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## Abstract

This paper uses Tobit models and data for union contracts to examine the extent of downward nominal-wage rigidity in Canada. To be consistent with important stylized facts, the models allow the variance of the notional wage-change distribution to be time-varying and test for menu-cost effects.

The empirical results confirm the importance of using a general specification with a time-changing variance and menu-cost effects. The variance of the notional distribution fell as inflation trended downward over the sample period, and there is evidence that menu-cost effects cause some contracts to have wage freezes rather than small wage increases. Each of these features reduces the estimated effect of rigidity on wage growth. The estimated net effect of downward rigidity and menu costs in the 1990s is approximately 0.4 percentage points for the average wage change in the first year of contracts, and less than 0.1 percentage point for the average annual change over the lifetime of contracts. On balance, the evidence suggests that the long-run trade-off between inflation and the unemployment rate is close to vertical at inflation rates of 2 per cent or more if productivity growth is near the average in recent decades.

*JEL classification: E24, E52, E61*
*Bank classification: Labour markets; Inflation targets*

## Résumé

Les auteurs utilisent les chiffres des accords salariaux et des modèles tobit pour étudier le degré de rigidité à la baisse des salaires nominaux au Canada. S’inspirant des principaux faits stylisés, ils permettent à la variance de la distribution théorique des variations salariales de fluctuer dans le temps et font intervenir dans leurs modèles les coûts d’étiquetage afin d’en déterminer l’influence.

Les résultats empiriques confirment qu’il est important d’utiliser une formulation générale qui admette une variance dynamique et tienne compte de l’effet possible des coûts d’étiquetage. La variance de la distribution théorique a diminué durant la période d’estimation alors que l’inflation accusait un recul, et il semble bien que les coûts d’étiquetage expliquent la présence de gels au lieu de faibles hausses de la rémunération dans certains cas. Chacun de ces facteurs réduit l’incidence estimative de la rigidité sur la croissance des salaires. Les auteurs estiment l’effet net de la rigidité à la baisse et des coûts d’étiquetage durant les années 1990 à environ 0,4 point de pourcentage pour ce qui est de la variation salariale moyenne enregistrée durant la première année de l’accord et à moins de 0,1 point dans le cas de la variation annuelle moyenne calculée sur la durée totale de l’accord. Les résultats donnent à penser que la courbe décrivant l’arbitrage à long terme entre l’inflation et le taux de chômage est quasi verticale lorsque le taux d’inflation se situe à 2 % ou plus, pour autant que la croissance de la productivité se maintienne autour de la moyenne observée ces dernières décennies.

*Classification JEL : E24, E52, E61*
*Classification de la Banque : Marchés du travail; Cibles en matière d’inflation*

# 1. Introduction

The recent trend toward low and stable inflation in many countries has focused attention on the factors determining the optimal inflation rate for an economy. This subject is a pressing issue for inflation-targeting countries who must compare the costs and benefits of lower inflation when choosing the appropriate level for their target.

One of the most common arguments for a positive inflation target is based on the premise that workers strongly resist cuts to their nominal-wage rates.[1] If nominal wages are downwardly rigid, and policy-makers pursue price stability, real wages cannot adjust downward following a negative shock to labour demand. Thus, it is argued that the combination of downward nominal-wage rigidity and price stability (or very low inflation) will magnify the employment losses from the negative shock. In contrast, nominal-wage floors will not constrain the adjustment process at some higher rate of inflation, as the decrease in real wage can be achieved with nominal wages rising less rapidly than prices. Accordingly, proponents of this hypothesis conclude that keeping inflation below some critical level will cause a permanent increase in unemployment, so policy-makers should target some positive inflation rate to facilitate real wage adjustments and avoid the employment costs of binding nominal-wage floors.

A growing literature has examined the extent of downward nominal-wage rigidity and its employment effects in Canada. Some of these studies, including Fortin and Dumont (2000) and Farès and Lemieux (2001), use a measure of aggregate wage growth to test the prediction that downward rigidity would cause the Phillips curve to become flatter at low inflation. Other studies have tested for rigidity using micro data such as individual union contracts. Recent contributions to the micro literature in Canada include Simpson, Cameron, and Hum (1998), Farès and Hogan (2000), and Crawford (2001).[2]

Simpson, Cameron, and Hum estimate the amount of downward nominal-wage rigidity by applying a Tobit model to data for union wage settlements in Canada. They conclude that resistance to pay cuts increased the average wage change in the first year of contracts by almost 0.7 percentage points over the 1993–95 period, and that this rigidity raised the unemployment rate by 2 percentage points. Farès and Hogan (2000) also use the Tobit methodology to study nominal-wage rigidity in Canada. In sharp contrast to the findings of Simpson, Cameron, and Hum, their results suggest that rigidity had no net effect on aggregate wage growth. One difference between

1. See Akerlof, Dickens, and Perry (1996) for a description of the hypothesis.
2. U.S. studies by Card and Hyslop (1997), Kahn (1997), and McLaughlin (1999) use a variety of micro techniques. Yates (1998) discusses international evidence on downward nominal rigidity.

these two studies is the variable used to proxy inflation expectations in the wage equation. Thus, a major focus in the first part of this paper is to examine the sensitivity of results to alternative assumptions about the formation of inflation expectations.

The second objective of this paper is to extend the Tobit literature in a more fundamental direction by modifying the standard model to include important features of the wage-change distribution. Previous studies using the Tobit model have used specifications that attribute all wage freezes to cases in which workers receive a wage change of zero rather than a wage cut (i.e., all freezes reflect downward nominal rigidity). An examination of the distribution of wage settlements provides evidence that this assumption is too restrictive: there are very few contracts with small wage increases *or* small wage decreases, which suggests that some freezes are caused by symmetric menu-cost effects rather than asymmetric downward rigidity.[3] If menu-cost effects exist, the traditional Tobit model—which constrains the censoring threshold to occur at zero—will overstate the effect of rigidity on wage growth.

Another feature of the distribution of wage settlements is a decrease in variance as the level of inflation trended downward from the late 1970s to the 1990s. One interpretation is that this decrease in variance simply reflects a thinning of the density in the left tail of the distribution owing to downward rigidity, rather than a change in the distribution that would be observed in the absence of rigidities (defined as the "notional" distribution). However, Crawford (2001) reports evidence that downward rigidity is not the only reason for the decrease in variance. While the percentage of contracts lying in the left tail did fall significantly as inflation trended downward, a very similar decline occurred on the *right side* of the distribution. Since similar movements occurred on both sides of the distribution, the evidence suggests that much of the decline in the observed variance can be attributed to a decrease in the variance of the notional distribution. Given the positive historical relationship between inflation and the variance of the wage-change distribution, constraining the notional variance to be constant is likely to result in an overstatement of rigidity in the low-inflation years.

In summary, these stylized facts suggest that empirical models of rigidity should: (i) allow the variance of the notional distribution to be time-varying, and (ii) test for menu-cost effects. These characteristics are important features of the models developed in this paper. Section 2 presents the characteristics of the standard Tobit model and summarizes empirical results from the two previous studies. Section 3 examines whether estimates of rigidity are sensitive to the way

3. If the level of the existing wage is still regarded as broadly appropriate given current conditions, the firm and workers may accept a wage freeze to avoid the costs of further negotiations over the size of a small wage change. See Crawford (2001) for a detailed discussion of the stylized facts from the distribution of wage settlements in Canada.

inflation expectations are modelled and the inclusion of a time-changing variance. To facilitate comparison with the findings of Simpson, Cameron, and Hum, the analysis in section 3 is based on a sample period ending in 1995. Section 4 extends the sample period to the end of 1999 and the model is made less restrictive by allowing for the possibility of menu-cost effects. Results from the extended Tobit model confirm the importance of using a general specification with a time-changing variance and menu-cost effects. Section 5 uses parameter estimates from Tobit models to study how downward rigidity and menu-cost effects might affect the shape of the long-run Phillips curve. Section 6 concludes, comparing the Tobit results with those from other Canadian studies of wage rigidity.

# 2. Tobit Models of Wage Growth

## 2.1 The standard model

The Tobit model's key feature is that it can be applied to markets in which censoring prevents the dependent variable from taking values below some threshold level (such as zero). This section begins with an outline of the standard Tobit model used in previous studies of nominal-wage rigidity.

Notional-wage growth in contract i, defined as the wage change that would have occurred in the absence of downward rigidity and menu-cost effects, is a function of a set of explanatory variables $X$ and a random variable $\varepsilon_{it}^n$ that varies across firms:

$$\Delta w_{it}^n = \beta X_t + \varepsilon_{it}^n . \tag{1}$$

Actual wage growth is equal to the notional level defined by equation (1) if there are pressures for a wage increase. If there are pressures for a wage cut, with $\Delta w_i^n < 0$, the wage cut is censored because of downward nominal rigidity and the contract provides a wage change of zero. Thus, in the standard Tobit model, wage growth at the micro level is

$$\Delta w_{it} = \begin{cases} \beta X_t + \varepsilon_{it}^n & \text{if} \quad \beta X_t + \varepsilon_{it}^n > 0 \\ 0 & \text{otherwise} \end{cases} \tag{2}$$

where $\Delta w_{it}$ is actual wage growth in contract i and the notional random variable $\varepsilon^n$ is normally distributed across firms with a zero mean and a constant variance $\sigma_n^2$. The parameters to be estimated are $\beta$ and $\sigma_n^2$. This specification of the model does not allow menu-cost effects because the threshold for censoring is zero in equation (2).[4]

4. Maddala (1983) describes the form of the likelihood function. Since the model assumes that there are no wage decreases, all contracts with pay cuts are excluded from our estimation.

