Wage Flexibility and Cyclical Unemployment in Canadian Provinces

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One of the enduring characteristics of the Canadian labour market has been substantial and persistent unemployment rate disparities across provinces. These disparities have long been a concern of government policy-makers and have led to a number of federal government initiatives designed with the intent of reducing these disparities. A small sample of these initiatives include the Atlantic Development Fund (introduced in 1962) designed to promote economic growth in the high unemployment region of the Atlantic provinces, the Fund for Rural Economic Development (introduced in 1966) designed to develop new industry in five high unemployment regions across Canada, the Department of Regional Economic Expansion (introduced in 1968) and, more recently, the Western Development Fund (introduced in 1980). For a comprehensive history of regional development programs in Canada, see Savoie (1992).

An examination of the provincial dispersion of unemployment in Canada over the past thirty years suggests three stylized facts. First, the regional pattern of high and low unemployment rate regions is remarkably stable. The Atlantic region, which comprises Newfoundland, Prince Edward Island, Nova Scotia and New Brunswick, have persistently had higher unemployment rates

than the rest of the country. In fact, with the exception of the 1970-71 period, the average unemployment rate in the Atlantic provinces has been higher than that in every other province every year since 1961. In addition, the other two high unemployment provinces, Quebec and British Columbia, have consistently experienced unemployment rates above that of the four low unemployment provinces (Ontario, Manitoba, Saskatchewan and Alberta). Second, as Swan and Serjak (1991) point out, while federal government policies since that time have resulted a decline in regional income inequalities, regional unemployment rate disparities have not narrowed. The pattern of dispersion is not markedly different in 1986 than it was twenty years earlier. Third, relative unemployment rates show a certain volatility, with substantial one-year narrowing occurring between 1969 and 1970, and again between 1981 and 1982, while in other periods there have been less rapid but no less substantial reductions (1962-66) or increases (1983-86) in the unemployment rate gaps.

Recently, a good deal of attention has been focused on the question of whether differences in real and nominal wage rigidity offer an explanation of the different unemployment rate experiences across countries (Grubb, Jackman and Layard 1983 and Bruno and Sachs 1985) and across regional labour markets within countries Hyclak and Johnes (1989). The purpose of this paper is to ask the same questions of provincial labour markets in Canada. Are there provincial differences in wage rigidity, and if so, can these differences offer an explanation of the different unemployment rate experiences of Canadian provinces?

The answers to these questions are important for policy-makers since differences in wage rigidity imply that nationally applied policies designed to influence cyclical unemployment will have different effects across provinces. This is of special concern to policy-makers in Canada since fluctuations in cyclical unemployment rates have been shown to be mainly due to changes in monetary variables.[1] Since monetary policies are by necessity equally applied across regions, differences in the degree of wage rigidity will result in some regions bearing a larger unemployment cost of disinflationary policies than others. The measures of wage rigidity we derive in this paper will allow us to measure these differential costs of contractionary monetary policies. Measures of these differential costs are of particular interest as the Bank of Canada has, in recent years, implemented a very restrictive anti-inflation tight money policy. While this has led to studies of the cost of this disinflation policy for the Canadian economy as a whole (Cozier and Wilkinson 1990, Lipsey 1990 and York 1990), we know of no effort to examine how these costs may differ

across provinces.

Differences in wage rigidity also offer a partial explanation of unemployment rate disparities, the causes of which have been the subject of much debate in Canada (Courchene 1981 and Matthews 1981). Money supply shocks will expand or contract cyclical unemployment rate disparities depending on whether high or low unemployment regions have relatively rigid wages. Our wage rigidity measures will therefore shed light on this on-going debate.

Despite the importance of having measures of provincial differences in wage rigidity for resolving these issues, little empirical work has been done to provide such measures.[2] This paper is an effort to fill this gap by providing measures of wage rigidity for each of Canada's ten provinces and by examining the extent to which provincial differences in cyclical unemployment experiences can be explained by differences in wage rigidity.

In section 2, we present a model that will provide provincial estimates of wage rigidity measures. In section 3, we discuss estimation issues and present the estimated coefficients of the model. Section 4 presents a discussion of the implications of these results and examines to what extent provincial differences in wage rigidity are correlated with provincial differences in variations in the cyclical unemployment rate. Finally, in section 5, we offer a conclusion.

The Model

In each period, households and firms are hypothesized to calculate a target rate of nominal wage inflation (w_t^*) based on an expected rate of price inflation (p_t^e), the degree of labour market tightness (CUR_t), the expected rate of growth in the marginal productivity of labour (q_t^e) and changes in tax burdens, government imposed wage controls, and other institutional and structural variables that might constrain the choice of w^* at full employment (X_t). Thus we specify the following reduced form target wage equation

$$w_t^* = p_t^e + \tau CUR_t + q_t^e + \beta X_t \qquad (1)$$

where CUR_t is the cyclical unemployment rate in period t defined as the observed minus the natural unemployment rate. Equation (1) imposes a number of restrictions on the long run relationship between rates of change in wages, prices and productivity. The unity coefficient on p_t^e reflects the assumption that anticipated shocks can have no effect on real wages or the equilibrium values

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1.   McCallum (1987) shows that over the period 1954-85 monetary variables were the principal source of the business cycle in Canada. This result accords with the theoretical result suggesting that economic fluctuations in a small open economy with a flexible exchange rate (like Canada) should be mainly due to monetary shocks.

