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Do Real Wages Respond Asymmetrically to Unemployment Shocks? Evidence from the U.S. and Canada

Using a set of cointegration and error correction models with asymmetric adjustment, this paper investigates aggregate labor market adjustment in the U.S. and Canada in the post 1973 period. Empirical results show real wages, productivity, and unemployment are cointegrated. Adjustment toward the long-run equilibrium seems to be linear for the U.S. and non-linear for Canada. The dynamic adjustment of real wages to unemployment and productivity shocks show markedly different responses to positive shocks than negative shocks in Canada. However, adjustment in the U.S. is mostly symmetric. This difference between the U.S. and Canadian real wage responses may provide an answer to the recent divergence of unemployment rates between the U.S. and Canada.

1. Introduction

A central issue between neoclassical and the new Keynesian economics concerns the downward flexibility of real wages. If real wages are downwardly rigid, then adverse shocks that reduce the demand for labor act to create sustained unemployment. An interesting comparison is the recent U.S. and Canadian experience concerning unemployment rates. Both nations should experience similar productivity shocks. However, a close look at unemployment rates (see Figure 1) suggests that the close association of the U.S. and Canadian unemployment rates ended around 1982. Nevertheless, Ashenfelter and Card (1986) focused on structural rigidities in the two labor markets and could not validate that real wage rigidity was a major source of the divergence. More recently, Card and Freeman (1994) attributed a good part of the divergence to the relatively market-oriented U.S.

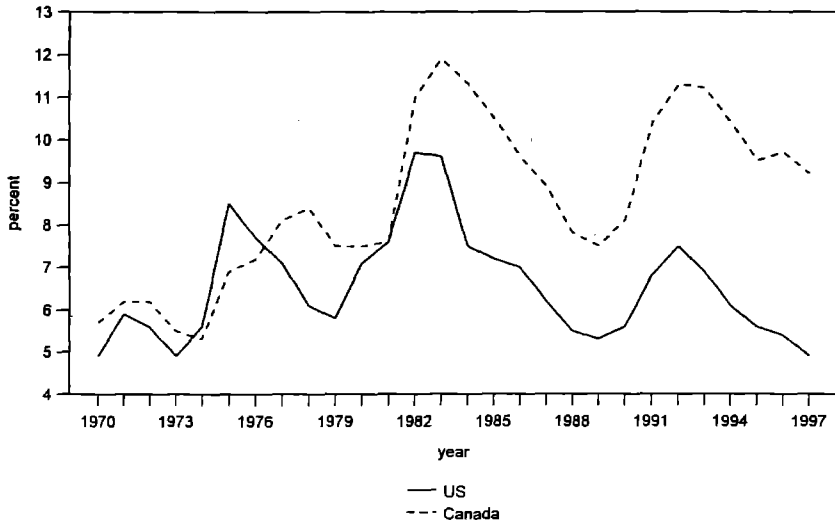


Figure 1.
Unemployment Rates in the U.S. and Canada

policies that allowed the incomes of less skilled workers to fall and poverty rates to rise, and Canadian policies that provided a safety net for the poor. Moreover, the two countries have different minimum wage laws, unemployment compensation schemes, regional/sectoral labor mobility, job security policies, and costs associated with hiring and dismissing workers.¹

The objective of this paper is to provide direct evidence on asymmetric adjustment in U.S. and Canadian labor markets using unit root and cointegration methods with non-linear adjustment. These methods are suitable for testing whether real wages behave asymmetrically in periods of high and low unemployment. If the wage adjustment is such that real wages increase more readily than they decrease, the wage structure fails to respond to unemployment in a market-clearing fashion. Enders and Granger (1998) and Enders and Siklos (2000) develop a number of unit-root and cointegration tests that do not presuppose that the variables follow a linear-symmetric adjustment mechanism. As such, the new tests should be more suitable to test asymmetry in the labor market and model the dynamic adjustment of wages.

¹For a recent discussion, see Nickell (1997). Hanes (1993) presents historical evidence linking decreases in wage flexibility to increases in worker's bargaining power starting in the late 19th century.

2. Theoretical Framework

Consider a wage bargaining framework where firms and unions bargain over wages and possibly over employment along the lines of Layard and Nickell (1986) and Nickell (1988). Firms consider the effect of wages (W) on profits (Π) when they bargain about wages. The profit function is of the form: $\Pi = \Pi(W/P, Y/N)K$, where Y is output, N is employment, K is the (predetermined) capital stock, and P is the price level. The union utility function is of a utilitarian form: $\Omega = NV + (N^S - N)V'$, where N^S is total labor supply, and V and V' are union members' utilities for the employed and the unemployed, respectively. These utilities in turn are functions of the real wage and factors that affect the wedge between consumption and product wage. The wage is set to maximize the function $[\Omega(W,N) - \Omega^0]^\alpha [\Pi(W,N) - \Pi^0]^{1-\alpha}$, where Π^0 and Ω^0 represent the status quo outcomes. The result of the Nash-bargaining solution is a general wage function of the form $W/P = W(Y/N, U, Z)$, where U is unemployment, Z is a vector of relevant exogenous variables which may include wedge elements, price of imported goods, union power, and variables that affect welfare if workers are unemployed. A compact log linear specification of labor market equilibrium is

$$\beta_0 + \beta_1 \omega_t + \beta_2 q_t + \beta_3 u_t = \mu_t, \quad (1)$$

where ω_t is the logarithm of the real wage, q_t is average labor productivity, u_t is the log unemployment rate, μ_t is a stochastic disturbance term.² The logarithmic specification for the unemployment rate represents the convex relationship between the unemployment rate and real wages. Notice that Equation (1) is not a structural equation in that it is impossible to determine the sign of the relationship between unemployment and the real wage assuming a market clearing framework. The relationship between unemployment and the real wage can be of either sign as labor supply and demand shifts may generate positive or negative correlations between the two variables, depending on the magnitude of these shifts.³

