Uncertainty and Labor Contract Durations

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Abstract

This paper provides an empirical investigation into the relationship between *ex ante* U.S. labor contract durations and uncertainty over the period 1970 to 1995. We construct measures of inflation uncertainty as well as aggregate nominal and real uncertainty. The results not only corroborate previous findings of an inverse relationship between contract durations and inflation uncertainty, but also document that this relationship extends to both measures of aggregate uncertainty. We also explore the robustness of this relationship to various measures of inflation uncertainty that have appeared in the literature.

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I. Introduction

It is widely agreed that the existence of labor contracts has potentially important implications for the behavior of the macroeconomy. The issue of contract length is particularly important for the efficacy of stabilization policies and the dynamic behavior of aggregate fluctuations. For example, labor contracts, by limiting the actions of agents during the duration of the contracts, can provide the monetary authority with an advantage in reacting to shocks in an economy and a role for stabilizing output [Fischer (1977)]. In addition, multi-period contracts can serve as a key propagation mechanism for shocks, with their structure being central to the adjustment process of output, prices and wages. This latter consideration is crucial to the design and conduct of disinflation policy [Taylor (1980)].

Recent union labor contract renewals have resulted in significantly longer contracts being signed than in the past. For example, the three major auto manufacturers and the United Auto Workers signed four-year national contracts in their last round of negotiations, breaking a long standing tradition of three-year contracts [Bloomberg (1999), Hyde (1999) and Phillips and Despeignes (1999)]. The United Steel Workers and Kennecott Copper Corporation signed a five-year contract in October of 1996 [Thomson (1996)]. A six-year contract with Goodyear Tire Company was ratified by the United Steel Workers in May of 1997 [Gamboa (1997)].

These examples illustrate a recent trend in the durations of union labor contracts. Figure 1 shows by year the 10^{th}, 25^{th}, 50^{th} , 75^{th} and 90^{th} percentiles of the distribution of *ex ante* contract duration for all major union contracts negotiated in that year and included in the *Bureau of Labor Statistics* (BLS) data.[1] Apart from some volatility in the early 1970s, the 10^{th} and 50^{th} percentiles have remained remarkably constant over this period. In contrast, the 90^{th} percentile shows a steady upward progression starting in the mid-1980s, with the 75^{th} percentile following suit beginning in the early 1990s.

There is a rich theoretical literature on the determinants of desired contract durations, largely based on the framework developed by Gray (1978) [see also Canzoneri (1980)]. In an extension of her earlier work on indexation, Gray argues that contract length should be positively

[1]The BLS defines a major union contract to be one which covers at least 1,000 workers. The *ex ante* duration as reported by the BLS is the number of months between the effective date of the contract and the planned expiration date. The *ex post* duration may differ from the *ex ante* durations due to early renegotiations or delayed settlements (which effectively extend the duration of the contract). For the remainder of this paper, we will refer to the *ex ante* duration simply as the contract duration.

related to transactions costs and inversely related to uncertainty, regardless of whether the uncertainty pertains to real or nominal shocks. Subsequent work suggests that the source of the uncertainty may matter for the desired contract duration. Danziger (1988) develops an implicit contract model where workers are risk averse and firms are risk neutral. Within this framework, contracts allow for efficient-risk-sharing between parties and provide workers with a means of insuring against income fluctuations due to aggregate productivity shocks. As a consequence, greater real uncertainty causes workers to seek increased insurance through longer contracts.

The theoretical literature on contract durations spawned a host of empirical studies aimed at identifying their key determinants. A primary focus of this research was understanding the impact of inflation uncertainty on contract durations. This line of research largely came to a halt in the early 1990s having failed to reach a consensus. There are several reasons why a return to this empirical question is worthwhile. First, as noted above, significant changes in contract durations emerged after the end of the sample periods used in the earlier investigations and may provide an important source of variation for identifying the underlying determinants of the desired contract duration. Second, developments in estimation techniques afford a more formal approach to deriving key variables of interest. For example, time series models of heteroskedasticity are now widely used to obtain time-varying measures of uncertainty [Engle (1982, 1983)]. In addition, the advent of more structural based procedures allow for a better integration of theory into the identification of nominal and real shocks [Blanchard and Quah (1989), Gali (1992)].

One objective of this paper is to clarify the empirical relationship between contract durations and inflation uncertainty. We will explore the sensitivity of this relationship to alternative methods for measuring inflation uncertainty using a consistent empirical specification and set of contract negotiations. A second objective of this paper is to extend the previous analysis and test the competing predictions of the Gray (1978) model and risk-sharing models regarding the sources of uncertainty and their effects on contract durations.

Our basic finding is that there is a significant inverse relationship between uncertainty and contract durations that is robust to two of the three methodologies for measuring inflation uncertainty. We further argue that the approach that does not provide evidence of a negative relationship may be problematic based on its method of construction and anomalies associated with its behavior. These considerations raise some concern about the reliability of results based on this particular measure of inflation uncertainty which has been used extensively in this literature. Importantly, when we extend the analysis to incorporate nominal and real uncertainty, we find that both types of uncertainty have a significant negative relationship with contract durations. Taken together, this evidence indicates that labor contract durations are endogenous to the economic environment prevailing at the time they are signed, but that risk-sharing concerns are not paramount.

In the next section of the paper, we discuss various measurement issues that arise when dealing with labor contract durations. We also outline the econometric framework we use in our estimation. Section III explores various methodologies that have been proposed in the literature to measure inflation uncertainty as well as nominal and real uncertainty. The variables used in the estimation are discussed in section IV along with the empirical findings. The paper concludes with a short summary of our findings.

II. Econometric Specification

The next four subsections discuss issues related to the econometric framework used to analyze the determinants of contract durations. These issues are: the measurement of the contract duration, censoring, indexation and scheduled reopenings, and the unbalanced nature of the panel data.

