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NOMINAL WAGE RIGIDITY: NON-PARAMETRIC TESTS BASED ON UNION DATA FOR CANADA

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Abstract

We study the wage-change distributions in union contracts reached in Canada between 1976-1999. We use non-parametric tests to check for nominal wage rigidity and find that it is present during low inflation periods.

KEYWORDS: Nominal wage rigidity, non-parametric tests.

JEL Classification: E31, J31.

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1. Introduction

The wage determination process is one of the most studied areas of empirical labour economics. Despite continuing work in this area, long-standing questions concerning the extent of downward nominal rigidity, an issue of fundamental importance to labour economics and industrial relations, remain. If pervasive, such rigidity would interfere with the functioning of the labour market, preventing the efficient re-allocation of labour from low to high-demand areas and inducing quantity adjustments and unemployment. Should nominal rigidity be more prevalent in some sectors than in others, similar shocks will have different price and quantity effects. For instance, if unions are more resistant to wage cuts than the non-union sector, real wage realignment may be more difficult to achieve in the union sector. This will be all the more so at times of low inflation because then inflation cannot 'grease' the wheels of the labour market.[1] Thus, recently achieved, exceptionally low, levels of inflation in countries such as Canada may have been attained at the expense of higher unemployment. Under these circumstances, low-inflation regimes may inject new sources of stress in industrial relations. These arguments suggest that more information on the extent and pattern of downward nominal rigidity would be valuable.

In conventional Keynesian models, downward rigidity is 'effective' when the real wage is too high, employment is on the labour demand curve, and unemployment prevails. Then shocks which raise the price level and lower the real wage increase employment. Thus, early attempts to gage the severity of downward nominal rigidity were macroeconomic in nature and investigated whether the real wage is countercyclical. Papers from Dunlop (1938) and Tarshis (1939)

[1] See Shultze (1959), Samuelson and Solow (1960) and Tobin (1972).

to Solon, Barsky and Parker (1994) and Abraham and Haltiwanger (1995) are in this tradition and a variety of results are available.

However, a new literature stemming partly from the availability of data at the micro level has emerged.[2] These studies typically start by constructing the cross-sectional nominal wage-change distributions from data such as the Panel Study of Income Dynamics or the Current Population Survey. Annual histograms are then used to study features of interest such as whether the mass to the left of zero is deficient relative to a no rigidity counterfactual, whether spikes at zero can be identified, the extent to which holes around zero may suggest the presence of 'menu costs', and whether wage-change distributions may be different in periods of high and low inflation. Issues of concern center around the extent to which periods of sufficiently low inflation have been examined, the role of measurement, timing, and rounding errors inherent in these surveys, the extent to which the visual evidence presented amounts to statistical tests, and whether such tests are best conducted using parametric or non-parametric techniques.

Parallel with this literature has been work[3] that seeks the reasons for nominal rigidity by interviewing the individuals who ought to know, e.g. executives and labour leaders. Bewley (1999) suggests that nominal wage cuts are shunned because of their likely impact on morale and that this is all the more likely where information flows are good. Bewley (1999) finds that, in the 'primary'

[2]The US literature includes, *inter alia*, McLaughlin (1994), Lebow, Stocton and Wascher (1995), Akerlof, Dickens and Perry (1996), Card and Hyslop (1997), Kahn (1997), Altonji and Devereux (1999), Groshen and Schweitzer (1999), Lebow, Sacks and Wilson (1999) and McLaughlin (1999a, b). Smith (1999) studies experience in the UK, Beissinger and Knoppik (2000) that in West Germany, and Fehr and Gotte (2000) that in Switzerland. The extant Canadian literature is reviewed below.

[3]See Blinder and Choi (1990), Agell and Lundborg (1995), Cambell and Kamlani (1997) and Bewley (1999).

sector[4], new employees are more likely to be hired at rates comparable to those of existing employees than is the case in the 'secondary' sector where short-term employees, often part-time, abound. This work suggests that the incidence of downward nominal rigidity should be most apparent in situations where long-term relations between a firm and its employees exist, where workers are organised into bargaining units where 'bad news travels fast' and particularly so when the bargaining unit is a union whose very existence and *modus operandi* stress wages over employment[5] and the prevention of outcomes such as nominal pay cuts. Some studies based on survey data have distinguished between the behaviour of union and non-union workers and some evidence has been provided that more rigidity exists in the union sector.[6]

A good source of information on outcomes in the union sector is collective bargaining agreements themselves. These are legally binding documents whose provisions are recorded electronically by federal authorities and are, therefore, very accurate. Thus, one of the controversies surrounding survey data, namely. whether measurement, rounding and timing errors exist, does not apply.[7] In addition, the Canadian data is available over a long period of time which includes periods of high inflation, a period of substantially reduced inflation, as well a period during which inflation was exceptionally low and much lower than in the

[4]'Primary-sector personnel include most factory, clerical, and secretarial workers, technical, professional, and managerial employees with permanent positions, and salespeople in stores and restaurants with regular customers whom the staff should know on a first-name basis.' Bewley (1999, 18).

[5]See Oswald (1993).

[6]For instance, McLaughlin (1999a, 129) finds that '... the skewness of union workers' wage changes is all atributable to nominal rigidity'.

[7]For the significance of these issues for the size of the spike at zero in the context of British data, see Smith (1999).

US. Thus, the issue of whether periods of exceptionally low inflation have been available for study does not arise either.

In this study we use a recently released version of the Canadian contract file[8] to study the implied distributions of nominal wage change over the period 1976-1999. We use a variety of non-parametric techniques and statistical tests. In section 2, we discuss the data set used and its basic features. In section 3, we discuss the salient features of the annual histograms constructed. In section 4, we consider the non-parametric test procedures used to examine the nominal wage-change distributions. Finally, in section 5, we present a summary of our results and our conclusions.

[8] Some other studies also use the Canadian contract data in this general context. Fortin (1996) argues that the Canadian recession of the early 1990s was deeper in Canada than in the US because of the conjunction of lower inflation and downward nominal rigidity. This last claim is based on 1992-94 histograms of the *first* year of wage settlements, a procedure criticised by Freedman and Maclem (1998). Simpson, Cameron and Hum (1998) estimate the increase in the unemployment rate that would be needed to moderate wage inflation by the amount attributed to wage rigidity. Their conclusion that this could be as high as 2% is questioned by Fares and Hogan (2000). Fares and Lemieux (2000) also focus on the macroeconomic consequences of nominal rigidity. Crawford and Harrison (1998) present histograms of nominal wage change in private and public sector union contracts. They calculate the skewness coefficients at times of high, medium and low inflation. Surprisingly, these coefficients become more negative at times of low inflation, though this result may not be statistically significant. Crawford and Harrison (1998) also apply hazard methods to their data and investigate whether the wage-change hazard depends negatively on the rate of inflation.

2. Data Sources

Wage agreements in the unionised sector are monitored by Human Resources Development Canada (HRDC) who made available to us[9] a detailed, monthly, file containing information on provisions for 10947 wage contracts signed in the Canadian unionised sector, both public and private, between 1976 and 1999. Because reporting requirements apply, this information is be very accurate. We detected inconsistencies in only two contracts which were excluded from the sample. The raw, monthly, file was processed to extract the information needed for the purposes of this study including the unique identifying code number for each contract, relevant dates[10], wage change that was due to a COLA clause and wage adjustment that was not contingent, as well as the duration[11] and sector[12] of each agreement. While the contract data pertain only to the unionized sector, this is much more significant in Canada than in the US.

The resulting data base involves settlements which range in duration from a few months to several years, and covers bargaining units involving 200 to nearly 80,000 employees. The average base wage rate paid to entry-level workers is \$12.40 at the beginning and \$13.49 at the end of these agreements, implying a rate of change of 8.79%. Since mean duration is approximately two years, the

[9] We are ndebted to Michel Legault of HRDC for providing us with the raw, monthly, data file.

[10] These include the settlement, effective and expiry date of each contract. The settlement date is used to date the contract in the histograms below and to relate its provisions to the rate of price inflation.

[11] Contract duration is defined as the expiry minus the effective date.

[12] The private-public sector distinction is based on the 1980 Standard Industrial Classification. The public sector is defined widely to include contracts in Health and Education because of the substantial involvement of governments in the provision of these services. Thus, contracts with SIC numbers between 800-830, inclusive, were deemed to be in the public sector.

annual rate of wage adjustment is approximately 4.4%. The increase in the base wage rate is, on average, \$1.09 and it consists of a \$0.97 non-contingent increase and a \$0.12 contingent increase through a COLA clause. Very few contracts contain COLA clauses[13]. We pursue our analysis using two definitions of wage adjustment, that is one that includes COLA adjustments and one that does not. Clearly, less rigidity will be displayed by the former series. It should be noted, however, that because the incidence and intensity of COLA clauses is limited, the results are not very sensitive to this distinction. To conserve space, the histograms in Figure 1 below refer only to the series which includes COLA adjustment as this is likely to provide a more complete characterisation, but histograms based on the alternative definition are very similar and are available on request. In this paper, wage change refers to what the negotiating parties implement over the whole contract *at annual rates* and appears in our sample as one observation for each contract.[14] Contract re-openers, lump-sum payments and profit-sharing are very rare and are not taken into account.

Table 1 below contains, for each year,[15] the number of all contracts, as well as the number of contracts by sector and contract duration. A total of 10945 contracts are spread over the 21 'years' in Table 1, with a low of 226 contracts

[13] The nature, incidence, and intensity of COLA clauses and their implications, particularly for modelling wage adjustment, are analyzed in, *inter alia,* Card (1983,1986), Christofides (1987,1990), Cousineau, Lacroix and Bilodeau (1983), Ehrenbrerg, Danziger and San (1984), Hendricks and Kahn (1985), Kaufman and Woglom (1984), Mitchell (1980), and Vroman (1984).