Different versions of Equation (1) have been estimated by many economists including Hall (1986, 1989), Mehra (1991), Nymoen (1992), Schoonbeek and Sterken (1995) and Chiarini and Piselli (1997). Since real wages, labor productivity and unemployment are usually found to be $I(1)$ processes,

²Similar wage equations are provided by the wage resistance model of Sargan (1964), insider-outsider models, and efficiency wage models; for structural wage equations see, Nickell (1988), and Manning (1993).

³Neftci (1984) and Sumner and Silver (1989) provide empirical evidence that unemployment and real wage correlations have been of either sign depending on the sample period.

the typical procedure is to estimate (1) and test for cointegration using the Engle and Granger (1987) or Johansen (1995) methodologies. The Engle-Granger procedure, for example, relies on the OLS estimate of ρ in the regression equation:

$$\Delta\mu_t = \rho\mu_{t-1} + \varepsilon_t, \tag{2}$$

where the estimated regression residuals from (1) are used to estimate (2). Cointegration implies $-2 < \rho < 0$, in which case μ_t in (2) is stationary with mean zero. Note that (2) implies symmetric adjustment in that $\Delta\mu_t$ equals ρ multiplied by μ_{t-1} , regardless of whether μ_{t-1} is positive or negative.⁴

Pippenger and Goering (1993) and Enders and Granger (1998) show that the standard tests for unit-roots and cointegration all have low power in the presence of misspecified dynamics. This is important since the linear relationship in Equation (2) is inappropriate if wages are sticky in the downward, but not upward, direction. For given values of q_t and u_t , suppose that ω_t is above the long-run value implied by Equation (1). Given downward wage rigidity, it will take longer to restore equilibrium than when ω_t is below its long-run value.

A formal way to introduce asymmetric adjustment is to let the deviations from the long-run equilibrium in Equation (1) behave as a Threshold Autoregressive (TAR) process. Thus, it is possible to replace (2) with

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \varepsilon_t, \tag{3}$$

where I_t is the Heaviside indicator such that

$$I_t = \begin{cases} 1 & \text{if } \mu_{t-1} \geq \tau \\ 0 & \text{if } \mu_{t-1} < \tau \end{cases} \tag{4}$$

and τ = the value of threshold.

Asymmetric adjustment is implied by different values of ρ_1 and ρ_2 ; when μ_{t-1} is above threshold, the adjustment is $\rho_1\mu_{t-1}$, and if μ_{t-1} is below the threshold, the adjustment is $\rho_2\mu_{t-1}$. A sufficient condition for stationarity

⁴Similarly, the Johansen (1995) methodology begins with a specification of the form:

$$\Delta x_t = \pi x_{t-1} + \varepsilon_t,$$

where x_t is the (3×1) vector $(\omega_t, q_t, u_t)'$, π is a (3×3) matrix, and ε_t is a (3×1) vector of normally distributed disturbances that may be contemporaneously correlated. Under the alternative hypothesis [i.e., $\text{rank}(\pi) \neq 0$] the adjustment process is symmetric around $x_t = 0$ in that for any $x_t \neq 0$, Δx_{t+1} always equals πx_t . Thus, πx_t can be viewed as an attractor such that its pull is strictly proportional to $\|x_t\|$.

of $\{\mu_t\}$ is: $-2 < (\rho_1, \rho_2) < 0$. Moreover, if the $\{\mu_t\}$ sequence is stationary, the least squares estimates of ρ_1 and ρ_2 have an asymptotic multivariate normal distribution (Tong 1983). Thus, suppose that the null hypothesis $\rho_1 = \rho_2 = 0$ is rejected. It is then possible to test for symmetric adjustment (i.e., $\rho_1 = \rho_2$) using a standard F -test. Since adjustment is symmetric if $\rho_1 = \rho_2$, the Engle-Granger test for cointegration is a special case of Enders-Granger (1998). If the errors in Equation (3) are serially correlated, it is possible to use an augmented TAR model for the residuals. Thus, Equation (3) is replaced by,⁵

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{t=0}^p \beta_t\Delta\mu_{t-1} + \varepsilon_t. \quad (5)$$

The critical values needed to test the null hypothesis for the hypothesis $\rho_1 = \rho_2 = 0$ depend on the number of variables used in the cointegrating vector. Enders and Siklos (2000) report critical values using cointegrating vectors containing only two variables. Since (1) contains three variables, we first needed to calculate the critical values for the three variable case. Our methodology is described in Appendix 1 and the critical values are reported in Table 1. To use the statistics in Table 1, perform the following 4 steps:

Step 1: Regress one of the variables on a constant and the other two variables and save the residuals in the sequence $\{\mu_t\}$. This equation is the estimated cointegrating vector.

Step 2: To find the consistent estimate of the threshold, order the $\{\mu_{t-1}\}$ sequence from smallest to largest. Although any value of $\{\mu_{t-1}\}$ is a potential threshold, for estimation purposes, it is necessary to have a reasonable number of observations in each of the two regimes. Hence, consider as a potential threshold, each μ_{t-1} between the lowest 15% and the highest 85% values of the series. Estimate regressions in the form of (5) using each potential value of μ_{t-1} as a threshold. The value resulting in the lowest residual sum of squares is the estimate of the threshold $\hat{\tau}$. Using $\hat{\tau}$ as the threshold, compare the F -statistic for the null hypothesis $\rho_1 = \rho_2 = 0$ with the appropriate critical value shown in Table 1.