1. Measuring Contract Durations

The empirical literature has considered two different definitions of the contract duration.[2] Wallace and Blanco (1991) define the contract duration to be the number of months between the prior contract expiration and the current contract expiration (Dur_1). The BLS in their publication *Current Wage Developments* (CWD) defines the contract duration as the number of months between the contract's effective date and its expiration date (Dur_2). This latter definition has been

[2]The term labor contract is a legal misnomer, since they are agreements and not contracts. One consequence of this distinction is that the portion of the agreement dealing with the terms and conditions of employment is deemed to survive the expiration of the agreement (so long as no new agreement has been negotiated and the parties have not reached a bargaining impasse). For the rest of the paper, however, we will ignore this distinction and will refer to these agreements as contracts.

adopted by Christofides and Wilton (1983), Christofides (1985, 1990), Vroman (1989), and Murphy (1992).

The theoretical literature stresses the importance of the information available to the bargaining unit (BU) at the time the contract is negotiated. Testing for the impact of uncertainty on contract durations may require careful attention to the timing of the actual negotiations. This suggests a third definition of the contract duration which is the number of months between the contract's negotiation date and its expiration date (Dur_3). In practice, the negotiation date can precede or follow a contract's effective date (and prior expiration date). For example, in an early settlement where the BU replaces a contract prior to its expiration (6% of negotiations), the negotiation date precedes the effective and prior expiration dates. Similarly, in a delayed settlement with backdating of the terms of the agreement (35% of negotiations), the negotiation date follows the effective date and prior expiration date.

There are two potential consequences of using either of the two existing definitions of the contract duration. First, the dependent variable in the analysis may be mis-measured. Second, there may be a timing mismatch between the uncertainty measure and the contract duration it is trying to explain. The extent of this mismatch depends on the time differences between the prior expiration date, the effective date and the negotiation date, as well as the degree of time aggregation used in the construction of the uncertainty measure.

Most of the prior empirical studies have used uncertainty proxies measured at a quarterly frequency. Table 1 shows the extent of this timing mismatch for our sample of contract negotiations. For each of the two existing duration measures, the table indicates the percent of the sample that would be mismatched for a given number of quarters. For example, the table shows that the use of the effective date would involve the same quarterly timing as the negotiation date for 75% of the negotiations, while 18% of the negotiations would involve timing that is either one quarter ahead or one quarter behind the negotiation quarter. It is also clear from the table that there is a much greater coincidence of timing between the effective date and the negotiation date (Dur_2 vs Dur_3), than between the effective date and the prior expiration date (Dur_1 vs Dur_3).

2. Censoring of the Contract Duration

The empirical literature on contract durations has assumed implicitly that the observed contract duration is the same as the desired contract duration.[3] Given the heterogeneity across BUs in the data, this would likely imply a smooth distribution of observed contract durations. Figure 1 suggests that this may be at variance with the data. Define a "timely settlement" to be one where the negotiation of the subsequent contract is concluded at the expiration of the current contract. Timely settlements have the feature that all three definitions of the contract duration discussed above are the same. Figure 2 shows a histogram of durations for all timely settlements pooled across the years of our sample. A clear feature of the data is the prominence of contracts that are multiples of twelve months in duration. Ninety-three percent of these settlements involve contracts durations that are multiples of twelve months.[4]

This pattern of contract durations illustrates the strong desire by many BUs to conduct their contract renewals during a specific time of the year. This interest (which is outside of the basic economic models under consideration in this paper) creates an inertia to the decisions of the BUs regarding the current contract duration. Variations in the prevailing uncertainty at the time of the negotiations may generate changes in the desired contract durations, but not in the observed contract durations. That is, the benefits arising from making small adjustments in the contract duration are outweighed by the costs of departing from the current seasonal timing of the negotiations.

In addition, when the BU does decide to adjust its contract duration, it is likely to adjust the duration in a way that maintains the seasonal timing. This can be seen in Figure 3, where we show the first difference in contract durations, again for the sample of timely settlements. For the cases where we have adjacent timely settlements, 80% involve no change in the duration. In 74% of the cases where the duration is adjusted, the change involves a multiple of a year.

[3]By desired we mean from the vantage point of the economic model to be tested.

[4]We include in this calculation contract durations that are 11, 23, 35, 47 and 59 months since these are generally the result of a prior expiration date being at the end of a month and the negotiation date being at the beginning of the next month.

Allowing for censoring of the underlying desired contract durations can be accommodated by assuming that the union and firm use the following simple rounding rule. If the BU chooses to only negotiate at a particular time of year, we assume that they round their desired contract duration up or down to satisfy that constraint in the way that minimizes the absolute deviation between the actual and their desired contract duration. For any censored contract duration, this implies that the unobserved desired duration lies in an interval corresponding to the six months prior to and the six months subsequent to the observed contract duration.

3. Indexation and Scheduled Contract Reopenings

The empirical literature has included an explanatory variable measuring the indexation of the contract to changes in the cost-of-living.[5] Because cost-of-living adjustment (COLA) clauses are jointly negotiated along with the length of the contract, the COLA clause is treated as endogenous.

There is a second mechanism by which the BU can build flexibility into the contract. A contract can specify one or more scheduled reopenings for the purpose of negotiating deferred wage increases. A scheduled reopening differs from an unscheduled reopening. An unscheduled reopening of the contract can occur at any time during the term of the contract with the mutual consent of the firm and the union. An unscheduled reopening requires the firm and the union to "negotiate in good faith" over all terms and conditions of employment. In contrast, a scheduled reopening occurs at a prespecified date, and only obligates the firm and the union to negotiate in good faith over that specific deferred wage increase.[6]

The existing literature has not controlled for the presence of a reopening clause. The relative importance of these two forms of contract flexibility for our data is illustrated in Table 2,

[5]For reasons of data availability, most of the literature has only included an indicator for whether a contract contains a cost-of-living adjustment clause. Christofides and Wilton (1983), Card (1986) and Christofides (1990) examine the degree of indexation.

