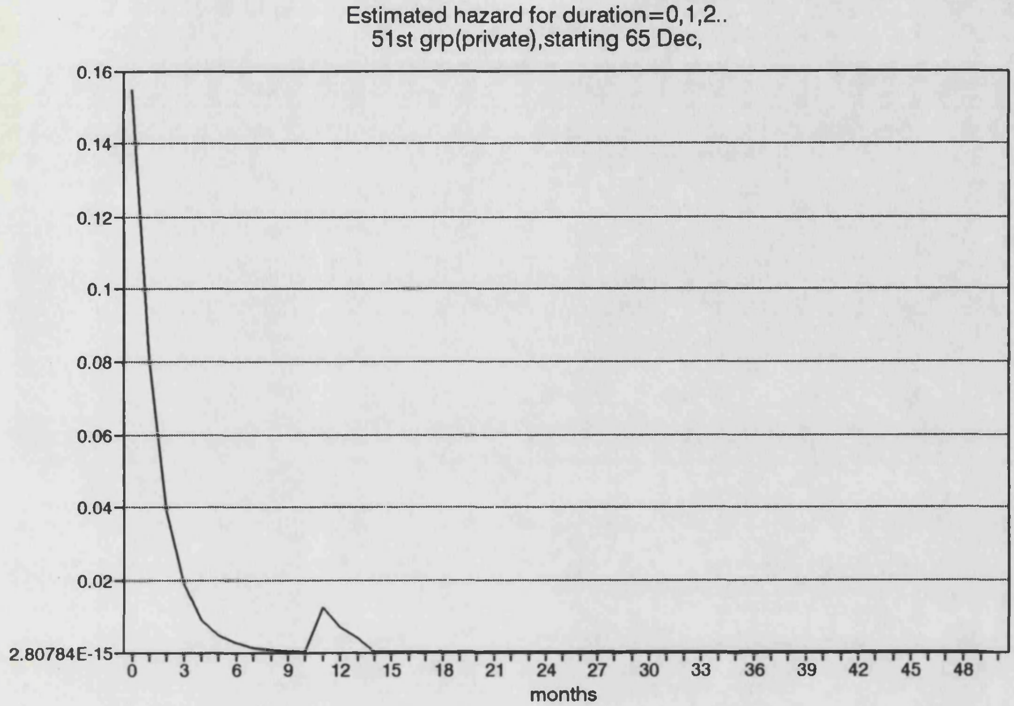


Figure 5.4

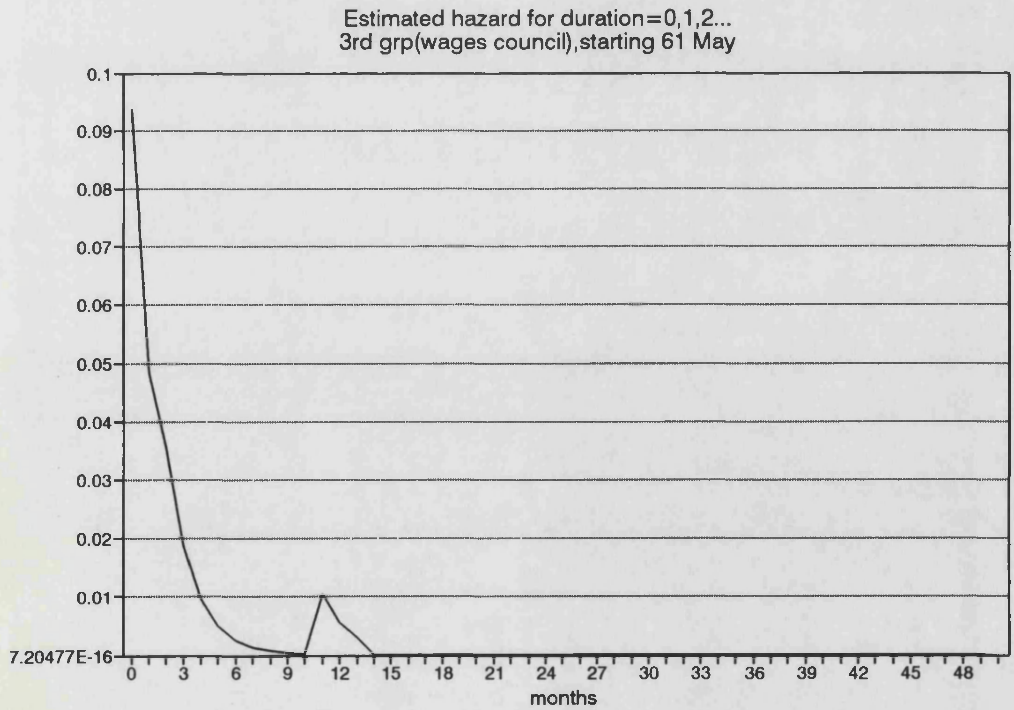


Figure 5.5

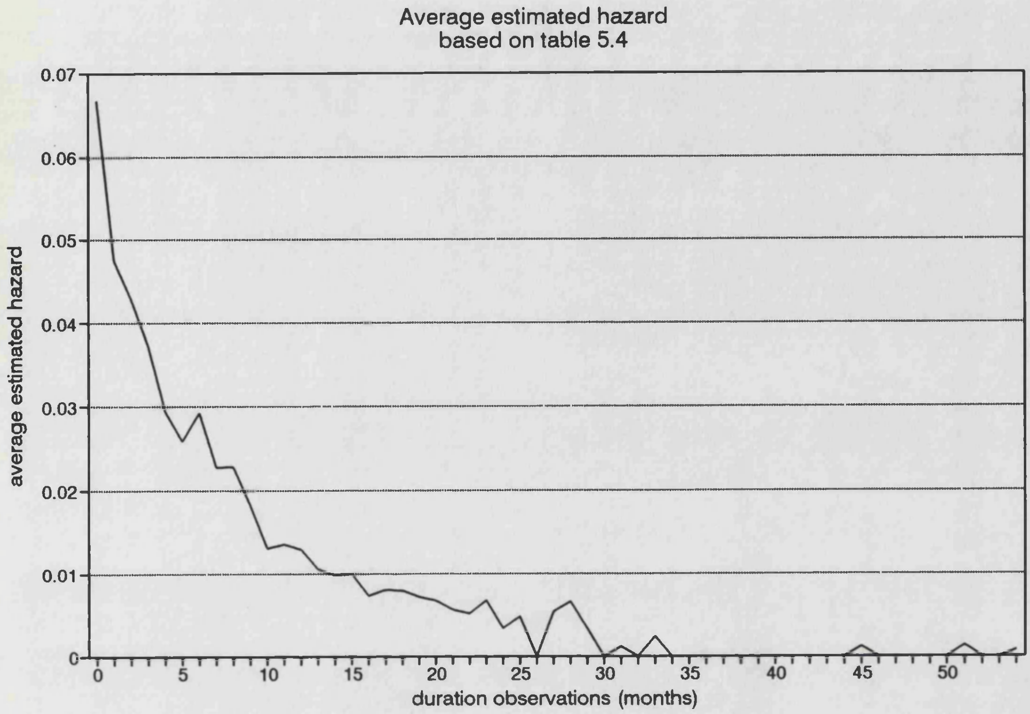


Figure 5.6

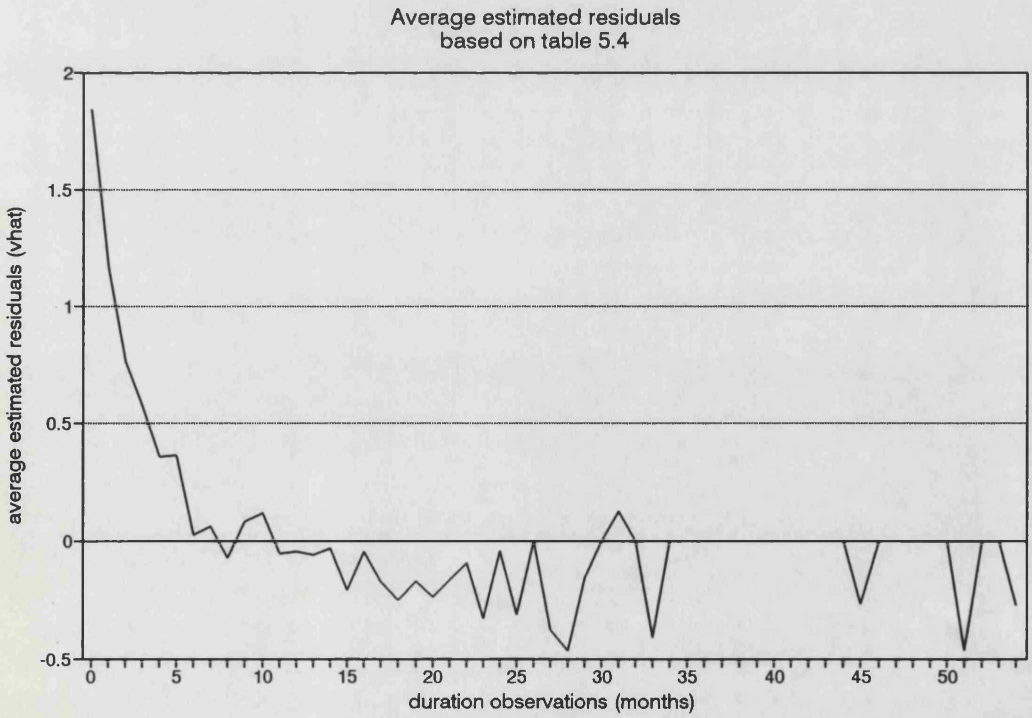


Figure 5.7

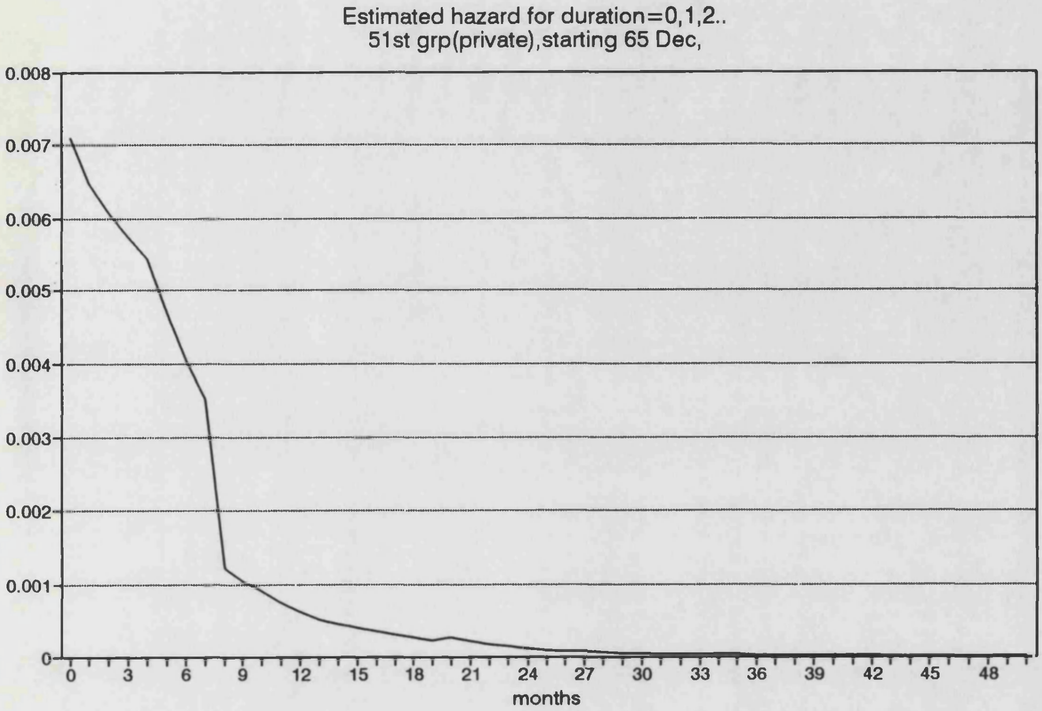


Figure 5.8

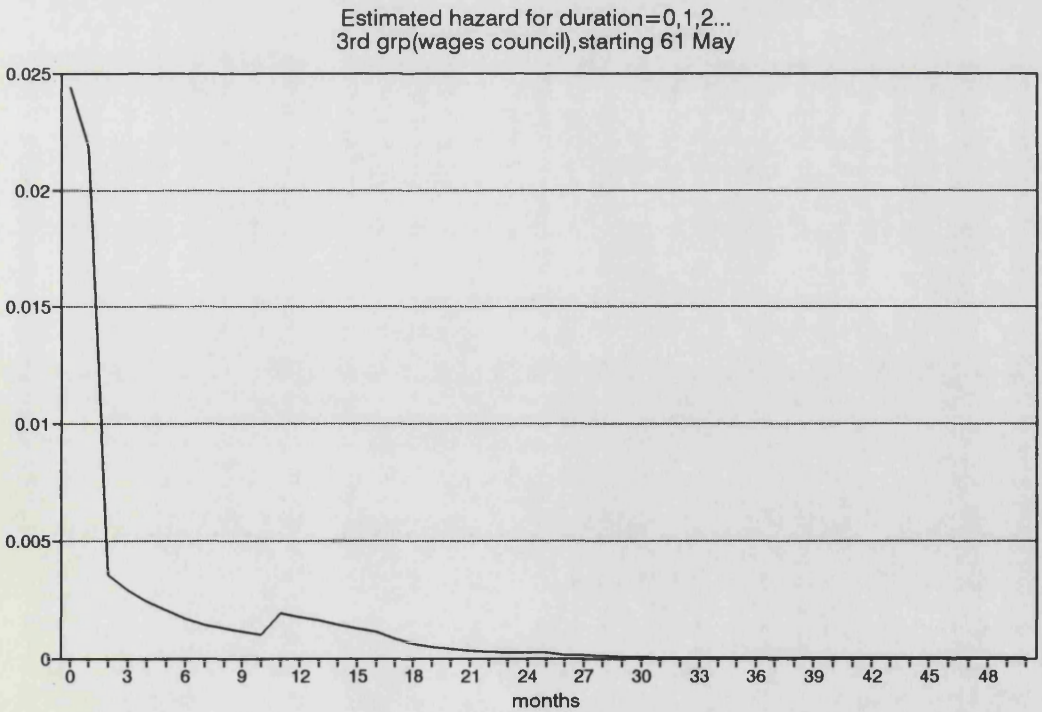


Figure 5.9

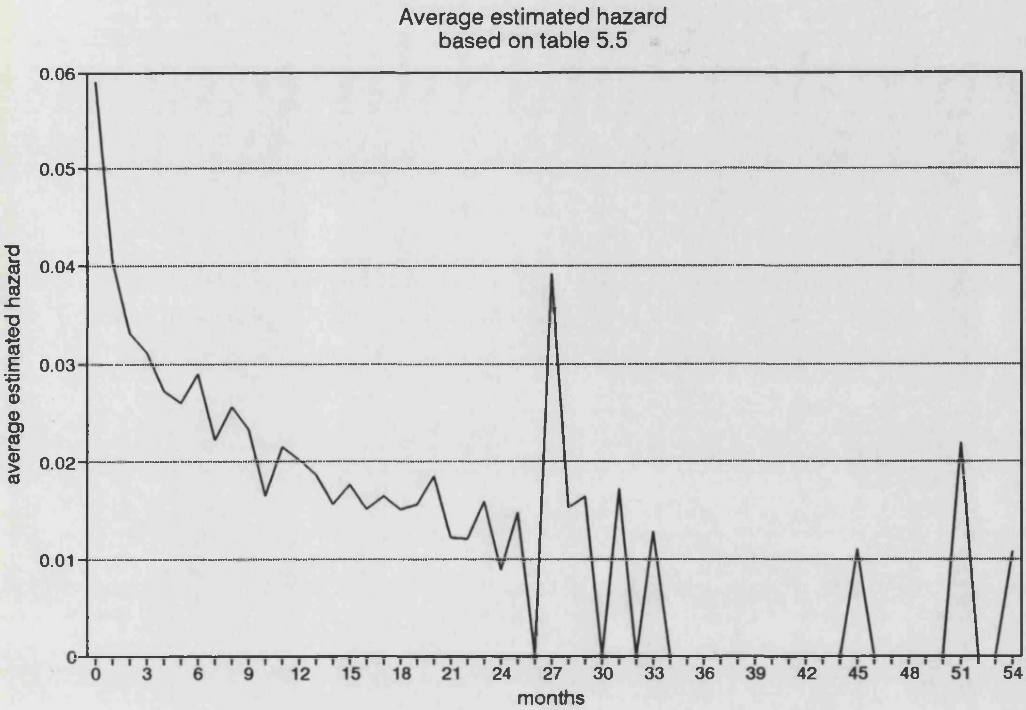


Figure 5.10

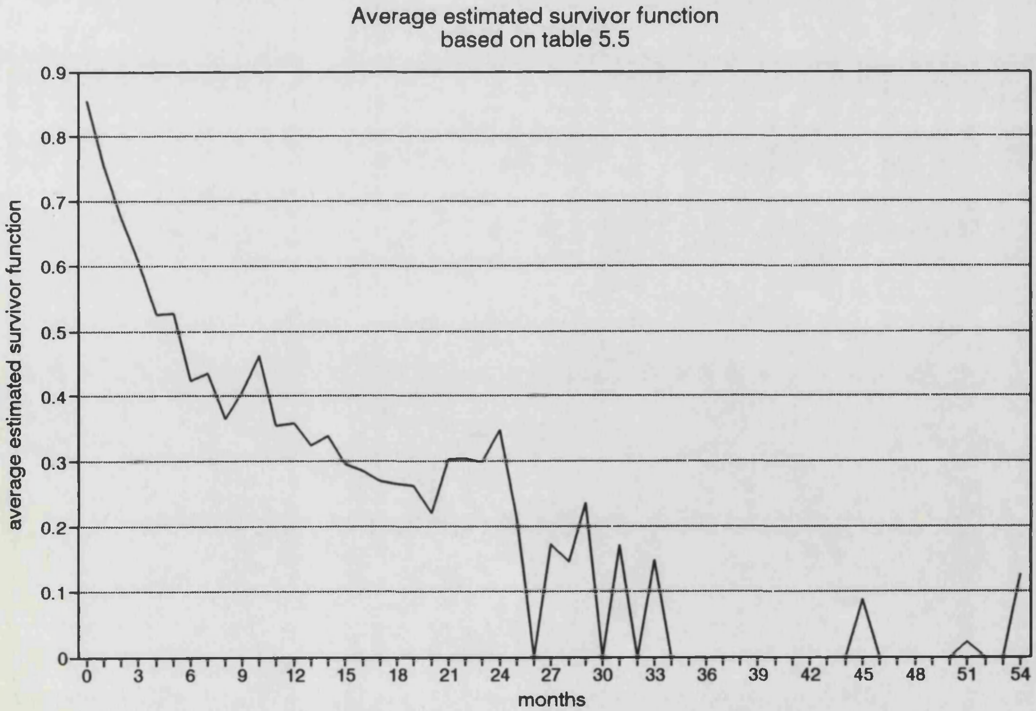


Figure 5.11

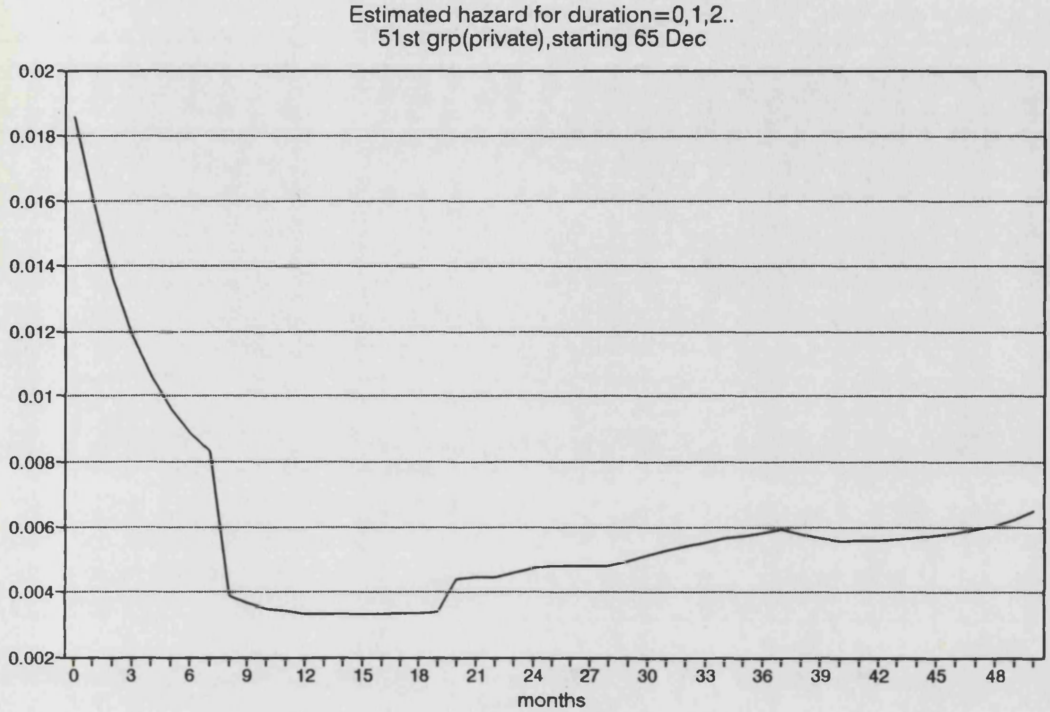


Figure 5.12

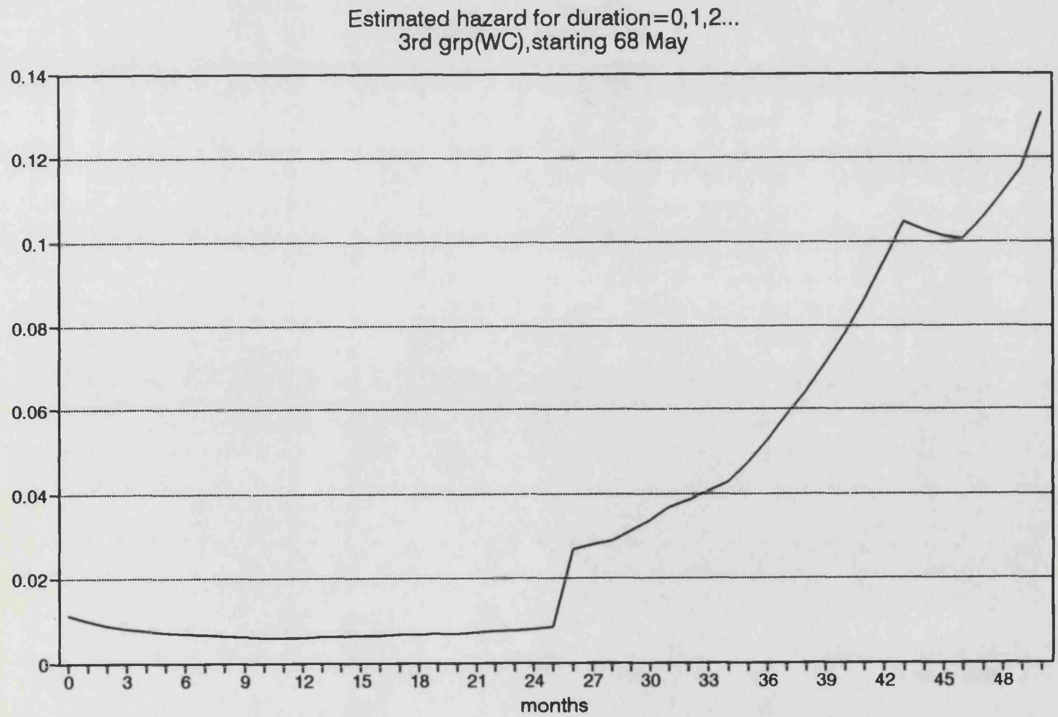


Figure 5.13

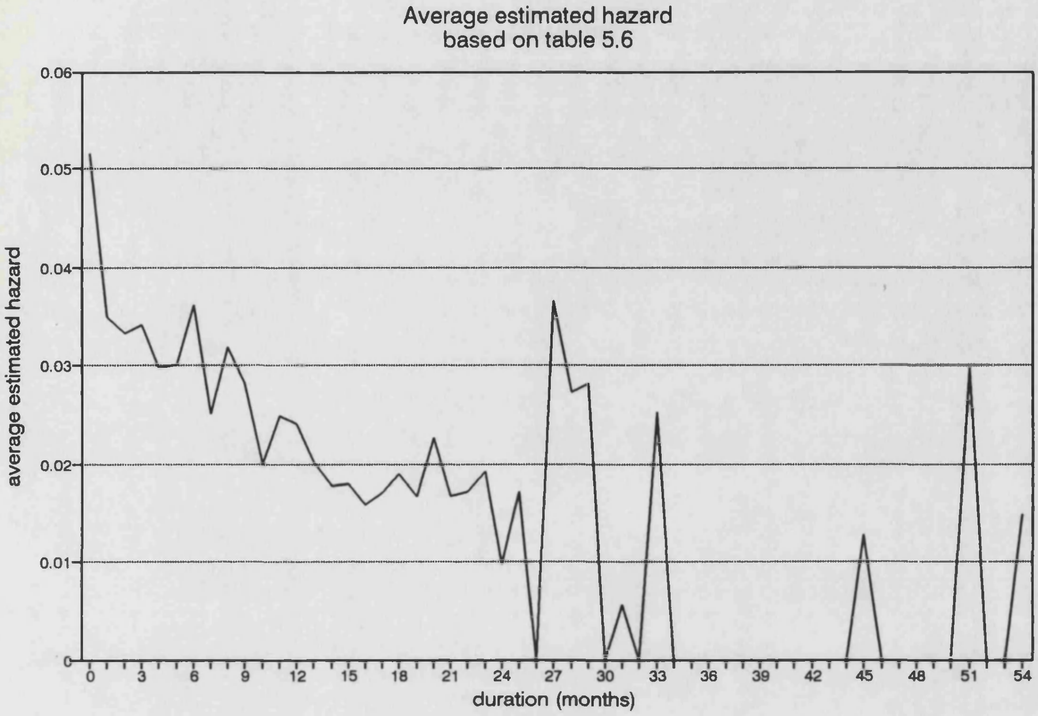


Figure 5.14

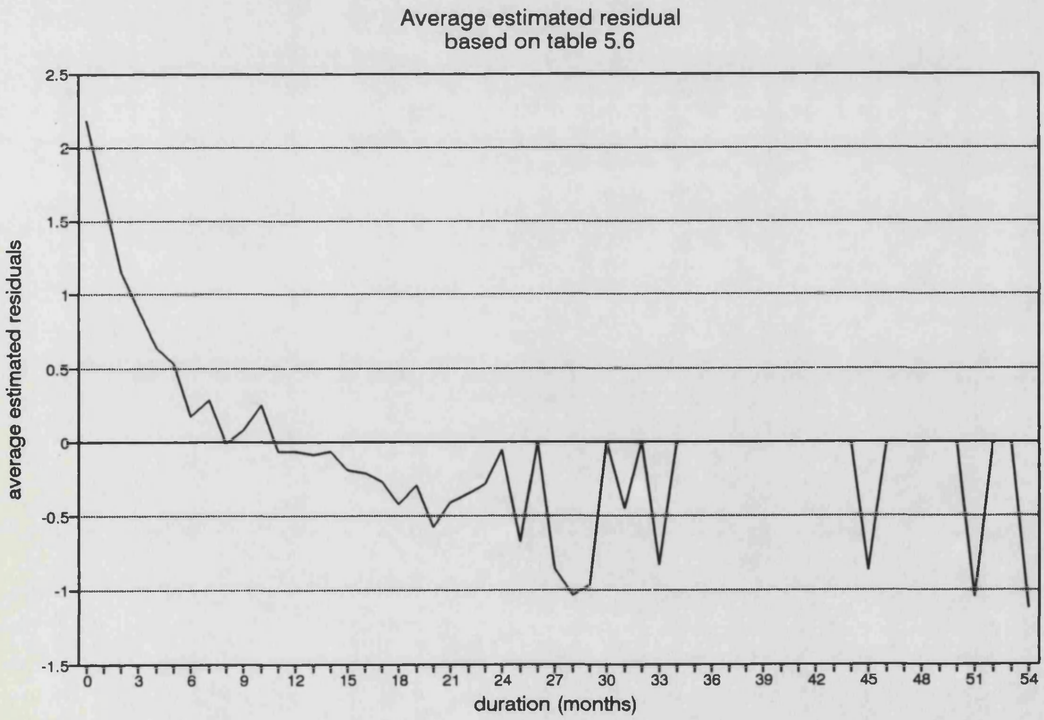
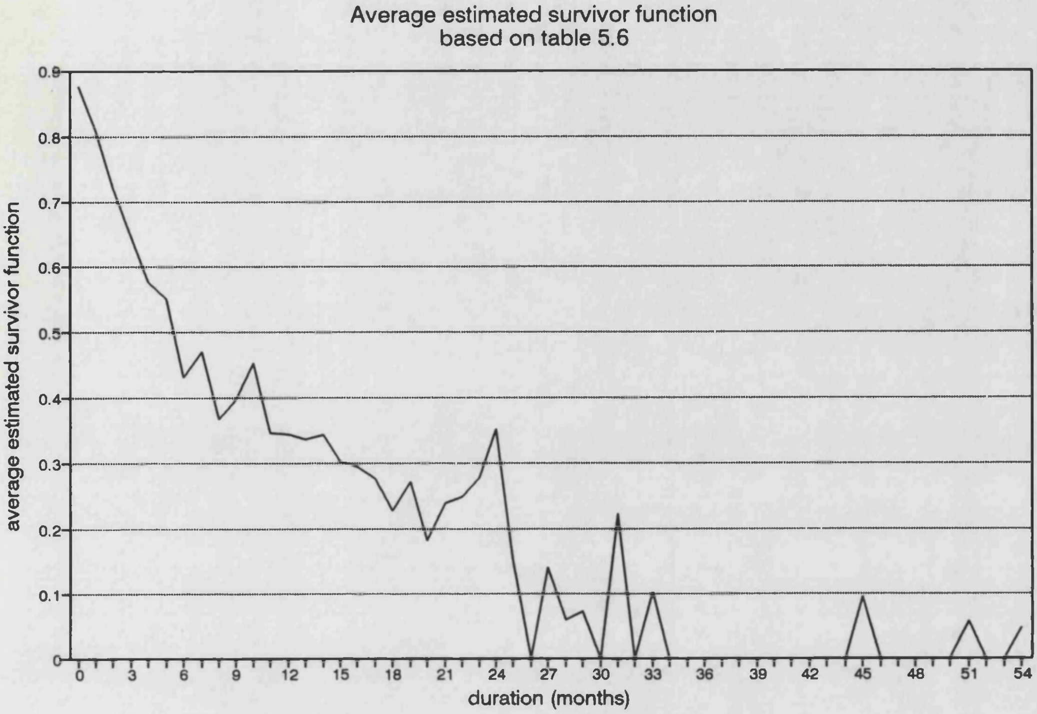