### *2.1.1 Quantifying rigidity*

Given the assumption that all contracts with pressures for nominal-wage cuts receive a wage freeze, the mean wage growth across all contracts $E(\Delta w_i)$ is equal to the probability that wage growth is positive multiplied by the mean wage change in contracts with $\Delta w_i > 0$. From equation (2),

$$E(\Delta w_i) = Prob(\varepsilon_i^n > -\beta X)E(\beta X + \varepsilon_i^n | \varepsilon_i^n > -\beta X) \qquad (3)$$

$$= \beta X F(z) + \sigma_n f(z)$$

where $F(\ )$ is the standard normal cumulative distribution, $f(\ )$ is the standard normal density, and $z \equiv \beta X / \sigma_n$. $F(z)$ is the proportion of contracts with wage increases (i.e., unaffected by downward rigidity).

The estimated effect of downward nominal rigidity on mean wage growth is the difference between the estimated mean in the presence of rigidity (equation (3)) and the estimated mean of the notional wage-change distribution ($\beta X$). This difference (RIG1) is:

$$RIG1 = -\beta X(1 - F(z)) + \sigma_n f(z) \geq 0 \ . \qquad (4)$$

Simpson, Cameron, and Hum calculate rigidity differently. They use the difference between the observed mean wage growth $\Delta w^m$ and the estimated mean of the notional distribution:

$$RIG2 = \Delta w^m - \beta X \ . \qquad (5)$$

RIG1 is the preferred measure of rigidity in this paper. It is calculated in a consistent manner in that *both* the notional outcome and the outcome with rigidity are obtained from the estimated model. Thus, by construction, estimates of rigidity from RIG1 are non-negative. In contrast, RIG2 is not constrained to be non-negative because it subtracts the *estimated* notional outcome from the *actual* mean wage growth.

### *2.1.2 Slope of the short-run Phillips curve*

For later discussion, it is useful to highlight how downward rigidity would affect the slope of the short-run Phillips curve in the standard Tobit model. From equation (3), the slope is the notional parameter $\beta_j$ if no contracts are subject to binding nominal wage floors. Conversely, if some contracts are affected by rigidity ($F(z) < 1$), the slope is

$$\frac{\partial}{\partial X_j}E(\Delta w_i) = \beta_j F(z). \tag{6}$$

Since $0 \le F(z) \le 1$, the Phillips curve is flatter in the presence of rigidity, and it becomes progressively flatter the higher the proportion of contracts affected by downward rigidity.

Equation (6) shows that the slope of the Phillips curve depends on *z*, which is the *ratio* between the mean and the standard deviation of the notional wage-change distribution. If this ratio decreases at lower rates of inflation, more contracts are constrained by nominal-wage floors and the Phillips curve becomes flatter at lower inflation. A decrease in mean wage growth is not a sufficient condition for rigidity to increase at lower inflation, because the standard deviation of the notional distribution may also vary with inflation. This observation illustrates the advantage of estimating rigidity using models that allow $\sigma_n$ to be time-varying.

## 2.2 Previous Tobit studies

The unemployment rate and the 12-month CPI inflation rate are the determinants of notional wage growth in the study of Simpson, Cameron, and Hum. The dependent variable is the wage change in the first year of union contracts signed from 1978 to August 1995. Their model predicts that wage settlements in the combined public and private sectors would have averaged - 0.11 per cent over the 1993–95 period in the absence of pay-cut resistance, whereas the observed average settlement was 0.56 per cent. Thus, they estimate that downward rigidity raised the average settlement by 0.67 percentage points over that period. Simpson, Cameron, and Hum report (without providing specific numbers) that a higher estimate of rigidity is obtained when only private sector contracts are included in the estimation.[5]

The CPI inflation variable is used by Simpson, Cameron, and Hum to capture the effect of inflation expectations on nominal-wage growth. The inflation rate in this measure fell from 1.8 per cent in 1993 to only 0.2 per cent in 1994, owing largely to a substantial decrease in tobacco tax rates, which was widely recognized at the time as a level shift in taxes that had little or no effect on expected future inflation.[6] This means that the variable used by Simpson, Cameron, and Hum probably understates inflation expectations considerably in 1994, resulting in a significant underestimate of the wage growth that would have occurred that year in the absence of

5. “We use the results for the combined public and private sectors, which provide more conservative estimates of the effects of pay-cut resistance according to the results in Table 5.” Simpson, Cameron, and Hum (1998, p. 304).
6. Consistent with this statement, there was little change in private sector forecasts of inflation for 1995 between the end of 1993 (before the tax change) and mid-1994 (after the tax change).

rigidity. Accordingly, their estimate of the average effect of rigidity on wage growth over the 1993–95 period (0.67 percentage points) is probably too high.

Farès and Hogan (2000) reach quite different conclusions from their application of the Tobit model to wage settlements in the manufacturing sector. Although statistically insignificant at the 10 per cent level, their point estimates imply that wage freezes were associated with *lower-than-expected* wage changes, which suggests that freezes tend to reflect wage increases being censored down to zero (consistent with a menu-cost effect) rather than wage cuts being censored up to zero.

The Simpson-Cameron-Hum and Farès-Hogan studies differ in a number of respects, including sectoral coverage, the measure of wage change, and the set of explanatory variables (Table 1). Some of the differences in their results could reflect the different ways in which inflation expectations enter their models. As noted previously, the CPI variable (inclusive of indirect tax effects) used by Simpson, Cameron, and Hum is a probable source of upward bias in their estimate of rigidity in the mid-1990s. The Farès-Hogan study avoids this bias because the effect of inflation expectations is incorporated implicitly through a set of year dummy variables.

Section 3 re-estimates the basic model of Simpson, Cameron, and Hum in order to investigate whether the estimates of rigidity are robust to alternative ways of modelling inflation expectations.

**Table 1: Comparison of Tobit models**

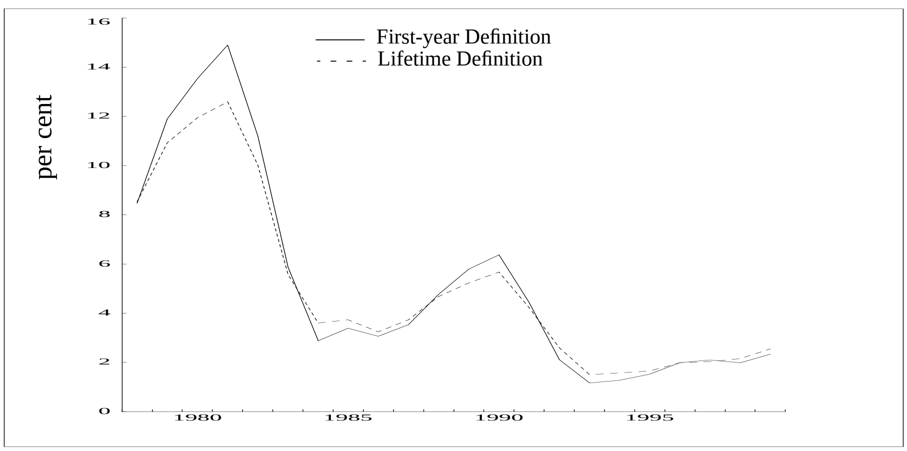
| | **Simpson, Cameron, and Hum (1998)** | **Farès and Hogan (2000)** |
|---|---|---|
| **Sectors** | (i) Combined public and private sectors (unionized)<br>(ii) Private sector (unionized) | Manufacturing (unionized) |
| **Sample period** | 1978:01–1995:08 | 1978–1997 |
| **Dependent variable** | Wage change in the first year of contracts | Average annual wage change over the lifetime of contracts |
| **Explanatory variables** | CPI inflation<br>National unemployment rate | Year dummy variables<br>Regional unemployment rate<br>Regional dummy variables<br>Lagged output growth |
| **Variance of notional distribution** ($\sigma_n^2$) | Either constant or a function of time | Constant |

# 3. Re-Examining the Tobit Estimates of Rigidity

## 3.1 Data

The wage data are private sector wage settlements from the data base of Human Resources Development Canada (HRDC). These data measure the percentage change in the base wage rate in unionized settlements for bargaining units with at least 500 members. To be consistent with the models of Simpson, Cameron, and Hum, this section uses the first-year definition of wage growth. That is, the negotiated wage change for the first year of a contract is used even if an agreement extended beyond one year. Models using the average annual wage change over the lifetime of each contract are estimated in section 4. Figure 1 shows that both measures of wage change averaged about 2 per cent during the 1992–99 period when inflation averaged about 1.5 per cent.