[6]A minority of reopening clauses are triggered by a prespecified movement in a cost-of-living index, instead of having their date prespecified. These are referred to as COLA reopenings. We lump these together with scheduled reopenings because there are too few of them to analyze separately.

which shows a cross-tabulation of these two contract provisions. Sixty-three percent of all contracts in our sample contain neither a COLA nor a reopener clause. The more prevalent type of flexibility is a COLA clause, which appears in 34% of the contracts. Scheduled reopening clauses appear in less than 5% of the contracts. Few contracts (less than one percent) contain both a COLA and a reopener clause.

4. Unbalanced Panel

The data consist of all major contract negotiations followed by the BLS from 1970 to 1995.[7] The BLS data provide a unique identification number for each BU, as well as a variable that tracks each negotiation. Together these two variables provide a panel structure to the data. The time period covered by different BUs can vary due to either changes in the scope of the coverage of major collective BUs by the BLS over the sample period, or changes in the size of the BU.

The nature of the sample implies that the number of contract negotiations that we observe for a specific BU is related to the average length of its contracts. For example, a BU that negotiates three year contracts can have a maximum of nine observations in the data. In contrast, a BU that negotiates one year contracts can have up to twenty-five observations. This not only leads to an unbalanced panel structure, but also implies that the length of each BU's panel is inversely related to the variable of interest.[8] Allowing each negotiation to have an equal weight in the estimation would result in over weighting short duration contracts relative to their importance at any point in time.[9] Table 3 shows the sample frequency, median, and mean contract duration for each panel length in the sample.

5. Econometric Framework

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[7]The data ends in 1995 due to the decision by the BLS to stop collecting data on major union contracts.

[8]The correlation between the contract duration and the number of contract negotiations by a BU is -0.33, which is statistically significant.

[9]Weighting each contract equally gives an overall mean duration of 32.8 months. Weighting each BU equally gives an overall mean contract duration of 33.8 months.

These four data considerations can be accommodated in the following econometric framework:

$$I_i = \begin{cases} 1 \text{ if the contracts for BU } i \text{ are censored} \\ 0 \text{ otherwise} \end{cases}$$

$$C_{it} = \begin{cases} 1 \ \text{ if the } t^{\text{th}} \text{contract for BU } i \text{ contains a COLA clause} \\ 0 \ \text{ otherwise} \end{cases}$$

$$R_{it} = \begin{cases} 1 \ \text{ if the } t^{\text{th}} \text{contract for BU } i \text{ contains a reopener clause} \\ 0 \ \text{ otherwise} \end{cases}$$

(1)

$$C^*_{it} = Z_{it}\gamma + \varepsilon_{cit}$$

$$R^*_{it} = M_{it}\delta + \varepsilon_{rit}$$

$$\left(\varepsilon_{cit}, \varepsilon_{rit}\right) \sim N(0, \Sigma) \quad \Sigma = \begin{bmatrix} 1 & \sigma_{cr} \\ \bullet & 1 \end{bmatrix}$$

$$C_{it} = \begin{cases} 1 \ \text{ if } C^*_{it} \geq 0 \\ 0 \ \text{ otherwise} \end{cases}$$

$$R_{it} = \begin{cases} 1 \ \text{ if } R^*_{it} \geq 0 \\ 0 \ \text{ otherwise} \end{cases}$$

$$D^*_{it} = X_{it}\beta + C_{it}\beta_c + R_{it}\beta_r + \varepsilon_{dit}$$

We estimate the parameter vectors (γ,δ) using a bivariate probit model. We replace the actual COLA and reopener indicators with their predicted values when estimating the duration specification.[10]

[10]We exclude the contract duration from the list of explanatory variables for the COLA and reopener specifications. See Appendix A for further discussion.

Assuming that ε_{dit} is distributed normally with mean zero and standard deviation σ_d, the contribution to the overall likelihood from the t^{th} contract for BU i is given by:

(2)
$$L_{it} = \phi\left(\frac{D_{it} - X_{it}\beta - \hat{C}_{it}\beta_c - \hat{R}_{it}\beta_r}{\sigma_d}\right)^{1-I_i} \bullet \left[\Phi\left(\frac{D^U - X_{it}\beta - \hat{C}_{it}\beta_c - \hat{R}_{it}\beta_r}{\sigma_d}\right) - \Phi\left(\frac{D^L - X_{it}\beta - \hat{C}_{it}\beta_c - \hat{R}_{it}\beta_r}{\sigma_d}\right)\right]^{I_i},$$

where $D^U = D_{it} + 6$(months) and $D^L = D_{it} - 6$(months).

To address the unbalanced nature of our panel, we set up the likelihood function so that each BU receives an equal weight. Let N_i denote the number of contract negotiations observed for the i^{th} BU. Then the contribution of this BU to the overall likelihood, L_i, is given by:

(3)
$$L_i = \prod_{t=1}^{N_i}\left(L_{it}^{1/N_i}\right)$$

III. Measuring Changes in Aggregate Uncertainty Over Time

This section outlines the methodologies used to construct time-varying measures of aggregate uncertainty. We initially discuss three alternative measures of inflation uncertainty, before turning our attention to the issue of structural-based estimates of nominal and real uncertainty.

1. Time-Varying Estimates of Inflation Uncertainty

While theoretical work has principally been concerned with the effects of aggregate nominal and real uncertainty on contract length, a number of empirical studies have narrowed the focus and investigated the relationship between inflation uncertainty and contract durations.

However, the lack of direct observations on inflation uncertainty has required researchers to adopt various methods to measure (and allow) shifts in the variance of inflation over time.