Figure 5.15



Average estimated survivor function
based on table 5.6
average estimated survivor function
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0
0
3
6
9
12
15
18
21
24
27
30
33
36
39
42
45
48
51
54
duration (months)

Figure 5.16

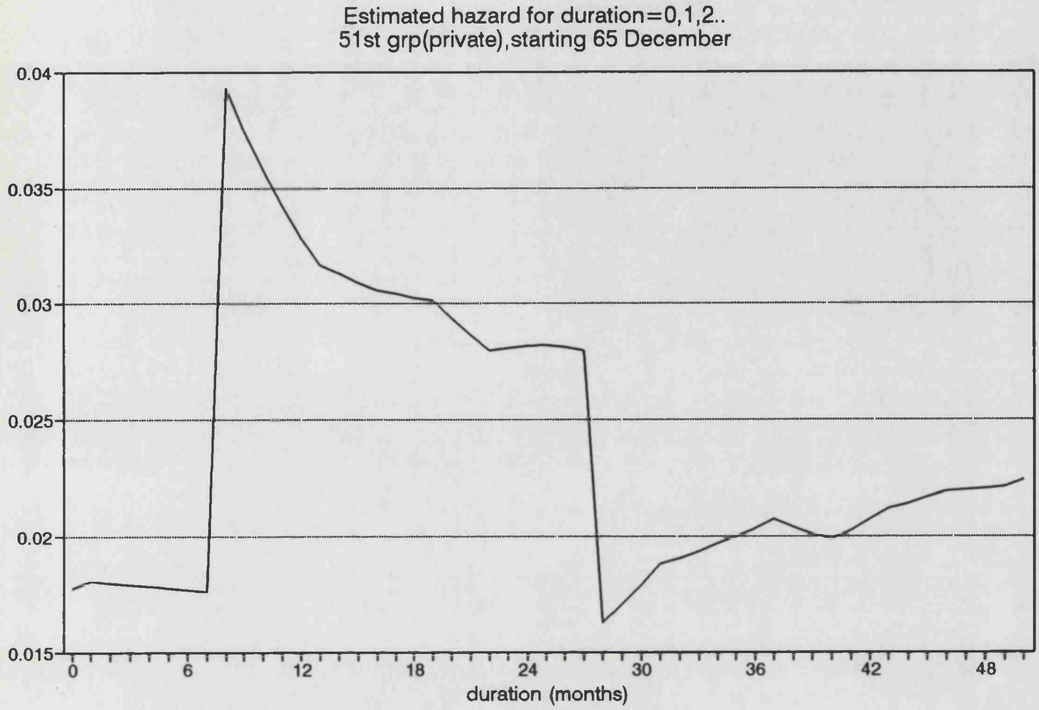


Figure 5.17

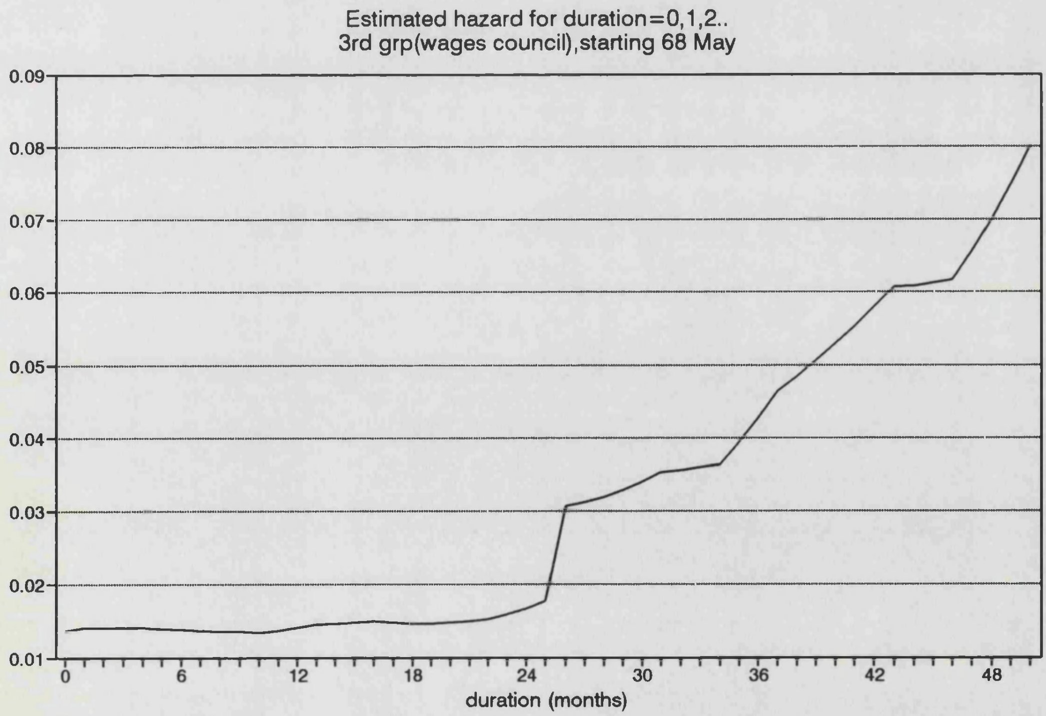


Figure 5.18

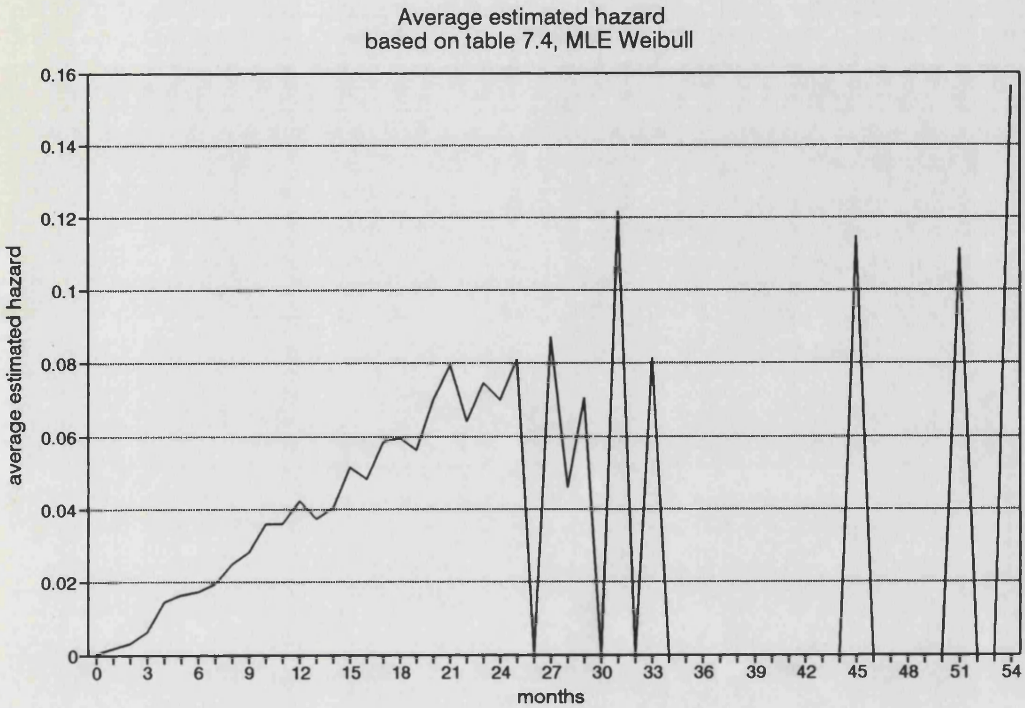


Figure 5.19

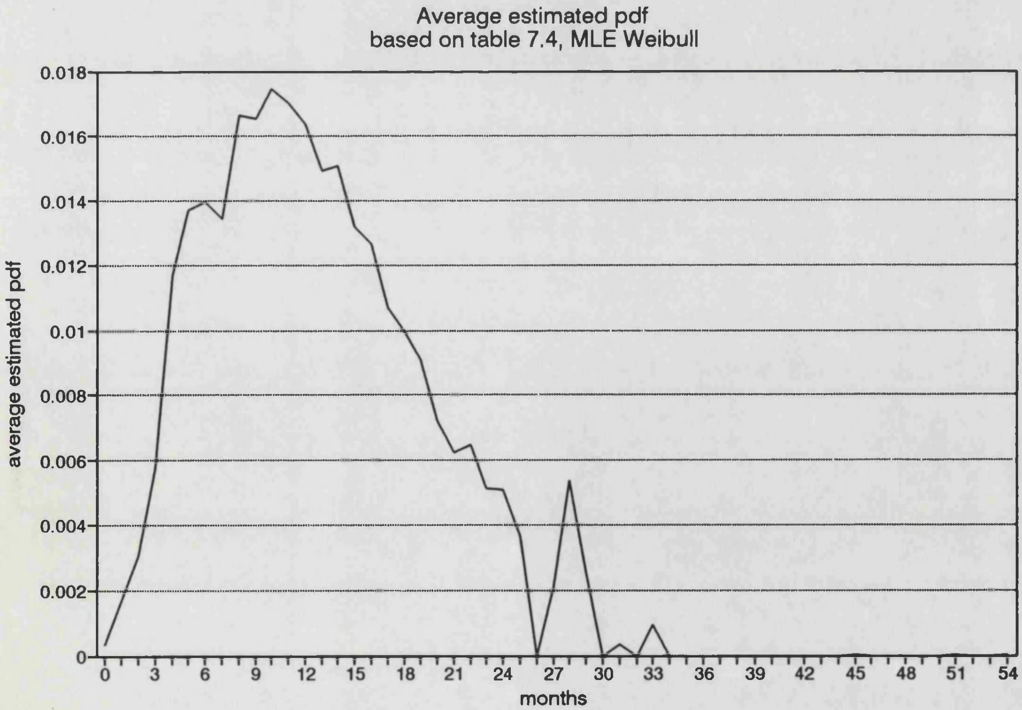


Figure 5.20

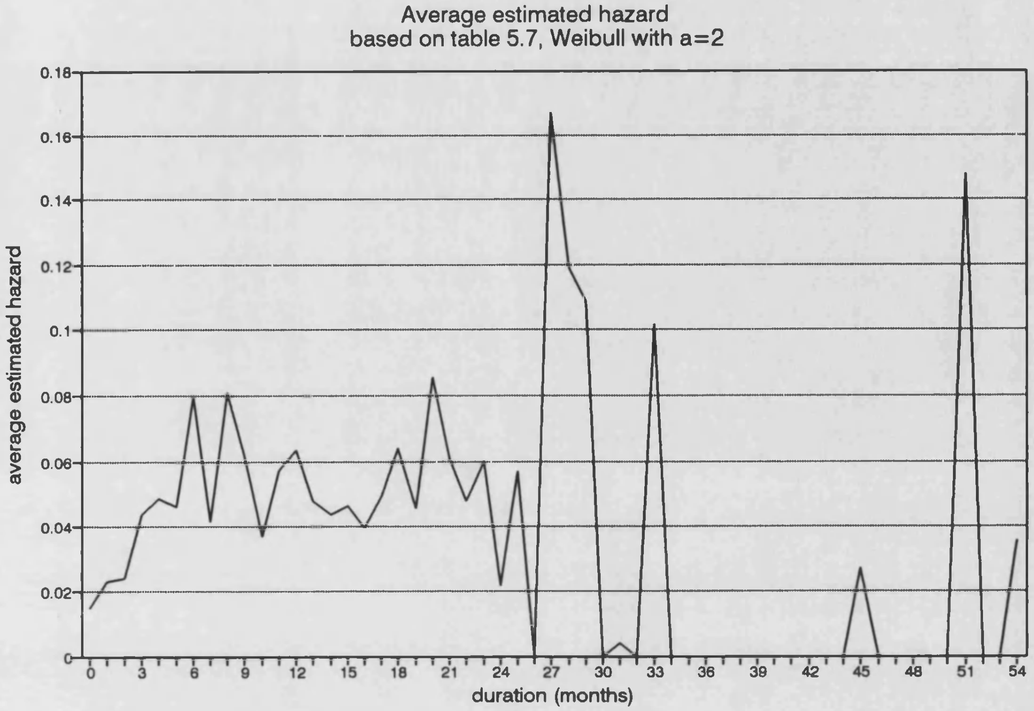


Figure 5.21

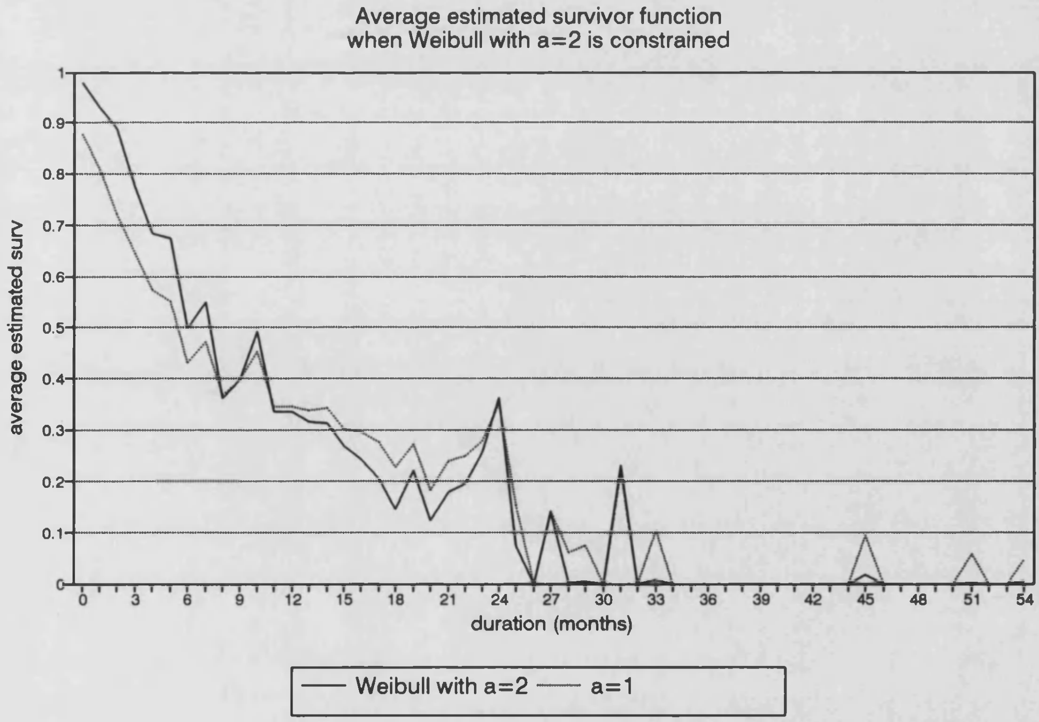


Figure 5.22

Table 5.1 Estimates of the duration equation[8]
(unskilled males; manufacturing and construction industry)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β and γ_2)			
$\ln(\delta_{s-1})$	0.168	0.055	3.07
Drw1	-1.860	1.112	1.67
$staged_{s-1}$	0.304	0.247	1.23
$dfix_{s-1}$	-0.366	0.216	1.70
$Drr^e(t)$	8.361	3.129	2.67
$Drr^u(t)$	-87.229	89.470	0.97
lnrw(t)	2.843	2.054	1.38
rel(t)	2.093	0.725	2.89
lnunin(t)	-0.696	0.144	4.82
lnprin(t)	-0.237	0.751	0.32
size(t)	502.70	819.10	0.61
α[9]	$-8.6*10^{-18}$	0.472	$1.8*10^{-17}$
$fd_1(t)$	-1.191	0.658	1.81
$fd_2(t)$	-0.945	0.316	2.99
$cd_1(t)$	-1.123	0.457	2.46
$cd_2(t)$	-1.116	0.345	3.23
$cd_3(t)$	-1.032	0.357	2.89
$cd_4(t)$	-0.271	0.327	0.83
$cd_5(t)+cd_6(t)$	-0.523	0.129	4.04

$N = 61$; $\Sigma S_n = 850$; Sargan $\chi^2(34) = 101.97$ p-value $=1.0*10^{-8}$

Variable	Coefficient	Standard error	t-ratio
group-specific effects (γ_0 and γ_1)			
constant	13.128	0.0744	17.65
W Council	0.013	0.612	0.02
ublic	0.625	1.425	0.43
TU	0.055	0.325	0.17

Table 5.2 Estimates of the duration equation[10]
(unskilled males; manufacturing and construction ind)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β and γ_2)			
$\ln(\delta_{s-1})$	0.037	0.013	2.92
Drw1	0.492	0.279	1.76
$staged_{s-1}$	0.241	0.068	3.55
$dfix_{s-1}$	-0.177	0.036	4.87
$Drr^e(t)$	-13.891	5.413	2.57
$Drr^u(t)$	-4.812	3.407	1.41
lnrw(t)	-0.586	0.337	1.74
rel(t)	0.092	0.172	0.54
prod(t)	6.623	5.565	1.19
lnunin(t)	-0.002	0.041	0.06
proin11(t)	0.098	0.088	1.11
size(t)	2365.3	666.78	3.55
α	-0.663	0.081	8.16
$fd_1(t)$	0.185	0.171	1.09
$fd_2(t)$	0.085	0.075	1.14
$cd_1(t)$	0.375	0.119	3.16
$cd_2(t)$	0.310	0.084	3.69
$cd_3(t)$	0.156	0.083	1.88
$cd_4(t)$	-0.020	0.048	0.41
$cd_5(t)+cd_6(t)$	-0.001	0.025	0.04
N = 61; $\Sigma\ S_n$ = 850; Sargan $\chi^2(33)$ = 32.864; p-value =0.474			
group-specific effects (γ_0 and γ_1)			
constant	-6.098	0.127	48.02
W Council	-0.084	0.087	0.97
Public	-0.074	0.595	0.12
TU	0.006	0.042	0.14

Table 5.3 Estimates of the duration equation[10]
(unskilled males; manufacturing and construction ind)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β *and* γ_2)			
$\ln(\delta_{s-1})$	0.023	0.011	1.99
Drw1	0.428	0.262	1.64
$staged_{s-1}$	0.226	0.062	3.64
$dfix_{s-1}$	-0.168	0.041	4.11
$Drr^e(t)$	-13.199	5.121	2.58
$Drr^u(t)$	-10.496	2.901	3.62
lnrw(t)	-1.172	0.395	2.97
rel(t)	0.334	0.193	1.74
prod(t)	9.488	5.061	1.87
lnunin(t)	0.001	0.050	0.03
prin11(t)	0.165	0.092	1.79
size(t)	1859.20	498.41	3.73
$\alpha_1(t)$	-0.728	0.079	9.26
$\alpha_2(t)$	4.641	1.077	4.31
$fd_1(t)$	0.249	0.166	1.50
$fd_2(t)$	0.093	0.082	1.13
$cd_1(t)$	0.389	0.096	4.07
$cd_2(t)$	0.320	0.085	3.77
$cd_3(t)$	0.148	0.890	1.66
$cd_4(t)$	-0.068	0.058	1.17
$cd_5(t)+cd_6(t)$	-0.035	0.028	1.25
$N = 61$; $\Sigma S_n = 850$; Sargan $\chi^2(32) = 31.992$ p-value =0.467			
group-specific effects (γ_0 *and* γ_1)			
constant	-10.066	0.144	69.90
W Council	-0.199	0.098	2.04
Public	-0.047	0.580	0.08
TU	0.006	0.051	0.12

Table 5.4 Estimates of the duration equation[11]
(unskilled males; manufacturing and construction ind)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β and γ_2)			
$\ln(\delta_{s-1})$	0.119	0.017	6.87
Drw1	-2.294	0.295	7.78
$staged_{s-1}$	0.445	0.082	5.42
$dfix_{s-1}$	-0.358	0.070	5.09
$Drr^e(t)$	26.85	2.745	9.78
$Drr^u(t)$	-3.366	2.066	1.63
lnrw(t)	3.313	0.908	3.65
rel(t)	2.080	0.319	6.51
lnunin(t)	-0.510	0.060	8.52
lnprin(t)	-1.081	0.337	3.21
size(t)	673.8	361.8	1.86
$fd_1(t)$	-0.769	0.250	3.08
$fd_2(t)$	-0.845	0.137	6.18
$cd_1(t)$	-0.450	0.183	2.46
$cd_2(t)$	-0.603	0.138	4.38
$cd_3(t)$	-0.592	0.137	4.33
$cd_4(t)$	-0.451	0.090	5.00
$cd_5(t)+cd_6(t)$	-0.421	0.051	8.26

$N = 61$; $\Sigma S_n = 850$; Sargan $\chi^2(35) = 47.889$ p-value $=0.072$

group-specific effects (γ_0 and γ_1)

Variable	Coefficient	Standard error	t-ratio
constant	20.59	0.645	31.94
W Council	-0.307	0.422	0.73
Public	0.406	0.331	1.23
TU	0.056	0.247	0.23

Table 5.5 Estimates of the duration equation[12]
(unskilled males; manufacturing and construction ind)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β *and* γ_2)			
$\ln(\delta_{s-1})$	0.079	0.052	1.54
Drw1	-2.11	0.785	2.69
$\ln rw_{s-1}$	-3.357	1.196	2.81
$staged_{s-1}$	0.257	0.221	1.17
$dfix_{s-1}$	-0.445	0.159	2.80
$Drr^u(t)$	-13.074	6.088	2.15
rel(t)	1.632	0.544	3.00
lnunin(t)	-0.563	2.437	0.23
prin11(t)	2.009	0.818	2.46
size(t)	29.158	1543.0	0.02
$\alpha_1(t)$	-32.106	61.51	0.52
$fd_1(t)$	-17.94	6413	0.003
$fd_2(t)$	-1.044	0.901	1.16
$cd_1(t)$	-14.92	708795	$-2.1*10^{-5}$
$cd_2(t)$	-38.14	33.39	1.14
$cd_3(t)$	-1.235	0.898	1.38
$cd_4(t)$	-0.860	0.932	0.92
$cd_5(t)+cd_6(t)$	-5.059	22.40	0.23

$N = 61$; $\Sigma S_n = 850$; Sargan $\chi^2(35) = 1.09*10^{11}$ p-value $=7.7*10^{-7}$

Variable	Coefficient	Standard error	t-ratio
group-specific effects (γ_0 *and* γ_1)			
constant	-9.701	0.142	68.32
W Council	-0.649	0.345	1.88
Public	0.560	0.319	1.76
TU	0.029	0.133	0.22

Table 5.6 Estimates of the duration equation[13]
(unskilled males; manufacturing and construction ind)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β and γ_2)			
$\ln(\delta_{s-1})$	0.090	0.028	3.18
Drw1	-2.806	0.435	6.45
$staged_{s-1}$	0.405	0.117	3.47
$dfix_{s-1}$	-0.625	0.101	6.21
$Drr^u(t)$	-4.673	1.041	4.49 — unexpected
rel(t)	1.035	0.413	2.50
lnunin(t)	-0.475	0.058	8.14
lnprin(t)	0.478	0.235	2.04
size(t)	16.35	38.70	0.42
$fd_1(t)$	-1.233	0.292	4.23
$fd_2(t)$	-0.730	0.134	5.46
$cd_1(t)$	-1.071	0.260	4.12
$cd_2(t)$	-1.050	0.240	4.38
$cd_3(t)$	-0.855	0.191	4.49
$cd_4(t)$	-0.031	0.091	0.337
$cd_5(t)+cd_6(t)$	-0.536	0.073	7.35

$N = 61$; $\Sigma S_n = 850$; Sargan $\chi^2(37) = 39.877$ p-value $=0.343$

Variable	Coefficient	Standard error	t-ratio
group-specific effects (γ_0 and γ_1)			
constant	-7.205	0.183	39.35
W Council	0.073	0.099	0.74
Public	0.575	0.169	3.40
TU	0.0379	0.060	0.63 R?

Table 5.7 Estimates of the duration equation[14]
(unskilled males; manufacturing and construction ind)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β and γ_2)			
$\ln(\delta_{s-1})$	0.145	0.058	2.48
Drw1	-6.335	0.980	6.47
$staged_{s-1}$	1.185	0.253	4.58
$dfix_{s-1}$	-1.259	0.216	5.82
$Drr^u(t)$	-9.617	1.343	7.16
rel(t)	0.098	0.712	0.138
lnunin(t)	-0.908	0.101	8.95
lnprin(t)	1.471	0.425	3.46
size(t)	86.70	54.29	1.60
$fd_1(t)$	-2.056	0.639	3.22
$fd_2(t)$	-1.220	0.265	4.61
$cd_1(t)$	-1.879	0.542	3.46
$cd_2(t)$	-1.893	0.479	3.95
$cd_3(t)$	-1.665	0.401	4.15
$cd_4(t)$	-0.209	0.201	1.04
$cd_5(t)+cd_6(t)$	-1.036	0.129	8.03

N = 61; $\Sigma\, S_n = 850$; Sargan $\chi^2(37) = 38.138$ p-value =0.417

Variable	Coefficient	Standard error	t-ratio
group-specific effects (γ_0 and γ_1)			
constant	-9.404	0.555	16.94
W Council	0.665	0.284	2.34
Public	0.969	0.336	2.88
TU	0.056	0.183	0.30

Table 5.8 Estimates of the wage equation[15]
(unskilled males; manufacturing and construction ind)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β and γ_2)			
Drr(T_s)	0.046	0.060	0.76
lnrw(T_s)	-0.348	0.075	4.64
Drw1	0.080	0.032	2.47
rel(T_s)	0.240	0.026	9.37
lnun(T_s)	0.047	0.020	2.35
lnunin(T_s)	-0.042	0.016	2.67
prod(T_s)	0.440	0.092	4.81
prin(T_s)	-0.036	0.018	1.96
size(T_s)	73.33	18.60	3.94
$staged_s$	-0.002	0.009	0.20
$sfix_s$	-0.015	0.011	1.31
$\ln(\delta_{s-1})$	-0.001	0.002	0.68
$\ln(\delta_s)$	-0.007	0.004	1.63
$fd_1(T_s)$	-0.024	0.037	0.65
$fd_2(T_s)$	-0.037	0.010	3.78
$cd_1(T_s)$	-0.062	0.016	3.83
$cd_2(T_s)$	-0.057	0.016	3.63
$cd_3(T_s)$	-0.055	0.014	3.95
$cd_4(T_s)$	-0.054	0.011	4.98
$cd_5(T_s)$	-0.077	0.008	9.29
$cd_6(T_s)$	-0.045	0.012	3.56
d12(T_s)	0.009	0.006	1.43

N = 61; ΣS_n = 850; Sargan $\chi^2(18)$ = 23.810 p-value =0.161

group-specific effects (γ_0 and γ_1)

Variable	Coefficient	Standard error	t-ratio
constant	-2.691	0.040	68.00
W Council	-0.068	0.023	3.01
Public	0.054	0.051	1.06
TU	0.002	0.013	0.17

Table 5.9 Estimates of the wage equation[15]
(unskilled males; manufacturing and construction ind)

Variable	Coefficient	Standard error	t-ratio
spell specific and continuous-time effects (β and γ_2)			
$\ln rw(T_s)$	-0.259	0.069	3.73
Drw1	0.057	0.023	2.50
$rel(T_s)$	0.204	0.020	10.02
$\ln un(T_s)$	0.027	0.019	1.378
$\ln unin(T_s)$	-0.021	0.014	1.54
$prod(T_s)$	0.309	0.064	4.82
$size(T_s)$	69.947	14.109	4.96
$staged_s$	-0.039	0.025	1.57
$fd_1(T_s)$	-0.037	0.032	1.15
$fd_2(T_s)$	-0.037	0.008	4.74
$cd_1(T_s)$	-0.065	0.014	4.63
$cd_2(T_s)$	-0.062	0.015	4.17
$cd_3(T_s)$	-0.052	0.012	4.22
$cd_4(T_s)$	-0.058	0.010	5.60
$cd_5(T_s)$	-0.079	0.007	11.58
$cd_6(T_s)$	-0.046	0.011	3.91
$d12(T_s)$	0.021	0.006	3.62

$N = 61$; $\Sigma S_n = 850$; Sargan $\chi^2(24) = 25.467$ P-value $= 0.381$

Variable	Coefficient	Standard error	t-ratio
group-specific effects (γ_0 and γ_1)			
constant	-2.022	0.032	63.60
W Council	-0.049	0.019	2.53
Public	0.052	0.055	0.95
TU	0.002	0.010	0.17

Table 5.10 Descriptive statistics for the spell specific variables used in the regressions (levels)

number of cross-section groups = 61 total observation = 850

variables	Means	Std Dev	Min	Max	observation with value 1
DW	0.066469	0.042248	-0.08683	0.319498	
LNDUR	3.8980	0.57978	0.00000	5.4553	
LNDUR-1	3.9071	0.55203	0.00000	5.4553	
DRW1	0.018306	0.040086	-0.12792	0.23181	
$DRR(T_{ns})$	-0.042866	0.030933	-0.17169	0.058915	
$REL(T_{ns})$	1.6691	0.21744	1.3021	2.3923	
$LNRW(T_{ns})$	-6.3795	0.10021	-6.6562	-6.0600	
$LNUN(T_{ns})$	6.0544	0.34729	5.2428	6.8337	
$LNUNIN(T_{ns})$	2.2480	0.68635	-0.48551	4.3053	
$PROD(T_{ns})$	0.028225	0.020491	-0.00800	0.14200	
$PRIN(T_{ns})$	0.055124	0.10678	-0.34558	0.53509	
$SIZE(T_{ns})$	-1.808E-005	0.00017	-0.00207	0.00079	
$STAGED_{ns}$	0.20824	0.40605	0.00000	1.0000	177
$STAGED_{ns-1}$	0.18588	0.38901	0.00000	1.0000	158
$DFIX_{ns-1}$	0.068235	0.25215	0.00000	1.0000	58
$SFIX_{ns}$	0.030588	0.17220	0.00000	1.0000	26
$FD1(T_{ns})$	0.0047059	0.068438	0.00000	1.0000	4
$CD1(T_{ns})$	0.024706	0.15523	0.00000	1.0000	21
$CD2(T_{ns})$	0.040000	0.19596	0.00000	1.0000	34
$CD3(T_{ns})$	0.028235	0.16564	0.00000	1.0000	24
$FD2(T_{ns})$	0.034118	0.18153	0.00000	1.0000	29
$CD4(T_{ns})$	0.051765	0.22155	0.00000	1.0000	44
$CD5(T_{ns})$	0.10824	0.31068	0.00000	1.0000	92
$CD6(T_{ns})$	0.035294	0.18452	0.00000	1.0000	30
$DT12(T_{ns})$	0.24588	0.43061	0.00000	1.0000	208