Figure 1: Private sector wage settlements



Following the specification of Simpson, Cameron, and Hum, notional wage growth is assumed to be a function of the national unemployment rate and a variable for inflation expectations.[7] Alternative series are used to proxy inflation expectations in this study:

(i) 12-month CPI inflation

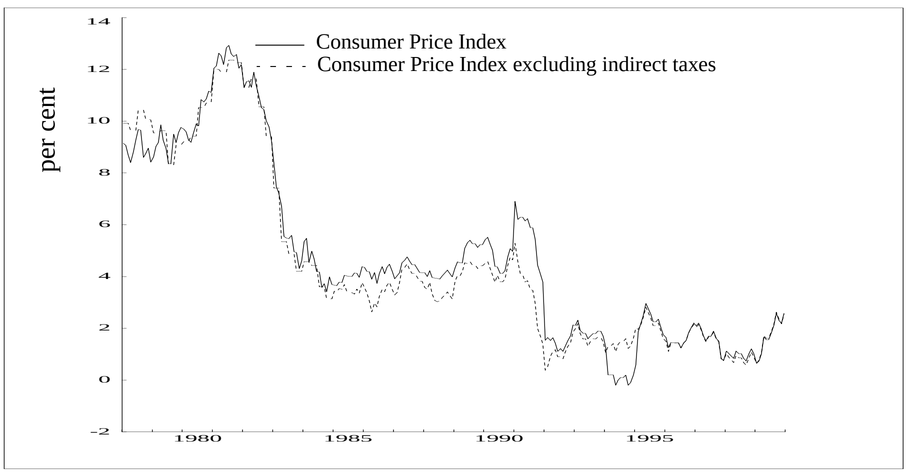
(ii) 12-month CPIxT inflation (CPI excluding the effect of changes in indirect taxes)

(iii) one-quarter-ahead forecasts from a Markov-switching model of inflation

(iv) forecasts of CPI inflation from the Conference Board of Canada's survey of private sector forecasters

Figures 2 and 3 plot these series.

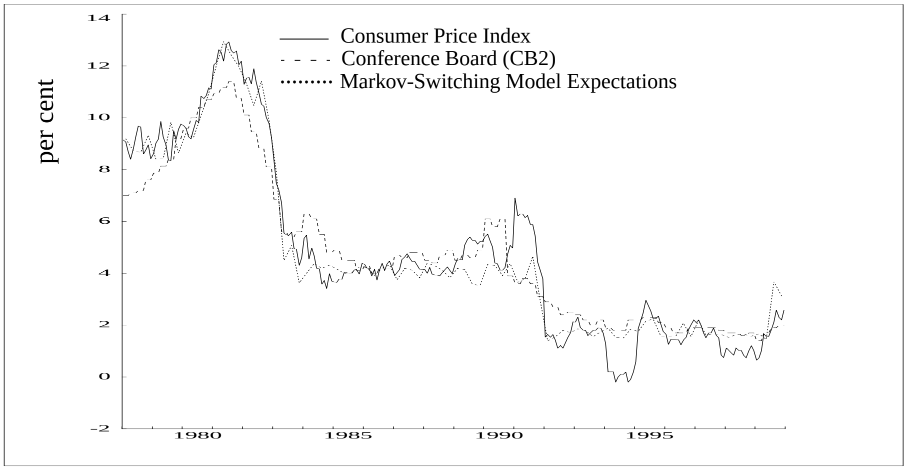
Figure 2: Inflation series




Our first proxy, 12-month CPI inflation, is the expectations proxy used by Simpson, Cameron, and Hum. The series for CPI excluding the effect of changes in indirect taxes is published by the Bank of Canada from January 1984 to the present on a monthly basis. There is also an unpublished quarterly series that begins in 1961. Beginning in 1985, the CPIxT inflation rate was calculated as the 12-month percentage change of the published series. Prior to this date, the inflation rate was calculated by imputing the four-quarter growth rate to each month in a quarter.

7. Each wage settlement is matched with the unemployment and inflation expectations data corresponding to the month the contract was settled. Other determinants of wage growth are incorporated implicitly through the random variable $\varepsilon_i^n$ in the notional wage equation.

Figure 3: Inflation series




The third series used to proxy inflation expectations (defined as MSM) is the one-quarter-ahead forecast from a 3-state Markov-switching model of CPI inflation excluding the effect of changes in the GST, QST, and tobacco taxes.[8] Each of the three states is represented by an inflation process with state-dependent values for the mean rate of inflation, inflation persistence, and the forecast-error variance.

Other proxies are obtained from the Conference Board of Canada's quarterly survey of private sector forecasters. These series are the average private sector forecasts of CPI inflation in the current year (CB1) and the following year (CB2). Unfortunately, there are some quarters prior to 1984 in which the survey was not conducted,[9] and therefore the series for these measures are not continuous over the entire sample period. To obtain continuous series, a linear interpolation of the data was used for missing observations.

On average, the series for inflation expectations are quite similar (Table 2). However, there can be divergences, most notably in 1994, when CPI inflation fell sharply as a result of the decrease in indirect taxes. Therefore, although the different series tend to follow similar paths, CPI inflation is a particularly poor proxy for inflation expectations in 1994, and use of this measure would tend to overstate average rigidity during the 1993–95 period emphasized in the study of Simpson, Cameron, and Hum.

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8. This series for inflation expectations was used by Dupasquier and Ricketts (1998). The data were updated to the end of 1999 by André Binette.
9. From 1978–84, two observations are missing from the forecast for current-year CPI inflation, while 12 data points are missing from the forecast for the next year's CPI inflation rate.

**Table 2: Mean values of proxies for inflation expectations**

| | 1978–99 | 1992–99 | 1993–95 |
|---|---|---|---|
| **CPI** | 4.78 | 1.45 | 1.40 |
| **CPIxT** | 4.51 | 1.50 | 1.73 |
| **MSM** | 4.69 | 1.82 | 1.79 |
| **CB1** | 4.80 | 1.63 | 1.67 |
| **CB2** | 4.79 | 2.01 | 2.12 |

The alternative series for inflation expectations are used to estimate two sets of models. In section 3.2, the variance of the notional distribution is assumed to be constant. The variance is allowed to be time-varying in section 3.3.

## 3.2 Models with a constant variance

*(i) CPI as the proxy for inflation expectations*

The first column in Table 3 reports the estimates of Simpson, Cameron, and Hum for the combined private and public sectors, while the second column lists their results for the private sector. The coefficient for CPI inflation is unusually high in both models. Contrary to theory, the estimated effect of inflation on nominal-wage growth is significantly greater than one.

Column 3 reports our attempt to replicate the Simpson, Cameron, and Hum model of the private sector with the CPI (inclusive of indirect tax effects) as the variable for inflation expectations. Consistent with their study, the sample period is 1978 to August 1995 and the dependent variable is the wage change in the first year of contracts. Despite the apparent use of data from only the first year of contracts in both studies, our model in column 3 is based on 3,736 observations, whereas Simpson, Cameron, and Hum report that over 9,500 private sector contracts were used in their study. The number of contracts used for our estimation is consistent with the information in tables produced by HRDC.[10]

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10. A February 2000 HRDC table indicates that 3,804 contracts were signed in the private sector from 1978 to the end of 1995. This total is consistent with the number of contracts used to estimate our models in columns 3 to 6 of Table 3 (28 contracts containing wage cuts in the first year were excluded from estimation, and 40 contracts were signed in the final four months of 1995). Another unexplained feature of the data used by Simpson, Cameron, and Hum is that they report (in their Table 5) that the private sector accounted for approximately 60 per cent of all contracts in the data base. The HRDC table shows that the correct proportion is about 40 per cent.

**Table 3: Models with constant variance and first-year definition (1978–95)[a]**

| | Simpson, Cameron, and Hum (1998) | | Private sector | | | |
|---|---|---|---|---|---|---|
| | Public and private sectors | Private sector | | | | |
| **Constant** | 2.286 (376.1) | 2.690 (392.3) | 3.92 (9.07) | 4.93 (11.45) | 4.60 (10.71) | 4.56 (10.78) |
| **Unemployment rate** | -0.387 (80.6) | -0.438 (69.7) | -0.52 (13.36) | -0.54 (13.69) | -0.57 (14.62) | -0.64 (16.90) |
| **CPI inflation** | 1.216 (164.7) | 1.243 (139.7) | 1.18 (60.43) | | | |
| **CPIxT inflation** | | | | 1.10 (58.63) | | |
| **MSM expectations** | | | | | 1.19 (59.94) | |
| **CB1** | | | | | | 1.27 (62.06) |
| $\sigma_n^2$ | na[b] | na[b] | 13.22 (40.07) | 13.65 (40.19) | 13.42 (40.03) | 12.97 (39.81) |
| **LLF** | | | -6259.99 | -6318.12 | -6292.67 | -6228.86 |
| **Observations** | 14,983 | 9,535 | 3,736 | 3,736 | 3,736 | 3,736 |
| **Rigidity[c]: 1991–92** | | | | | | |
| **RIG1** | | | 0.87 | 0.97 | 0.89 | 0.93 |
| **RIG2** | 0.94 | na[b] | 0.72 | 1.67 | 1.52 | 1.03 |
| **Rigidity: 1993–95** | | | | | | |
| **RIG1** | | | 1.42 | 1.00 | 1.13 | 1.50 |
| **RIG2** | 0.67 | na[d] | 1.04 | 0.11 | 0.50 | 1.27 |

a. Dependent variable is the nominal-wage change in the first year of the contract. The variance of the notional wage-change distribution is constrained to be constant. T-statistics are in parentheses. LLF is the value of the log-likelihood function.

b. Not reported in Simpson, Cameron, and Hum.

c. "Rigidity" is the estimated effect of downward nominal rigidity on the average wage change in the first year of contracts (percentage points). The two alternative measures, RIG1 and RIG2, are defined in section 2.1.

d. Not reported in Simpson, Cameron, and Hum. See footnote 11.