[14] An alternative approach involves defining sub-periods of the contract and establishing wage adjustment over each of these. For a discussion of this issue, see Fortin (1996) and Freedman and Macklem (1998). We prefer the current specification because it summarises the overall intentions of the contract.

[15] Because of the smaller number of contracts, the first two and the last three years in the sample are considered together in everything that follows.

in 1977 and a high of 676 contracts in 1984. Despite the broad definition of the public sector, it includes less than half the agreements reached in the private sector. Considerably more contracts are long than short and the modal length is two years. Further descriptive statistics, including the rate of Consumer Price Index inflation as well as mean and median wage adjustment appear in Table 2 below. Figure 1 and Tables 1 and 2 are considered in detail in the next section.

3. Features of Wage-Change Distributions

Figure 1 presents wage-adjustment histograms for each of the 21 year groups in the sample. In constructing these, care was taken to centre the bins on zero.[16] During the high-inflation years of 1977 to around 1983, the histograms are centered well to the right of zero. Until 1982, average wage growth exceeded the rate of CPI inflation - see Table 2. The histograms are reasonably symmetric and display no pronounced spikes at zero. A substantial portion of the wage settlements in each year imply negative *real* adjustments which can be quite substantial. For instance, in 1977 when inflation was 7.55% and the average wage change including COLA was 8.69%, nearly 40% of the contracts entailed real wage reductions, some of which were as high as nearly 8%. Few contracts involved *nominal* wage reductions. In 1983, when CPI inflation was 10.8% and the average wage adjustment including COLA was only 4.89%, approximately 95% of the contracts entailed real wage reductions, some as high as nearly 11%. Again, few nominal wage reductions were implemented.[17]

[16] That is, the zero interval is -0.5 to 0.49999. Further intervals increase and decrease in 1% units.

[17] It should be noted, however, that not all high-inflation years involved a substantial proportion of real wage reductions. During 1980 and 1981, less than 10% of settlements entailed real wage reductions.

When inflation began to abate during and after 1983, but before the period 1988-90 when average real wages began to increase again, the general appearance of the histograms changed noticeably: In this period, they are characterized by noteworthy mass and censoring at zero and virtually no nominal wage decreases. In 1987, the percentage of contracts which involved real wage decreases was, at 44%, lower than in 1983, a year of higher price inflation; the maximal real wage decrease was of the order of 4%, much lower than in the high-inflation, period of 1977-83.[18]

During 1988-90, average wage adjustment increased and actually exceeded $C\dot{P}I$ (Table 2). Histograms for these three years are quite symmetric and the descent to zero reasonably smooth. Despite the fact that wage and price inflation are considerably lower during 1988-90 than during 1977-83, these histograms are similar in general appearance to that for 1978, for example, and appear to have been substantially influenced by the easing of labour market conditions. The proportion of contracts that involved real wage decreases was, on average, 14% during this period, a number smaller than in either of the earlier periods. The maximal real wage decrease was of the order of 5 to 6%.

After 1991, wage and price inflation declined to levels which are unprecedented in recent decades and much lower than those in the US. It is histograms like those for 1991 and 1992 that led Fortin (1996) to argue that extensive nominal wage rigidity was present in the Canadian economy.[19] These histograms display considerable mass and very strong censoring at zero. Indeed, the concentration of

[18] On average, the proportion of contracts which involved real wage decreases was 52% during 1984-87 as opposed to 41% during the earlier high-inflation period of 1977-83. We return to this point below.

[19] Fortin (1996) notes that the the decline in Canadian wage and price inflation may afford a much better opportunity to study low-inflation behaviour than is possible using US data.

mass at zero is so pronounced that even though the rate of CPI inflation was extremely low, a very substantial proportion of contracts experienced real wage declines. The average for the period 1991-97 is approximately 53%, a number higher than that prevailing during any of the earlier sub-periods, including the high-inflation period of 1977-83. Naturally, the extent, as opposed to the incidence, of real wage reductions was limited by the fact that CPI inflation was exceptionally low; real wage reductions were at most 1-2%. Nominal wage reductions were the exception rather than the rule, though it should be noted that, in the exceptional year of 1994, some 12% of contracts involved nominal wage reductions.

The apparent absence of nominal wage reductions from the histograms of Figure 1, may raise the concern that the -0.5 to 0.4999 bin may hide a substantial number of very small nominal wage reductions. This is not the case. As Table 2, columns 3, 4, 9 and 10 show, most of the mass in the zero bin is in fact at zero itself. For instance, in 1995, the year with the lowest CPI inflation rate, 38% of the mass in the wage-change distribution that includes COLA is in the zero bin and 35% is at zero itself.

The histograms in Figure 1 and, in particular, their evolution through different inflation periods, show clearly that nominal wage rates are generally downwardly rigid so that the 'inflation as grease' argument is worthy of serious consideration. Our discussion above suggests that the proportion of contracts which entail real wage reductions does not decrease as inflation abates. Indeed, in our data, it is highest during 1991-97 precisely because of the substantial concentration of nominal wage-change mass at zero. Nevertheless, the *extent* of real wage reductions that is possible in the low inflation period is limited by the conjunction of a low level of CPI inflation and downward nominal rigidity. Thus, while real wage reductions of the order of 10 to12% were possible in the high-inflation years of 1981-83, these reductions were limited to about one tenth of this amount during the low inflation period. One suspects that, unless a case can be made that the need for real wage realignment is lower during periods of low inflation, this reduced scope for real wage reductions may have some impact on the smooth functioning of the labour market.

In view of the fact that the visual evidence in the histograms of Figure 1 is consistent with the notion that nominal wages may be rigid downward, it is important to turn to some more formal statistical tests of features of interest in the wage-change distributions. These deal with whether the visual indications, in Figure 1, of concentration of mass at zero could have arisen by chance. In the next section, we consider the role of the symmetry of wage-change distributions, as well as behaviour at and below zero. Following a discussion of the relevant issues and tests in the first two subsections, we present the econometric results obtained.

4. Test Statistics and Results Obtained

4.1. Is There a Role For Symmetry?

A feature of the new literature dealing with nominal wage rigidity is its concern with the symmetry of the wage-change distribution. During a period of high inflation, the nominal wage change distribution may be symmetric around some measure of inflation plus average productivity growth.[20] By contrast, when for given average productivity growth inflation is low, some sectoral shocks may re-

[20] Whether this distribution will be symmetric will depend largely on the distribution of idiosynchratic shocks and on whether the overall mean is high enough to make nominal wage cuts unnecessary.

quire decreases in nominal wage rates. *If* nominal wage rates are rigid downward, the wage-change distribution may display considerable mass at zero and is likely to be asymmetric, given that nominal wage decreases will be censored at zero.[21]

While nominal wage rigidity is likely, at times of low inflation, to induce asymmetries in the wage-change distribution, rejection of symmetry need not be due to rigidities and care must be taken in interpreting results dealing with this issue. To begin with, it is conceivable that the distribution of sectoral productivity shocks is systematically altered by inflation so that it is symmetric during high-inflation periods and asymmetric during low-inflation periods. In that case, low-inflation asymmetries may be due to changes in the density function of the underlying productivity shocks. We are not aware of empirical or theoretical work which demonstrates that this is the case. Thus, if nominal rigidity exists, we would expect low-inflation distributions to be more asymmetric than high-inflation distributions and we consider how best to test for this possibility. A second possible complication may be that some high-inflation distributions may themselves be asymmetric[22]. Although this point cannot be ruled out, we do not believe it renders testing for symmetry uninformative. Rather, it is a matter of following an appropriate methodological approach. We would expect statistical tests to produce more emphatic rejections of symmetry during periods of low-inflation. If this does not occur, and we cannot reject symmetry at times of very low inflation, then the case for nominal rigidity is weak or non-existent.

A number of papers have considered the issue of symmetry using the skewness

[21] As Crawford and Harrison (1997) note, "Asymmetries in the wage change distribution have been proposed as another test of nominal rigidity."

[22] McLaughlin (1999a, 130) provides possible reasons for asymmetries. On the other hand, Card and Hyslop (1997, 86) note that '... most conventional models of wage determination imply symmetry'.

coefficient, median-mean differences, and other such measures - see McLaughlin (1999a,1999b). We, too, consider a number of these measures. However, histograms may be sensitive to bin choices, the skewness coefficient depends on distributional details and may change sign with minor changes in density[23], and the median-mean test treats the median as a known parameter. For these reasons, we prefer to use the kernel-based, non-parametric, test proposed by Ahmad and Li (1997). It does not rely on any distributional assumptions and it is a consistent symmetry test against all possible alternatives.

Let $X_1,...,X_n$ denote a random sample drawn from F, an absolutely continuous distribution with its density function f defined on the real line R. For simplicity of exposition and without loss of generality we assume, for now, that the density function is centered on zero.[24] The null hypothesis is expressed as $H_0 : f(x) = f(-x)$ for all $x \in R$, whereas the alternative is that H_0 is false, i.e., $H_1 : f(x) \neq f(-x)$ for some $x \in R$. This is an intuitively appealing and very direct way of constructing a test statistic for the symmetry of a distribution. The procedure we use tests the hypothesis that a distribution is symmetric about its median. The test is based on kernel estimation and, when appropriately centered and scaled, is asymptotically standard normal under the null hypothesis.

