How Certain Are We About the Role of Uncertainty in the Labour Contract Duration Decision? Evidence for Canada and Implications

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Introduction

One of the stylized facts of Canadian labour markets in the 1990s and into the new millennium is a lengthening in labour contract duration. Various authors have analyzed the underlying factors influencing the duration of labour contracts in Canada, but they have typically covered only periods up to the late 1980s or early 1990s. A particular area of interest to the Bank of Canada is the role played by inflation uncertainty. Since the adoption of an inflation target by the Bank of Canada in 1991, Longworth (2002) documents that based on a variety of measures, both the level and variability of inflation have fallen. Moreover, a lengthening in contract duration has been cited as an indicator of increased monetary policy credibility by many authors (e.g., Amano, Coletti, and Macklem 1999; Jenkins and O'Reilly 2001; and Longworth 2002). With about 10 years under the new inflation regime, it is therefore an opportune time to re-examine the impact of inflation uncertainty on the duration of labour contracts. The limited empirical evidence for Canada suggests a negative relationship, i.e., that

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greater inflation uncertainty tends to reduce contract length. This is also a finding from the U.S. literature in the area. In addition, with inflation low and stable, bargaining units (BUs)—unions and firms—may have become relatively more concerned about changes in the real economy (for example, whether a recession is likely) and take this into account. Thus, another goal of this paper is to extend the literature in the Canadian context by examining the role of real uncertainty. Since there are several ways that uncertainty may be measured, this paper discusses three methods: statistical-based, those derived from surveys, and regression-based measures. The paper also examines the robustness of the results against various specifications of the uncertainty measures.

Although several uncertainty measures are created and discussed, the empirical analysis focuses on two regression-based methodologies: uncertainty measures derived from either the estimation of autoregressive conditional heteroscedasticity (ARCH) models or a structural vector autoregression (SVAR). The results suggest that inflation uncertainty is significantly related to contract duration, and this finding holds for both methodologies. There is no evidence of a significant link between real uncertainty and contract duration.

The paper begins with an overview of the theoretical link between uncertainty and contract duration. It then discusses the various uncertainty measures that have been used in studies on the determinants of labour contract duration. This is followed by a critique of their relative merits. The paper then provides a summary of the previous empirical research in the area, a description of the estimation techniques used in the analysis, and a presentation of the estimation results. The final section concludes and provides some implications of the results. The appendixes provide additional information on the estimation methodology and results.

1 Theoretical Work

The theoretical research on the duration of labour contracts began with a paper by Gray (1978), who developed a model to determine the optimal labour contract length and the degree of indexation. She shows that, on the one hand, longer contracts allow firms and workers to amortize the costs of negotiation over a greater time period. On the other, shorter contracts minimize the deviations of wages, output, and employment from their desired values because of unforeseen shocks. Thus, the optimal contract duration balances the costs of contracting with potential deviations of variables from their desired levels. Gray argues that for a given degree of indexing, contract length is a decreasing function of uncertainty (no matter what the source) and an increasing function of the costs of contracting.

Canzoneri (1980) came to the same conclusion as Gray by developing a policy evaluation model. In his model, if monetary policy reduces uncertainty in the economy, then contract length becomes longer, and employment, output, and prices all have longer cyclical lags. Indeed, he notes on page 254 of his paper that the result would be a "more inertia ridden cyclical structure with a dampened amplitude."

Danziger (1988) argues that it is important to distinguish between nominal and real uncertainty. In his model, an increase in uncertainty due to changes in real factors such as oil-price shocks leads to longer contract duration, and not shorter, as suggested by Gray and Canzoneri. Basically, workers seek to protect themselves partially or fully against the effects of real shocks by signing longer contracts, which act as a form of insurance. He refers to this as the efficient risk-sharing hypothesis. Moreover, Harris and Holmstrom (1987) suggest that in periods of high uncertainty, information-gathering costs may escalate rapidly, boosting negotiation costs and making longer-term contracts more likely. On the other hand, it is well known that when inflation is very high and variable, contracting ceases to exist. Clearly, disentangling these effects is an empirical question.

1.1 How to measure uncertainty

Uncertainty is unobserved and measures therefore have to be constructed. Our definition of inflation uncertainty is taken from Stuber (2001), who argues that it is the degree to which the future inflation rate is unknown in the sense of not being predictable, given past performance. The literature, however, does not have a well-defined measure of real uncertainty. We therefore use one similar to that for inflation, i.e., we assume that it is the uncertainty faced by economic agents in the sense that the rate of future real output growth is uncertain, given past performance.

Most empirical work has proxied uncertainty in three main ways, using statistical-based measures, those derived from surveys, and measures created through regression analysis. We discuss each method below.

1.1.1 Statistical measures

Statistical measures of inflation uncertainty are usually a simple filter of the data. Those used in various studies include a moving average of the standard deviation of core inflation[1] (Stuber 2001; Longworth 2002) and its coefficient of variation (Kanago 1998). As regards real uncertainty,

1. Core inflation is defined as the consumer price index (CPI) excluding the eight most volatile components and the effects of indirect taxes on the remaining items (see http://www.bankofcanada.ca/en/press/background.pdf).

statistical indicators can be developed similar to those for inflation, using a moving average of the standard deviation and of the coefficient of variation of real GDP (Table 1).

1.1.2 Survey-based measures

Survey-based measures have also been used in previous studies, namely the Livingston Survey in the United States (Vroman 1989). In Canada, roughly similar survey data are available from the Watson Wyatt (formerly KPMG) annual survey of medium-term inflation expectations of economists and portfolio managers.[2] One way to create a measure of inflation uncertainty is to calculate the difference between the upper and lower quartile forecasts among respondents (Stuber 2001; Longworth 2002). Similarly, real uncertainty measures can be derived from the Watson Wyatt Survey of real GDP medium-term forecasts.

1.1.3 Regression-based measures

There has been a considerable evolution in how to measure uncertainty using regression techniques. One of the earliest measures of inflation uncertainty was created by Christofides and Wilton (1983), and it has become the standard in empirical work on contract duration determinants. They ran a rolling regression of the CPI (quarterly growth at annual rates) (P_t) on a polynomial lag of degree 3 over an 11-quarter period. More formally, this can be stated as:

$$P_t = \alpha_0 + \alpha_1 P_{t-1} + \ldots + \alpha_{11} P_{t-11} + dGST_t + e_t .$$

Note that we have included a dummy variable (GST_t) to control for price changes due to the introduction of the GST.[3] Its introduction was well known by agents, and the impact on the price level should have been predictable. Christofides and Wilton found that a third-degree polynomial best fit the data:

$$\alpha_i = a_0 + a_1 i + a_2 i^2 + a_3 i^3 .$$

Uncertainty for a specific period is obtained by the square root of the standard error of the estimate (SEE):

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2. In particular, participants in the Watson Wyatt Survey are asked to provide their medium-term average forecasts for a set of variables, including real GDP and inflation, over a horizon of two to five years ahead (see Stuber 2001).
3. The specification of the dummy variable is as in Crawford and Kasumovich (1996).

Table 1
The pros and cons of the various uncertainty measures

Measure	Advantages	Disadvantages
Moving average of the standard deviation	• relatively easy to calculate	• omits expectations • variability does not infer uncertainty
Moving average of the coefficient of variation	• relatively easy to calculate • relative measure of uncertainty	• omits expectations • variability does not infer uncertainty
Dispersion measures from surveys of expectations	• ex ante measure based on forecasts	• annual survey • forecast horizon not necessarily in line with contract duration • does not take into account the uncertainty attached to each forecast • might be considered a proxy of disagreement
Rolling regression	• explicitly models the forecasting problem • widely used in the literature	• choice of the lag structure is arbitrary in a polynomial lag equation • variance is assumed constant
ARCH/GARCH models	• explicitly models the forecasting problem • uncertainty increases with forecast errors	• relatively sophisticated form of uncertainty • measure of uncertainty is symmetric, although this assumption can be relaxed • potentially sensitive to model specification
Structural VAR model	• uses a set of variables to identify nominal and real shocks instead of only a single variable	• relatively sophisticated form of uncertainty • potentially sensitive to model specification

$$SEE = \left(\left(\frac{1}{T-K}\right)\sum_{t=1}^{n} e_t^2\right)^{\frac{1}{2}},$$

and the equation is then re-estimated for each quarter by adding an additional observation at the end of the sample and dropping the first observation at the beginning of the sample.[4]

Murphy (2000) used a similar technique to create a real uncertainty measure based on the U.S. unemployment rate. Adapting his method to the Canadian

4. For example, if a contract was signed in 1978Q4, then the square root of the SEE of the regression up to 1978Q3 is the estimate of inflation uncertainty. The regression was run using a rolling window, with the initial sample from 1960Q4 to 1978Q3.

context, it can be presented as a rolling regression of the quarterly unemployment rate:

$$\mu_t = \delta_0 + \sum_{i=1}^{11} \delta_i \mu_{t-1} + \sum_{i=1}^{11} \lambda_i \pi_{t-1} + v_t,$$

where μ_t is the aggregate unemployment rate and π_t is the growth in inflation as measured by the CPI. The unemployment rate on the right-hand side of the equation is a polynomial lag of degree 3 over an 11-quarter period, while inflation is an 11-lag, second-degree polynomial. The CPI is included in the equation to control for changes in the unemployment rate due to nominal factors. The real uncertainty measure is the SEE from the regression.[5]

An alternative strategy to calculate inflation uncertainty is to use an ARCH-type model. We draw on work by Crawford and Kasumovich (1996), who created a measure of inflation uncertainty based on an autoregressive (AR) model with ARCH effects. The conditional mean function is defined as an AR(4) process:[6]

$$\pi_t = \beta_0 + \beta_1 \pi_{t-1} + \ldots + \beta_4 \pi_{t-4} + dGST_t + u_t,$$

where π_t is the year-over-year rate of change of the CPI, excluding food and energy and the effects of changes in indirect taxes, and GST_t is a dummy variable to control for price changes due to the introduction of the GST. The conditional variance of inflation is estimated as an ARCH(1) process:

$$h_t = \phi_0 + \phi_1 u^2_{t-1} + \phi_2 \pi_{t-1},$$

where π_{t-1} is included in the variance specification to capture the link between uncertainty and the level of inflation. Inflation uncertainty is measured by the conditional variance of the one-quarter ahead forecast.

