Nominal Wage Rigidity in Contract Data: A Parametric Approach

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July 11, 2002

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Abstract

Using wage agreements reached in the Canadian unionized sector during 1976-99, a period of high as well as exceptionally low inflation, we consider how histograms of wage adjustment change as inflation reaches the low levels of the 1990s. The histograms and parametric tests suggest that wage adjustment is characterized by downward nominal rigidity and significant spikes at zero. There is some evidence of modest menu-cost effects. We examine whether the rigidity features of wage adjustment are sensitive to indexation provisions and investigate whether the distinction between short and long contracts is useful.

JEL Classification: E31

Keywords: Low wage inflation, nominal rigidity, menu-costs

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1 Introduction

In Keynesian economics, downward rigidity in the nominal wage rate is typically associated with movements along the labour demand curve and a countercyclical real wage rate. A number of papers, from Dunlop (1938) and Tarshis (1939) to Solon, Barsky and Parker (1994) and Abraham and Haltiwanger (1995), consider rigidity indirectly by focusing on the cyclicality of the real wage rate.

Recent concern with 'menu costs' has broadened the context within which rigidities are generated[1] and the availability of data at the micro level has led to more direct examinations of the extent and nature of nominal wage rigidity. McLaughlin (1994) examines the wage-change experience of household heads who report a wage or salary for consecutive years. Using data from the Panel Study of Income Dynamics (PSID) for 1976-1986, he reports that real and nominal wage cuts are reasonably common and that, although there is some evidence of nominal rigidity in the form of a spike at zero in the log-wage change distribution, the degree of rigidity is limited. Lebow, Stockton and Wascher (1995) use the PSID data over the period 1970-1988 and also report limited evidence for downward nominal wage rigidity.[2] Card and

[1] See, for instance, Akerloff and Yellen (1985), Mankiw (1985), Caplin and Spulber (1987), and Ball and Romer (1990). Akerlof, Dickens and Perry (1996) review US evidence of downward wage rigidity and, persuaded by its pervasiveness, construct a model which predicts that attempts to lower the inflation rate by restraining aggregate demand entail permanently, rather than temporarily, higher unemployment and a long-run trade-off.

[2] Possible asymmetries in the nominal wage-change distribution and the extent to which this asymmetry itself varies with the inflation rate may shed some light on the degree of nominal wage rigidity. The measure of asymmetry used by these authors is negatively

Hyslop (1997) examine wage change in individual data drawn from the 1979-1993 Current Population Survey and the 1976-1979 and 1985-1988 PSID. Their real wage-change histograms show considerable concentration of mass at minus the rate of inflation and some visual evidence of asymmetry. They conclude that a 1% increase in the inflation rate reduces the fraction of workers affected by downward nominal rigidity by 0.5%. Kahn (1997) studies the wage change experienced by household heads who were in the same job for at least two contiguous years, during the period 1970-1988, using the PSID. Histograms for each year in this period are constructed and regression analysis based on this data shows substantial spikes around zero for a number of years, considerable evidence of menu costs and of downward wage rigidity particularly among wage, as distinct from salary, earners. Smith (2000), using British Household Panel Study data from the period 1991-96, concludes that wage rigidity is limited.

Results based on survey evidence have been greeted with some scepticism because of the possibility of data errors. Many of the studies cited above deal extensively with this issue and, indeed, Smith (2002) measures the extent to which accounting for various kinds of errors softens initial evidence of nominal rigidities. By contrast, wage data drawn from collective bargaining agreements appear to be accurate because they are generated from the contracts themselves by the relevant government agencies. The fact that they are drawn from the unionized sector is both a disadvantage and an advan-

related to the inflation rate only for employees who do not change jobs and are paid an hourly wage. For this group, the spike at zero is about 8% of the sample but only half of that is attributed to downward nominal wage rigidity.

tage. It is a disadvantage because union coverage is not complete. It is an advantage because it offers evidence from a set of labour market transactions where the rigidity forces at work may be rather different. For instance, it is standard practice for contract negotiations to begin toward the end of an existing agreement. Given this, one argument for nominal rigidity, that is the cost of determining whether an adjustment should occur at all, collapses and we are left with the pure accounting cost of revising wages as the only source of menu-cost rigidity. Thus, we might expect more mass immediately above and below zero and less mass at zero in wage-change distributions based on contract data than is the case in situations where both kinds of menu-costs[3] may apply. Another possibility is that union leadership might wish to demonstrate success by securing, under adverse bargaining conditions, even modest nominal wage increases. This argument, if valid, suggests that wage-change distributions in the unionised sector might display more mass immediately above zero than would appear in the non-unionised sector. For all these reasons, menu-cost behaviour may be more muted in union contracts than elsewhere. It is often argued that unions must demonstrate their prowess to the membership by successfully resisting nominal wage cuts. If so, thinning below zero and spikes at zero might be more pronounced than is the case in the non-union sector. We refer to these latter phenomena as downward nominal wage rigidity (DNWR). Finally, the nature of wage-change distributions at and around zero may be very different in the case of one-year and

[3]We refer to menu-cost behaviour as that which results in unusually low mass immediately above and below zero, combined with a corresponding concentration of density at zero.

multi-year union contracts: It is possible that menu-cost behaviour will be more evident in the case of short contracts because the desired nominal wage change will be smaller overall and not worth implementing if each nominal adjustment is costly. Moreover, DNWR may be more prevalent in short contracts because unions may be more willing to settle for a wage freeze over a short period of time.

Despite the obvious wealth of information that might be gleaned from wage-change distributions based on collective bargaining agreements, more econometric work is needed to extract it. What *has* been done so far is based on data from Canada, where union coverage is considerably higher than that in the US. Fortin (1996) reviews possible reasons for the high, relative to the US, level of the Canadian unemployment rate and pays particular attention to the possible role of wage rigidity. A histogram of nominal wage change in Canadian union contracts which contain no Cost-of-Living-Allowance (COLA) clauses and were signed between 1992-1994 indicates a large spike at zero and virtually no mass below that point. Based on visual evidence from this histogram, he concludes that considerable rigidity is evident. These findings are criticized by Freedman and Macklem (1998) on the grounds that he uses information only from the first year of contracts, regardless of contract length, and that most rigidity emanates from the public sector. Crawford and Harrison (1998) examine the evidence from Canadian union contracts in greater detail. They present histograms of nominal wage change in private and public sector union contracts and 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. Christofides and Stengos (1998) use non-parametric techniques to examine specifically the symmetry of nominal wage-change distributions, drawn from both survey and contract data. They report more asymmetry at times of low inflation. Simpson, Cameron and Hum (1998) pursue the themes in Fortin (1996) further. They consider the increase in the unemployment rate that would be needed to moderate wage inflation by the amount they attribute to nominal wage rigidity. They conclude that this increase in the unemployment rate could be as high as two percentage points, a conclusion that is questioned by Fares and Hogan (2000). Despite these efforts, there is no parametric econometric study, based on union contract data, which examines the nature of wage-change distributions with particular reference to behaviour at and in the neighbourhood of zero. How big are spikes at zero and how do these differ in years of low and high inflation? Is there evidence of thinning in the area of the distribution below zero relative to some benchmark? What evidence is there that the wage-change distribution contains spikes at zero in conjunction with holes around zero, behaviour which may be indicative of menu costs, and how important is this evidence quantitatively? These are all matters on which we hope to shed some light, although we do not attempt to select among competing explanations for nominal wage rigidity.