2.   Thirsk (1973) is often cited as providing evidence of provincial differences in the responsiveness of wage to labour market tightness. However, Fortin and Newton (1982) have provided a good deal of evidence to suggest that the measure of labour market tightness employed by Thirsk is an inappropriate measure. Further, Thirsk does not distinguish between types of wage rigidity.

of real variables. The unity coefficient on q_t^e reflects an assumption that, in a long run equilibrium characterized by steady-state inflation, real wages should grow at the same rate as labour productivity implying stable income shares. The estimate of τ measures the response of target nominal wage inflation to changes in the cyclical unemployment rate.

There are two possible sources of labour market rigidities. The first occurs when, following a shock, target real wages do not adjust to clear the labour market. This form of rigidity is usually called real wage rigidity (RWR). There are a number of reasons why such rigidity might exist. According to the efficiency wage literature, real wage rigidity arises because labour productivity is a positive function of the real wage. Thus, in a recession, firms are reluctant to cut real wages and instead opt to reduce employment. Insider-outsider theories suggest firms are less able to adjust real wages in response to recessions due to the currently employed (insiders) using the power they hold due to the costs of hiring and firing to prevent the hiring of the currently unemployed (outsiders) at a lower wage. Implicit contract theories similarly predict real wage rigidity and labour hoarding over the business cycle. $(1/\tau)$ provides a measure of real wage rigidity. As the absolute value of τ falls in value, a real shock causing an increase in labour market tightness has a smaller effect on real wage inflation. Thus, real wages are increasingly rigid as the absolute value of $(1/\tau)$ rises.

The other source of labour market rigidity is the failure of observed nominal wages to adjust instantaneously to changes in nominal target wages. A measure of nominal wage rigidity (NWR) can be created by positing a partial adjustment model of nominal wage inertia.

$$w_t - w_{t-1} = \alpha(w_t^* - w_{t-1}) \quad 0 \leq \alpha \leq 1 \tag{2}$$

The size of α measures the speed with which nominal wage inflation adjusts to the target level. If $\alpha = 1$, nominal wages are perfectly flexible since $w_t = w_t^*$ in every period. If $\alpha = 0$, nominal wages are fixed. Thus $(1/\alpha)$ provides a measure of the degree of nominal wage rigidity (NWR). The size of α has been hypothesized to be a function of a large number of factors. Staggered wage contracts imply that at any point in time some fraction of the total labour force is subject to predetermined wage rates. As a consequence, adjustment of the whole labour market toward the target nominal wage will occur only over a number of periods. In part, then, the size of α reflects the average length of wage contracts. The magnitude of α will also reflect the degree of wage indexation. The higher the degree of indexation of the nominal wage to price inflation the more quickly can wages adjust to changes in the desired nominal wage that result from changes in the expected rate of inflation. Nominal wage rigidity may also arise due to the failure of labour suppliers to immediately adjust price expectations when a nominal shock occurs. Provincial differences in the size of α might also be expected to differ because of differences in the degree of industry competition within those provinces. As

competitive pressures are greater in more open economies we would expect α to be positively related to the openness of the provincial economy.

Substituting (1) into (2), we have

$$w_t = (1 - \alpha)w_{t-1} + \alpha p_t^e + \alpha\tau CUR_t + \alpha q_t^e + \alpha\beta X_t \tag{3}$$

as the basic equation to be estimated. Estimation of (3) requires that we have provincial values of the expected rate of inflation, the cyclical rate of unemployment, the expected rate of productivity growth and the X variables.

The expected rate of inflation is derived by assuming price expectations are formed adaptively. The assumption of adaptive expectations has been offered in the literature as a reasonable specification in a world where information relevant for determining price expectations is expensive to obtain and to process and where knowledge of the true size of shocks is uncertain (see, for example, Friedman 1979). It is also useful to note that deriving rational price expectations in regional economies is made difficult by the lack of data on international and, especially, interregional import price inflation. In small open economies, changes in import prices should play an important role in determining rational price forecasts.[3] Fortin (1991) further notes that over the 1955-90 period the annual change in the CPI inflation rate has rarely exceeded two percentage points, making it likely that economic agents find a simple forecasting rule of thumb to be a cost effective strategy. In recognition of these arguments, we follow Fortin (1991) in assuming agents adopt the following forecasting rule of thumb: $p_t^e = p_{t-1}$. Using the same price expectation assumption, Fortin and Newton (1982) tested the stability and predictive ability of a national version of (3) and found that it tracked history well, was stable, and performed well in *ex-ante* predictions. The assumption that $p_t^e = p_{t-1}$ seems well supported by Canadian data.

Estimation also requires data on the cyclical rate of unemployment for each region.[4] In recent work (Johnson and Kneebone 1991), we have produced provincial estimates of natural and cyclical unemployment rates and we use the cyclical unemployment rate estimates here as our measure of labour market tightness in each province. Our estimates of provincial cyclical rates were

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3.   Note that Blackley (1989) estimates a rational price expectation by regressing rates of CPI inflation for U.S. states against lagged CPI inflation, changes in the money supply and changes in GDP but omits changes in international and interregional import prices as explanatory variables. It may be for this reason that Blackley's results, using his rational price expectation, were unsuccessful (see his footnote 10).