Step 3: If the alternative hypothesis is accepted, it is possible to test for symmetric versus asymmetric adjustment since ρ_1 and ρ_2 converge to multivariate normal distributions. As such, the restriction that adjustment is symmetric (i.e., the null hypothesis: $\rho_1 = \rho_2$) can be tested using the usual F -statistic. Nevertheless, Enders and Falk (1998) and Hansen (1997) show

⁵Enders and Granger (1998) and Enders and Siklos (2000) suggest that lag length should be determined by a model selection criterion such as the AIC or BIC.

TABLE 1. *Critical Values for the Φ Statistic*

	No Lagged Changes			1 Lagged Change			4 Lagged Changes			8 Lagged Changes		
	10%	5%	1%	10%	5%	1%	10%	5%	1%	10%	5%	1%
T = 50	7.472	8.830	11.888	7.616	8.976	11.997	7.241	8.531	11.615	NA	NA	NA
T = 100	7.269	8.454	11.113	7.261	8.477	11.147	7.143	8.267	10.841	6.767	7.861	10.212
T = 250	7.144	8.262	10.567	7.225	8.343	10.721	7.162	8.281	10.566	7.044	8.146	10.451

NOTES: NA: We do not provide the critical values for the model with 8 lags using only 50 observations. On average, there are 21 observations available to estimate each value of p . As such, we cannot recommend estimating the TAR in these circumstances.

TABLE 2. *Cointegration Statistics for the Wage Model*

Country	lags	$\rho_1 \mu^+$	$\rho_2 \mu^-$	Φ	$\rho_1 = \rho_2$	Estimated long-run relation
U.S.	8	-0.249 (-3.687)	-0.333 (-3.580)	9.581*	0.829	$\omega = -1.507 + 0.303 q - 0.011 u + \mu$
Canada	4	-0.061 (-0.857)	-0.344 (-4.197)	9.079*	6.945*	$\omega = 0.264 + 0.279 q + 0.091 u + \mu$

NOTES: 1. Quarterly data from 1973:i-1998:ii

2. (*) significant at 10% or below.

3. t -statistics are given in parentheses.

that small sample properties of the OLS estimates of ρ_1 and ρ_2 have inflated standard errors and the convergence properties of the OLS estimates can be poor. Hence, inference concerning the individual values of ρ_1 and ρ_2 is problematic.

Step 4: Diagnostic checking of the residuals should be undertaken to ascertain whether the $\{\varepsilon_t\}$ series can reasonably be characterized by a white-noise process. If the residuals are correlated, return to Step 2 and re-estimate the model augmented with lagged changes of $\{\mu_t\}$. Lag lengths can be determined by an analysis of the regression residuals and/or using a number of widely used model selection criteria such as the AIC or BIC.

3. Empirical Results

The wage Equation in (1) is estimated using quarterly data from 1973:i–1998:ii for the U.S. and Canada. Data descriptions and sources are given in Appendix 2. We use a three-quarter moving average measure of labor productivity, as this rules out measurement errors. Table 2 presents test results for μ_t based on Equation (1) assuming threshold adjustment. The table reports values of the adjustment coefficients ρ_1 and ρ_2 , their t -values, and the Φ -statistic for the null hypothesis of a unit root in μ_t (no cointegration) against the alternative of cointegration with asymmetric adjustment. The F -test for symmetric adjustment $\rho_1 = \rho_2$ and the underlying long-run relations are also presented in the table. The lag length is selected such that the Akaike Information Criterion (AIC) is minimized. As such AIC selects 4 lags for Canada and 8 lags for the U.S.⁶

The estimated Φ statistic for the null hypothesis $\rho_1 = \rho_2 = 0$ is 9.581 for the U.S. The critical value reported in Table 1 at the 5% significance level with 100 observations and 8 lags is 7.861. As such, we reject the null hypothesis of a unit root in favor of stationarity with asymmetric adjustment in the U.S. As for Canada, the estimated Φ statistic for the null hypothesis $\rho_1 = \rho_2 = 0$ is 9.079. Given that the critical value of the Φ at the 5% significance level is 8.267, we also find that there is a stationary long-run equilibrium in the Canadian labor market.

The F -statistic for symmetric adjustment ($\rho_1 = \rho_2$) given in Table 2 strongly rejects symmetric adjustment for Canada at less than a 1% significance level. The statistic for the U.S. tends toward symmetric adjustment. Notice that in both cases negative deviations from the long-run equilibrium are eliminated much faster than positive deviations. Specifically in the U.S. 24.9% of a positive deviation is eliminated within a quarter while 33.3% of a negative deviation is eliminated within the same time frame. The point

⁶The estimated threshold is -0.01024 for the U.S. and 0.01739 for Canada.

estimate of ρ_1 in Canada indicates almost three times faster adjustment for negative deviations than for positive deviations.

It is known that coefficients of cointegration relations cannot be interpreted as elasticities (see Lutkepohl 1994); this is because the *ceteris paribus* assumption may not be meaningful. The error correction specification and impulse response functions based on this specification can be more informative.