Table 5.11 Descriptive statistics of instruments used for the duration equation (levels)

number of cross-section groups= 61 total observation= 850

Variable	Mean	Std Dev	Min	Max
DRR2	-0.041937	0.025250	-0.17332	0.01047
RR52	-0.039873	0.020772	-0.092162	0.016741
DRR104	-0.042788	0.021485	-0.095781	0.0069123
DRW2	0.014333	0.034816	-0.12842	0.17983
DW2	0.058989	0.033135	-0.086826	0.26203
DP2	0.038274	0.023050	-0.005370	0.16588
RPI2	0.038274	0.023050	-0.0053706	0.16588
RPI104	0.040742	0.021884	-0.0021482	0.10121

Table 5.11 Descriptive statistics of instruments used for the duration equation (contd')

number of cross-section groups= 61 total observation= 850

Variable	Mean	Std Dev	Min	Max
RPI52	0.036450	0.019346	-0.010672	0.099644
DRR2	-0.0058327	0.0071845	-0.041420	0.012516
DRR52	-0.0052515	0.0075792	-0.023840	0.020474
DRR104	-0.0042247	0.0082945	-0.023704	0.017387
UNEM2	0.090275	0.31625	-0.49061	1.3508
UNEM52	0.10797	0.27334	-0.31572	1.0257
UNEM104	0.14253	0.27882	-0.28703	0.83987
REL2	1.6079	0.18958	1.2940	2.3713
REL52	1.6212	0.19806	1.2762	2.2623
PROD2	0.025454	0.020356	-0.0080357	0.12394
PROD52	0.020343	0.013969	-0.0039604	0.072835
PROD104	0.025434	0.015505	0.0019920	0.079282
EARN2	0.069466	0.043125	-0.075749	0.32132
EARN52	0.059940	0.039289	-0.047351	0.21010
EARN104	0.068445	0.044298	-0.047351	0.21010
LNEARN2	2.6260	0.27837	2.1294	3.2438
LNEARN52	2.6836	0.29025	2.2094	3.3534
LNEARN104	2.7489	0.31189	2.2915	3.4965
LNRPI2	4.6418	0.15591	4.3656	5.0330
LNRPI52	4.6775	0.16529	4.3969	5.0795
LNRPI104	4.7172	0.17946	4.4625	5.1602
LNRET2	-0.17030	0.028316	-0.23193	-0.13239
LNRET52	-0.17559	0.029250	-0.23350	-0.13202
LNRET104	-0.17986	0.028590	-0.23350	-0.13643
LNDUR2	3.9143	0.54046	0.00000	5.4161
LNDUR3	3.9354	0.51313	0.00000	4.9836
LNRW2	-6.3470	0.098058	-6.5937	-6.0165
SIZE2[16]	0.0012954	0.0015283	4.7318E-005	0.011382
SIZE52	0.0012830	0.0014908	4.6046E-005	0.010718
SIZE104	0.0012601	0.0014491	4.1667E-005	0.009874
LNUNIN2	2.0879	0.66848	-0.69315	3.7954
LNUNIN52	2.1641	0.68293	-0.69315	4.4232
LNUNIN104	2.2654	0.69471	-0.69315	4.4232
PROIN52	0.045525	0.10197	-0.29851	0.50394
PROIN104	0.046660	0.11141	-0.41667	0.47253
STAGED2	0.16588	0.37197	0.00000	1.0000
DFIX2	0.057647	0.23307	0.00000	1.0000
FD1(T_{ns-2})	0.0047059	0.068438	0.00000	1.0000
C1(T_{ns-2})	0.024706	0.15523	0.00000	1.0000
C2(T_{ns-2})	0.038824	0.19317	0.00000	1.0000
C3(T_{ns-2})	0.028235	0.16564	0.00000	1.0000
FD2(T_{ns-2})	0.036471	0.18746	0.00000	1.0000
C4(T_{ns-2})	0.050588	0.21916	0.00000	1.0000
C5(T_{ns-2})	0.098824	0.29842	0.00000	1.0000
C6(T_{ns-2})	0.028235	0.16564	0.00000	1.0000
DT1(T_{ns-2})	0.19882	0.39911	0.00000	1.0000

Table 5.12 Descriptive statistics for the instruments used for the wage equation (levels)

number of cross-section groups= 61 total observation= 850

variable	Mean	Std Dev	Min	Max
LNDUR1	3.9071	0.55203	0.00000	5.4553
LNDUR2	3.9143	0.54046	0.00000	5.4161
LNDUR3	3.9354	0.51313	0.00000	4.9836
DRR1	-0.041804	0.029247	-0.17169	0.058915
DRW2	0.017258	0.036477	-0.12792	0.21989
RET1	-0.0059445	0.0076599	-0.041420	0.022472
RET2	-0.0058327	0.0071845	-0.041420	0.012516
PROD1	0.026336	0.020807	-0.008000	0.14160
PROD2	0.025453	0.020358	-0.008000	0.12390
REL1	1.6391	0.20375	1.2940	2.3714
REL2	1.6079	0.18959	1.2940	2.3714
SIZ1	0.001284	0.001487	5.05E-5	0.001026
SIZ2	0.002395	0.000153	4.73E-05	0.011382
LNUN1	5.9636	0.34045	5.0907	6.8337
LNUN2	5.8899	0.33368	5.0907	6.5597
LNWHI1	3.5323	0.096725	3.4250	3.7887
LNWHO1	3.7888	0.12587	3.5957	4.1311
FD1	0.0047059	0.068438	0.00000	1.0000
C1	0.025882	0.15878	0.00000	1.0000
C2	0.038824	0.19317	0.00000	1.0000
C3	0.028235	0.16564	0.00000	1.0000
FD2	0.035294	0.18452	0.00000	1.0000
C4	0.051765	0.22155	0.00000	1.0000
C5	0.10588	0.30769	0.00000	1.0000
C6	0.034118	0.18153	0.00000	1.0000
D12	0.23882	0.42636	0.00000	1.0000
STAGED1	0.18588	0.38901	0.00000	1.0000
STAGED2	0.16588	0.37197	0.00000	1.0000
SFIX1	0.029412	0.16896	0.00000	1.0000
LNUNIN1	2.1572	0.67352	-0.69315	4.3053
LNUNIN2	2.0879	0.66848	-0.69315	3.7954
PROIN1	0.051248	0.11015	-0.34558	0.53509
PROIN2	0.055961	0.11459	-0.34558	0.53509
LNRW2	-6.3470	0.098058	-6.5937	-6.0165
LNRW3	-6.3606	0.098290	-6.6458	-6.0631
EARN1	0.071638	0.045494	-0.075749	0.32132
EARN2	0.069466	0.043125	-0.075749	0.32132
RPI1	0.041293	0.025509	-0.0053706	0.19183
RPI2	0.038274	0.023050	-0.0053706	0.16588
DRR2	-0.041937	0.025250	-0.17332	0.010147

Footnotes to chapter 5

1. This is required in determining appropriate lagged explanatory variables as instruments.
2. Our sample is large in terms of the number of bargaining groups, N, but small in terms of the number of observed bargains (which is random and variable across groups); thus, the relevant sampling theory is asymptotic in N.
3. Compute the average of $\hat{h}(\delta_{ns})=\exp(\hat{\gamma}_2'z_{ns}+\hat{\gamma}_1'c_n+\hat{\gamma}_0+\hat{u}_n+\hat{\beta}'x(\delta_{ns}))$ over all observations of duration in nearest months.
4. With respect to the way workers update their information in making the price/retention ratio expectations, we have carried out a similar estimation with one month lag as opposed to no updating, but the overall estimates are very similar including the parameters of the duration dependence. Only difference is seen in the coefficient of Drr^u which is no longer as significant, and is much larger in magnitude. This makes sense since Drr^u under the expectation with 1 month lag is very small and hardly varies over t. On the other hand, if workers have rather slow flow of information, their expectation apt to include significant expectational error (i.e., Drr^u), and such error plays a negative significant role in the hazard.
5. In order to make sure that there is no cycle, possibly annual, exhibited by Drr^e, we seasonally adjusted the retail price index by using TSP's SAMA command and use the adjusted series in the following analysis. Casual glance at the figure (4.1) and (4.2) of Drr^e's movement for arbitrary observations revealed a jump at around a year, which could have been taking the effect of the annual jump. However, such spikes turned out to be not seasonal, and the estimates hardly changed for the specifications of table 5.6.
6. Our economic theory does not explicitly predict the forms at which each variable enter the hazard. Nonetheless, as far as the time-varying covariates are concerned, it is the changes since the start of the spell, and not their levels, that builds pressure towards a failure once the spell starts, since the hazard is conditioned on the starting event of the spell. Their levels at the start of a spell, however, may carry a different implication altogether as the initial condition of a spell reflecting the era in which the observation took place. It maybe that in the 70's, bargaining groups might have been more predisposed towards annual contract as a norm than they were in the 50's due to factors not captured by already existing explanatory

variables. If so, this is sufficiently represented by a single trend variable that is spell specific. Since our data is a panel, time t is measured from the different origin in calendar time not only for the different bargaining groups but also for the different spells of the same group. In order to avoid difficulties in interpretations as well as for the numerical stability, it seems better to let the secular trend be included in the spell specific variable while x(t) should include variables that carry similar magnitude overtime. In our specification, log of real wage may fall into such category of secular trend. Hence, we either omit such variable altogether or we leave its value at the start of the spell in z_{ns}.

7. Our hazard equation embodies the mechanism of economic pressure that accumulates to trigger the failure, as predicted in the theory of chapter 4. However, a decision to call for a short completed spell is reasonably considered to base on a very different criteria from what our theory depicts. Shorter durations tend to accompany staged settlement, which is already taken into account by the stage dummy. Still, there are many short durations that do not involve staged settlement (there are 19 observations started by a staged settlement and also 19 started by a non-staged settlement that have lasted for less than 15 weeks). In view of this, we merged shorter duration spells with either the preceding or succeeding spells and treated them as a single spell so as to avoid breaking the chain of the multiple spells for each group. The resulting estimates, however, hardly changed for any of the variables.
8. First step heteroscedastic consistent estimates. $Drr^e(\tau^i)$ is based on information available at τ^{i-1} or τ^0.
9.

$$\text{Baseline hazard, } h_0(t;\alpha) = \begin{cases} 1 & t < 48 \\ \alpha + 1 & 48 < t < 72 \\ 1 & t > 72 \end{cases}$$

where α is constrained to be positive by setting $\alpha = \exp(b)$ and iterating the values of b. Estimated value of b was -39.2945.
10. Second step estimates. $Drr^e(\tau^i)$, $Drr^u(\tau^i)$ is based on information set available either at τ^{i-3} or τ^0. Seasonally adjusted RPI used to derive Drr variables.
11. Second step estimates. $Drr^e(\tau^i)$, $Drr^u(\tau^i)$ is based on the information set available at τ^0. Seasonally unadjusted RPI used to derive Drr.
12. First step heteroscedastic consistent estimates. $Drr^u(\tau^i)$ is based on information set available either at τ^{i-3} or τ^0. Seasonally adjusted retail price index.

13.Second step estimates. $Drr^u(\tau^i)$ is based on the information set available at τ^{i-3} or τ^0. Seasonally adjusted RPI used to derive Drr.

14. Second step estimates. $Drr^u(\tau^i)$ is based on the information set available at τ^{i-3} or τ^0. The baseline hazard is constrained to be: $h_0(t)=\alpha t^{\alpha-1}$ with $\alpha=2$. Seasonally adjusted RPI used to derive Drr.

15.Second step estimates.

16.Used as instruments are :

size2 = number of workers negotiating (T_{ns-2})/total employment (T_{ns-2}), and not its difference since the previous negotiation, as is used in the actual hazard equation. Same holds for size52 and size104.

Chapter 6: Simulation and incomes policy extension

6.1 Incomes policy sub-model

Incomes policy is put into effect by a statement made by the government over a description of a policy which consists mainly of two elements: the degree of allowance and enforcement. The first usually imposes a ceiling to the rate of wage changes. They include forms such as wage freezes (i.e., zero pay norm), flat increases, fixed rate increases, increases accordance with productivity, and those induced by the cost of living allowances (COLA). The degree of enforcement varies from mere advisory to statutory, and they can be targeted towards different sectors of the economy such as public, wages council or private.

The method we have been adopting so far captures the effect of individual incomes policy episode by simple on/off dummies. But it is not capable of gauging the effectiveness of the particular degree of norm and/or enforcement on the timing and degree of wage settlements that is consistent throughout the sample period. Practically speaking, even though a policy itself is made up of these two components, their effectiveness is said to largely vary with a political climate at a time and the appropriateness of the monetary and fiscal policies mix simultaneously adopted with the incomes policy. Nonetheless, it is of an interest to find out if there exists any consistent responses in the settled wages to the degree of enforcement and the ceilings set by the government throughout our sample period. Separate dummies can differentiate the effect of separate policies that took place at a certain point in time, but they cannot differentiate the effect of different degrees of allowances and enforcement. In other words, with the dummy formulation, it is not possible to measure the effect of changes in the degree of norms or enforcement, from that of a norm n% to (n-1)%, for instance. If a huge data set were available, one could use separate dummies for every possible degree of policy criteria, although in practice, a number of different combinations would be too large for such an analysis to be feasible. Therefore, instead of the individual policy dummy approach, we propose a competing model with a self contained incomes policy sub function whose arguments include both a degree of pay norm and enforcement so that we can represent any episodes of incomes policy by varying a degree of these arguments. In this way, we can find out if there exists any consistent impact of the norm and enforcement on the timing and magnitude of wage changes in our sample. Also, by adjusting the size of a norm according to the length of a contract just terminated, we can measure the annualized impact of the norm; per unit effect of the norm on the rate of wage changes per unit of time (i.e., 1 year). This analysis, to an extent, enables us to see the combined impact of the ceiling policy, on both timing and magnitude of wage changes. Having such a self-contained sub function capable of expressing any forms of the incomes policy is particularly useful in the context of simulation. For example, it is possible to simulate what would have happened to the wage levels if the policies took place at different times.

This idea was introduced in the work of Pudney and Boyle (1986), in which they incorporated such a sub model into the average earnings equation. A major complication in their macro model involved a procedure to express a sequence of complex pay norms in terms of the approximate percentage increase over each quarter in a sample. Since the macro data they have used was observed quarterly while the policy was announced on a monthly basis, it was necessary to incorporate the policy-off as well as the policy-on segment of any quarter to construct a new pay norm series corresponding to each quarter. With respect to the degree of enforcement towards the public sector, they used a share of employees in such sector out of total number of employees, which always lies in the [0,1] domain. This is a rather arbitrary but sensible proxy, considering that the analysis was done at such an aggregated level. In our case, however, the analysis can be far simpler to implement. We know exactly in which month of the year a wage change of a particular group took place, and the hazard equation takes into account of the monthly observed explanatory variables, into which the incomes policy norm is straightforward to introduce. In addition, the degree of enforcement can also be easily incorporated since our analysis distinguishes if a group unit belongs to the public, wages council or the private sector.

6.1.1 The model: wage equation

Our aim is then to incorporate such a self contained function of incomes policy into the system of timing and size of wage changes and to find out if there is any consistent impact of a degree of pay norm and enforcement on these two variables.

The purpose of the government in setting a norm is to deflect the amount of pay rise away from what would have been without the policy towards a norm stated by the policy. Then, we would be more interested in the impact of the different levels of norms on the amount of wage changes rather than on the probability of negotiation. It is not too clear if the negotiation probability should be affected by whether the norm is 2% or 3%, but rather, whether the norm is zero or nonzero. In addition, complication arises due to the introduction of incomes policy sub function in the hazard equation. Hence, we consider incorporating such sub-function mainly in the context of wage equation.

Consider the predicted rate of wage change without the presence of an incomes policy at calendar time, T_{ns}, denoted as f_{ns}. And compare such value with the target of the government: the norm. For the wage change observed at T_{ns}, we have:

$$f_{ns} = \gamma_1^{*}{}'\xi_{ns}^{*} + \gamma_2{}'C_n \qquad (6\text{-}1\text{-}1)$$

where ξ_{ns}^{*} is the vector of covariates used in the wage equation (5-3-2) without the incomes policy dummies. Then introduce the incomes policy variables, N(t) and $D_n(t)$. N(t) is a time series representing a degree of ceiling in terms of a rate of nominal wage change set out by the government. It is equal to f_{ns} when there is no policy at T_{ns}. $D_n(t)$, on the other hand, also represents a norm but only for the groups towards which such policy is enforced. In this sense, this variables is affixed with a group identifier, n. Hence, this variable equals $N(T_{ns})$ when there is an enforced policy at T_{ns}, and equals f_{ns} when there is no such policy. The wage equation can now be written as:

$$\dot{w}_{ns} = f_{ns} - \phi_1(f_{ns} - N(T_{ns})) - \phi_2(f_{ns} - D_n(T_{ns})) + u_n + v_{ns} \qquad (6\text{-}1\text{-}2)$$

where, at T_{ns};

$$N(T_{ns}) = \begin{cases} N(T_{ns}) & \text{if there exitsts a policy with norm } N(T_{ns}) \\ f_{ns} & \text{otherwise} \end{cases}$$

$$D_n(T_{ns}) = \begin{cases} N(T_{ns}) & \text{if there exists an enforced policy with norm } N(T_{ns}) \\ f_{ns} & \text{otherwise} \end{cases}$$

Hence, in the absence of any incomes policy at T_{ns}, the wage equation becomes:

$$\dot{w}_{ns} = f_{ns} + u_n + v_{ns} \qquad (6\text{-}1\text{-}3)$$

The government's aim of the incomes policy is to draw the rate of wage increase, $\dot{w}_{ns}$, away from f_{ns} towards $N(T_{ns})$. ϕ_1 measures the impact of a policy with a norm $N(T_{ns})$ in the absence of enforcement and ϕ_2 measures the additional effect of an enforced policy. Hence, the total effect of a full incomes policy is

$\phi_1+\phi_2$. In the case of an enforced policy with a norm, $N(T_{ns})$, the equation (6-1-2) can be written as:

$$\dot{w}_{ns} = (1-\phi_1-\phi_2)f_{ns} + (\phi_1+\phi_2)N(T_{ns}) + u_n + v_{ns} \qquad (6\text{-}1\text{-}4)$$

The closer $(\phi_1+\phi_2)$ is towards 1, the more effective the policy. If we assume a priori that a voluntary policy has no effect on the rate of wage changes, we can ignore ϕ_1. On the other hand, if we assume that a degree of enforcement has no impact on the rate of wage changes, $\phi_2=0$. This representation is only a local approximation that assumes a 1% change in a norm, whether it is a change from 5% to 4% or 1% to 0%, to provide the same effect on $\dot{w}_{ns}$. This makes it seem as if the government could lower the value of $\dot{w}_{ns}$ to any level by setting N_{ns} arbitrarily low. More satisfactory specification would have ϕ_1 and ϕ_2 as the increasing functions of N_{ns}, each with an upper asymptote. This means that the more strict the norm, the lower its per unit effectiveness. Nonetheless, we assume here that the government has an incentive not to assign norms unduly low so that an approximation such as (6-1-2) is good enough.

This equation can be differenced to eliminate a need to make any distributional form assumptions for the random group specific effect, u_n, hence:

$$\Delta\dot{w}_{ns} = \Delta f_{ns} - \phi_1(\Delta f_{ns} - \Delta N(T_{ns})) - \phi_2(\Delta f_{ns} - \Delta D_n(T_{ns})) + \Delta v_{ns} \quad (6\text{-}1\text{-}5)$$

This can then be estimated by the GMM using the same set of instruments that was used in the 2SIV estimation with the individual incomes policy dummies (chapter 5). A list of the norms used for creating the series, $N(T_{ns})$, is in Table 6.1 above. We have considered the upper end of their announced range as most relevant to the unions in demanding their wage claim, hence, applicable to $N(T_{ns})$. Amongst all the policies in the sample period, only cd4 was vague on their announcement of the norm, stating only that "target considerably less than

Table 6.1 Incomes policy and the norms

policy	period	enforced sector	norm
fd1	1961 Jul-62 Mar	non-private	0
cd1	1962 Apr-63 Mar	non-private	0.025
cd2	1963 Apr-65 Apr	non-private	0.035
cd3	1965 Apr-66 Jul	non-private	0.035
fd2	1966 Jul-67 Jun	all sectors	0
cd4	1967 Jul-68 Mar	all sectors	0.0175
cd5	1968 Mar-69 Dec	all sectors	0.035
cd6	1970 Man-70 Jun	all sectors	0.045

previous 3-3.5%". We have opted for the mid point between 3.5 and 0, the norms of the adjacent policies.

6.1.2 The model: hazard equation

It is not straightforward to incorporate the level of different norms in the hazard equation since the norm is not explicitly comparable with the hazard but the rate of wage changes. Intuitively speaking, what we are interested in is the impact of a norm on the target wage, the level of wage that workers aspire given the economic conditions at a time. This reflects the imaginary level of wages were the unions to strike a bargain at every month during a contract. Negotiation is triggered whenever a difference between such a target wage and the actual wage exceeds the cost of negotiation. Then, our question is really how the degree of norm affects the level of target wage so as to change a negotiation probability. Since such target wage is only implicit in the hazard equation, it cannot be compared with the norm directly. Nonetheless, considering that the norm does not have too wide a range unless it is a zero norm, it seems reasonable to consider only the effect of zero or nonzero norm towards the negotiation probability rather than try deriving the impact of every different degree of norm. Therefore, we pursue estimation of the hazard equation with two incomes policy dummies representing zero and non-zero norms. In addition, this framework can easily incorporate their interactions with the degree of enforcement. The hazard equation at i-th month during the n,s-th contract which started at calendar time T_{ns-1} becomes:

$$h(\tau^i_{ns}-T_{ns-1}) = \exp(f(\tau^i_{ns}) + ip(\tau^i_{ns}) + u_n + \varepsilon_{ns})\, h_0(\tau^i_{ns}-T_{ns-1}) \qquad (6\text{-}1\text{-}6)$$

where;

$$f(\tau^i_{ns}) = \beta' x^*(\tau^i_{ns}) + \gamma_2' z_{ns} + \gamma_1' c_n$$

$$ip(\tau^i_{ns}) = \mu_1 F(\tau^i_{ns}) + \mu_2 C(\tau^i_{ns}) + \mu_3 E(\tau^i_{ns}) F(\tau^i_{ns}) + \mu_4 E(\tau^i_{ns}) C(\tau^i_{ns})$$

$$F(\tau^i_{ns}) = \begin{cases} 1 & \text{if there exists a freeze policy at } \tau^i_{ns} \\ 0 & \text{otherwise} \end{cases}$$

$$C(\tau^i_{ns}) = \begin{cases} 1 & \text{if there exists a ceiling policy at } \tau^i_{ns} \\ 0 & \text{otherwise} \end{cases}$$

$$E(\tau^i_{ns}) = \begin{cases} 1 & \text{if a policy at } \tau^i_{ns} \text{ is enforced} \\ 0 & \text{otherwise} \end{cases}$$

$x^*(\tau_{ns})$ is the vector of within-spell varying covariates used in the hazard equation (5-3-1) without the incomes policy dummies. μ_1 measures the impact of a non-enforced ceiling policy on the negotiation probability while $\mu_1+\mu_3$ measures their effect when fully enforced. Likewise, μ_2 measures the impact of a non-enforced freeze policy and $\mu_2+\mu_4$ is that of an enforced freeze policy.

6.1.3 Estimation : wage equation

The differenced equation (6-1-5) is estimated by GMM. The second step estimates and t-ratios based on their heteroscedasticity consistent standard errors are listed in Table 6.2, column (1)-(3). The second step estimators use the estimated variance covariance matrix based on the first step estimator as a weighting matrix in the objective function to be minimized.

Real and relative wages, a wages council dummy and productivity variables are again found to be the main determinants. Significant role played by the number of negotiators, SIZE, and the last consecutive real wage changes, DRW1, in the 2SIVE (Table 5.8, 5.9) are not observed here. Also, there seems to be a finding consistent with the hysteresis evidence where the global unemployment exerts a negative excess supply effect while the industry unemployment gives a positive effect. Opposite finding was seen in table 5.8 and 5.9 although none of them were significant. Staged, pre-determined wage changes are exerting a stronger negative impact here than in table 5.8, although they are still not precisely estimated. Moreover, there seems to be a correlation between ϕ_1 and ϕ_2, which causes ϕ_2 to be insignificant. This is possibly due to: (1) a lack of variation in the group characteristic variables which failed to be identified, and (2) a lack of variation between $N(T_{ns})$ and $D_n(T_{ns})$. But the same regression without ϕ_1, ignoring the effect of a voluntary policy, also revealed insignificantly negative ϕ_2. Moreover, the regression with $\phi_2=0$ suggests a very strong influence of ϕ_1 as listed in column (2). We have also run the same regression while fixing the parameters of wages council and constant term. Here, too, ϕ_2 is found insignificant (Table 6.2, column (3)).

These results so far suggest a definite impact of the incomes policy, in particular, that of the non-enforced policy. Additional impact of the enforcement seems negligible and insignificant. Notably, the effect of having a wage changes during the non-enforced wage freeze policy as compared to having it during the no-policy regime is a reduction of as much as 41-46 percent in the

Table 6.2 Estimates of the wage equation
(unskilled males; manufacturing and construction ind)

Variable	(1)	(2)	(3)	(4)
spell specific effects				
size(T_{ns})	27.716	30.587	26.753	34.293
	(0.69)	(0.86)	(0.75)	(1.59)
lnrw(T_{ns})	-0.197	-0.202	-0.198	-0.201
	(2.96)	(3.13)	(21.0)	(32.1)
Drw1	0.003	-0.007	0.002	-0.007
	(0.10)	(0.25)	(0.06)	(0.28)
rel(T_{ns})	0.227	0.233	0.229	0.198
	(8.77)	(11.12)	(9.81)	(11.1)
lnun(T_{ns})	-0.053	-0.057	-0.055	-0.025
	(1.85)	(2.22)	(2.86)	(1.64)
lnunin(T_{ns})	0.038	0.041	0.040	0.023
	(1.83)	(2.16)	(3.02)	(1.96)
prod(T_{ns})	0.169	0.152	0.172	0.098
	(2.23)	(2.12)	(2.50)	(1.98)
staged$_{ns}$	-0.013	-0.013	-0.013	-0.013
	(1.21)	(1.16)	(1.36)	(1.45)
sfix$_{ns}$	-0.047	-0.051	-0.049	-0.021
	(1.95)	(2.93)	(1.96)	(1.95)
d12(T_{ns})	0.014	0.012	0.014	0.017
	(1.81)	(1.99)	(2.22)	(3.85)
f_{ns}-N(T_{ns})	0.417	0.464	0.428	0.141
	(2.03)	(5.23)	(2.19)	(1.88)
f_{ns}-D_n(T_{ns})	0.036		0.030	0.133
	(0.30)		(0.25)	(3.39)
group-specific effects				
constant	-1.238	-1.262	-1.240	-1.240
	(2.45)	(2.58)		
W Council	-0.249	-0.256	-0.249	-0.249
	(4.38)	(4.75)		
Sargan test	35.11	36.23	35.33	35.26
(df)	(26)	(27)	(28)	(28)
(p-value)	(0.11)	(0.11)	(0.16)	(0.16)

prevailing rate of wage increase, which seems rather large. An additional effect of the enforcement is shown to be roughly 3 percent, making the impact of full incomes policy freeze to be around 44-49 percent. This result also implies that a 1% point decrease in the non-enforced norm induces 0.44% point reduction in

$\dot{w}_{ns}$. Recall the wage equation with additive incomes policy dummies in chapter 5:

$$\dot{w}_{ns} = f_{ns} + \gamma_1 fd1(T_{ns}) + \gamma_2 cd1(T_{ns}) + \gamma_3 cd2(T_{ns}) + \gamma_4 cd3(T_{ns}) + \gamma_5 fd2(T_{ns}) + \gamma_6 cd4(T_{ns}) + \gamma_7 cd5(T_{ns}) + \gamma_8 cd6(T_{ns}) + u_n + v_{ns} \qquad (6\text{-}1\text{-}7)$$

Unfortunately, this model is not nested in the nonlinear model of the equation (6-1-2), which makes direct comparison of their coefficient estimates impossible. A major difference lies in the fact that the coefficients, γ's, of the individual policy dummies represent the combined effects of per unit change in the norm and the deviation between the target norm and the prevailing wage at a time, which the government wants to reduce to zero. In addition, such effect of per unit change of a norm is allowed to vary from one policy to another. On the other hand, ϕ's in the incomes policy sub-function model represent the effectiveness of per unit change in the norm where such impact is assumed to be fixed across the policy regimes. Nonetheless, we can make a rough comparison of their implied incomes policy effects on wages by calculating the rate of change in $\dot{w}_{ns}$ between the policy-off and policy-on regimes. Computing the rate of change of $\dot{w}_{ns}$ makes this comparison as robust as possible to the differences in f_{ns} predicted by (6-1-2) and (6-1-7). For example, we can compare $\hat{\gamma}_1/\overline{(\hat{f}_n+\hat{u}_n)}$ from (6-1-7) for fd1 with $-(\hat{\phi}_1+\hat{\phi}_2)\overline{\hat{f}}_{ns}/\overline{(\hat{f}_n+\hat{u}_n)}$, where $\overline{\hat{f}}_{ns}$ or $\overline{(\hat{f}_n+\hat{u}_n)}$ is the sample average of estimates over observations that took place during the relevant policy regime, using the coefficient estimates based on (6-1-7) and (6-1-2), respectively. Likewise, for cd1, we can compare $\hat{\gamma}_2/\overline{(\hat{f}_n+\hat{u}_n)}$ from (6-1-7) with $-(\overline{\hat{f}}_n-0.025)(\hat{\phi}_1+\hat{\phi}_2)/\overline{(\hat{f}_n+\hat{u}_n)}$ from (6-1-2), and so on. In doing this comparison, a set of explanatory variables that are used to predict f_{ns} are made identical to each other, to Table 6.2. If per unit effect of the norm is constant throughout the sample period, they should roughly be the same. Note that this is not a formal test. If one wants to test the invariance across regimes of per unit impact of the norm, one has to build a model that allows for a separate impact of the norm per unit for different policies, in which equation (6-1-2) nests. A set of linear restrictions can then be tested on such a model, but a number of parameters will be too large for it to be feasible.