Our parameter estimate for CPI inflation in column 3 is lower than the estimate of Simpson, Cameron, and Hum, but it is still significantly greater than one. According to their measure of rigidity (RIG2), downward rigidity raised mean wage growth in the private sector by 0.72 percentage points in 1991–92 and by 1.04 percentage points in 1993–95. Despite the large discrepancy in the number of observations, this result is qualitatively consistent with their finding of significant rigidity when CPI inflation is the proxy for inflation expectations.[11]

*(ii) Other proxies for inflation expectations*

The final three columns in Table 3 show results when the other series are used to proxy inflation expectations. The estimates of rigidity are generally high in these models, which assume a constant variance for the notional wage-change distribution and no menu-cost effects. These restrictions are tested in the remainder of the paper.

## 3.3 Models with a time-changing variance

The observed distribution of wage settlements shows a significant decrease in variance as inflation trended downward from an average of close to 11 per cent in 1978–82 to approximately 1.5 per cent over the 1992–99 period.[12] This stylized fact suggests that empirical models should test whether the variance of the notional distribution $\sigma_n^2$ changes over time. The Tobit estimates will be inconsistent if the variance is constrained to be constant when the true variance is heteroscedastic (Maddala 1983).

Simpson, Cameron, and Hum estimate a model in which the notional variance is a function of a time trend $t$:

$$\sigma_n^2 = \sigma_0^2 e^{\alpha t} \quad (7)$$

Since inflation trended downward over their sample period, a negative estimate for $\alpha$ would indicate that the notional variance fell as the economy moved to the lower-inflation years. However, although they do not report their parameter estimates for $\alpha$ or $\sigma_0^2$, they indicate that their results "suggest that this variance is slightly increasing, not decreasing, with time for the private sector" (Simpson, Cameron, and Hum 1998).

The positive time trend for the notional variance in the study by Simpson, Cameron, and Hum is quite surprising, given the strong downward trend in the variance of the actual data. Table 4 shows our results when we use the same specification for the notional variance (equation (7)). In contrast

11. As noted in section 2.2, Simpson, Cameron, and Hum do not report their estimate of rigidity in the private sector during the 1993–95 period, but they do indicate that it is higher than the 0.67 percentage point estimate for the combined public and private sectors.
12. Crawford (2001) presents evidence from a hazard model that much of this trend in the variance reflects a decrease over time in the variance of the notional distribution (i.e., it cannot be attributed exclusively to a thinning of the density in the left tail of the distribution owing to downward nominal rigidity). In that study, the downward trend in the notional variance is attributed to decreases in inflation uncertainty.

to the findings of Simpson, Cameron, and Hum, each of our models has a statistically significant *negative* parameter for the time trend.[13] Thus, relaxation of the constant-variance assumption gives much lower estimates of the variance of the notional wage-change distribution in the low-inflation years of the 1990s (Figure 4).

Since the decrease in variance reduces the density in the left tail of the notional distribution, the models with a time-changing variance (Table 4) have significantly lower estimates of rigidity than their constant-variance counterparts in Table 3. Using the preferred measure of rigidity RIG1, the average estimate of rigidity across the four models is reduced from 1.26 percentage points (Table 3) to 0.71 percentage points (Table 4) for the period from 1993–95. Estimates of the average effect of rigidity on wage growth in 1991–92 are also reduced significantly.

Finally, the estimates of average rigidity in the 1993–95 period are quite sensitive to the measure of inflation expectations. Rigidity averages approximately 0.5 percentage points in the models using proxies that exclude the effect of changes in indirect taxes (CPIxT and MSM), compared to almost 0.9 percentage points in the other two cases. A significant part of this difference occurs because CPI inflation (and CB1) understates forward-looking inflation expectations considerably in 1994, owing to the large decrease in indirect tax rates (described earlier).[14] Models presented in the remainder of the paper use proxies that are not affected by this problem.

13. A potential explanation for the difference in results is the discrepancy between the number of contracts in the HRDC data base and the number of observations reported by Simpson, Cameron, and Hum (see section 3.2).

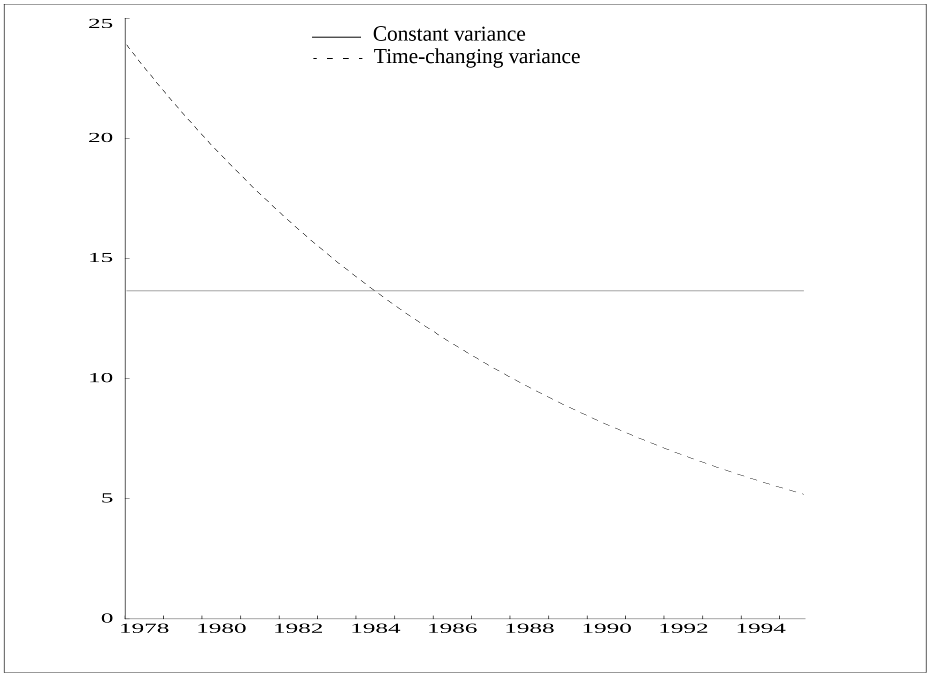
14. Another model was estimated with CPI inflation as the expectations proxy and a 1994 dummy variable to control for the bias in this proxy in 1994. The dummy variable was highly significant and the estimate of rigidity for 1994 was reduced by almost 0.8 percentage points.

### Table 4: Models with time-changing variance and first-year definition (1978–95)[a]

| | Simpson et al. (1998) private sector | Private sector | | | |
|---|---|---|---|---|---|
| **Constant** | 2.685 (510.1) | 4.17 (10.35) | 4.71 (12.61) | 4.52 (10.97) | 4.53 (11.57) |
| **Unemployment rate** | -0.502 (74.6) | -0.50 (13.57) | -0.50 (14.33) | -0.54 (14.46) | -0.59 (16.15) |
| **CPI inflation** | 1.276 (129.9) | 1.12 (57.47) | | | |
| **CPIxT inflation** | | | 1.09 (62.25) | | |
| **MSM expectations** | | | | 1.16 (54.42) | |
| **CB1** | | | | | 1.20 (60.32) |
| $\sigma_0^2$ | na[b] | 21.49 (21.07) | 24.06 (22.57) | 23.60 (21.91) | 21.31 (20.59) |
| $\alpha$ | na[b] | -0.006 (13.38) | -0.007 (17.41) | -0.007 (15.67) | -0.006 (13.10) |
| **LLF** | | -6170.19 | -6182.87 | -6170.94 | -6141.63 |
| **Observations** | 9,535 | 3,736 | 3,736 | 3,736 | 3,736 |
| **Rigidity[c]: 1991–92** | | | | | |
| **RIG1** | | 0.55 | 0.57 | 0.50 | 0.58 |
| **RIG2** | na[b] | 0.50 | 1.47 | 1.29 | 0.74 |
| **Rigidity: 1993–95** | | | | | |
| **RIG1** | | 0.87 | 0.48 | 0.57 | 0.90 |
| **RIG2** | na[d] | 0.71 | -0.08 | 0.26 | 0.86 |

a. Dependent variable is the nominal-wage change in the first year of the contract. T-statistics are in parentheses.
b. Not reported in Simpson, Cameron, and Hum.
c. "Rigidity" is the estimated effect of downward nominal rigidity on the average wage change in the first year of contracts (percentage points). The two alternative measures, RIG1 and RIG2, are defined in section 2.1.
d. Not reported in Simpson, Cameron, and Hum. See footnote 11.

Figure 4: Variance of the notional wage-change distribution
(CPIxT model)




# 4. Models with Menu-Cost Effects

The previous models were based on the assumption that all wage freezes represent contracts with pressures for a notional wage cut. With menu-cost effects, however, the critical threshold for the censoring of notional outcomes would occur at some positive level, rather than zero. The Tobit model is now extended to consider a case in which there is a stochastic non-zero threshold at the micro level. In addition, we lengthen the sample period to include data up to 1999, and consider both the first-year and lifetime measures of wage change.