As one approach, Christofides and Wilton (1983) and Christofides (1985, 1990) specify an autoregressive process for inflation, and then reestimate the model sequentially by adding individual observations to the sample. The square of the standard error of the estimate from the sliding regressions is then used as a proxy for inflation uncertainty. This approach can be described as follows:

$$(4) \qquad \begin{aligned} \pi_{t+1} &= X_t \beta_t + \varepsilon_{t+1}, \\ \sigma_t^2(\pi) &= E[(\varepsilon_l^2)], \quad l = 1, 2, \ldots, t. \end{aligned}$$

where π_{t+1} is the inflation rate between period t and period t+1, X_t is a vector of explanatory variables (lagged inflation rates) available through period t, β_t is the coefficient vector of the model through period t, ε_{t+1} is the error term of the model, and $\sigma_t^2(\pi)$ is the measure of inflation uncertainty in period t.

Another approach is to examine survey data on price expectations. This provides direct measures of inflation expectations, which circumvents possible errors in specifying how people form their inflation expectations. Researchers have used the dispersion of forecasts across respondents as a proxy for the variance of inflation.[11] Vroman (1989) and Kanago (1998) adopt this procedure and construct a measure of inflation uncertainty based on the cross-sectional variance of predicted price changes from the Livingston survey. This approach can be described by:

$$(5) \qquad \sigma_t^2(\pi) = (1/N) \sum_{i=1}^{N} (\pi_{it}^j - \bar{\pi}_t^j)^2$$

[11] There is a large empirical literature that has employed forecast dispersion measures from survey series to quantify the effect of inflation uncertainty on aggregate economic activity. These include Cukierman and Wachtel (1979), Levi and Makin (1980), Mullineaux (1980) and Holland (1986, 1993).

where π^j_{it} is individual i's forecast of inflation between period t and period $t+j$, $\bar{\pi}^j_t$ is the consensus (average) inflation forecast across the N survey respondents at time t, and j is defined as the ratio of the forecasting horizon to the sampling interval of the survey data.[12]

The validity of using proxies based on equation (5) depends critically on the relationship between disagreement and uncertainty. Underlying this approach is the assumption that consensus among individuals' point predictions for inflation is indicative of a high degree of predictive confidence on the their part. It is important to note, therefore, that Pagan, Hall and Trivedi (1983) and Zarnowitz and Lambros (1987) argue that there is little reason to believe that inferences from the (observed) distribution of point predictions should be informative about the (subjective) probability distribution of possible outcomes used by individuals to generate their point predictions. Thus, the relationship between forecast dispersion and forecast uncertainty remains an open empirical question.

As an alternative to (5), Rich, Raymond and Butler (1992) propose a survey-based measure of inflation uncertainty that is directly linked to the predictability of the inflation process. Drawing on the work of Engle (1982, 1983), Rich, Raymond and Butler model the conditional variance of the consensus forecast errors an Autoregressive Conditional Heteroskedasticity (ARCH) process. Specifically, they consider the following two-equation system to generate a time-varying measure of inflation uncertainty:

$$\begin{aligned} E[\varepsilon_{t,j}|I_t] &= E[\pi_{t+j} - \bar{\pi}^j_t|I_t] = 0 \\ \sigma^2_t(\pi) &= E[\varepsilon^2_{t,j}|I_t] = h_{t,j} = \alpha_0 + \sum_{i=j}^{p} \alpha_i \varepsilon^2_{t-i,j} \end{aligned} \tag{6}$$

where π_{t+j} is the rate of inflation between period t and period $t+j$, $\varepsilon_{t,j}$ is the consensus inflation forecast error from the survey conducted in period t, I_t denotes an information set that includes

[12]Equation (5) accounts for possible overlapping data, where the forecast horizon of the survey series exceeds the sampling interval. Overlapping data results in the respondents' predictions acting as a multi-step-ahead forecast and a value of j greater than unity. While this point has no direct bearing on the construction of the forecast dispersion measure in (5), it is relevant for our subsequent discussion.

all information available through time t, and $h_{t,j}$ is the variance of $\varepsilon_{t,j}$ conditional on I_t.[13] The key feature of equation (6) is that it explicitly models the conditional variance process of the forecast errors, and thereby provides a more natural measure of inflation uncertainty from survey data on expected price changes.[14]

Figure 4 illustrates estimates of inflation uncertainty for the period 1969:Q2-1995:Q3 based on the approaches described in equations (4)-(6). The upper panel plots an estimate of inflation uncertainty using the sliding regression technique. For purposes of comparison, we also include a rolling variance measured with a five-year window with equal weights. The five-year rolling window is applied to the updated residuals from the sliding regressions.

The middle and lower panels of Figure 4 depict estimates of inflation uncertainty using measures of forecast dispersion and forecast uncertainty from the Survey Research Center (SRC) expected price change series. The survey series data are quarterly observations on the one-year CPI inflation forecasts (this implies that $j = 4$ in equation (6)). The measures of forecast dispersion and forecast uncertainty correspond, respectively, to the cross-sectional variance of the survey forecasts and an ARCH estimate of inflation uncertainty.[15]

Figure 4 illustrates that the behavior of the forecast dispersion and forecast uncertainty measures is broadly similar. Specifically, both rise rather dramatically with the food and oil price shocks of 1973-74, and then generally decline through the middle 1980s and into the early 1990s.

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[13]The formulation in (6) can accommodate the use of overlapping data. Overlapping data places restrictions on the specification of the ARCH process due to informational considerations, and does not preclude autocorrelation of the model's disturbance terms. Rich, Raymond and Butler (1992) provide details on how estimation of (6) can be conducted within a generalized method of moments framework.

[14]Kanago (1988) also proposes an ARCH-based measure of inflation uncertainty. However, he uses a reduced-form forecasting equation to model the conditional mean of the inflation process rather than a survey consensus estimate.