We use the latest release of the Canadian wage contract data which contains information from the period 1976 to 1999. This period is longer and contains eras of much lower inflation than those studied in North America to this point. This latter characteristic is particularly important in the context of the literature which deals with the nature of wage adjustment in periods of very low inflation. We examine the evidence for the whole sample but also deal with the importance of distinctions involving short and long contracts as well as contracts with and without indexation provisions. Because we wish to focus on the measurement of spikes at and around zero we adopt parametric techniques, such as those used in Kahn (1997), but also explicitly consider the influence of the wage inflation median on the nature of the wage-change distributions and the extent to which our results are sensitive to particular parameterisations. More details on the data and sources appear in section 2. The methodology used and the results obtained are described in section 3. Concluding observations appear in section 4.

2 Data and Sources

The contract data used in this paper are compiled by Human Resources Development Canada (HRDC is the federal ministry responsible for monitoring agreements between firms and unions), are based on agreements for which reporting requirements apply, and are, therefore, thought to be very accurate.[4]

The data base involves settlement dates as early as 1976 and as late as 1999, includes agreements which range in duration from a few months to several years, and covers bargaining units involving 200 to nearly 80,000

[4]Examination of the data revealed only two observations out of the 10,947 supplied to us by HRDC which did not satisfy basic consistency criteria. These observations have been excluded.

employees. The base wage rate, paid to entry-level workers, in the 10,945 available contracts is on average \$12.40 at the beginning and \$13.49 at the end of these contracts. The implied rate of change of 8.79% applies to contracts for which the mean duration is about two years and is roughly half this amount at annual rates. The average increase in the base wage rate of \$1.09 consists, subject to rounding, of a \$0.96 non-contingent increase (WNC) and a \$0.12 contingent increase which is the result of cost-of-living-allowance (COLA) clauses. Very few contracts contain COLA clauses.[5] In this paper, the variables WNC and $WNC + COLA$ are the non-contingent and total, respectively, wage change defined over the whole of the jth contract *at annual* rates. They appear in our sample as a single observation for each contract.[6] Table 1 contains, for each year,[7] the number of contracts, the corresponding average of the non-contingent wage adjustment (WNC) at annual rates, the average of the total wage increase ($WNC + COLA$) over the life of contracts at annual rates, and the annual rate of Consumer Price Index inflation. Wage adjustment and price inflation clearly move together, with total wage adjustment exceeding non-contingent adjustment noticeably in

[5] 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).

[6] An alternative approach involves defining sub-periods of the contract and establishing WNC and COLA over each of these. For a discussion of this issue, see Fortin (1996) and Freedman and Macklem (1998).

[7] 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.

the high-inflation years, when the yield on the COLA clauses is substantial. We proceed to analyse total ($WNC + COLA$) wage adjustment though we also report some findings pertinent to WNC only.[8] It should be noted that, because the incidence and intensity of COLA clauses is limited, the results below are similar to those which obtain when contingent increases are not included. To conserve space, we report only histograms for the high-inflation years of 1981 and 1982, the connecting medium inflation years of 1983, 1984, 1989 and 1990, and the low-inflation years of 1991 and 1992; histograms for other years in the sample are available on request.

Figure 1 presents histograms for $WNC + COLA$ for 1981-1984, while the companion Figure 2 presents histograms for 1989-1992. In constructing these, care was taken to standardize the 'bins' and clearly centre them on zero.[9] During the high inflation years of 1977-82, the histograms appear reasonably symmetric and they tend to display no noticeable spikes at zero or overwhelming evidence of menu-cost behaviour - see, for example, the histogram for 1981 in Figure 1. When inflation begins to abate after 1982 but before it increases again somewhat during 1988-91, the general appearance of the histograms changes substantially: During 1984-87 (see, for example, the histograms for 1983 and 1984, Figure 1), the histograms are characterized by

[8] Since our main aim is establishing the general patterns of nominal wage rigidity, total wage adjustment is preferable to WNC only. Analysing non-indexed contracts only would raise selection issues, for it is well-known that the incidence and intensity of indexation provisions are endogenous. A referee notes that the *ex post* COLA is, under rational expectations, an unbiased estimator of the expected COLA that agents would normally be concerned with.

[9] That is, the zero interval is -0.49 to 0.5. Further intervals increase and decrease in one percentage units. Note that almost all density in the zero bin is actually at zero itself.

considerable density at and immediately above zero, virtually no wage decreases and some indications of possible menu-cost behaviour - as it happens in the two illustrative years of 1983 and 1984 only. As average wage adjustment increases during 1988-90, the general appearance of the histograms changes noticeably: The histograms for these three years (see, for example, those for 1989 and 1990 in Figure 2) 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 1977-82, these histograms are similar in rough form to that for 1981, for example, and appear to have been substantially influenced by the easing of labour market conditions during this period - see Table 1. Beginning in 1992, wage and price inflation declines to levels which are unprecedented in recent decades and are much lower than those in the US. It is histograms like those for 1991 and 1992, Figure 2, albeit it for WNC only, that led Fortin (1996) to argue that extensive nominal wage rigidity was present.[10] Note the hole immediately above the zero bin in the histogram for 1991, a fact which is consistent with menu-cost behaviour. In summary, Figures 1 and 2 suggest that, as inflation moderates, wage adjustment becomes concentrated at and above zero with virtually no nominal wage decreases in evidence. Since wage decreases are negligible, menu cost behaviour to the left of zero is not an issue but some indication of menu-cost behaviour may be present, in selected years, above zero.

However, visual evidence such as that in Figures 1 and 2 does not amount

[10] Fortin (1996) notes that the Canadian recession in the 1990s was more severe than that in the US and that 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.

to statistical statements and work in this area is beginning to attempt this next step. How can the evidence in these figures be marshalled to bear in a statistical sense on the issue of nominal wage rigidities? Suppose that nominal wage adjustment in a contract is given by the rate of monetary expansion plus an idiosyncratic productivity shock. There is no reason to believe that the density function of the idiosyncratic shocks, which may be symmetric around the rate of monetary expansion during periods of high-inflation, should become asymmetric and produce asymmetric, censored, histograms, such as those in Figure 2, during the low-inflation periods. Thus, the lack of any noticeable density for negative values of $WNC+COLA$ after 1991, combined with substantial mass at zero, may be viewed as evidence of DNWR. One way of proceeding is to consider a broad historical period, such as that covered by the contract data, and use this information to estimate the shape of a representative histogram. Points of particular interest in the domain of the wage-change distribution (e.g. zero) can be examined in greater detail and be related to the average height of histograms for the period as a whole. In the next section, the information in Figures 1 and 2 is analyzed using parametric techniques, similar to those in Kahn (1997), which focus particularly on behaviour around zero and which take into account the median of the wage-change distribution itself. We wish to measure the size of spikes in histograms below, at and immediately above zero and thus to quantify the degree and nature of nominal rigidities. The robustness of our findings to alternative parameterisations and samples is also considered.