4. The Dynamic Adjustment of Wages and Unemployment: Canada vs. the U.S.

Canada and the U.S. form a large trading bloc and share common institutions and similar standards of living. Empirical results indicate asymmetric adjustment in the Canadian labor market, while adjustment in the U.S. seems to be symmetric. In this section, we examine the dynamic adjustment of wages and unemployment using an error-correction model.

Using the long-run relations reported in Table 2, the fitted error-correction equations for Canada assuming threshold adjustment (with *t*-statistics in parentheses) are

$$\Delta q_t = A_{11}(L)\Delta q_{t-1} + A_{12}(L)\Delta \omega_{t-1} + A_{13}(L)\Delta u_{t-1} - 0.0471 z_{plus,t-1} - 0.0441 z_{minus,t-1}; \quad (6)$$

(-1.582) (-1.123)

$$\Delta \omega_t = A_{21}(L)\Delta q_{t-1} + A_{22}(L)\Delta \omega_{t-1} + A_{23}(L)\Delta u_{t-1} - 0.0724 z_{plus,t-1} - 0.2893 z_{minus,t-1}; \quad (7)$$

(-0.805) (-2.445)

$$\Delta u_t = A_{31}(L)\Delta q_{t-1} + A_{32}(L)\Delta \omega_{t-1} + A_{33}(L)\Delta u_{t-1} + 0.0192 z_{plus,t-1} + 0.1443 z_{minus,t-1}; \quad (8)$$

(0.484) (2.766)

where $z_{plus_t} = I_t(\omega_t - 0.279q_t - 0.091u_t - 0.264)$,
 $z_{minus_t} = (1 - I_t)(\omega_t - 0.279q_t - 0.091u_t - 0.264)$,
 $I_t =$ threshold Heaviside indicator function,

$A_{ij}(L)$ is a polynomial in the lag operator L , and the lag length is selected using the multivariate version of *AIC* which selected 8 lags.

It is interesting to note that the adjustment coefficients of z_{plus} and z_{minus} are markedly different for real wages and unemployment. Specifically, Equation (7) indicates that real wage growth is faster when there is a negative deviation from long-run equilibrium than when there is a positive deviation. The point estimates imply that the real wage in Canada adjusts

Equations (12)–(14) indicate that the U.S. productivity and real wages close the gap with adjustment coefficients that are significant while unemployment is not responsive to deviations from the long-run equilibrium. In that sense the U.S. unemployment rate is “exogenous” with respect to the long-run equilibrium. Adjustments toward long-run equilibrium are accomplished via changes in labor productivity and real wages in the U.S.

Due to the non-linearity of the moving-average representation of a TAR process, the impulse responses do not have the usual interpretation. As discussed in Koop, Pesaran and Potter (1996), the impulse responses of a non-linear system depend on the initial conditions and the nature of the shocks. We assume that the system is in long-run equilibrium and consider the impulse responses from 1-standard deviation shocks obtained using a Choleski decomposition with an ordering $q_t \rightarrow \omega_t \rightarrow u_t$ (i.e., labor productivity is prior to the real wage which is prior to unemployment). The response of the real wage to positive and negative shocks for Canada is given in Figure 2.⁷

As shown in the upper panel of the figure, following a unit *positive* unemployment shock, the real wage initially declines up to five quarters and then starts to rise, reverting itself after eight quarters. Overall a positive unemployment shock seems to have a positive long-run effect on real wages in Canada. However, after a *negative* unemployment shock, the real wage increases up to seven quarters and starts to decline slowly with no long-run effect. While the short-run response of the real wage to positive and negative unemployment shocks is consistent with market clearing, it is evident that real wage adjustment is not linear. The real wage increases more readily in response to “excess demand” in the labor market, while in response to “excess supply,” the response is much smaller and reverts in sign in the long run.

The middle panel of Figure 2 gives the response of the real wage to productivity shocks. The real wage increases in response to a positive productivity shock with some oscillations. The response to a negative productivity shock is mostly positive except for the negative initial impact effect. The bottom panel, which gives the response of the real wage to own shocks, tells a similar story. Real wages increase in response to a positive shock but they increase after an initial decline in response to a negative shock. All of the responses are qualitatively consistent with a non-linear behavior of real wages in Canada. The real wage responses indicate upward as well as downward change in the very short run, but the responses are positive in the long run.

The response of unemployment in Canada is examined in Figure 3.

⁷The results are not sensitive to the ordering in the Choleski decomposition as the maximum correlation coefficient between any pair of reduced form innovations is 16% in Canada.

Do Real Wages Respond?