A univariate GARCH (generalized ARCH) model was used to create a real uncertainty proxy. Like inflation, real Canadian output growth $(ycan_t)$

5. The sample period is the same as for the inflation uncertainty measure constructed with the rolling regression. Wallace (2001) also created a measure of real uncertainty using similar techniques, though this latter measure was based on oil-price shocks. Specifically, the measure was the standard error from a one-period-ahead quarterly oil-price forecast, using an autoregressive model of the first difference of oil prices over the 1947–80 time period.

6. The model is estimated over the sample period 1962Q3 to 2001Q2.

exhibits significant conditional heteroscedasticity. Our conditional mean equation can be considered a reduced-form IS curve:

$$ycan_t = \gamma_0 + \gamma_1 ycan_{t-1} + \gamma_2 yus_t + \gamma_3 spread_{t-4} + \varepsilon_t,$$

where yus_t is real U.S. GDP growth and $spread_{t-4}$ is the Canadian term structure.[7] The conditional variance of GDP growth is a GARCH (1,1) process:[8]

$$h_t = \Phi_0 + \Phi_1 \varepsilon^2_{t-1} + \Phi_2 h_{t-1}.$$

The final regression method used to create uncertainty measures comes from Rich and Tracy (2000), who developed measures of uncertainty derived from the estimation of an SVAR based on a model developed by Galí (1992). They estimated a four-variable SVAR of real output growth, the change in the yield on the three-month Treasury bill, the ex post real return on the three-month Treasury bill, and the growth rate in the real money supply. The real aggregate shock was linked to real output growth and formed the real uncertainty measure. The remaining three variables comprised an aggregate demand shock and were used to create a nominal uncertainty measure.

The SVAR estimated in this paper is similar in spirit to that of Rich and Tracy. It contains five variables: real GDP, the inflation rate, the real short- and long-term interest rates, and the unemployment rate. From it, we create two aggregate uncertainty measures: a real uncertainty proxy using one permanent shock (real GDP) and one demand shock (unemployment rate), and a nominal uncertainty proxy, using the inflation and interest rate variables. This latter variable is therefore more broad than the inflation uncertainty measure derived from the ARCH model. The sample period used is 1964Q2 to 2001Q4 with a three-year rolling window. Appendix 1 outlines the SVAR methodology in greater detail.

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7. We found that a four-quarter lag of the term structure seemed to perform best, which is the same finding as Cozier and Tkacz (1994).

8. Lagrange multiplier (LM) tests show the volatility clustering of an ARCH process of order one and thus reject the null hypothesis of constant conditional variance over a sample period 1965Q4 to 2001Q2. Lee (1991) showed that the LM tests for ARCH disturbances are identical to an LM test for GARCH disturbances. This leads us to adopt the GARCH (1,1) process, which is found to provide a good description of the data.

1.2 Description of the pattern of the uncertainty variables

The different measures of inflation and real uncertainty are portrayed in Figures 1 to 4. Beginning with the statistical measures of inflation uncertainty, they each suggest that inflation uncertainty has tended to decline over time (Figure 1). Thus, the spike in the moving average of the standard deviation of inflation was relatively high in the 1980s when inflation was also relatively high, and then moved down over the 1990s. The coefficient of variation followed a similar pattern, although it had a larger spike in the early 1990s than in the 1980s. This is not surprising. One might think that a one percentage-point change in inflation in the 1990s could generate more uncertainty than a similar variation in the 1980s, when the level of inflation was much higher.

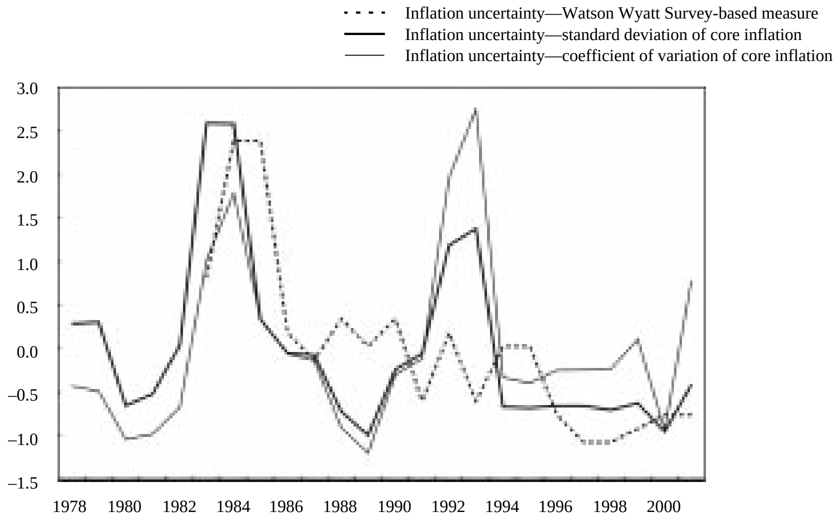
By contrast, the survey-based measure has more or less remained constant at a low level from the mid-1980s, although there was an obvious shift downwards in the latter part of the 1990s. Overall, both the statistical and the survey-based measures point to relatively lower inflation uncertainty in the latter part of that decade and into the new millennium.

The regression measures of inflation uncertainty in Figure 2 typically reveal a more clear downward trend than the other uncertainty variables: the uncertainty variable derived from the SVAR model is an exception. The latter may be due to a number of factors, including the relatively high government debt-to-GDP ratio that drove up long-term real interest rates in the early to mid-1990s, and which are included in the SVAR. The GARCH measure suggests that uncertainty has fallen significantly when inflation appears to have shifted to another regime (Ricketts and Rose 1995), while the Christofides and Wilton (1983) measure has declined more or less continuously over time.

Statistical measures of real uncertainty (Figure 3) do not show as marked a decline as the inflation uncertainty measures. In addition, they are more narrowly dispersed around zero, although they show relatively more volatility around turning points. The survey-based measure also rose in the late 1990s around the time when the Canadian economy was being buffeted by the Asian and Russian crises.

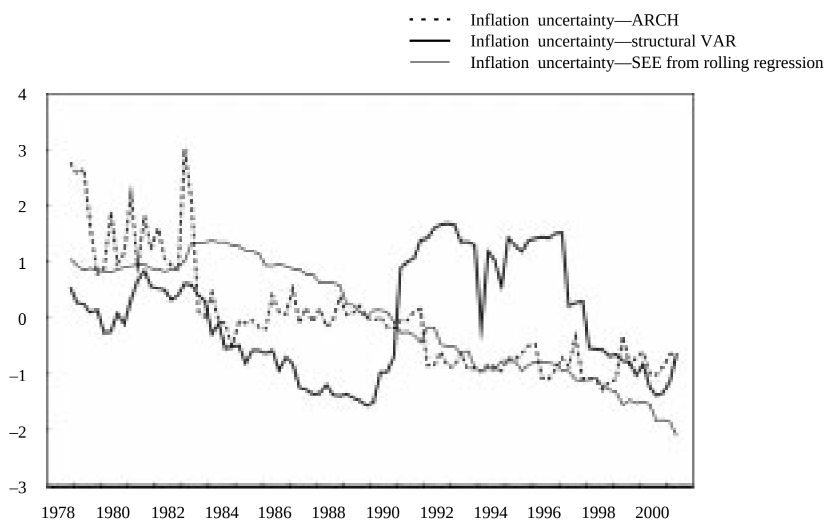
Real uncertainty measures derived from econometric techniques exhibit relatively more dispersion than the statistical ones. Debs (2001) discovered a structural break in the variability of Canadian GDP growth in the second quarter of 1991. This appears consistent with the behaviour exhibited by the

Figure 1
Statistical measures of inflation uncertainty*



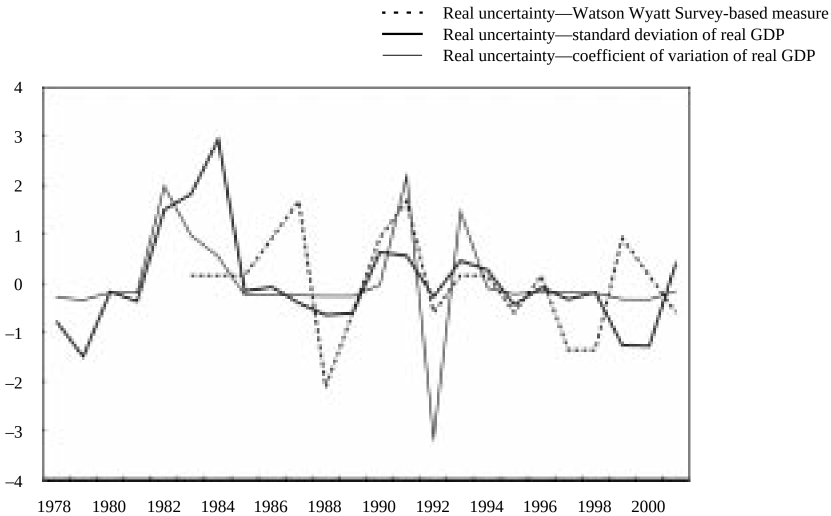
* Standardized to mean 0 and to 1 standard deviation.