4.   Blackley (1989) uses state total, rather than the cyclical, unemployment rates as a measure of labour market tightness in U.S. states. Unless natural rates of unemployment are constant over time, there is little reason to believe variation in the total unemployment rate is an accurate measure of variations in labour market tightness. Fortin (1989) concludes that, for Canadian data at the national level, the total unemployment rate is not a good measure of labour market tightness.

derived in such a way as to make them consistent with the national estimates gen-erated by Fortin (1989). Using Fortin's estimates in another context, McCallum (1987) has found them to be an accurate measure of the natural rate in Canada.[5]

Our measure of the expected rate of increase in the productivity of labour (q_t^e) reflects an assumption that agents respond only to long run productivity changes. We assume that the expected rate of increase in the productivity of labour is insensitive to short run variations in actual productivity growth. Thus it can be proxied by a variable measuring long-run trend productivity growth.[6] An examination of the time paths of real GDP per employed worker in each of the ten provinces revealed that, in each case, productivity had been growing at a slightly decreasing rate over the time period under investigation. However, in each province there is clear evidence that the decline in the rate of growth accelerated in the middle to late 1970s. Several different models were tried. The preferred model involved fitting a linear spline regression of time on real GDP per worker. The switching point was dictated by the data and allowed to vary from province to province. In every province the null hypothesis that the slope was the same in both sub-periods was firmly rejected by the data.

The fitted values from this regression give, if plotted, a positively sloped straight line with a kink at the knot. The slope of this line is the absolute change in real output per person per year. Since the absolute change is a constant the percentage change must be falling. This formulation fit better than one that allowed the percentage change to be a constant rather than the absolute change to be a constant -- undoubtedly because real productivity was growing at a decreasing rate over that time period. After the knot, the slope of the line declines indicating an even slower rate of growth in productivity. *PRD* (the variable designed to proxy (q_t^e)) was created as the percentage change in the fitted (that is, trend) values of the linear spline regression.

Finally, the *X* variables represent other constraints on the rate of change in the target money wage. Changes in these variables affect the ability of labour to win real wage gains in a full employment market. Although a number of variables have been suggested in the literature, only two proved to be significant in any of the regressions. *UI* measures the rate of change in an index measuring the generosity of unemployment insurance and *AIB* is a dummy variable to control for the effects of the Anti-Inflation Board (1976-78). The

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5.   Using a somewhat different approach, Burns (1990) has produced provincial natural rate estimates similar to those in Johnson and Kneebone (1991).

6.   A common assumption made in the literature is to assume the rate of productivity growth has been constant. Thus q^e can be represented by a constant term. It is commonly recognized, however, that the trend rate of growth of productivity has significantly slowed since the 1960s (see, for example, the discussion in Grubb, Jackman and Layard 1983; Kahn 1984; and Fortin 1989). Thus, the assumption of a constant rate of productivity growth over our estimation period (1962-86) would not seem to be appropriate.

exact definitions of these variables are provided in the data appendix.[7]

The equation estimated for each province is given by

$$w_t = (1-\alpha)w_{t-1} + \alpha p_{t-1} + \alpha\tau CUR_t + \alpha PRD_t + \alpha\beta_1 AIB_t + \alpha\beta_2 UI_t \qquad (4)$$

Estimation Issues and the Results

Equation 4 was estimated for each of Canada's 10 provinces using data covering the period 1961-1986. Because of contemporaneous correlation between provinces, Zellner efficient estimators (SUR) were used to derive parameter estimates.

In order to test the linear homogeneity restriction, the equations were first estimated by unconstrained SUR. The sum of the unconstrained point estimates of the coefficients on w_{t-1} and p_{t-1} proved to be very close to unity in all provinces but one, ranging between 0.960 (Newfoundland) and 1.053 (Ontario). The one exception was P.E.I. where the point estimates summed to 0.875. The null hypothesis that the coefficients summed to unity in each province was tested and could not be rejected.[8]

A maintained hypothesis of the model is that, in a steady-state, real wages should increase at the same rate as trend productivity growth. This hypothesis was tested by estimating the model with a constant term to pick up any trend change in real wages not explained by the other variables. The null hypothesis that the constant was equal to zero in each equation could not be rejected, individually or jointly, at the 1 percent confidence level. Furthermore, the coefficient estimates of the other parameters proved to be quite robust with respect to the inclusion or exclusion of the constant.

In order to obtain final estimates of the coefficients equation (4) was re-

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7.   In addition to *UI* and *AIB*, we also included *TAX* (the percentage change in the effective indirect tax rate) and *MWAGE* (the percentage change in the legislated nominal minimum wage). One would expect a positive sign on *TAX* and, because increases in minimum wages and the generosity of unemployment insurance increase the reservation wage of labour, one would expect a positive sign on *MWAGE* and *UI*. The *AIB* dummy variable is expected to have a negative sign as wage controls reduce the ability of labour to win real wage gains. Previous studies on the impact of the Anti-Inflation Board (referred to in the text) indicate the program reduced wage inflation by approximately 2 percentage points in each year of the program. The estimated coefficients on *MWAGE* and *TAX* proved to be insignificant in every formulation and thus they were dropped from the estimating equation.