The key elements of the modified Tobit model are given by equations (8) to (10). Notional wage growth is specified in the same way as previously, with the random variable $\varepsilon_{it}^{n}$ allowed to have a variance $\sigma_{nt}^{2}$ that is time-varying. The modification is that some agents with underlying pressures for a small wage increase may receive a wage freeze instead if there are menu-cost effects. Thus, the observed outcome at firm i ($\Delta w_i$) is a wage change of zero if notional growth is less than some threshold level $k_i$. These thresholds equal a constant $k_0$ plus a contract-specific random term $\varepsilon_i^k$, which is normally distributed across firms with a zero mean and variance $\sigma_k^2$.[15] It is assumed that

15. Models with stochastic thresholds are discussed in Nelson (1977).

$\varepsilon_i^k$ and $\varepsilon_i^n$ are uncorrelated. Thus, actual wage growth at the micro level is now given by equation (10):

$$\Delta w_i^n = \beta X + \varepsilon_i^n \tag{8}$$

$$k_i = k_0 + \varepsilon_i^k \tag{9}$$

$$\Delta w_i = \begin{cases} \beta X + \varepsilon_i^n & \text{if } \beta X + \varepsilon_i^n > k_i \\ 0 & \text{otherwise} \end{cases} \tag{10}$$

The standard Tobit model with downward nominal-wage rigidity (and no menu-cost effects) is a special case of equation (10) in which the threshold $k_i$ is constrained to be zero for all firms. By relaxing this constraint, the notional distribution is estimated in a way that allows for the possibility that some wage freezes are cases in which agents receive a wage change of zero rather than a small wage increase. If menu-cost effects do exist, the standard model will overstate the amount of the notional distribution in the range of wage cuts (and, therefore, tend to overstate downward rigidity). Parameter estimates from the modified model can be used to estimate the net effect of downward nominal rigidity and menu-cost effects on aggregate wage growth.

The short-run wage Phillips curve implied by the Tobit model with menu-cost effects is calculated from equations (8) to (10) as

$$\Delta w = E(\Delta w_i) = \beta X F(\tilde{z}) + \sigma_n f(\tilde{z}) \tag{11}$$

where *F*( ) is the standard normal cumulative distribution, *f*( ) is the standard normal density, $\tilde{z} \equiv (\beta X - k_0)/\sigma_v$ and $\sigma_v = \sqrt{\sigma_{nt}^2 + \sigma_k^2}$ . Rigidity in this model is estimated as the difference between equation (11) and the notional mean, $\beta X$ .

Following Simpson, Cameron, and Hum and the models in section 3, the empirical results reported below are obtained from specifications in which the variance of the notional distribution is a function of a time trend *t*:

$$\sigma_{nt}^2 = \sigma_0^2 e^{\alpha t} \tag{12}$$

As stated previously, a negative parameter estimate for $\alpha$ would indicate that the notional variance fell as inflation trended downward over the sample period.

Table 5 lists the parameter estimates for the models using the wage change in the first year of contracts. Table 6 shows results using the average annual wage change over the lifetime of contracts (the "lifetime" measure of wage growth). Three sets of results are reported: (i) the standard Tobit model with a constant notional variance and no menu-cost effects, (ii) the model with a non-constant variance but no menu-cost effects, and (iii) the model with both a time-changing variance and menu-cost effects.

These results provide empirical support for both extensions to the standard model:

- Parameter estimates for the time trend in the notional variance ($\alpha$) are negative and statistically significant. Thus, relaxation of the constant-variance assumption gives much lower estimates of the variance of the notional wage-change distribution in the low-inflation years of the 1990s. A similar pattern was observed in alternative versions of the model (not reported in the table) in which the level of inflation or inflation uncertainty replaced the time trend in equation (12).

- The menu-cost parameters, $k_0$ and $\sigma_k^2$, are highly significant. Contrary to the key assumption in the standard Tobit model of section 3, this result implies that wage freezes come from *both* sides of zero in the notional distribution.

The estimates of rigidity in Tables 5 and 6 refer to the *net* effect of asymmetric downward nominal rigidity and symmetric menu-costs on aggregate wage growth. These results demonstrate that significantly lower estimates of rigidity are obtained when the model is extended to include a time-changing variance and menu-cost effects:

- Since the decrease in variance reduces the notional density in the left tail of the distribution in the low-inflation years, models with a time-changing variance have significantly lower estimates of rigidity than their constant-variance counterparts. When menu-cost effects are included, the average net effect on wage growth over the 1990s is approximately 0.4 percentage points with the first-year definition of wage change (Table 5), and 0.07 percentage points using the lifetime measure (Table 6).

- Similar estimates of rigidity (not shown in the table) were obtained when the sample was restricted to include only contracts without a cost-of-living clause. We also estimated separate models for short-term contracts (duration up to 12 months) and longer-term contracts, to determine whether rigidity is related systematically to contract length. The estimates of rigidity from models restricted to the longer-term contracts were only marginally lower than in the models based on the full set of contracts.

- The much higher estimates from the first-year models imply that downward rigidity tends to be concentrated in the first year of contracts and is partially reversed in later years of the same contracts. That is, the average effect on the level of wages by the end of contracts (0.07 multiplied by the average contract duration of about 3 years in the 1990s) is *less* than the estimated average effect in the first year (about 0.4 percentage points).

- The estimated effects of rigidity on wage growth are significantly greater in the standard Tobit models, which impose a constant variance for the notional distribution and no menu-cost effects. Since these constraints are rejected in the more general models, we conclude that the standard model overstates the effects of rigidity by a wide margin.

**Table 5: Tobit model (1978–99; first-year measure)[a]**

| | Constant variance | | Time-changing variance | | | |
|---|---|---|---|---|---|---|
| | **CPIxT** | **MSM** | **CPIxT** | **MSM** | **CPIxT + Menu costs** | **MSM + Menu costs** |
| **Constant** | 4.46 (12.16) | 3.91 (10.53) | 4.31 (12.57) | 3.60 (10.84) | 4.29 (13.39) | 3.56 (11.22) |
| **Unemployment rate** | -0.50 (14.24) | -0.51 (14.40) | -0.47 (13.76) | -0.46 (13.85) | -0.45 (13.98) | -0.44 (13.90) |
| **CPIxT inflation** | 1.11 (70.68) | | 1.11 (70.41) | | 1.09 (70.85) | |
| **MSM expectations** | | 1.21 (69.94) | | 1.19 (73.44) | | 1.17 (70.39) |
| $k_0$ | | | | | 0.61 (12.96) | 0.61 (12.99) |
| $\sigma_k^2$ | | | | | 0.18 (4.74) | 0.18 (4.21) |
| $\sigma_0^2$ | 12.92 (44.98) | 12.76 (43.33) | 20.98 (24.48) | 20.26 (24.30) | 20.92 (26.23) | 20.20 (25.14) |
| $\alpha$ | | | -0.005 (17.09) | -0.005 (15.15) | -0.006 (21.07) | -0.005 (18.27) |
| **LLF** | -7306.06 | -7285.30 | -7170.08 | -7169.34 | -7090.18 | -7091.21 |
| **Observations** | 4,426 | 4,426 | 4,426 | 4,426 | 4,426 | 4,426 |
| **Rigidity[b]: 1991–99** | 0.87 | 0.89 | 0.53 | 0.56 | 0.39 | 0.42 |

a. Dependent variable is the nominal-wage change in the first year of private sector contracts. T-statistics are in parentheses. LLF is the log of the likelihood function.

b. “Rigidity” is the net effect of downward nominal rigidity and menu costs on the average wage change in the first year of contracts (percentage points).

**Table 6: Tobit model (1978–99; lifetime measure)[a]**

| | Constant variance | | Time-changing variance | | | | |
|---|---|---|---|---|---|---|---|
| | **CPIxT** | **MSM** | **CPIxT** | **MSM** | **CPIxT + Menu costs** | **MSM + Menu costs** | **CB+ Menu costs** |
| **Constant** | 3.78 (15.81) | 3.35 (13.89) | 3.31 (15.87) | 2.62 (13.02) | 3.28 (16.61) | 2.57 (12.81) | 2.72 (13.25) |
| **Unemployment rate** | -0.32 (13.93) | -0.33 (14.29) | -0.28 (13.23) | -0.26 (12.59) | -0.27 (13.49) | -0.25 (12.26) | -0.32 (15.46) |
| **CPIxT inflation** | 0.92 (90.62) | | 0.93 (89.29) | | 0.92 (98.76) | | |
| **MSM expectations** | | 0.99 (95.02) | | 1.00 (90.73) | | 0.99 (92.22) | |
| **CB survey (CB2)** | | | | | | | 1.09 (89.82) |
| $k_0$ | | | | | 0.43 (11.80) | 0.43 (11.53) | 0.42 (11.09) |
| $\sigma_k^2$ | | | | | 0.10 (3.54) | 0.11 (3.50) | 0.10 (3.84) |
| $\sigma_0^2$ | 5.42 (45.79) | 5.39 (46.28) | 10.17 (25.65) | 10.17 (25.19) | 10.21 (25.78) | 10.21 (26.21) | 10.54 (24.79) |
| $\alpha$ | | | -0.006 (23.18) | -0.006 (22.86) | -0.007 (24.64) | -0.007 (24.33) | -0.007 (24.51) |
| **LLF** | -5907.57 | -5895.76 | -5658.96 | -5653.45 | -5605.98 | -5599.22 | -5585.21 |
| **Observations** | 4,428 | 4,428 | 4,428 | 4,428 | 4,428 | 4,428 | 4,428 |
| **Rigidity[b]: 1991–99** | 0.25 | 0.26 | 0.10 | 0.10 | 0.07 | 0.07 | 0.07 |

a. Dependent variable is the average annual wage change over the lifetime of private sector contracts. T-statistics are in parentheses. LLF is the log of the likelihood function.
b. "Rigidity" is the net effect of downward nominal rigidity and menu-costs on the average annual wage change (percentage points).