Figure 2
Econometric measures of inflation uncertainty*



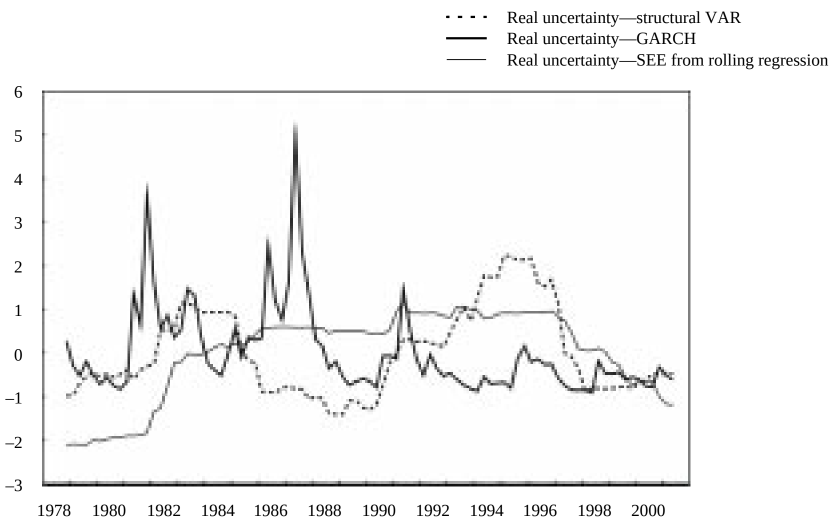
* Standardized to mean 0 and to 1 standard deviation.

Figure 3
Statistical measures of real uncertainty*



* Standardized to mean 0 and to 1 standard deviation.

Figure 4
Econometric measures of real uncertainty*



* Standardized to mean 0 and to 1 standard deviation.

GARCH measure of real uncertainty.[9] By contrast, the SVAR measure of real uncertainty suggests relatively greater uncertainty in the 1990s. The rolling regression measure suggested by Murphy (2000) actually rose somewhat over time—consistent with the behaviour of the unemployment rate—and then fell sharply in the latter part of the 1990s, once again as the unemployment rate declined. By the late 1990s, all of the regression-based measures showed a relatively low level of uncertainty.

1.3 The pros and cons of each set of uncertainty indicators

The distinct behaviour among the uncertainty measures raises the obvious question of which are the most relevant to use in our empirical work. Unfortunately, there is no clear consensus in the literature. There are, however, clear advantages and disadvantages attached to each one (Table 1).

First, survey measures provide an obvious ex ante—what was known by the economic agents at the time of signing the contract—and a direct measure of uncertainty, since the survey measures are not revised and relate to the expectations of forecasters. Nevertheless, an important shortcoming of the Watson-Wyatt data is their annual frequency. In addition, the data range is limited, going back only to 1983. A more important limitation of the data is that they relate to the views of professional forecasters, and can be very sensitive to one forecast outlier. For this reason, we use the difference between the high and low quartile, and not the difference between the high and low forecasts. Nevertheless, there may also be respondents in the survey who do not give their best forecast because of strategic considerations or an incentive to publish a forecast that stands out. An implicit assumption for the use of this measure would be that it is widely known among bargaining parties, who use it as the basis for their negotiations.

Second, the statistical measures are easy to compute. But, as many authors have noted, a standard criticism of statistical measures of uncertainty is that at least a part of their variability is predictable (Stuber 2001). For example, some of the rise in statistical variability in the CPI in the first half of the 1990s was due to federal sales tax reform, which was announced well in advance of its implementation and therefore highly predictable. For that reason, we have used a measure of core inflation that excludes the impact of indirect taxes. Nevertheless, this measure still appears relatively volatile. On the positive side, given that the CPI is rarely revised, these measures can be

9. The real uncertainty measure based on GARCH techniques also reveals relatively high uncertainty in early 1987, which may be explained by the significant variability in real GDP growth at this time period. More precisely, real GDP quarterly growth (annual rates) went from –2.2 per cent in 1986Q4 to 9.0 per cent in 1987Q1.

Table 2
Summary statistics

Variable	Mean	Standard deviation	Mininum	Maximum
Duration (months)	30.4	10.59	5.0	72.0
Number of employees in BU	1,703	3,462.4	500	61,930
Change in number of employees	–28.3	667.4	–7,000	10,210
Incidence of cost of living agreements	0.30	0.46	0.0	1.0
Unemployment rate	9.3	1.64	6.7	12.9
Union density	36.0	0.02	24.0	38.0
Strike	0.11	0.08	0.02	0.28
Inflation uncertainty—3-year moving average of the standard deviation	0.7	0.7	0.02	2.5
Inflation uncertainty—3-year moving average of the coefficient of variation	0.2	0.1	0.006	0.5
Inflation uncertainty—survey-based measure	1.1	0.6	0.3	2.5
Inflation uncertainty—rolling regression	2.2	0.3	1.6	2.5
Inflation uncertainty—ARCH	1.4	0.8	0.4	3.8
Inflation uncertainty—SVAR	0.5	0.2	0.2	0.7
Real uncertainty—3-year moving average of the standard deviation	1.8	0.9	0.4	4.4
Real uncertainty—3-year moving average of the coefficient of variation	1.0	2.7	–7.1	6.6
Real uncertainty—survey-based measure	0.6	0.1	0.3	0.8
Real uncertainty—GARCH	0.4	0.2	0.3	1.4
Real uncertainty—SVAR	0.2	0.09	0.1	0.4

considered ex ante measures of uncertainty. This is not the case, however, for the GDP measure, which is a variable that is typically revised substantially over time.[10]

Third, although statistical measures are relatively easy to calculate, they are naive in the sense that they ignore the essential aspect of the agent's problem; that is, forecasting inflation. Regression methods explicitly model the forecasting problem of the agent and thus provide a better framework to measure inflation uncertainty. Still, it is unclear whether various econometric specifications capture how uncertainty is perceived by agents. Thus, a number of methods have been proposed.

The econometric measure developed by Christofides and Wilton (1983) is easy to calculate and varies over time as inflation changes. But an underlying assumption is a constant variance over time—which is not consistent with the data-generating process—and the choice of the lag

10. This problem could be overcome with the use of real-time GDP data, but at present we do not have a long enough time series for our work.

structure in the polynomial lag equation is basically arbitrary. A more refined measure can be created using GARCH techniques, but this methodology was not available at the time of the original Christofides and Wilton paper. It explicitly allows the modelling of the variance of inflation. The model used in this paper, however, assumes that positive and negative shocks on inflation raise the conditional variance by the same amount.[11] The inflation uncertainty measure is also sensitive to the model specification.[12] By contrast, the real uncertainty GARCH-based measure appears to be robust to different mean specifications of real output growth.[13] Rich and Tracy argue that the SVAR approach has the relative merit of relying on measures that are aggregate in nature rather than on those based on single variables that may be too narrow in scope to be reliable proxies.

For these reasons, our preferred measures of uncertainty are those derived using regression techniques. Within these, the GARCH and SVAR techniques are chosen for the empirical analysis because they explicitly model the conditional variance and contain more of the variables that may play a role in the decision-making of agents. However, it is not clear which better captures uncertainty. We therefore present estimates for each in the analysis below. In Appendix 3, we provide information on the impact of the other uncertainty measures on contract duration.

2 Previous Empirical Research on Contract Duration Determinants

Empirical work on the determinants of labour contract duration began in earnest in Canada with Christofides and Wilton (1983). They used Canadian microdata from Labour Canada on wage contracts signed over the period 1966 to 1975 and found a negative relationship between contract duration and inflation uncertainty. The authors concluded that this negative relationship was most apparent when contracts were not indexed, but that it was still evident even when contracts had a cost of living adjustment (COLA). Christofides (1990) also found evidence for a negative relationship between inflation uncertainty and contract duration, although it was not significant at conventional levels.

11. In fact, more weight can be attributed to a positive shock to inflation uncertainty by using an asymmetric GARCH or a threshold GARCH.

12. Crawford and Kasumovich (1996) found relatively lower inflation uncertainty using a reduced-form Phillips curve model instead of an autoregressive process for the conditional mean function.

13. Adding the real exchange rate or a real non-energy commodity price index to the conditional mean function for real GDP does not significantly change the real uncertainty measure.