8.   The test of the joint hypothesis that the coefficients sum to unity in every province has a $\chi^2(10)$ distribution. The test statistic generated a value of 13.51 which implies the null hypothesis cannot be rejected even with the probability of a Type I error set as high as 20%.

estimated by constrained SUR. The results are presented in Table 1.[9]

On the whole, the results must be judged to be quite good. The point estimates of α range between 0.260 (Quebec) and 1.06 (New Brunswick). The coefficient on the labour market tightness variable ($\alpha\tau$), although insignificantly different from zero in three of the smaller provinces, has the expected negative sign in all provinces.[10] The magnitude of the coefficient on *CUR* is also generally consistent with the results of recent studies of the national labour market suggesting a value of between -0.36 (Rose 1988) and -0.56 (Coe 1985). The estimated effect of the wage and price controls of the Anti-Inflation Board are also consistent with previous Canadian estimates (see Rose 1988; Fortin 1989; and Riddell and Smith 1982) that these controls reduced wage inflation by approximately two percentage points in each year in which the controls were in effect. Finally, the estimated coefficients on the index of unemployment insurance generosity are significantly different from zero in four provinces and these coefficients have the expected positive effect on wage inflation.

Implications of the Results

Previous work on regional labour market differences in Canada (Johnson and Kneebone 1989) has found significant differences in labour market adjustment between regions in Canada, but some substantial homogeneity within regions. To investigate the provincial and regional homogeneity of the wage determination mechanism in Canada we tested the null hypotheses that the coefficients are equal in all provinces in Canada, in all provinces within the Atlantic region (Newfoundland, P.E.I., Nova Scotia, New Brunswick), and in all provinces within the Prairie region (Manitoba, Saskatchewan, Alberta). The relevant Wald Chi-square statistics, along with the resulting p-values, are presented in Table 2.

Differences in the estimated value of α describe differences in the degree of nominal wage flexibility. The greater the absolute value of α, the greater is the adjustment of nominal wage inflation to target wage inflation and the smaller the degree of nominal wage rigidity. The hypothesis that α is the same in all provinces in Canada was strongly rejected and was rejected as well for all provinces in the Prairie region at the 5 percent level of significance. A

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9.   The equation was estimated in the form:

$(w_t - w_{t-1}) = \alpha(p_{t-1} + PRD_t - w_{t-1}) + \alpha\tau CUR_t + \alpha\beta_1 AIB_t + \alpha\beta_2 UI_t.$

However, estimating it in the form of equation 4 provided almost identical estimates of coefficients and standard errors. The two approaches are equivalent asymptotically.

10. In studies of regional (rather than provincial) labour markets, Wilton and Prescott (1990) found the unemployment rate had a statistically insignificant effect on wage inflation in the Atlantic provinces. Our results show this is true only for two of the four Atlantic provinces suggesting it is useful to examine provincial rather than regional labour markets.

TABLE 1 SUR Estimates of Equation (4), Annual data 1961-1986

Province	α	$\alpha\tau$	$\alpha\beta_1$	$\alpha\beta_2$
Newfoundland	0.800[a]	-0.667[a]	-3.00[b]	0.0038
	(8.06)	(4.16)	(2.50)	(1.29)
Prince Edward Island	0.717[a]	-0.382	-1.74	0.0106[b]
	(6.06)	(1.58)	(1.02)	(2.47)
Nova Scotia	0.773[a]	-0.815[a]	-0.75	0.0046[b]
	(7.92)	(4.09)	(0.91)	(2.18)
New Brunswick	1.06[a]	-0.219	-2.35[c]	0.0081[b]
	(8.53)	(0.77)	(2.02)	(2.69)
Québec	0.260[c]	-0.346[a]	-1.77[c]	0.0008
	(1.78)	(3.29)	(2.02)	(0.45)
Ontario	0.461[a]	-0.341[b]	-1.52[b]	0.0010
	(3.61)	(2.82)	(2.23)	(0.64)
Manitoba	0.750[a]	-0.203	-2.20[b]	0.0063[a]
	(6.25)	(1.37)	(2.65)	(3.15)
Saskatchewan	0.464[a]	-0.395[c]	-2.08[c]	0.0035
	(3.89)	(1.88)	(1.78)	(1.66)
Alberta	0.364[a]	-0.329[a]	-2.87[a]	-0.0007
	(4.80)	(3.86)	(2.99)	(0.38)
British Columbia	0.534[a]	-0.406[a]	-1.58	-0.0007
	(3.36)	(5.03)	(1.69)	(0.34)

Note: Figures in parentheses are t-statistics.
a.   significantly different from zero at 1% significance level, two-tailed test
b.   significantly different from zero at 5% significance level, two-tailed test
c.   significantly different from zero at 10% significance level, two-tailed test

somewhat greater degree of homogeneity with respect to α is evident in the Atlantic region where the hypothesis that α is the same in all four Atlantic region provinces was not rejected at the 5 percent level of significance. Our point estimates clearly show a higher degree of nominal wage adjustment in the Atlantic region than in the rest of the country. In fact, in none of the four Atlantic provinces is the estimated value of α significantly less than unity.

The anti-inflation program was designed as a national program to be applied equally in all provinces, thus there was no expectation that the effect of that program on wage inflation would be different from province to province. The tests of the hypotheses of an equal *AIB* coefficient in all provinces in Canada, and in all provinces within the Prairie region, could not be rejected, as expected. The null hypothesis of equal coefficients within the Atlantic region was rejected at the 5% level.