3 Results and Sensitivity Analysis

3.1 Main Results

The information analyzed in this section is contained in histograms such as those displayed in Figures 1 and 2. In our first model in equation (1) below, it is supposed that

$$H_{it} = \sum_{i=1}^{n} \alpha_i D_i + u_{it}, \ \forall i = 1, ..., n \ and \ \forall t = 1977, ..., 1997 \tag{1}$$

where H_{it} is the relative frequency in the histogram bar which is located i percentage points below (or above) the median of the distribution for year t, the α_i are parameters which are constant across time, D_i is a dummy variable which equals 1 when the relative frequency refers to a histogram bar which is i percentage points below the median, and u_{it} is an error term with classical properties. Implied parameterisations include the size of bins, which is assumed to equal one percentage point, and n, the number of bars required to adequately describe the representative histogram. The latter must reflect the empirical distributions, for a large value of n could involve D_i which are never positive when i is large and a small value of n could entail histogram bars which are never taken into account. The parameter n is initially set equal to eight, although other values are also considered. The parameter constancy across t is imposed by stacking equation (1) for the twenty-one years in the sample and estimating using OLS. In effect, the $\hat{\alpha}_i$ are the mean values of the histogram heights at the respective i, thus producing a representative histogram.

The literature on inflation as a lubricant places considerable stress on the symmetry of the wage-change distribution as inflation changes. This distribution is expected to be reasonably symmetric at high rates of inflation but, when inflation is low, DNWR may produce little mass below zero, considerable mass at zero, and declining densities for higher rates of wage adjustment. In Table 2, we begin by investigating the wage-change ($WNC + COLA$) distribution over the entire sample period. Once stacked, there are 168 observations (21 years times 8 histogram bars in each year) on each side of the median. Columns 1 and 2, Table 2, report the results for the combined sample, pooling observations above and below the median, while columns 3 to 6 report the results for the sub-samples below and above the median. The choice of $n = 8$ appears reasonable[11], the fit of the equation is satisfactory, and the representative histogram falls off to zero as n increases to 8.[12] A test for structural homogeneity above and below the median, which tests the symmetry of the wage-change distribution without imposing any distributional restrictions, does not reject symmetry,[13] suggesting that ef-

[11] We have experimented with other values of n, the number of percentage-point bars included in the data. The lower value of $n = 5$ failed to capture needed density when the median rate of wage inflation was high. A value of $n = 13$ produced insignificant coefficient estimates for $i > 8$ but our results below were not very different. Clearly, n should be large enough to capture needed density but not so large that it weakens the statistical integrity of the model through the estimation of superfluous parameters.

[12] The average value of the dependent variable for the sample as a whole is 0.942. That is, some 0.058 of the density is contained in the median bar which is excluded by construction. The sum of the coefficients (the predicted density) on each side of the median is 0.471 which, of course, amounts to the actual density. Subject to rounding, similar comments hold for the sub-samples.

[13] The calculated F value is 0.794 while the critical one is 1.94.

fects specific to the low-inflation years wash out over the 21-year period as a whole and that an explicit search for patterns peculiar to the low-inflation years needs to be conducted. To that end, interactions with the wage median are considered below. It should be noted that, while symmetry tests have been used as indicators of DNWR, these must be carried out in such a way that they take account of the possibility that no-rigidity wage-change distributions may themselves be asymmetric. A strength of the present approach is that it provides more powerful tests of DNWR and the presence of menu-costs.

Equation (1) is, therefore, augmented with four dummy variables whose coefficients modify the densities below, at, and immediately around zero:

$$H_{it} = \sum_{i=1}^{n} \alpha_i D_i + \beta_1 DN_{it} + \beta_2 DN1_{it} + \gamma D0_{it} + \beta_3 DP1_{it} + u_{it} \qquad (2)$$

where DN_{it} equals 1 if, for given i and t, the histogram bar involves a decrease in the base wage rate and is zero otherwise, $DN1_{it}$ equals 1 if the histogram bar is one percentage point below the zero bar and is equal to zero otherwise, $D0_{it}$ equals 1 if the histogram bar is actually located at zero and is equal to zero otherwise, and $DP1_{it}$ equals 1 if the histogram bar is one percentage point above the zero bar and is equal to zero otherwise. Thus, consistent with the notion of DNWR, the overall histogram can have a spike at zero ($\gamma > 0$) and uniformly lower densities to the left of zero ($\beta_1 < 0$), as well as menu-cost effects, i.e. unusually low densities, or holes, immediately below and above zero (β_2 and β_3 negative). These effects, if statistically significant, would introduce detail in the area below the median that would render the wage-change distribution asymmetric. A variant of this equation involves forcing the spike at zero to reflect the possible decrease (as measured by β_1) in each of the i densities below zero and possible holes (as measured by β_2 and β_3) in the densities just below and just above zero:

$$H_{it} = \sum_{i=1}^{n} \alpha_i D_i + \gamma D0_{it} + \beta_1 \left[DN_{it} - (8 - i)\, D0_{it} \right]$$

$$+ \beta_2 \left(DN1_{it} - D0_{it} \right) + \beta_3 \left(DP1_{it} - D0_{it} \right) + u_{it}. \qquad (3)$$

The within-equation parameter constraints are clear from equation (3). As in the case of equation (1), the data for the twenty one years are stacked in order to impose the cross-equation parameter constraints implied by equations (2) and (3). The results appear in Table 3, columns 1-4, under the labels Model E2 and Model E3, or 'E2' and 'E3', for short respectively. The base histogram for equation (2), that is the one that would hold when the wage-change distribution is well to the right of zero, is indicated by the estimated coefficients $\alpha_1 - \alpha_8$ which are analogous to those in column 3, Table 2. The area to the left of zero now has histogram bars which are lower by 0.035 and there is an additional height at zero of 0.038 - see Table 3, column 1. These effects are statistically significant at the 1% level in two-tailed tests. The additional height at zero appears small relative to the spikes evident in Figures 1 and 2, but it must be remembered that it applies to the sample as a whole, not just the low-inflation period. We return to this issue in the context of the more complex models below. The evidence for holes around zero is more mixed: There is evidence of such effects below ($\hat{\beta}_2 = -0.033$ with a t-statistic of -2.33) but not above ($\hat{\beta}_3 = -0.016$ with a t-statistic of -1.12) zero. This is surprising, given the histograms in Figures 1 and 2. We return to this point below.

The more tightly specified model E3, of equation (3), is shown in columns 3-4, Table 3. The interpretation of the β coefficients is straight-forward and the estimates similar to those in column 1, Table 2. The coefficient on $D0$ is given by

$$\gamma - ((8 - i)\beta_1 + \beta_2 + \beta_3) \tag{4}$$

and is plotted, in Figure 3, for both the E2 and E3 models as well as for the additional models (E2m and E3m) introduced below. Clearly, for model E2, γ is the constant 0.038. For model E3, Figure 3 shows how the size of expression (4) changes with different values of i, given the estimates of γ, β_1, β_2 and β_3. This coefficient, which is only relevant when the histogram includes zero, is equal to 0.169 when the median is equal to zero (as in 1993 and 1994), 0.138 when $i = 1$ implying that the median is within one percentage point of the bar containing zero, 0.107 when $i = 2$ and the median is within two percentage points of the bar containing zero, and so on. That is, the height of the additional mass at zero (due to $D0$) is larger when the median is close to zero, as indeed is suggested by the histograms in Figures 1 and 2. In this sense, the problem noted in model E2, namely that the additional density at zero of 0.038 is rather small, disappears.