For now, look at table 6.3 and make a rough comparison between the implied effectiveness of each policy derived from $\overline{f}_{ns}$'s and $\hat{\phi}$'s from equation (6-1-2) and $\overline{f}_{ns}$'s and $\hat{\gamma}$'s from (6-1-7). For example, the effect of fd1 in non-private groups is -0.623 in column (2) which is comparable to -0.597 in column (7), since the fd1 dummy is only made applicable towards the enforced groups. We can make similar comparisons for other policies by looking at column (2) and (7). These comparisons suggest that the first four policies appear to have as large an effect towards non-enforced sector as they do towards the enforced ones. In other words, the policy seem to have no additional enforcement effect over the non-private groups during cd1, cd2 and cd3. In general, earlier policies seem to have a stronger per unit impact than the latter as can be seen in column (7), which failed to be captured by the parameters ϕ_1 and ϕ_2 since they assume per unit effectiveness of a norm to be invariant across the policies. This divergence, in addition to the fact that the estimates $\hat{\phi}_1$ and $\hat{\phi}_2$ are not as precisely estimated as their individual dummy counterparts, may suggest that factors other than the norm and enforcement are important in determining the effectiveness of the incomes policy. A better comparison can be done by looking at the dynamically simulated wage levels based on these specifications, results of which will be reported later.

Introduction of the incomes policy sub-function in the wage equation enables us to see the impact of norms on the *timing* of negotiations in a way that is not possible when the individual policy dummies are used. So far, we have seen the impact of the norms set by the government on the rate of wage changes without taking into account how long the previous contract has lasted. The norm stated by the government usually refers to the rate of wage changes over a course of one year. This is clear when a policy states a norm in conjunction with a requirement that at least 12 months should separate the settlements. However, whether workers assume such norm to be an entitlement per year or per settlement is not clearly known. If they consider it as an entitlement per settlement,

Table 6.3 Comparison of individual policy effect

(1) policy	(2) enforced	(3) non-enf	(4) nob	(5) $\widehat{f_{ns}+u_n}$ (6-1-2)	(6) $\widehat{f_{ns}+u_n}$ (6-1-7)	(7) $\gamma/(\widehat{f_{ns}+u_n})$	(8) enforced	(9) non-enf
fd1	-0.623	-0.582	20	0.068	0.062	-0.597	-0.785	-0.404
cd1	-0.203	-0.189	53	0.055	0.061	-1.059	-0.562	-0.292
cd2	-0.267	-0.250	106	0.069	0.075	-0.826	-0.515	-0.282
cd3	-0.338	-0.316	90	0.082	0.080	-0.653	-0.552	-0.296
fd2	-0.211		29	0.051	0.076	-0.486	-0.569	
cd4	-0.486		44	0.121	0.132	-0.439	-0.524	
cd5	-0.363		92	0.120	0.137	-0.578	-0.434	
cd6	-0.351		30	0.150	0.146	-0.312	-0.397	

except for the zero-norm, a higher wage can be achieved by having more frequent settlements while still abiding by the policy, as long as the outcome of each settlement is below the norm. If this is true for *all* the ceiling policies, we should be able to see the difference in their impact once we take into account of the contract length which has just terminated. For this purpose, equation (6-1-2) is simply reformulated as:

$$\dot{w}_{ns} = f_{ns} - \phi_1{*}(f_{ns} - N(T_{ns})\frac{\delta_{ns}}{52}) - \phi_2{*}(f_{ns} - D_n(T_{ns})\frac{\delta_{ns}}{52}) + u_n + v_{ns} \qquad (6\text{-}1\text{-}7)$$

In (6-1-2), ϕ_1 represents the impact of a norm, $N(T_{ns})$ on $\dot{w}_{ns}$ at each settlement while $\phi_1{*}$ in (6-1-7) represents its effect per annum. Hence, if the norms are abided by at each settlement but it occurs more frequently than once a year, ϕ_1 should become larger than $\phi_1{*}$, consequently making equation (6-1-2) look as if a policy is effective.

Look at column (4) of Table 6.2 (this is comparable to the result of column (3)) and also, column (8) and (9) of table 6.3 which are all based on equation (6-1-7) above. We find distinct changes, first of all, in the relative significance of ϕ_1 and ϕ_2, then, in their magnitude. The total impact of the enforced policy on $\dot{w}_{ns}$ has declined, which supports our prediction that the policies are, after all, not as effective in reducing $\dot{w}_{ns}$ per annum as it seemed they are in column (3). It shows a reduction of 0.27% points in $\dot{w}$ per annum for every 1% point decrease in the enforced policy. Even though the total impact of the full policy has declined, the additional enforcement effect has increased and is now strongly significant. Impact of the policy per annum more or less doubles when it is an enforced one, as is also seen from column (8) and (9) of table 6.3. On the other hand, the impact of non-enforced norms has declined. This indicates that the reduction in $\dot{w}_{ns}$ per settlement is partly compensated by having $\dot{w}_{ns}$ more frequently while abiding by the policy norm at each settlement, at least among the non-enforced private sectors. This is indeed the case when one looks at the average rate of wage changes occurred and the average contract length terminated in each quarter over the period plotted in figure 6.1 and 6.2. During cd1, cd2 and cd3, non-private groups had larger rate of wage changes on average than the private groups, and during cd2 and cd3, they had a longer contract length terminating than the private groups on average. While the policies still had some impact, private groups who experienced non-enforced cd1 cd2 and cd3 policies had, on average, more frequent settlements with lower $\dot{w}$, hence $\phi_1^* < \phi_1$. On the other hand, non-private groups, towards whom the first 3 ceiling policies were enforced, had terminated longer contracts, longer than one year, with higher $\dot{w}$ on average, hence, $\phi_2^* > \phi_2$. The average length of contract terminated during cd1, cd2 and cd3 are 66.67, 71.29 and 52.84 for the non-private groups and 63.9, 58.36 and 41.89 for all the groups. When equation (6-1-7) is estimated without fixing the coefficients of wages council and constant terms, these variables in addition to ϕ_1, ϕ_2 failed to be identified (very large coefficients with large standard errors). This is probably due to strong collinearity between ϕ_1, ϕ_2 and wages council dummy since the variations originally found in f's were smoothed by the introduction of contract length.

This finding is consistent with the finding from the hazard equation estimation (table 5.6) where the ceiling policies cd1, cd2 and cd3 are found to have comparatively large positive effect (as much as 242% increase for the fd1) towards enforced sectors on the average duration.

6.1.4 Estimation: hazard equation

Estimates of the hazard equation specification (6-1-6) has revealed that there is a significant impact of the voluntary freeze policy in reducing the hazard rate (column (1) table 6.4). Additional impact of the enforced freeze policy is in fact significantly positive, although the net effect is still negative. The ceiling policy is insignificant when not enforced, but becomes significantly effective in reducing the hazard when enforced. In addition, being a wages council group significantly prolongs, while a public sector group significantly reduces the contract length.

Figure 3.24 shows that the number of negotiations occurred each quarter during the sample period is dramatically reduced during the freeze policies. The average wage changes observed during the corresponding period, however, provides a mixed message: during fd1, it actually rises, while during fd2 it is sharply reduced. Also, according to the analysis in the previous section that estimates the annualized impact of norms for the ceiling policies, we find that during the first three ceiling policies, the non-private sectors had particularly infrequent wage changes while that was not the case during the latter ceiling polices enforced to all sectors. This may suggest that there is a difference in the impact of enforcement on the contract length depending on the parties towards which the policy is enforced. In order to separate their effect, we redefine $E(\tau_{ns}^i)$ to represent all-sector enforcement and add another dummy variable, $PE(\tau_{ns}^i)$, for the non-private sector enforcement, so that:

$$E(\tau^i_{ns}) = \begin{cases} 1 & \text{if a poilcy at } \tau^i_{ns} \text{ is enforced in all sectors} \\ 0 & \text{otherwise} \end{cases}$$

$$EP(\tau^i_{ns}) = \begin{cases} 1 & \text{if a poilcy at } \tau^i_{ns} \text{ is enforced in non-private sectors} \\ 0 & \text{otherwise} \end{cases}$$

Hence, the incomes policy effect in the hazard equation is now represented by:

$$\begin{aligned} ip(\tau^i_{ns}) = \mu_1 F(\tau^i_{ns}) + \mu_2 C(\tau^i_{ns}) + \mu_3 E(\tau^i_{ns})F(\tau^i_{ns}) + \mu_4 E(\tau^i_{ns})C(\tau^i_{ns}) \\ + \mu_5 EP(\tau^i_{ns})F(\tau^i_{ns}) + \mu_6 EP(\tau^i_{ns})C(\tau^i_{ns}) \end{aligned}$$

Then, μ_1 provides the impact of a non-enforced freeze policy, $\mu_1+\mu_3$ gives the total impact of a freeze policy enforced to all sectors, likewise, $\mu_1+\mu_5$ represents the impact of a freeze policy enforced only to the non-private sectors (including wages council). As can be seen from column (2) of Table 6.4, there is a strong effect of the non-enforced freeze policy in reducing the hazard. Additional effects of enforcement are positive for both fd1 and fd2, but their full effects are still negative and about the same size. Enforcement of the ceiling policy significantly reduces the probability to negotiate. In particular, ceiling policies enforced towards non-private sectors are found to have stronger restraint on the occurrence of wage changes than those enforced to all sectors, which is a consistent finding to our result in the former section; during cd1, cd2 and cd3, the non-private sector groups had a longer contract length on average.

6.1.5 Conclusion

An introduction of the non-linear incomes policy sub-function has turned out to be of little improvement over the individual dummy variable approach in terms of their significance. Nevertheless, it was possible to measure the impact of incomes policy that are different from those derived from the policy on/off dummies; the effect of different degrees of norms and enforcement on both the timing and the size of wage settlements.

In the hazard equation, we found that the non-enforced freeze policy plays an important role in reducing the hazard. Negative impact of the ceiling policy is severed with enforcement. In particular, the ceiling policies, cd1, cd2 and cd3 had the strongest and significant impact. They increase the contract length on average by as much as 75%. In the wage equation, a particularly interesting

Table 6.4 Estimates of the hazard equation[1]
(unskilled males; manufacturing and construction ind)

Variable	(1)	(2)
spell specific effects		
$\ln(\delta_{ns-1})$	0.062	0.060
	(2.75)	(2.50)
Drw1	-2.816	-2.667
	(7.15)	(5.78)
$staged_{ns-1}$	0.461	0.461
	(4.61)	(4.30)
$dfix_{ns-1}$	-0.442	-0.478
	(5.25)	(5.34)
$Drr^{u}(t)$[2]	-6.002	-5.376
	(5.56)	(5.05)
lnunin(t)	-0.426	-0.441
	(7.68)	(7.61)
lnprin(t)	0.152	0.056
	(0.60)	(0.21)
size(t)	3.721	2.993
	(1.01)	(0.74)
rel(t)	0.887	1.080
	(1.98)	(2.36)
Freeze(t)	-2.042	-1.811
	(3.39)	(3.59)
Ceiling(t)	-0.108	-0.103
	(1.34)	(1.09)
EFreeze(t)	1.507	1.172
	(2.46)	(2.33)
ECeiling(t)	-0.383	-0.346
	(4.51)	(3.51)
EPFreeze(t)		1.120
		(1.86)
EPCeiling(t)		-0.648
		(3.94)
group-specific effects		
constant	-5.195	-4.969
	(42.6)	(39.8)
W Council	-0.233	-0.220
	(3.07)	(2.67)
Public	0.312	0.406
	(5.28)	(6.59)
TU	0.034	0.038
	(0.86)	(0.88)
Sargan test	41.67	39.72
(df)	(40)	(38)
(p-value)	(0.40)	(0.39)

finding was the dramatic shift in the relative effect between enforcement and non-enforcement policies when we considered a norm as a limit per annum (i.e., adjusted the norm according to the elapsed contract length). When the effectiveness of the policy was determined on the basis of the deviation between expected $\dot{w}$ and the norm, it seemed that the policy had a huge effect on the non-enforced groups with hardly any additional effect towards the enforced groups. However, when the impact was based on the measure between the expected $\dot{w}$ and the norm adjusted for the length of the contract just terminated, the effect on the non-enforced groups has largely declined while the additional enforcement effect has increased. A 1% point decrease in the non-enforced norm was found to reduce the rate of wage increase per annum by around 0.14% with an additional 0.13% reduction when enforced. This implies that during the cd1, cd2 and cd3, non enforced groups enjoyed relatively frequent wage changes, thus making it seem as if they were abiding by the policy. And this is totally consistent with our finding from the hazard equation estimates.

Nonetheless, comparing roughly the implied impact of each policy between the individual dummy and the policy sub-function formulation implies the effectiveness of the norm and enforcement that is not quite consistent throughout our sample period. Moreover, the parameters in the sub-function were not as precisely estimated as the individual policy dummies. These results suggest one point: factors other than the norm and enforcement are important in determining the effectiveness of each incomes policy. Hence, varying the elements of incomes policy tool does not necessarily change the degree of effectiveness, but rather, its timing - the political climate, economic condition not captured in our models, the appropriateness and the consistency of the other fiscal, monetary and exchange rate policies adopted at the time - seems to play a big role. A proper comparison of these models have to be made by dynamic simulations, to which we now turn.

6.2 Dynamic Simulation

Given the estimated coefficients of the hazard and the wage equations we have analysed so far, it is of an interest to investigate more fully the dynamic impact of external influences on this system, particularly of incomes policies, by means of a dynamic simulation technique. This enables us to clearly look at two different effects of policies on wage changes: delay and moderation. Moreover, based of these results, we can make comparisons amongst the various specifications we have considered for this system.

In a usual regression model, dependent variable can be simulated using the values of exogenous variables from the data and the lagged endogenous variables substituted for the lagged simulated values. The hazard equation, however, implies a probability distribution of the contract length but not the actual point at which a contract terminates. Since the hazard depends on the time varying explanatory variables in our formulation, there is no unique contract length that can be determined from such a duration distribution function. The problem is that the rate of change of wages cannot be simulated unless the timing of the settlement is known.

In view of this, we have considered the discrete approximation of the process where the decision to terminate the spell is made each month according to the probability given by the hazard equation. More specifically, we pick a starting date, say T_0, for an arbitrarily chosen observation from the data and derive the estimated hazard for the first month given the group's characteristics and other spell specific factors known at the start of this particular contract, $\hat{h}(1)$. Having derived $\hat{h}(1)$, it is straightforward to compute the corresponding survivor function, $\hat{S}(1)$. Then, at T_0+(1 month), a Bernoulli trial takes place whose probability to exit between T_0 and T_0+1 is $\text{Prob}(0 \leq \delta \leq 1 \mid \delta \geq 0)$, which is equal to $(\hat{S}(0)-\hat{S}(1))/\hat{S}(0)=(\hat{S}(0)-\hat{S}(1))$. The instantaneous exit rate at particular duration, δ, is equal to $h(\delta)$. Since we are approximating the truly continuous decision sequence to take place monthly, a discrete probability of exiting at the δ-th month given that the last decision was made at the (δ-1)-th month has to include the possibility of exiting sometime between (δ-1) and δ, which is approximately equal to the sum of the hazard over a one-month interval. If the "exit" is chosen, this contract is terminated with a one-month duration at T_0+(1 month), calendar time. At this point, the agreed wage is computed and the spell specific factors are updated for the new contract starting at T_0+(1 month). If "non-exit" is chosen instead, this contract continues into the second month. In either case, the next step is to consider a similar choice problem at T_0+(2 months), then at T_0+(3 months), and so on. As a result, we will have the simulated monthly hazard and the level of wages prevailing at each months for as long as there are data of exogenous variables. This computation can be done repeatedly, and consequently their averaged values can be compared with the similarly simulated values with the incomes policy coefficients set to zero – a simulation of what would have happened in the absence of the incomes policy. Dynamic simulation of the hazard generates the timing of negotiation, on the basis of which the wages to be claimed is determined. The effect of incomes policy is not only to reduce the level of the wage claim, but also to alter the timing of negotiations which in turn affect the amount of wage changes. Hence, a dynamic simulation of the wage equation taking the timing of wage changes as fixed would not provide a fully satisfactory analysis of the impact of incomes policy, since it fails to incorporate their effect on the magnitude of wage changes as a result of their altering the timing of settlements.

In this section, we first dynamically simulate the wage changes keeping the timing of settlements fixed at their observed values. Then we will derive the average wage levels by simulating both the timing as well as the amount of wage changes using the procedure described above. In so doing, our aim is to divide the effect of incomes policies on wage changes into two separate effects: delay and moderation.

6.2.1 Dynamic Simulation: wage equation

Here, we simulate the average simulated wage level keeping the timing of wage changes as observed in our data. Due to the nature of the data, the starting and the ending period differ across groups. We use the settlement dates observed during the incomes policy period, namely, between 1961 and 70.

During the course of this simulation, we are keeping the price level as well as the aggregate average earnings level fixed at their historical values. In a full model, we should allow prices to respond to the general variation of wage level in the form of a dynamic price equation. However, it is not clear how the variation in the negotiated wages of a single bargaining unit affects the general wage level. As long as we conduct our simulation conditional on the historical relation between prices and aggregate earnings, we consider it reasonable to assume that the negotiated (i.e., simulated) wage rise applicable to a single bargaining group is not significant in influencing the general earnings level, and hence the general price level.

Figure 6.3 plots the simulated average wage levels based on the estimates listed in table 5.9 which includes individual policy dummies. The policy-on locus of simulated values has a very close correspondence with the actual prevailing average wage level. These wage levels are derived from the rate of change of wage equation where the simulated values are substituted for the relevant explanatory variables such as DRW1, Rel and Lnrw. In this sense, they are not the estimated wage levels but the dynamically derived values. Hence, there always is a possibility of cumulative errors building up to move the simulated values away from the actual values. Comparison of the simulated average wage levels with and without the incomes policies clearly shows that the wage level without the policy is never below the policy-on level. The difference between policy on and off regimes gradually widens as they enter the late 60's, and it is at its widest during the policy cd5 and cd6. In practice, there was an enforced freeze policy starting in November 72 and a vague voluntary policy was announced between the times but were not included in our wage equation. According to figure 6.3, the differences in their simulated wage levels narrows substantially by the end of 71.

We have done a similar dynamic simulation based on the estimates of wage equation with the incomes policy sub function analysed in section 6.1. Figure 6.4 is based on wage equation (6-1-2), without any adjustment for the elapsed durations, and figure 6.5 is based on equation (6-1-8) that has a level of norm adjusted for the elapsed contract length. In both cases, policy-off wage levels are simulated equally well and they never fall below the policy-on levels. The gap in the wage levels between them also narrows by 71' but not as much as it does in figure 6.3. The strong impact of cd5 and cd6 found in figure 6.3 is more clearly seen in figure 6.5, which incorporates the adjusted policy sub-function, but their effect in pulling down the hazard is simulated slightly more strongly than their actual effect.

6.2.2 Dynamic Simulation: hazard and wage equation

We have argued that the dynamic simulation of the wage equation, while treating their timings as those observed in the data, is not a satisfactory way to measure the impact of the incomes policy. This is indeed true since these policies not only affect the level of wages directly, but also the timing of wage changes. Hence, a certain incomes policy may be able to reduce the frequency of wage changes, and by so doing, pull the wage levels below what would have been without the policy. In this section, we use the hazard as well as the rate of wage change equation to dynamically simulate average wage levels during the incomes policy period, between 1961 and 71.

The average over 50 replications of the simulated wage levels and the monthly conditional exit rate (i.e., the hazard summed over a monthly interval, called, monthly hazard hereafter)[3] based on the estimates of table 5.5 and 5.9, for an arbitrary chosen group starting from an arbitrary observation point, are plotted in figure 6.6-6.7, 6.10-6.11, 6.14-6.15 and those based of table 5.6 and 5.9 are in figure 6.8-6.9, 6.12-6.13, 6.16-6.17. For the sake of comparison, the groups and the timings of their starting points used for these simulations include those used to derive the estimated hazard in figure 5.28-5.29 and 5.34-

5.35 in chapter 5. Even though differences exist between different specifications of the hazard, these figures clearly suggest the significant effect of the incomes policies in reducing the wage levels, as well as the hazard rate in general. Note that their movement over time depends on the particular group and the timing chosen as a starting point of the simulation.

Specifically, figures 6.6-6.9 are for a wages council group with a starting period of May 1961. The first tick on the X-axis corresponds to June 61 and they are ticked monthly thereafter. Since the group belongs to the wages council, it is exposed to the enforced incomes policy fd1, cd1 and cd2, period of which are indicated below the graphs. There is a mixed evidence on the level of the monthly hazard during the beginning of fd1 policy. A common feature is an acute decline in the hazard as soon as it enters the fd1 policy. In figure 6.6, it is not certain what makes the policy-off regime of the simulated hazard to drop suddenly whose timing coincides with the introduction of fd1. Thereafter, the probability to negotiate as well as the wage level is lower under the policy-on regime. The simulation of figure 6.10-6.13, on the other hand, starts in May 1968 of the same wages council group. This simulation period covers the ceiling policies, cd5 and cd6, which are enforced to all sectors. They suggest a significant impact of these policies in reducing the conditional exit probability. In particular, as soon as cd6 is terminated, we find a discrete jump in the hazard creating more frequent wage changes leading to sharper rise in wages. The absence of further policies after cd6 seems to move the wage level eventually back to its original path that experienced no incomes policies.

Figures 6.14-6.17 are for a private sector group with a starting date of December 1965. During this period, this group is exposed to the enforced fd2, cd4 and cd5 policies. cd3 is not enforced towards this group and our policy dummy is allowed only for the non-private groups. The freeze policy, fd2, has an immediate effect in reducing the hazard and keeping it lower while virtually halting further pay rises. As soon as fd2 is over, the simulated monthly hazard increases to bring about more frequent wage changes throughout cd4. But the introduction of cd5 again pulls the hazard down, which subsequently lowers the rate of wage rise. In the data, number of negotiations acutely increased right after the fd2 and reached a peak in the middle of cd4 policy. Consequently, the simulated wage level is constantly lower as long as the enforced policies are in effect.

In order to see the effect of the non-enforced policies, we have done another simulation on the same private sector group starting in August 1960. So hence it will experience the non-enforced fd1, cd1 and cd2 policies. Since the model with the individual policy dummies does not allow the non enforced policies to have any impact, this time, we have used the estimates of table 6.4 column(2), which has 6 dummies to differentiate the effect of incomes policy depending on the degree and coverage of the enforcement. For the wage equation, we continue using the result of table 5.9. They are plotted in figure 6.18-6.19. Under this specification, fd1 is found to be more effective in reducing the hazard towards the non-enforced private sector groups. While cd1, cd2 and cd3 were particularly effective towards enforced non-private sector groups. The figures show a sudden decrease in the hazard during the non-enforced freeze policy, fd1. As a result, the simulated wage rise slows down as it enters fd1 and its level stays almost invariant during the freeze. After fd1, we see hardly any effect of the non enforced ceiling policies on the hazard, nonetheless, the wage level continues to be lower than the policy-off level. As we have seen from table 6.4, these figures suggest the significance of the non-enforced policies, particularly of fd1, on both the magnitude and the timing of wage changes in the private sector.

So far, all of these figures (6.6 - 6.19) have shown a significant impact of the incomes policies. It is not possible, however, to gauge the effect of the incomes policies on the *average* wage levels across groups over the entire sample period from them. These figures depend on their initial condition, the choice of the starting date and the movement of x(t) during the period covered by the simulation, all of which are unique to each. In order to capture the effect of incomes policies on the average level of negotiation probability and wages throughout the period during which the incomes policies were in effect, we need to have a dynamic simulation analysis done over the entire sample groups that actually experienced wage rises during such period. For this purpose, we have rearranged the simulation so that it calculates the averages of 30 replications for each group that had a negotiation during the beginning of 1961. Conditional on the initial observation point and the corresponding spell specific (z_{ns}, although Staged and Dfix are set to zero) and group specific factors (including bargaining system employed and the group specific random factor estimates, $\hat{u}_n$), it simulates the monthly hazard rate and the wage level month by month for a 100-month period. Consequently, we are able to derive the average simulated conditional negotiation probability and the wage levels between mid 1961 to 1971. In this way, this simulation captures not only the direct impact but also the indirect impact of the incomes policies, through their effect on the timing of wage changes, on the level of wages. We do this dynamic simulation on the basis of the estimates reported in table 5.5, 5.6, 5.7, and 7.4, for the sake of comparison. Table 5.7 restricts the duration distribution to follow Weibull with parameter 2. Table 7.4 assumes the Weibull duration distribution with no time-varying covariates and no heterogeneity and is estimated by the Maximum Likelihood. The average simulated wage levels are compared with the simulated wage levels of the former section derived with the timing of negotiations fixed as in the data. By so doing, we hope we can differentiate two kinds of impact, delay and moderation, on the average wage levels.

Figure 6.21 depicts the average wage level derived from the dynamic simulation of the timing as well as the rate of wage change based on the estimates of tables 5.5 and 5.9. Incomes policy consistently and significantly reduces the level of wages. However, this specification yields unrealistically low conditional exit rates (figure 6.20). As a consequence, simulated wage level is averaged out to be much lower than the actual level.

Figure 6.23, on the other hand, depicts the simulated average wage level based on table 5.6 and 5.9, and the result is far more reasonable than that of figure 6.21. Under this specification, the incomes policy consistently reduces the level of wages, and in general, reduces the conditional exit probability (figure 6.22). In particular, the fd2 freeze policy has a severe restraining effect on the negotiation probability, during which the wage rises are also suppressed. As they enter the following ceiling policy, cd4, the probability of negotiation acutely increases, which raises the rate of wage changes. However, such rate of wage rise is quickly restrained as it enters the next ceiling policy that pulls the negotiation probability down. The effect of cd5 and cd6 in reducing the frequency of wage changes, therefore, lowering the rate of pay rise seems too strongly simulated. The actual wage level increased much faster during the 70's. Nonetheless, the monthly hazard shoots up as soon as the last ceiling policy terminates, in June 70, and the presence of no further incomes policy makes the policy-on wage level to almost catch up with the non-policy locus by the early 71.

Figure 6.25 is the similarly simulated wage levels using the hazard specification of table 6.4. This formulation of the hazard allows for the separate effect of non-enforced policies, in which we found a significant negative impact of fd1 towards the private sector groups. Accordingly, the simulated conditional exit probability (figure 6.24) now clearly suggests a significant decrease during fd1, and less pronounced effect of all other policies. However, the resulting simulated policy-on wage locus are hardly different from figure 6.23.

The hazard specification of table 5.6 or table 6.4 assumes a constant hazard were it not for the movement of x(t) during the spell. In other words, its baseline hazard has the Weibull duration distribution with α constrained to be 1 a priori. The same dynamic simulation based on the estimates of the Weibull hazard specification with a constraint, $\alpha=2$, (table 5.7) are listed in figure 6.27-28. Note that for a given value of α, the comparative static effect on the log duration can always be deduced from the coefficient estimates, which between $\alpha=1$ and $\alpha=2$ are found broadly similar with an exception of the relative wage effect. The simulated conditional exit probability is markedly different during the fd1, but the overall movement from then onwards is similar. More importantly, the resulting simulated wage level is very close, including the under-predicted rate of wage increase in the 70's (figure 6.26). These evidence implies that the set of explanatory variables incorporated in the hazard specification of table 5.6, even though it failed to allow for a pure duration dependence, adequately explains the duration dependence: the way the implied negotiation probability evolves over the sample period.

We have also conducted a dynamic simulation on the basis of the maximum likelihood estimates that assumes no heterogeneity and no time varying covariates. Hence, once the spell starts, a timing of the next negotiation is determined purely by the passage of time, in particular, according to the Weibull duration distribution. Even though the ML estimation lacks important factors such as the group random effect or any time varying covariates other than the elapsed duration itself, their simulated wage level has succeeded in capturing the faster growth in the 70's, the phenomena the former simulation failed to capture. Since there is no time varying covariates, the dynamic simulation can also be done by generating the values of completed durations for given values of the start-of-a-spell conditions, rather than repeating a Bernoulli trial month by month. Nonetheless, for the sake of comparison with the other specifications, we have proceeded to do this simulation also according to the Bernoulli trials, so that the average simulated conditional exit probability will be computed. Figure 6.29 is the simulated average monthly hazard based on the MLE. We find hardly any effect of the incomes polices, only a small increase during fd1 and a slight decrease during cd4, both of which are findings different from any of the previous figures. This specification only predicts the effect of incomes policy status at the start of a spell, and does not take into account the length of such policy, nor any changes in the policy regime during the contract spell. In this sense, their incomes policy impact are not instantaneous. As a result, there are no clear policy effect seen in the simulated values of the monthly hazard. On the other hand, in the dynamic hazard specification, such as table 5.6 or 6.4, fd1 is found to have a significantly negative effect and cd4, insignificantly negative effect. From the data, we have seen the number of negotiations go down acutely during fd1 and decrease after the initial increase during the cd4 policy. Since the coefficient estimates in the dynamic specification represent the instantaneous effects of the policies, it makes sense to find their effect negative for fd1 and ambiguous for cd4. Indeed, they are reflected on the simulated monthly hazard in ways that are consistent with the data. These are the examples of how incorrectly the static hazard formulation simulates the effect of incomes policies on the timing of negotiation. The overall conditional probability to negotiate between 1963 to late 71 is almost invariant to whether the policy is on or off. This implies that the simulated wage level depicted in figure 6.30, although they are close to the actual values, is merely a result of the dynamic simulation of the wage equation with the timing of wage rises taking place at random.