Studies of downward rigidity using aggregate data (such as Farès and Lemieux 2001) test for evidence that the Phillips curve becomes flatter at low inflation. For comparison, we consider the Phillips curve implied by the Tobit model. Figures 5a and 6a show the estimated slopes of the short-run Phillips curves from the CPIxT models with menu-cost effects.[16] The changes in slope reflect variations over time in the degree to which downward rigidity and menu costs affect wage

16. Relative to the slope in the standard Tobit model (equation (6)), the Phillips curve is steeper in the presence of menu costs.

outcomes. As Figures 5b and 6b show, the estimated Phillips curve becomes flatter when there is an increase in the predicted incidence of wage freezes. The predicted series follow the actual data for wage freezes closely, although there is a tendency to overstate the number of freezes during the cyclical downturn in the early 1990s. This suggests that the models overstate the amount of rigidity (and, therefore, the change in slope) over that period.

Figure 5a: Slope of the short-run Phillips curve
(First-year measure)

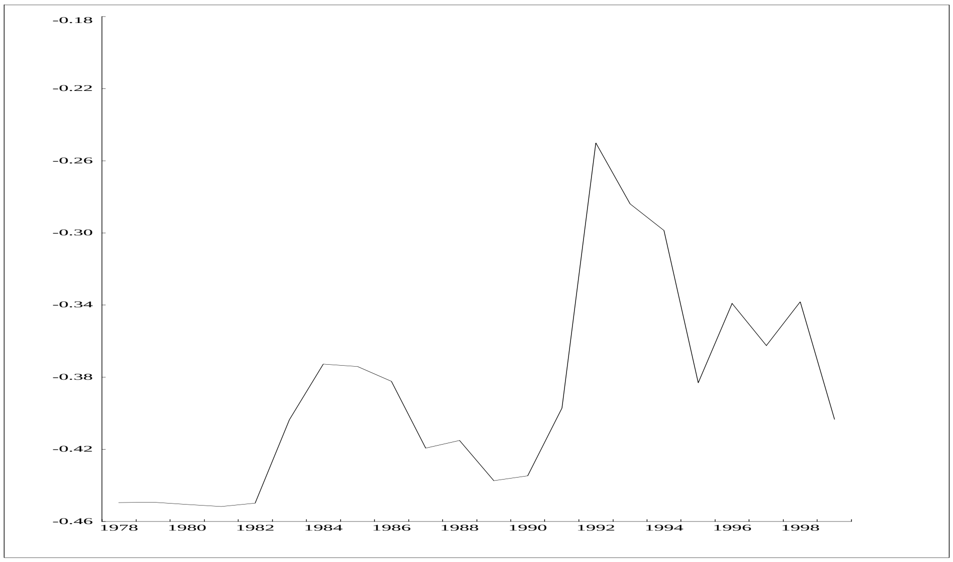



Figure 5b: Actual and predicted freezes
(% of contracts; First-year measure)

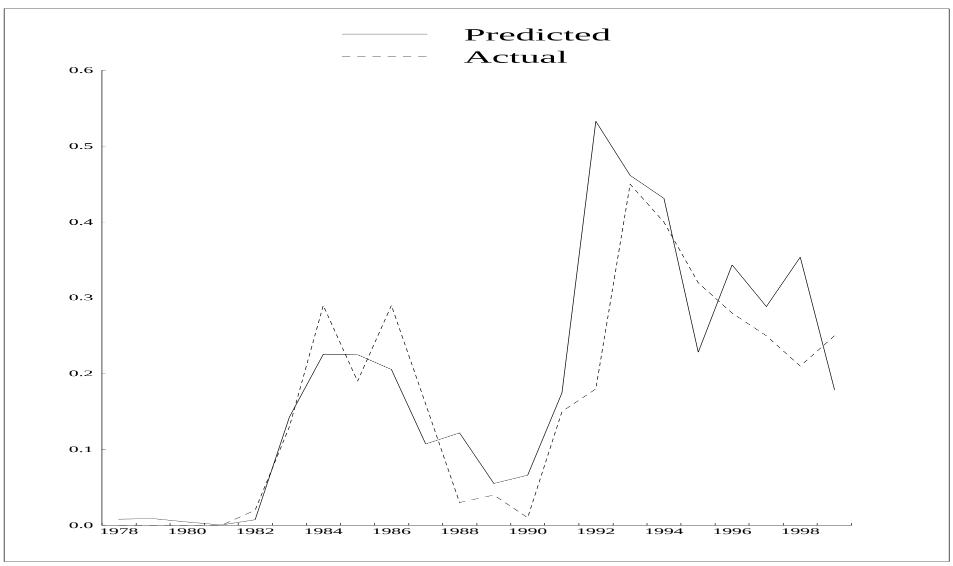


Figure 6a: Slope of the short-run Phillips curve
(Lifetime measure)

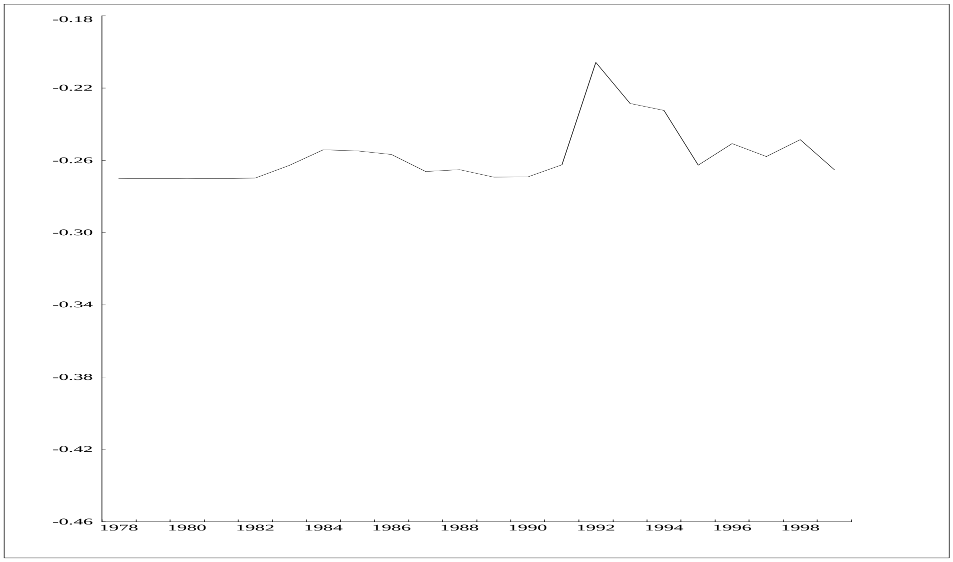
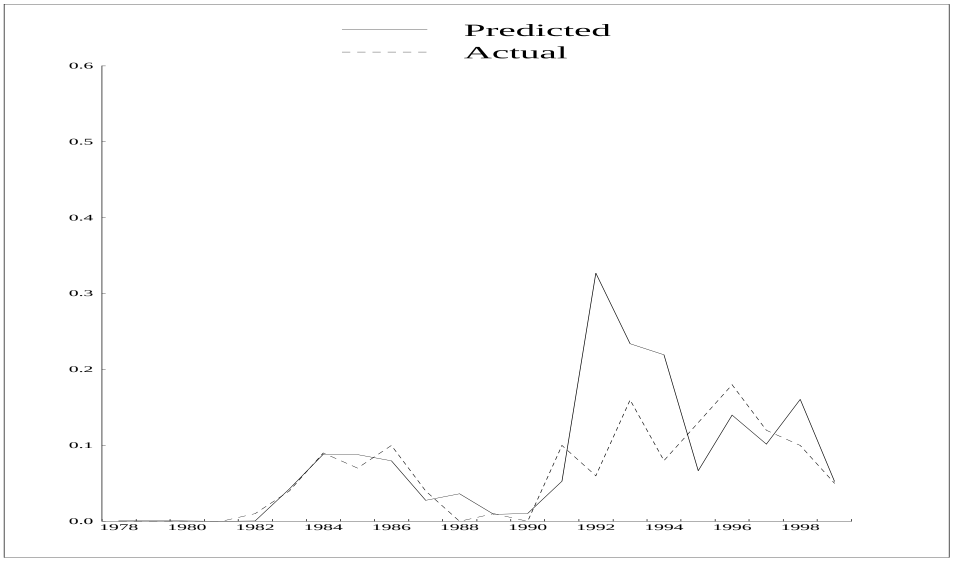



Figure 6b: Actual and predicted freezes
(% of contracts; Lifetime measure)



# 5. Estimates of the Long-Run Phillips Curve

One way to assess the effect of wage rigidity on employment is to examine the long-run trade-off between inflation and the unemployment rate. If nominal wages are downwardly rigid, the long-run Phillips curve will become non-vertical at low rates of inflation. A key issue in assessing the employment implications of rigidity is how changes in the trend rate of productivity growth would affect the long-run trade-off at low inflation. To provide some evidence on this question, productivity growth is added as an explicit determinant of the mean rate of notional wage growth:[17]

$$\mu \equiv \beta X = \beta_0 + \beta_1 U + \beta_2 \Pi^e + \beta_3 \Delta Prod \qquad (13)$$

where $U$ is the unemployment rate, $\Pi^e$ is inflation expectations, and $\Delta Prod$ is the growth rate of labour productivity.

