Regional Wage Determination
and the Interregional Transmission of Wage Changes

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In the Canadian regional economics literature, a tendency toward wage parity across the country is presented as a possible explanation for persistent regional disparities in unemployment rates. Regional differences in resource endowments, productivity, transportation costs, and product prices tend to produce an equilibrium set of real wage differentials between various regions. But if workers in a low-wage region are successful in obtaining wage parity with workers in more prosperous regions, the "equilibrium" regional wage differentials will dissolve and be replaced by persistent regional unemployment rate differentials—see, for example, Anderson (1988:Chaps. 3 and 6).

In their 1981 study of regional disparities, Swan and Kovacs used Newfoundland data to examine the evidence in favour of regional wage parity. "Our maintained hypothesis is that numerous social pressures would exist, trying to push wages toward parity with the more fortunate parts of the country. This assumption about wage parity pressures is a second key aspect of our theory . . . . [U]nions in bargaining will cite wage rates for comparable employees across the country" (see Swan and Kovacs 1981:Chap. 3). In the Swan and Kovacs model, regional wage changes are assumed to be determined by market forces

(such as provincial unemployment rates) and by comparisons of wages in the rest of the country.

A tendency toward wage parity across regions also has implications for national anti-inflation policies. If one region of the country (for example, central Canada) is experiencing excess demand in the labour market, higher wage settlements in central Canada would be transmitted to other regions where labour market conditions might be very different. If the government believes that wage inflation in one region will permeate the rest of the country, then anti-inflation monetary policy is likely to be more restrictive than if inflationary pressures remain localized. Whether wage settlements in one sector of the economy are transmitted to other sectors of the economy may have important implications for both unemployment and inflation.

In this article, we will (1) analyze the wage determination process in different regions of Canada, (2) measure the extent to which real wage changes are transmitted from region to region, both through industry channels and through general spillover effects, and (3) measure wage spillover effects within regions. We also present evidence on the hysteresis hypothesis which implies that wage changes depend on the change in unemployment rather than on the level of unemployment.

Theoretical Background

The issue of interrelationships and interdependencies of wage rates in the labour market has a long history in economics. In Chapter 2 of *General Theory*, Keynes points to relative wage effects to explain why nominal wage rates might be rigid in a downward direction: "Any individual or group of individuals, who consent to a reduction in money wages relatively to others, will suffer a *relative* reduction in real wages, which is sufficient justification for them to resist it" (Keynes 1936:14). Throughout the 1940s and 1950s, much of the literature emphasized the relative distribution of wage rates and the importance of maintaining an appropriate set of wage differentials between occupational-industrial groups in the labour market. An aberrant settlement by one bargaining group could "spill over" into wage settlements of other groups as each bargaining group attempted to restore the historical pattern of wage differentials.

The subsequent literature developed a number of explanations for wage spillover mechanisms. For example, Phelps (1968) emphasized the important role of turnover costs (recruitment, hiring, training, and firing costs) in the wage determination process. The profit-maximizing firm will establish an optimal wage differential between its own

wage rate and the wages paid by competing firms. The maintenance of this optimal wage differential will give rise to spillover effects between firms. More recently, labour turnover cost models have been subsumed within the growing efficiency wage literature.

The basic contention of efficiency wage models is that worker effort depends on the real wage paid by firms. Higher wages not only increase individual effort but also reduce turnover costs—see Stiglitz (1974) and Salop (1979). In addition, Shapiro and Stiglitz (1984) have stressed the connection between reductions in shirking costs (slacking-off) and higher wages, and Stiglitz (1976) has pointed out that higher wages give firms access to higher-quality applicants. Akerlof (1984) recognizes that higher wages can lead to higher morale among workers. Although efficiency wage models have been developed primarily to explain persistent, involuntary unemployment, they also imply that the real wage paid by a firm depends on the wages paid by other firms—that is, efficiency wage models are consistent with wage spillover effects. Assuming worker effort is positively related to the wage rate, firms have an incentive to pay workers a markup over the worker's expected alternative wage. Thus, efficiency wage models provide an explanation for wage spillovers between firms and a role for the unemployment rate in determining wage *levels*.

Another recent development in macroeconomics is the staggered wage contract literature. Taylor (1980) explained the persistence of unemployment and inflation in terms of a rational expectations, staggered contracts model. The key feature of Taylor's model is that "when making wage decisions, firms (and unions) look at the wage rates that are set at other firms and which will be in effect during their own contract period. . . . In effect, each contract is written relative to other contracts" (Taylor 1980:2). Taylor proposes a "simple and plausible wage setting procedure" in which current wage settlements depend on (1) past and future wage settlements, with wage settlements more distant in time given less weight, and (2) excess demand in the labour market. In this article, we explicitly account for expectations of future price movements rather than future wage settlements. In addition, we measure the influence of "past wage settlements" in a number of ways, and our empirical results provide considerable support for both of Taylor's assumptions.

Model Specification and Data

As discussed in Christofides et al. (1980a, 1980b), there are a number of important econometric advantages in using *micro* wage contract data (before aggregation) to study wage movements and wage spillover ef-

fects.[1] In this article, we analyze 1,841 individual wage contracts signed in the Canadian private sector between January 1, 1979, and March 1989[2]—a period of very diverse macroeconomic conditions (Table 1).[3]

Following the tradition of the empirical wage spillover literature—see, for example, Christofides et al. (1980b), Cousineau (1987), Drewes (1987), Lacroix and Dussault (1984), Lacroix and Robert (1986), Rabeau (1989), and Vroman (1984)—we analyze the determinants of the annualized percentage change in the nominal wage rate over the life of each contract. Our main focus is on the role of wage spillover effects of recent wage settlements in current wage changes.

In any examination of the evidence for regional linkages in wage determination, it is important to distinguish between nominal and real wage changes. Clearly, if all wages rise at a common rate of inflation, we would observe identical wage growth in all regions with no change in relative real wages (or in real wage differentials). The issue of regional wage spillovers therefore concerns the linkages between real wage changes. Moreover, these linkages can follow alternative routes. For example, the emphasis on "comparable workers" in the Swan-Kovacs hypothesis could be interpreted to mean that the direct linkages are confined to specific industries. Workers in industry X in region A may be able to negotiate real wage increases similar to those received by "comparable employees across the country"—that is, workers in industry X in region B. If, in addition to these industry links, real wage changes in region A depend on other real wage changes in region A (including industry X), then there will be indirect links from industry X in region B to all industries in region A. If these potential real wage interdependencies are strong, they would provide an explanation for persistent regional disparities in unemployment rates. We test for wage spillovers (net of expected inflation) both from within the industry and from various geographical locations.