The test by Ahmad and Li (1997) is based on a measure of departure from the null hypothesis provided by the $L_2 - norm$. This is defined as

$$\int_{-\infty}^{\infty} [f(x) - f(-x)]^2 dx. \tag{4.1}$$

To test for symmetry, define the functional:

[23] See Mood, Graybill and Boes (1974).

[24] Ahmad and Li (1997) show that the asymptotic properties of the test statistic are valid when the data are centered around an estimated mean or median.

$$I = \frac{1}{2}\int_{-\infty}^{\infty}[f(x) - f(-x)]^2 dx = \int_{-\infty}^{\infty}[f(x) - f(-x)]dF(x). \qquad (4.2)$$

The above is zero if and only if $f(x) = f(-x)$ almost everywhere and positive otherwise. Hence, it constitutes a reasonable measure of departure from the null hypothesis, when the latter is false. The test statistic is based on a direct estimator of (2) above

$$\begin{aligned} \widehat{I}_n &= \int_{-\infty}^{\infty}[\widehat{f}_n(x) - \widehat{f}_n(-x)]dF_n(x) = \\ & (n^2a)^{-1}\sum_{i=1}^{n}\sum_{j=1}^{n}[K(\frac{X_i - X_j}{a}) - K(\frac{X_i + X_j)}{a})], \end{aligned} \qquad (4.3)$$

where $\widehat{f}_n(x) = (na)^{-1}\sum_{i=1}^{n} K(\frac{x-X_i}{a})$ is the kernel estimate of $f(x)$ and $F_n(x)$ is the empirical distribution function based on the sample. The function $K(.)$ is the kernel function, a known density symmetric about zero, and $a = a_n$ is a sequence of smoothing parameters (bandwidths) such that a_n approaches zero as the sample size n approaches infinity. The test statistic is asymptotically normally distributed, see Ahmad and Li (1997) for details. We present results involving this test-statistic in section 4.3.

4.2. Tail Behaviour

Another feature of nominal wage-change distributions that is of considerable interest is the extent to which mass at and below zero is unusually high relative to some benchmark. Card and Hyslop (1997) assume that the area above the median may be used as the no-rigidity counterfactual for the area below the median. They then measure the extent of nominal wage rigidity by subtracting an appropriate integral of this counterfactual area from that of the actual density function to the left of the median. McLaughlin's (1999a) symmetrically differenced histograms are similar in spirit except that they refer to the entire range of the distribution. Lebow, Stockton and Wascher (1995) also propose a similar descriptive device that measures the difference between the value of the cumulative density function at zero and the integral to the right of the point $2 \times Median$.

A test statistic based on these ideas is the difference of the empirical distribution function above twice the median and below zero. Under the null hypothesis of no downward rigidity, observations above the twice the median point will occur with the same probability as observations below zero and the test statistic will be centered around zero. Using as an estimate of $F_n(x)$, $\widehat{F}_n(x) = (nn_x a)^{-1} \sum_{i=1}^{n} \sum_{j=1}^{n_x} [K(\frac{X_i - X_j}{a})]$, where n_x denotes the observations up to point x, we define the test statistic as

$$\widehat{D}_n = \{1 - F_n(2medx)\} - F_n(0)$$

$$\widehat{D}_n = \{1 - [n(n_{2medx})a]^{-1} \sum_{i=1}^{n} \sum_{j=1}^{n_{2medx}} [K(\frac{X_i - X_j}{a})]\} - [n(n_0)a]^{-1} \sum_{i=1}^{n} \sum_{j=1}^{n_0} [K(\frac{X_i - X_j}{a})], \tag{4.4}$$

where n_{2medx} and n_0 denote the number of observations up to the point of twice the median and zero respectively. The distribution of $\widehat{D}_n$ based on the comparison of two population proportions is straightforward to construct and follows the standard normal variate. The smoothing parameter is chosen as $a = s_x n^{-\frac{1}{\alpha}}$, where s_x denotes the standard deviation of the sample data.

In standard density estimation α is chosen usually to be 5. However, evidence from simulations by Ahmad and Li (1997) suggest a larger value of α than 5, since using the latter results in test statistics that tend to reject the null hypothesis of symmetry too often. A larger value of α results in greater smoothing, and hence favors the null hypothesis of symmetry. Hence, in the following section, we present results for $\alpha = 8$ and 10 for both of the above tests. The results are based on the normal kernel. Note also that the symmetry test-statistic $\widehat{I}_n$ is a one-tail test while $\widehat{D}_n$ is two-tailed. However, in the latter case only rejections with the negative values of $\widehat{D}_n$ would signify evidence of wage rigidity, since in that case the tail at and below zero would dominate the tail above twice the median.

4.3. Results Obtained

Table 2 presents some of the features of the histograms in Figure 1 in greater detail as well as results based on the $\widehat{I}_n$ and $\widehat{D}_n$ test-statistics. Column 3, Table 2, shows the mass in the zero bin of the COLA histograms in Figure 1. As already noted, most of this mass is at precisely zero (column 4, Table 2). Columns 9 and 10, Table 2, show similar information for histograms similar to those in Figure 1 which refer to wage settlements that do not include a COLA clause. The respective columns contain similar results because indexed wage adjustment is small. It is clear that the amount of mass at zero increases dramatically in the low-inflation years. Indeed, in 1993, over half the mass of the wage-change distribution is at zero. Unlike the results from surveys of individuals analysed in previous studies, the incidence of censoring at zero in union contracts is very high. It is conceivable that some of this may be weakened through wage drift, benefit changes and other such measures but the morale stories in Bewley (1999) would suggest that, if such pressures were at all important, they would be effected through the more visible wage rate.

For the reasons noted earlier, we have not calculated skewness coefficients but we report medians and means because they have already been supplied as part of Figure 1. When COLA adjustments are included, the mean exceeds the median in all years except during 1982 to 1986. When COLA adjustments are excluded, the suggestion of a right-skewed distribution in the low-inflation years continues but, in most earlier years, the relative size of the mean and the median is quite erratic.

The Ahmad and Li (1997) statistic rejects symmetry most decisively in the low-inflation years of 1991-97 and quite decisively when wage inflation subsided after 1982. It suggests that the null hypothesis of symmetry can be rejected in some other years as well but the predominant pattern is the one just noted. This pattern receives support from a regression, using the 21 'years' in the sample, of the $\widehat{I}_n$ statistic on a constant and the rate of price inflation. When all contracts are used (columns 4-8, rows 1 and 5, Table 3) the slope in this simple regression equation is negative and significantly different from zero at the 5% level.[25] The results suggest, therefore, that the accumulation of mass and censoring at zero result in asymmetries that are strong enough to be picked up by the Ahmad and Li (1997) test statistic and that this is all the more so in periods of low inflation.

Having dealt with this particular concern, we now turn to the issue of whether the mass at and below zero[26] is larger than what might be suggested by the right tail of the distribution. If so, the statistic $\widehat{D}_n$ should, as measured, be negative and we might expect it to be all the more so as inflation abates.[27] Columns 7 and 13, Table 2, indicate that this is so. The null hypothesis of equality of mass in the

[25] The suggestion that as price inflation subsides the $\widehat{I}_n$ statistic increases is not supported in the smaller samples that result when the data is broken down by sector and contract duration.

[26] As already noted, with the exception of 1994, evaluating the cumulative density function at zero is essentially tantamount to measuring mass at zero.

[27] That is, in the simple regressions of Table 3 a positive slope coefficient should be expected.

two tails is rejected in both columns only during the declining (1984-86) and low (1991-97) inflation years.[28] In 1995, for instance, the distribution that includes COLA indicates that the left-hand side contains 11.72 percentage points more mass than the counterfactual, right-hand, tail. The results in Table 3 suggest, even in the sub-samples, that as inflation moderates, $\widehat{D}_n$ declines. Thus, the accumulation of mass at zero is statistically significant and this is all the more so when inflation is low.[29]

These results are independent of the indexation provisions in the contracts under study because indexation is not prevalent and, where it exists, it is not strong. The results, in Tables 2 and 3, are similar whether cost-of-living adjustments are included in the definition of wage adjustment or not. In 1993, for instance, when the spike at zero was 51% with COLA included, it is 53% when COLA is excluded. In addition, the patterns in the $\widehat{D}_n$ and $\widehat{I}_n$ test-statistics are similar across the two definitions; as one might expect, however, the definition which excludes COLA generally tends to produce absolute values of the test-statistics which are larger than when COLA is included.

When the results, in Table 3, are broken down by sector (public and private)

[28] Note that when the median of the wage-change distribution is itself zero, as in 1993-94, the $\widehat{D}_n$ statistic is not defined. We have added dummy variables, which equal 1 in relevant years and equal 0 otherwise, to the simple regression equations underlying Table 3 - see the notes in that table. The results are almost identical when the observations for relevant years are omitted.

[29] McLaughlin´s (1999a, 127-8) conjecture that the relation between $\widehat{D}_n$ and the rate of inflation will probably be too strong cannot be convincingly evaluated. One could easily establish the extent to which a measure such as $\widehat{D}_n$ changes as the asymmetric distribution, which purportedly characterises the no-rigidity case, shifts to the left. This information could then be included as an additional regressor in Table 3. However, the asymmetric distribution chosen is arbitrary and may not apply at each point in time. Note that McLaughlin´s (1999a) conjecture does not apply to symmetric distributions.

and length of contract (one year or less and over one year), the symmetry test $\widehat{I}_n$ is calculated for samples which are much smaller (especially short contracts) and, in the simple regression equations of Table 3, the coefficient on CPI inflation is not significantly different from zero. As far as the tail test $\widehat{D}_n$ is concerned, there does not appear to be much of a difference between behaviour in the total sample, the public and private sub-sectors and short and long contracts. In all cases, the coefficient on the rate of inflation is positive and significantly different from zero for all but short contracts.