A relatively larger number of studies have been carried out with U.S. data. One of the earliest empirical papers to examine the impact of inflation uncertainty was Vroman (1989). She constructed a measure of inflation uncertainty using the coefficient of variation of inflation expectations derived from the Livingston Survey. Her estimation uncovered a significant negative link between it and contract duration. Using a measure of inflation uncertainty similar to that of Christofides and Wilton, Murphy (1992) also found a significant negative relationship between inflation uncertainty and contract duration. The first empirical study to address the issue of real uncertainty was by Wallace and Blanco (1991), who developed industry-specific measures of uncertainty based on producer prices at the sectoral level relative to the producer price index. Their measure of inflation uncertainty was derived from the standard error of a one-period-ahead money-supply forecast based on a simple autoregressive structure. They did not find a significant link between contract duration and either type of uncertainty. Wallace (2001) also did not find evidence of a clear-cut relationship between uncertainty and contract duration, regardless of the source of uncertainty. Murphy's (2000) results supported Danziger's efficient risk-sharing hypothesis, i.e., that real uncertainty leads to contracts of longer duration. In addition, greater inflation uncertainty (using the Christofides and Wilton measure) appeared to reduce the length of labour contracts.

Rich and Tracy (2000) expanded the analysis of contract duration to include a variety of real and nominal uncertainty measures to test the robustness of previous results with U.S. data on labour contracts. They found that the significance of uncertainty varied with the measure of uncertainty. For example, using the Christofides and Wilton measure, they discovered no evidence of a significant link between inflation uncertainty and contract duration. In contrast, survey-based measures of inflation uncertainty similar to those used by Vroman (1989) yielded a significant negative relationship, as did those based on either an ARCH or an SVAR methodology. As regards real uncertainty, contrary to the hypothesis of Danziger (1988), there was no evidence of a significant relationship with contract duration.

In summary, the existing literature, mainly in a U.S. context, reveals a lack of consistent results, suggesting that the impact of various forms of uncertainty on contract duration remains an open question. There are few studies to determine whether this is also the case in Canada. Moreover, it is unclear whether the inconsistent results are due to how uncertainty is constructed or to the role of uncertainty itself.

3 Stylized Facts on Contract Duration

The data for the study are derived from the Department of Human Resources Canada (HRDC) wage-settlements file. The data set contains 11,743 collective agreements in the private and public sectors covering 500 or more employees and beginning in 1978Q1. For our analysis, the data up to the second quarter of 2001 are used. The data set also includes the bargaining unit identifier, the number of employees, whether the contract includes a COLA clause, provincial and industry codes, and settlement, effective and expiry dates.[14]

In this paper, we focus only on the private sector and have excluded public sector contracts, which reduces the sample size to 4,644. The data for the public sector are substantially affected over time by institutional factors (such as wage freezes and suspension of collective bargaining), making them difficult to interpret. Owing to the inclusion of lagged data, the sample size is further reduced to 3,631 private sector agreements, beginning in 1978Q4.[15] Figure 5 presents the annual average contract duration over time. While the trend is clearly upwards, there is a significant increase in the annual average duration in the private sector, particularly from 1996 to 2000. In comparison with the private sector, agreements in the public sector are approximately five months shorter.

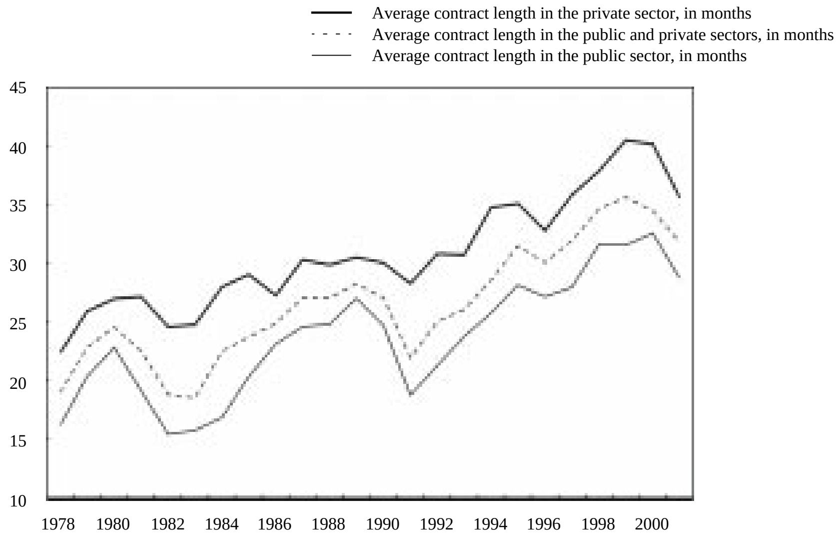
Contracts are typically one, two, or three years in duration (77 per cent of all contracts fall into one of these groups). There is substantial evidence of staggered wage setting over time (not all contracts are one year in duration) and also within time periods (not all contracts are signed on the same date each year). The share of one-year contracts drops by almost 50 per cent between the pre- and post-inflation-targeting periods and longer contracts (greater than 36 months) have become more evident over the decade to 2001. Indeed, the average contract duration has risen by about eight months over the sample period.[16] The annual average duration weighted by employees is almost the same as the unweighted duration (30.1 and 30.4

14. The settlement date is defined as the date when the contract is signed by the BU. The effective date is defined as the date when the provisions of the contract take effect, which can be either before, at the same time, or after the settlement date. In this paper, contract duration is calculated as the difference between the expiry date and the effective date, in months. For example, if a contract expires in $t + 36$ and $t + 60$, then the duration is 36 and 24 months, respectively.

15. The data set contains 3,631 agreements, representing 811 firms and union bargaining pairs, which means that on average each BU is present about 4.5 times in the data.

16. A study by HRDC (2001) found that small BUs (100 to 499 workers) had the highest proportion of collective agreements, with durations exceeding 36 months. Note that our data set does not include these small BUs.

Figure 5
Annual average contract duration



months, respectively) over the sample period (Table 2). On average, there are about 1,700 employees represented by a BU in the private sector and this number has fallen very slightly over time.

Contract length in the Atlantic region is slightly higher than the national average (32.9 versus 30.4 months) (Table 3). By industry, manufacturing represents 49 per cent of all agreements signed in the private sector (see Table 4). Both it and the primary industries sector have the longest average duration, about 31 months.

About 30 per cent of all private agreements contain a COLA clause, and these contracts are approximately five months longer than all other contracts. Roughly 70 per cent of COLA clauses are in relatively long duration contracts (36 months or more), which is consistent with the hypothesis that indexation and longer contracts go hand in hand. Figure 6 shows that the proportion of contracts with a COLA clause has decreased only slightly over time, with most of the reduction taking place in the 1980s. Interestingly, the incidence of a COLA clause did not change very much in the 1990s in the private sector. This suggests that unions, once having bargained for such a clause, would be reluctant to give them up even in a low inflationary environment.

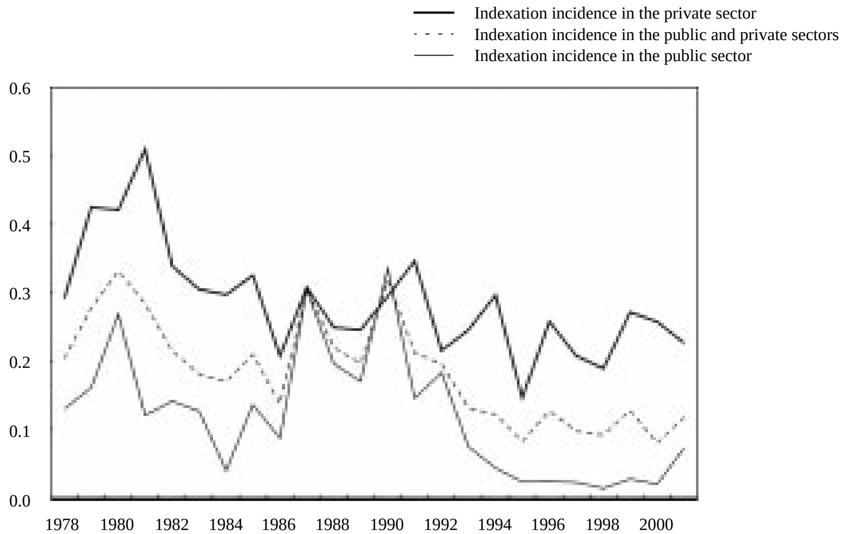
Table 3
Regional distribution of agreements

Region	Proportion
Atlantic	0.060
Newfoundland	0.015
Prince Edward Island	0.001
Nova Scotia	0.026
New Brunswick	0.020
Quebec	0.212
Ontario	0.379
Prairies	0.110
Manitoba	0.033
Saskatchewan	0.017
Alberta	0.058
British Columbia	0.142
More than one province	0.097

Table 4
Industry distribution of agreements

Industry	Proportion
Primary industries	0.06
Utilities	0.03
Construction	0.14
Manufacturing	0.49
Wholesale and retail trade	0.10
Transportation	0.08
Information and culture	0.03
Finance, real estate, and management services	0.02
Education, health, and social services	0.01
Entertainment and hospitality	0.04

Figure 6
Annual proportion of COLA agreements



4 Estimation Techniques

The decision on the length of a labour contract signed by BUs is likely to depend on whether the agreement also contains a COLA clause. Thus, the estimation technique must account for this fact. Typically, the two-stage least-squares technique has been used (Vroman 1989; Murphy 1992), though more recent work has focused on the estimation of a simultaneous equation probit model, as proposed by Heckman (1978) and discussed in Maddala (1983). This type of model has several variations and two are examined in this paper.