Previous work on the regional effects of unemployment insurance in Canada (see, for example, Maki 1977; Johnson and Kneebone 1991; and Rose 1988) suggest significant regional differences in the UI impact of unemployment insurance on labour market adjustment. The findings of this paper re-

TABLE 2 Wald Chi-square Statistics for the Null Hypothesis of Equal Coefficients Within Regions

Coefficient	Canada $\chi^2(9)$	Canada p-value	Atlantic $\chi^2(3)$	Atlantic p-value	Prairie $\chi^2(2)$	Prairie p-value
α	41.95	0.0000	6.39	0.0973	7.05	0.0294
$\alpha\tau$	19.50	0.0213	6.53	0.0885	0.79	0.6722
$\alpha\beta_1$	14.85	0.0950	9.94	0.0190	0.83	0.6594
$\alpha\beta_2$	23.38	0.0044	4.38	0.2237	10.12	0.0063

Note: The p-values give the probability of having obtained the observed value of the χ^2 statistic under the assumption that the null hypothesis is true. A p-value of 0.10 is equivalent to saying that the null is rejected at the 10% confidence level.

affirm this result as the null hypothesis of equal *UI* coefficients across all provinces was rejected at the 1 percent confidence level. We can also reject the hypothesis for the Prairie provinces with p-value of less than 1 percent. In the Atlantic provinces, however, we again find a somewhat higher degree of homogeneity than is evident elsewhere in the country. That is, the null hypothesis of equal *UI* coefficients for provinces in the Atlantic region is not rejected even at the 10 percent confidence level. It is also interesting to note that, in three of the four Atlantic provinces, increases in the generosity of unemployment insurance have a statistically significant positive effect on real wage inflation whereas this is true in only one of the remaining six provinces. This suggests that labour markets within the Atlantic region are more seriously affected by changes in the federally administered unemployment insurance program than labour markets elsewhere in the country.[11] In particular, it implies that real wages in that region are sensitive to changes in the unemployment insurance program. An interesting implication of this result is that the federally administered unemployment insurance program has an important role to play in affecting unemployment rate disparities across provinces. That is, an increase in the generosity of unemployment insurance increases real wage inflation in the Atlantic region relative to elsewhere and thus creates a market incentive that tends to worsen unemployment rate disparities.[12]

Differences in the estimated value of $\alpha\tau$ describe differences in the slope of the short run trade-off between cyclical unemployment and wage inflation. The greater the absolute value of $\alpha\tau$, the smaller is the ability of a provincial labour market to absorb unemployment shocks without wage inflation. Point es-

timates are noticeably higher in two provinces (Newfoundland and Nova Scotia) than elsewhere. In particular, they are higher than the values in the two largest labour markets, Quebec and Ontario, both of which have quite similar values of $\alpha\tau$. Based on the point estimates, wage inflation in Newfoundland and Nova Scotia is, therefore, much more sensitive to unemployment shocks than is the case for the greatest part of the Canadian labour market.

The magnitude of $\alpha\tau$ can also be used to infer the unemployment rate costs imposed by disinflationary policies of the central bank. Specifically, central bank disinflationary policy will impose fewer costs, in terms of increases in the unemployment rate, the larger is $\alpha\tau$. The magnitude of this cost can be measured by a 'sacrifice ratio' the size of which is determined by the degree of nominal (NWR) and real (RWR) wage rigidity in each provincial labour market.

Our provincial estimates of measures of NWR, RWR and sacrifice ratios are presented in Table 3. The sacrifice ratio is defined as the cumulative percentage points of unemployment that results from a permanent one percentage point reduction in the rate of price or wage inflation (these will be the same in the steady-state). From (4), this sacrifice ratio is given by the absolute value of $1/\alpha\tau$.[13] Sacrifice ratios are an especially interesting measure for regional labour markets within a national economy, for all such regions are subject to the same monetary policy. Ignoring the three provinces with insignificant estimates of the coefficient on CUR_t, our results show that efforts by the Bank of Canada to permanently reduce the rate of inflation have a smaller cost, in terms of increased cyclical unemployment rates, for the Atlantic provinces than for the other provinces, all of which have very similar sacrifice ratios.[14] It is useful, however, to recognize that the sacrifice ratio cuts both ways. That is, the relatively small sacrifice ratio for the Atlantic provinces also indicates that

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11. Maki (1977) derived similar results in his study of the effects of changes in the generosity of unemployment insurance on the duration of benefit claims. He found that the effect of changes in the replacement rate on the duration of unemployment insurance claims was very similar in the four Atlantic provinces and significantly higher in the Atlantic region than in the rest of Canada.

12. See Johnson and Kneebone (1991) for estimates and discussion of this effect.

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13. To see this, re-write (4) as
$CUR_t = (1/\alpha\tau)[-w_t + (1-\alpha)w_{t-1} + \alpha p_{t-1}]$. Thus $\partial CUR_t/\partial w_t = -(1/\alpha\tau)$.
Since we are interested in the cumulative effect on the unemployment rate, we should also note that
$CUR_{t+1} = (1/\alpha\tau)[-w_{t+1} + (1-\alpha)w_t + \alpha p_t]$ which can be simplified to
$CUR_{t+1} = (1/\alpha\tau)[-\alpha w_t + \alpha p_t]$ by noting that if this is a permanent reduction in w, then
$w_{t+1} = w_t$. If we further assume $p_t = w_t$, then we have $\partial CUR_{t+1}/\partial w_t = 0$ and the cumulative effect of the permanent change in wage inflation is simply $(1/\alpha\tau)$. Thus our measure of the sacrifice ratio assumes an immediate and complete transmission of wage to price inflation.

