Wage Indexation in Labor Contracts
and the Measurement of Escalation Elasticities*

By

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ABSTRACT

It is now widely acknowledged that the process of wage determination in long term labor contracts may be critical to the determination and persistence of inflation and unemployment in North America. This paper uses Canadian contract data to analyse one important feature of long term contracts with cost of living indexing: the responsiveness of contracted wage rates to changes in prices within the contract period. Surprisingly, there is wide variation in this responsiveness across industries and across contracts. In about 15 percent of contracts average wage earners are over-indexed to inflation, receiving cost of living wage adjustments that increase their wages faster than the rate of inflation. On the other hand, in another 30 percent of contracts average wage earners receive cost of living increases that respond to each percentage increase in prices with a .7 percent or smaller increase in wages. Much of this variation is attributable to systematic industry effects, suggesting that industry characteristics may have a major role in determining the degree of indexation.

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Long term labor contracts are a major institutional feature of North American labor markets. Their presence has been linked to cyclical movements in union-non-union relative wage rates,1/ and more recently to the correlation of inflation and unemployment and the observed persistence of both.2/ While the general nature of these contracts is well known, including the tendency to combine a schedule of wage increments with an indexation rule linking further wage adjustments to movements in the Consumers' Price Index, the wide diversity in actual details between contracts is not. Using a sample of major union contracts from Canada, this paper presents and analyses one important property of contracts with indexation provisions: the responsiveness of contractual wage rates to changes in the price level during the contract period.

Although the theory of optimal indexation has received considerable attention from economists, empirical analysis of the issue has been limited.3/ Furthermore, most previous empirical work has concentrated on measuring the ratio of realized wage increases due to escalation to the increase in prices over the life of the contract.4/ In the presence of substantial non-contingent wage increases, this concern over the total yield of escalation provisions is

1/ Lewis [5], Chapter VI.
2/ See in particular Taylor [9].
3/ Theories of optimal indexation are pursued in Gray [4] and Blanchard [1], for example.
4/ The major American study is Douty's [2] 1975 monograph. Canadian contract data has previously been analysed by Wilton [10].
largely misplaced. On the other hand, more appropriate measures of the elasticity of contractual wage rates with respect to prices reveal several challenging features for the theory of optimal wage indexation, including the presence of elasticities in excess of unity, and a wide variation in elasticities across industries. For example, average wage earners in about 15 percent of contracts are over-indexed to inflation, while in another third of contracts they receive cost of living wage increases that respond to a one percentage point increase in prices with less than a .7 percent increase in wages.

The balance of the paper is divided into three sections. In the first section, the structure of typical indexation provisions in labor contracts is characterized. This leads directly to a class of measures of the elasticity of escalation. In the second section, evidence on the distribution of these measures is presented and analysed. The implications of the findings for the theory of indexation are briefly discussed in the conclusion.

2/ It has lead to a paradox in explaining the responsiveness of wage increases in the major union sector to contemporaneous price increases -- see Mitchell [6], pp. 132-135. Mitchell finds that wage increases are close to unit-elastic with respect to current period price increases. On the other hand, he notes that the ratio of escalation induced wage increases to price increases over the contract is typically one half. However these two numbers do not measure the same thing. This point is discussed further in Section 1 below.
I. **Observed Escalation Formulas: Toward an Appropriate Measure of Ex-Ante Elasticities**

In looking at actual escalated wage contracts, a number of empirical regularities stand out. First, in most escalated contracts, wage indexation provisions are incorporated with substantial non-contingent wage increases. In a study of major American union contracts negotiated between 1968 and 1974, Douty [2] found that the proportion of total realized wage increases due to escalation was about 20% for contracts written in 1973 and 1974, and 10% or less for the earlier contracts.\(^{6/}\) A study of escalated contracts from the Canadian manufacturing sector by Wilton [10], confirms this finding: in none of the years from 1968 to 1975 did the share of realized increases due to escalation exceed 50%.\(^{7/}\)

A second finding is that in a majority of contracts, escalation clauses provide for a fixed cents per hour increment to each workers' wage in response to a given change in the price level. The incidence of proportional percentage formulas (relating a percentage increase in wage rates to each percentage-point increase in prices) is relatively low. In other words, escalation provisions in union contracts often specify the derivative of wage rates with respect to prices, rather than the logarithmic derivative. Furthermore, the same derivative is typically

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\(^{6/}\) Douty [2], Table 9, p. 31.

\(^{7/}\) Wilton [10], Table 7, p. 19. The contracts studies in Section II below are a subset of those investigated by Wilton.
applied to all wage rates in the contract, implying a lower elasticity of indexation for higher wage rates. For instance, a tabulation of union contracts written in New York State in 1979 shows that 19% of all escalated contracts contained provisions relating percentage changes of wages to percentage changes in the Consumer's Price Index. This number was only 11% for contracts in the manufacturing sector, where the traditional formula offering equal cents per hour increases for each point change in the price level was found in 63% of all contracts. Douty does not report strictly comparable figures for the years covered by his study, however some indication of the importance of non-proportional formulas is provided by the fact that 60% of the contracts in 1975 adjusted wages to prices according to one of only four formulas, all non-proportional. Table I shows that only 6 percent of contracts studied in this paper specify proportional percentage escalation formulas.\footnote{8}

A third finding is that in a substantial number of indexed contracts, the escalation formula is restricted in some way: either by specifying the amount by which it can increase wages, or by restricting the circumstances under which it is operative. A specification of an absolute limit on the wage increase provided by indexation (known as cap) was found by Douty in 30 percent of all major contracts in effect in the U. S. in 1975 with provisions for indexation in that year. Table I shows that 35 percent of the Canadian contract sample contain similar specifications.