The role of inflation generally and the median of the wage-change distribution in particular can be made more explicit. Figures 1 and 2 suggest that the size of spikes at zero and the amount of mass in the negative quadrant depend on the location of the histogram. Accordingly, we modify equations (2) and (3) by introducing interactions involving the median (M_t) of each t

distribution, and we refer to the resulting models as E2m (Equation (2) with median interactions) and E3m (Equation (3) with median interactions). For instance, E3m becomes:

$$H_{it} = \sum_{i=1}^{n} \alpha_i D_i + (\gamma + \gamma_m M_t) D0_{it} + (\beta_1 + \beta_{1m} M_t) \left[DN_{it} - (8 - i)\, D0_{it} \right]$$

$$+(\beta_2 + \beta_{2m} M_t) (DN1_{it} - D0_{it}) + (\beta_3 + \beta_{3m} M_t) (DP1_{it} - D0_{it}) + u_{it} \quad (5)$$

The equations are still stacked to maintain the cross-equation constraints.

Columns 5-8, Table 3 report the results obtained and an improvement in fit is evident.[14] In the E2m model in column 5, Table 3, all coefficients but β_3 and the interactions with the median for β_1 and β_3 are significant at the 1% level. The results suggest a statistically significant thinning of the mass in the negative quadrant ($\hat{\beta}_1 = -0.042$), significant additional mass at zero ($\hat{\gamma} = 0.145$) which decreases as the median increases, a significant hole ($\hat{\beta}_2 = -0.111$) in the bar immediately below zero, and a hole in the bar immediately above zero which is neither significant nor dependent on the median. The combination of all these effects produces a predicted histogram (below the median) which matches the salient facts in Figures 1 and

[14] Kahn (1997) considers a variant of equation (3), where the dummy variables operate on the α_i in multiplicative, rather than linear, fashion - her 'proportional' model. For salaried workers, the proportional model must be supplemented with time trends before reasonable predictions can be obtained. Our own equation (5) estimates more basic parameters; these additional parameters render time trends unnecessary and provide an economic interpretation for temporal changes in the coefficients (they are functions of the median wage change or, more broadly, the overall inflationary environment).

2 extremely well; we postpone discussion of this feature until our next model has been considered. Equation (5), which incorporates the median interactions as well as the within-equation constraints (the E3m model), appears in column 7, Table 3, and has a slightly better overall fit. This equation again suggests a thinning in the mass below zero which diminishes as the median increases. There is an additional hole in the bar immediately below zero which decreases in size as the median increases. The level and interaction effects are all statistically significant at the 5% level. The level and interaction involving γ are significantly different from zero at the 5% level. The coefficient and interaction involving $DP1$ are not significantly different from zero.

The implications of these results for nominal wage rigidity and behaviour in the neighbourhood of zero are explored in Figures 3 to 5. In Figure 3, the general pattern noted for model E3, namely that at low rates of inflation the coefficient on $D0$ is larger than at high rates of inflation, is also evident in models E2m and E3m. For E3m, the relationship is not linear because the value of all coefficients in expression (4) are functions of the median. In this, most complex specification, the spike at zero reaches 0.21 when the median is zero and declines to zero for values around the average median ($M = 5.14$) in the sample.

In Figures 4 and 5, we present the implications of these results for the shape of the predicted histogram below the median when $M = 5.14$ and for the low-inflation value of $M = 2$, respectively. Models E2m[15] and E3m

[15]To calculate the histogram bars for the E2m model we would proceed as follows. When $M = 5.14$, the first bin below the one containing the median is 3.5-4.5, the second is 2.5-3.5 and so on. At four percentage points below the median, $DP1 = 1$ and its

result in similar predictions so we proceed using model E3m only.[16] When inflation is reasonably high and $M = 5.14$, the predicted histogram in Figure 4 declines reasonably smoothly, with some indication of menu-cost behaviour above zero.[17] However, in the low inflation period, when, as an example $M = 2$, the bar at the $0.49 - 1.5$ bin is equal to 0.208 while that at the $-0.49 - 0.5$ bin is equal to 0.209. Between them, these two heights essentially exhaust the density below the median (see Table 2, where the area below the median is shown to be 0.449) and the six bars to the left of the zero bin

coefficient is $(0.006 - 0.005 \times 5.14) = -0.020$. At five points $D0 = 1$ with a coefficient of $(0.145 - 0.028 \times 5.14) = 0.001$, while at six points $D1N = 1$ with a coefficient of $(-0.111 + 0.024 \times 5.14) = 0.012$. By this stage, $DN = 1$ as well, with a coefficient of $(-0.042 + 0.004 \times 5.14) = -0.021$. To these effects must be added the base heights for the histograms appearing in column 5, Table 3, and the sum of all these would constitute the predicted histogram heights to the left of the median. The procedure would be repeated with $M = 2$, beginning in the bar immediately above zero where the coefficient on $DP1$ would apply right away.

[16] Again, we evaluate at $M = 5.14$, using procedures analogous to those described in the previous footnote. The overall coefficients for β_1, β_2, γ, and β_3, taking the median interactions into account, are respectively $-0.009, 0.002, -0.047$ and -0.016. The coefficient on $D0$ is given by the median-interacted equivalent to expression (4) and is plotted in Figure 3 for different values of the median or i. When $i = 5$, as implied by the value $M = 5.14$, the coefficient on $D0$ is equal to -0.006 . For these values, the predicted histogram to the left of the median is plotted in Figure 4. The procedure is repeated for $M = 2$, obtaining new values of β_1, β_2, γ, β_3 and $D0$ which are equal to $-0.028, -0.067, -0.151, -0.003$, and 0.087 respectively, and resulting in the histogram in Figure 5.

[17] That is, the estimated coefficients and interactions blend into a combined picture on menu costs which suggests only mild effects above zero. Recall that holes around zero were observed for only a few years in the histograms and then only above zero.

have heights which hover around zero[18], showing essentially no wage cuts. The predicted low-inflation histogram matches the general character of the actual low-inflation histograms, in Figure 2, well. The histogram for 1996 (not shown in Figure 2), for example, when the median is in the $0.49 - 1.5$ bin, is very similar to that in Figure 5.

In summary, the parameterization in equation (5), appears to capture the salient features of histograms such as those in Figures 1 and 2 extremely well. In particular, during high-inflation periods, the histograms descend reasonably smoothly to zero, while during low-inflation periods, there is little mass below zero, there is a large spike at zero, and weak to no evidence for holes above zero. It remains to check the robustness of these results for other parameterisations and other samples.