14. This result is consistent with the speculation offered by Howitt (1990) that contractionary monetary policy, by increasing interest rates, might have its greatest impact on durable good producing industries (most of which are located in Ontario and Quebec) and on the grain industry (located in Manitoba, Saskatchewan and Alberta) by exacerbating the debt problems of grain farmers. The national value of the sacrifice ratio implied by our estimates (approximately 2.75) is also roughly consistent with Howitt's estimate of 2.35 (assuming an Okun's Law coefficient of 2.0). In another recent study, Cozier and Wilkinson (1990) derive an estimate of about 2.0.

TABLE 3 Estimated Values of RWR, NWR, and the Sacrifice Ratio: Canadian Provinces

	RWR	NWR	Sacrifice Ratio
Newfoundland	1.199	1.250	1.499
Prince Edward Island	1.877	1.395	2.618
Nova Scotia	0.948	1.294	1.227
New Brunswick	4.840	0.943	4.566
Québec	0.751	3.846	2.890
Ontario	1.352	2.169	2.933
Manitoba	3.695	1.333	4.926
Saskatchewan	1.175	2.155	2.532
Alberta	1.106	2.747	3.067
British Columbia	1.315	1.873	2.463

expansionary monetary policies have their smallest effect on unemployment rates in that region.

Our estimates of the RWR measure show six provinces with values tightly bunched between 0.95 and 1.35. Prince Edward Island has a slightly larger value while Quebec has a slightly lower value. Only New Brunswick and Manitoba have noticeably different (larger) values. A national average value for RWR would seem to be approximately unity suggesting that, on average in Canada, a one percentage point unemployment rate shock reduces the target rate of real wage growth by one percentage point.

Our estimates of NWR are well-determined in all provinces given the low estimated standard errors for our estimators of α. Our results suggest a lower value of the NWR for the Atlantic provinces and Manitoba than for the rest of the country. This indicates that nominal wages adjust much more quickly toward the target level in the Atlantic provinces than elsewhere and this explains the smaller than average sacrifice ratio for this region. That is, the greater flexibility of nominal wages implies that the effects of contractionary monetary policies are more quickly translated from reductions in the target rate of wage change into reductions in the actual rate of wage change. The unemployment rate costs of disinflationary policies are consequently smaller than in other provinces due to this faster than average response.

Another interesting implication of finding wages to be relatively flexible in the high unemployment region of Atlantic Canada is that a fall in aggregate demand will increase the unemployment rate by less in the Atlantic region than elsewhere due to the relatively quick response of wages to the fall in demand. As a result, cyclical unemployment rate disparities between provinces will shrink. In an expansion, however, the opposite will be true as wages rise more quickly in the Atlantic provinces than elsewhere. This result is consistent with the findings of Raynaud (1987) who used vector auto-regression techniques to provide evidence suggesting a narrowing (widening) of regional unemployment rate disparities in an economic downturn (upturn). As Raynaud notes, his tech-
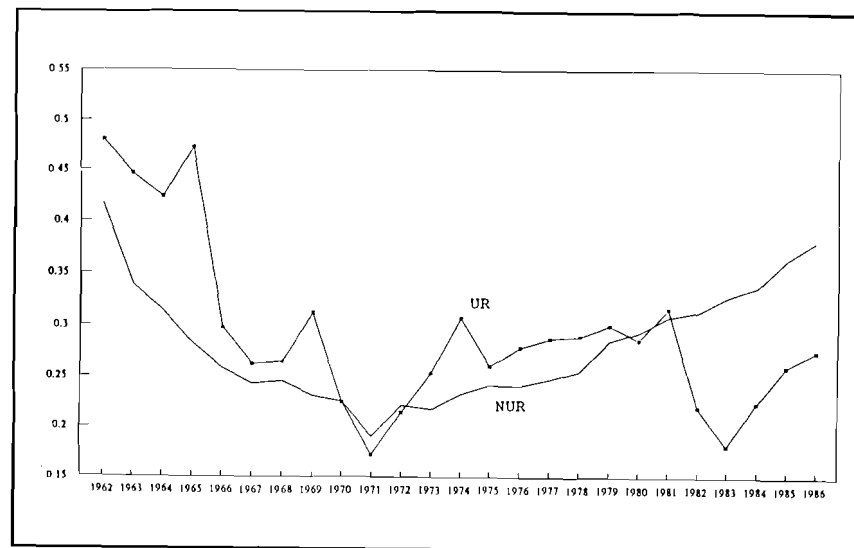

FIGURE 1 Coefficients of Variation

nique only allows a description of correlations and does not identify the reasons for these patterns. Our results identify the reason for this relationship as being the greater flexibility of wages in the Atlantic provinces.