\footnote{8}{It is interesting to note that most contracts with proportional percentage escalation formulas exhibit unit elasticity of the wage with respect to the price level.}
TABLE 1

Characteristics of Escalated Contracts:
286 Major Contracts Written between 1967 and 1975\(^a/\)

<table>
<thead>
<tr>
<th>Provision</th>
<th>Proportion of Contracts with provision (%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Percentage formula</td>
<td>6.4</td>
</tr>
<tr>
<td>2. Restriction on maximum increase provided by indexation (cap)</td>
<td>35.2</td>
</tr>
<tr>
<td>3. Restriction of escalation to price increases over and above some future price level</td>
<td>31.3</td>
</tr>
<tr>
<td>4. Restriction of escalation to price increases over and above some minimum (trigger)</td>
<td>12.6</td>
</tr>
</tbody>
</table>

\(^a/\) Data Source: the contract data are extracted from details published in various issues of the Collective Bargaining Review. I am grateful to David Wilton for supplying copies of the relevant pages of the Review.
Restrictions on the initiation of indexation are also widespread: 31 percent of the Canadian contracts limit escalation provisions to price increases over and above the price level at some date later than the contract signing date. Alternatively or in addition to these restrictions, some 13 percent of the contracts specify a price increase that must occur before the escalator is operative.

In accordance with this description of typical indexed labor contracts, we give a general form for the path of nominal wages in contracts with non-proportional escalation. Let $w(t)$ represent the nominal wage $t$ periods after the signing date of the contract. For the moment, assume that the contract contains a single wage rate. Then the wage rate at time $t$ is given by

$$w(t) = w(0) + n(t) + \min(c, \alpha \max(0, p(t) - \bar{p})),$$

where $w(0)$ is the wage rate at the signing date, $n(t)$ is the (non-contingent) deferred wage increase in effect at $t$, $c$ is the cap amount or maximum wage increase allowed by the escalator, $\alpha$ is the derivative of wages with respect to prices, $p(t)$ is the price level at $t$, and $\bar{p}$ is the price level at which the escalator becomes effective (the trigger price). The price $\bar{p}$ may be specified at the signing date as a fixed number, or it may be specified as the price level at some future date $t^*$. A reasonable interpretation of this formula is that it represents a piece-wise linear approximation to the desired constant elasticity trajectory for wages. Specifically, suppose that the optimal elasticity of nominal wages with respect to the price level is constant and equal to $\lambda^*$. Then, along the optimal path, the wage rate at time $t$ is given by
\[(2) \quad \log w(t) = \log w(0) + \lambda^* \log p(t) - \lambda^* \log p(0).\]

Treating (1) as an approximation to (2) requires that the two formulas give approximately the same wage rate and elasticity of wages with respect to prices, at least in the interval of prices over which the escalator in (1) is operative. In particular, it will be assumed that \(\sigma, \bar{p}, \) and the non-contingent increments \(n(t)\) are chosen such that both the level of wages and the elasticity of wages with respect to prices given by equation (1) correspond to the optimal values given by equation (2) at the level of prices that trigger the escalator. Let \(\bar{w}\) give the wage rate when indexation starts. By assumption, the contractual elements are chosen to satisfy

\[(3) \quad \Upsilon = w(0) \left(\frac{\bar{p}}{p(0)}\right)^{\lambda^*},\]

and

\[(4) \quad \lambda^* = \sigma \frac{\bar{p}}{\bar{w}}.\]

Equation (3) expresses the requirement that the contractual wage rate given by formula (1) is equal to the wage rate along the constant elasticity wage path, at the price level or date that initiates indexation. Equation (4) equates the elasticity of wages with respect to prices in formula (1) with the ex-ante elasticity \(\lambda^*\) at the price level or date when the escalator becomes operative. The relationship between the desired wage-price trajectory and the linearized trajectory is illustrated in Figure 1. Note the tangency condition at \(p = \bar{p}, \ w = \bar{w}\) imposed by (3) and (4).