Since recent wage settlements will be correlated with the dependent variable because of the effects of such common influences as inflation and excess demand conditions, we include measures of inflation expectations, inflation catch-up, and provincial unemployment rates in our regression models. Our proxy for the expected rate of inflation, PE,

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1. For example, the use of aggregate quarterly or annual data discards valuable information on the timing, identity, and location of recently negotiated wage settlements in other firms.
2. Information on the wage contracts signed in this period was obtained from a Labour Canada user tape, and all contracts for which a complete set of explanatory variables could be obtained were used in the empirical analysis.
3. This sample period begins one year after the termination of wage and price controls, which were initiated in 1975. For an analysis of wage determination in the aftermath of controls, see Christofides and Wilton (1985).

TABLE 1 Average Annual Percentage Changes in Wage Rate (W) in Indexed (COLA) and Non-indexed (Non-COLA) Wage Contracts and Unemployment Rates (UR) by Region, 1979-1989

	1979	1980	1981	1982	1983	1984	1985	1986	1987	1988	1989
Atlantic											
W Non-COLA	9.4	13.1	13.7	11.3	6.0	3.4	4.4	2.5	3.1	3.3	NA
UR	12.9	11.0	11.6	14.2	16.3	16.8	15.9	18.0	14.7	13.8	12.3
Quebec											
W COLA	12.2	11.5	13.9	11.4	5.8	4.2	4.1	3.5	3.9	4.3	NA
Non-COLA	9.9	11.9	12.6	9.8	5.8	4.7	3.7	4.0	3.9	3.7	NA
UR	9.6	9.9	10.4	13.8	13.9	12.8	11.9	11.0	10.3	9.4	9.4
Ontario											
W COLA	11.3	11.4	13.3	12.3	6.7	5.4	4.6	4.2	4.3	4.5	5.0
Non-COLA	8.9	11.9	11.5	10.4	6.6	4.3	4.1	3.6	4.1	4.5	4.4
UR	6.5	6.8	6.6	9.7	10.4	9.1	8.0	7.0	6.1	5.0	5.0
Prairies											
W Non-COLA	9.7	11.4	14.6	11.4	5.3	2.9	3.1	2.4	2.5	3.6	2.8
UR	4.3	4.3	4.5	7.5	9.7	9.8	9.2	8.9	8.6	7.9	7.4
British Columbia											
W COLA	11.5	13.8	14.1	10.3	5.9	2.2	2.2	2.5	2.8	4.7	NA
Non-COLA	9.6	10.6	15.2	10.8	4.3	2.8	2.2	1.85	2.1	4.7	4.5
UR	7.7	6.8	6.7	12.1	13.8	14.7	14.1	12.5	11.9	10.4	9.2
National inflation	9.2	10.2	12.5	10.8	5.8	4.3	4.0	4.2	4.4	4.0	5.0

Note: NA = not available. "National inflation" is the annual percentage change in the Consumer Price Index (CPI) for Canada.

is based on the average of 17 different forecasts of the future inflation rate in Canada, regularly published by the Conference Board of Canada. The inflation catch-up variable measures the amount of uncompensated inflation that occurred during the previous contract because of errors in inflation expectations. (A precise definition of the catch-up variable is presented in the next section.) If firms and workers take account of past forecast errors, then this inflation catch-up variable will play a role in current wage changes. Empirical support for the importance of inflation catch-up is presented in Christofides et al. (1980a) and Prescott and Wilton (1992).

As noted above, Taylor's model of wage determination includes a measure of labour market conditions. Traditional Phillips curve approaches have used the unemployment rate to quantify the state of excess demand in the labour market. The efficiency wage literature, however, argues that the *level* of unemployment influences the *level* of wages rather than the change in wages. In a similar vein, Blanchard and Summers (1986) propose an insider-outsider model of

the labour market to account for the persistence of unemployment—the so-called hysteresis effect. This model also predicts that wage changes are a function of the first difference of the unemployment rate rather than the level of the unemployment rate. Because we have followed the literature in estimating the wage equation in difference form, we have included both the current and the lagged provincial unemployment rates in our estimating equation. The lagged provincial unemployment rate is measured at the time the previous contract was signed. Thus, we allow the data to determine the manner in which unemployment affects wage changes. To the extent that the coefficient on the lagged unemployment rate is positive and comparable in size to the negative coefficient on the current unemployment rate, this would be evidence of a hysteresis effect—that is, expected real wage growth would depend on the change in the unemployment rate, not the level of unemployment.[4]

Wage spillover variables can be measured in two different ways. As discussed earlier, Swan and Kovacs emphasized the importance of "comparable" workers in the wage transmission process. The most direct comparisons can be made between workers in the same industry since they are likely to have similar levels and types of education and training. Given that a large part of human capital is acquired through on-the-job training, it is likely that labour mobility is greatest within industries rather than between industries. To the extent that labour markets are defined at the industry level, we would expect to observe stronger wage spillover effects between contract settlements within an industry than between settlements that are made in different industries. Accordingly, our sample of 1,841 contracts is segregated into 38 different industry groupings (by three-digit SIC code), and we generate separate measures of past wage settlements for each of the 38 industries.

A second definition of wage spillovers is based on geographic considerations. Since search and moving costs likely depend positively on distance, it seems plausible that spillover effects will weaken with distance. We expect wage settlements in a given city or local market (such as Calgary) to be most influenced by settlements in the same local market. Wage settlements farther away, such as in all areas of the province (Alberta) or region (the Prairies) or in Canada as a whole, will have a much smaller effect on local wage settlements.

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4.   Nested within our estimation equation is a conventional price expectations-augmented Phillips curve. In this context, if the unemployment rate variable was specified as the deviation from the natural rate, the presence of hysteresis would be detected by the inclusion of the lagged unemployment rate (for the natural rate) with an expected positive coefficient of comparable magnitude to the negative coefficient on the current unemployment rate.