Interest in nominal and real wage rigidity measures arise due to the fact that such rigidities worsen the impact on unemployment rates of adverse shocks. An interesting question to ask, then, is how well do provincial differences in the rigidity measures explain provincial differences in cyclical unemployment rate experience. To illustrate this point the weighted coefficients of variation (across provinces) for the observed and natural unemployment rates have been calculated and graphed in Figure 1.[15] A high coefficient of variation indicates a large amount of unemployment dispersion across provinces (high degree of unemployment rate disparity). A coefficient of variation equal to zero indicates all 10 provinces have exactly the same unemployment rate. Referring to Figure 1, it is apparent that the results described above aid us in explaining the changes in unemployment rate disparities relative to changes in

15. The weighted coefficient of variation is calculated as σ / UR_n where UR_n is the national unemployment rate and σ is the standard deviation of provincial unemployment rates (UR_i) about the national average. That is, $\sigma = [\Sigma\alpha_i(UR_i - UR_n)^2]^{.5}$ where α_i are provincial labour force weights. The coefficient of variation is used, rather than the ordinary weighted standard deviation, because we wish to measure relative disparities. If unemployment rates rise in all provinces by 10%, the absolute disparity in provincial rates will increase, and this will be reflected by a corresponding increase in the standard deviation. However, the relative disparity in rates will be unchanged, as will be the coefficient of variation.

natural rate disparities. The 1960s was a period of strong and steady economic expansion during which time our wage rigidity results would suggest a widening of cyclical unemployment rate disparities. However, as this was a period of small cyclical unemployment rates, we would expect any change in total unemployment rate disparities to be mainly explained by changes in natural rate disparities. This is the pattern suggested by Figure 1. At the other extreme is the pattern of change in disparities during the 1981-83 recession. Our wage rigidity results would suggest a narrowing of cyclical rate disparities and this influence would be strongly felt in a measure of total unemployment rate disparities because of the growing share of the cyclical in the total unemployment rate during recessions. Again, Figure 1 supports this interpretation as disparities between total unemployment rates plummeted at the same time natural rate disparities continued to climb.

To shed further light on this question, we investigated the relationship between wage rigidity and the effect of shocks on cyclical unemployment. In order to provide a measure of cyclical unemployment fluctuation, two variables were created. The first is the standard deviation of cyclical unemployment rates over the whole estimation period (σ).[16] We would expect that provinces exhibiting more wage rigidity would experience greater variation in cyclical unemployment over the quarter-century for which we have data. The second measure is the total point years of cyclical unemployment for the years 1982-85 inclusive (TPYCU). It was hypothesized that a province with high wage rigidity would suffer a deeper and more prolonged adverse employment effect as a result of the 1981-82 recession. Both of these measures were calculated for every province and the results are presented in Table 4. Interestingly, there is a strong correlation between these two measures (the correlation coefficient between TPYCU and σ is 0.9629) suggesting that there has been little change in relative degrees of wage rigidity across provinces over the 1961-86 period. The general ranking of the provinces is the same using both measures with British Columbia, Alberta, and Québec (in that order) showing the greatest degree of cyclical variability. The relationship between NWR and the TPYCU measure of cyclical unemployment variation is illustrated in Figure 2. The correlation between NWR and TPYCU is 0.53.

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16.   In this case the standard deviation, rather than the coefficient of variation is used because, for each province, the expected value of the *cyclical* unemployment rate over the period is zero. Thus, the standard deviations are directly comparable.

TABLE 4 Analysis of Provincial Differences in Cyclical Unemployment Rates

Province	Mean (62-86)	σ (62-86)	Maximum	Minimum	TPYCU (82-85)
Newfoundland	-0.032	1.92	4.1 (1985)	-2.7 (1966)	10.75
Prince Edward Island	-0.012	1.62	3.0 (1982)	-2.7 (1978)	6.46
Nova Scotia	0.019	1.05	2.4 (1982)	-1.3 (1974)	7.19
New Brunswick	-0.004	1.14	2.1 (1985)	-1.7 (1967)	7.20
Québec	-0.004	2.25	4.7 (1982)	-3.0 (1966)	16.79
Ontario	0.006	1.75	4.2 (1983)	-1.8 (1974)	13.59
Manitoba	-0.099	1.91	4.1 (1983)	-2.5 (1969)	13.83
Saskatchewan	-0.025	1.58	3.4 (1984)	-2.1 (1967)	11.05
Alberta	0.024	3.21	7.1 (1984)	-3.0 (1969)	23.68
British Columbia	-0.027	3.44	7.8 (1984)	-3.5 (1969)	26.78

Conclusion

The purpose of this paper was to answer two questions. First, we asked whether differences exist in the degree of wage rigidity across provincial labour markets in Canada. Our results suggest that such differences do indeed exist and that, in particular, the Atlantic provinces exhibit a significantly lower degree of nominal wage rigidity than the other provinces. The second question we asked was how useful are these differences in wage rigidity in explaining the different cyclical unemployment rate experiences across provinces. We found evidence to suggest that differences in nominal wage rigidity play a significant role in explaining these differences.

The recent interest in wage rigidity measures is explained by the fact that increasingly rigid wages worsen the impact of adverse shocks on unemployment rates. The cost, in terms of unemployment, of increasingly rigid wages is measured by sacrifice ratios. An important implication of our estimates is that the sacrifice ratios of the Atlantic provinces are significantly less than those in the other provinces, all of which have reasonably similar sacrifice ratios. As a result, adverse shocks, including those imposed by central bank disinflationary policies, have a much smaller negative effect in the Atlantic provinces than elsewhere. Since in recent years the Bank of Canada has pursued a vigorous anti-inflation policy of tight money our results suggest that the Atlantic provinces have suffered least from this policy while the remaining provinces have shared roughly equally in these costs.