3.2 Cost of Living Allowance Clauses

The data in Figures 1 and 2 are based on total wage adjustment for reasons that were noted in section 2. However, since most contracts do not have COLA clauses, it is useful to check how our results change if wage adjustment is defined to exclude the contingent increase. The extent to which the results can be expected to differ is conditioned by the following general considerations. First, only 1644 out of the 10945 contracts contain COLA clauses so that total wage adjustment inclusive of COLA increases cannot be very different from WNC. Second, the yield of the COLA clause will depend on

[18]The implication, in Figure 5, that some of the densities are marginally negative could perhaps be avoided with even more complex and flexible specifications, an improvement that would come at the cost of estimating more parameters.

the inflation rate and will be greater in the high-inflation years of 1980-1982 than during the 1990s; indeed, after the early 1980s, when inflation subsided, some of the 'triggers' may not have been exceeded and COLA clauses may not have been activated.

These considerations help explain why, in general, the histograms of wage change exclusive of COLA adjustments are very similar to those which include COLA adjustments in Figures 1 and 2.[19] Notable differences are few in number and are confined to the high inflation years. For instance, in 1979, when average wage adjustment is 8.41% without COLA and 10.64% with COLA, the height at zero is 0.05 when COLA is excluded but there is virtually no mass at zero when COLA is included. Apparently, contracts with zero change in non-contingent wages had some COLA adjustment. A similar pattern holds in 1982 and 1983 when the spike at zero is about 0.03 and 0.10 respectively when COLA is excluded but only 0.005 and 0.06, respectively, in Figure 1. However, in most years, including transition years such as 1984, the histograms are very similar indeed.

Histograms, similar to those in Figures 1 and 2, but excluding COLA adjustments are used to re-estimate the results in Table 3. While some differences in detail are present, the results are, for the reasons noted two paragraphs earlier, generally similar. In Table 4, columns 1 and 2, we repeat, for the reader's convenience, the results for the most complex specification (E3m), reported in Table 3, and compare them to the results which obtain when COLA adjustments are excluded - columns 7 and 8, Table 4. To see

[19]Figures, similar to Figures 1 and 2, but for wage change exclusive of COLA adjustments and for all years in the sample are available on request.

what the parameters mean for the case of the average median ($M = 4.77$) when COLA is excluded, we re-calculated the histogram in Figure 4; this is not reported to conserve space. If we were to report it, the reader would see that the general appearance of the histogram without COLA is similar to that in Figure 4, although the bar at zero is somewhat higher (0.053 rather than 0.039 in Figure 4). Thus, COLA clauses mute rigidity but only mildly so given their low incidence and intensity. We also re-calculated the histogram when $M = 2$, but with COLA excluded; again, in the interests of conserving space, we do not report it. If we were to report it, the reader would see that it is very similar to that in Figure 5, except that the bars just above and at zero are now 0.213 and 0.324 respectively (instead of 0.208 and 0.209 respectively). Thus, at low inflation, the indication of wage freezes is stronger with COLA excluded.

In summary then, our findings are generally similar whether the yield on COLA clauses is included or excluded, although the overall results suggest greater nominal rigidity when COLA is excluded. The incidence and intensity of indexation clauses both declined somewhat during the recent low-inflation period, thereby setting in motion a mechanism which further increases nominal rigidity somewhat.

3.3 Short and Long Contracts

To examine experience in short and long contracts, we defined as short contracts those of less than, or equal to, twelve months in duration[20] and as

[20] Duration is defined as the time interval between the expiry and effective date of each contract. Note that duration increased somewhat during the low-inflation period, thereby

long contracts those with more than twelve months in duration. Figures analogous to Figures 1 and 2, which are not included here but are available on request, suggest that short contracts have histograms which are less symmetric than long contracts in periods of low inflation with more thinning below zero and considerably higher spikes at zero.[21] In Table 4 we report estimates of equation (5) for short (columns 3 and 4) and long (columns 5 and 6) contracts.[22]

The results for the sub-samples of short and long contracts are qualitatively similar to those in columns 1-2, Table 4. The implications of these estimates for histograms evaluated at the respective average values of the medians and at $M = 2$ can be examined in figures analogous to Figures 4 and 5 - not presented here to conserve space. If we were to present the equivalent to Figure 4, the reader would see a striking difference between histograms for short and long contracts - when these are drawn for the average medians of $M = 4.65$ and $M = 5.38$ for short and long contracts respectively. Long contracts have virtually no mass at zero,[23] while short contracts have mass

making it less likely that additional wage flexibility was achieved via more frequent negotiations and wage adjustments.

[21] For instance, during the low-inflation years of 1993-1996, the spike at zero is around 0.7 in short contracts and roughly half that in long contracts.

[22] Note that conventional structural homogeneity tests cannot be based on the results in columns 1-6, Table 4, because the dependent variable for columns 1-2, Table 4, consists of the heights below the median in the histogram for all observations; it is not the result of stacking the data of columns 3-6, Table 4.

[23] The reader is reminded that the wage change analyzed is annualized and takes into account all the years in the contract. Thus, a zero increase in the wage rate is not definitionally less likely in long contracts.

at zero equal to 0.10. This is the highest estimate of wage rigidity in all the sub-cases considered. The histograms would also indicate that rigidity in the sense of a drop-off in density below zero is a feature of short rather than long contracts. A possible explanation was offered in the introduction: Unions are more likely to agree to a freeze over a short period of time. It is interesting that a number of long contracts signed during the low-inflation period contain freezes in the first year, followed by positive wage adjustment thereafter. There is virtually no suggestion in the histograms of holes around zero, suggesting that the menu-cost effects and interactions in the estimated equations are rather weak. When $M = 2$, figures for short and long contracts analogous to Figure 5 would indicate that the pile-up at zero is more intense in the case of short contracts.

4 Conclusion

The parametric approach above captures the salient features of the wage-change distributions in Canadian collective bargaining agreements. The results obtained suggest that nominal wage rates are rigid downwards at times of low inflation and that wage-change distributions, which are largely symmetric at times of high inflation, become asymmetric with significant thinning below zero and mass accumulating at zero during periods of low inflation. The estimated equations contain significant evidence of menu-cost behaviour in several coefficients and median interactions. When the most complex specification is used to generate the predicted histogram at average values of the median, there is some indication of a hole immediately above zero, as suggested for selected years by our histograms. The method used in this paper is similar to that applied by Kahn (1997) to US survey data. She found evidence of thinning below zero and spikes at zero for wage but not for salary earners. Our data relate to wage earners only, and, in this sense, our results are similar. However, the effects identified here for downward nominal wage rigidity are much stronger, perhaps reflecting the fact that our data refer to union contracts only - see the arguments made in the introduction. Kahn (1997), too, finds menu-cost behaviour in her estimated equations.

These results, which are based on wage-change distributions that include COLA adjustments, are not substantially altered when indexation provisions are not taken into account. The reason is twofold: First, only 15% of collective bargaining agreements contain COLA clauses and, second, many of these clauses yield very small or no increase in wages, particularly during periods of low inflation. Where differences in the results can be noticed, they naturally suggest that indexation attenuates the effects of nominal rigidity.