To capture this geographic dimension of wage spillover effects, we first define 38 distinct local labour markets. In many cases, these local labour markets can be fairly narrowly defined in terms of geographic area (such as Montreal, Ottawa-Hull, Toronto, Windsor-Sarnia, Brandon, Regina, and Calgary). In a limited number of cases, we were forced to group towns in wider regions.[5] We then create a set of geographic spillover variables based on the location of recent wage settlements. From a small geographic area (the local labour market), we widen the geographic bounds to include the entire province or region and then all of Canada. Finally, to test the hypothesis that wage settlements in the southern Ontario labour market spill over into the various regions of Canada (where labour market conditions may be quite different), we also have created a geographic spillover based on recent wage settlements in southern Ontario.

In summary, five spillover variables are constructed in the following way. For each observation of the dependent variable, we calculate the average size of all wage settlements signed during the previous 12 months (1) within the same specific industry (WSIC), regardless of geography; (2) within the same local labour market (WGEO1), regardless of industry; (3) within the same province or region (WGEO2), regardless of industry; (4) within all of Canada (WGEO3), regardless of industry; and (5) within southern Ontario (WGEO4), regardless of industry. Obviously, the value of these spillover variables will depend on the expected inflation rate, PE, in the economy. As discussed earlier, our wage spillover model is specified in real terms; the expected real wage change in firm A is determined in part by recent real wage changes in alternative firms in the same industry or the same geographic area. Given that the expected inflation rate is included in the wage change equation as a separate explanatory variable (with an expected coefficient of unity), we have expressed all spillover variables in real terms by subtracting PE from the relevant measure of recent nominal wage changes.

In an earlier study, Christofides et al. (1980b) found that the spillover effect from previous wage settlements in the same industry depended on the number of days that had elapsed since the previous wage contracts were signed. We expect that, ceteris paribus, more recent changes in industry wage rates will be more informative than earlier changes and will therefore have a larger effect on current wage changes. Given the costs of a job search and the costs of either moving or commuting, it seems plausible that wage spillover effects will also weaken with distance (wages paid by firms a great distance away are less relevant). Thus, the weight attached to each previous settlement

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5.   A complete description of the 38 locations can be obtained from the authors.

within the industry is postulated to decline linearly with the number of elapsed days (DAYS) since the previous contract was signed and the distance (DIST) between the previous settlement and the current (dependent variable) settlement.

Our sample of 1,841 Canadian wage contracts signed during the 1979-1989 period contains 647 contracts that contained a cost-of-living allowance (COLA) clause. When modelling wage-setting behaviour, it seems reasonable to allow for the possibility that the structure of the wage equation might depend on whether or not the contract is indexed for inflation. In a recent study, Prescott and Wilton (1992) found evidence that supports this contention. By separating the indexed (COLA) and non-indexed (non-COLA) contracts in this study, we can examine whether spillover effects are different in these alternative types of wage contract.

Estimation Equations

As discussed in Prescott and Wilton (1992), the division of a set of wage contracts into COLA and non-COLA groups may result in non-random sampling and thus sample selectivity bias in parameter estimates. The samples of contracts included in the separate non-COLA and COLA regressions were not chosen in a random manner since the COLA decision is taken at the level of the individual bargaining unit. It can be shown that the non-random allocation of observations between two regressions leads to inconsistent parameter estimates or so-called sample selectivity bias. Since the inconsistency can be expressed in terms of a missing variable problem, we can obtain consistent estimates of our wage equations by correcting for the missing variable. This is done through a two-step procedure. The first step is to estimate a probit model of the decision to include or exclude a COLA clause in the wage contract. From the estimated probit equation, SELECTIVITY variables are generated (one each for the COLA and non-COLA wage equations). When these are included in the wage equations, the usual non-linear least-squares estimator is consistent. In the interest of space, we do not report the five estimated regional probit equations, although the variables used in the probit analysis are described in the Appendix. For a more complete discussion of the sample selectivity bias problem and an example of our estimated COLA incidence equation, the reader is referred to Prescott and Wilton (1992).

For contracts that are *not* indexed to the inflation rate (the non-COLA set), the wage equation to be estimated takes the following form:

$$\hat{W} = \beta_0 + \beta_1 PE + \beta_2 CU + \beta_3 UR + \beta_4 UR_{-1}$$
$$+ \sum_{j=1}^{J} (\beta_5 + \beta_6 DAYS_j + \beta_7 DIST_j)(WSIC_j - PE) / J$$
$$+ \sum_{k=1}^{K} \beta_8 (WGEO_k - PE)/K + \beta_9 SELECTIVITY \qquad (1)$$

where $\hat{W}$ is the negotiated annual percentage change in wages over the contract period. The first line reflects the impact of the expected inflation rate, PE (its coefficient β_1 is expected to have a value of unity in the non-COLA regression), the inflation catch-up variable that we referred to earlier (CU), and the current (UR) and lagged (UR_{-1}) unemployment rate values.

The CU variable, which has been found to play a role in wage determination (see Christofides et al. 1980a and Prescott and Wilton 1992), measures the extent of uncompensated inflation during the previous contract. If the previous wage contract had 100 percent COLA coverage, there would be no uncompensated inflation to carry over into the next contract. Generally, however, the COLA elasticity is less than unity, and in most non-COLA contracts the previous contract also had no COLA clause. Thus, errors in past inflation forecasts generally lead to uncompensated inflation. To the extent that firms and workers take previous forecast errors into account during current negotiations, our inflation catch-up variable will play a role in the wage change equation. Following the earlier work cited above, uncompensated inflation in the previous contract is measured by

$$CU = (1 - \theta_{-1})[PA_{-1} - \beta_1 PE_{-1}]L_{-1}/L$$

where θ is the COLA elasticity; PA and PE are actual and expected annual inflation rates, respectively; L is the current contract length; and β_1 is the coefficient attached to current inflation expectations in the wage equation (β_1 is the proportion of expected inflation built into current contracts). The subscript (-1) indicates that the variable refers to the previous contract. The ratio L_{-1}/L ensures that uncompensated past inflation is measured on an annual average basis over the life of the current contract. Notice that if the previous COLA elasticity is unity, then the catch-up term is zero. In cases where the previous COLA elasticity is zero, the catch-up expression simplifies to