It is useful to examine the ex-post elasticity measure in light of this discussion. In a contract of length \(L\), conditional on an end
Figure 1
of contract price level \( p(t) \), the ex-post elasticity is:

\[
e = \min(c, \frac{\alpha \max(0, p(t) - \bar{p})}{p(t) - p(o)} \cdot \frac{p(o)}{w(o)})
\]

If there is no delay in indexation, that is if \( \bar{p} = p(o) \), and if the cap is not effective ex-post then

\[
e = \frac{\alpha p(o)}{w(o)} = \lambda^*
\]

If the cap is effective ex-post, \( e \) will underestimate \( \lambda^* \), according to the interpretation of formula (1) as an approximation to a constant elasticity wage rule. Alternatively, if escalation is delayed (\( \bar{p} > p(o) \)), then even if the cap is not effective by the end of the contract,

\[
e = \frac{\alpha (p(t) - \bar{p})}{p(t) - p(o)} \cdot \frac{p(o)}{w(o)}
\]

\[
= \lambda^* \frac{w(o)}{p(o)} \cdot \frac{w}{\bar{p}} \cdot \left(1 - \frac{\bar{p} - p(o)}{p(t) - p(o)}\right)
\]

Whenever escalation is substantially delayed, the ex-post elasticity \( e \) will be an underestimate of the ex-ante elasticity \( \lambda^* \).

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2/ Since the ex-post elasticity is an arc concept, the choice of dates at which to convert the realized wage change into an elasticity is arbitrary. Conventionally, beginning of contract wages and prices are chosen.

10/ In principle, real wage increases from the signing date to the start of indexation could exactly offset the bracketed term \( \left(1 - \frac{\bar{p} - p(o)}{p(t) - p(o)}\right) \). However, typical increases in real wages are not of sufficient magnitude to offset the effects of a delay in indexation of one year in a three year contract, for example.
The explanation for this result is straightforward. In typical contractual arrangements, (non-contingent) deferred wage increases adjust nominal wages to a level consistent with some amount of price inflation. Increases in prices over and above the level incorporated into non-contingent wage increments raise wages via the indexation function, at least to the maximum allowed by the cap provisions of the escalator. Thus the ratio of the percentage change in the wage effected by the escalator to the total percentage change in prices over the contract is an understatement of the ex-ante elasticity: the correct denominator is the percentage change in prices in excess of the price change anticipated in the non-contingent wage increments. Only when these increments are zero is the ex-post elasticity likely to be a useful estimate of the ex-ante concept.

So far we have not discussed the determination of the various structural elements of the escalation formula given by (1). To do so it is necessary to distinguish between contracts that initiate indexation when the price level reaches a pre-determined level, and contracts that initiate indexation at a pre-specified date. For the former, let $p^T$ give the price level that triggers indexation. The level of wages at the point when $p(t) = p^T$ is uncertain. Let $\tau > 0$ solve $E[p(\tau)] = p^T$; the expected wage rate when indexation commences is $w(0) + n(\tau)$. Equations (3) and (4) become
(3a) \[ w(0) + n(\tau) = w(0) \left( \frac{E}{p(0)} \right)^{\lambda^*} \]

and

(1a) \[ \lambda^* = \alpha \frac{E}{w(0) + n(\tau)} \]

On the other hand, for contracts that initiate indexation on a preannounced date \( t^* \), the price level at the start of indexation is uncertain, although the wage rate is known. Equations (3) and (4) become

(3b) \[ w(0) + n(t^*) = w(0) \left( \frac{E}{p(0)} \right)^{\lambda^*} \]

and

(1b) \[ \lambda^* = \alpha \frac{E}{w(0) + n(t^*)} \]

It is apparent that equations (3) and (4) do not fully specify all the elements of formula (1). For example, contracting parties opting for an indexation rule with a pre-specified trigger price must choose a non-contingent wage schedule \( n(t) \), a wage derivative \( \alpha \), and the trigger price level \( p^T \). One way of viewing this indeterminacy is to note that non-contingent wage increases can be set to compensate for some, none or all of expected price inflation, and the wage derivative and trigger price can then be chosen to satisfy (3) and (4). Unfortunately, a satisfactory explanation for the choice of non-contingent wage increments is not available. Although in the absence of labor turnover and in the presence of perfect capital markets the timing of inflationary wage increments is arbitrary, violation of
either assumption leads to a preference for wage paths that closely track prices and for a reduction in the relative share of wage adjustment accomplished through non-contingent increases. None the less, nearly all escalated labor contracts contain sizeable non-contingent wage increases, and the facts seem to indicate that a majority of inflation-offsetting wage increases are non-contingent. One explanation for the retention of pre-announced increases in escalated contracts is that these increases provide concrete evidence of union leaders' efforts on behalf of membership, and are therefore highly valued by union leaders.\footnote{The problem posed for union leadership by automatic wage adjustment formulas was apparently recognized very early in the history of escalation clauses: Garbarino [3] notes that in the second escalated contract signed by General Motors and the U. A. W., management agreed to a form of the union shop to protect the position of the union.}