5. Conclusions

To the extent that downward nominal rigidity is present, it is more likely to be prevalent in the union sector. We use data from collective bargaining agreements reached in Canada between 1976 and 1999 to study the implied wage-change distributions using a variety of non-parametric techniques and tests.

The period under study may be divided into sub-periods of high (1976-82), medium (1983-90) and very low inflation (1991-99). In the high-inflation period, the wage-change distributions contain no spikes at zero, have left and right-hand tails which contain about the same mass and tend to be symmetric. Moreover, any evidence of asymmetry that we find in that period is not consistent with wage rigidity. In the low-inflation period the picture is decidedly different: To begin with, very substantial spikes at zero are in evidence, the most pronounced of which, in 1993, has mass in excess of 50%. In addition, the left-hand tail, which includes zero, is significantly heavier than the right-hand tail that serves as the no-rigidity counterfactual. Finally, symmetry is rejected very convincingly in this period. During the medium inflation period, when the inflation rate is halved, spikes at zero begin to appear, mass begins to concentrate in the left-hand tail and the distributions for some years are asymmetric. Looking at all these results in the context of the simple regressions, in Table 3, the Lebow, Stockton and Wascher (1995) measure, which compares the left and right-hand tails, and the symmetry test statistic are systematically related to the rate of inflation.

Our results shed light on the inflation as 'grease' hypothesis. While the *incidence* of real wage reductions was very high during the high-inflation period, it remained generally high throughout the period under study. In the low-inflation period, real wage reductions were frequent because of the substantial concentration of mass at exactly zero, just lower than the average rate of CPI inflation. Nevertheless, the *magnitute* of real wage reductions during the low-inflation period was about one tenth of the maximal adjustments possible during the high-inflation period. Thus, the degree of real wage re-allignment during the low-inflation period was modest.

It should be noted that the results above are independent of the indexation provisions in these collective bargaining agreements. Thus, the results, in Tables 2 and 3, are very similar whether cost-of-living adjustments are included in the definition of the wage-change variable or not.

When the results are broken down by sector (public and private) and length of contract (one year or less and ever one year), tail behaviour is not affected, though the symmetry test, which is then calculated using smaller samples, is more erratic and fails to show the clear relation to the inflation rate that is apparent in the whole sample.

References

[1] Abraham K. G. and J. C. Haltiwanger (1995) 'Real Wages and the Business Cycle', *Journal of Economic Literature*, XXXIII, September, 1215-1264.

[2] Agell J. and P. Lundborg (1995) 'Theories of Pay and Unemployment: Survey Evidence from Swedish Manufacturing Firms', *Scandinavian Journal of Economics,* 97, 295-307.

[3] Ahmad I. A. and Q. Li (1997) 'Testing the Symmetry of an Unknown Density Function by the Kernel Method', *Journal of Nonparametric Statistics*, 7, 279-293.

[4] Akerlof G. A., W. T. Dickens and G. L. Perry (1996) 'The Macroeconomics of Low Inflation', *Brookings Papers in Economic Activity*, 1-59.

[5] Altonji J. and P. Devereux (1999) 'The Extent and Consequences of Downward Nominal Rigidity', Working Paper 7236, National Bureau of Economic Research.

[6] Beissinger T. and C. Knoppik (2000) 'Downward Nominal Rigidity in West-German Earnings 1975-1995', University of Regensberg, Discussion Paper 344.

[7] Bewley T. F.(1999) *Why Wages Don't Fall During A Recession* (Cambridge, Harvard University Press).

[8] Blinder A. S. and D. H. Choi (1990) 'A Shred of Evidence on Theories of Wage Rigidity', *Quarterly Journal of Economics,* 105, 4, 1003-1015.

[9] Campbell C. M. and K. S. Kamlani (1997) 'The Reasons for Wage Rigidity: Evidence from a Survey of Firms', *Quarterly Journal of Economics,* 112, 759-789.

[10] Card D. (1983) 'Cost of Living Escalators in Major Union Contracts', *Industrial and Labour Relations Review,* 34-48.

[11] Card D. (1986) 'An Empirical Model of Wage Indexation Provisions in Union Contracts', *Journal of Political Economy*, S144-173.

[12] Card D. and D. Hyslop (1997) 'Does Inflation Grease the Wheels of the Labor Market?', In *Reducing Inflation: Motivation and Strategy*, C. Romer and D. Romer, eds., National Bureau of Economic Research Studies in Business Cycles, vol. 30, (Chicago, University of Chicago Press), 114-121.

[13] Christofides L.N. (1987) 'Wage Adjustment in Contracts Containing Cost-of-Living Allowance Clauses', *The Review of Economics and Statistics*, August, 531-536.

[14] Christofides L.N. (1990) 'The Interaction Between Indexation, Contract Duration and Non-Contingent Wage Adjustment', *Economica,* 395-409.

[15] Cousineau J., R. Lacroix and D. Bilodeau (1983) 'The Determinants of Escalator Clauses in Collective Agreements', *The Review of Economics and Statistics*, 196-202.

[16] Crawford A. and A. Harrison (1998) 'Testing for Downward Rigidity in Nominal Wages', in *Price Stability, Inflation Targets and Monetary Policy,* 179-225, (Ottawa, Bank of Canada).

[17] Dunlop J. T. (1938) 'The Measurement of Real and Money Wages' *The Economic Journal,* September, 413-434.

[18] Ehrenberg R. G., L. Danziger and G. San (1984) 'Cost-of-Living-Adjustment Clauses in Union Contracts', *Research in Labor Economics,* 1-63.

[19] Fares J. and S. Hogan (2000) 'The Employment Costs of Downward Nominal-Wage Rigidity', Bank of Canada Working Paper 2000-1.

[20] Fares J. and T. Lemieux (2000) 'Downward Nominal Rigidity: A Critical Assessment and Some New Evidence for Canada', Mimeograph.

[21] Fehr E. and L. Gotte (2000) 'Robustness and Real Consequences of Nominal Wage Rigidity', CESifo Working Paper Series, No. 335.

[22] Fortin P. (1996) 'The Great Canadian Slump', *Canadian Journal of Economics*, November, 761-787.

[23] Freedman C. and T. Macklem (1998) 'A Comment on "The Great Canadian Slump"', *Canadian Journal of Economics,* Vol. 31, No. 3, 646-665.

[24] Groshen E. L. and M. E. Schweitzer (1999) 'Identifying Inflation´s Grease and Sand Effects in the Labor Market', in M. Feldstein, ed., *The Costs and Benefits of Price Stability, (Chicago, The University of Chicago Press),* 273-313.

[25] Hendricks W. E. and L. E. Kahn (1985) *Wage Determination in the United States: Cola and Uncola*, (Cambridge, MA, Ballinger Press).

[26] Kahn S. (1997) 'Evidence of Nominal Wage Stickiness from Microdata', *American Economic Review*, December, 993-1008.

[27] Kaufman R. T. and G. Woglom (1984) 'The Effects of Expectations on Union Wages', *American Economic Review,* 3, 418-32.

[28] Lebow D. E., D. J. Stocton and W. L. Wascher (1995) 'Inflation, Nominal Wage Rigidity, and the Efficiency of Labor Markets', Working Paper 95-45, Federal Reserve Board, Washington, D.C.

[29] Lebow D. E., R. E. Saks and B. A. Wilson (1999) 'Downward Nominal Rigidity: Evidence from the Employment Cost Index', Board of Governors of the Federal Reserve System, mimeograph.

[30] McLaughlin K. J. (1994) 'Rigid Wages?', *Journal of Monetary Economics,* 383-414.

[31] McLaughlin K. J. (1999a) 'Are Nominal Wages Skewed Away From Wage Cuts?', *Review,* Federal Reserve Bank of St. Louis May/June, Vol. 81, No. 3, 117-132.

[32] McLaughlin K. J. (1999b) 'Testing for Asymmetry in the Distribution of Wage Changes', Manuscript.

[33] Mitchell D. J. B. (1980) *Unions, Wages and Inflation* (Washington, D. C.: The Brookings Institution).

[34] Mood A. M., F. A. Graybill and D. Boes (1974) *Introduction to the Theory of Statistics* (New York, McGraw-Hill).

[35] Oswald A. (1993) 'Efficient Contracts are on the Labour Demand Curve: Theory and Facts', *Labour Economics,* 1, 85-114.

[36] Samuelson P. A. and R. M. Solow (1960) 'Problems of Achieving and Maintaining a Stable Price Level: Analytical Aspects of Anti-Inflation Policy', *American Economic Review, Papers and Proceedings*, 2, 177-194.

[37] Shultze C. L. (1959) 'Recent Inflation in the United States', Study Paper 1, Joint Economic Committee, 86 Congress, 1st Session, (September).

[38] Simpson W., N. Cameron and D. Hum (1998) 'Is Hypoinflation a Good Policy?', *Canadian Public Policy-Analyse de Politiques,* 291-308.