[15]Rich, Raymond and Butler (1992) find evidence of a positive and statistically significant relationship between the cross-sectional variance of the survey forecasts and the ARCH estimate of inflation uncertainty for the SRC survey series. We interpret this evidence as partial justification for the use of the forecast dispersion measure as a proxy for inflation uncertainty. Unlike Vroman (1989), our analysis does not employ data from the Livingston price expectations series because that survey is conducted on a semi-annual basis and does not coincide with the quarterly frequency of our data on contract durations.

There is an additional episode of low consensus coinciding with the second round of adverse supply shocks toward the latter part of the 1970s.

In contrast to the survey-based measures, the estimate of inflation uncertainty from the sliding regression technique trends slightly upward over the bulk of the sample period. The movements in the series reflect its inherent long-term memory process, where the estimate of inflation uncertainty is constructed as a weighted average of *all* past squared forecast errors.[16] As a consequence, the increased unpredictability of the inflation process during the 1970s and early 1980s leads to a gradual rise in its value over this time period. The increased predictability of inflation toward the latter part of our sample period only translates into a slight decline in its value.

While a formal evaluation of the merits of the inflation uncertainty measures is beyond the scope of this paper, casual observation suggests that the sliding regression technique may be problematic. This method yields a measure of inflation uncertainty that differs markedly from the survey-based measures and it displays little correspondence with identifiable events over the sample period. This likely reflects the weighting scheme used in the sliding regression technique, which appears to mask important variation in the predictability of the inflation process and generates excessive smoothness in the resulting estimate of inflation uncertainty.[17] The choice of the inflation uncertainty measure and its implications for the robustness of the results will be explored further in the discussion of the empirical findings.

2. Nominal Uncertainty, Real Uncertainty and Structural Vector Autoregressive Models

While the measures of inflation uncertainty in equations (4)-(6) differ in terms of their construction, there is an issue that concerns each proxy and its inclusion in empirical models of contract durations. Specifically, the previous discussion did not draw a distinction between aggregate nominal and real shocks. If contract length depends in different ways on the source of

[16]The weights are inversely related to the degrees of freedom in the regression equation.

[17]Becasue the rolling variance and ARCH estimates of inflation uncertainty depend on the more recent history of squared forecast errors, their movements are more responsive to changes in the predictability of the inflation process.

uncertainty in an economy, then any attempt to interpret the relationship between contract duration and inflation uncertainty is problematic due to the influence of both nominal and real shocks on the behavior of inflation. This consideration would not only be relevant for previous empirical work, but would also pertain to any study using aggregate uncertainty and inflation uncertainty interchangeably.

Due to the recognition of this problem, or because of their interest in providing an analysis closer in spirit to the theoretical literature, some empirical studies on contract durations have tried to incorporate nominal and real shocks in their analysis. For example, Wallace and Blanco (1991) use residual-based estimates of money supply shocks and industry-specific productivity shocks to construct proxies for nominal uncertainty and real uncertainty, respectively. Wallace (1999) extends this earlier work by including a measure of real uncertainty based on oil price shocks.[18] While these two papers account for different types of aggregate uncertainty, a potential drawback of their analysis is that the estimates are based on single variables that may be too narrow in scope to serve as reliable proxies. What is needed, therefore, is an estimation strategy that can be applied to a set of economic variables, and which offers a more appealing scheme for identifying nominal and real shocks.

The structural vector autoregression (SVAR) approach offers one such methodology. This approach uses restrictions based on economic theory to uncover the structural disturbances in an economy. Within the SVAR framework, Gali (1992) has proposed a four-variable system that allows for the identification of four structural shocks corresponding to an aggregate supply shock, an IS shock, a money demand shock, and a money supply shock. Specifically, Gali's model is based on the following reduced form VAR:

$$B(L)X_t = \mu_t \quad , \tag{7}$$

[18]Wallace and Blanco (1991) and Wallace (1999) adopt the sliding regressions technique to construct time-varying estimates of nominal uncertainty and real uncertainty.

where X is a vector containing the four variables of the system, $B(L) = I_n - B_1L - \ldots - B_pL^p$ is a pth-order lag polynomial matrix, I_n is the (4 x 4) identity matrix and $\mu_t = [\mu_{1t},\ldots,\mu_{4t}]'$ is the vector of reduced-form errors.[19]

The structural VAR representation is given by:

$$A(L)X_t = \varepsilon_t \quad , \tag{8}$$

where A_0 is nonsingular and normalized to have 1's down the main diagonal, $A(L) = A_0B(L)$, and $\varepsilon_t = A_0\mu_t$ is the (4 x 1) vector of structural disturbances to the system.

The ability to disentangle the structural shocks from the reduced-form errors hinges critically on identification restrictions that allow for the estimation of the matrix A_0. In contrast to other SVAR models that use a recursive system to achieve identification, Gali (1992) considers a combination of short- and long-run restrictions that focus on the effects of the structural shocks on particular variables.[20] The popularity of the Gali model partly stems from its nonrecursive structure. In addition, the model is capable of generating reasonable responses on the part of its variables to the structural shocks within a relatively low dimensional system.

For our investigation into the determinants of contract duration, the SVAR approach provides a particularly attractive framework to isolate different sources of uncertainty in the economy. The IS shock, money demand shock, and money supply shock can be combined into a composite aggregate demand shock to characterize the behavior of nominal disturbances, while the aggregate supply shock can be interpreted as a real shock. Following Friedman and Kuttner (1996), we use a rolling window procedure to obtain time-varying estimates of the variance of nominal and real shocks.

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[19]The four-variable system is comprised of real output growth, the change in the yield on 3-month Treasury bills, the *ex post* real return on 3-month Treasury bills, and the growth rate in real money balances (measured by M1).

[20]The normalization of the diagonal elements in A_0 as well as the assumption that the structural shocks are mutually uncorrelated provide a subset of the restrictions used for identification purposes. See Gali (1992) for further discussion of the identifying restrictions.