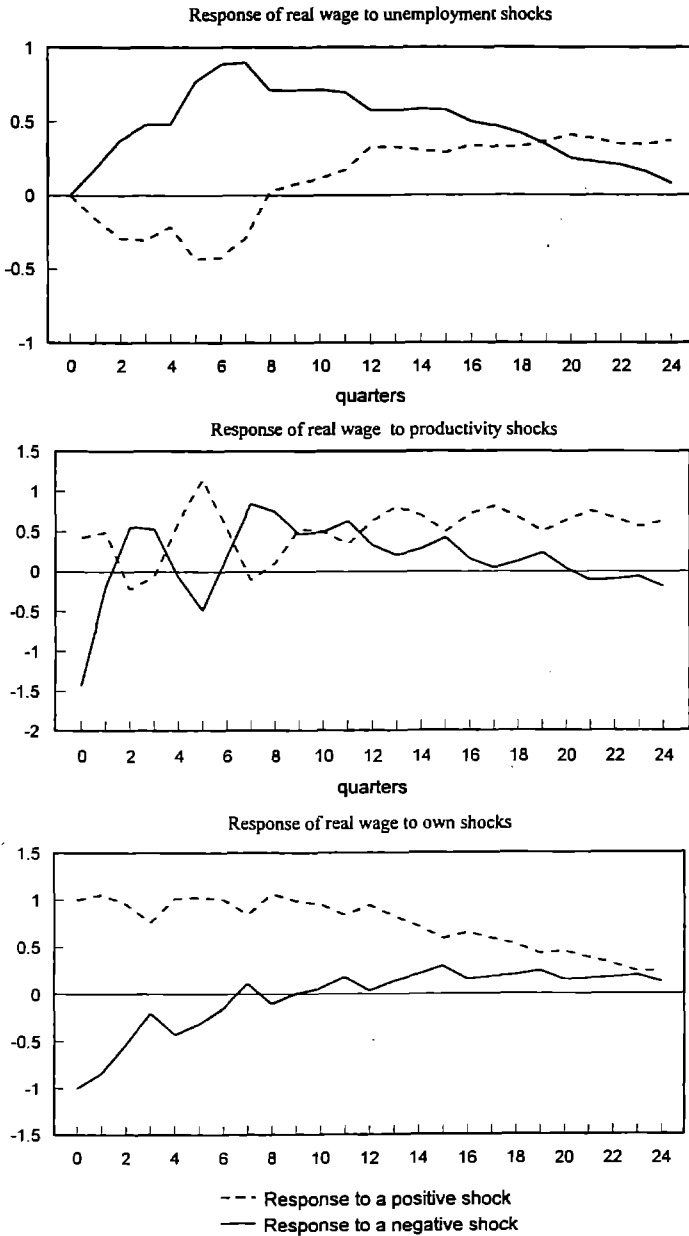


Figure 2.
Real Wage Responses: Canada

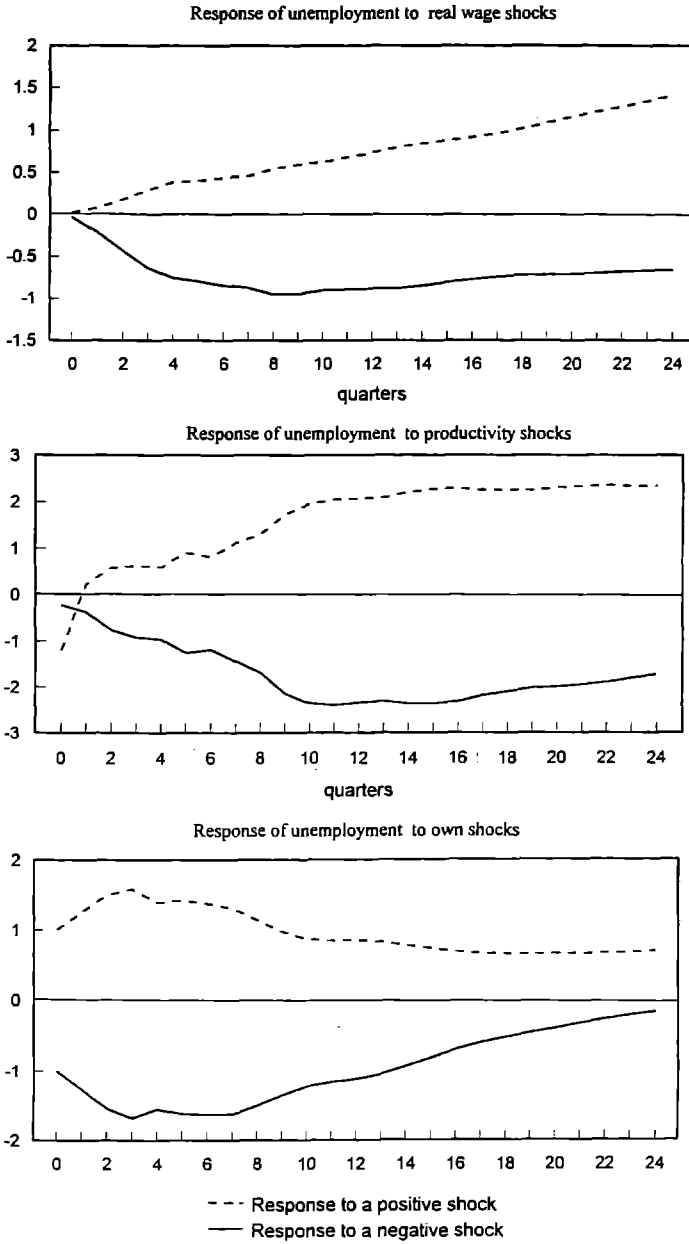


Figure 3.
Unemployment Responses: Canada

The upper panel of the figure gives the response of unemployment to real wage shocks. Unemployment rises in response to positive real wage shocks and declines in response to negative real wage shocks. Figure 3 reveals that some unemployment responses are asymmetric as well. For example, after a positive real wage shock, the unemployment rate rises steadily, whereas in response to a negative real wage shock, it declines and starts to level off. The responses of unemployment to productivity and own shocks are largely symmetric. Unemployment rises in response to a positive productivity shock after an initial negative impact effect while it falls in response to a negative productivity shock.