The dynamic simulation that fixes the timing of wage changes as observed in the data has succeeded in simulating the average wages that are almost equivalent to the actual wage levels. When the timing is also simulated, the wage levels are very close to the actual values between 62 and 68 (figure 6.31), but the discrepancies exist between them during the rest of the sample period. A comparison of the simulated wage levels between the two reveals some interesting points. If the movement of these simulated values differ, we know that is due to the simulated frequency of wage changes.

As can be seen in figure 6.31, prior to 62, it is higher than the actual due to more frequent wage changes simulated during fd1 (since this diversion is not found in any of figure 6.3-6.5). Surprisingly, this is the case even when the simulation is based on the hazard formulation that showed a strong negative impact of fd1 on the private groups (table 6.4). With such a negative effect of fd1 amongst all groups, less frequent wage changes should be simulated resulting in lower, closer to the actual, simulated wage levels than those based on table 5.9. This does not seem to be the case, however. At the termination of fd1, we see a slight rebound in the wage level which was not seen in the actual wage level. From then onwards, the wage levels are well simulated up to the first all-sector-enforced ceiling policy, cd4. During cd4, cd5 and cd6, the wage level actually increased more rapidly than the simulated ones. That is, the impact of these policies in reducing the frequency of negotiations were too strongly represented compared to their actual effect. This resulted in the lower than the actual rate of wage increase to be simulated during the early 70's. The huge rebound at the termination of cd6 observed in the actual wage level, on the other hand, has been successfully picked up. This is mainly caused by an increased frequency of wage changes that led to a higher rate of increase in the average wage level. And such rate of increase is accurately simulated. Eventually, we find the policy-on locus of the simulated wage levels catching up with the policy-off locus by the early 71, at the time, their simulated wage level also converges to the actual wage level.

Comparison with the simulation using the wage equation only (figure 6.3-6.5) reveals the misspecified portions of the policy-off average wage levels. We have already found out that there are significant differences between the locus of policy-on and policy-off negotiation probabilities. For example, a significant rebound in the hazard seen when cd6 terminated was not found in the policy-off simulated hazard. Considering that the figure 6.3-6.5 are based on the historical settlement dates that actually experienced all the incomes policies, the wage levels depicted there are likely to be overstated. This is because the actual impact of the incomes policies in altering the frequency of wage changes are kept intact even in the simulation without any policies. In these figures, wage changes take place equally frequently throughout the sample period whether the policy is on or off. In general, the simulated wage level is much smoother when the timing is fixed. And the gap in the wage levels between policy-on and policy-off regimes is more uniform across episodes. When the timing is also simulated, their impact become more diverse since they now include the combined effect on the wage levels, that is, they are now additionally picking up the effect of the changes in the frequency of negotiations. According to them, cd5 and cd6 had the largest combined impact in reducing the average wage levels, although, as we have seen, they have overstated the extent of the actual effect.

Failure to correctly simulate the rate of wage rise during the 70's are seen in all the hazard specifications reported so far except for the static one. At least, this implies that the identification problem regarding the pure duration dependence in the hazard specification is not responsible for this problem. What is, then, causing cd4, cd5 and cd6 to suppress the hazard too much? It may be that some factors that offset the negative impact of the latter incomes policies on the frequency of wage changes are being omitted from the hazard specification of table 5.6, or that these policies did not have an uniform degree of impact on the hazard throughout their policy period. The observation that the simulated policy-off locus is lower than the actual implies the strong possibility of the former. In view of this, we have additionally incorporated the log of real wage level, which showed a sudden slight increase in the early 70, into the spell specific component and allowed the policy cd5 and cd6 to have separate effects on the hazard. However, the resulting simulated wage level has turned out to be hardly different.

6.2.3 Conclusion

In order to investigate the impact of the incomes policies more fully on the dynamic system determining the timing and magnitude of wage changes, we have experimented with the dynamic simulations.

As long as the specification for the hazard contains time varying covariates, a unique length of a completed duration spell given its starting point cannot be determined by the hazard function. In a full dynamic simulation, we need to simulate when the wage changes take place (i.e., completed durations), since the rate of wage change equation is conditioned on the timing of the settlement and the average wage levels are simulated on the basis of the wage equation.

In order to simulate the timing of settlement, we approximated the continuous sequence of decision making to take place monthly. Then, this problem reduces to a repeated monthly Bernoulli trials with the conditional probability to exit at the δ-th month of the elapsed duration being $(S(\delta-1)-S(\delta))/S(\delta-1)$, where S(.) denotes a survivor function which can be easily deduced from the hazard rate. Provided that the timing of negotiation is simulated, the explanatory variables that determine the wages to be claimed at such simulated timing also become available. Thus, the wage levels can be simulated accordingly. The dynamic simulation done this way enables us to investigate the full impact of external influences, particularly of incomes policies, on this system, particularly, on the timing and the magnitude of wage changes.

The figures show lower negotiation probability during all the policy episodes with a possible exception of cd4, which took effect right after the mandatory freeze policy, fd2. The first freeze policy, fd1, was found effective in reducing the frequency of negotiation, thus, lowering the rate of wage rise even amongst the non-enforced groups. Simulated wage levels, whether the timing is also simulated or not, were lower under all the policy episodes. In particular, the average wage level, simulated with the timing of settlements fixed as in the data very closely replicated the actual wage level. Also, amongst the average simulated wages based on several specifications of the wage and hazard equations, the most restrictive static specification of the hazard estimated by the MLE replicated the ones closest to the actual levels. Nonetheless, we found that :(1) additionally simulating the timing of wage changes enabled us to see the impact of policies in suppressing or increasing the frequency of wage changes and their resulting impact on the realized wage claim, (2) allowing the external influences to affect the hazard continuously throughout the spell enabled us to derive the instantaneous impact of these variables, which were reflected in the simulated probability to negotiate in a way consistent with the data. The static representation failed to predict any effect of the incomes policies on the negotiation probability. Hence, their simulated wage levels, even though they were closer to the actual average than those based on the time-varying hazard, was a result of the simulated settlements taking place at random, since their timing was not affected by the external influences. In particular, our simulation succeeded in replicating most of the average wage level prior to 1970, and the sudden increase in the negotiation probability in the aftermath of fd2 and cd6, which led to a sharper increase in the wage level. However, the impact of the latter ceiling policies, cd4, cd5 and cd6, in reducing the hazard rate was much too strongly simulated than their actual effect. This may be due to some omitted factors that induced more frequent settlements during the 70's. At least, the failure to allow for the pure duration dependence in our final specification of the hazard is not responsible for this. Otherwise, the time varying covariates incorporated for the specification can be said to have adequately explained the dynamic dependence for much of our sample period.

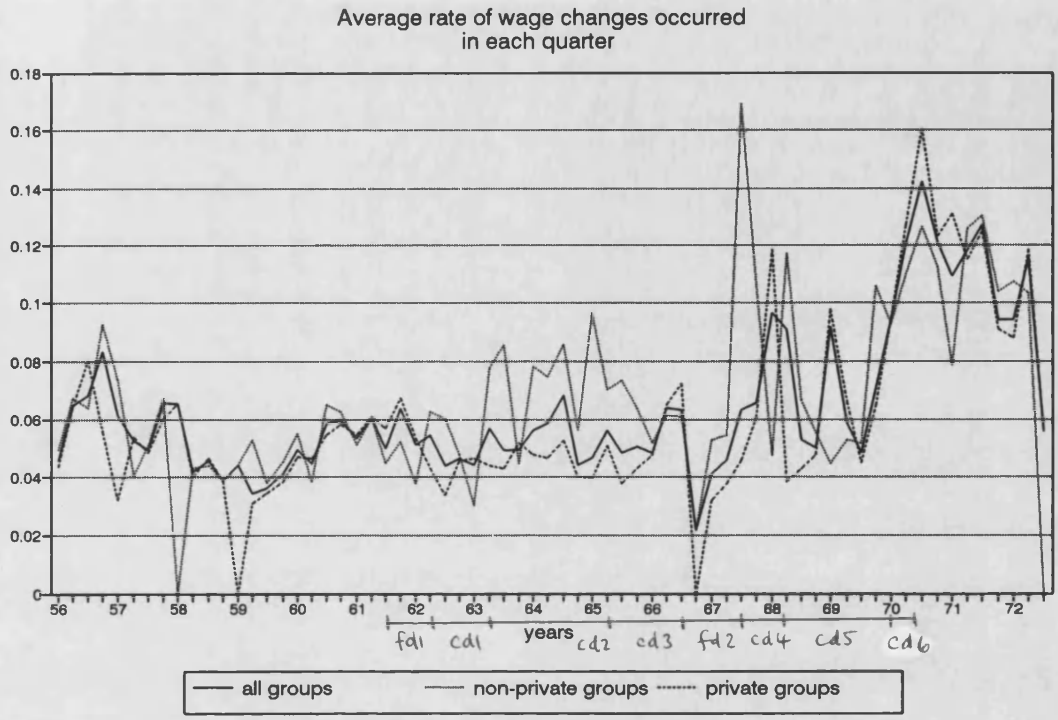


Figure 6.1

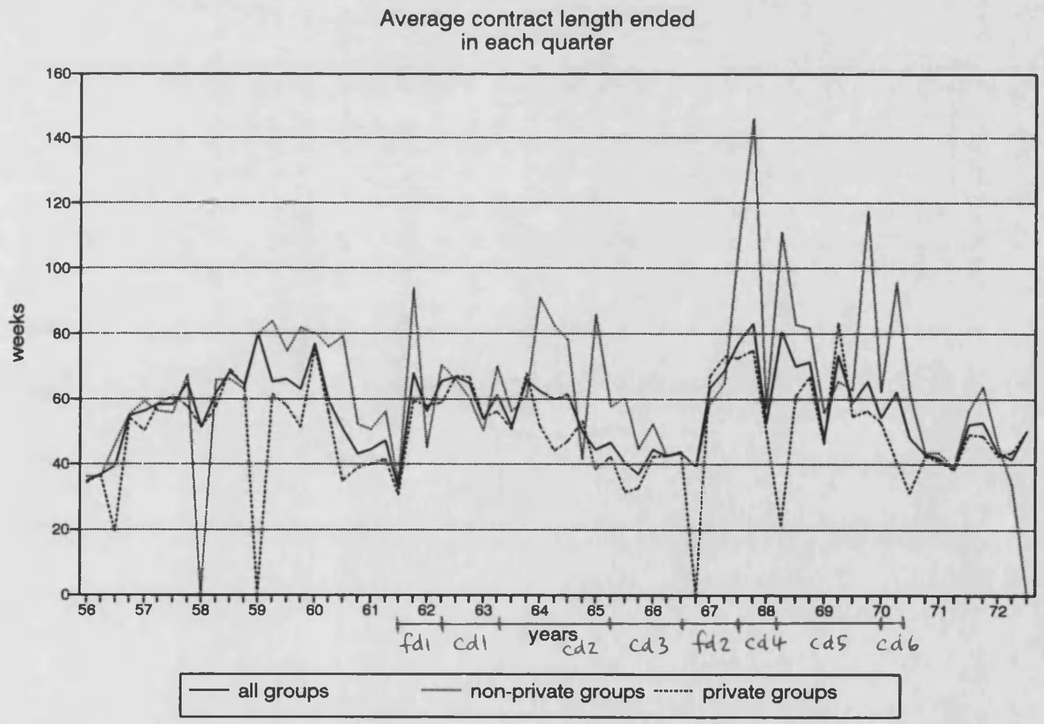


Figure 6.2

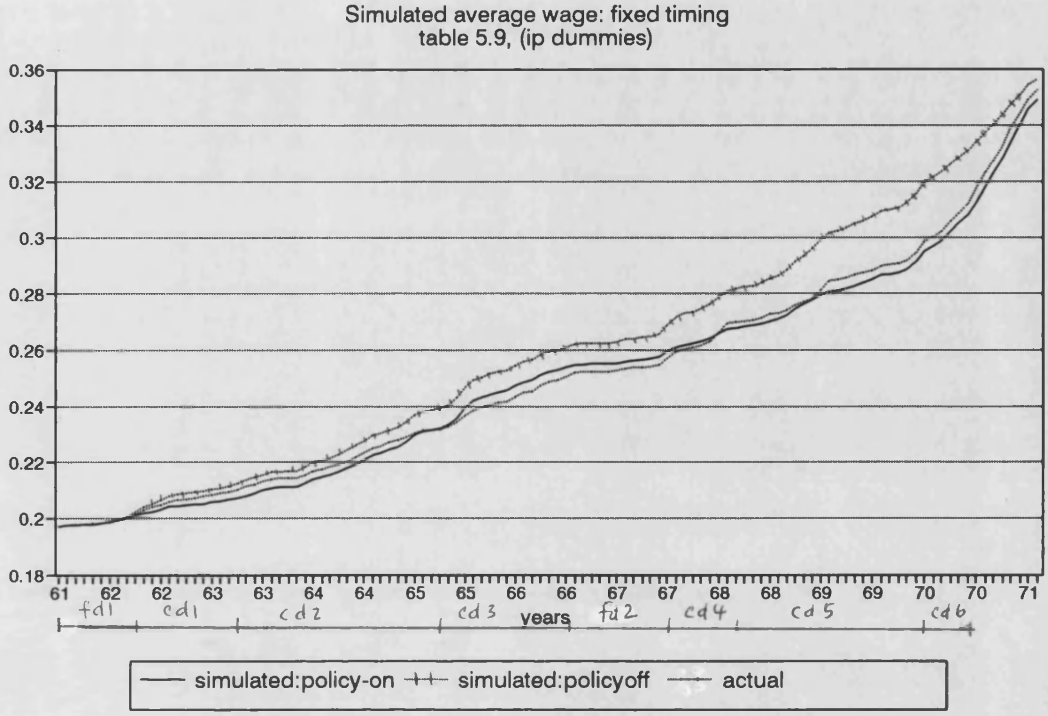


Figure 6.3

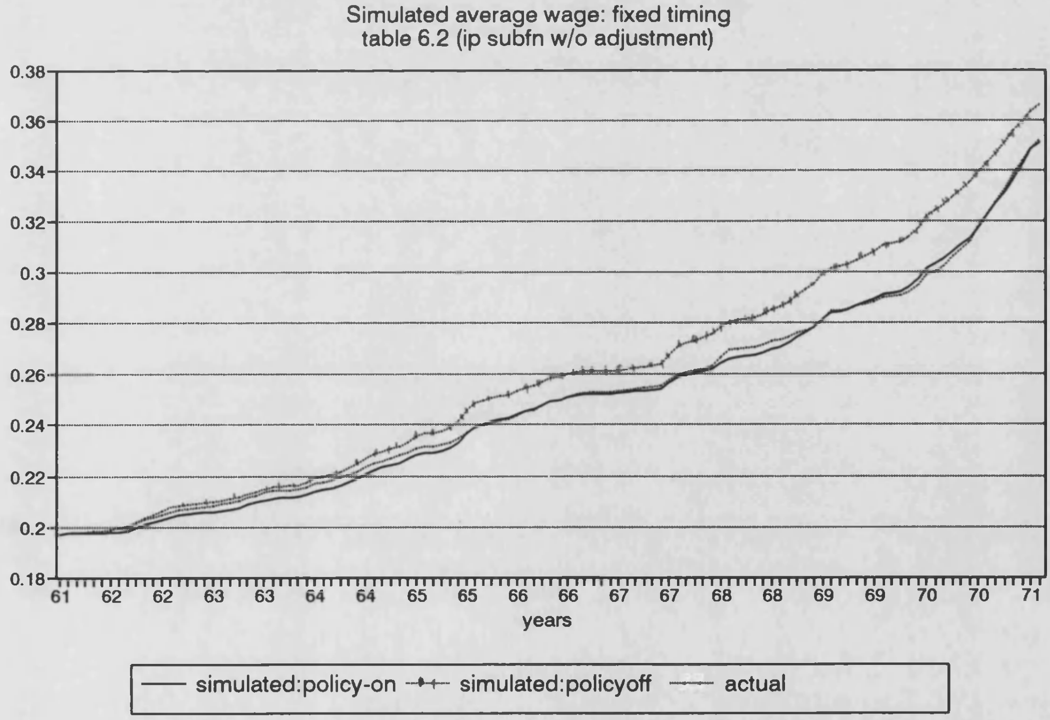
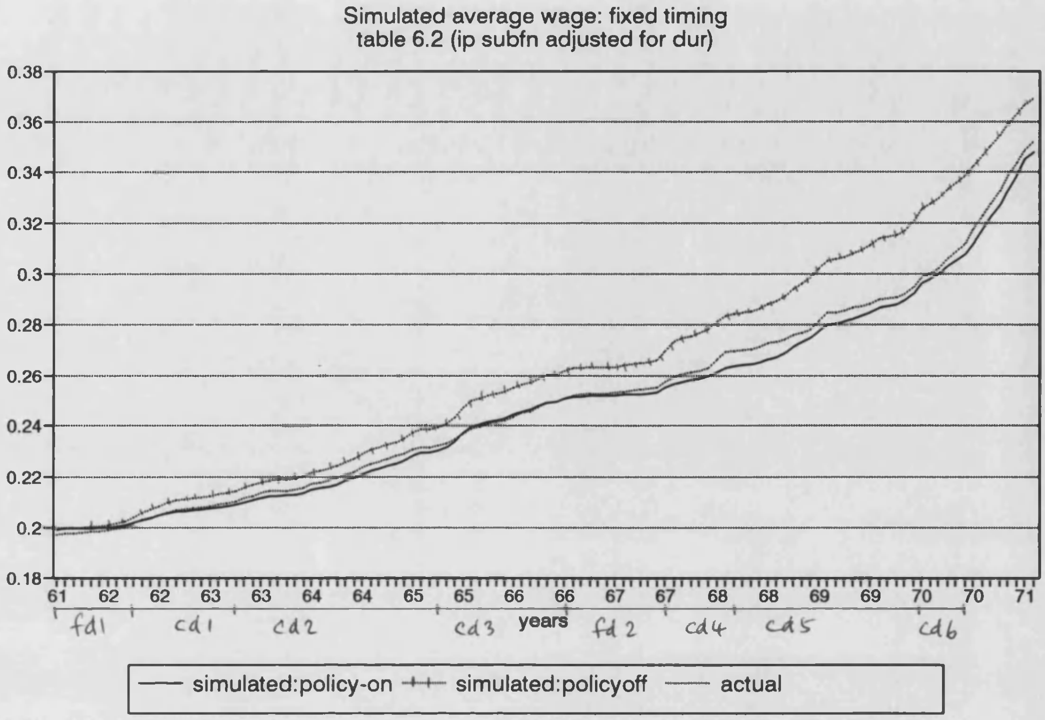


Figure 6.4



Simulated average wage: fixed timing
table 6.2 (ip subfn adjusted for dur)
0.38
0.36
0.34
0.32
0.3
0.28
0.26
0.24
0.22
0.2
0.18
61 62 62 63 63 64 64 65 65 66 66 67 67 68 68 69 69 70 70 71
years
fd1 cd1 cd2 cd3 fd2 cd4 cd5 cd6
simulated:policy-on
simulated:policyoff
actual