Figure 5 illustrates the estimates of nominal and real uncertainty for the period 1969:Q2-1995:Q3 from the Gali (1992) model. The moving-average variances are constructed using a five-year rolling window with equal weights. As is evident from the plots, the variances of these structural shocks do change over time, as well as relative to each other. Aggregate demand shocks became increasingly more variable in the late 1970s and early 1980s, and declined sharply at the outset of the current expansion. In the case of aggregate supply shocks, they were more variable during the early and mid-1970s, before declining and remaining fairly steady since the mid-1980s. The patterns in Figure 5 also imply that the uncertainty of nominal shocks relative to real shocks rose steadily and dramatically through the middle 1980s, before returning to levels comparable to those at the beginning of the sample period.

An important question concerns the relationship of real and nominal uncertainty to inflation uncertainty. To address this issue, we regressed our ARCH inflation uncertainty measure on the Gali aggregate demand and aggregate supply uncertainty measures. While the results indicate that inflation uncertainty is positively related to both aggregate demand (nominal) uncertainty and aggregate supply (real) uncertainty, only the latter yielded evidence of a statistically significant relationship.[21]

IV. Empirical Specification and Results

In this section, we discuss the additional explanatory variables which we use in estimating the econometric model outlined in section two. We begin with the BU's decision to adopt a COLA and/or a reopener clause in the current contract. We then examine to its decision regarding the desired duration of the contract.

1. COLA and Contract Reopeners

The theoretical work on COLA clauses has not explicitly modeled the bargaining process between a union and a firm. Rather, the literature has adopted the simplifying assumption that the outcome of this bargaining will be well approximated by an optimal risk-sharing model

[21] In the case of forecast dispersion, the evidence indicates that forecast dispersion is positively and statistically significantly related to both nominal and real uncertainty.

[Ehrenberg *et al* (1984) and Card (1986)]. There are several predictions from this approach that have guided the empirical literature.

The starting point in the modeling is to assume that there is a fixed cost to writing a COLA clause into the contract, which suggests that scale economies exist in the use of indexed contracts. While there are a number of ways to proxy for these scale economies, we will follow the literature and control for the size of the BU [Cousineau *et al* (1983), Ehrenberg *et al* (1984)].[22] To the extent that the costs of writing a COLA vary systematically by industry, the inclusion of industry fixed-effects may also be an effective way to capture this cost variation [Cousineau *et al* (1983) and Vroman (1989)]. Unions as well may develop some expertise in negotiating COLA clauses, which can be passed on to their local affiliates. In addition to controlling for industry fixed-effects, we also include union fixed-effects in our specification. Similarly, if a BU had a COLA in its prior contract, the cost of indexing the current contract is likely to be lower. In most cases, the basic structure of the COLA is maintained between contracts. This motivates the frequent use of a lagged COLA indicator as an instrument for current indexation [Vroman (1989) and Christofides (1990)].

Gray (1976) develops a model in which the optimal degree of indexation depends on the source of shocks to the economy. Aggregate demand shocks tend to move prices and spot market wages in the same direction. Indexation, then, helps to maintain the contract wage relative to the spot market wage in response to demand shocks. In contrast, aggregate supply shocks tend to move prices and spot market wages in opposite directions. For example, a negative supply shock will tend to push up prices, while lowering the marginal productivity of labor. Indexation will push the contract wage up when spot market wages are facing downward pressure. Gray's analysis suggests that the optimal degree of indexation is increasing in the variance of aggregate demand shocks relative to aggregate supply shocks. We test this prediction by separately controlling for nominal and real uncertainty.

Assuming that an indexed contract incorporates the optimal degree of indexation, Ehrenberg *et al* (1984) show that the decision to adopt a COLA depends on the degree of

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[22] Alternatives would be to scale by some measure of the size of the firm, or the number of BUs at the firm.

inflation uncertainty, but not on the expected inflation rate over the next contract. Ehrenberg *et al* (1984) and Murphy (1992) include both a proxy for expected inflation and inflation uncertainty to test this hypothesis. Vroman (1989) controls for the unexpected inflation rate over the prior contract and current inflation uncertainty. Cousineau *et al* (1983) and Christofides (1990) control for inflation uncertainty, but not the expected inflation rate. We will include proxies for both the expected inflation rate and inflation uncertainty. In addition, we will explore the sensitivity of the results to the choice of proxies for inflation uncertainty, and the robustness of the inflation uncertainty effect when additional proxies for nominal and real uncertainty are included.

A central prediction of the risk-sharing framework is that the decision to index and the degree of indexation should depend in part on the degree of correlation between the COLA index and the firm's output and input prices.[23] Card (1986) tests if the marginal elasticity of indexation in his sample of Canadian contracts varies with measures of both correlations. Hendricks and Kahn (1983), Ehrenberg *et al* (1984), Christofides (1990) and Murphy (1992) include measures of the correlation between the consumer price index and the producer price index for the firm's industry. A problem with using this control variable in our analysis is that producer price data only exists for some of the industries in our sample, and for some of the years covered by our sample. Since the literature has not allowed these correlations to be time-varying, their effects will be picked up by the industry fixed-effects.

Wage and price controls may affect a BU's decision to index a contract. Cousineau *et al* (1983) argue that in the case of the Canadian controls, parties have an incentive to index a contract that has an expiration date following the likely end date of the control period. Wage increases generated by the COLA clause after the control period ends would not be subject to review. This suggests that early phases of wage/price controls should have little if any impact on the decision to index a contract, but that indexation may rise as the expectation of the cessation of controls begins to set in. Vroman (1989) in her analysis of U.S. labor contracts includes indicators for the Nixon and Carter control periods.

[23]More specifically, indexation should depend on the correlation between the unexplained variance in the index and the unexplained variation in the firm's output and input prices.