Although equations (3) and (4) do not fully specify the elements of (1), they are sufficient to permit an expression for \( \lambda^* \) in terms of observable contract parameters. To ease notation, let \( \lambda^0 = \alpha \frac{p(o)}{w(o)} \). \( \lambda^0 \) is the elasticity of indexation relative to the beginning of contract wage rate and price level. In contracts without delayed escalation, \( p(o) = \bar{p} \) and \( w(o) = \bar{w} \), and hence \( \lambda^* = \lambda^0 \). For contracts with delayed escalation, \( \lambda^0 = \lambda^* \frac{\bar{w}}{\bar{p}} \frac{p(o)}{w(o)} \); \( \lambda^0 \) and \( \lambda^* \) differ to the extent that real wages change over the period from the signing date to the start of indexation. For contracts with a specific trigger price, (3a) and (3b) can be solved to yield

\[
(5a) \quad \lambda^0 = \lambda^* \left[ \frac{\bar{p}}{p(o)} \right] \lambda^*-1.
\]
If \( \lambda^* < 1 \), \( \lambda^0 \) is an under-estimate of \( \lambda^* \), while if \( \lambda^* > 1 \), \( \lambda^0 \) is an over-estimate of \( \lambda^* \), although the error is not large if \( p^T \) represents a small markup on \( p(o) \).

For contracts with a specific starting date for indexation, (4a) and (4b) can be solved to yield

\[
\lambda^0 = \lambda^* \left( \frac{\lambda^* - 1}{\lambda^0} \right)^{\frac{1}{\lambda^* - 1}}. 
\]

Again, \( \lambda^0 \) is an over or under-estimate of \( \lambda^* \) as \( \lambda^* \) is greater than or less than unity.

The effects of the transformation from \( \lambda^* \) to \( \lambda^0 \) are illustrated in Figure 2, which shows a hypothetical sample distribution of \( \lambda^* \) together with the corresponding sample distribution of \( \lambda^0 \). In contracts with delayed escalation, a given \( \lambda^* \) translates into a value of \( \lambda^0 \) further from unity; as a result the distribution of \( \lambda^0 \) is dispersed about the unit elasticity ordinate, relative to the distribution of \( \lambda^* \).

To obtain an estimate of \( \lambda^* \) for escalated contracts with \( p^T \) known, take logarithms of (5a) to yield

\[
\log \lambda^* + \lambda^* \pi - \log \lambda^0 - \pi = 0, \tag{6a}
\]

where \( \pi = \log \left( \frac{p^T}{p(o)} \right) \), and use is made of the approximation \( \log (1 + \pi) = \pi \). For given \( \pi \) and \( \lambda^0 \), (6a) can be solved by Newton’s method for \( \lambda^* \). For escalated contracts with indexation scheduled to commence at a known date \( t^* \), the same manipulations on (5b) yield

\[12/ \] For the set of 36 contracts with pre-announced trigger prices in the sample examined in Section II, the mean ratio of \( p^T \) to \( p(o) \) is 1.083.
(6b) \[ \log \lambda^* - \frac{1}{\lambda^*} g + \log \lambda^0 = 0, \]

where \( g = \log \frac{w(t^*)}{w(c)} \). Again, this equation can be solved by an interactive procedure for \( \lambda^* \), given \( \lambda^0 \) and \( g \).\(^{13/} \)

Equations (6) provide estimates of \( \lambda^* \) from the observable contract parameters consistent with the assumption that the linear escalation rule represents an approximation to an optimal constant elasticity wage-price trajectory. On the other hand, as an estimator of \( \lambda^* \) the number \( \lambda^0 \) is not without appeal, primarily because of its simplicity, but also because it summarizes one aspect of contractual escalation in a model-free way. For these reasons evidence on the distributions of both \( \lambda^* \) and \( \lambda^0 \) are presented in Section II.

Given an actual path for wages as described by formula (1), the target elasticity \( \lambda^* \), the optimal wage path given by (2) and the restrictions imposed by equations (3) and (4) between \( \alpha \) (the derivative of wages with respect to prices) the non-contingent wage increases in the contract and the trigger price level \( p^T \) or the delay to the start of indexation, what can be said about the decision to impose a ceiling on escalation-based wage increases?

Some insight is provided by Figure 1, which shows a less than unit elastic wage-price profile, approximated by a piece-wise linear formula with escalation triggered at the price level achieved at \( t^* \). It is clear from this figure that if \( \lambda^* \) is low enough, a cap on escalated wage increases is desirable. It is also clear that a limit on wage escalation will only be imposed if \( \lambda^* < 1 \).

\(^{13/} \) The mean value of \( g \) for contractual base wage rates is .074 in the subset of contracts with escalation delayed to a future date studied in Section II.
Two issues remain in the discussion of observed escalation formulas. The first of these concerns the widespread use of non-proportional indexation rules that fix the derivative of wages with respect to prices, rather than the logarithmic derivative. Essentially, one should ask why linearized formulas like (1) are used if the optimal wage-price relationship is a constant-elasticity one. However, given the size of nominal wage increases effected by escalators in typical union contract, the error in a non-proportioned formula is likely to be small if the target elasticity is reasonably close to unity. Of course, if the target elasticity is exactly unity and if there are no non-contingent increases after the initiation of indexation (which is the case in some contracts studied below) then a formula like (1) satisfying the appropriate choice of equations (3) or (4) exhibits unit elasticity at every price level. The fact that nonproportional formulas are used so extensively can be taken as evidence that actual target elasticities are closer to unity then might have been suggested by earlier empirical work using ex-post elasticities.