When the sample is split into short contracts, whose duration is less than or equal to twelve months, and long contracts, there is clear evidence of considerably more downward nominal rigidity in short than in long contracts, a fact which may reflect the willingness of unions to accept wage freezes for short periods of time only. Indeed, several of the long contracts contain a wage freeze in the first year of the contract, but this is followed by wage increases thereafter. The estimated equations for short and long contracts do show a number of significant menu-cost effects but these are not strong enough to show up in predicted histograms for average values of wage inflation.

Our results, based on Canadian union contract data, complement what has been primarily a US literature and add to the growing international evidence on the extent of nominal wage rigidity contained in Agell and Lundborg (1995), Beissinger and Knoppik (2000), Fehr and Gotte (2000) and Smith (2000). The fact that Canada experienced a rather long period of exceptionally low inflation which is included in our sample helps clarify the processes at work during such periods.

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Table 1

Number of Contracts, Average Wage and Price Inflation by Year

Year	# contracts[1]	WNC	$WNC + COLA$	$\dot{CPI}$[2]
1976/77	226	6.48	8.69	7.55
1978	673	7.12	8.16	8.01
1979	569	8.41	10.64	8.95
1980	520	11.15	12.39	9.13
1981	450	12.76	13.64	10.16
1982	562	9.85	10.31	12.43
1983	643	4.47	4.89	10.8
1984	676	3.45	3.76	5.86
1985	519	3.44	3.78	4.3
1986	551	3.44	3.65	3.96
1987	557	3.56	3.90	4.18
1988	556	4.61	4.92	4.34
1989	493	5.41	5.68	4.05
1990	547	5.43	5.79	4.99
1991	530	3.69	3.89	4.76
1992	632	2.11	2.16	5.62
1993	516	0.65	0.75	1.49
1994	471	0.51	0.60	1.86
1995	460	0.82	0.86	0.16
1996	448	1.14	1.22	2.16
1997/99	346	1.76	1.87	1.62
Total	10945			

[1]Data from Human Resource Development Canada. WNC is the percentage change in the base wage rate over a contract at annual rates. $COLA$, defined similarly, is the wage change arising from indexation.
[2]Data label P100000, obtained from CANSIM. This column is the percentage annual change in the all-items Consumer Price Index.

Table 2

Equation (1) Above and/or Below the Median: $WNC + COLA$

	Above and Below		Below the Median		Above the Median	
	Coef	t-stat	Coef	t-stat	Coef	t-stat
α_1	0.210	30.04	0.211	18.37	0.209	26.04
α_2	0.132	18.89	0.115	10.04	0.149	18.55
α_3	0.063	9.00	0.060	5.23	0.066	8.20
α_4	0.031	4.40	0.034	2.94	0.028	3.45
α_5	0.021	3.07	0.021	1.87	0.021	2.68
α_6	0.008	1.19	0.007	0.58	0.010	1.25
α_7	0.004	0.58	0.001	0.08	0.007	0.89
α_8	0.002	0.27	0.000	0.04	0.003	0.42
($\sum$ Dep. Var.)/Years	0.942		0.450		0.492	
$\sum \alpha$	0.471		0.449		0.493	
Standard error	0.045		0.053		0.037	
R^2	0.710		0.642		0.800	
R^2 adjusted	0.704		0.626		0.791	
Mean of Dep. Var.	0.059		0.056		0.062	
Observations	336		168		168	

Table 3

Intervals Below the Median Only: $WNC + COLA$

	Model E2		Model E3		Med. Interact. Model E2m		Med. Interact. Model E3m	
	Coef	t-stat	Coef	t-stat	Coef	t-stat	Coef	t-stat
α_1	0.215	20.59	0.205	20.72	0.214	21.94	0.211	21.72
α_2	0.122	11.66	0.114	11.72	0.123	13.20	0.122	13.26
α_3	0.075	7.05	0.074	7.49	0.078	8.33	0.078	8.36
α_4	0.043	3.86	0.041	4.02	0.044	4.41	0.044	4.51
α_5	0.033	2.85	0.036	3.34	0.034	3.28	0.038	3.54
α_6	0.028	2.37	0.034	3.09	0.027	2.46	0.026	2.40
α_7	0.032	2.63	0.029	2.82	0.021	1.79	0.018	1.58
α_8	0.025	1.98	0.027	2.52	0.025	2.07	0.018	1.53
β_1	-0.035	-3.69	-0.031	-5.77	-0.042	-4.13	-0.040	-4.41
β_2	-0.033	-2.33	-0.036	-2.84	-0.111	-4.776	-0.111	-4.87
γ	0.038	2.71	-0.129	-3.99	0.145	5.40	-0.217	-2.85
β_3	-0.016	-1.12	-0.014	-1.08	0.006	0.16	0.005	0.13
β_{1M}	–	–	–	–	0.004	1.27	0.006	2.53
β_{2M}	–	–	–	–	0.024	3.52	0.022	3.51
γ_M	–	–	–	–	-0.028	-4.38	0.033	2.50
β_{3M}	–	–	–	–	-0.005	-0.67	-0.004	-0.64
Standard error	0.046		0.044		0.041		0.041	
R^2	0.729		0.757		0.794		0.798	
R^2 adjusted	0.710		0.740		0.774		0.778	
Mean of Dep. Var.	0.056		0.056		0.056		0.056	
Observations	168		168		168		168	

Table 4

Median Interaction Model E3m: Short Versus Long Contracts; $WNC + COLA$ Versus WNC

	All Contracts $WNC + COLA$		Short Contracts $WNC + COLA$		Long Contracts $WNC + COLA$		All Contracts WNC Only	
	Coef	t-stat	Coef	t-stat	Coef	t-stat	Coef	t-stat
α_1	0.211	21.72	0.250	20.23	0.194	20.61	0.182	23.10
α_2	0.122	13.26	0.080	7.14	0.143	15.13	0.102	13.50
α_3	0.078	8.36	0.043	3.75	0.078	8.03	0.065	8.04
α_4	0.044	4.51	0.038	3.12	0.052	5.09	0.040	4.70
α_5	0.038	3.54	0.029	2.30	0.035	3.22	0.044	5.08
α_6	0.026	2.40	0.021	1.59	0.025	2.26	0.034	3.67
α_7	0.018	1.58	0.021	1.50	0.017	1.48	0.031	3.21
α_8	0.018	1.53	0.024	1.61	0.020	1.70	0.029	2.89
β_1	-0.040	-4.41	-0.028	-2.59	-0.041	-4.00	-0.047	-5.84
β_2	-0.111	-4.87	-0.197	-7.66	-0.085	-3.53	-0.072	-4.05
γ	-0.217	-2.85	-0.248	-2.24	-0.166	-2.01	-0.147	-2.34
β_3	0.005	0.13	0.050	1.12	0.047	1.58	0.055	2.09
β_{1M}	0.006	2.53	-0.001	-0.48	0.007	2.19	0.005	2.43
β_{2M}	0.022	3.51	0.043	6.51	0.016	2.40	0.012	2.83
γ_M	0.033	2.50	0.040	2.36	0.019	1.12	0.015	1.45
β_{3M}	-0.004	-0.64	-0.012	-1.39	-0.013	-1.87	-0.012	-2.29
Standard error	0.041		0.050		0.041		0.033	
R^2	0.798		0.721		0.798		0.834	
R^2 adjusted	0.778		0.694		0.778		0.818	
Mean of Dep. Var.	0.052		0.048		0.060		0.052	
Observations	168		168		168		168	