$$CU = [PA_{-1} - \beta_1 PE_{-1}]L_{-1}/L$$

The second line of equation (1) specifies the industry spillover variable (WSIC) and its interactions with time (DAYS) and distance (DIST). The parameter J refers to the number of contracts in the same industry that fall within the 12-month period before the current con-

tract. Of course, J varies from observation to observation, but to keep our notation as simple as possible we have omitted the observation index in equation (1). The term $\Sigma_{j=1}^{J}(WSIC_j - PE)/J$ therefore represents the average real wage change within the industry during the previous year. The parameter β_5 represents the effect this average wage change has on the wage change in the current contract. If β_6 and β_7 are found to be negative and significant, contracts signed longer ago and farther away have less impact on the current contract than contracts that are closer in time and space. The third line includes the generic geographic wage spillover variable WGEO, which represents the four alternatives, WGEO1-WGEO4, described above. The parameter K refers to the number of contracts signed in the previous 12-month period, which are therefore included in WGEO1-WGEO4. As with J, K varies from observation to observation. Finally, the SELECTIVITY variable is generated by the first-stage probit analysis of COLA incidence.

COLA contracts differ from non-COLA contracts in that inflation compensation is broken into non-contingent and contingent components. The nominal value of the non-contingent part of the wage package is determined at the time of signing, but the nominal value of the contingent part depends on the size of the negotiated COLA elasticity and the actual inflation that emerges during the contract. In general terms, the COLA contract wage equation can be expressed as

$$\tilde{W}_i = \alpha^i PE + \theta^i PE + h(X_i) \tag{2}$$

where $\tilde{W}_i$ is the expected wage increase in the ith contract, inclusive of COLA;

$\alpha^i PE$ is non-contingent inflation compensation;

$\theta^i PE$ is contingent inflation compensation; and

X_i is a vector of explanatory variables.

Since the two inflation coefficients are potentially different in every contract, it is obvious that an equation such as [2] cannot be estimated unless some restrictions are imposed on the α^i and θ^i coefficients. In Prescott and Wilton (1992), it was shown that a simple linear restriction is flexible enough to allow variation in total ex ante inflation compensation across contracts and variation in the composition of ex ante inflation compensation between contingent and non-contingent components. Further, within the general family of models, statistical tests showed that the following special case could not be rejected:

$$\hat{W} = \beta_0 + \beta_1 (1 - \theta)PE + h(X) \tag{3}$$

where $\hat{W}$ is the negotiated wage increase exclusive of the COLA (and we have omitted the contract index).

Thus, the COLA and non-COLA equations are identical except for the specification of the inflation expectations term. Multiplying PE by $(1 - \theta)$ makes equation (1) directly applicable to the set of COLA contracts. It should be emphasized that for the set of 647 current COLA wage contracts, the dependent variable is expressed as the average annual negotiated wage change *excluding all COLA adjustments*. In the COLA specification, we do not expect β_1 to be unity precisely. But we would expect the coefficient representing total ex ante inflation compensation, $\beta_1 + (1 - \beta_1)\theta$, to lie between zero and unity. Given that observed COLA elasticities also lie in this interval, we expect $0 \le \beta_1 \le 1$.

The existence of COLA also has implications for the construction of the spillover variables (WSIC and WGEO1-WGEO4). In calculating these variables, we have estimated the total wage change in a previous COLA contract by adding the negotiated non-contingent wage change to the product of the expected annual inflation rate and the COLA elasticity. Thus, in constructing our spillover variables we have assumed that both past COLA and past non-COLA wage settlements spill over into current COLA (and non-COLA) wage contracts with equal effect.

Given the absence of information on the COLA wage-price elasticity before 1978, earlier Canadian spillover studies (for example, Christofides et al. 1980b and Drewes 1987) had to confine their analysis to wage contracts without COLA clauses. Since many of the larger and more powerful unions have negotiated COLA clauses in their contracts, studies based on the subset of non-COLA wage data from the 1960s and 1970s may have introduced measurement errors into the spillover variables and sample selectivity bias into the estimated coefficients. Unlike these earlier studies, the econometric results presented in the next section fully incorporate COLA contracts into the analysis, including a correction for sample selectivity bias arising from endogenous switching between COLA and non-COLA contracts. Our sample also includes data drawn from the severe recession and rapid disinflation of the early 1980s.

Empirical Results

Non-indexed Contracts

Tables 2 and 3 present econometric results for private sector, non-COLA wage contracts signed during the 1979-1989 period in Quebec (Table 2) and in Ontario, the Atlantic provinces, the Prairie provinces, and British Columbia (Table 3). For a micro study, the overall fit of these

TABLE 2 Wage Change Regressions: Non-indexed Contracts, Quebec, 1979-1989

Independent Variable	(1)	(2)	(3)	(4)	(5)
Constant	3.421 (1.61)	3.334 (1.51)	3.537 (1.50)	3.471 (1.64)	4.186 (0.877)
PE	1.142 (0.084)	1.123 (0.083)	1.101 (0.073)	1.154 (0.069)	1.105 (0.059)
UR	-0.418 (0.088)	-0.415 (0.075)	-0.416 (0.072)	-0.433 (0.113)	-0.436 (0.067)
UR$_{-1}$	0.040[a] (0.092)	0.051[a] (0.093)	0.043[a] (0.090)	0.041[a] (0.091)	
SELECTIVITY	0.507 (0.266)	0.520 (0.266)	0.519 (0.262)	0.502 (0.265)	0.532 (0.259)
WSIC	0.778 (0.125)	0.765 (0.124)	0.736 (0.121)	0.775 (0.127)	0.743 (0.121)
WSIC x DAYS	-0.107 (0.043)	-0.107 (0.043)	-0.095 (0.043)	-0.105 (0.044)	-0.099 (0.043)
WSIC x DIST	-0.150 (0.038)	-0.148 (0.037)	-0.156 (0.037)	-0.150 (0.038)	-0.155 (0.037)
WGEO3	0.064[a] (0.199)				
WGEO2		0.158* (0.209)			
WGEO1			0.240 (0.102)		0.246 (0.102)
WGEO4				0.088[a] (0.243)	
CU	0.052[a] (0.061)	0.050[a] (0.093)	0.046[a] (0.061)	0.051[a] (0.061)	
SEE	2.087	2.086	2.071	2.084	2.067
R^2	0.727	0.727	0.731	0.727	0.732

Note: Sample size = 355. Figures in parentheses are standard errors. See Appendix for definitions of variables.
a. Coefficient has t-statistic with a probability value greater than 0.1.