The long-run trade-off between price inflation and unemployment is obtained by assuming that prices are a mark-up over unit labour costs. With this assumption, the long-run trade-off is calculated by subtracting productivity growth from wage growth (defined by equations (11) to (13)) and imposing the long-run condition that $\Pi = \Pi^e$. If the coefficients on expected inflation and productivity growth in equation (13) are constrained to equal one,[18] the long-run trade-off is

$$\Pi = \frac{(\beta_0 + \beta_1 U)F(\tilde{z}) + \sigma_n f(\tilde{z})}{1 - F(\tilde{z})} - \Delta Prod \qquad (14)$$

An equivalent form of the long-run Phillips curve is

$$\Pi = \frac{\beta_1 (U - U_{NR})F(\tilde{z}) + \sigma_n f(\tilde{z})}{1 - F(\tilde{z})} - \Delta Prod$$

where $U_{NR} = -\beta_0 / \beta_1$ is the natural rate of unemployment.[19] The long-run trade-off becomes vertical at the natural rate only when inflation is sufficiently high that downward rigidity and menu-cost effects are not binding (i.e., when $F(\tilde{z}) = 1$). The curve is downward-sloping at lower

17. Implicitly, the constant terms in the models of sections 3 and 4 captured the mean impact of productivity growth as well as other omitted determinants of the mean level of notional wage growth.
18. This assumption is consistent with the long-run equilibrium condition in a competitive labour market.
19. We investigated alternative versions of the model with an index of employment insurance (EI) generosity as a determinant of the natural rate. This variable was significant only in the first-year models. In this section, we report the models without the EI index (the slope of the long-run trade-off was similar or somewhat steeper when the EI variable was included in the model).

rates of inflation if downward rigidity occurs over this range of inflation. A decrease in the variance of the notional wage-change distribution, $\sigma_n^2$, or an increase in productivity growth, would lessen the influence of downward rigidity on wage growth and shift the Phillips curve to the left over its non-vertical range.

The long-run Phillips curve is constructed by evaluating equation (14) using parameter estimates from Tobit models with menu-cost effects and a time-changing variance. Various proxies were used for inflation expectations. In this section, we focus primarily on the results from models with the Markov-switching proxy. Table 7 reports the parameter estimates from models with wage growth measured by the percentage change in the first year of private sector contracts. The first column lists the results from the model in section 4. The second column shows the new results when productivity growth is included explicitly as a determinant of wage growth and the coefficients on expectations and productivity growth are constrained to equal one.[20] Relative to the previous version of the model, there is a significant worsening of the value of the log-likelihood function, with a lower (constrained) parameter on inflation expectations and a stronger effect from the unemployment rate. Estimates of the net effect of downward rigidity and menu costs on the average wage settlement in the 1990s tend to be fairly similar in the two models.

20. The series for (trend) productivity growth is a 5-year moving average of the growth rate of output per person-hour in the business sector. This variable averaged 1.15 per cent over the 1978–99 estimation period. The data are from the annual productivity measures of Statistics Canada (Cansim no. i602502).

**Table 7: Tobit model (1978–99; first-year measure)[a]**

| | | |
|---|---|---|
| **Constant** | 3.56<br>(11.22) | 3.60<br>(11.46) |
| **Unemployment rate** | -0.44<br>(13.90) | -0.49<br>(14.55) |
| **Inflation expectations** | 1.17<br>(70.39) | 1.0 |
| **Productivity growth** | -- | 1.0 |
| $\sigma_0^2$ | 20.20<br>(25.14) | 22.40<br>(23.89) |
| $\alpha$ | -0.005<br>(18.27) | -0.006<br>(17.90) |
| $k_0$ | 0.61<br>(13.00) | 0.60<br>(13.13) |
| $\sigma_k^2$ | 0.18<br>(4.21) | 0.18<br>(4.19) |
| **LLF** | -7,091.21 | -7,294.18 |
| **Observations** | 4,426 | 4,426 |
| **Rigidity[b]: 1991–99** | 0.42 | 0.41 |

a. Dependent variable is the nominal-wage change in the first year of private sector contracts. T-statistics are in parentheses. LLF is the log of the likelihood function. The variance of the notional wage-change distribution is time-varying with the form $\sigma_{nt}^2 = \sigma_0^2 e^{\alpha t}$ .

b. "Rigidity" is the net effect of downward nominal rigidity and menu-costs on the average wage change in the first year of contracts (percentage points).

Table 8 lists the parameter estimates when the same models are estimated using data for the average annual wage change over the lifetime of contracts. Once again, there is a deterioration in the value of the log-likelihood function in the model with productivity growth and the parameter constraints. In this case, the main influence on other parameters is a large decrease in the value of the constant.

**Table 8: Tobit model (1978–99; lifetime measure)**[a]

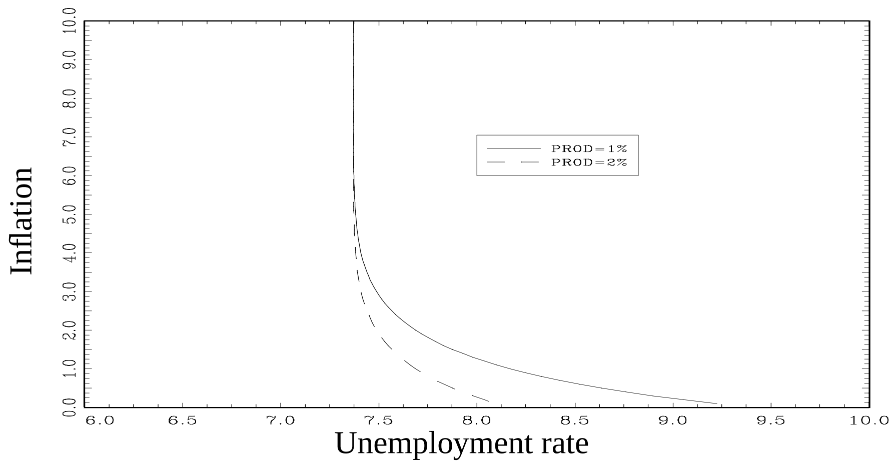
| | | |
|---|---|---|
| **Constant** | 2.57<br>(12.81) | 1.27<br>(6.22) |
| **Unemployment rate** | -0.25<br>(12.26) | -0.23<br>(10.58) |
| **Inflation expectations** | 0.99<br>(92.22) | 1.0 |
| **Productivity growth** | -- | 1.0 |
| $\sigma_0^2$ | 10.21<br>(26.21) | 10.89<br>(25.70) |
| $\alpha$ | -0.007<br>(24.33) | -0.006<br>(21.99) |
| $k_0$ | 0.43<br>(11.53) | 0.39<br>(10.34) |
| $\sigma_k^2$ | 0.11<br>(3.50) | 0.08<br>(4.46) |
| **LLF** | -5,599.22 | -5,856.83 |
| **Observations** | 4,428 | 4,428 |
| **Rigidity**[b]**: 1991–99** | 0.07 | 0.11 |

a. Dependent variable is the average annual wage change over the lifetime of private sector contracts. The variance of the notional wage-change distribution is time-varying with the form $\sigma_{nt}^2 = \sigma_0^2 e^{\alpha t}$.

b. "Rigidity" is the net effect of downward nominal rigidity and menu-costs on the average annual wage change (percentage points).

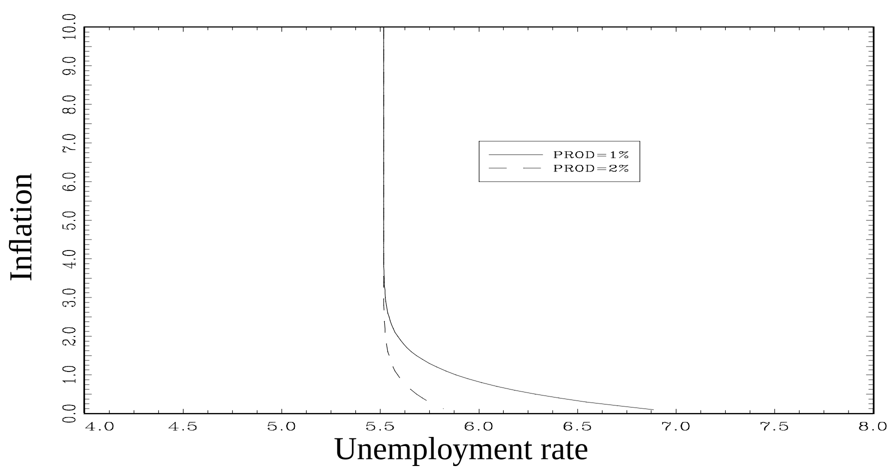
Figures 7 and 8 show the long-run Phillips curve based on parameter estimates from the final columns of Tables 7 and 8 and alternative assumptions for the growth rate of labour productivity (1 per cent or 2 per cent). The shape of the curve depends on the value assumed for the variance of the notional wage-change distribution $\sigma_n^2$. In Figures 7 and 8, this variance is fixed at its average estimated level over the period of low inflation in the 1990s.