Figure 6.5

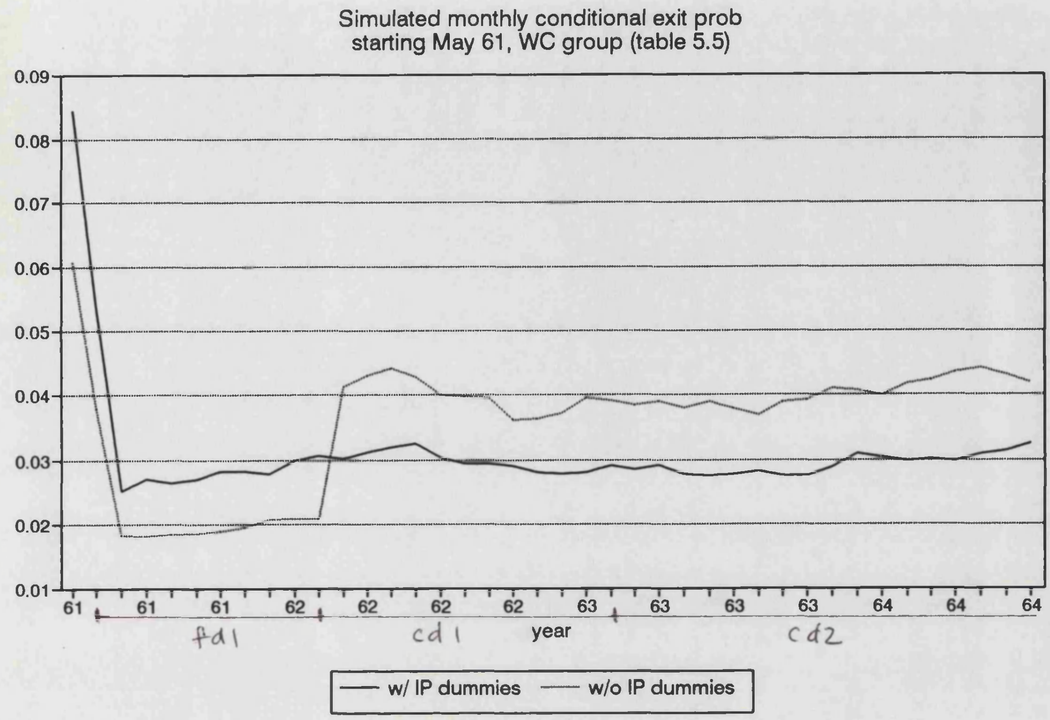


Figure 6.6

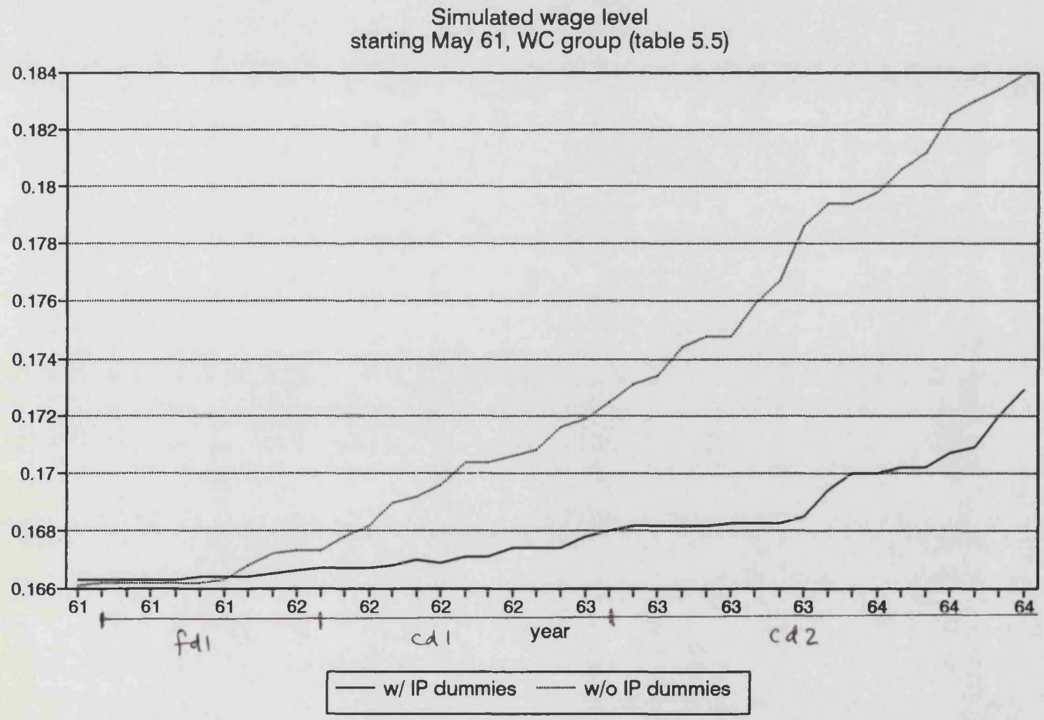


Figure 6.7

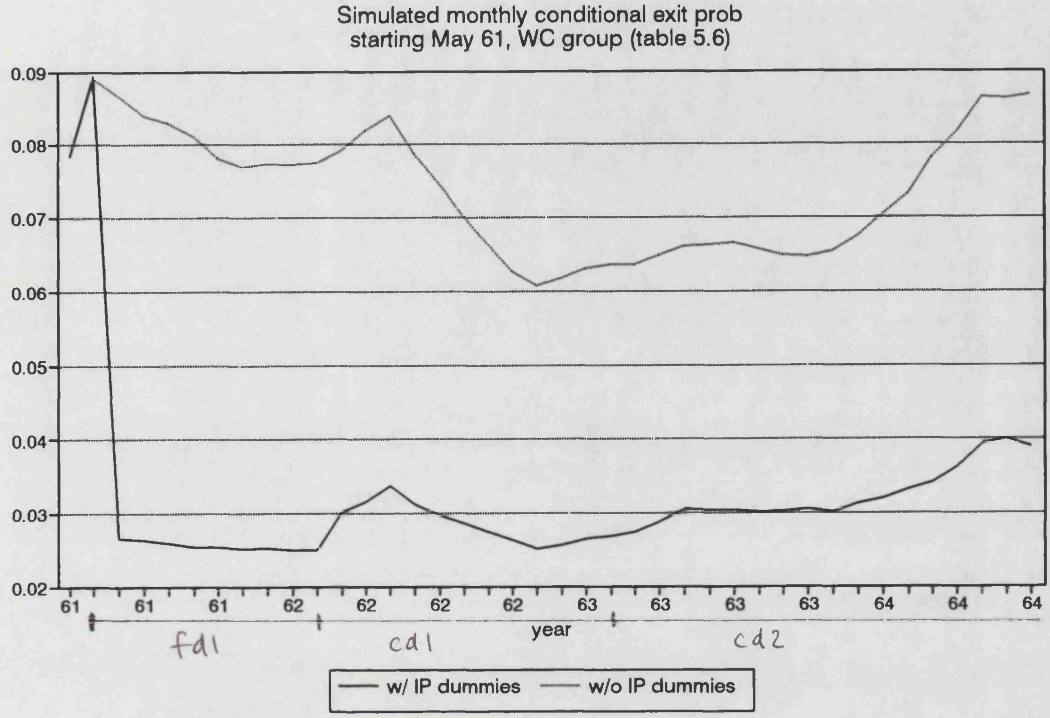


Figure 6.8

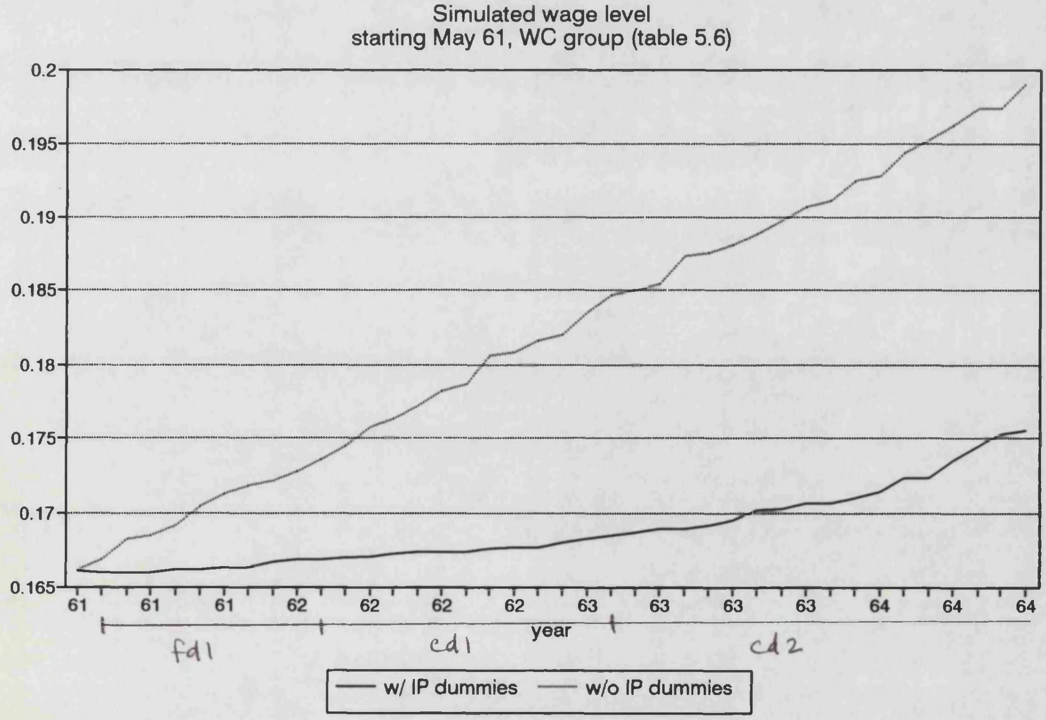


Figure 6.9

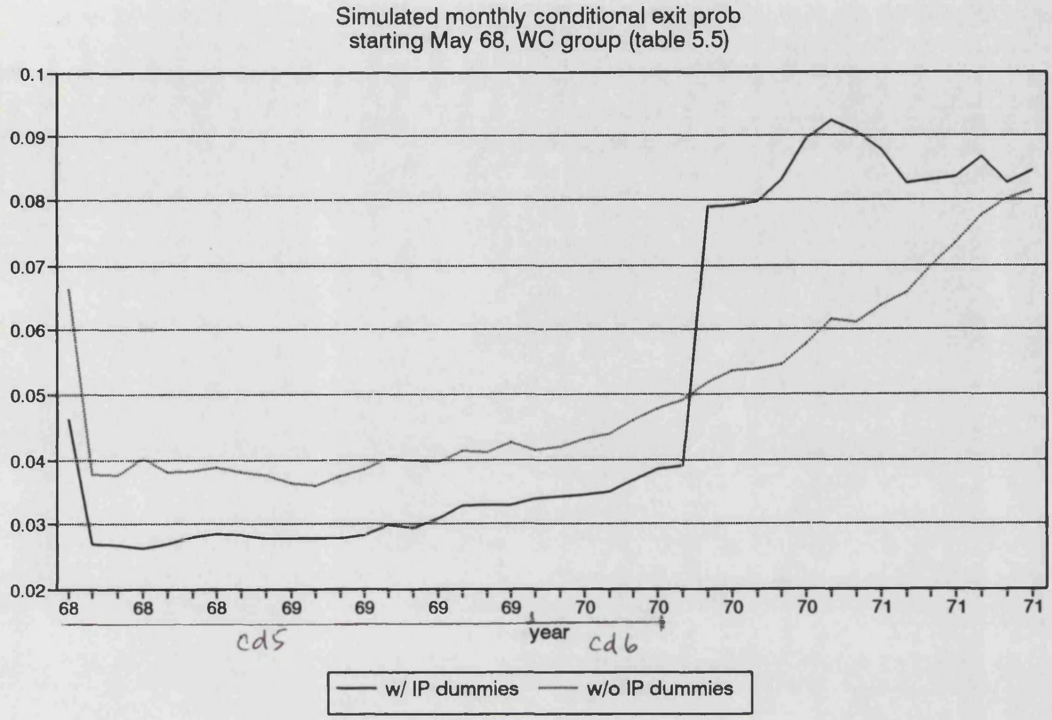


Figure 6.10

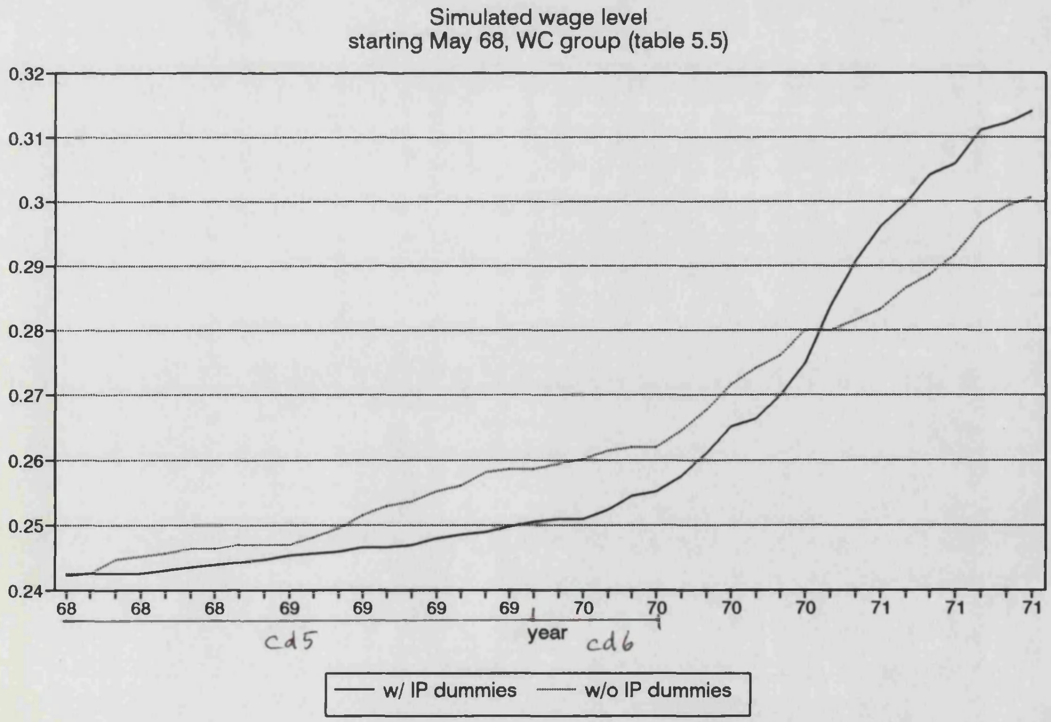


Figure 6.11

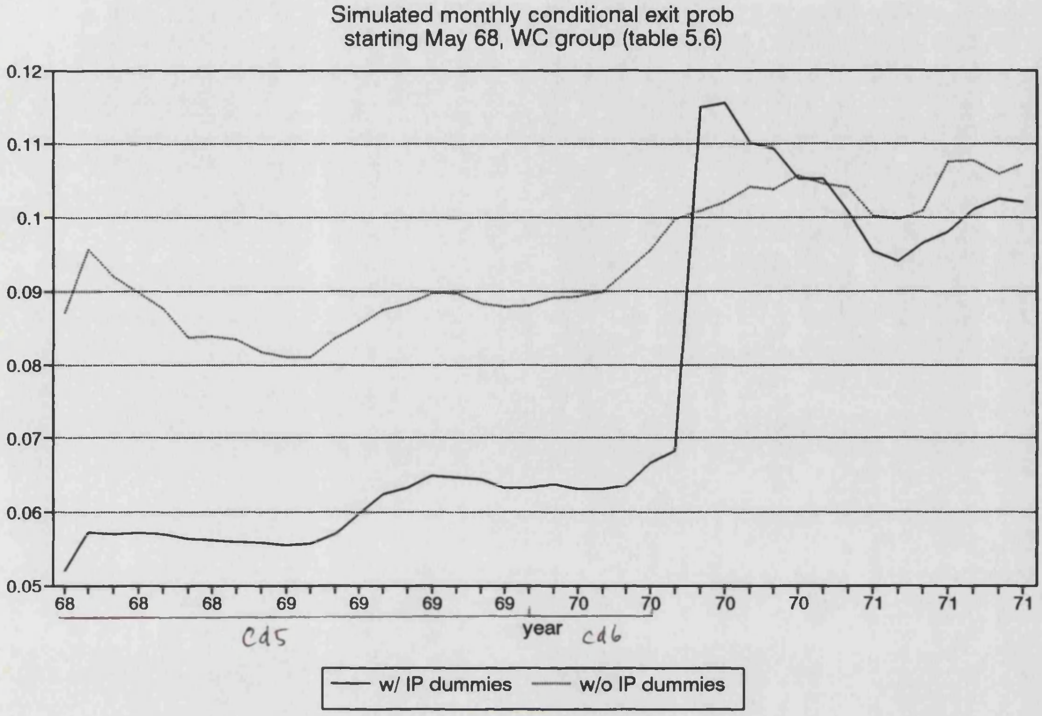


Figure 6.12

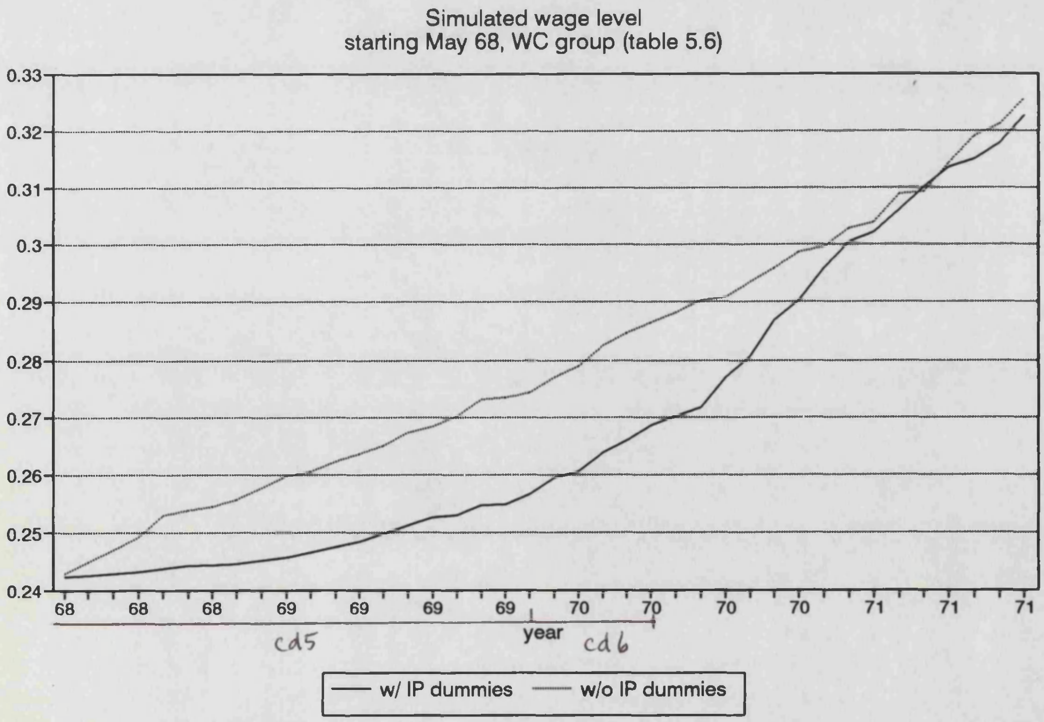


Figure 6.13

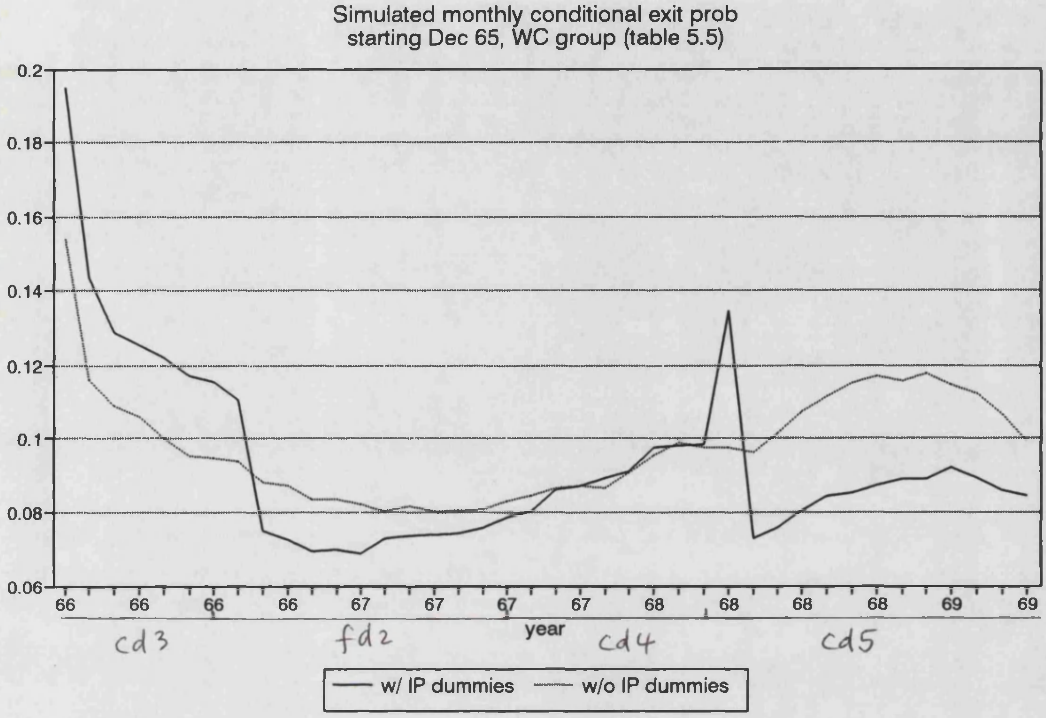


Figure 6.14

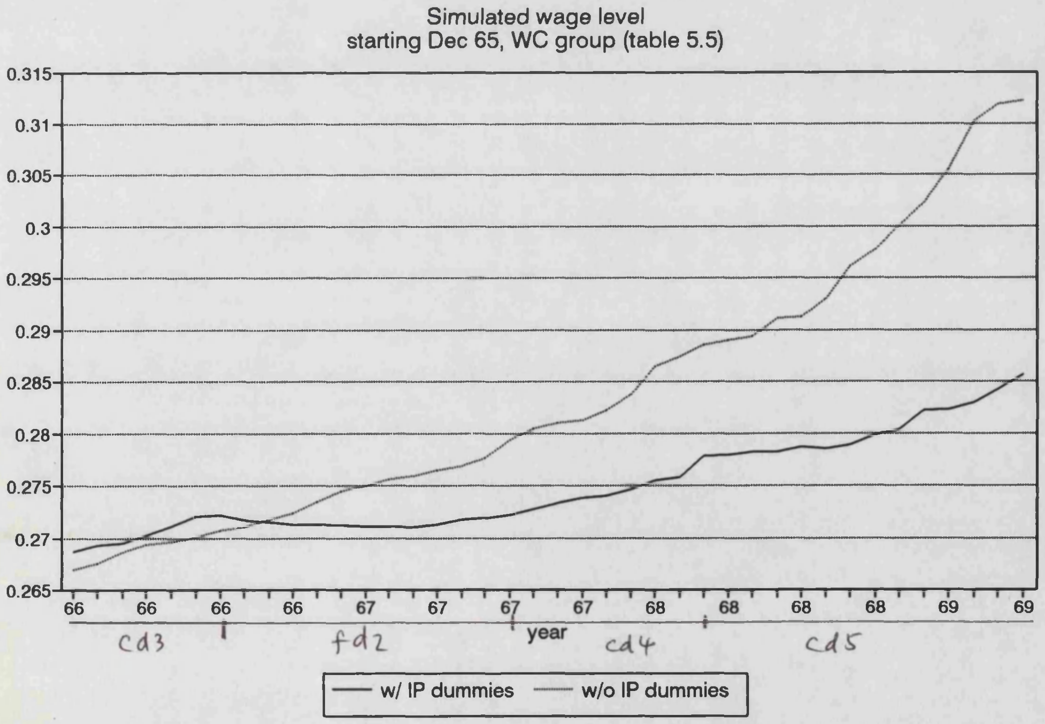


Figure 6.15

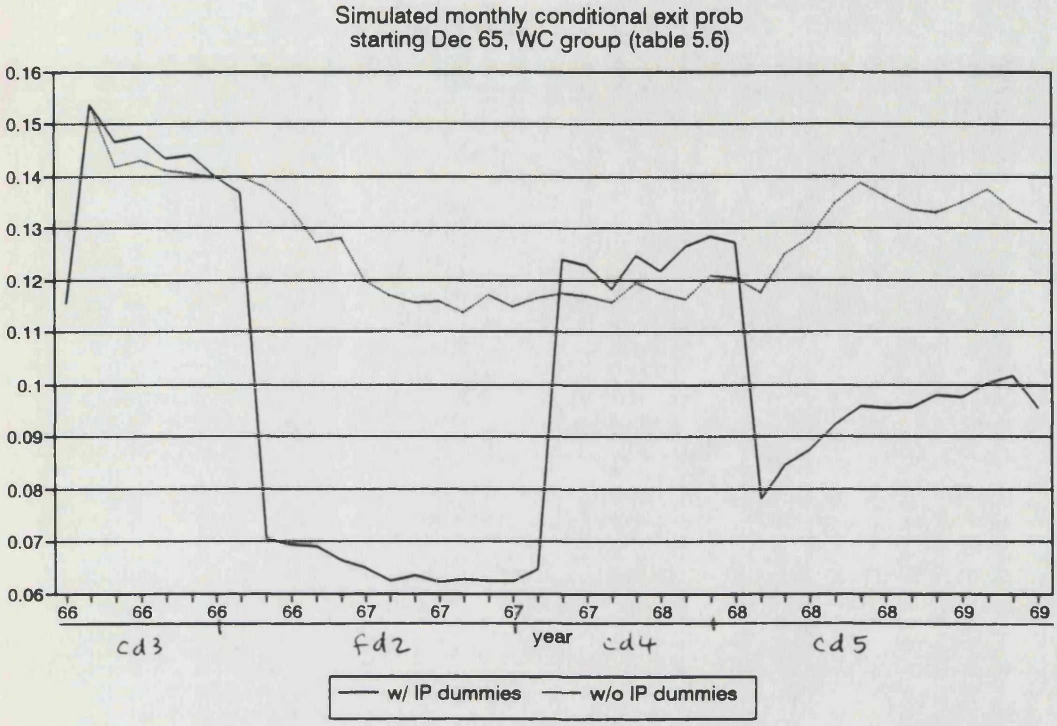
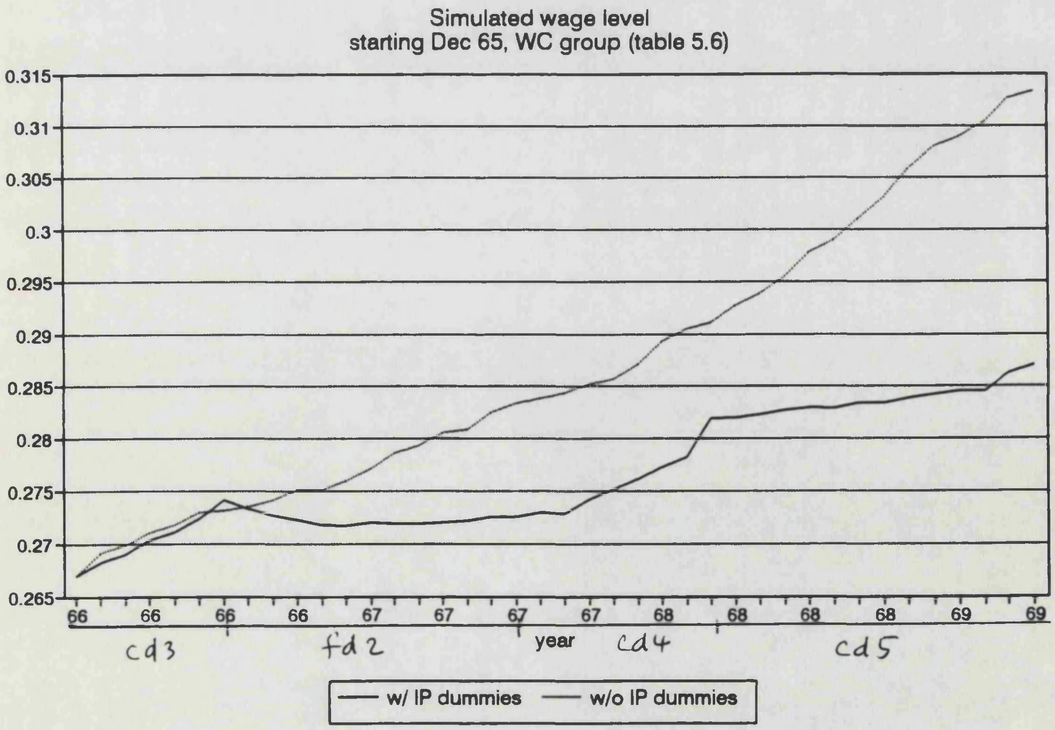


Figure 6.16



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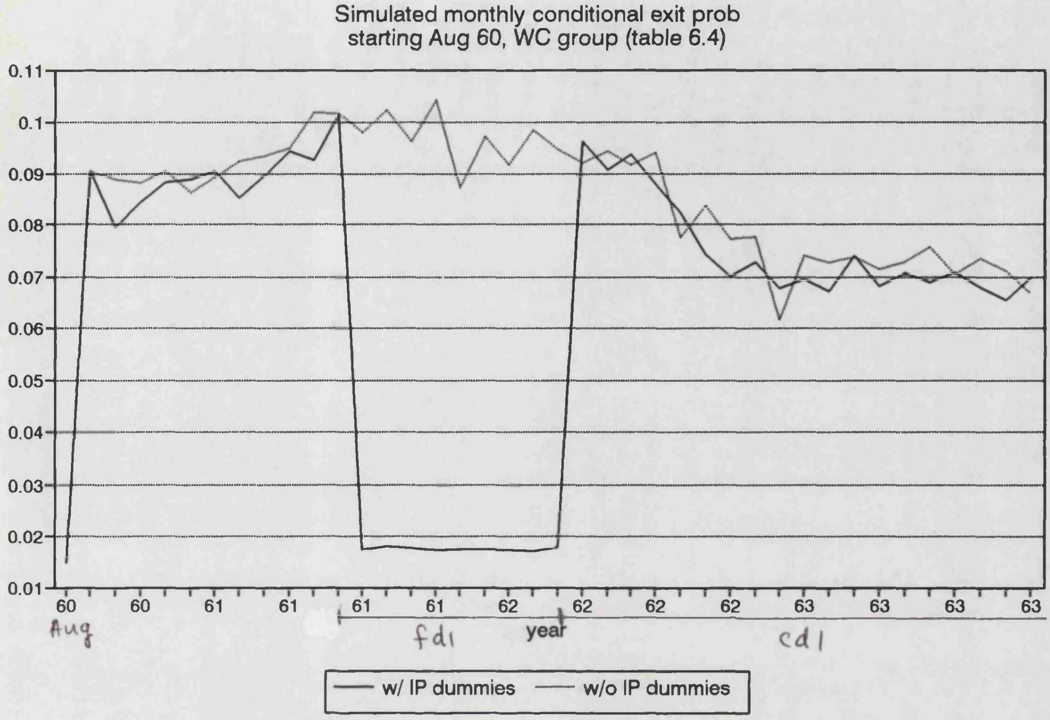