A final detail in the measurement of escalation elasticities is the treatment of aggregation over different wage rates in contracts for which the escalator provides equal absolute increases to all workers. As previously observed, a large majority of escalated contracts have this characteristic. In fact however, it should not

\({1/2}\) This is particularly true of the contract sample studies below, which is heavily weighted toward manufacturing industries.
be surprising that escalator clauses yield equal absolute wage increments in contracts that traditionally specify non-contingent wage increases in this way. Stieber [8] has noted that escalator clauses yielding equal increases have been a major source of narrowing skill differentials in basic steel industries and the same conclusion is surely true in other industries. It seems reasonable to attribute non-proportional cost of living adjustment formulas to the same forces that have lead to narrowing skill differentials in non-escalated union contracts.15/ One fruitful hypothesis in this regard is that equal wage increments represent a dominant social choice option under the democratic voting scheme used by most unions. However, a thorough treatment of the aggregation problem is beyond the scope of this research. The resolution adopted by earlier studies is to confine attention to escalation of base wage rates.16/ It is also possible and useful to consider the elasticity of escalation of other contractual wage rates, and of averages of wage rates within the contract. Because of data limitations, only three wage rates are dealt with below: the base wage rate, the highest contractual wage rate, and a skill-weighted average contractual wage rate.

15/ In fact, in the sample of escalated contracts under study, skill differentials have been narrowed by non-contingent wage increases as well as by cost of living increases. The mean reduction in the ratio of highest skilled workers' wages to the contractual base rate effected by non-contingent wage increases over the life of the contract was 2.91%. Interestingly enough, 3/4 of this relative narrowing was accomplished by wage increases effective at the signing date.

16/ An exception is Wilton [10], who also reports elasticities with respect to a measure of average contractual wage rates, based on 3-digit SIC average hourly earnings.
II. Empirical Evidence on Escalation Elasticities

In this section, the distribution of escalation elasticities for a sample of 281 contracts written in Canada between 1967 and 1975 is presented and analyzed. The purpose is not to test any theory of the determinants of wage elasticities, but rather to present the data in a straightforward manner, so that it is clear exactly what a useful theory must explain. In keeping with the analysis of Section I, new but natural measures of the elasticity of escalation are used.\footnote{So far as I am aware all previous empirical studies have utilized ex-post elasticities.} The results contrast with those obtained by Doughty [2] and Wilton [10] for example, and also present some interesting problems for future theoretical work on contractual escalation.

As a starting point, it is useful to present the sample distribution of the elasticity measure $\lambda^0$, which gives the elasticity of escalation relation to the beginning of contract wage rate and price level. As argued in Section I, this measure is a good estimator of $\lambda^*$, the ex-ante elasticity, for contracts with no delay in the initiation of indexation. For other contracts, $\lambda^0$ and $\lambda^*$ diverge to the extent that real wages change between contract signing and the start of indexation. Since this divergence is likely to be small, $\lambda^0$ provides a simple and reasonably accurate first approximation of $\lambda^*$. 

\footnote{So far as I am aware all previous empirical studies have utilized ex-post elasticities.}
Let $w_b(t)$ represent the contractual base wage rate at time $t$, let $w_s(t)$ represent the highest contractual wage rate at $t$ (paid to the most highly skilled workers) and let $\delta$ represent the proportion of unskilled workers in the contract.\(^{18}\) An average contractual wage rate is given by

$$\bar{w}(t) = (w_b(t))^\delta (w_s(t))^{1-\delta}.$$  

Corresponding to the three wage rates $w_b$, $w_s$, and $\bar{w}$, are three measures of the escalation elasticity of a particular contract based on $\lambda^0$: let

$$\lambda^0_b = a \frac{p(o)}{w_b(o)},$$

$$\lambda^0_s = a \frac{p(o)}{w_s(o)},$$

and

$$\lambda^0 = a \frac{p(o)}{\bar{w}(o)}.$$  

For the small number of contracts in which wages are indexed proportionally to prices, $\lambda^0_b$, $\lambda^0_s$, and $\lambda^0$ are defined to be equal to the factor of proportionality between percentage changes in prices and percentage changes in wages effected by the escalator.\(^{19}\)

\(^{18}\) For each contract, $\delta$ is defined as the proportion of laborers and operatives in total employment of laborers operatives and craftsmen, in the appropriate 3-digit SIC industry. Since breakdowns of industry employment into these classifications are not available from the Canadian census, the industry figures were obtained from the 1970 American Census and SIC codes were matched between Canada and the U.S. The sample mean of $\delta$ is .69, with a maximum of .91, a minimum of .20, and a standard deviation of .14.