TABLE 3 Wage Change Regressions: Non-indexed Contracts by Region, 1979-1989

Independent Variable	Ontario (1)	Atlantic (2)	(3)	Prairies (4)	British Columbia (5)
Constant	1.499 (0.586)	-0.127 (1.88)	0.572 (1.58)	1.311 (1.14)	3.485 (1.06)
PE	1.037 (0.060)	1.169 (0.152)	1.160 (0.110)	1.256 (0.102)	1.007 (0.087)
UR	-0.243 (0.067)	-0.123 (0.079)	-0.096 (0.068)	-0.331 (0.092)	-0.352 (0.078)
UR$_{-1}$		0.087[a] (0.080)			
SELECTIVITY	0.000 (0.186)	1.393 (0.600)	1.507 (0.604)	1.126 (0.594)	-0.318 (0.536)
WSIC	0.747 (0.106)	0.967 (0.265)	1.371 (0.216)	1.206 (0.233)	0.275 (0.147)
WSIC x DAYS	-0.127 (0.037)	-0.424 (0.116)	-0.483 (0.099)	-0.089 (0.051)	-0.085 (0.051)
WSIC x DIST	-0.107 (0.033)	0.069[a] (0.056)		-0.337 (0.075)	
WGEO2		0.022[a] (0.176)			
WGEO1	0.172 (0.085)				
CU	0.158 (0.064)	0.109[a] (0.095)		0.224 (0.057)	0.422 (0.102)
SEE	1.773	1.787	1.809	1.849	2.007
R^2	0.778	0.846	0.842	0.860	0.842
Sample size	362	93	93	159	225

Note: The figures in parentheses are standard errors. See Appendix for definitions of variables.
a. Coefficient has t-statistic with a probability value greater than 0.1.

estimated wage change equations is surprisingly good,[6] particularly when one considers the diverse macroeconomic conditions found from 1979 to 1989 (see Table 1). Even though the selectivity variable is significant at the 0.05 level only in Quebec and in the Atlantic provinces, we leave this variable in all equations to correct for any potential biases from partitioning wage contracts into separate non-COLA and COLA data sets.

Space considerations do not allow a full set of wage spillover results to be reported for every province.[7] Instead, Table 2 provides complete details on the wage spillover variables for the province of Quebec, the region with the second largest set of observations. Our discussion of Table 2 will indicate how we arrive at the "preferred" regression specifications. Before turning to the spillover variables, we note that price expectations and the current unemployment rate both exert a significant effect on wage changes. The price expectations coefficient is not significantly different from unity (as theory predicts) and the unemployment coefficient suggests that, ceteris paribus, a

6. See, for example, Christofides et al. (1980b), Drewes (1987), Riddell (1979), and Vroman (1984).

7. Readers interested in more detailed results can request Working Paper 9003, Waterloo Economic Series, Department of Economics, University of Waterloo.

2.3 percentage point increase (decrease) in the unemployment rate results in a 1 percentage point decrease (increase) in wages. Variables to measure past unexpected inflation and labour market conditions at the time of the previous settlement are not significant (and are dropped in column [5] of Table 2). There is no evidence of price "catch-up" wage demands or a labour market hysteresis effect in Quebec.

The average value of past wage settlements within the industry across Canada is a highly significant determinant of wage changes in Quebec. As hypothesized, industry wage settlements closest in time and space have the largest effect on current negotiations. In each case, the days and distance interactions have a negative coefficient, which is significantly different from zero. According to Christofides et al. (1980b), industry wage settlements that took place some time ago have a much smaller effect on current negotiations. For example, wage settlements that occurred 365 days ago have only one-half of the spillover effect of settlements that took place one day ago. And the farther away the industry wage settlement, the less effect it has on current wage settlements in Quebec. To illustrate, a Montreal wage settlement 182 days ago would have a 0.56 spillover effect on a current Montreal wage settlement in the same industry. A wage settlement 182 days ago and 3,600 kilometers away (for example, in Alberta) would have a zero effect on a Montreal wage settlement in the same industry (see Figure 1). The spillover effect from an industry wage settlement in Toronto (which is 554 kilometers from Montreal) is diminished by less than 0.1 by the distance interaction variable.

The geographic spillover variables are defined in terms of all contracts signed in the previous year, regardless of industry. Column [1] of Table 2 presents a spillover variable based on all settlements signed in Canada during the last year; successive columns narrow the geographic boundaries to the province of Quebec, the local Quebec labour market, and southern Ontario. Rather interestingly, wage settlements from across Canada, from the entire province of Quebec, and from southern Ontario have no significant effect on Quebec wage settlements (the t-scores for these three geographic spillover variables in columns [1], [2], and [4] are 0.32, 0.76, and 0.36, respectively). The only geographic spillover variable that is significant is from the local Quebec labour market (a t-score of 2.4 in column [3]). Column [5] of Table 2 drops the two insignificant variables (the lagged unemployment rate and price catch-up) and presents our preferred spillover equation for the province of Quebec. The preferred specification retains the most significant spillover coefficients and excludes insignificant coefficients.
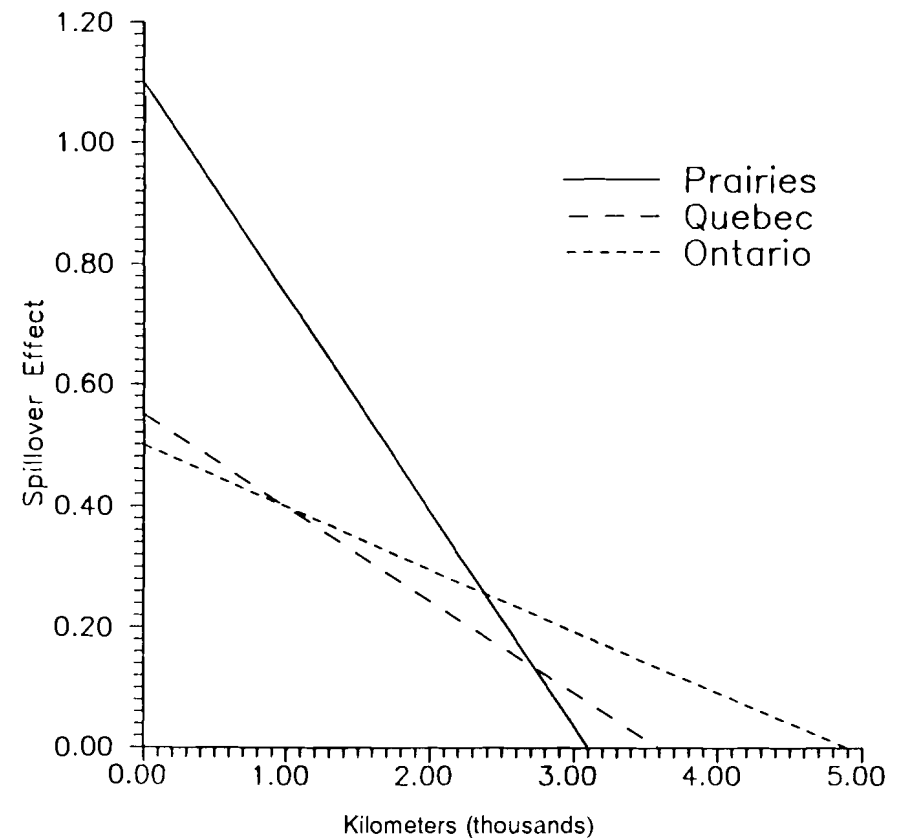
FIGURE 1 Industry spillover effects in non-indexed contracts by distance between wage settlements