Figure 6.18

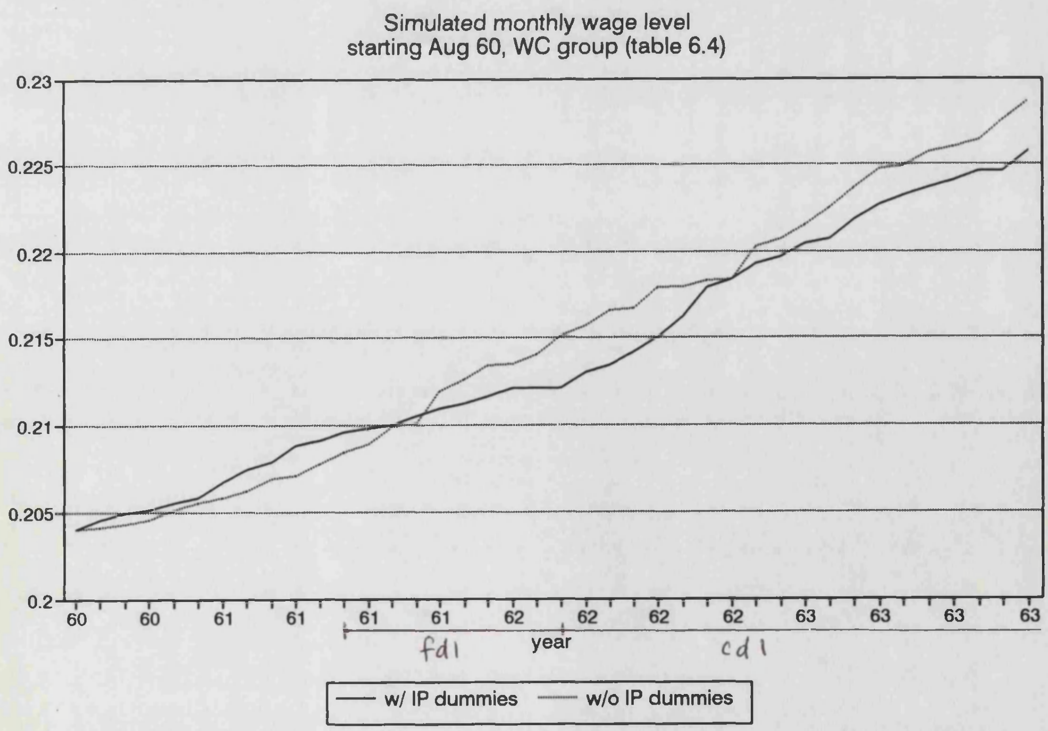


Figure 6.19

Simulated average conditional exit prob
for all groups, based on table 5.5

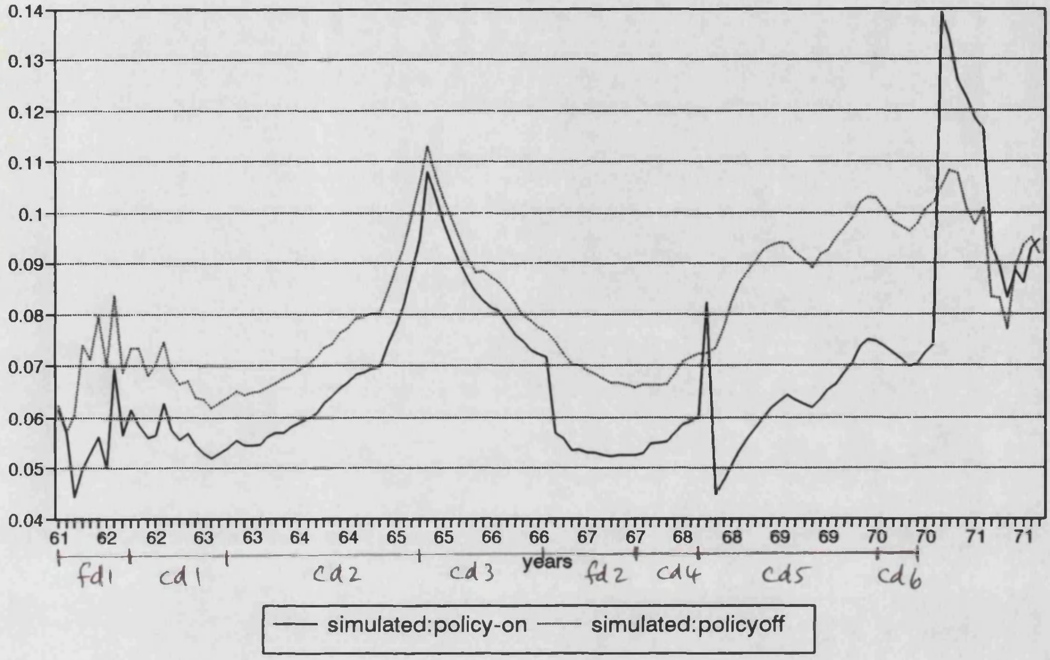


Figure 6.20

Simulated average wage for all groups
based on table 5.5 and 5.9

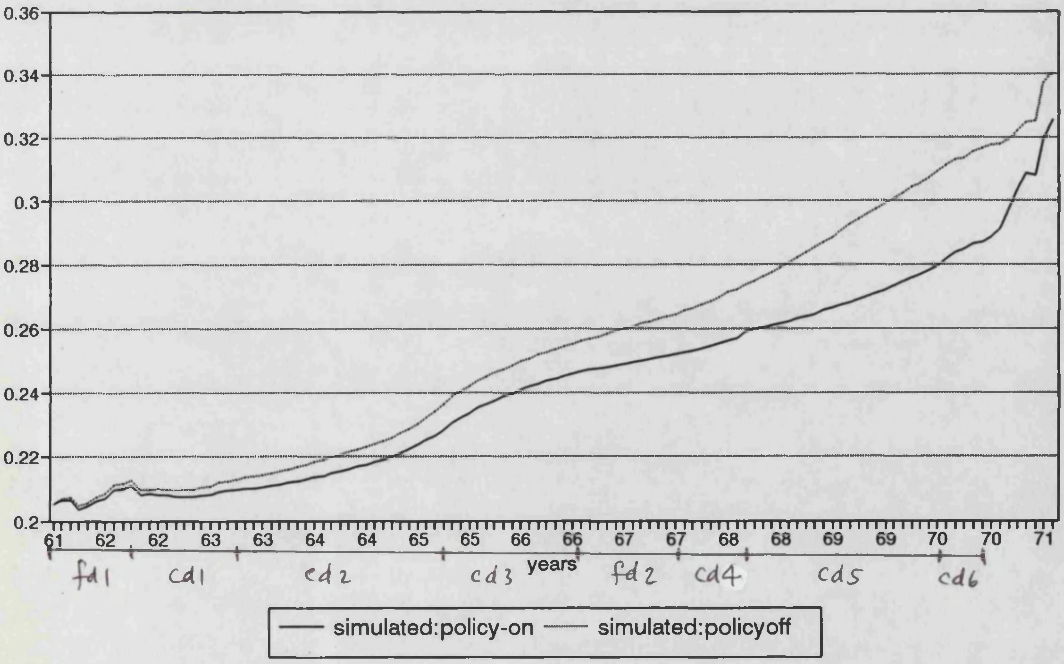


Figure 6.21

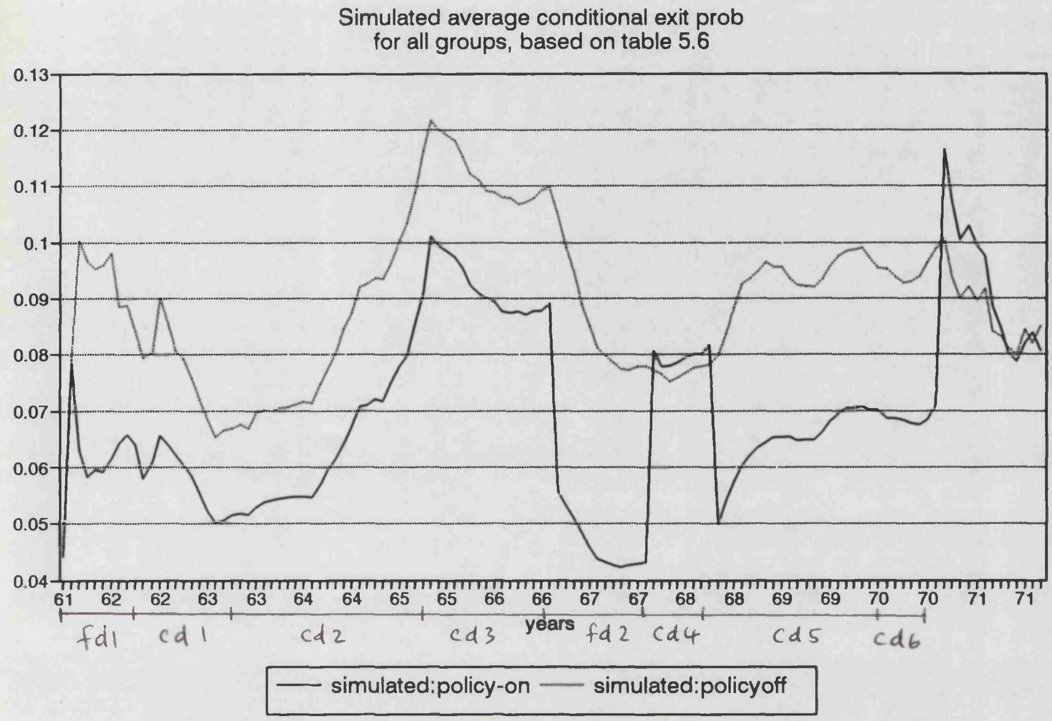


Figure 6.22

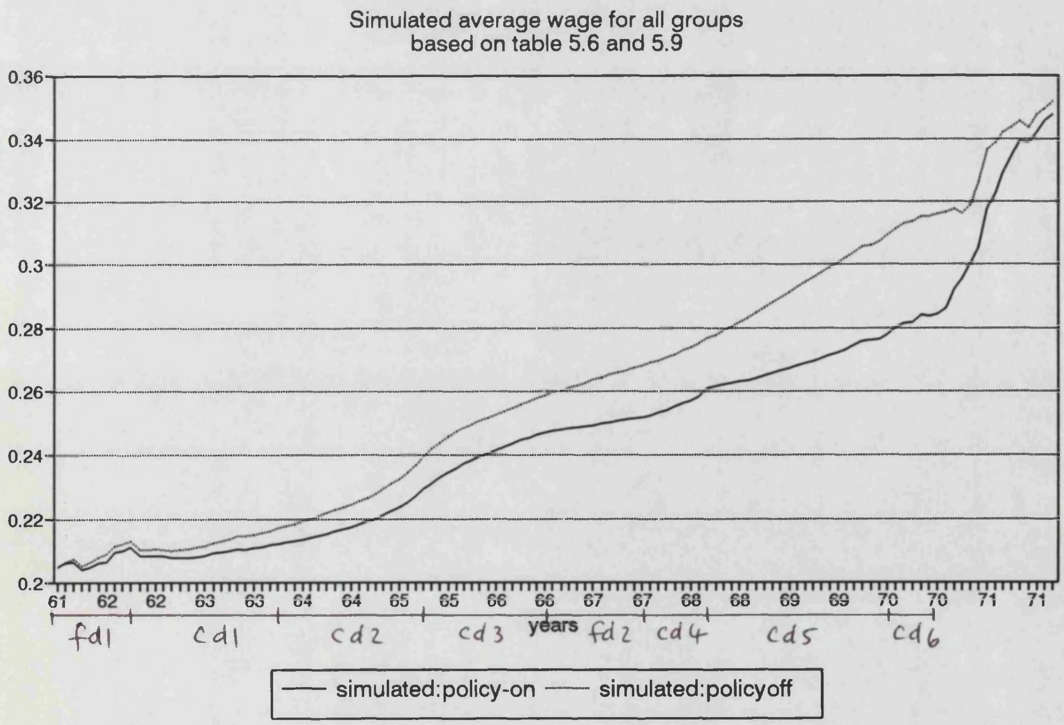


Figure 6.23

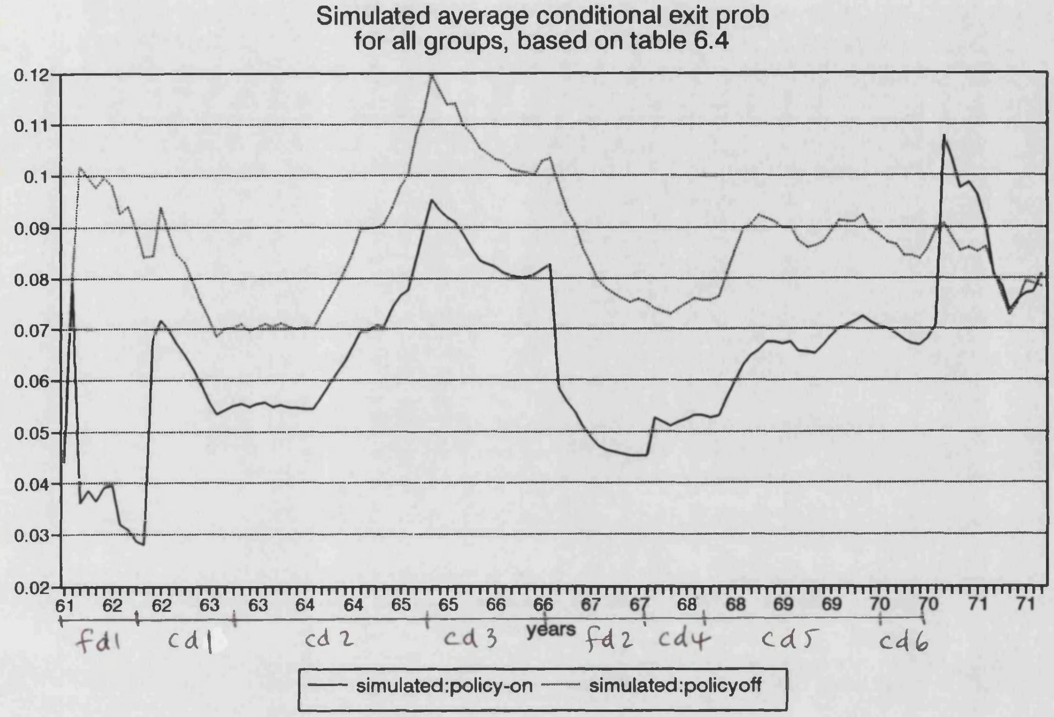


Figure 6.24

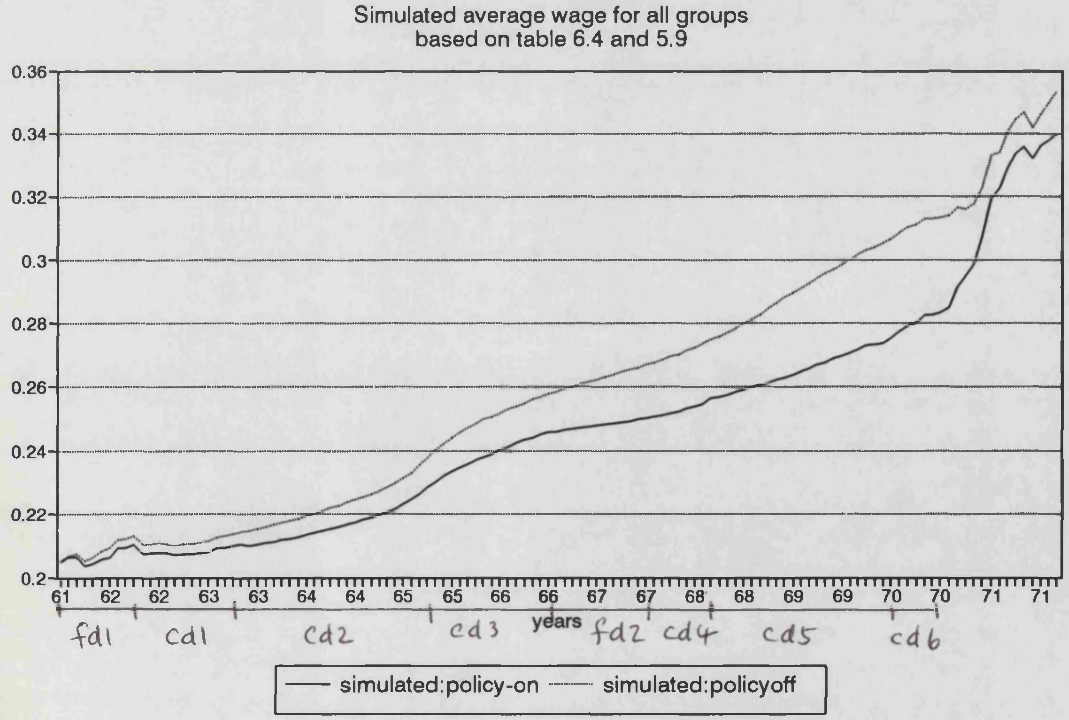


Figure 6.25

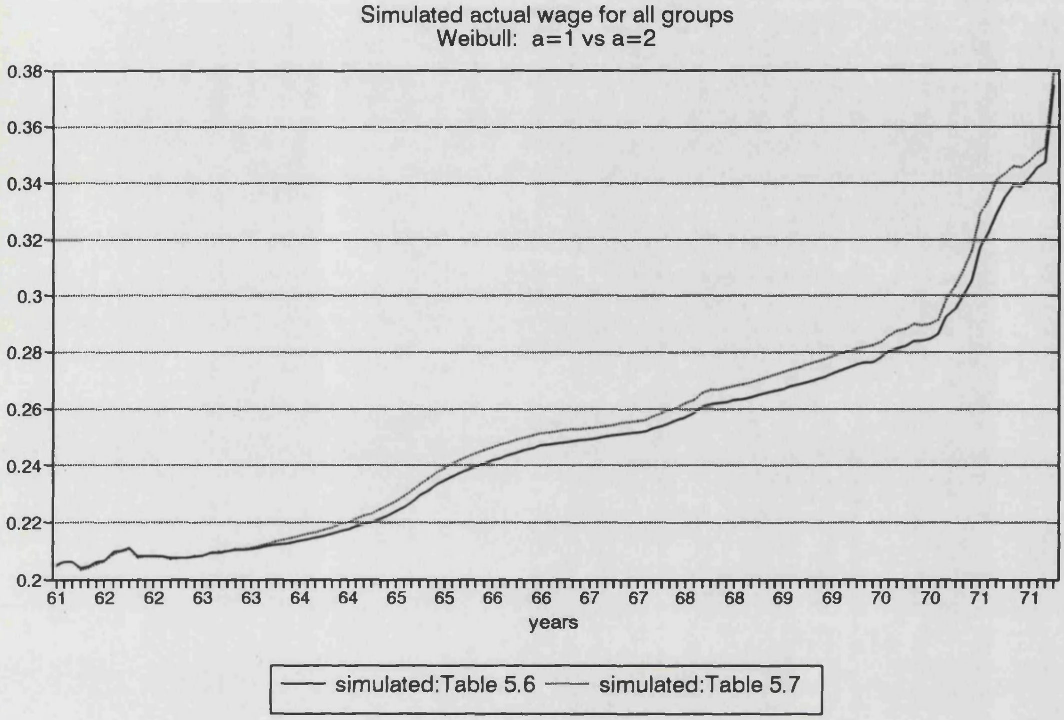


Figure 6.26

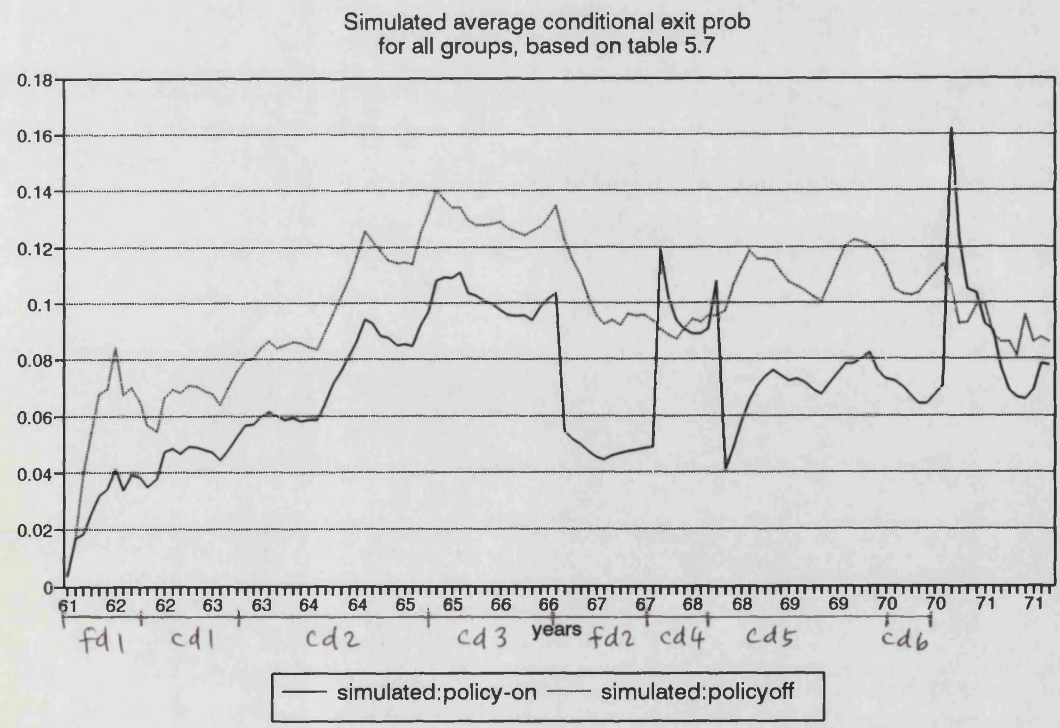


Figure 6.27

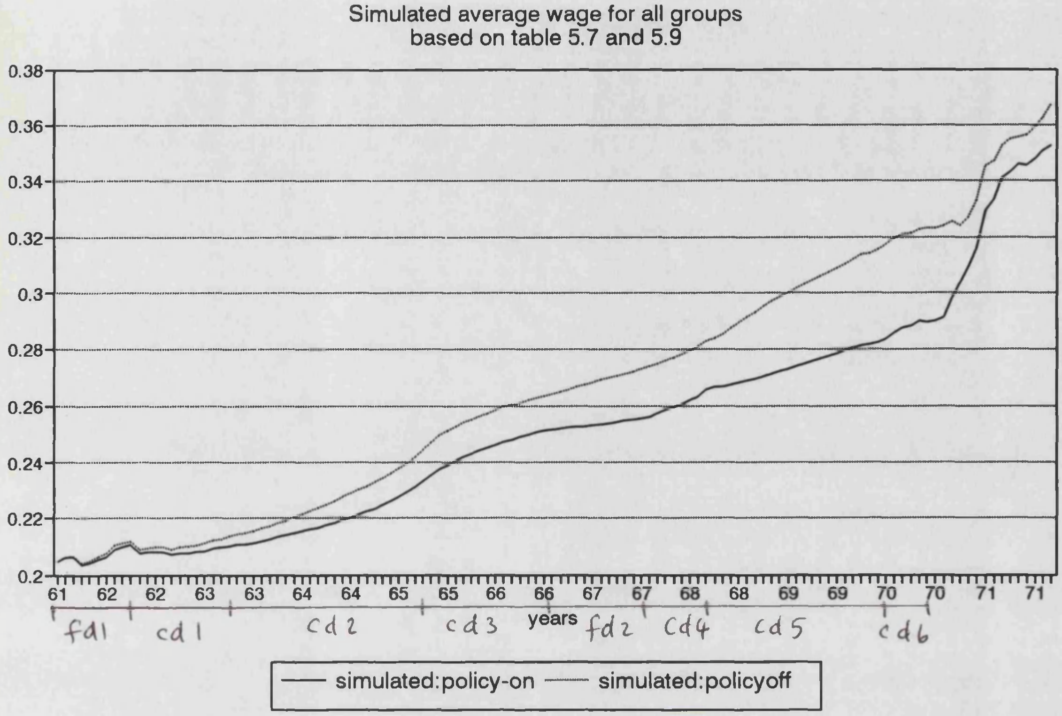


Figure 6.28

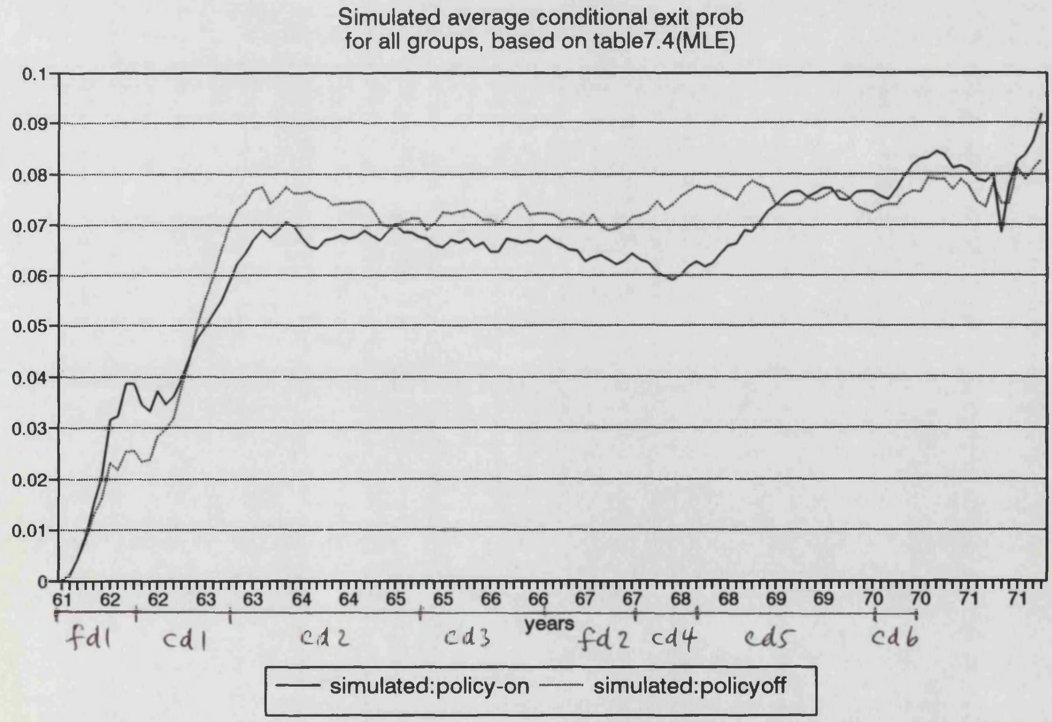


Figure 6.29

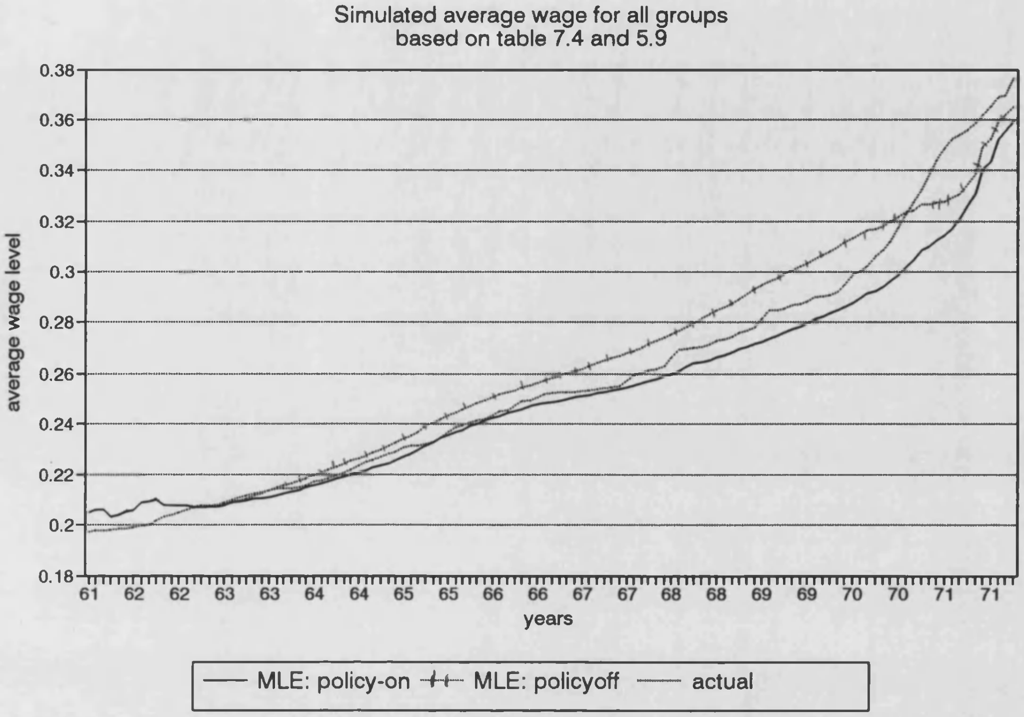


Figure 6.30

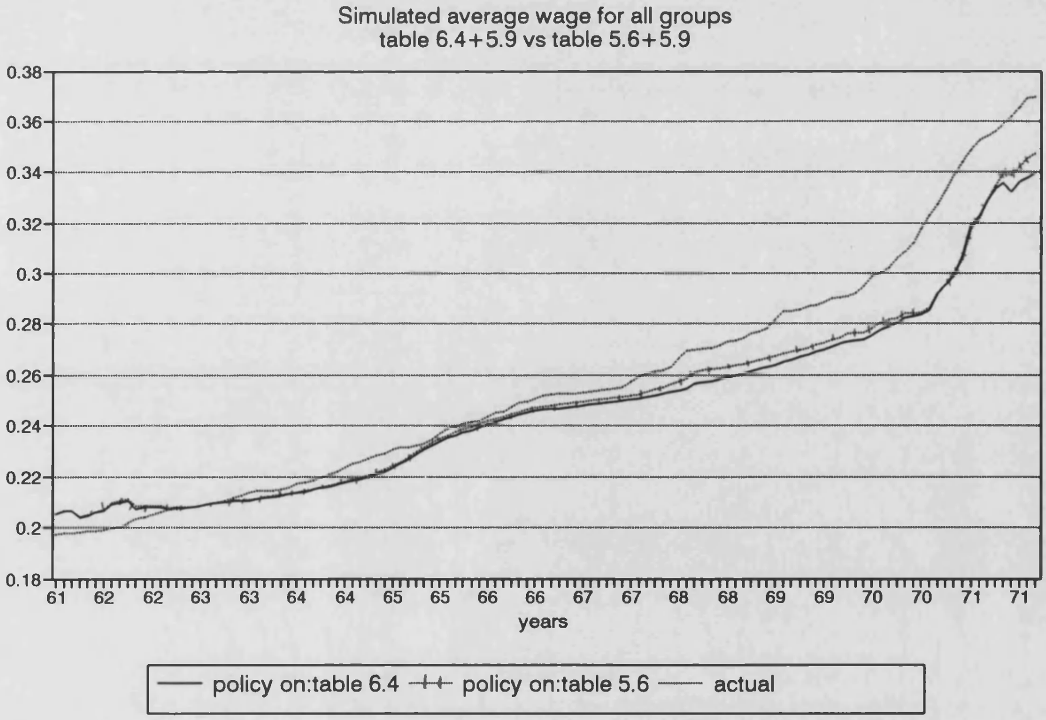


Figure 6.31

Footnote to chapter 6

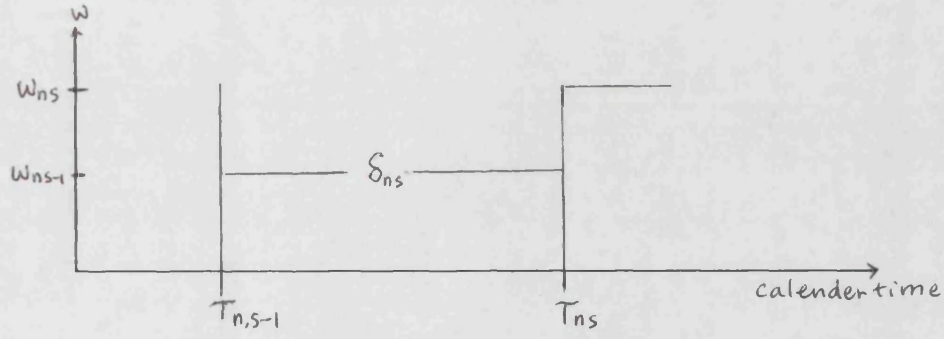
1. First step heteroscedastic consistent estimates, t-ratio in parenthesis.
2. $DRR^e(\tau^i)$ is based on information set available either at τ^{i-3} or τ^0.
3. Plot of the estimated density function would be appropriate if plotting it against the elapsed duration. However, since the dynamically simulated durations corresponding to each point in calendar time largely vary, we plot the conditional monthly exit probability against the calendar time.

Chapter 7: Comparison with the other estimation methods

7.1 Introduction

The paramount uniqueness of the model discussed in chapter 4 and 5 is its capability in explaining both the timing and the degree of negotiated wage changes simultaneously. The bargaining intervals and the negotiated wages are interrelated with each other in a sense that the wages to be negotiated depend on the length of the contract just terminated and its timing. In explaining their interdependence, our model is able to incorporate not only the length of the contract, but also the entire time path of economic environment during such contract. In practice, a simultaneous model like ours becomes particularly useful in explaining any pairs of variables bearing a similar relation where one measures an elapsed spell length while the other is observed at the end of such spell, therefore, depends on the length and the timing of spell termination.

Look at the figure below for a typical s-th duration for n-th cross-section group (ie. δ_{ns}) which started at T_{ns-1} and ended at T_{ns}, calendar time. Let the corresponding measure be wage, w_{ns}, which is determined at the end of δ_{ns}, at T_{ns}.



(Figure 7.1)

The simultaneous relation is represented in terms of a joint density function which can be decomposed into two components: the spell duration distribution and the conditional wage distribution, where the latter is conditioned on the entire history up to and including a termination point of the current spell. This is to say that a size of wage change is assumed to be determined after a decision to terminate a previous contract is made. Once the joint density is decomposed in this way, it is then possible to model and estimate them separately. The conditional wage distribution is represented in terms of a linear wage equation where the explanatory variables include factors known at a time of new negotiation, hence, its estimation becomes quite straightforward. On the other hand, the duration distribution is represented by the hazard rate which has one to one relation with the duration density function. We have chosen the hazard rate for the analysis essentially because it is more intuitive compared to the the duration distribution itself. But the advantages of using the duration analysis is far more than that. It can accommodate censored data, time-varying regressors, random heterogeneity, and moreover, can identify a form of duration dependence. As one can imagine, different restrictions and assumptions over these complex issues yield different models, therefore different ways to estimate them. Obviously, the estimation techniques grow more complex as the model becomes more generalized and less restrictive. Essentially, estimation method varies according to the assumptions over how to deal with a time path of the explanatory variables, the duration distribution (i.e., a form of the baseline hazard) and the heterogeneity distribution.

In this chapter, we set forth several conventional methods of estimating the hazard equation and actually estimate them using the same data set we have used for our model in the previous chapters. In the following section, we categorize these different estimation methods into three groups and discuss briefly their pros and cons in general. They are then compared to the nonlinear instrumental variable estimates of our dynamic hazard model. In section 7.3, the actual models to be estimated are described and their results discussed. Since most of the existing studies have avoided the complication by fixing the values of explanatory variables at the start of a spell, we pay particular attention to the effect of such variables being spell specific instead of varying throughout the spell, and see how the direction of duration dependence is affected by such restrictions. For this purpose, the time-varying covariates are not considered here.

7.2 Methods of estimation

We assume the proportionate hazard formulation throughout this chapter. For the s-th duration observation of the n-th group, δ_{ns}, observed at calendar time $T_{ns-1}+\delta_{ns}$, the most general hazard given the explanatory variables z_{ns}, the group-specific random heterogeneity u_n and the error term ε_{ns} is:

$$h(\delta_{ns} \mid z_{ns}, u_n, \varepsilon_{ns}) = \exp(\gamma' z_{ns} + u_n + \varepsilon_{ns})\, h_0(\delta_{ns};\alpha)$$

h_0 captures its dependence on time that is common to all groups, while the remaining part shifts the hazard for different negotiation groups for each observation. Followings are the estimation methods for the above hazard function most widely used in the conventional literature:

(1) Linear regression
(2) Fully parametric maximum likelihood
(3) Partial likelihood/ semi-parametric maximum likelihood

Our methodology in chapter 5 and 6 can be regarded as a family of regression models, though, these referred under (1) here are with a simple dependent variable such as a logarithm of duration. Hence they cannot take into account of the within spell varying regressors nor can it identify a parameter of duration dependence. As is the case in the dynamic method in chapter 5 and 6, censored data cannot be handled. On the other hand, the group or the spell specific random terms, and the possibility of correlations between these terms and the explanatory variables can be straightforwardly incorporated. Its simplicity makes it a valuable starting point before embarking on to more complicated estimations.

In our context of multiple spell setting, the hazard is conditioned on its past history of bargaining. Hence, the joint density for each duration is a product of conditional densities each of which is conditioned on the unknown heterogeneity term, possibly, both group and spell specific. In order to adopt a maximum likelihood estimation (ie. (2)), it requires a multiple integration over such heterogeneity terms to derive the appropriate marginal likelihood. Hence, to estimate parametrically, we must assume distributional forms for all of these random terms. However, not only will these assumptions be inevitably *ad hoc*, since the economic theory rarely predicts their forms, but also their computation of the marginal distribution be complicated and difficult, particularly when both group and spell specific random terms are allowed. Also, the individual error term may not be independent but correlated with the lagged durations. Moreover, introduction of time varying explanatory variables would further complicate such computation. One can, on the other hand, ignore the random terms and start off by introducing the time-varying covariates. But the *ad hoc* assumption over the parametric forms of the baseline hazard still has to be made, while, this time, the difficulty arises in identifying between the time-varying covariates and the pure duration dependence. The correctness of the parametric assumptions is inherent problem with the parametric maximum likelihood which can be avoided, to a certain extent, by adopting the semi-parametric method.

Unlike (1) and (2), the partial likelihood method in (3) does not require any assumption over the form of the baseline hazard since they cancels out in the process of building a "partial" likelihood. This makes the method more robust to the specification error. However, it cannot practically accommodate a random heterogeneity term and the inclusion of time-varying covariates makes computation rather messy. The semi-parametric method, on the other hand, can handle the random heterogeneity rather easily. The semi-parametric method maintains the parametric forms of how the explanatory variables enter the hazard (i.e., $\exp(\gamma' z)$), while either the baseline hazard or the heterogeneity or both can be estimated non-parametrically. Essentially, they treat the mixing distribution or the baseline hazard as a step function and estimate the values of their steps along with the structural parameters (ie. γ). In this case, the baseline hazard estimator should ideally posses as many steps as there are distinct duration observations, although this induces inefficiency as such number increases. For example, this may not be suitable for our sample which contains 118 distinct duration observations. In practice, a number of steps to be estimated is allowed to be substantially smaller, but this time, the intervals between the steps and the number of steps may become a source of specification error. The random heterogeneity term, on the other hand, is considered estimable for up to 4 points of increase, according to the Monte Carlo experiments (Heckman and Singer 1984). The same simulation results also suggest that such nonparametric mixing distribution is poorly estimated compared to the structural parameters. At least, these estimators can be used to check the plausibility of the structural parameter estimates of the full parametric model, although there is no formal test statistics to compare the two. In practice, estimating the mixing distribution as well as the baseline hazard non-parametrically seems rather difficult (Meyer 1990). Moreover, they are not immune to the usual problem with the likelihood estimation of a mixture model: sensitivity to the starting values and multiple local maxima. Nonetheless, these methods are important because of their robustness to the specification errors, which too frequently overshadow the other estimation methods.

Considering the advantages and disadvantages of the several estimation methods discussed above, our methodology still bears the superior robustness in a sense that it is free of unnecessary assumptions over the random heterogeneity distributions both group and spell specific, and the duration distributions. It can also accommodate complications such as heterogeneity and time-varying explanatory variables in great ease. The latter is important since it allows the spell length to be influenced by the events occurring at any time after a negotiation, hence the hazard is not only a function of a duration but also of the entire time path of the explanatory variables during the course of a spell. In this way, out model embodies a functional relation rather than a function.