Several researchers include controls for tightness in the labor market. The choice of the level of labor market aggregation to use varies across studies. Vroman (1989) uses the aggregate unemployment rate, Christofides (1990) uses a regional unemployment rate, while Murphy (1992) uses a state unemployment rate. We follow a hybrid of these approaches by including measures of tightness in both the local labor market (measured at the state level) and the national labor market (using the firm's industry as the reference point).

We control for both trend and cyclical conditions in the state and industry where the BU is located. We assume that the employment process follows a quadratic time trend, where we allow for seasonal employment effects and up to a second-order autocorrelation in the errors. We use BLS state employment data and national industry employment data measured at a quarterly frequency to estimate the parameters. Letting E_{it} denote the employment in state or industry i in period t and Q_{jt} denote quarterly seasonal dummy variables, we estimate the following:

$$
\begin{aligned}
&\log E_{it} = \beta_{i0} + \beta_{i1}t + \beta_{i2}t^2 + \sum_{j=1}^{3}\delta_{ij}Q_{jt} + \upsilon_{it}, \\
(9)\qquad &\upsilon_{it} = \rho_1\upsilon_{it-1} + \rho_2\upsilon_{it-2} + \varepsilon_{it} \\
&\varepsilon_{it} \sim N(0,\sigma_i^2)
\end{aligned}
$$

We proxy long-run employment trends in the state or industry by the implied employment growth, $\beta_{i1} + 2\beta_{i2}t$. The composite employment residual, ν_{it}, provides a proxy for cyclical conditions in the state or industry, with tighter labor market conditions represented by larger residuals.

We use the same set of control variables to model the decision to include a reopening clause in the contract. As in the case of a COLA clause, we use an indicator for a reopening clause in the prior contract as an instrument for the presence of a reopening clause. The fixed costs of writing a reopener clause are likely to be less than the fixed costs of writing a COLA clause. However, we expect that the presence of a reopener in the prior contract will still increase the likelihood of having a reopener in the current contract, since the parties are familiar with the process of negotiating under such a clause.

The results comparing different measures of expected inflation and the degree of inflation uncertainty on the decision to include an indexation and/or reopening clause in the contract are given in Table 4. The extent to which the indexation decision is sensitive to the source of aggregate uncertainty is explored in Table 5. Variable sources and summary statistics are provided in the Data Appendix.

The likelihood of indexation increases with both of our measures of expected inflation. Higher expected inflation, though, does not raise the likelihood of a reopener clause. In contrast, the inflation uncertainty results are sensitive to the methodology used to construct the uncertainty measure. The sliding regression methodology generates an uncertainty measure which has a significant negative effect on the likelihood of a COLA, a finding that is opposite to the prediction from the theoretical literature. Using the dispersion of inflation forecasts as the uncertainty measure produces a positive but insignificant relationship with the indexation decision. Finally, the ARCH approach produces a quantitatively similar effect as the dispersion approach, but is much more precisely estimated. The results using the survey-based measures of inflation uncertainty are consistent with those of Cousineau *et al* (1983), Ehrenberg *et al* (1984) and Christofides (1990). Increased inflation uncertainty as proxied by either the dispersion or ARCH measure also raises the likelihood that the contract will contain a scheduled reopening clause.

A BU which had a COLA in its prior contract is much more likely to have a COLA in its current contract, which is similar to the finding in Christofides (1990) using Canadian data. Likewise, a BU which had a reopening clause in its prior contract is much more likely to have a reopening clause in its current contract. Larger BUs are more likely to index, but there is no effect of size on the likelihood of a reopening clause. This is consistent with the view that there is some fixity to the costs of writing a COLA clause, and accords with the findings in Cousineau *et al* (1983) and Ehrenberg *et al* (1984). The industry and union fixed-effects are highly significant.

There is some evidence of cyclicality in the use of indexation and reopenings clauses. The probability that a contract is indexed increases when employment conditions are tight in the industry. The probability that a contract contains a scheduled reopening falls when employment conditions are tight in the local labor market. In addition, BUs that are organized in industries with higher trend employment growth are more likely to use indexed contracts.

We also explore Gray's (1976) hypothesis that the source of uncertainty should affect the degree of indexation. It is important to note, however, that we are limited to an indirect test because we can only observe the decision to index, and not the degree of indexation conditional on the decision to index. The first two columns of Table 5 replace inflation uncertainty with our measures of aggregate demand and supply uncertainty, with the data indicating that both sources of uncertainty lead to a higher likelihood of indexation. With the previous caveat in mind, the results do not support the prediction from Gray's (1976) model as the the standardized marginal effect from supply uncertainty is larger in magnitude than for demand uncertainty. The third and fourth columns add back in our ARCH proxy for inflation uncertainty. Controlling for aggregate demand and supply uncertainty, we find no independent evidence of a connection between inflation uncertainty and the indexation decision. Finally, controlling for all three measures of aggregate uncertainty we find no additional role for the expected inflation rate in predicting indexation, which supports the prediction from Ehrenberg *et al* (1984).

2. The Contract Duration Decision

The Gray (1978) model suggests that the two key determinants of the desired contract duration are the fixed costs of negotiating a contract and uncertainty. One difficulty in testing this hypothesis is that no direct measures of negotiation costs are available. A simple analysis of variance indicates that 15% of the variation in contract durations (treated as uncensored) is explained by a set of 2-digit industry fixed-effects.[24] One interpretation is that there is a strong industry-specific component to the variation in contracting costs across different BUs.[25]

Several alternative proxies for contracting costs have appeared in the literature. If some of the costs of negotiating contract renewals are fixed in nature, then there may exist important scale economies. Typically, these scale economies have been proxied by controlling for the size of the

[24]Adding union fixed-effects increases the explained variance to 19%.

[25]Several researchers have included controls for the demographic characteristics of the union workers in a particular industry. To the extent that these vary mainly across industries rather than within industries over time, these demographic effects will be picked up by the industry fixed-effects.