Figure 7: Long-run Phillips curve, First-year measure*




* Figures 7 and 8 hold $\sigma_n$ constant at its average level from 1992–99.

Figure 8: Long-run Phillips curve, Lifetime measure




Information from the figures is summarized in Tables 9 and 10. The first columns show the estimated long-run trade-off between inflation and the unemployment rate when productivity growth is assumed to be 1 per cent (close to the average over the sample period). In this scenario, a decrease in long-run inflation would have a relatively small effect on unemployment over the range of the curve above 2 per cent inflation, particularly in the model using the lifetime measure of wage change. As inflation falls to lower levels, the model predicts that a higher proportion of agents are subject to binding nominal wage floors, and the long-run trade-off becomes flatter. Reducing inflation from 2 per cent to 1 per cent is estimated to increase the unemployment rate by

0.48 percentage points in the model for the first-year wage change. The corresponding rise in the unemployment rate is smaller (0.29 percentage points) in the model based on the lifetime measure, with most of this increase (0.20 percentage points) concentrated over the range below 1.5 per cent inflation.[21] In the alternative models using expectations proxies other than the Markov variable, the estimated employment effects were either similar (in the case of the CPIxT proxy) or smaller (Conference Board).[22]

**Table 9: First-year measure**

| **Inflation (per cent)** | **Unemployment rate** | | |
|---|---|---|---|
| | ΔProd = 1.0 | ΔProd = 1.5 | ΔProd = 2.0 |
| **4** | 7.41 | 7.39 | 7.38 |
| **3** | 7.49 | 7.44 | 7.41 |
| **2** | 7.69 | 7.57 | 7.49 |
| **1.5** | 7.88 | 7.69 | 7.57 |
| **1** | 8.17 | 7.88 | 7.69 |

**Table 10: Lifetime measure**

| **Inflation (per cent)** | **Unemployment rate** | | |
|---|---|---|---|
| | ΔProd = 1.0 | ΔProd = 1.5 | ΔProd = 2.0 |
| **4** | 5.52 | 5.52 | 5.52 |
| **3** | 5.53 | 5.52 | 5.52 |
| **2** | 5.59 | 5.54 | 5.53 |
| **1.5** | 5.68 | 5.59 | 5.54 |
| **1** | 5.88 | 5.68 | 5.59 |

21. The estimated curve becomes quite flat in Figures 7 and 8 as the long-run inflation rate falls below 1 per cent. This range of the curve covers inflation rates below recent historical experience.

22. Estimates of the natural rate in the lifetime models were higher than shown in Table 10 when expectations were measured by the CPIxT or Conference Board series.

An increase in trend productivity growth would cause the trade-off to become more favourable at a given inflation rate. If productivity growth is 1.5 per cent, reducing the long-run inflation rate from 2 per cent to 1 per cent would raise the unemployment rate by 0.14 percentage points, according to the model for the lifetime measure of wage change (versus 0.29 percentage points with productivity growth of 1 per cent).

How representative of the aggregate economy are these results from models estimated with the wage-settlements data? To evaluate this question, recall that the long-run Phillips curve was derived under the assumption that inflation is the difference between wage growth and productivity growth (i.e., prices are a mark-up over unit labour costs). The shape of the Phillips curve depends on the frequency of binding wage floors at different levels of inflation. Therefore, if the wage-settlements data are representative of rigidity in the overall economy, the estimated model would tend to be a good indicator of the slope of the aggregate Phillips curve. Conversely, if the wage-settlements data overstate (understate) rigidity in the broader economy, the estimated long-run Phillips curve will tend to overstate (understate) the trade-off at low inflation. Several points can be made.

First, the estimates of rigidity are significantly higher in the models using the first year of contracts rather than the average wage change over the lifetime of contracts. This finding implies that the first-year models will overstate the impact of rigidity on the average wage growth of all contracts in effect during a given year. As a result, the long-run trade-off at low inflation would be steeper than shown in Table 9 and Figure 7.

Second, the wage-settlements data include union contracts covering at least 500 employees. These data have several potential limitations as an indicator of rigidity in the overall economy. They do not include non-unionized employees or workers at small and medium-sized firms. In addition, since they measure changes in the base wage rate, they exclude variable forms of compensation such as bonuses and profit-sharing. The long-run trade-off would be steeper than suggested by the wage-settlements data if rigidity is less widespread in the non-union sector or at smaller firms, or if variable compensation is a source of additional flexibility in total labour costs. Some evidence consistent with this view is provided by informal analysis of the frequencies of wage freezes and wage cuts in alternative micro data bases (see Crawford and Harrison 1998).

Third, the long-run trade-offs shown in Figures 7 and 8 were constructed holding the variance of the notional wage-change distribution constant at its average level during the low-inflation years of the 1990s. In contrast to this assumption, the negative parameter estimates for $\alpha$ imply that the notional variances trended downward as inflation fell in Canada. If the notional variance is positively related to the inflation rate, the long-run trade-off should be calculated with $\sigma_n$

endogenous rather than constant, and the long-run Phillips curve would be steeper than indicated in these figures.[23]

On balance, this discussion suggests that the long-run trade-off is close to vertical at inflation rates of 2 per cent or more if productivity growth is 1 per cent. The estimated trade-off becomes flatter (at an increasing rate) as inflation falls below 2 per cent.

## 6. Conclusions

This paper has reported results from Tobit models that estimate the impact of downward nominal-wage rigidity and symmetric menu-cost effects on private sector wage settlements in Canada. The models were structured to incorporate several stylized facts from the distribution of wage settlements. The results provide evidence that the variance of the notional distribution fell as inflation decreased in Canada, and that menu costs caused some contracts to have wage freezes rather than small wage increases. Each of these findings reduces the number of wage cuts that would be expected in the absence of downward nominal rigidity. Using various proxies for inflation expectations, the net effect of downward rigidity and menu costs in the 1990s is estimated to be approximately 0.4 percentage points for the average wage change in the first year of contracts, and less than 0.1 percentage point for the average annual change over the lifetime of contracts. The evidence suggests that the long-run trade-off between inflation and unemployment is close to vertical at inflation rates of 2 per cent or more if productivity growth is near the average in recent decades.

As a cross-check on these conclusions, it is useful to compare the results from Tobit models with those from other approaches. The first of these comparisons involves another study from the micro literature for Canada. Crawford (2001) studied downward nominal rigidity using a hazard model that provides an estimate of the notional wage-change distribution without imposing strong assumptions on the shape of the distribution.[24] The variance of the notional distribution was allowed to change over time depending on a set of variables, including inflation uncertainty and the sectoral mix of contracts, and the distribution was not constrained to be symmetric. Although the structure of the hazard model is quite different from the Tobit model, these two methodologies

23. Intuitively, if the notional variance falls at lower inflation, a decrease in inflation would lead to a smaller increase in rigidity (and a steeper long-run trade-off) than incorporated in Figures 7 and 8. This relationship follows from the fact that the impact of rigidity on aggregate wage growth depends on both the mean and the variance of the notional wage-change distribution.

24. The wage variable is measured as the deviation from the median wage-change in the current period. An earlier version of the hazard model (Crawford and Harrison 1998) used the level of nominal-wage growth.

give similar estimates of the net effect of downward rigidity and menu costs on wage growth during the low-inflation period of the 1990s.

Another cross-check on the Tobit findings is possible by considering several studies that use reduced-form employment equations (at either the sectoral or firm levels) to test whether downward nominal-wage rigidity has reduced employment in Canada. In addition to their Tobit work, Simpson, Cameron, and Hum (1998) estimated an employment equation in which sectoral employment growth is a function of the incidence of wage freezes in the sector (a proxy for rigidity) and sectoral output growth. They concluded that rigidity had a significant negative effect on employment over the 1993–95 period. Faruqui (2000) extended their model in various ways to better control for the effects of demand shocks. In most of his specifications, the wage-freeze proxy for rigidity has no significant effect on employment growth. A study of the manufacturing sector by Farès and Hogan (2000) found a similar result. The low estimated effects of rigidity on aggregate wage growth in Tobit and hazard models may help to explain why the studies of Faruqui and Farès-Hogan do not find a systematic impact on employment.

Finally, the micro results can be compared with the evidence from other Canadian studies that use a measure of aggregate wage growth to determine whether the Phillips curve shows evidence of downward nominal-wage rigidity at low inflation. Fortin and Dumont (2000) conclude there is a significant long-run trade-off between inflation and unemployment at low and moderate rates of inflation.[25] However, their results imply frequencies of wage freezes that are much greater than observed in any of the micro data bases, which suggests their model significantly overstates the effects of downward rigidity. Farès and Lemieux (2001) find no change in the slope of the short-run Phillips curve at low inflation using a measure of aggregate wage growth that adjusts for compositional shifts in employment. The Farès-Lemieux result is broadly consistent with our conclusion that the long-run trade-off is close to vertical over the range of inflation rates observed in recent years.

In conclusion, our interpretation of the evidence from the Tobit models and other sources suggests that any effect of downward nominal-wage rigidity on employment was small during the low-inflation years of the 1990s.

25. Fortin and Dumont's model includes both downward nominal-wage rigidity and the "near rationality" effects of Akerlof, Dickens, and Perry (2000).

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