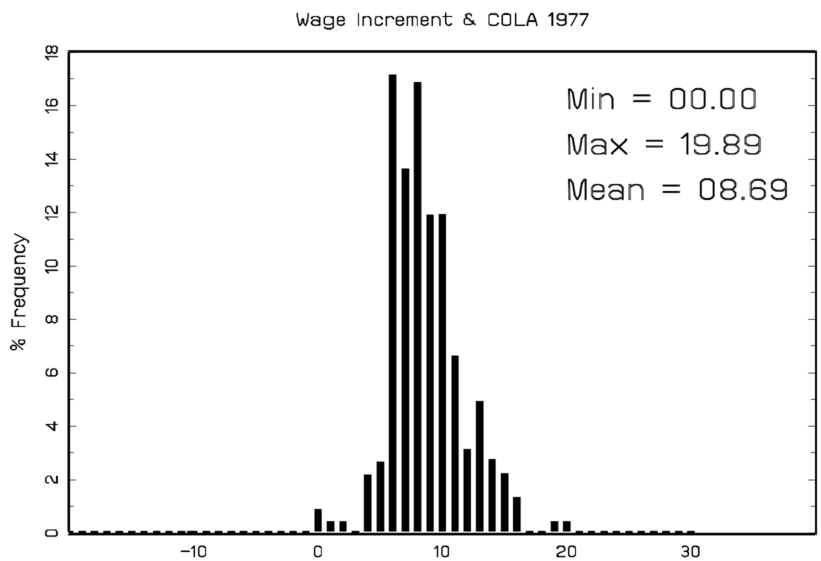
[39] Smith J. C. (1999) 'Nominal Wage Rigidity in the UK', University of Warwick, Mimeograph.

[40] Solon G., R. Barsky and J. Parker (1994) 'Measuring the Cyclicality of Real Wages: How Important is Composition Bias', *Quarterly Journal of Economics*, February, 1-25.

[41] Tarshis L. (1939) 'Changes in Real and Money Wages', *Economic Journal,* March, 150-154.

[42] Tobin J. (1972) 'Inflation and Unemployment', *American Economic Review,* 1, 1-18.

[43] Vroman W. (1984) 'Wage Contract Settlements in US Manufacturing', *The Review of Economics and Statistics,* November, 661-665.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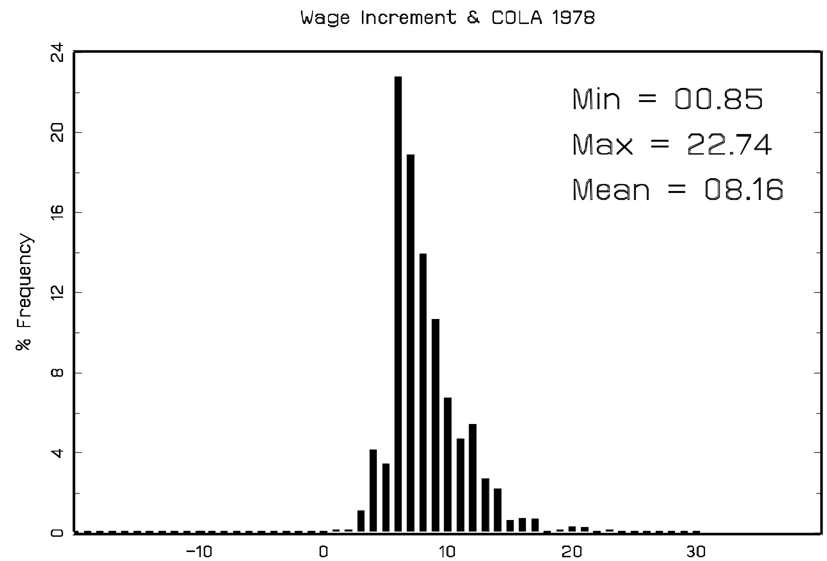
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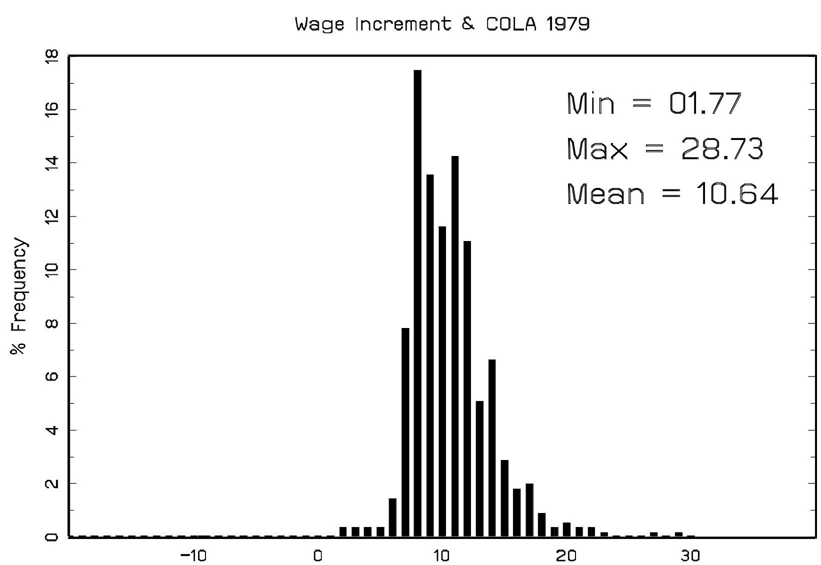
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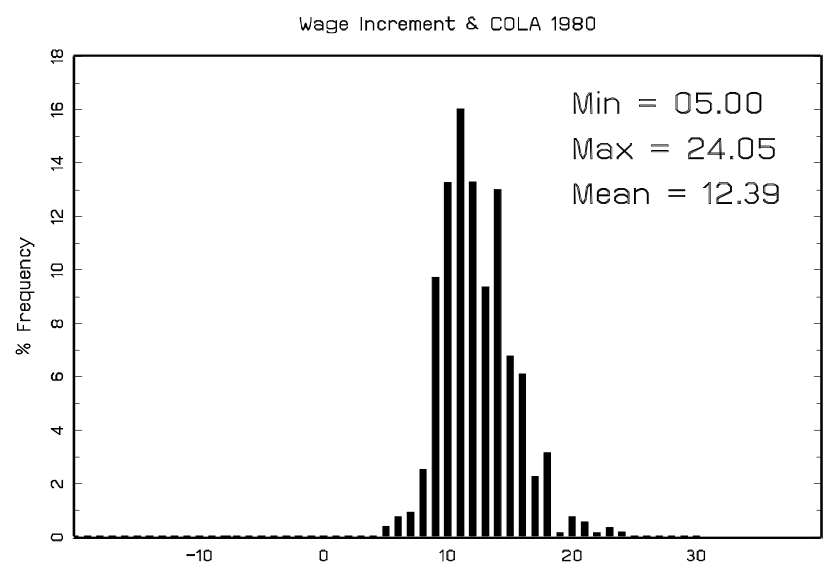
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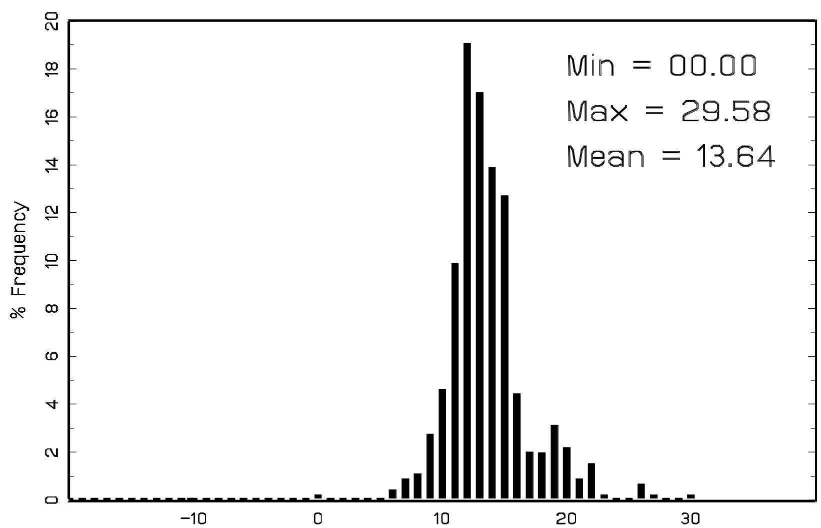

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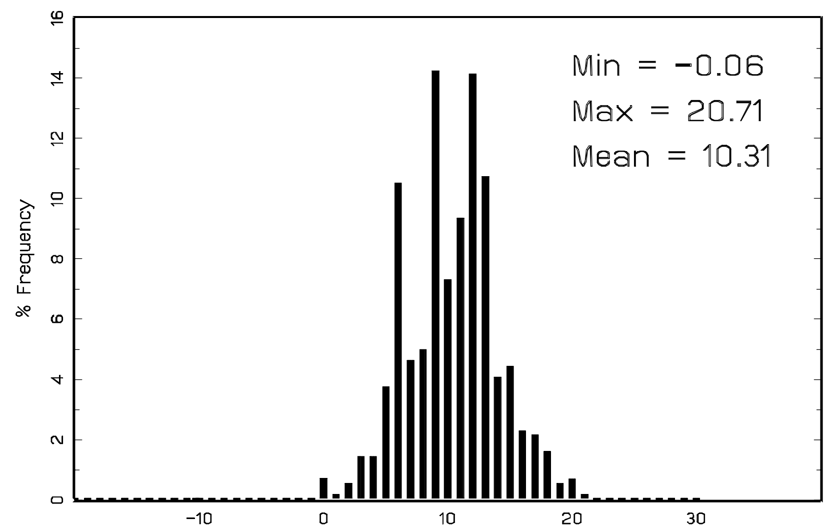