To contrast with U.S. results consider impulse response functions based on the system in (11)–(13).⁸ Figure 4 gives the response of U.S. real wage to positive and negative shocks. It is evident from the figure that the U.S. real wage responses are mostly symmetric. The upper panel in the figure indicates that real wage initially increases in response to positive unemployment shocks, then declines with no evident long-run effect. The response of the real wage to productivity and own shocks is also symmetric. However, in contrast to Canada, the real wage in the U.S. seems flexible in the upward as well as downward direction. Moreover, the effects of positive as well as negative shocks to the real wage in the U.S. show some evidence of being eliminated in the long run.

The response of unemployment in the U.S. is given in Figure 5. The upper panel gives the response of unemployment to real wage shocks. Unemployment in the U.S. rises in response to positive real wage shocks. However, the response to positive real wage shocks seems greater than the response to negative shocks. The middle and lower panels of the figure indicate mostly symmetric responses of unemployment to productivity and own shocks.

Overall, the real wage—unemployment relationship is mostly linear in the U.S. but not in Canada. Moreover in Canada, the most dramatic difference between responses to positive and negative shocks occurs for real wages. Another major difference between the U.S. and Canada is that both positive and negative shocks seem to have positive effects on the real wage in Canada. Although the real wage seems to decline on impact in the short run in Canada (particularly after negative productivity and real wage shocks), the reversion occurs almost immediately and the real wage response has an upward trend thereafter. This suggests limited downward real wage flexibility in Canada.

⁸Since statistical tests indicate a linear adjustment for the U.S., the use of (9)–(11) instead of (12)–(14) is for illustrative purposes only.

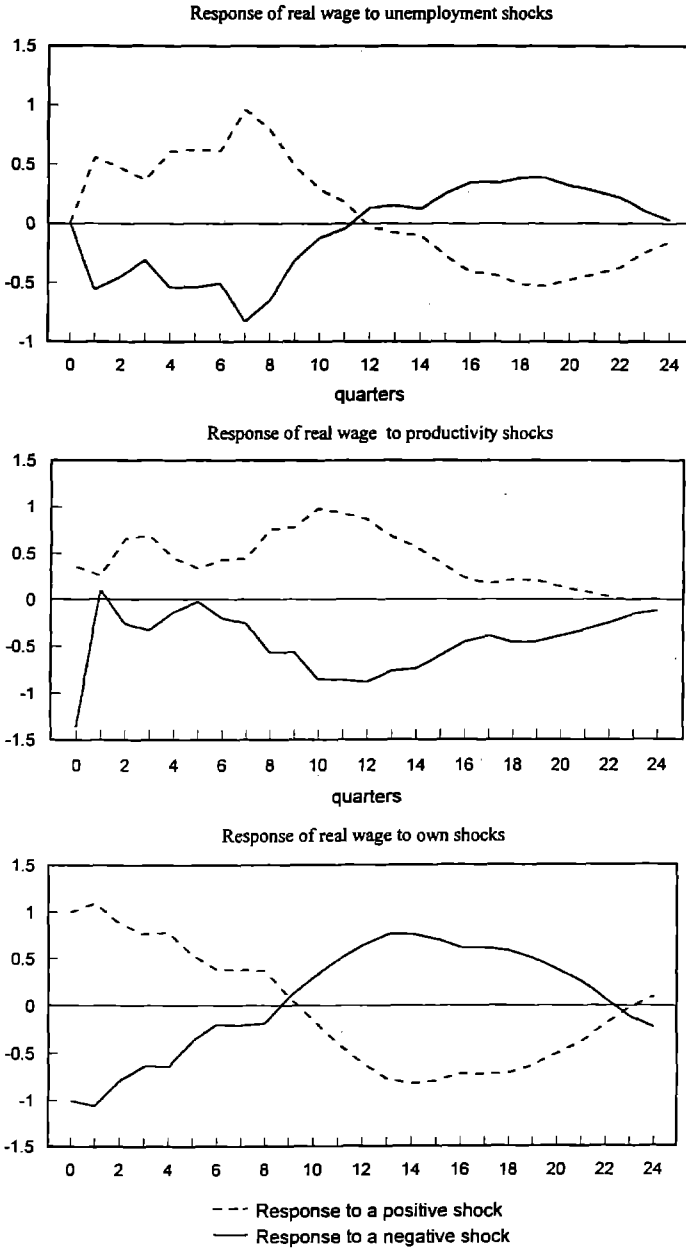


Figure 4.
Real Wage Responses: U.S.

Do Real Wages Respond?