The first model can be expressed as the following set of equations:

$$Dur_{it} = X_{1it}\alpha_1 + Cola_{it}\beta_1 + u_{1i}, \tag{1}$$

$$Cola^*_{it} = X_{2it}\alpha_2 + Dur_{it}\beta_2 + u_{2i}, \tag{2}$$

$$Cola_{it} = \begin{matrix} 1 \;\; if \;\; Cola^*_{it} > 0 \\ 0 \;\; if \;\; Cola^*_{it} \le 0 \end{matrix}, \tag{3}$$

where Dur_{it} is a continuous variable of the duration of an agreement in months based on the settlement date; Cola^*_{it} is a latent dependent variable measuring the propensity for a contract to be indexed, such that if $Cola_{it} = 1$, then the contract is indexed to the cost of living, and if $Cola_{it} = 0$, then it is not; X_{1it} is a vector of exogenous variables thought to affect the duration of an agreement, including the uncertainty measures; X_{2it} is a vector of exogenous variables for the cost of living decision (including the relevant uncertainty measures); and u_{1i} and u_{2i} are random disturbance terms.

As shown in Appendix 2, to estimate this model, a restriction must be placed on the coefficients (see Maddala 1983). This restriction is $\beta_1\beta_2 = 0$. Given that the goal of this paper is to examine contract duration, this restriction is imposed by setting $\beta_2 = 0$ in equation (2). The indexation equation is first estimated via a probit using maximum likelihood techniques. Then the duration equation (equation (1)), including the predicted values from equation (2), is estimated by OLS. This type of model has been used by authors such as Vroman (1989) and Rich and Tracy (2000).

In the second formulation, equation (1) is replaced by:

$$Dur_{it} = X_{1it}\alpha_1 + Cola^*_{it}\beta_1 + u_{1i}. \tag{4}$$

In this case, the latent dependent variable—the desire to index—enters into the equation directly. No similar coherency restriction is required in this model. The reduced form for equation (2) can then be estimated as a probit model using maximum likelihood techniques and equation (4) by ordinary least squares (OLS). The fitted values are then substituted back into the structural equations, which are estimated by OLS. This is the type of model estimated by Murphy (2000). Which model should be preferred is not clear. The second model does, however, allow the degree of indexation to affect the duration of a contract and thus may contain relatively more information than the binary variable. Further information on the methodologies is presented in Appendix 2.

In practice, the two methodologies generate very similar results. For that reason, we present the results for the first model below and outline the results of the other methodology in Appendix 3, Table A3.1. All continuous exogenous variables were standardized to zero mean and one standard deviation, while all of the regressions contain robust standard errors using the Huber-White estimator to correct for heteroscedasticity, which was present in the residuals.[17] Moreover, the estimation of the indexation equation contains additional variables thought to affect it, but not the duration decision. These are discussed further in Appendix 4.

5 Exogenous Variables in the Regression

In addition to uncertainty measures, a fairly standard set of variables is usually included in the analysis of the determinants of contract length. They typically proxy either some element of the costs of negotiation or the business cycle.

A set of aggregate variables was attached to each contract based on the settlement date of the contract. They include the national unemployment rate and the share of employed workers who are union members (union density).[18] These variables take on their value for the quarter in which the contract was signed.

The inclusion of the unemployment rate serves two purposes. Christofides and Stark (1996) used it to proxy union bargaining power. A lower unemployment rate should reflect greater union bargaining power, and therefore reduce contract duration. On the other hand, Vroman (1989) included the unemployment rate as a measure of tightness of the labour market and found that contract duration was procyclical, i.e., workers seek to lock in gains in tighter labour markets through longer contracts.

The union density variable captures union bargaining strength. Indeed, Murphy (1992) argues that a key concession by a union is a longer contract, thus a decline in union bargaining strength should lead to relatively longer contracts.

Uncertainty measures are appended using the methodologies outlined above. Additional variables, described below, come from the wage data set itself.

The number of employees covered by each BU is often used as a proxy for several facets of the negotiation process, and as such, the expected sign on its coefficient is indeterminate. First, it is used to proxy negotiation or transactions costs. Larger BUs may have scale economies in negotiation, suggesting a negative coefficient. On the other hand, larger BUs may also have more complex negotiations and thus higher costs, suggesting a positive

17. We also corrected the standard errors for any non-independence of observations across BUs. Although the standard errors were different for some of the coefficients, this did not affect inference, i.e., variables that were significant remained significant, so we do not report these results.

18. The unemployment rate by industry would better relate to specific industry conditions, but data only begin in 1987.

coefficient. Second, it may reflect bargaining power. Murphy argues that unions are more risk-averse than firms.[19] Greater bargaining power by unions would therefore tend to lead to shorter contracts. The change in the number of employees in the BU between two agreements is included to account for membership changes. For example, declines are expected to be associated with longer contracts as unions (with less bargaining power) attempt to lock in gains over a longer period of time (Rich and Tracy 2000).

Indexation is an alternative way of dealing with uncertainty. Results from previous studies (Vroman 1989; Christofides 1990; Rich and Tracy 2000) have found a significant positive relationship between cost of living indexation and contract duration, consistent with the view that longer contracts are more likely to be indexed, helping to ensure the bargaining unit against unfavourable price movements.

A trend variable is included in the regressions to account for the general upward movement in contract duration in the sample period. We include this as the year a contract was signed. Nevertheless, given that this upward trend was most apparent in the 1990s, in one specification, this variable is replaced with a dummy variable taking on the value of one for the period 1992 to present, and zero otherwise. Use of this variable is consistent with work done by Ricketts and Rose (1995), who estimated a Markov-switching model and found that the Canadian inflation process can be split into three distinct regimes, one of which is a low-inflation period beginning in 1992. In addition, the inclusion of this variable is also consistent with the notion that uncertainty is related to the inflation regime as well, and not solely to the level of inflation (O'Reilly 1998).

Finally, two different sets of dummy variables are used in the regression analysis. First, it is likely that contract duration varies across industries. A set of industry dummy variables is therefore included to capture aspects of contracting costs not taken into account by other variables (fixed effects).[20] Industry dummies also capture variation in bargaining power across sectors. Production functions differ among industries, and the bargaining power of a BU probably depends positively on the importance of labour input in the production process. Manufacturing is the excluded sector. Second, a set of

19. Larger firms may have more resources and staff to bargain with unions. In addition, small firms might be more risk-averse than large firms, and prefer to sign shorter contracts that allow them the option of revising the terms of agreements more frequently.

20. The industries are entertainment and hospitality; finance, insurance, and real estate; health and education; information and communications; primary, retail, and wholesale trade; manufacturing; and transportation.

regional dummy variables is included to capture the heterogeneity in settlements across Canada, with Ontario being the excluded province.[21]

Due to the nature of our data set, certain variables that could help clarify the contract duration decision are unavailable. Most of these proxy economic conditions at the firm level.[22] Nevertheless, to the extent that the variables are industry specific, they should be picked up in the industry fixed effects (dummy variables).

6 Estimation Results

Table 5 presents the estimation results for the duration equation of the first model. As a first step, we began by including the inflation/nominal uncertainty measures only (in addition to the other exogenous variables). The coefficients on the inflation/nominal uncertainty measures are negative and significant, suggesting that declining uncertainty of this form has been associated with greater contract length. This negative relationship is consistent with the findings of many other authors. In particular, the regression with the GARCH measure of inflation uncertainty indicates that a one standard deviation increase in inflation uncertainty decreases contract length by about one month. The SVAR uncertainty measure suggests an effect of about two-thirds that of the GARCH uncertainty measure.

Table 6 presents the same equations, but now including the real uncertainty measures. There is little change in the coefficients on the inflation/nominal uncertainty measures. The real uncertainty measures, however, are not significant. Note that for the SVAR measure, we have removed the unemployment rate variable, which was very highly correlated with the real uncertainty measure. In general, these results provide little support for the notion that real uncertainty—at least as defined here—plays an important role in contract duration.

Turning to the other explanatory variables, the coefficient on COLA (predicted values) is positive and significant, indicating that indexation leads to longer contracts, ceteris paribus. The impact is similar across specifications, and indicates that contracts with COLA clauses are close to five months longer than those without such a clause. Longer contract duration for agreements with COLA clauses has also been found by other authors (Christofides and Wilton 1983; Vroman 1989; Rich and Tracy

21. The regions are the Atlantic provinces, Quebec, Ontario, the Prairie provinces, and British Columbia. In addition, a dichotomous variable is included to capture contracts that are in effect in more than one region.

22. An obvious candidate is corporate profits by industry, but data begin only in 1988, not long enough a time period for our analysis.