In previous papers (Johnson and Kneebone 1989, 1991) we have remarked that in a country often concerned about unemployment rate disparities across provinces and regions, it is surprising that greater attention has not been paid to the question of what causes these disparities. In Johnson and Kneebone
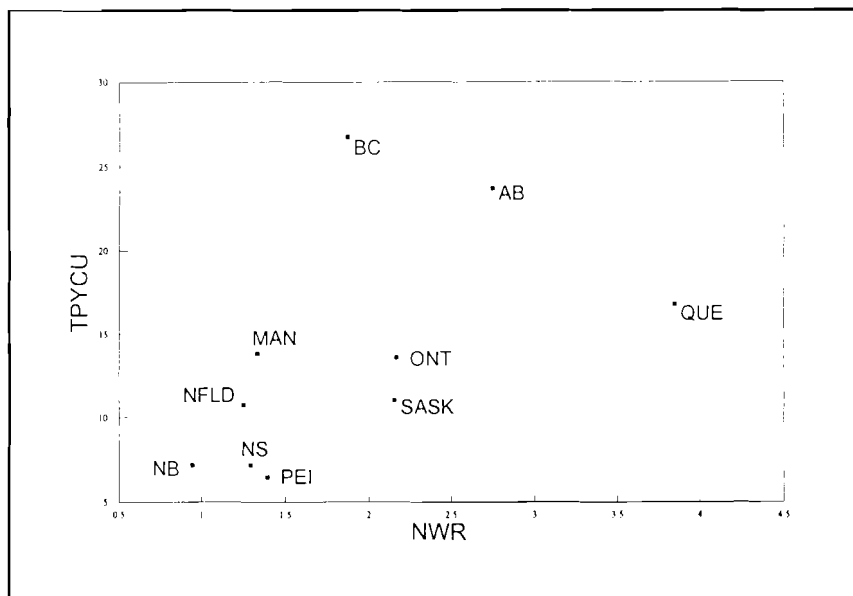
FIGURE 2 TPYCU vs. NWR

(1991), we found that much of the unemployment rate disparities across provinces was due to differences in natural unemployment rates (a result confirmed by Burns 1990). In that paper, we identified minimum wage laws and unemployment insurance policies to be important determinants of these natural unemployment rate disparities. This paper finds that much of the remaining (cyclical) unemployment rate differences arise from differences in wage rigidity. We have not endeavoured to explain the reasons for these differences in wage rigidity but this is clearly an important avenue for future research. Klau and Mittelstädt (1986) suggest that government labour market policies may be a contributing factor by affecting labour mobility and real labour cost. If so, our results suggest that changes in government labour market policies may go a long way to eliminating regional and provincial unemployment rate disparities.

Appendix: Definition of Variables

$w_{i,t}$ The percentage change in nominal average weekly gross industrial wages and salaries of all employees in province i in year t. Source: Statistics Canada 72-002.

$p_{i,t}$ The percentage change in the consumer price index in pro-

vince i in year t. Provincial consumer price indices were calculated as the population-weighted average of city consumer price indices for all cities in the province for which Statistics Canada collects data. Source: Statistics Canada 62-002 and 62-010.

$CUR_{i,t}$ Cyclical total unemployment rate in province i in year t. Calculated as actual total provincial unemployment rate (Source: Statistics Canada 71-529 and the Conference Board of Canada) minus the provincial natural unemployment rate (Source: Johnson and Kneebone (1989), Table A2).

$PRD_{i,t}$ Expected percentage change in productivity in province i in year t. Calculated as the fitted values of a linear spline regression of time on real provincial GDP (Source: Statistics Canada 13-213S) per employed worker (Source: Statistics Canada 71-001).

AIB_t Dummy variable which equals unity for the years 1976-78, zero otherwise.

$UI_{i,t}$ Percentage change in an index of unemployment insurance generosity in province i in year t. Following Fortin (1984), it is calculated as the product of three features of unemployment insurance: Coverage, replacement rate, and duration ratio. Coverage measures the proportion of the labour force covered by unemployment insurance. The replacement rate measures the ratio of after-tax unemployment insurance benefits to after-tax employment income. The duration ratio measures the ratio of the maximum number of weeks one is entitled to collect benefits to the minimum number of weeks one must work in order to be entitled to collect benefits. Thus the index of unemployment insurance generosity in province i (UIG_i) is calculated as;

$$UIG_i = (CON_i/LF_i)(AWB_i/AW_i(1-TX_i/PI_i))(MAX/MIN_i)$$

where:

CON_i = persons covered by unemployment insurance in province i (Statistics Canada 73-201).
LF_i = labour force (Statistics Canada 71-529 and Conference Board of Canada).
AWB_i = average weekly after-tax unemployment insurance benefits (Statistics Canada 73-001).
AW_i = average weekly gross wages and salaries of all employees (Statistics Canada 72-002).
TX_i = direct taxes paid by persons to all levels of govern-

ment (Statistics Canada 13-213S and 13-001).

PI_i = personal income (Statistics Canada 13-213S).

MAX = maximum number of weeks one may collect unemployment insurance benefits (Statistics Canada 73-201).

MIN_i = minimum number of weeks of employment necessary to qualify for benefits (Statistics Canada 73-201).

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