Figure 1
Percentage Wage Change Histograms (1981-1984)

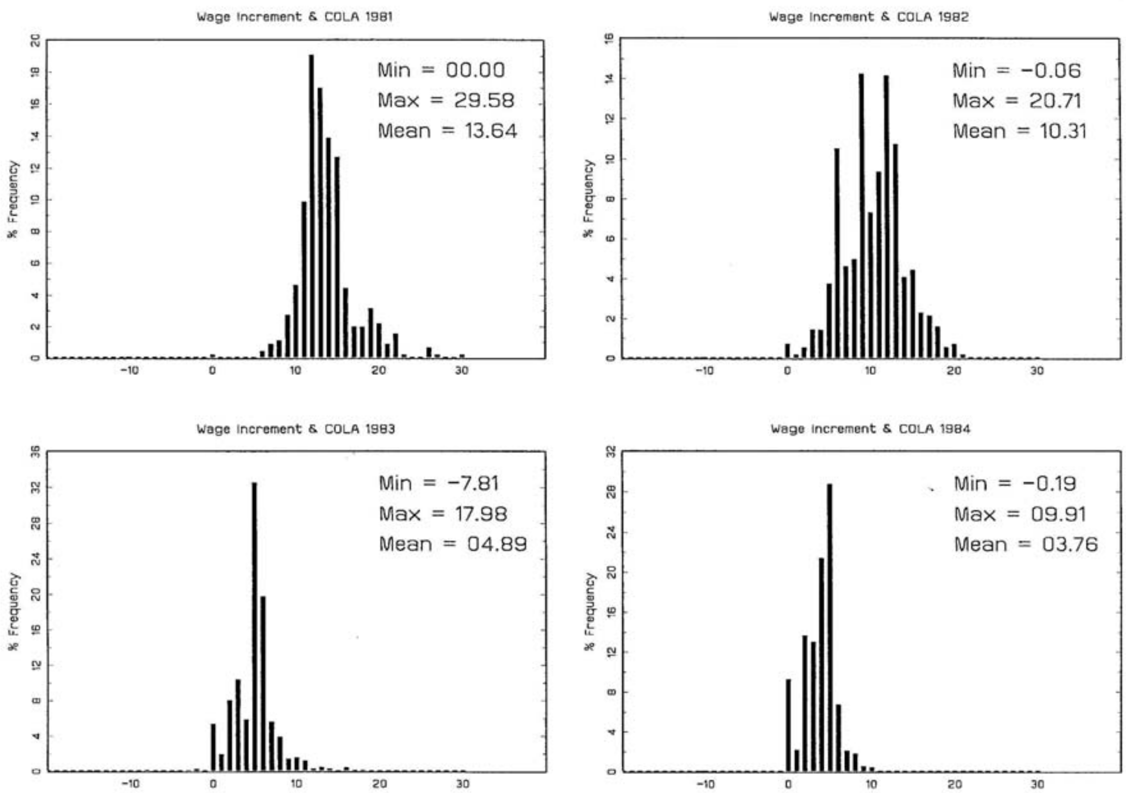

Figure 2
Percentage Wage Change Histograms (1989-1992)

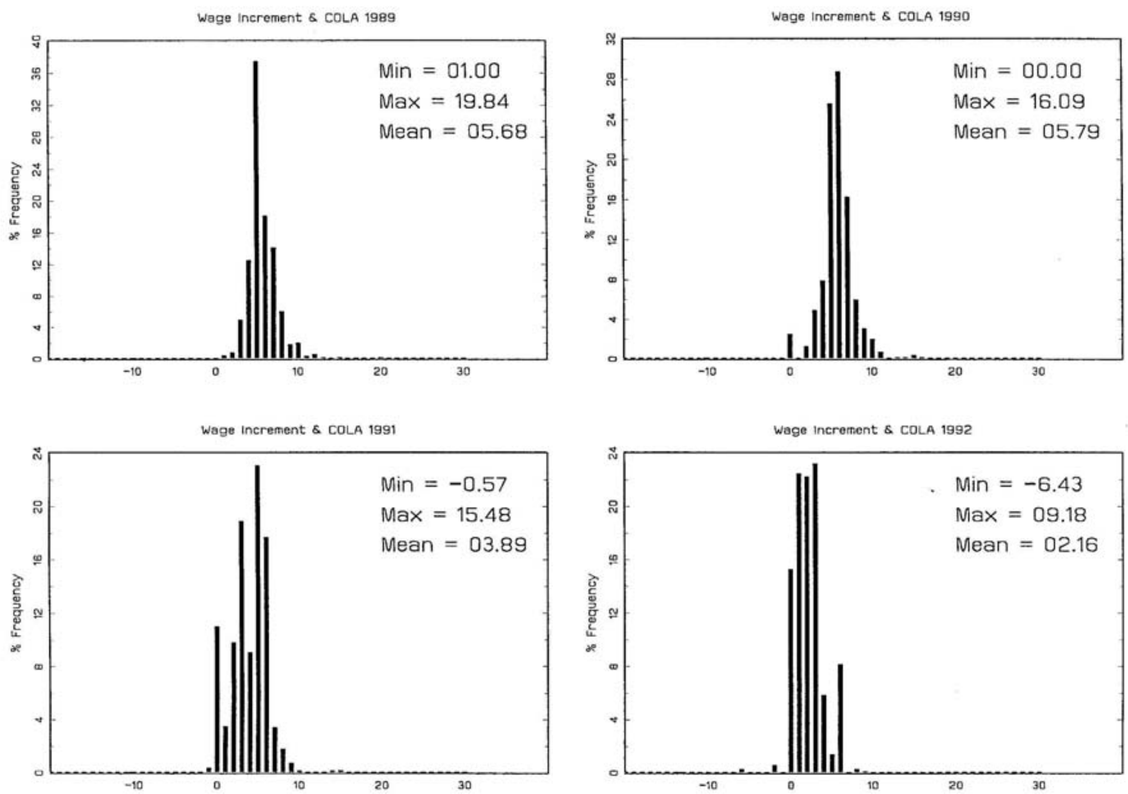

Figure 3
Additional Mass at Zero as a Function of the Median Percentage Wage Change