Table 5
Estimation results: inflation/nominal uncertainty

	1		2		3		4	
COLA	4.8	(0.49)**	4.7	(0.49)**	4.8	(0.49)**	4.7	(0.49)**
Unemployment rate	–1.3	(0.17)**	–0.88	(0.20)**	–2.1	(0.19)**	–1.2	(0.20)**
Union density	0.33	(0.19)	0.46	(0.19)**	0.50	(0.21)**	1.2	(0.19)**
Number of employees	–0.26	(0.19)	–0.25	(0.19)	–0.28	(0.19)	–0.30	(0.19)
Change in employment levels	0.23	(0.15)	0.23	(0.15)	0.23	(0.15)	0.19	(0.15)
Trend	0.60	(0.05)**	0.75	(0.03)**				
Dummy variable for low inflation period					4.8	(0.46)**	10.3	(0.45)**
GARCH inflation uncertainty	–0.99	(0.26)**			–2.4	(0.21)**		
SVAR nominal uncertainty			–0.64	(0.17)**			–2.4	(0.21)**
F-statistic (industry)	9.4		9.3		8.3		7.8	
F-statistic (regional)	12.4		12.7		11.4		11.6	
R squared	0.24		0.24		0.24		0.24	

Notes: Huber-White robust standard errors reported.
** denotes significance at the 5 per cent level.

2000). The coefficient on the unemployment rate is negative and significant, consistent with the finding of Vroman that when labour markets are tight, workers seek to lock in gains through longer contracts. The size of the BU, and its change, appear to have no significant impact on contract duration. This suggests that there are few economies of scale in the bargaining process and that membership changes have not led unions to seek longer contracts. The union density variable is positive and its coefficient typically significant, lending support to the notion that greater union strength leads to a lengthening in contract duration (the opposite of what Murphy (1992) hypothesizes). The coefficient on the trend variable is positive and significant, indicating that contract length has risen about 0.7 months per year. In addition, this trend variable is highly significant and accounts for about half of the explanatory power of the regressions. Moreover, it is noteworthy that the inflation uncertainty variable is still significant, even when a trend variable is included.

While the coefficients on the dummy variables are not reported, they are jointly significant as a group (i.e., by industry and by region). They reveal

Table 6
Estimation results: all uncertainty measures

	1		2		3		4	
COLA	4.8	(0.49)**	4.6	(0.49)**	4.8	(0.49)**	4.7	(0.49)**
Unemployment rate	–1.3	(0.17)**			–2.1	(0.19)**		
Union density	0.36	(0.19)	0.47	(0.19)**	0.50	(0.21)**	1.2	(0.18)**
Number of employees	–0.26	(0.19)	–0.25	(0.19)	–0.28	(0.19)	–0.30	(0.19)
Change in employment levels	0.23	(0.15)	0.22	(0.15)	0.23	(0.15)	0.19	(0.15)
Trend	0.59	(0.05)**	0.73	(0.03)**				
Dummy variable for low inflation period					4.8	(0.49)**	10.0	(0.48)**
GARCH inflation uncertainty	–0.95	(0.26)**			–2.4	(0.21)**		
GARCH real uncertainty	–0.23	(0.15)			0.01	(0.15)		
SVAR nominal uncertainty			–1.1	(0.21)**			–2.9	(0.23)**
SVAR real uncertainty			0.29	(0.23)			0.04	(0.24)
F-statistic (industry)	9.3		9.2		8.3		7.7	
F-statistic (regional)	12.4		12.5		11.4		11.4	
R squared	0.24		0.24		0.23		0.23	

Notes: Huber-White robust standard errors reported.
** denotes significance at the 5 per cent level.

that (relative to Ontario) contract duration is significantly greater in Quebec and the Atlantic region, but significantly lower in Alberta, after controlling for other factors. By industry (relative to manufacturing), contract duration is significantly lower in construction, health and education, as well as utilities, but not significantly different for the remaining industries.

As noted, we replaced the trend variable with a dummy variable taking on the value of one for the time period from 1992. These results are presented in equations (3) and (4) in Table 5. The coefficient on this dummy variable is positive and highly significant (similar to the trend measure), indicating an upward shift in contract duration in the inflation-targeting period. Although this coefficient is well determined within each regression, its amplitude varies depending on the set of uncertainty measures. It is lowest with the GARCH uncertainty measures and highest with the SVAR measure, the latter indicating an increase of about 10 months in the inflation-targeting period.

To check for differences in the slope coefficients, we interacted the dummy variable for the low-inflation period with the inflation uncertainty measures. In the case of the GARCH measure, the coefficient on the interaction term is insignificant. In the case of the SVAR measure, however, the coefficient is significant and negative. This suggests that the relationship between contract duration and uncertainty has become more negative over time.

In summary, the results of the estimation provide considerable support for the notion that lower inflation/nominal uncertainty leads to contracts of longer duration. By contrast, there appears to be little link between real uncertainty and contract duration in the Canadian context. In addition, the results also suggest that contract duration has become longer over the period in which the inflation-targeting regime has been in place. Appendix tables A3.2 and A3.3 provide some sensitivity analysis, using the other measures of uncertainty outlined above. The results vary depending on whether the trend term or the dummy variable is included in the regressions as well as how uncertainty is measured.

Finally, the overall explanatory power of the regressions is not high and it is worth highlighting that more information at the BU level might help to increase the explanatory power of the regressions. One shortcoming of the data set is that it pertains only to large firms where differences across BUs may not be large. Similarly, the sample period is relatively short and does not include the oil-price shocks of the 1970s when inflation began to rise significantly.

7 Conclusions and Some Implications of Longer Contract Duration

Our results confirm the findings of other studies in both Canada and the United States that falling inflation uncertainty leads to contracts of longer duration. And similar to papers examining U.S. data, there remains doubt about the role for real uncertainty in the duration decision. When the list of uncertainty measures is broadened to include those derived from simple filters or survey-based data, the results become much less certain. There are good reasons, however, to prefer the results from uncertainty measures developed through regression techniques.

Nevertheless, additional work could be done to refine the uncertainty measures. For example, the inflation and real uncertainty measures could be jointly estimated in a GARCH-in-mean framework (e.g., Grier and Perry

2000), which may lead to more efficient estimates, although it is not clear that this would change the behaviour of the measures, or the results that much. As noted above, the sample period could also be extended back to the 1960s to obtain more variation in the data, and to help verify the role of uncertainty in contract duration. This may be possible in the Canadian context by linking the database used in this study with those used in previous Canadian studies, such as Christofides and Wilton (1983). It would also be useful to increase the richness of the database with other variables concerning the bargaining process and economic conditions at the time a contract is signed, to the extent that it is possible. This might help to boost the explanatory power of the regressions.

As regards the implications of our results, we confine them to three areas: endogeneity of the contract duration decision, generalizing the results to price-level targeting, and economic welfare.

7.1 Endogenous vs exogenous contract duration

The duration of labour agreements is important for the efficiency of monetary policy because such contracts limit the ability of firms and workers to respond to adverse shocks. As Fischer (1977) noted, this provides the monetary authority with the opportunity to stabilize output even in a rational-expectations setting. In particular, he shows that a role for stabilization policy is created by long-term wage contracts where wages are predetermined in a rational-expectations framework. Thus, longer contracts emerging from falling inflation uncertainty since the early 1980s have potentially increased nominal rigidity in the economy and the stabilization role for monetary policy.

A key assumption by Fischer, however, is that contract length is exogenous. This is also the assumption by authors of papers that use Calvo (1983)-style contracts to model the wage-setting decision. Our results, however, suggest that the decision on the duration of a contract, and thus when wages will be renegotiated, is contingent on the amount of inflation uncertainty in the economy, as suggested by Gray (1978) and Canzoneri (1980). Therefore, when inflation uncertainty is high, BUs are more likely to agree to shorter contracts, ceteris paribus. This would imply that wages would be set more frequently. Consequently, the results of this paper lend support to the notion that contract length is endogenous in the Canadian context. Taylor (1999) summarizes papers that have come to the same conclusion for other countries, but there appears to have been little work incorporating endogenous contract length. He notes, however, that there has been a relatively recent move to create models with “state-dependent pricing,” i.e., where contract duration is contingent on the state of the economy. This is an area that could be explored further by examining the role of nominal-wage rigidities in the presence of endogenous contract duration and the implications for inflation dynamics.

In addition, our results indicate that the duration of contracts varies slightly by sector and while the paper does not address the underlying factors that account for this directly, this nevertheless implies that the impact of monetary policy by sector would be different. While it may appear obvious that monetary policy affects sectors differently (see Farès and Srour 2001), contract duration provides an additional reason why this might be so. A surprising finding, however, is that contract duration is not markedly different among each of the sectors. Indeed, it was only found to be significantly different from the manufacturing sector solely for construction, health and education, and utilities.

An interesting question is that to the extent that nominal rigidity in the economy has increased because of longer duration contracts, why has there not been relatively more output variability? Indeed, even if nominal rigidity has increased, this suggests that it may not pose a problem for the conduct of monetary policy. In the Canadian monetary policy regime, a key role is undoubtedly played by the flexible exchange rate, which tempers the declines in output and employment that would have to come about because of negative economic shocks. But other factors have probably been at play as well, including better inventory management, fewer relative price shocks, the changing structure of the Canadian economy, and better conduct of monetary policy (Debs 2001; Longworth 2002). Of course, nominal-wage rigidity may not have increased all that much. Indeed, the contracts examined cover only a small portion of employment since the unionized sector makes up only about 30 per cent of paid employees, and this share has changed very little over the sample period. Moreover, the focus in this paper was only on the private sector. Wage-setting may also have become more synchronized, which would tend to reduce inertia stemming from staggered contracts. Nevertheless, to the extent that these agreements are similar in nature to the private non-unionized sector would suggest that the results might hold more generally (Taylor 1999). Unfortunately, there is a lack of data that would allow a more detailed look at wage-setting practices. More research is needed in this area, as well as an examination of the link between output and inflation variability.