\(^{19}\) Among the 18 contracts in the sample with proportional indexation of wages to prices, 15 exhibit unit elasticity.
The sample distribution of $\lambda^o_b$ is shown in Figure 3, and summary statistics are provided in Table 2. The mean of $\lambda^o_b$ is .92 with a sample standard deviation of 0.22. The maximum observed value of $\lambda^o_b$ is 1.91 and the minimum observed value is 0.33; 24% of contracts exhibit values of $\lambda^o_b$ in excess of unity. The sample distribution of $\lambda^o_s$ is shown in Figure 4. The use of the higher wage rate shifts the distribution of elasticities to the left: the sample mean of $\lambda^o_s$ is 0.67, the maximum observed value is 1.34, and the minimum observed value is 0.18. Unlike the distribution of base rate elasticities, the distribution of $\lambda^o_s$ is not single peaked: the presence of the proportionally escalated contracts induces a secondary peak at the unit-elastic ordinate. Finally, the sample distribution of $\lambda^o$ is shown in Figure 5. As expected, this distribution lies between the distribution of $\lambda^o_s$ and that of $\lambda^o_b$. The sample mean of $\lambda^o$ is 0.83 and 14 percent of the contracts exhibit beginning of contract elasticities in excess of unity.

To the extent that the wage rate elasticity defined with beginning of contract wages and prices is a good measure of the desired ex-ante elasticity, the distributions of $\lambda^o_b$, $\lambda^o_s$, and $\lambda^o$ provide interesting input to a theory of wage escalation. Two perhaps surprising findings are the presence of a substantial fraction of elasticities greater than one, and the large overall dispersion in observed elasticities. Recall however that a priori $\lambda^o$ is more dispersed about the unit-elastic ordinate than the ex-ante elasticity $\lambda^*$. To check whether these two findings are a result of the use of beginning of contract wages and prices in computing elasticities, we turn to the adjusted measure of ex-ante elasticities suggested in the discussion of Section I.
TABLE 2

Summary Statistics for the Distributions
of Elasticities shown in Figures 1-6.

<table>
<thead>
<tr>
<th>Variable Name</th>
<th>Mean</th>
<th>Standard Deviation</th>
<th>Minimum</th>
<th>Maximum</th>
<th>Percent in Excess of Unity</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\lambda_0$</td>
<td>.92</td>
<td>.22</td>
<td>.33</td>
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<td>.23</td>
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<td>.18</td>
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<td>$\bar{\lambda}$</td>
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<td>.19</td>
<td>.33</td>
<td>1.67</td>
<td>.14</td>
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<td>$\lambda_0^*$</td>
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<td>.33</td>
<td>1.84</td>
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<tr>
<td>$\lambda_s^*$</td>
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<tr>
<td>$\bar{\lambda}$</td>
<td>.84</td>
<td>.18</td>
<td>.33</td>
<td>1.62</td>
<td>.14</td>
</tr>
</tbody>
</table>
Corresponding to $\lambda^0_b$, $\lambda^0_s$ and $\lambda^0$, define the corrected ex-ante elasticity measures $\lambda^*_b$, $\lambda^*_s$ and $\lambda^*$ as the solutions to the appropriate version of equations (6) with the substitution of the contractual base wage rate, the highest contractual wage rate, and the geometrically weighted average of the two wage rates, respectively. The distributions of these three measures are illustrated in Figures 6, 7, and 8. As expected, the distributions of the beginning of contract elasticities are more dispersed about the unit elasticity ordinate than the distributions of the corresponding ex-ante elasticity estimators. This observation is confirmed by the summary statistics in Table 2: note for example that max ($\lambda^0_b$) $\geq$ max ($\lambda^*_b$), min ($\lambda^0_b$) $\leq$ min ($\lambda^*_b$) and standard deviation ($\lambda^0_b$) $>$ standard deviation ($\lambda^*_b$).

Apparently, neither the dispersion in elasticities nor the presence of elasticities greater than unity is due to the use of beginning of contract wages and prices in defining elasticities for contracts with delayed escalation. If one accepts the skill-weighted geometric mean wage as a good indicator of the average contractual wage rate, then the best estimator of the ex-ante elasticity of escalation for average wage earners in the contracts under study ranges from a low of .33 to a high of 1.62 (Table 2, bottom row). Figure 8 shows that in 26 percent of the contracts, $\lambda^*$ lies

20/ If $\lambda^*$ is symmetrically distributed about unity, and if the transformation factor from $\lambda^*$ to $\lambda^0$ is distributed independently of $\lambda^*$, then $\lambda^0$ represents a mean preserving spread from $\lambda^*$. If the mean of $\lambda^*$ is far from unity, then the transformation to $\lambda^0$ does not necessarily increase the sample variance, as a comparison of the standard deviations of $\lambda^0$ and $\lambda^*$ reveals.
in the interval (.65, .75). However, in 14 percent of the contracts \( \bar{\lambda}* \) exceeds one and in 15 percent of contracts, \( \tilde{\lambda}* \) is less than one half.

An obvious question with relevance for testing theories of the determination of escalation elasticities concerns the relative shares of systematic and unsystematic variation in \( \lambda* \). Prior to formal testing of any particular model of escalation elasticities, one might ask how much of the dispersion in \( \lambda* \) could possibly be explained by such a model. If the variance in the estimators of \( \lambda* \) is due solely to measurement error, then a theory of cross-contract variation in \( \lambda* \) is superfluous and tests of such a theory will never give rise to 'statistically significant' results.