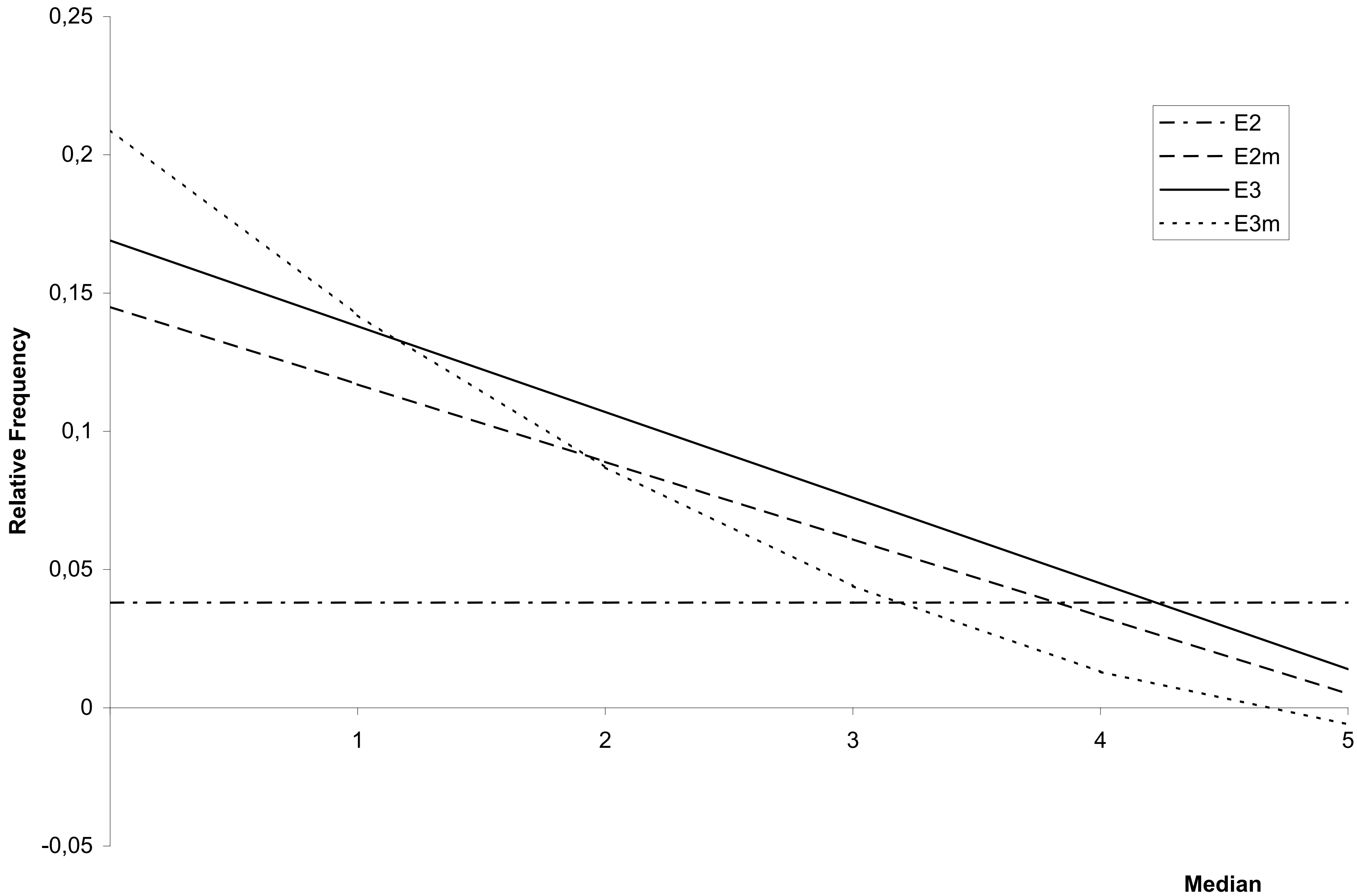

Figure 4
Predicted Histogram: Median Percentage Wage Change = 5.14

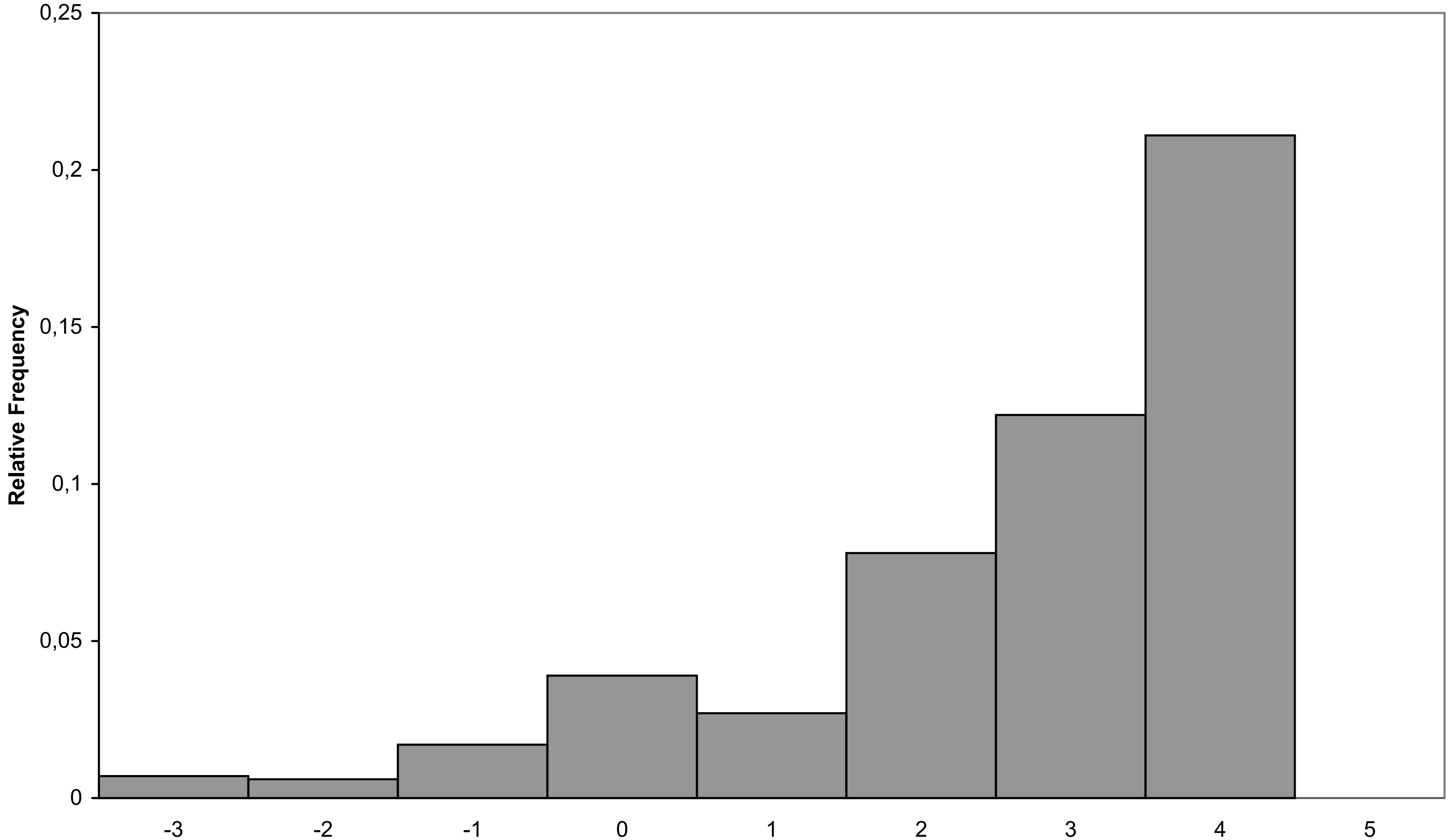

Figure 5
Predicted Histogram: Median Percentage Wage Change= 2