Indexation also helps to offset nominal inertia. However, as discussed in Appendix 4, there has been little change in the incidence of indexation in the private sector, and we find few statistically significant variables that help to explain the indexation decision. Moreover, it is important to note that while price indexation may reduce nominal inertia, it also aggravates the impact of real shocks on the economy. For example, in the case of a negative supply-side shock, prices rise and output falls. In an indexed contract, wages would rise at the same time that output is falling, exacerbating the negative impact of the shock and likely requiring greater employment declines than would otherwise be the case. There is insufficient information on BU agreements to determine whether they contain clauses linked to real factors.

Besides the macroeconomics factors discussed above, there are several other reasons why workers and firms sign agreements, either explicit or implicit. For example, it is not solely bargaining costs associated with uncertainty that lead firms and workers to contract. Indeed, large fixed costs due to hiring and firing encourage firms and workers to agree to labour contracts of greater duration, especially for skilled workers. Firms may also set wages and labour contracts based on efficiency wage concerns to solicit work effort or reduce shirking. To the extent that it is possible, incorporating more features of why firms and workers agree to contracts into the analysis would also be useful and would allow a check on the robustness of the results related to the uncertainty variables.

7.2 The policy framework

Given that under the Bank of Canada's inflation-targeting regime, uncertainty about inflation appears to have fallen and contract duration lengthened, would a move to a price-level target lead to results similar to those in this paper, i.e., would a price-level target reduce uncertainty further and thus boost contract duration? Unfortunately, there is no clear-cut answer to that question. Conventional wisdom has been that to stabilize the price level, higher than average inflation must be followed by lower than average inflation to meet the price-level target. This would then result in higher inflation variability than inflation targeting, since under the latter, bygones are bygones. Indeed, as Stuber (2001) notes, it may be easy to predict the inflation rate under certain circumstances, but the degree of unpredictability of the price level over longer time periods would remain quite high because of base period drift. A number of different authors have, however, challenged this view, arguing that it depends on a number of factors, for example, the amount of persistence in output, the extent to which expectations are forward- or backward-looking, assumptions made about the type of Phillips curve, and so on (Svensson 1996; Kiley 1998; Maclean and Pioro 2001; Srour 2001).

7.3 Economic welfare

One of the key gains from longer contract duration is likely to be lower transactions costs for the economy, ceteris paribus, and thus a welfare gain. Indeed, with longer contract duration, economic agents spend less time, effort, and money to deal with more predictable inflation. Another benefit is the reduction in costs from labour market disputes such as the loss of output due to strikes. And as Longworth (2002) notes, there is less need to protect oneself against unexpected inflation, which means another savings in resources.

Appendix 1
Structural VAR Estimation

An SVAR model is adopted to calculate measures of nominal and real uncertainty. This methodology has been proposed by Galí (1992) and Rich and Tracy (2000). This framework allows for the identification of structural shocks by decomposition. The SVAR includes five variables to identify both nominal and real shocks, which is an advantage over other measures of uncertainty that are based on a single variable. Obviously, our variable of interest, real chain-weighted GDP, is included. The nominal variable is total CPI year-over-year, which allows us to distinguish between nominal and real shocks. The unemployment rate is included because of its information content for economic activity. Finally, the real day-to-day loan rate and the real 10-year-plus government bond yield are included because of their importance in helping to explain both monetary and fiscal policy changes and long-term investment decisions. All variables are transformed if necessary in order to be stationary and are ordered in the SVAR. Eight lags are used to eliminate serial correlation from the residuals.

The shocks and the variables for the structural model can be summarized as follows:

$$\varepsilon_t = \begin{bmatrix} \varepsilon_S \\ \varepsilon_{d1} \\ \varepsilon_{d2} \\ \varepsilon_{d3} \\ \varepsilon_{d4} \end{bmatrix} \text{ and } Z_t = \begin{bmatrix} \Delta GDP \\ \Delta\pi \\ \Delta u \\ rr_{st} \\ rr_{lt} \end{bmatrix}, \tag{A1.1}$$

where it is assumed that real GDP can be decomposed into one permanent component and four transitory components. Thus, a minimum of identifying restrictions are imposed on the variance-covariance matrix of the vector of structural innovations. It is assumed that only the supply shock has a permanent effect on the level of real GDP, while demand shocks have a transitory effect and no long-run impact on real GDP. This hypothesis gives us four restrictions. We add others to separate the four demand shocks. Under such assumptions, the matrix of long-run effects is lower triangular. The structural model is therefore identified using the Blanchard and Quah decomposition:

$$\begin{array}{cc} & \begin{bmatrix}\varepsilon_S & \varepsilon_{d1} & \varepsilon_{d2} & \varepsilon_{d3} & \varepsilon_{d4}\end{bmatrix} \\ \begin{bmatrix} GDP \\ \pi \\ u \\ rr_{st} \\ rr_{lt} \end{bmatrix} & \begin{bmatrix} r_{11} & 0 & 0 & 0 & 0 \\ r_{21} & r_{22} & 0 & 0 & 0 \\ r_{31} & r_{32} & r_{33} & 0 & 0 \\ r_{41} & r_{42} & r_{43} & r_{44} & 0 \\ r_{51} & r_{52} & r_{53} & r_{54} & r_{55} \end{bmatrix} = \Gamma(1)\,. \end{array}$$

Impulse-response functions of real GDP from this SVAR reveal that transitory shocks have a standard hump-shaped form, and the effect of those shocks dissipates over time. The supply shock has an effect on the level of real GDP, which cumulates steadily over time.

More importantly for uncertainty, the SVAR is an attractive model since uncertainty measures can be calculated by using the five types of shocks. The real uncertainty measure is defined to be a weighted sum of ε_S and ε_{d2}. The nominal uncertainty variable is defined as a weighted sum of three shocks: the monetary shock, ε_{d1}, and the two remaining demand shocks, ε_{d3} and ε_{d4}. Note that the sum of different shocks is weighted by their respective variance (not normalized), i.e., each shock is weighted by the diagonal elements of the contemporaneous variance-covariance matrix. Then, a three-year moving-average standard deviation of this weighted sum gives the uncertainty measures. The choice of a three-year rolling window is based on the response of the level of real GDP to a permanent shock.

Appendix 2
Econometric Methodology

A method used widely in the literature is the simultaneous probit developed by Heckman (1978), which treats the COLA decision as a latent dependent variable. The framework is as follows:

$$Dur_{it} = X_{1it}\alpha_1 + Cola_{it}\beta_1 + u_{1i}, \tag{A2.1}$$

$$Cola^*_{it} = X_{2it}\alpha_2 + Dur_{it}\beta_2 + u_{2i}, \tag{A2.2}$$

$$Cola_{it} = \begin{matrix} 1 \;\; if \;\; Cola^*_{it} > 0 \\ 0 \;\; if \;\; Cola^*_{it} \le 0 \end{matrix}. \tag{A2.3}$$

Because of the latent dependent variable, Heckman notes that for this system to be identified (or coherent), a restriction must be imposed. He calls this the *principle assumption* and it is $\beta_1\beta_2 = 0$. To see why this must be the case, substitute Dur_{it} into $Cola*_{it}$, which gives:

$$Cola^*_{it} = Cola_{it}\beta_1\beta_2 + X_{1it}\beta_2\alpha_1 + X_{2it}\alpha_2 + u_{1i}\beta_2 + u_{2i}.$$

When $Cola_{it} = 0$, then

$$-(X_{1it}\beta_2\alpha_1 + X_{2it}\alpha_2) \ge u_{1i}\beta_2 + u_{2i}.$$

When $Cola_{it} = 1$, then

$$-(X_{1it}\beta_2\alpha_1 + X_{2it}\alpha_2) - \beta_1\beta_2 < u_{1i}\beta_2 + u_{2i}.$$

For logical consistency, this implies that $\beta_1\beta_2 = 0$. This restriction is imposed for estimation, i.e., we assume that $\beta_2 = 0$.

In the first stage, equation (A2.2) is a probit equation estimated by maximum likelihood techniques, and the predicted values are generated for $Cola_{it}$. In the second stage, they are substituted into equation (A2.1), which is then estimated by OLS (see Maddala 1983).

Another way to estimate the model is to allow the latent dependent variable, $Cola^*_{it}$, to enter the system directly. Thus, equation (A2.1) becomes:

$$Dur_{it} = X_{1it}\alpha_2 + Cola^*_{it}\beta_1 + u_{1i}. \tag{A2.4}$$

The methodology (see Maddala 1983) is to then estimate the reduced forms of equations (A2.2) and (A2.4):

$$Dur_{it} = X\Pi_{dur} + v_{1i}, \tag{A2.5}$$

$$Cola^{*}_{it} = X\Pi_{cola} + v_{2i}, \tag{A2.6}$$

where the vector X contains all exogenous variables in the system. Equation (A2.5) is estimated by OLS, and equation (A2.6) is a probit equation estimated by maximum likelihood techniques. The predicted values for Dur_{it} and $Cola^{*}_{it}$ are then substituted back into the structural equations, which can be estimated by OLS. There is, however, an additional complication. Because $Cola^{*}_{it}$ is unobserved, equation (A2.2) must be normalized and estimated by OLS.

$$Dur_{it} = Cola^{*}_{it}(1/\beta_2) - X_{2it}(d_2/\beta_2) - u_{2i}/\beta_2, \tag{A2.7}$$

where $Cola^{*}_{it}$ can be derived from equation (A2.6).