Furthermore, our regression format enables a simple test of lagged duration dependence, since our model naturally incorporates a multiple spell nature of our sample without having have to go through a complicated multiple integration to derive the marginal joint density. The drawbacks are the difficulties in identifying the pure duration dependence and in dealing with the time varying variables that are collinear with the elapsed durations during the spell. It is important to find some time varying explanatory variables that have sufficient variations across observations, yet posses a monotonic movement, so that they are identified to represent the underlying duration dependence. Censored duration observations are not considered and cannot be handled in our model. Nonetheless, for the purpose of analyzing a joint determination of a pair of variables, this system of the hazard and wage equation seems most adequate. The robust semi-parametric method, which estimates the baseline hazard and the heterogeneity non-parametrically, would be most compatible to our model in terms of the robustness with respect to the specification errors, but its estimation is very complicated by itself, hence, will not be estimated here.

7.3 Estimation

We have discussed the several different estimation methods, their likely shortcomings, their a priori assumptions, and restrictions. In this section, we carry out the actual estimation and find out the effect of the assumptions restricting each estimation method.

The explanatory variables used in the estimations hereafter are somewhat different from those used under our dynamic model in chapter 5 and 6, for they are constrained to be spell specific, so that they are not allowed to vary during the spell. The explanatory variables corresponding to each duration observations are those observed at the start of a spell. We assume, unless stated otherwise, that each observation are homogeneous in the unconditional probability distribution over the duration times except for these covariates, z_{ns}, which is invariant within a spell. Hence, there is no random nor fixed unobservable factor that represents group specific effect upon pooling of observations, and the group characteristics are assumed to be wholly captured by the Wage council, public sector dummies and the number of trade unions involved. Other explanatory variables involve spell specific factors representing: real take home pay, employers capacity to pay, outside influence, incomes policy, and other influences such as the dummies for the staged settlement. Following lists are the specific variables representing above factors actually used in the estimation. For the incomes policy dummies, we list the criteria for which a value 1 is assigned. W refers to wages, R to retention ratio and P to retail price index.

(1) *Group characteristic variables*

WC = 1, if the group is in the Wages Council sector

Pub = 1, if the group is in the public sector

TU = number of trade unions involved in the negotiating group

Spell specific variables

(2) *Real take home pay*

lnrw = $\log(W(T_{ns-1})R(T_{ns-1})/P(T_{ns-1})$

Drw1 = $W(T_{ns-1})R(T_{ns-1})P(T_{ns-2})/W(T_{ns-2})R(T_{ns-2})P(T_{ns-1}) - 1$

(3) *Employer's capacity to pay*

lnprin = log of industry's trading profit at T_{ns-1}

(4) *Outside influence*

rel = the ratio of average UK earnings at time T_{ns-1} to the group's wage, $W(T_{ns-1})$

lnunin = log of industry specific unemployment rate

(5) *Incomes policy*

Wage freeze

fd1 = 1, for public sector groups, if $T_{ns-1} \in$(Jul 1961 - Mar 1962)

fd2 = 1, if $T_{ns-1} \in$(Jul 1966 - Jun 1967)

Wage ceiling

cd1 = 1, for public sector groups if $T_{ns-1} \in$(Apr 1962 - Mar 1963)

cd2 = 1, for public sector groups if $T_{ns-1} \in$(Apr 1963 - Apr 1985)

cd3 = 1, for public sector groups if $T_{ns-1} \in$(Apr 1965 - Jul 1966)

cd4 = 1, if $T_{ns-1} \in$(Jul 1967 - Mar 1968)

cd5 = 1, if $T_{ns-1} \in$(Mar 1968 - Dec 1969)

cd6 = 1, if $T_{ns-1} \in$(Jan 1970 - Jun 1970)

Twelve-months policy

d12 = 1, if $T_{ns-1} \in$(Jul 1967 - Dec 1970)

(6) *Other factors*

size = number of workers covered by the n-th bargaining group as a proportion of total UK employment at T_{ns-1}

staged = 1, if the wage increase at T_{ns-1} is a part of a staged settlement

dfix = 1, if the date of wage increase at T_{ns-1} is fixed in advance as part of a staged settlement

lagdur = log of previous duration, δ_{ns-1}

In the following sections, we present the methodology and the actual model to be estimated for each method introduced in a former section and discuss their estimation results.

7.3.1 Linear regression

First of all, conduct a simple linear regression estimation. Assume a constant hazard function with a vector of explanatory variables, z_{ns}, unique to the n,s-th observation so that: $h(\delta_{ns}|z) = \exp(z_{ns}'\gamma)$. We are assuming also that these explanatory variables are time-invariant, hence, the hazard is fixed at any length of elapsed duration given a set of explanatory variables. The relationship between the hazard and the duration density is:

$$f_{\delta}(t) = h(t) (\exp (- \int_0^t h(\tau) d\tau)) \quad (7\text{-}3\text{-}1)$$

Transformation of a variable from t to u where $u = -\log(\int_0^t h(\tau)d\tau)$ gives a density of the transformed variable as:

$$f_u(u) = \exp(-u-\exp(-u)) \quad (7\text{-}3\text{-}2)$$

hence, u follows a type I Extreme value distribution with mean $\psi(1)$ and variance $\psi'(1)$, where $\psi(\xi)$ is the digamma function: $\psi(\xi)=d\ln\Gamma(\xi)/d\xi$. In this special case of a constant hazard with time-invariant explanatory variables, $\log t = -\gamma'z - u$, where u follows such an ancillary distribution which doesn't depend on the value of γ[1]. Then, allowing a conventional spell specific random error term, v_{ns}, a linear regression model to be estimated is:

$$\log \delta_{ns} = \gamma^{*\prime} z_{ns} + \varepsilon_{ns} \quad (7\text{-}3\text{-}3)$$

where $\varepsilon_{ns}=-u-\psi(1)+v_{ns}$, which is constructed to have a zero conditional mean. The estimated result is listed in Table (7.1). Note that variables that reduces duration length on average increases the hazard (i.e., $-\gamma=\gamma^*$).

The overall fit of the model is poor, but the directions to which the spell specific variables exert their effects on the hazard are rather similar to the dynamic model where (Table 5.6 or 5.4). Relative pay as well as the staged dummy exerts a strong shortening effect on the average durations. Staged settlement with fixed date has significant net negative effect as predicted. Spells started during the incomes policy in general have a tendency to last longer with one significant exception of fd_1. Since the incomes policy dummies indicate their status at the start of a spell and does not incorporate how long such policy have continued to last nor any possibility of other policy taking place during the contract, comparing their estimated coefficients with those derived from the dynamic model can be misleading. We will talk about this later in the conclusion section. We find no evidence of a strong lagged duration dependence. Puzzling effect of the log of real take home pay exists with a long run elasticity (-0.377). It implies that the lower net real pay at the start of the contract induces longer contract *ceteris paribus*, which is contradictory to the theory. Under this equation, it is not possible to differentiate the unexpected from the expected real take home pay, and how these covariates evolved once a spell has started is ignored.

A natural extension of the simple OLS estimates is the introduction of a group-specific effect, u_n, which is easily introduced under this regression framework:

$$\ln \delta_{ns} = \gamma^{*\prime} z_{ns} + u_n + \varepsilon_{ns} \qquad (7\text{-}3\text{-}4)$$

Some groups are predisposed towards longer negotiation intervals than the others often in a way not totally explained by the observable factors such as a degree of unionisation or a bargaining structure. Hence, it is important to allow for such a random group-specific effect although for other estimation methods, it is not so easy to implement. To make the matter more complicated, the existing explanatory variables, z_{ns}, which include lagged dependent variable, is likely to be correlated with u_n in our context. In order to avoid making an *ad hoc* distributional assumption over such a random term, we have estimated the differenced equation by using the same set of instruments as those in the dynamic hazard equation and achieve the consistent estimates (asymptotic is over the number of cross section observations). Recall the argument in selecting the appropriate set of instruments: a timing of observation at T_{ns-1} is endogenous to ε_{ns-1} (one of the error component in a differenced equation), and the exogenous variables observed at the endogenous timing should be treated as endogenous, therefore not suitable as instruments. The result is shown in Table (7.2). The estimated effects of the time-varying covariates are quite different from Table (7.1). In addition to the increase in efficiency, a strong positive lagged duration dependence and the significant and positive effect of the real take home pay at the start of the spell are observed. Relative wage still has a highly significant positive effect and Fd1 at T_{ns-1} no longer shortens the contract that followed. However, the Sargan's test statistic barely satisfies 5%

critical level for the correct specification. In particular, their strong effect of relative wages, overall incomes policies, and lagged positive duration dependence are all commonly observed with the estimates from the dynamic model. They both allow for a random spell specific as well as a group specific error term but no direct duration dependence. In the dynamic constant hazard specification, the hazard was allowed to vary with the passage of time through the time-varying explanatory variables. Here, all the variables are measured at the start of a spell, including Lnrw which should have a tendency to prolong a spell, *ceteris paribus*.

7.3.2 Parametric maximum likelihood

Under this estimation method, a parametric likelihood function for each observation is required which is then maximized with respect to the unknown parameters. This means that the complete parametric form of density function has to be assumed if not known a priori.

As before, we assume the proportionate hazard formulation: $h(t,z) = \exp(z_{ns}'\gamma)\ h_0(t;\alpha)$. Using the relationship between the hazard and the duration distribution, we derive the likelihood for the n,s-th duration observation as:

$$f_\delta(\delta_{ns}) = \exp(z_{ns}'\gamma)\ h_0\ \exp\left[-\exp(z_{ns}'\gamma)\ \left(\int_0^{\delta_{ns}} h_0\ dt\right)\right] \qquad (7\text{-}3\text{-}5)$$

Elements of the vector z are those known at T_{ns-1} (ie. start of the n,s-th spell), and their paths during the spell are considered invariant.

There are several choices one can adopt for the parametric form, h_0, which has to be somewhat arbitrarily determined to carry out the parametric ML estimation. The exponential duration distribution, $h_0=1$, is equivalent to the constant hazard where the hazard is fixed throughout the spell but its level is determined entirely by the covariates, z. This was also the case in the simple regression models of the former section. Then the log likelihood is:

$$L = \sum_n \sum_s [\ z_{ns}'\gamma - \exp(z_{ns}'\gamma)\ \delta_{ns}\ ] \qquad (7\text{-}3\text{-}6)$$

Assumption of the constant hazard is obviously too restrictive, all the more since the covariates which are in fact varying continuously are considered fixed during the spell. A monotonic duration dependence can be incorporated by assuming the Weibull duration distribution, $h_0 = \alpha\ t^{\alpha-1}$, which nest the exponential case when $\alpha=1$. For $\alpha>1$, the hazard is monotonically increasing. This parameter cannot be identified when a regression method such as those described in the former section is used. For now, neither of these allow for a group nor a spell specific random error term. With the Weibull assumption, the likelihood becomes:

$$L = \sum_n \sum_s [\ z_{ns}'\gamma + \log\alpha + (\alpha-1)\log\delta_{ns} - \exp(z_{ns}'\gamma)\ \delta_{ns}^{\alpha} \qquad (7\text{-}3\text{-}7)$$

The resulting estimates of the exponential and the Weibull duration distribution models are listed in Table (7.3) and (7.4), respectively. The estimated coefficients of the exponential hazard specification, Table (7.3), although very poorly estimated, strongly resemble that of the OLS estimates, Table (7.1). The latter does not allow for a random group specific term but a spell specific error term, while the former allows none. Note that their coefficients should be opposite in signs to imply the same direction of effects. Most remarkable feature of Table (7.3) is its strong positive duration dependence identified by $\hat{\alpha}=2.68$, although the overall effect of other variables, particularly of those significant, have remained rather similar in all three cases (ie. simple OLS, exponential MLE and Weibull MLE). Note also that under the Weibull baseline hazard, the direct effect of z on the log of duration is represented by γ/α rather than γ alone. Compared to the exponential MLE, the coefficients are more precisely estimated, and the major determinants are found to be staged, dfix, lnunin, WC and the incomes policies fd1, cd3 and cd4. Fd1 is the only incomes policy which gives a tendency to shorten a spell length on average. This finding was consistent in all three specifications, where its immediate effect on log duration is 2.1177/2.679=0.79 in the Weibull MLE, 0.697 in the exponential MLE and 0.567 in the OLS.

In the dynamic specification, explanatory variables were allowed to vary throughout the spell, amongst which the unexpected real take home pay, industry unemployment and most of the incomes policy variables gave particularly strong push towards ending the contract spell. Having taken into account of the time-varying economic environment, the average hazard computed at the completed duration observation was shown to decline, at least up to 2 years, with duration. Nonetheless, the way the hazard varies over time depended hugely on the time-path of covariates. Under the current specification, the way the hazard varies over time is solely determined by the Weibull duration distribution, a monotonic relation with the elapsed duration. Considering it was mainly the time path of variables *after* the negotiation that contributed in raising the hazard as the spell lengthened, it is likely that their omission creates a bias particularly in a direction of duration dependence. Look at figure 5.19 which plots average estimated hazard against the elapsed durations in months. This is based on the result of Table 6.2.4, the Weibull duration distribution without heterogeneity. Dotted lines are the interpolated locus of average estimated hazard, since there is no observation corresponding to some of the longer duration length. The figure suggests an overall positive duration dependence. Also, the average estimated density of duration in fig 5.20 depicts a very plausible peak at around 9-15 months, as suggested by the data. Comparing this to the figure 5.17, for example, which shows decreasing estimated average pdf at the completed duration observations, it may seem that the specification of the dynamic hazard model is unreasonable. The result of dynamic simulations reported in chapter 6.2, however, captures the incomes policy effect on the exit probability (ie. negotiation probability) far more responsively and reasonably than the prediction based on the Weibull duration distribution. The simulated exit probabilities based on the result of the MLE with Weibull are broadly invariant throughout the sample period, irrespective of the presence of incomes policy. Hence, their simulated wage levels are the result of settlements taking place almost randomly, with no clear impact of the policies on the timing of wage changes.

Apart from the potential misspecification caused by treating all the covariates as fixed, MLE estimates reported in Table (7.4) are liable to heterogeneity bias since the likelihood function (7-3-7) did not take into account of a term representing omitted spell specific attributes. As far as the finding of a positive duration dependence in Table (7.4) is concerned, however, the omitted heterogeneity, even if existed, would have given a downward bias to $\hat{\alpha}$ (Lancaster (1979)). If anything, the positive duration dependence is would only be strengthened.

Nonetheless, let us see what happens with the inclusion of such heterogeneity term in this formulation. First, we introduce a multiplicative heterogeneity term, μ, in the hazard, so that:

$$h_{ns} = \exp(z_{ns}'\gamma)\ h_0\ \mu$$

Allowing for such factor again requires an ad-hoc parametric assumption for its distribution, which, for the sake of convenience, is considered to follow the Gamma with an unit mean and variance, σ^2. Then, its distribution (ie. the mixing distribution), $f(\mu)$, is:

$$f_{\mu}(\mu) = \mu^{\sigma^{-2}-1} \exp(-\mu\, \sigma^{-2})\, \sigma^{-2} \qquad (7\text{-}3\text{-}8)$$

In order to derive the unconditional distribution of the duration, which is what we observe, it is necessary to integrate out for u, so that:

$$f(\delta_{ns} | z_{ns}) = \int_0^{\infty} f(\delta_{ns} | z_{ns}, \mu)\, f_{\mu}(\mu)\, d\mu \qquad (7\text{-}3\text{-}9)$$

Solving a differential equation after integrating by parts yields:

$$f(\delta_{ns} | z_{ns}) = e^{z_{ns}'\gamma}\, h_0\, [1 + \sigma^2 e^{z_{ns}'\gamma}\, I_0(\delta_{ns})]^{-\sigma^{-2}-1} \qquad (7\text{-}3\text{-}10)$$

Our data only contains the completed spells, hence, the log likelihood to be maximized is simply:

$$\sum_{n=1}^{N} \sum_{s=1}^{S_n} \ln f(\delta_{ns} \,| z_{ns})$$

where $f(\delta_{ns} | z_{ns})$ takes a form of equation (7-3-10). In table (7.5), we show the resulting estimates of the Weibull hazard with Gamma mixing distribution. This time, the log likelihood to be maximized is:

$$\log L = \sum_{n}\sum_{s} \left\{ z_{ns}'\gamma + \log(\alpha) + (\alpha-1)\log(\delta_{ns}) + (-\sigma^{-2}-1)\log(1 + \sigma^2 e^{z_{ns}'\gamma}\, \delta_{ns}^{\alpha}) \right\} \qquad (7\text{-}3\text{-}11)$$

As expected, a duration dependence parameter has increased, probably due to the significant heterogeneity term which was omitted in the previous model. Estimates of the structural parameters and their significance have not changed much, which should be the case if the covariates are independent of the omitted regressors. Apart from the heterogeneity variance and the duration dependence, major determinants are again, staged, dfix, lnunin, fd1, fd2, cd3, cd4 and the wages council dummy. Significance of the spell specific random error term is clearly seen, but the group specific random term is still not accounted for in this formulation.

At this point, with a spell-specific heterogeneity and a monotonic duration dependence allowed for, the static impact on log duration of potentially time-varying covariates (which are constrained to be fixed here) as well as the spell-specific covariates differ significantly from the dynamic results stated in table 5.6, or even with table 5.7 which assumes Weibull duration distribution with $\alpha=2$. It is hard to see if their differences are due to time-invariance of the covariates, lack of group-specific random terms or the misspecification of duration and/or mixing distributions. For example, starting a spell with staged, fixed settlement significantly shortens the contract duration on average in the MLE as much as 45% in MLE with Gamma heterogeneity-Weibull specification, 14% with Weibull specification. However, under the dynamic specification of chapter 5 (table 5.6), starting the contract with the staged settlement does shorten the contract by 41% but if its date is also fixed, it actually increases the contract length by 22%. The impact of the staged variable in the formulation of tables 7.3-7.6 may include the random group specific effect, since the lengthening effect of the staged dummies are also seen in table 7.2, which is the only model with the group specific random effect.

7.3.3 Partial Likelihood

Partial likelihood is a semi-parametric method in a sense that no parametric assumption needs to be made over a form of the baseline hazard since it cancels out as we write the "partial" sample likelihood. Here, we only utilize rank information on the duration observations in order to derive the inference on the structural parameters for given values of covariates. Each element of the partial likelihood represents a conditional probability that j-th spell ends at a certain duration, say t, given all the other observations which is at least as long as t. This conditional probability can be written in terms of the hazard as: $h(t|x_j(t))/\sum_{k\in R_t} h(t|x_k(t))$, which, in case of the proportional hazard, becomes: $\exp(x_j(t)\gamma)/\sum_{k\in R_t}\exp(x_k(t)\gamma)$, where R_t is the risk set just prior to the spell length, t. The risk set at t, R_t, contains all the observations with spell length of at least t. Log likelihood to be maximized for a set of observations (x_j,t_j) where $j = 1....M$ and $M = \sum_{n=1}^{N} S_n$, is:

$$\log L = \sum_{i=1}^{n} \sum_{j(t_j=t_i)} \left\{ x_j(t_i)\gamma - \ln \left[\sum_{k\in R_{t_i}} \exp(x_k(t_i)\gamma) \right] \right\} \quad (7\text{-}3\text{-}12)$$

where n represents a number of distinct duration observations in a sample. The second summation over j accommodates a possibility of ties in the observation, in which case, there exists more than one j with a duration observation with length t_i. Our sample has 118 distinct duration lengths out of 850 total observations. Estimation is carried out using the survival analysis routine of the LIMDEP programming package and is listed in Table (7.6).

Results are strikingly similar to Table (7.5) or (7.4). Considering that the partial likelihood is immune to the specification error involved in the baseline hazard, this may imply the appropriateness of the Weibull assumption, in particular, of the monotonic increasing duration dependence. Moreover, since the heterogeneity term in the MLE with the Weibull distribution specification only affected the parameter estimates of the duration dependence but not the other structural parameters, it seems that these partial likelihood estimates are not affected much by its omission of the spell specific heterogeneity term. Nonetheless, this does not preclude the possibility of: (1) duration dependence being produced primarily via the movement of time-varying covariates rather than the elapsed duration itself, (2) the significant omitted group specific effect that is creating bias in the structural parameters. What we seem to know instead, is that given the static covariates and no other group effect, the Weibull duration distribution is, after all, not a bad approximation.

7.4 Conclusion

The estimation of the static hazard models have, not surprisingly, revealed very different effect of the start-of-a-spell covariates as compared to their continuous counterparts. We have found the significantly increasing hazard overtime as well as the significant spell specific random heterogeneity factor. Starting with the OLS estimation of log durations, we have gradually generalized a model by introducing (1) the Weibull duration distribution, (2) the Gamma heterogeneity, and finally, (3) adopting partial likelihood thereby dropping an *ad hoc* assumption over the duration dependence. The structural parameter estimates are mostly consistent throughout these models, in particular, there seems to be a strong spell specific random factor and the positive duration dependence. Moreover, the assumption of the Weibull monotonic duration dependence assumed in the parametric MLE is supported by the result of the partial likelihood estimation, although this is still within the framework of the time-invariant covariates with no group-specific random factors. Major influences are as follows: 1% rise in the industry unemployment at the start of the spell shortens duration by 0.08%, the planned staged negotiation shortens the following contract length by as much as 45%. Most effective incomes policies are fd1, cd3 and cd4, with their immediate effect being 76% increase for the fd1 and 24% decline for either cd3 or cd4 on average duration. These figures are very different from the estimates derived from the dynamic hazard formulation of chapter 5.

These static models have revealed monotonically increasing hazard for all the duration observations. On the other hand, the dynamic models of the former chapters resulted in the hazard whose movement is largely affected by the time path of the explanatory variables during the spell. Whether this seemingly monotonic hazard is in fact generated via the movement of covariates overtime is never made clear under these static models. As a consequence, the effect of the spell specific structural parameters become biased. The implications of the time varying covariates can differ enormously whether it is the impact of their entire time path during the spell or their values at particular point in time. For example, consider the incomes policy dummies. The static model represents the effect on the termination probability of a contract which is started off during the policy-on period. On the other hand, the dynamic model depicts the instantaneous effect of the policy status any time during the contract. Hence, the latter takes into account of the timing of events that occur after the negotiation, such as a policy's termination or an introduction of another policy, while the former gives information of policy on/off at only one point, namely, at the start of the observed spell. For instance, the freeze policy, fd1, is considered successful towards the non-private sector groups for the first half of the policy period, namely, between July 61 and Jan 62. But a month prior to the end of the policy some groups started having wage rises. Then, as soon as it terminated, most of these groups enjoyed another wage rises together with many other groups who kept oath during the policy. As a result, those who had their wage rises *during* the fd1 had them towards the end of such policy and it was not long before they had another one. This has been reflected in the positive impact of fd1 on the hazard in the static model. The truth of a matter, however, is that the policy was unsuccessful only during the latter part of the policy period. Its overall impact on the hazard is likely to be ambiguous as we have found from the result of the dynamic specification model. In this sense, the static effect of the policy status at the beginning of a spell on the hazard can be largely misleading.

There seems to be another source of bias coming from a misspecification of the model, namely, the omission of a group specific random effect. A marked departure of the IV estimates on the differenced duration equation, which was the only equation with both group and spell specific random terms (ie. Table 7.2), from the other models calls for the significance of the group specific random effect that should not be ignored. In theory, existence of such effect is justified strongly. This brings about another complexity in the estimation procedure since the lagged durations and wage changes will not be independent of such group-specific effect. Unlike the maximum likelihood or any other parametric estimation methods, the regression format used in chapter 5 and Table 7.2 is particularly useful in dealing with this problem.

(Table 7.1)
Estimation of the duration equation[2]
(unskilled males; manufacturing and construction ind)

dependent variable : $\log(\delta_{ns})$

Variable	Coefficient	t-ratio
spell specific variables		
$\ln(\delta_{ns-1})$	-0.007	0.20
drw1	-0.986	1.58
staged	-0.454	5.04
dfix	0.380	3.58
lnrw	-0.380	2.43
rel	-0.721	3.69
lnunin	-0.060	1.60
lnprin	0.042	1.39
size	-14.19	0.87
fd_1	-0.567	3.59
fd_2	0.310	3.61
cd_1	0.120	1.16
cd_2	0.206	1.49
cd_3	0.286	2.18
cd_4	0.268	2.64
cd_5+cd_6	0.094	0.99
d12	0.052	0.46
group-specific effects		
constant	0.326	0.18
W council	0.145	2.37
Public	-0.048	0.31
TU	-0.032	1.93

$N = 61$; $\Sigma S_n = 850$; adjusted R-sq $= 0.169$; D-W st $= 2.018$

(Table 7.2)
Estimation of the duration equation[3]
(unskilled males; manufacturing and construction ind)

dependent variable : $\log(\delta_{ns})$

Variable	Coefficient	t-ratio
spell specific variables		
$\ln(\delta_{ns-1})$	-0.193	7.07
drw1	-0.287	0.60
staged	-0.120	0.90
dfix	0.467	4.54
lnrw	2.498	2.55
rel	-2.989	12.60
lnunin	0.133	2.97
lnprin	-0.534	1.77
size	208.649	2.66
fd_1	0.498	2.08
fd_2	0.564	3.08
cd_1	0.304	2.02
cd_2	0.669	4.95
cd_3	0.139	1.62
cd_4	0.256	3.10
cd_5+cd_6	0.510	6.46
dt12	-0.150	1.94

$N = 61$; $\Sigma S_n = 850$; Sargan's test $= 49.570$; p-value$=0.066$
Rssq $= 0.848$

group-specific effects

constant	27.113	68.38
W council	0.782	2.50
Public	-0.949	0.74
TU	-0.060	0.47

(Table 7.3)
Estimation of the duration equation[4]
(unskilled males; manufacturing and construction ind)

dependent variable : $\log(\delta_{ns})$

Variable	Coefficient	t-ratio
spell specific variables		
$\ln(\delta_{ns-1})$	0.016	0.09
drw1	-0.091	0.03
staged	0.351	1.36
dfix	-0.222	0.56
lnrw	0.380	0.30
rel	0.413	0.54
lnunin	0.069	0.36
lnprin	-0.230	0.18
size	2.753	0.05
fd_1	0.697	0.35
fd_2	-0.218	0.28
cd_1	-0.122	0.30
cd_2	-0.148	0.36
cd_3	-0.253	0.51
cd_4	-0.243	0.39
cd_5+cd_6	-0.069	0.13
d12	0.024	0.05

$N = 61$; $\Sigma S_n = 850$; log likelihood $= -4252.3$

group-specific effects

constant	-2.361	0.30
W council	-0.165	0.55
Public	0.027	0.04
TU	0.023	0.25

(Table 7.4)
Estimation of the duration equation[5]
(unskilled males; manufacturing and construction ind)

dependent variable : $\log(\delta_{ns})$

Variable	Coefficient	t-ratio
spell specific variables		
$\ln(\delta_{ns-1})$	0.054	0.82
drw1	1.249	1.35
staged	0.754	5.41
dfix	-0.379	2.06
lnrw	0.594	1.24
rel	0.415	1.37
lnunin	0.242	3.61
lnprin	-0.066	1.20
size	-21.52	0.67
fd_1	2.118	5.90
fd_2	-0.429	2.64
cd_1	-0.675	1.84
cd_2	-0.739	1.38
cd_3	-0.595	3.37
cd_4	-0.705	3.92
cd_5+cd_6	-0.178	1.21
d12	0.231	1.46
α	2.679	28.99

$N = 61$; $\Sigma S_n = 850$; $\log L = -3825.0$

Variable	Coefficient	t-ratio
group-specific effects		
constant	-8.285	2.70
W council	-0.459	4.36
Public	0.056	0.23
TU	0.032	0.98

(Table 7.5)
Estimation of the duration equation[6]
(unskilled males; manufacturing and construction ind)

dependent variable : $\log(\delta_{ns})$

Variable	Coefficient	t-ratio
spell specific variables		
$\ln(\delta_{ns-1})$	0.056	0.71
drw1	1.142	1.06
staged	0.893	5.04
dfix	0.478	2.18
lnrw	0.714	1.33
rel	0.786	1.95
lnunin	0.236	2.91
lnprin	-0.067	1.09
size	-11.23	0.31
fd_1	2.304	5.10
fd_2	-0.546	2.70
cd_1	-0.273	0.67
cd_2	-0.284	0.70
cd_3	-0.729	3.29
cd_4	-0.739	3.40
cd_5+cd_6	-0.154	0.89
d12	0.166	0.87
α	3.028	16.16
sigma	0.425	4.94

$N = 61$; $\Sigma S_n = 850$; $\log L = -3816.5$

group-specific effects

Variable	Coefficient	t-ratio
constant	-9.462	2.69
W council	-0.522	4.13
Public	0.017	0.06
TU	0.057	1.55

(Table 7.6)
Estimation of the duration equation[7]
(unskilled males; manufacturing and construction ind)

dependent variable : $\log(\delta_{ns})$

Variable	Coefficient	t-ratio
spell specific variables		
$\ln(\delta_{ns-1})$	0.073	1.03
Drw1	2.285	2.14
staged	0.758	6.06
dfix	-0.370	2.15
lnrw	0.552	1.09
rel	0.645	2.07
lnunin	0.263	3.55
lnprin	-0.089	1.66
size	-19.87	0.81
fd_1	2.334	4.44
fd_2	-0.411	1.78
cd_1	-0.175	0.85
cd_2	-0.199	0.84
cd_3	-0.624	3.09
cd_4	-0.771	3.42
cd_5+cd_6	-0.162	0.83
d12	0.253	1.26

$N = 61$; $\Sigma S_n = 850$; $\log L = -4820.4$

group-specific effects

Variable	Coefficient	t-ratio
W council	-0.521	4.30
Public	0.009	0.04
T Union	0.041	1.23

Footnotes to chapter 7

1. Although it depends on the correctness of the formulation for the hazard from which this transformation originated.
2. Ordinary Least square estimation. White's heteroscedasticity robust standard errors reported.
3. Instrumental fixed effect estimation (differenced) using the same set of IV used to estimate the dynamic hazard function. 2nd step estimates reported.
4. Maximum likelihood estimation, exponential duration dependence assumed.
5. Maximum likelihood estimation, Weibull duration distribution assumed so that:
   $h_0 = \alpha \, t^{\alpha-1}$
6. Maximum likelihood estimation, Weibull hazard with Gamma heterogeneity.
7. Partial likelihood estimation.

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