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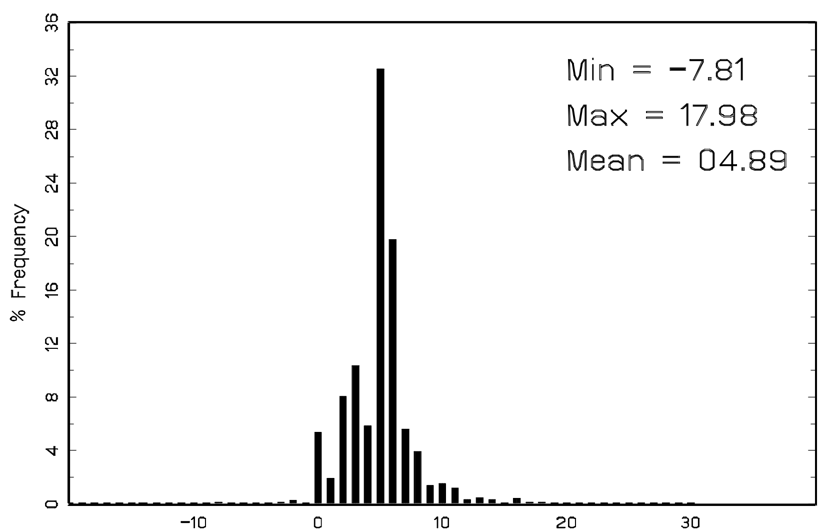

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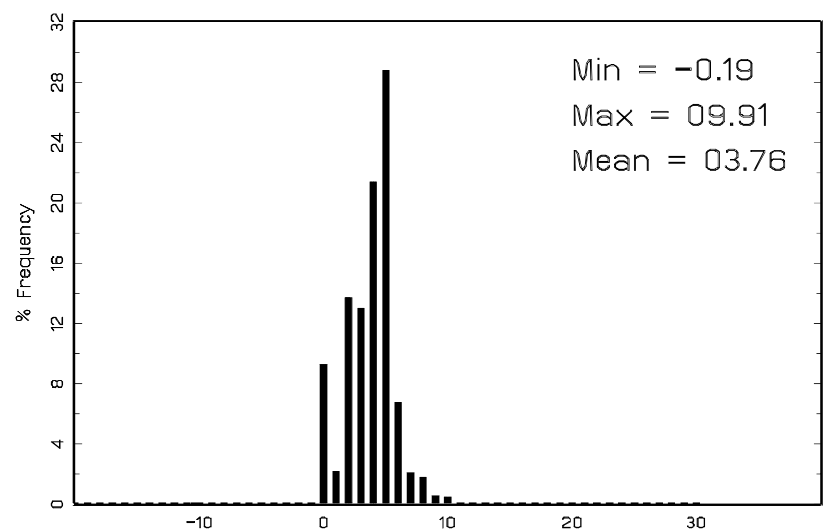

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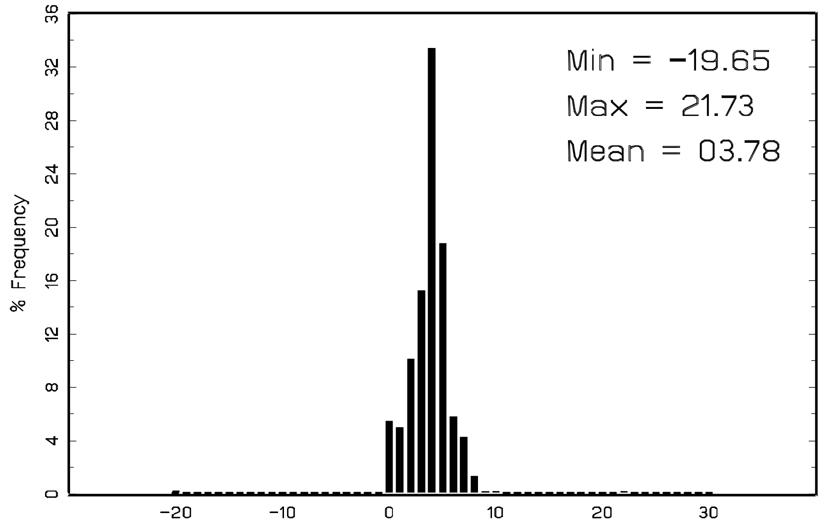

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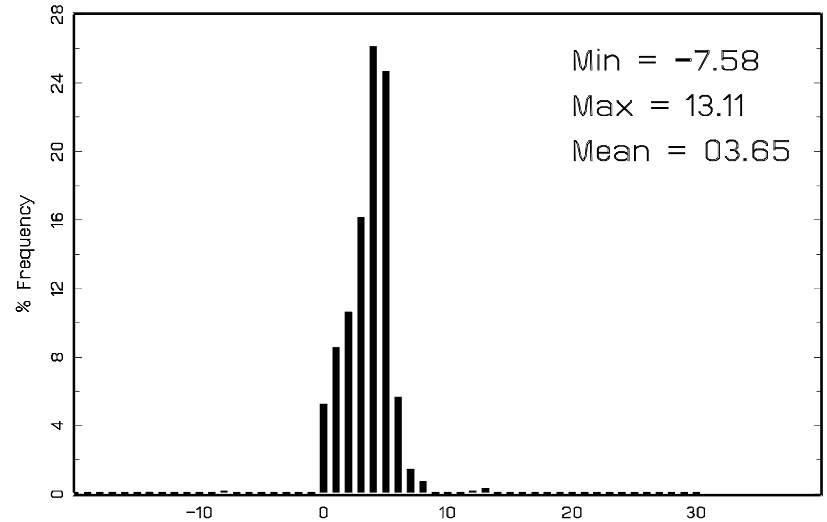

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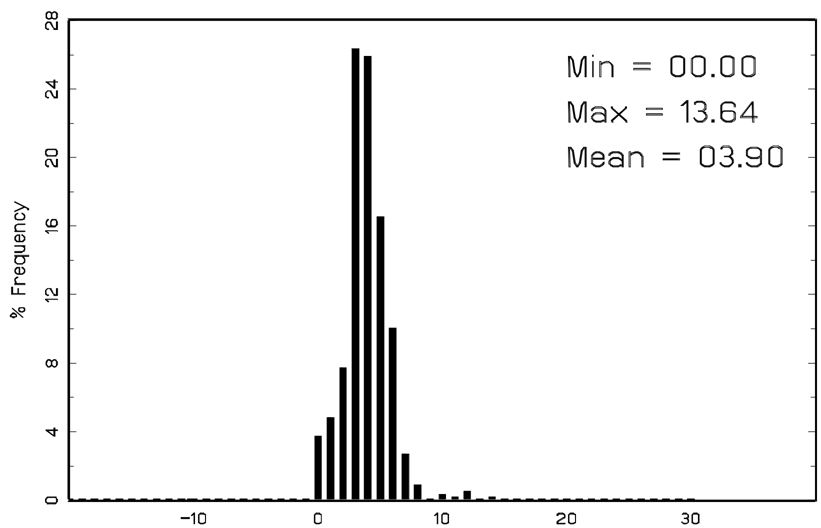

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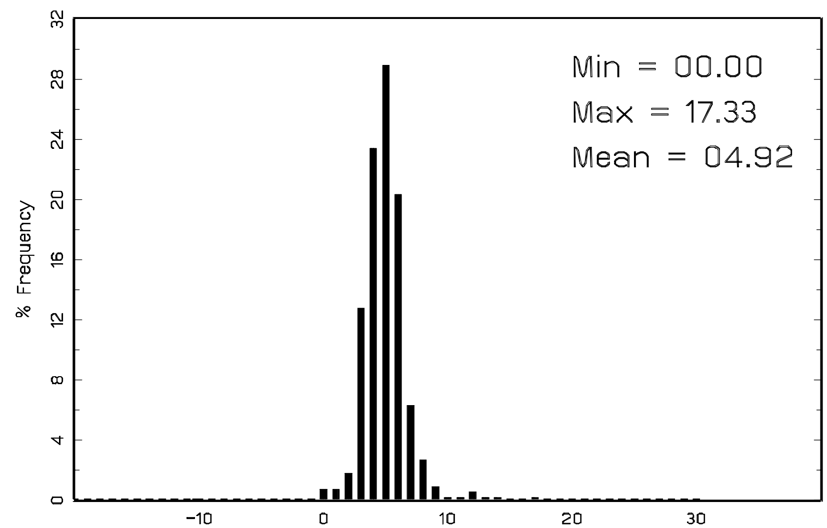