Equation (A2.7) can then be solved for the coefficients α_1/β_2 and $1/\beta_2$ to generate the structural parameters of equation (A2.2). Since the focus of this paper is on contract duration, this second step was not carried out for the COLA equation.

Another issue in the methodology is the extent to which an additional—potentially endogenous—variable, wages, should be entered into the system. Its exclusion would lead to biased estimates to the extent that it should be included. Adding a wage equation into the system is a relatively straightforward extension to each of the models outlined above. We attempted to incorporate it into the system containing the GARCH measures of uncertainty. However, we encountered problems with multicollinearity between the wage variable and the GARCH measure of inflation uncertainty. Not surprisingly, nominal-wage changes have fallen in step with lower inflation uncertainty. Indeed, the correlation between the two variables is over 0.7. Thus, each measure basically tells the same story when included in the duration equation.

Appendix 3
Additional Estimation Results

Table A3.1
Estimation results of duration equation: Model 2 formulation

	1		2		3		4	
COLA	4.8	(0.48)**	4.5	(0.49)**	4.9	(0.49)**	4.7	(0.49)**
Unemployment rate	–1.2	(0.17)**			–2.0	(0.19)**		
Union density	0.27	(0.19)	–0.15	(0.16)	0.41	(0.21)	0.43	(0.16)
Number of employees	–0.26	(0.19)	–0.25	(0.19)	–0.28	(0.19)	–0.30	(0.19)
Change in employment levels	0.23	(0.15)	0.25	(0.15)	0.23	(0.15)	0.24	(0.15)
Trend	0.59	(0.05)**	0.76	(0.03)**				
Dummy variable for low inflation period					4.8	(0.49)**	10.6	(0.47)**
GARCH inflation uncertainty	–0.99	(0.26)**			–2.4	(0.21)**		
GARCH real uncertainty	–0.23	(0.15)			0.02	(0.16)		
SVAR inflation uncertainty			–1.2	(0.21)**			–3.0	(0.22)**
SVAR real uncertainty			0.43	(0.24)			0.17	(0.24)
R squared	0.24		0.24		0.23		0.23	

Notes: Hubert-White standard errors reported.
** denotes significance at the 5 per cent level.

Table A3.2

Estimation results: Non-regression-based uncertainty measures

	1		**2**		**3**		**4**		**5**		**6**	
COLA	4.6	(0.49)**	4.7	(0.49)**	4.1	(0.57)**	4.7	(0.50)**	4.5	(0.50)**	4.2	(0.57)**
Unemployment rate	–1.5	(0.21)**	–1.3	(0.21)**	–0.29	(0.40)	–2.4	(0.23)**	–2.2	(0.21)**	–1.4	(0.43)**
Union density	0.47	(0.21)	0.53	(0.19)**	–1.4	(0.74)	0.90	(0.21)**	1.3	(0.19)**	–2.0	(0.72)**
Number of employees	–0.25	(0.19)	–0.24	(0.19)	–0.28	(0.19)	–0.25	(0.19)	–0.25	(0.20)	–0.31	(0.19)
Change in employment levels	0.22	(0.15)	0.23	(0.15)	0.24	(0.16)	0.19	(0.16)	0.19	(0.16)	0.22	(0.16)
Trend	0.75	(0.03)**	0.70	(0.04)**	0.79	(0.07)**						
Dummy variable for low inflation period							9.2	(0.47)**	9.8	(0.71)**	6.3	(0.57)**
Moving average: std deviation CPI	0.47	(0.21)					–1.4	(0.24)**				
Moving average: std deviation GDP	0.07	(0.27)					1.4	(0.32)**				
Coefficient of variation: CPI			0.21	(0.22)					–0.93	(0.33)**		
Coefficient of variation: GDP			–0.37	(0.15)					0.40	(0.16)**		
Survey: CPI					0.91	(0.34)**					0.17	(0.31)
Survey: GDP					–0.09	(0.19)					0.32	(0.21)
R squared	0.21		0.24		0.21		0.22		0.21		0.21	

Notes: Hubert-White standard errors reported.
** denotes significance at the 5 per cent level.

Table A3.3
Estimation results: Rolling-regression-based uncertainty measures

	1		2	
COLA	4.6	(0.49)**	4.6	(0.49)**
Unemployment rate	–1.1	(0.18)**	–2.0	(0.21)**
Union density	0.47	(0.19)**	0.78	(0.20)**
Number of employees	–0.24	(0.19)	–0.25	(0.20)
Change in employment levels	0.22	(0.15)	0.21	(0.16)
Trend	0.86	(0.10)**		
Dummy variable for low inflation period			3.4	(0.74)**
Rolling regression: inflation	0.77	(0.50)	–2.3	(0.39)**
Rolling regression: real	–0.22	(0.24)	1.1	(0.18)**
R squared	0.24		0.23	

Notes: Hubert-White standard errors reported.
** denotes significance at the 5 per cent level.

Appendix 4
The COLA Equation

While not the focus of this paper,[1] the estimation methodology allows us to examine the determinants of the inclusion of COLAs in contracts. The indexation equation contains additional variables thought to affect it and not the duration decision. These regressors act as instruments that help to disentangle the COLA and duration decisions. To some extent, the decision on which variables to include in each stage (and to exclude from the other) is arbitrary, and economic theory provides only partial guidance. Bearing this in mind, the unemployment rate was excluded, and additional instruments that were included in the first-stage COLA equation (but not in the duration equation) were: whether the previous contract contained an indexation clause, the year-over-year rate of inflation at the time the contract was signed, a measure of strike activity, whether the agreement was the first one signed between the firm and union, and whether it was the first one signed in the 1990s.

The latter two variables were included to proxy transactions costs. If it is the first time a contract has been signed between bargaining parties, then they might be more likely to put in a COLA clause. On the other hand, if it is the first contract signed in the 1990s, then the coefficient might be negative, given lower inflation uncertainty. Whether the previous contract contained a COLA clause was also added to account for bargaining costs. If the previous contract contained a COLA clause, then bargaining costs are likely to be lower. The inflation variable serves as a proxy for expected inflation. Various authors have argued that it should play no role in this equation because what matters is not the expected level of inflation, but its variability (Ragan and Bratsberg 2000). Nevertheless, since the empirical evidence is mixed, we incorporate it in the equation. The strike activity variable is defined as the percentage of estimated working time loss due to strikes[2] and is also included to proxy bargaining costs. Finally, the COLA equation contains only a measure of inflation uncertainty, and no real uncertainty variable, because a COLA clause relates primarily to inflation uncertainty.

As seen in Appendix Table A4.1, to a large extent the main explanatory power for COLA comes from the lagged dependent variable and the union variable. In addition, the industry-regional dummy variables are jointly significant. Thus, the incidence of a COLA clause is likely to be lower

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1. A few empirical studies have dealt with the indexation issue, such as Cousineau, Lacroix, and Bilodeau (1983) and Christofides and Stark (1996).
2. More specifically, the percentage of estimated working time is paid employees divided by workers involved in strikes, weighted by the days of work during a year.

(relative to manufacturing) in construction, retail and wholesale trade, and entertainment and hospitality industries. By region, the incidence is significantly lower in Alberta. Unlike the case of the duration equation, however, the inflation uncertainty measures are seldom significant, but where they are significant, the coefficient is negative, a surprising finding because it indicates that indexation rises as uncertainty falls. At first glance, these results may appear surprising. But it should be borne in mind that over the sample period used in this study, the incidence of a COLA clause in *private* sector contracts was virtually unchanged.

Table A4.1
Estimation results of the COLA equation

	1	2	3	4	5	
Number of employees	0.04 (0.03)	0.04 (0.03)	0.04 (0.03)	0.04 (0.03)	0.01 (0.03)	0.04 (0.03)
Change in employment levels	0.03 (0.03)	0.03 (0.03)	0.03 (0.03)	0.03 (0.03)	0.02 (0.04)	0.03 (0.03)
CPI	0.04 (0.03)	0.02 (0.03)	0.00 (0.03)	0.02 (0.03)	–0.03 (0.04)	0.01 (0.03)
Union	–0.12 (0.03)**	–0.14 (0.03)**	–0.11 (0.03)**	–0.11 (0.03)**	–0.07 (0.07)**	–0.10 (0.03)**
Strike	0.06 (0.05)	0.10 (0.04)**	0.10 (0.04)**	0.06 (0.04)	–0.10 (0.06)	0.15 (0.05)**
Lag COLA	2.4 (0.07)**	2.4 (0.07)**	2.4 (0.07)**	2.4 (0.07)**	2.5 (0.09)**	2.4 (0.07)**
GARCH inflation uncertainty	0.09 (0.05)					
SVAR inflation uncertainty		0.05 (0.03)				
Moving average std deviation CPI			–0.06 (.04)			
Moving average coefficient of variation CPI				–0.09 (0.04)**		
Survey CPI					–0.08 (0.05)	
Rolling CPI						–0.09 (0.05)**
F-statistic (industry)	91.9	53.9	56.2	54.8	36.1	54.9
F-statistic (region)	36.8	22.9	23	23.3	20.5	22.6

Notes: Hubert-White standard errors reported.
** denotes significance at the 5 per cent level.

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