One way of answering this question is to group contracts in some meaningful way and then compare within-group and across-group variation in the measure of \( \lambda* \). Of course, the role of a theory of escalation elasticities is to provide insight into the relevant grouping schemes. However, in this case the data suggest a natural scheme. The contracts are drawn from a set of approximately 60 3-digit industries and involve some 150 different unions. A simple test for the presence of systematic variations in the ex-ante elasticity estimators is a test of the significance of industry and union dummy variables in a cross-contract linear regression. Should such a regression prove significant, then the data provide at least some support for a theory of escalation elasticities based upon inter-industry and inter-union differences. In fact, if a theory relates escalation elasticities to particular inter-industry and inter-union differences, then a strong test of that theory is a comparison of a fixed-effects regression using industry and union dummy variables with a
regression that explains the dispersion in $\lambda^*$ by a restricted set of
industry and/or union characteristics.\textsuperscript{21}

Table 3 lists the distribution of the contract sample 19 industry
groups. The majority (71 percent) of contracts are drawn from the manu-
facturing sector, but contracts from mining, forestry, transportation,
public utilities and service industries are also included. Table 3 also
gives the distribution of contracts by union. Note the large proportion
of contracts signed by two unions: the Autoworkers and the Steelworkers.
This concentration reflects the substantial contribution of automotive,
mining, iron and steel, machinery, and electrical industries to the sample.

Table 4 reports summary statistics for the regressions of the various
elasticity increases on 19 industry-group variables and 3 union variables.\textsuperscript{22}
In every case, the 22 explanatory variables are highly jointly significant,
although in no case does the $R^2$ of the regression exceed .25. The results
suggest that at this level of aggregation by industry there is a statisti-
cally significant inter-industry/inter-union dispersion in elasticity
estimates, although the share of observed dispersion in the elasticity
measures attributable to the industry and union groupings is not overwhelm-
ing.

\textsuperscript{21}For example, a theory might relate the elasticity measure $\bar{\lambda}_{ij}^*$ from the
$1^{th}$ contract in the $i^{th}$ industry to a set of industry characteristics $x_j$.
Comparing the regression $\bar{\lambda}_{ij}^* = x_j' \beta + \varepsilon_{ij}$ with the regression $\bar{\lambda}_{ij} =
\frac{1}{n_j} \sum_{i,j} \bar{\lambda}_{ij}^* + u_{ij}$, where $n_j$ is the number of contracts in the $j^{th}$ industry,
allows a test of the hypothesis that inter-industry differences in $\bar{\lambda}^*$ are
indeed explained by the factors $x_j$.

\textsuperscript{22}The unions specified are the Autoworkers, the Steelworkers, and inde-
pendent or non-affiliated unions and employee associations.
### TABLE 3
Distribution of Contract Sample by Industry Group and by Union

<table>
<thead>
<tr>
<th>Industry Group:</th>
<th>Percent of Contracts</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. forestry, pulp and paper, wood prod's</td>
<td>4.3</td>
</tr>
<tr>
<td>2. furniture and fixtures</td>
<td>1.8</td>
</tr>
<tr>
<td>3. food processing</td>
<td>6.7</td>
</tr>
<tr>
<td>4. automobiles, auto-parts, tires</td>
<td>16.0</td>
</tr>
<tr>
<td>5. clothing and textiles</td>
<td>2.5</td>
</tr>
<tr>
<td>6. iron and steel, smelting</td>
<td>10.6</td>
</tr>
<tr>
<td>7. metal fabrication</td>
<td>6.0</td>
</tr>
<tr>
<td>8. agricultural equipment</td>
<td>2.1</td>
</tr>
<tr>
<td>9. aircraft</td>
<td>3.9</td>
</tr>
<tr>
<td>10. rolling stock and shipbuilding</td>
<td>1.8</td>
</tr>
<tr>
<td>11. electrical appliances and equipment</td>
<td>11.7</td>
</tr>
<tr>
<td>12. non-metallic mineral processing</td>
<td>4.3</td>
</tr>
<tr>
<td>13. miscellaneous manufacturing</td>
<td>2.1</td>
</tr>
<tr>
<td>14. automobile carriers</td>
<td>3.2</td>
</tr>
<tr>
<td>15. public transportation</td>
<td>5.3</td>
</tr>
<tr>
<td>16. public utilities</td>
<td>5.3</td>
</tr>
<tr>
<td>17. food stores</td>
<td>2.1</td>
</tr>
<tr>
<td>18. retail, service</td>
<td>2.5</td>
</tr>
<tr>
<td>19. metal mining</td>
<td>7.5</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Union:</th>
<th>Percent of Contracts</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. United Auto. Workers</td>
<td>26.0</td>
</tr>
<tr>
<td>2. United Steelworkers</td>
<td>18.5</td>
</tr>
<tr>
<td>3. Teamsters</td>
<td>3.6</td>
</tr>
<tr>
<td>4. United Electrical Workers</td>
<td>3.6</td>
</tr>
<tr>
<td>5. Machinists and Aerospace Workers</td>
<td>2.5</td>
</tr>
<tr>
<td>6. United Rubber Workers</td>
<td>2.1</td>
</tr>
<tr>
<td>7. International Chemical Workers</td>
<td>1.4</td>
</tr>
<tr>
<td>8. Textile Workers</td>
<td>1.4</td>
</tr>
<tr>
<td>9. Other national/international unions</td>
<td>50.8</td>
</tr>
<tr>
<td>10. Unaffiliated locals, employee assoc.</td>
<td>11.4</td>
</tr>
<tr>
<td></td>
<td>$\lambda_{0}$</td>
</tr>
<tr>
<td>-------</td>
<td>---------------</td>
</tr>
<tr>
<td>$R_{a}^{2}$</td>
<td>.23</td>
</tr>
<tr>
<td>$F_{b}$</td>
<td>3.68</td>
</tr>
<tr>
<td>Prob. value $^{c}$</td>
<td>.0001</td>
</tr>
</tbody>
</table>