Table 3 presents our preferred wage equation for the other four regions of Canada.[8] The price expectations variable is highly significant in all regions, and the coefficient for price expectations is not significantly different from one in Ontario, the Atlantic provinces, and British Columbia. In contrast, the coefficient for price catch-up is much lower, ranging from 0.16 in Ontario to 0.42 in British Columbia (it is not significantly different from zero in the Atlantic provinces). In the non-indexed sector of the economy, workers and firms appear to be largely *forward-looking* when it comes to price inflation.

In all regions except for the Atlantic provinces (where our sample size is relatively small), the provincial unemployment rate is a highly significant determinant of wage changes (with t-scores in the 3.5-4.5 range). In the Atlantic provinces, the unemployment rate has

8.   A full set of spillover regressions is available in Prescott and Wilton (1990).

the expected negative sign but a t-score of only 1.4.[9] Ignoring the Atlantic provinces, movements in the unemployment rate exert the smallest effect on wage settlements in Ontario, where the estimated coefficient is about one-quarter, compared to the Prairie provinces and British Columbia, where the estimated unemployment coefficient is about one-third (in Table 2 the estimated coefficient on the unemployment rate in Quebec is closer to one-half). In every region, the coefficient on the lagged provincial unemployment rate is not significantly different from zero. There is little evidence of a hysteresis effect in Canadian labour markets. Real wage growth depends on the level of unemployment, not the change in the unemployment rate.

Turning to the wage spillover variables, past wage settlements within the specific industry to which the dependent variable belongs have a significant effect on current wage changes (in British Columbia it is a close call with a t-score of 1.9 for the industry spillover variable). In each region the days interaction with the industry spillover variable has the expected negative sign (with a t-score ranging from 1.7 in British Columbia and the Prairie provinces to 4.9 in the Atlantic provinces). With the exception of the Atlantic provinces, the coefficient for the days interaction variable is quite similar in all regions of Canada.

The distance interactions with the industry spillover variable are more interesting. It was hypothesized that industry wage settlements that are farther away should have less impact on new wage settlements. The distance interaction variable turns out to be insignificant for the two coastal regions, the Atlantic provinces and British Columbia. For these regions, it appears that industry wage settlements a long way away (for example, central Canada) are just as important as industry wage settlements within the region. Although the coefficient on the distance interaction variable in the Atlantic provinces has a t-score of 1.2 (see column [2], Table 3), we note that this coefficient is *positive*. More distant industry wage settlements may be given a greater weight in the Atlantic provinces. There is support for the hypothesis that *within a given industry* settlements in central Canada affect wage changes in the same industry in the Atlantic provinces.

In Ontario and the Prairie provinces (and in Quebec in Table 2), the distance interaction with the industry spillover variable has a significant negative coefficient. More distant settlements within the same industry have a smaller effect on current wage changes. Figure 1 plots the industry wage spillover effect in terms of the kilometer distance between the current wage settlement and the previous settlement

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9. Drewes (1987) also found the labour market variable to be insignificant in the Atlantic Region.

within the industry (we assume that this previous settlement took place 182 days ago). The distance gradient is steepest in the Prairie provinces. If the previous industry wage settlement was located 3,100 kilometers away (and 182 days ago), it would have no spillover effect for wage settlements in the same industry in the Prairie provinces. Given the distances involved (Canada is about 7,500 kilometers wide), our results suggest that wage settlements in Ontario and Quebec have little effect on wage settlements in the Prairie provinces in the same industry. Similarly, wage settlements outside Ontario and Quebec have relatively small effects on wage settlements within the same industry in central Canada. For example, a Winnipeg wage settlement (182 days ago) produces an industry wage spillover effect of only 0.29 for a Toronto wage settlement and 0.19 for a Montreal wage settlement. Wage settlements farther west than Winnipeg would have even smaller spillover effects on industry wage settlements in central Canada.

The pattern for the geographic spillover variables in Ontario is the same as in Quebec (see Table 2). Wage spillovers from across Canada or from the entire province of Ontario are insignificant. As found in Quebec, the only significant geographic spillover variable in Ontario arises from within the local labour market (see column [1], Table 3).

Outside of Ontario and Quebec, we could detect no significant geographic wage spillover effects. When price expectations, provincial unemployment rates, and industry-specific wage spillover effects are included in the regression, there is no statistical evidence for a spillover effect based on all Canadian wage settlements for new contracts signed in the Atlantic provinces, Prairie provinces, and British Columbia. Similarly, when one narrows the geographic spillover effect down to the specific region where the dependent variable is located, the regional wage spillover variable is also found to be insignificant in the Atlantic provinces, Prairie provinces, and British Columbia.

Finally, wage settlements in southern Ontario have no significant effect on wage settlements throughout the rest of the country. In all regions of the country, the southern Ontario spillover variable is insignificant. There is no statistical evidence that wage settlements in the booming southern Ontario labour market of the late 1980s "spilled" into other regions of Canada.