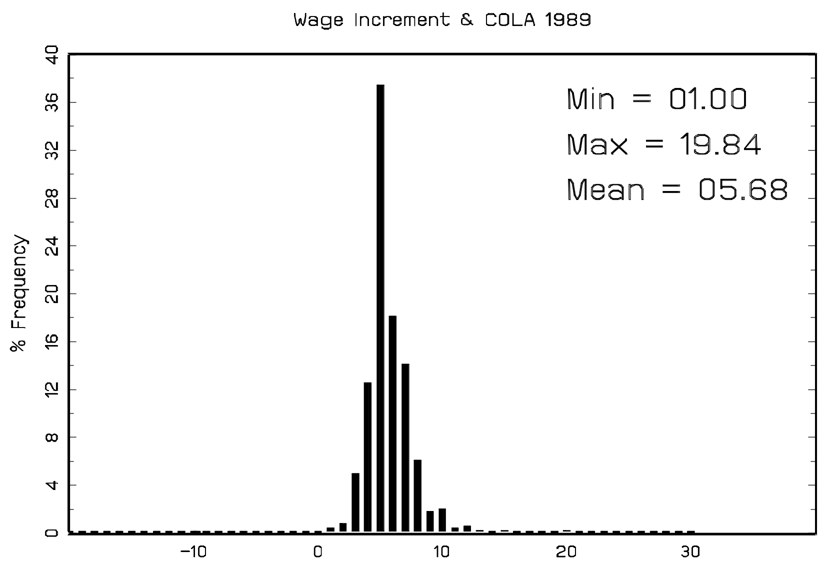
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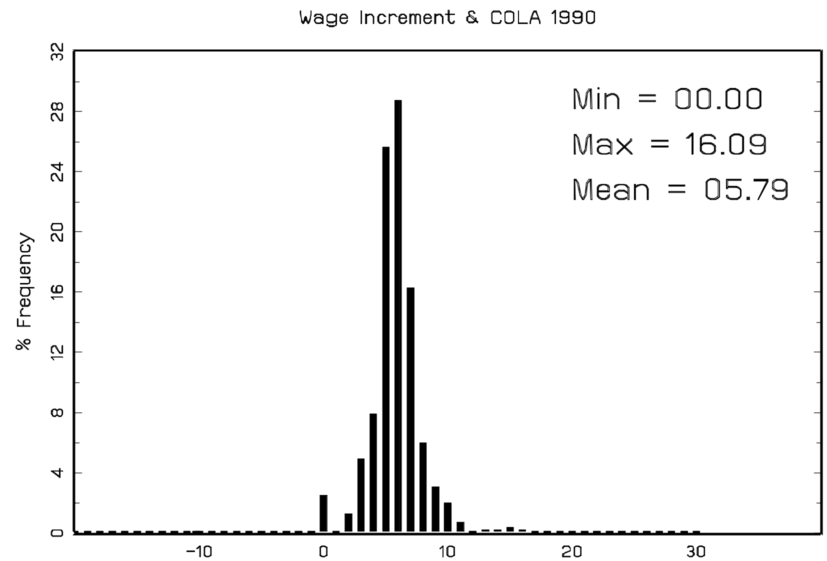
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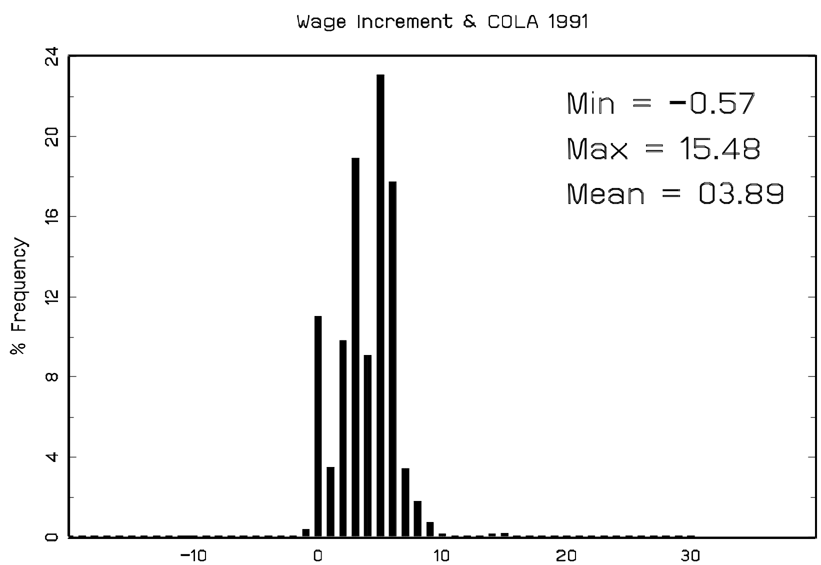
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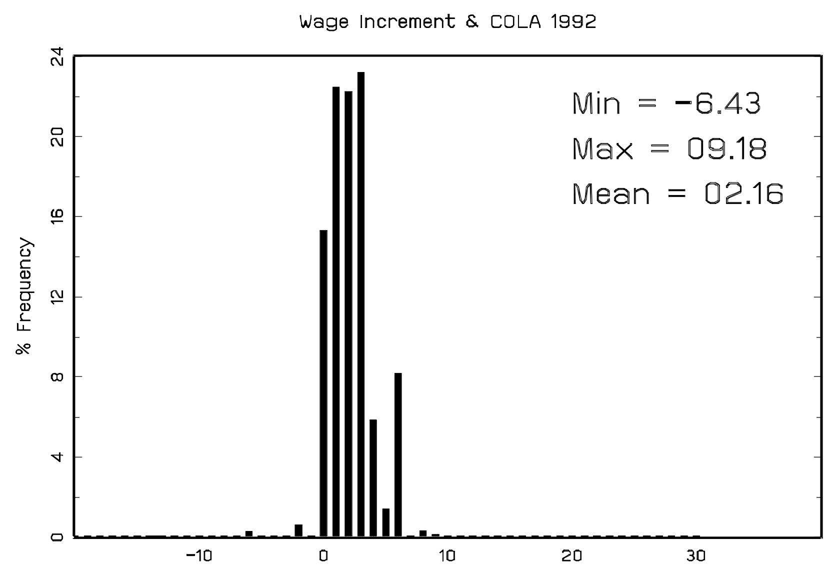
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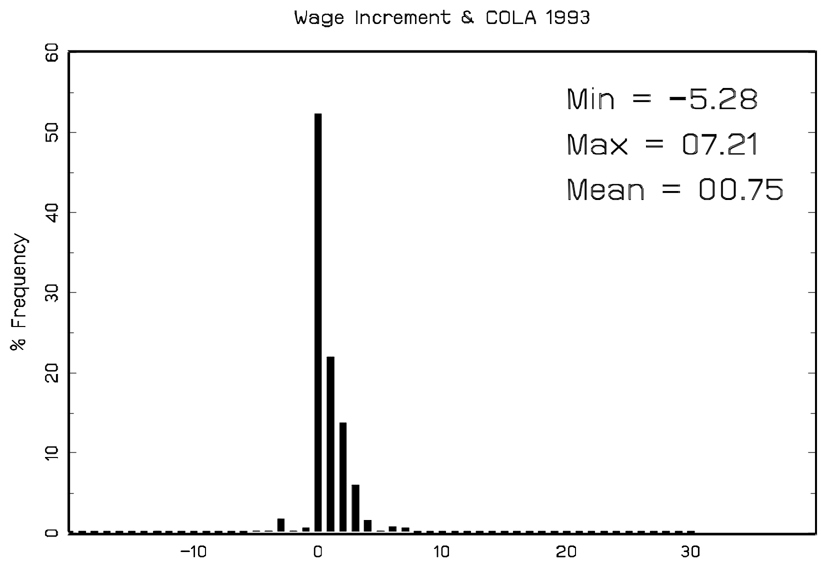
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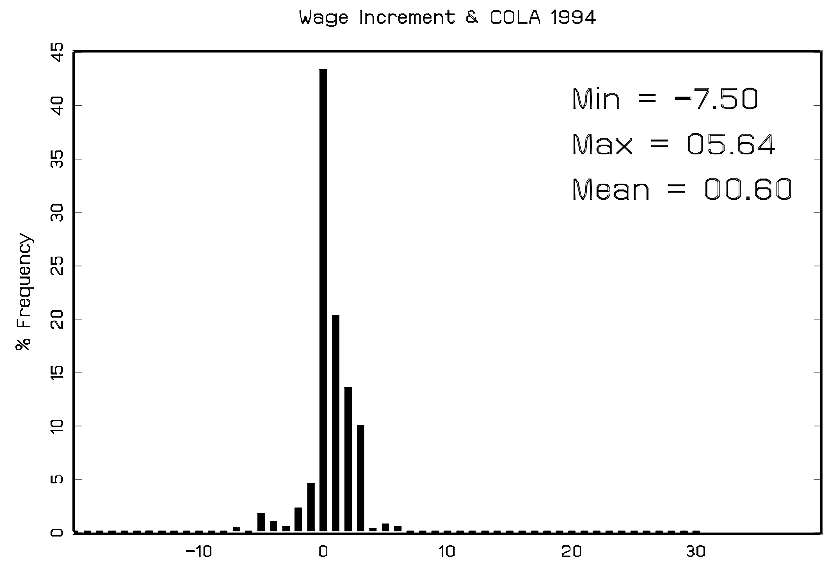
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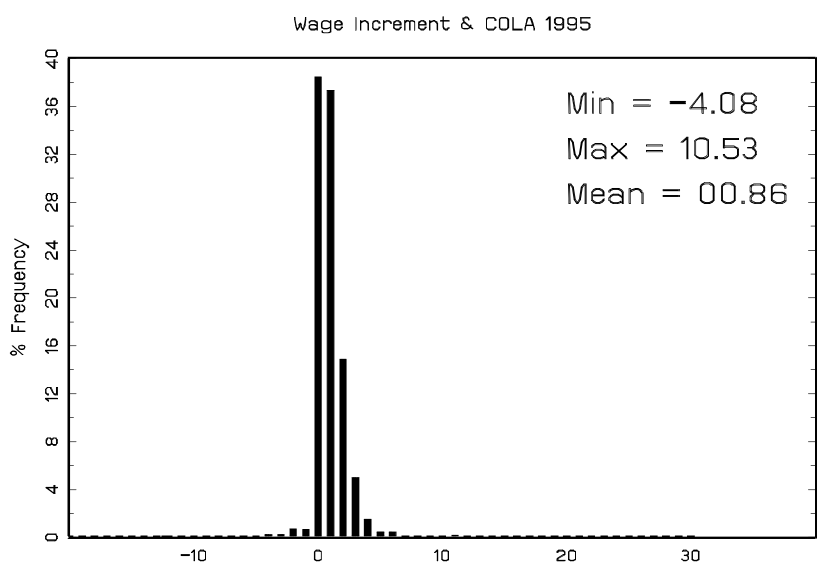
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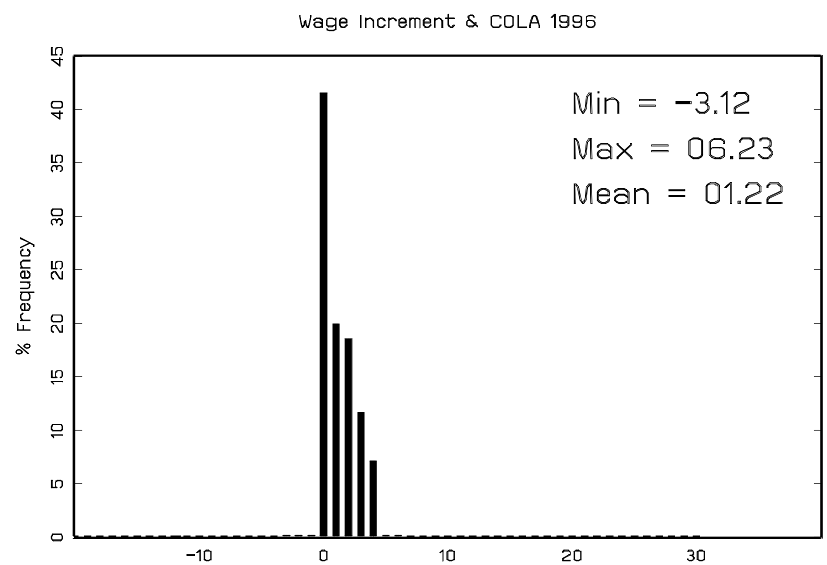
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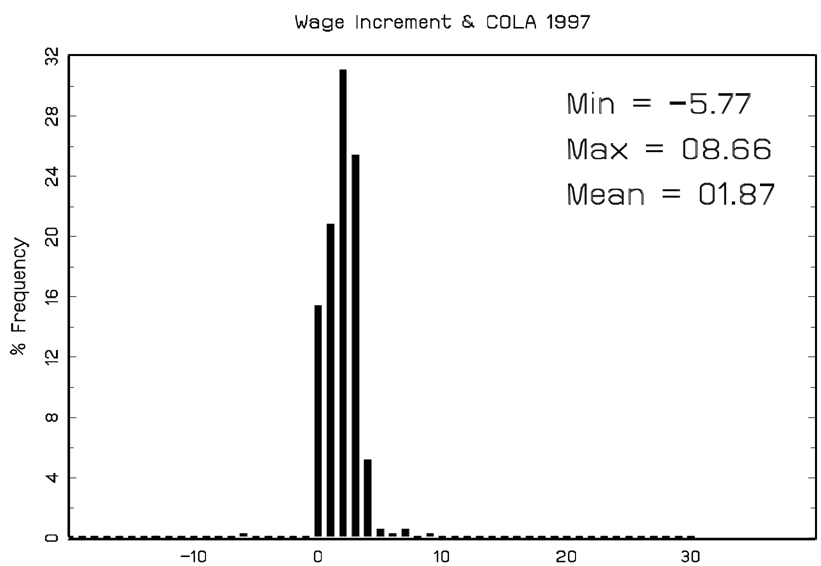
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Table 1