*a/* \( R^2 \) of the regression of dependent variable on 19 major industry group dummy variables and 3 union variables.

*b/* F-statistic for the null hypothesis that all coefficients are zero. The F-ratio has 22 and 259 degrees of freedom.

*c/* Probability of obtaining a larger F-ratio under the null hypothesis.
### TABLE 5

**Explanatory Power of Industry Variables: Disaggregated**

<table>
<thead>
<tr>
<th></th>
<th>$\lambda_b^*$</th>
<th>$\lambda_s^*$</th>
<th>$\lambda^*$</th>
<th>$\lambda_b^*$</th>
<th>$\lambda_s^*$</th>
<th>$\lambda^*$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$R^2$ a/</td>
<td>.38</td>
<td>.34</td>
<td>.39</td>
<td>.37</td>
<td>.31</td>
<td>.39</td>
</tr>
<tr>
<td>$F$ b/</td>
<td>3.21</td>
<td>2.70</td>
<td>3.35</td>
<td>3.08</td>
<td>2.35</td>
<td>3.35</td>
</tr>
<tr>
<td>Prob. Value c/</td>
<td>.0001</td>
<td>.0001</td>
<td>.0001</td>
<td>.0001</td>
<td>.0001</td>
<td>.0001</td>
</tr>
<tr>
<td>$F$ d/</td>
<td>2.22</td>
<td>2.17</td>
<td>2.30</td>
<td>2.17</td>
<td>1.85</td>
<td>2.27</td>
</tr>
<tr>
<td>(19 vs. 45 controls)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

---

a/ $R^2$ of the regression of dependent variable on 45 industry dummy variables.

b/ $F$-statistic for the null hypothesis that all coefficients are zero. The $F$-ratio has 45 and 236 degrees of freedom.

c/ Probability of obtaining a larger $F$-ratio under the null hypothesis.

d/ $F$-statistic for the loss in explanatory power in going from 45 to 19 industry controls. The $F$-ratio has 26 and 236 degrees of freedom: $F_{0.05}^{26,236} = 1.55$. 
Regressions controlling for a larger number of different unions yield about the same explanatory power as those reported in Table 4: even at moderate levels of aggregation by industry the effects of individual unions on escalation elasticities are small. This should perhaps not be surprising given that most unions are concentrated in particular industries.

To test whether a finer disaggregation of contracts by industry is warranted by the data, regressions were also performed on a set of 45 industry dummy variables. Because of collinearity problems, union variables were excluded. The results, reported in Table 5, show that at this lower level of aggregation, about 35 percent of the total variation in each of the elasticity measures is accounted for. The F-ratios reported in the bottom row of the table test the restrictions implicit in moving from 45 to 19 industry variables. In every case, the data reject the higher level of aggregation. These tests imply that a successful theory of the determination of escalation elasticities must distinguish rather carefully between industries, in order to account for the loss of information in aggregating from 45 to 19 industry controls.

Summary and Conclusions

This paper has discussed the interpretation of escalation provisions in major union contracts, and has provided evidence on the distribution of one important descriptive statistic across such contracts — the elasticity of indexation of nominal wages to prices. The major empirical findings are first, that escalation elasticities vary widely across contracts
and are by no means restricted either to the interval between zero and unity or to a narrow band around unity; and second, that a significant portion of this variation is due to differences in average escalation elasticities between industries. These findings present a serious challenge to the model of optimal indexation proposed by Gray [4], which predicts elasticities between zero and one, and unit-elastic indexation only in economies with negligible real shocks. They are less damaging for an alternative model of Blanchard's [1], which gives the deviation of the optimal escalation elasticity from unity as a function of both the elasticity of labor demand with respect to raw materials prices and the correlation of aggregate and raw materials prices. It remains an interesting question for future research whether or not the dispersion in a particular set of industry characteristics can account for the observed pattern of escalation elasticities across industries.

While the discussion in this paper has concentrated almost exclusively on the responsiveness of contractual wage rates to changes in the price level during the contract period, a number of other provisions also vary widely across contracts. Among these are the length of contract, the relative sizes of contingent and non-contingent wage increases, and changes in intra-contract relative wage rates. A complete understanding of the wage determination process in long term contracts will require empirical and theoretical analysis of these and other provisions.
REFERENCES