Indexed Contracts

Within our sample period (1979-1989), there are 647 contracts with COLA clauses. These are distributed over the five regions in such a

TABLE 4 Wage Change Regressions: Indexed Contracts in Selected Regions, 1979-1989

Independent Variable	Ontario				Quebec	British Columbia
	(1)	(2)	(3)	(4)	(5)	(6)
Constant	2.364 (0.864)	2.743 (0.882)	2.780 (0.860)	2.467 (0.861)	6.494 (0.970)	8.488 (1.45)
PE	0.655 (0.036)	0.685 (0.036)	0.680 (0.035)	0.647 (0.036)	0.600 (0.050)	0.629 (0.100)
UR	-0.360 (0.068)	-0.335 (0.082)	-0.356 (0.069)	-0.364 (0.067)	-0.479 (0.083)	-0.622 (0.106)
UR_{-1}	0.161 (0.079)	0.070 (0.078)	0.075 (0.072)	0.155 (0.079)		
SELECTIVITY	0.415 (0.264)	0.434 (0.270)	0.489 (0.263)	0.453 (0.262)	0.406 (0.285)	0.639[a] (0.620)
WSIC	0.479 (0.110)	0.573 (0.110)	0.553 (0.106)	0.428 (0.100)	0.697 (0.157)	0.756 (0.302)
WSIC x DAYS	-0.080 (0.028)	-0.091 (0.029)	-0.093 (0.028)	-0.079 (0.028)	-0.264 (0.066)	
WSIC x DIST	-0.055[a] (0.049)	-0.044[a] (0.050)	-0.059[a] (0.050)			-0.127 (0.076)
WGEO3	0.494 (0.132)			0.482 (0.131)		
WGEO2		0.248 (0.181)			1.136 (0.190)	0.525 (0.194)
WGEO1			0.319 (0.093)			
CU	0.346 (0.062)	0.417 (0.062)	0.431 (0.060)	0.338 (0.061)	0.286 (0.065)	
SEE	1.708	1.746	1.714	1.709	1.923	1.900
$\bar{R}^2$	0.738	0.727	0.736	0.738	0.782	0.889
Sample size	277	277	277	277	233	75

Note: Figures in parentheses are standard errors. See Appendix for definition of variables.
a. Coefficient has t-statistic with a probability value greater than 0.1.

way, however, that it is not feasible to estimate COLA wage equations for the Atlantic Region and the Prairies. In Quebec and Ontario, there are 233 and 277 COLA contracts, respectively. British Columbia has considerably fewer (75 contracts), but enough for some conclusions to be drawn. The results of estimating wage equations for these three regions are reported in Table 4. More detailed results are reported in Prescott and Wilton (1990).

The main focus of our attention is on the wage spillover variables, which are defined over industries and over space. The most detailed spillover results are reported for Ontario, which has the largest sample of COLA contracts. These Ontario results are given in the first four

columns of Table 4, with the preferred regression appearing in the fourth column of Table 4. The last two columns present the preferred regressions for Quebec and for British Columbia.

Before we look at the wage spillover effects for Ontario, consider the other factors in the wage equations. First, the coefficients on the price expectations variable are remarkably stable across the three provinces. In the preferred regressions (the last three columns), the price expectations coefficients range between 0.60 and 0.65. Since these are COLA contracts, there is no prior expectation that these coefficients should all be unity. The unemployment coefficients do not show the same stability. In all three provinces, the unemployment coefficient is negative and highly significant in a statistical sense, but the coefficients range from a low (in absolute terms) of -0.36 in Ontario to a high of -0.62 in British Columbia. Since none of the lagged unemployment rates is statistically significant, there is no evidence of a hysteresis effect in Canadian COLA wage contracts. On the contrary, both the indexed and non-indexed wage contract samples support a traditional Phillips curve view of wage dynamics. However, given the sharp difference between the unemployment coefficient in Ontario and its counterpart in Quebec and British Columbia, these data imply that the provincial Phillips curves are far from parallel. In particular, the Phillips curve for Ontario is considerably flatter than those for the other two provinces.

Wage changes in COLA contracts are in part determined by past mistakes in inflation forecasts. This is shown by the coefficients for the catch-up terms, which are of similar magnitudes in the three provinces. The catch-up effect is strongest in Ontario, where 34 percent of uncompensated inflation during the previous contract is built into the current contract. In Quebec and Ontario, the small standard errors on the catch-up coefficients imply that they are precisely estimated. But the small sample in British Columbia leads to a wide confidence interval around a similar point estimate (not reported). Finally, the selectivity variables are not statistically significant in any of the three provinces.

As for the wage spillover variables, in Ontario and Quebec there are strong and statistically significant wage spillover effects from previous contracts signed in the same industry. In British Columbia, the industry spillover effect is not so well defined (this may stem from the relatively small sample size). As with the non-indexed contracts, there is a negative temporal effect in the industry spillover in Ontario and Quebec, but we found no such effect in British Columbia. At least in Ontario and Quebec, the more distant in time the previous wage increase, the weaker is its impact on the current contract. But there seems to be no strong evidence for a distance effect on the industry wage

spillover in the set of COLA contracts. Although we found all distance/industry interaction effects to be negative, in none of the three provinces did we find that distance had a statistically significant modifying effect on industry spillovers. In other words, the evidence suggests that in indexed contracts previous own-industry wage increases are important, regardless of where the contracts were signed, and that more recent own-industry wage settlements have the greatest spillover effect.

In the first three columns of Table 4, we introduce alternative regional spillover variables into the Ontario regression. These are the Canada-wide, regional, and local spillover variables. Note the differences between the indexed and non-indexed contracts. The conclusion that emerges from the COLA sample is that the spillover coefficient declines as the geographical region is more narrowly defined. Thus in Ontario, Canada-wide wage spillovers give the lowest standard error of regression. Indeed, the local spillover variable is not statistically significant in Ontario (nor is it in Quebec, although we do not report this regression here). Thus, in all three provinces the results favour the broader geographical spillover variables over the local spillover variable. In addition, we found no evidence in the COLA contracts to suggest that wage changes in southern Ontario play a key role in Canadian wage changes.