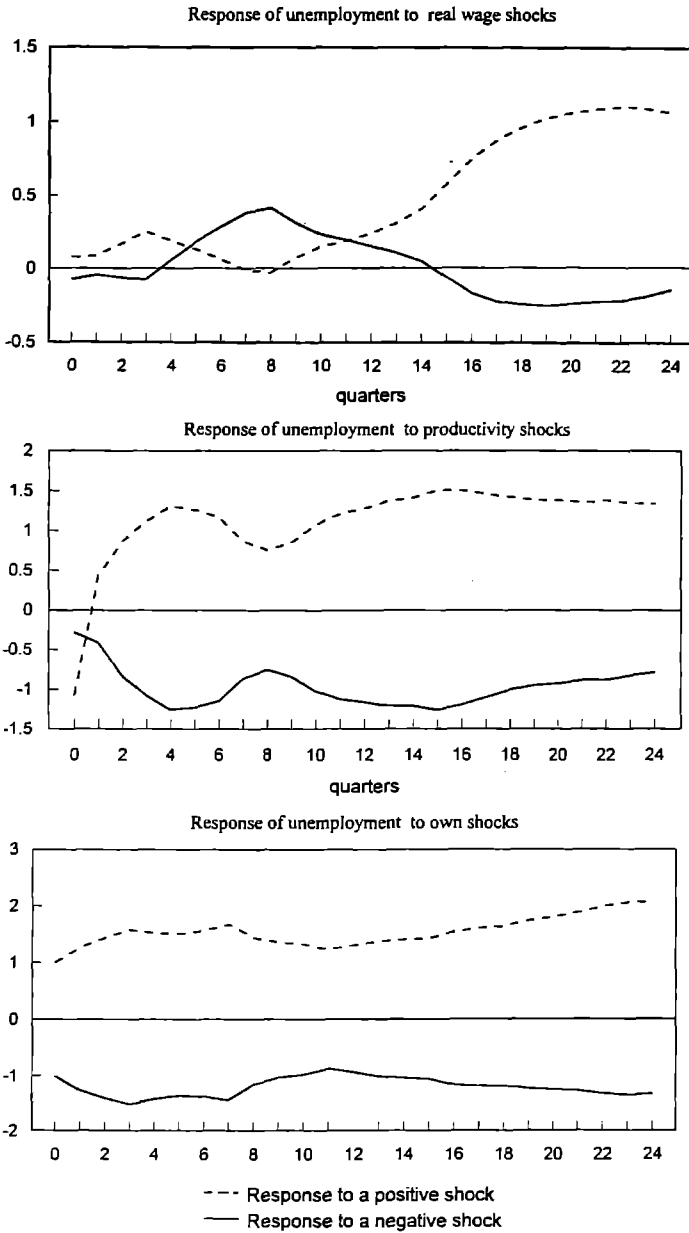


Figure 5.
Unemployment Responses: U.S.

5. Aggregate Labor Market Characteristics: U.S. vs. Canada

To better interpret the results, consider aggregate labor market characteristics presented in Figure 6, which presents annual labor force, employment, and unemployment indicators for both countries broken down by sex since 1970. The data are from the *U.S. Bureau of Labor Statistics*, and for comparison purposes, Canadian figures are chosen such that they approximate U.S. definitions.

Aggregate unemployment figures indicate that there is a divergence of unemployment rates in the two countries starting in 1982, with Canada having substantially higher unemployment rates thereafter. However, labor force participation rates and employment population ratios exhibit no major divergence in the two countries. Thus, Figure 6 indicates that the recent divergence in unemployment rates between the U.S. and Canada does not appear to be driven by a surge in labor supply in Canada.

To examine the age distribution of unemployment in the U.S. and Canada, consider Table 3. It is evident that the unemployment gap is more evident for working ages 25 and over than for young workers. Specifically, unemployment rates for teenagers appear similar in the two countries and are three to four times the unemployment rates of adults. However, unemployment rates for 25 years and over age group diverge substantially in the two countries.

The recent unemployment divergence poses a challenge. As a possible explanation, the social safety nets in the two countries attract the most attention. Canada's unemployment insurance system offers longer duration of benefits than comparable U.S. programs with less restrictive eligibility requirements. The availability of benefits for maternity leaves, sickness, and training in Canada may be contributing factors as well. While one-third of the unemployed collect unemployment benefits in the U.S. more than 90% of those unemployed in Canada collect benefits (Card and Freeman 1994). There is also a decline in unionization in the U.S. relative to Canada, which has been cited as a contributing factor to wage inequality and labor market flexibility in the U.S. (Lemieux 1993; Card and Freeman 1994). These differences notwithstanding, benefit patterns in the two countries have existed in the last 30 years (Ashenfelter and Card 1986) and may not be sufficient to explain the *recent* divergence in unemployment rates in the two countries. In that regard, evidence presented above regarding asymmetric adjustment in Canada and symmetric adjustment in the U.S. labor market may provide an important clue.

6. Conclusions

Using a set of cointegration and error correction methods that do not assume a linear adjustment, this paper investigates labor market adjustment

Do Real Wages Respond?