Number of Contracts by Year, Sector and Contract Length

	Total	Private	Public	Long	Short
1977[1]	226	179	47	125	101
1978	673	490	183	373	300
1979	569	378	191	415	154
1980	520	382	138	407	113
1981	450	306	144	309	141
1982	562	383	179	282	280
1983	643	407	236	296	347
1984	676	468	208	425	251
1985	519	344	175	394	125
1986	551	339	212	449	102
1987	557	419	138	450	107
1988	556	379	177	484	72
1989	493	309	184	426	67
1990	547	377	170	462	85
1991	530	358	172	386	144
1992	632	357	275	450	182
1993	516	341	175	445	71
1994	471	306	165	399	72
1995	460	283	177	390	70
1996	448	293	155	382	66
1997[2]	346	271	75	292	54
Total	10945	7369	3576	8041	2904

Notes: [1]Includes 1996 contracts.

[2]Includes 1998 and 1999 contracts.

Table 2
Descriptive and Test[1] Statistics

	$\dot{P}$	Wage Change Includes COLA						Wage Change Excludes COLA					
		MRW	SDRW	Median	Mean	Skew	$\widehat{D_n}$	MRW	SDRW	Median	Mean	Skew	$\widehat{D_n}$
1977[2]	9.58	-0.90	2.75	8.20	8.69	0.57	-0.09	-3.10	2.60	6.70	6.48	0.21	-0.80
1978	9.93	-1.81	2.70	7.42	8.16	1.30	1.02	-2.82	2.75	6.78	7.12	1.11	0.62
1979	11.53	-0.91	3.01	10.09	10.64	1.26	0.45	-3.11	3.74	8.63	8.41	0.21	0.36
1980	12.23	0.17	3.03	11.94	12.39	0.71	0.13	-1.04	3.35	11.03	11.15	-0.06	0.30
1981	9.51	4.14	3.38	13.09	13.64	1.03	0.12	3.27	3.98	12.87	12.76	-0.16	-0.09
1982	5.93	4.39	3.01	10.63	10.31	-0.01	-0.44	3.93	3.47	10.02	9.85	-0.25	-1.35
1983	4.46	0.44	2.67	5.00	4.89	0.60	-1.67	0.02	2.81	5.00	4.47	0.16	-4.83
1984	4.16	-0.40	1.87	4.00	3.76	-0.17	-4.44	-0.69	2.03	3.82	3.45	0.02	-6.03
1985	4.35	-0.56	2.17	4.04	3.78	-1.44	-2.34	-0.90	2.26	3.79	3.44	-1.12	-3.43
1986	4.41	-0.76	1.84	4.09	3.65	-0.07	-2.48	-0.97	1.90	3.76	3.44	0.08	-2.88
1987	4.65	-0.75	1.76	3.82	3.90	0.83	-1.29	-1.09	1.92	3.40	3.56	0.83	-2.33
1988	5.16	-0.24	1.78	4.89	4.92	1.44	0.03	-0.56	1.99	4.68	4.61	1.00	-0.79
1989	5.01	0.68	1.87	5.22	5.68	1.84	0.54	0.42	1.95	5.12	5.41	1.39	0.53
1990	3.90	1.88	2.16	5.77	5.79	0.47	-0.90	1.53	2.34	5.65	5.43	0.29	-1.58
1991	1.50	2.39	2.19	4.82	3.89	0.15	-5.20	2.20	2.20	3.90	3.69	0.28	-5.76
1992	1.50	0.66	1.68	2.00	2.16	0.40	-5.07	0.61	1.70	1.97	2.11	0.37	-5.69
1993	1.14	-0.39	1.41	0.00	0.75	1.00	n/a	-0.49	1.30	0.00	0.65	0.90	n/a
1994	1.80	-1.20	1.75	0.00	0.60	-0.79	n/a	-1.29	1.71	0.00	0.51	-0.77	n/a
1995	1.56	-0.69	1.20	0.68	0.86	1.98	-11.72	-0.74	1.15	0.68	0.82	1.99	-12.91
1996	1.62	-0.42	1.32	0.86	1.22	0.66	-10.20	-0.51	1.27	0.76	1.14	0.78	-10.55
1997[2]	1.72	0.14	1.33	1.87	1.87	0.16	-4.94	0.03	1.29	1.71	1.76	0.07	-4.96

Notes: [1] $\widehat{D_n} = \{1 - F_n(2 \cdot Median)\} - F(0)$,where F is the empirical distribution function. Only values for $\alpha = 8$ are reported, those for $\alpha = 10$ are very similar.

[2] Contracts for 1976 and 1977 have been merged into '1977' and those for 1997 to 1999 into '1997'.

Table 3

Regression Results: Test-Statistics on $\dot{CPI}$[1]

	$\widehat{D_n}$				$\widehat{I_n}$			
	$\alpha = 8$		$\alpha = 10$		$\alpha = 8$		$\alpha = 10$	
	coeff	t-stat	coeff	t-stat	coeff	t-stat	coeff	t-stat
$\dot{W}$ Includes COLA								
All Contracts	0.72	3.53	0.67	3.57	-0.70	-1.97	-0.68	-2.00
Private Sector	0.62	3.76	0.59	3.90	0.17	1.06	0.15	1.36
Public Sector	0.46	2.39	0.80	3.93	-0.31	-1.37	-0.31	-1.58
Long Contracts	0.40	2.49	0.36	2.44	-0.40	-1.36	-0.37	-1.39
Short Contracts	0.28	1.54	0.19	1.03	-0.22	-0.91	-0.20	-1.03
$\dot{W}$ Excludes COLA								
All Contracts	0.80	3.87	0.75	3.93	-1.65	-2.52	-1.51	-2.32
Private Sector	0.64	4.23	0.59	4.23	-0.04	-0.29	0.02	0.25
Public Sector	0.56	2.89	0.55	2.84	-0.25	-1.02	-0.22	-0.98
Long Contracts	0.40	2.39	0.36	2.34	-0.84	-1.34	-0.82	-1.32
Short Contracts	0.29	1.58	0.29	1.60	-0.22	-0.88	-0.19	-0.95

Note: [1] Based on results, such as those in Table 1, from the twenty-one years 1977-97. The regression equation involves $\widehat{I_n}$ or $\widehat{D_n}$ as the regressand and a constant plus $C\dot{P}I$ as regressors. When the median equals zero, $\widehat{D_n}$ is not appropriate and a dummy variable which equals 1 in 1993-94 (All), 1993-94 and 1996 (Public), 1993 (Long) and 1993-97 (Short) and is zero otherwise, is also included.