Conclusions

This article presents an econometric analysis of the determinants of regional wage changes in Canadian private sector wage settlements over the period 1979-1989. Our main purpose has been to explore the nature of wage spillovers between past and current wage settlements. In non-indexed contracts, the average wage settlement within a specific industry is a highly significant determinant of wage changes within the same industry. In addition, industry spillover effects decline with time, so that more recent industry settlements have a stronger spillover effect than earlier wage settlements. In the three interior regions of Canada, the industry spillover effect also declines with distance. The role of geographic spillovers, regardless of industry, was explored as well. In the sample of non-indexed contracts, broadly defined geographic spillover variables were found to be insignificant. But in Ontario and Quebec, where the data allowed us to define local labour markets most precisely, we found a significant spillover from wage settlements within the local labour market to current wage settlements in the same local market. Finally, there was no evidence of direct wage spillovers from southern Ontario to other regions of the country.

For non-indexed (private sector) wage contracts, there is little statistical evidence to support the assumption of regional wage parity forces. During the 1980s, a period of increasing regional unemployment disparities, wage settlements in central Canada did not have much effect on wage settlements in other regions. While industry-specific wage spillovers do exist, such industry spillovers tend to weaken with time and distance and, more important, do not cross over into other industries.

The analysis of indexed contracts was limited to Ontario, Quebec, and British Columbia. The industry spillover variable was found to be highly significant, but its interactions with time and distance generally were not. In contrast to the non-indexed contracts, the COLA settlements appear to be influenced by broadly defined geographic spillovers. In Quebec and Ontario, the local spillover variable was found to be significant, but t-statistics favoured the more broadly defined geographic spillover variables. Finally, there was no evidence in the indexed sample to suggest that wage changes in southern Ontario play a key direct role elsewhere in the country.

Addison and Burton (1979) have discussed some of the difficulties in distinguishing between neoclassical and institutional models of wage spillover effects. In this article, we have not attempted to test formally alternative theories. Our empirical results are broadly consistent, however, with the neoclassical interpretation. Our findings that wage spillovers are measurable and significant within industries and within local markets are consistent with the hypothesis that industry and local demand or supply disturbances affect all wage settlements within the industry or within the local market. Moreover, the differences we find between the indexed and non-indexed samples regarding the geographic spillovers do not conflict with the neoclassical model. Firms that agree to indexed wage contracts are probably subject to more broadly defined shocks. Christofides (1990) found that indexed contracts are more likely to be found in industries in which industry output prices are more highly correlated with the consumer price index. In our study, this shows up in linkages between (geographically) broadly defined wage spillover variables and current wage changes. But we find no evidence for direct wage spillovers from Ontario to other provinces. Had we found such evidence it would be much more difficult to rationalize the results in terms of a neoclassical model of labour markets.

As for the role of the unemployment rate, Canadian contract data from the 1980s indicate that wage changes depend significantly on the current unemployment rate but not on the first difference of the unemployment rate. Thus, we find no evidence in support of the hysteresis hypothesis. Finally, price inflation plays a significant role

in wage determination. In both indexed and non-indexed contracts, inflation expectations proved to have a positive and significant effect on wage changes. In non-indexed settlements, the coefficient for inflation expectations is not significantly different from unity. The effect of past errors in inflation forecasts was captured by an inflation catch-up variable which was found to be statistically significant in all three COLA samples and in three of the five non-COLA samples.

Appendix: Definition of Variables

The wage data used in this study were taken from a Labour Canada user tape, which provides information on contracts signed in Canada. For each wage contract, the following was recorded: the signing, opening, and closing dates of the contract; the negotiated (non-contingent) wage increases during the contract; the amount paid in a (contingent) cost-of-living allowance (COLA) during the contract; the identity of the union and firm; a three-digit Standard Industrial Classification (SIC) code; and geographic codes. After considerable editing to ensure consistency through time and across bargaining units, the following contract-specific variables were created from the Labour Canada tape:

$\hat{W}$ The average annual percentage change in the negotiated (non-contingent) base wage rate over the life of the current contract.

θ The ex post COLA wage-price elasticity—the percentage change in the base wage rate attributable to the COLA clause divided by the percentage change in the Consumer Price Index during the life of the current contract.

θ_{-1} The ex post COLA wage-price elasticity for the previous contract (for the same bargaining unit).

L Length of the current contract in years.

L_{-1} Length of same bargaining unit's previous contract in years.

PA_{-1} The average annual percentage change in the Consumer Price Index over the life of the previous contract.

PE The average annual inflation rate forecast (by 17 financial or consulting firms, as published by the Conference Board of Canada) at the signing date of the current contract.

UR The seasonally adjusted provincial unemployment rate in the three months immediately preceding the signing of the current wage contract.

UR_{-1} A similarly defined variable for the previous contract.

SELECTIVITY   A variable generated by a first-stage probit analysis of COLA incidence. As discussed in Prescott and Wilton (1992), the following variables are used in the regional probit equations of COLA incidence:
— a dummy variable that has a value of unity if the current contract includes a COLA clause, 0 otherwise (dependent variable);
— a similarly defined dummy variable for the previous contract;
— the standard deviation of the monthly changes in the Consumer Price Index over the five years immediately preceding the signing of the contract;
— the number of employees in the bargaining unit (in thousands); and
— the standard error of estimate from a regression of the log of the industry selling price index on two lagged values of the dependent variable and two lagged values of the log of the Consumer Price Index (using annual data on 79 three-digit SIC industries over the 1961-1986 period, data permitting).

WSIC   The average annual percentage change in the base wage rate for all (spillover) contracts signed in the same three-digit SIC industry as the reference contract during the previous year.

WSIC x DAYS   A weighted average of industry wage changes in the year before the reference contract, the weights being the number of days between the reference and spillover contracts.

WSIC x DIST   A weighted average of industry wage changes in the year before the reference contract, the weights being the number of kilometers separating the reference and spillover contracts.

WGEO1-WGEO4   The average annual percentage change in the base wage rate for all contracts signed in the same geographical labour market during the previous year. These are: 1, the local market; 2, the regional or provincial market; 3, all of Canada; and 4, southern Ontario.

CU   A measure of uncompensated inflation in the previous contract so that $CU = (1 - \theta_{-1})[PA_{-1} - \beta_1 PE_{-1}]L_{-1}/L$.

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