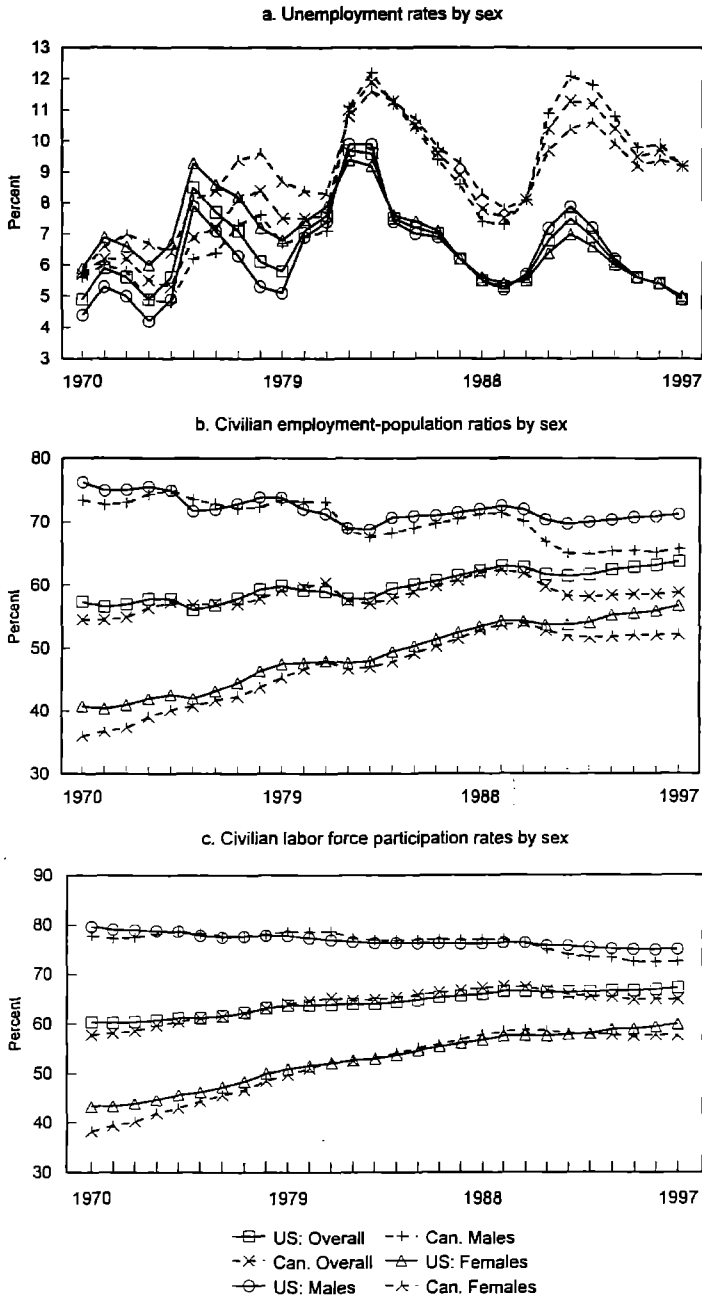


Figure 6.
Selected Labor Force, Employment, and Unemployment Indicators, 1970-97

TABLE 3. *Civilian Unemployment Rates by Age, 1993–1997*

Age Group	1993		1994		1995		1996		1997	
	US	CA	US	CA	US	CA	US	CA	US	CA
All working ages	6.9	11.2	6.1	10.4	5.6	9.5	5.4	9.7	4.9	9.2
Under 25 years	13.4	17.7	12.5	16.5	12.1	15.6	12.0	16.1	11.3	16.7
Teenagers*	19.0	20.0	17.6	18.9	17.3	18.5	16.7	20.1	16.0	21.8
20–24 years	10.5	16.2	9.7	15.0	9.1	13.7	9.3	13.6	8.5	13.6
25 years and over	5.6	9.9	4.8	9.2	4.3	8.3	4.2	8.5	3.8	7.8

NOTES: (*) 16- to 19-year-olds in the United States; 15- to 19-year-olds in Canada
Source: U.S. Bureau of Labor Statistics, International Labor Statistics Division.

in the U.S. and Canada in the post 1973 period. Aggregate data show that real wages, productivity, and unemployment are cointegrated. Adjustment toward the long-run equilibrium seems to be linear for the U.S. and asymmetric for Canada. The paper then contrasts the Canadian experience with the U.S. The dynamic adjustment of wages to unemployment and productivity shocks show markedly different responses to positive shocks from negative shocks in Canada. Economic shocks that require real wage decreases in Canada have short-lived decreases followed by real wage increases. However, adjustment in the U.S. is mostly symmetric. This difference between the U.S. and Canada may provide a clue to the recent divergence of unemployment rates between the two countries.

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Appendix 1.

Critical Values of the Cointegration Test

To develop critical values that can be used to test for cointegration, we generated 50,000 random-walk processes of the following form:

$$x_{it} = x_{it-1} + v_{it}, \quad i = 1, \dots, 3, \quad t = 1, \dots, T. \quad (A1)$$

For $T = 50, 100,$ and 250 , three sets of T normally distributed and uncorrelated pseudo-random numbers with standard deviation equal to unity were drawn to represent the $\{v_{it}\}$ sequences. Randomizing the initial values of $\{x_{it}\}$, the next T values of each were generated using (A1). For each of the 50,000 series, the TAR model given by (1), (4) and (5) was estimated. Since the value of the threshold τ is typically unknown, for each of the 50,000 replications, we used Chan's (1993) method for obtaining the consistent estimate of the threshold. For each estimated equation, we estimated ρ_1 and

ρ_2 and recorded the F -statistic for the joint hypothesis $\rho_1 = \rho_2 = 0$. This F -statistic—called the Φ -statistic—is reported in Table 1 for various values of sample sizes (T) and lag lengths p . For example, for $T = 100$, Table 1 shows that the Φ -statistic for the null hypothesis $\rho_1 = \rho_2 = 0$ exceeded 8.477 in approximately 5% of the 50,000 trials using a model augmented with 1 lagged change in $\{\mu_t\}$.

Appendix 2

Data and Sources

ω : log (Nominal Wage/CPI)

q : In Canada, three-year moving average of log output per worker (Output/Employment). In the U.S. q is output per hour in the nonfarm business sector (Fed. Res. Bank of St Louis Online FRED Database: OPHNFB)

u : log of total civilian unemployment rate, seasonally adjusted (*OECD Main Economic Indicators* various issues)

Employment: Total civilian employment index, seasonally adjusted (*OECD Quarterly Labor Force Statistics*, and *Main Economic Indicators* various issues)

Output Index: Real GDP (*International Financial Statistics*, IFS CD-ROM); updated using *OECD Main Economic Indicators*, various issues.

CPI: (*IFS* CD-ROM; updated using *OECD Main Economic Indicators*)

Nominal wage series refer to the following:

U.S.: Hourly compensation in the nonfarm business sector, seasonally adjusted (Fed. Res. Bank of St Louis Online FRED Database: COMPNFB)

Canada: Hourly earnings in manufacturing (*IFS* CD-ROM: 65ey)