BU [Christofides and Wilton (1983), Christofides (1985), and Vroman (1989)].[26] To the extent that national unions provide assistance to their affiliates, there may be systematic differences in contracting costs across unions. We control for this by including a set of union fixed-effects.

One component of the costs of contract renewals is the associated costs of a strike if one occurs during the negotiations. Vroman (1989) controls for whether there is a strike in the current contract. She argues that a strike in the current contract will raise the likelihood of a strike in the subsequent contract renewal, thereby increasing the overall fixed costs associated with the current contract. This would suggest that, all else constant, the parties should agree to a longer contract. However, Schnell and Gramm (1987) using U.S. data find that the probability of a strike declines with the incidence of a strike in the prior contract negotiation. Card (1988) and Cramton, Gunderson and Tracy (1999) examine Canadian data and find that the impact of a current strike on the likelihood of a strike at the next negotiation varies with the length of the current strike. Short strikes (less than 28 days) are associated with a higher strike incidence at the contract renewal, while long strikes (more than 50 days) are associated with a lower strike incidence at the contract renewal. There is no clear connection, then, between the occurrence of a strike in the current and subsequent contract negotiations.

The other key aspect for testing the Gray (1978) model is controlling for uncertainty. Most of the literature has focused exclusively on using inflation uncertainty as a proxy for nominal uncertainty [Christofides and Wilton (1983), Christofides (1985), Vroman (1989) and Murphy (1992).[27] We will again consider the three alternative methodologies for inflation uncertainty previously examined in section three as well as our structural-based measures of aggregate nominal and real uncertainty. This allows us to conduct two important tests. First, we can determine if the effect of aggregate uncertainty on contract durations differs depending on the source of the uncertainty. Second, we can assess the extent to which inflation uncertainty captures the underlying level of nominal uncertainty.

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[26] Murphy (1992) controls for the percent of the firm's employment covered by the BU.

[27] As previously noted, exceptions are Wallace and Blanco (1991) and Wallace (1999) who attempt to control for both nominal and real uncertainty.

Christofides (1985) argues that a BU's desired contract duration will depend on whether the contract is negotiated during a period of wage and price controls. If there is a perception that the controls will lapse in the near future, then there is an incentive for the parties to negotiate a shorter contract that will expire following the end of the controls. This will permit the parties at the contract renewal to reset the terms without the controls imposing any constraints. Christofides (1985) tests this prediction using Canadian contract data, while Vroman (1989) and Wallace and Blanco (1991) investigate this hypothesis using U.S. contract data.

We also control for conditions in the industry and local labor markets. As we discussed earlier, we control for trend employment growth in the industry and state, as well as for the degree to which current employment in the industry and state is above or below its trend. Vroman (1989) controls for the overall tightness in the labor market using the aggregate unemployment rate. Murphy (1992) controls for the employment growth rate in the BU's industry in the year prior to the contract renewal.

The impact of inflation uncertainty on contract durations is explored in Table 6. As was evident by the results reported in Table 4 regarding the decisions to index or include a reopening clause, the results are not robust to the choice of the inflation uncertainty measure. Moreover, the disparity in the findings is along the same lines. Specification (1) uses the sliding regression methodology and provides no evidence of a significant relationship between inflation uncertainty and contract durations. In contrast, both measures derived from the SRC survey data generate inflation uncertainty estimates that have a negative and significant relationship with desired contract durations. The dispersion measure displays a larger standardized impact than the ARCH measure. There is clear evidence, then, of a negative relationship between inflation uncertainty (appropriately measured) and contract durations.[28]

In specification (4) of Table 6 we substitute our two aggregate uncertainty measures for the inflation uncertainty measure. Consistent with the Gray (1978) model and in contrast to the

[28]Our results contrast with Kanago (1988) who analyzes the effects of various inflation proxies on contract durations in a sample of 1,000 large contracts expiring during the period from 1954 to 1980. He finds that an ARCH measure of inflation uncertainty is positively related to contract durations, while an SEE measure and a mean square error of inflation estimates from the SCR survey are negatively related to contract durations.

predictions from Danziger (1988), we find that both nominal and real uncertainty are associated with shorter contract durations. The real uncertainty effect is slightly larger in absolute value than the nominal uncertainty effect, while both are larger than the previous inflation uncertainty effects. When we extend the model to include the ARCH measure of inflation uncertainty [specification (5) of Table 6], the coefficients on the aggregate uncertainty measures are relatively unaffected, while the coefficient on inflation uncertainty is no longer significant.

The aggregate demand uncertainty measure used in specification (4) is a composite built up from three underlying components: an IS shock, a money demand shock, and a money supply shock. Table 7 explores the question of whether the negative effect of the composite aggregate demand uncertainty is being driven primarily by any of its underlying components. The data indicate a negative, significant and roughly equal effect of uncertainty over each of the three components on desired contract durations.

Our findings also contrast with Wallace (1999). He constructs a proxy for aggregate real uncertainty using oil price shocks and the sliding regression technique. As in Wallace and Blanco (1991), he controls for nominal uncertainty using a proxy based on money supply shocks.[29] In a sample of manufacturing contracts ending in 1980, he finds that contract durations have a positive and significant connection with real uncertainty, and a negative and significant connection with nominal uncertainty. When we construct estimates of the Wallace (1999) real and nominal uncertainty measures and incorporate them into our sample and specification, we find that both measures of aggregate uncertainty have a positive and significant effect on contract durations. This contrast with our findings underscores the importance of the methodology used to construct the aggregate uncertainty measures.

A prediction from some models is that after controlling for the uncertainty over future inflation, the expected inflation rate should have no independent influence on contract durations. The impact of the expected inflation rate on the desired duration is again sensitive to how we

[29]Specifically, Wallace and Blanco (1991) use a nominal uncertainty measure derived from a sliding regression of M1 on its lags and seasonal dummy variables. We estimated two uncertainty measures from their model. The first averages the squared residuals over the entire sliding estimation period, while the second averages the squared residuals over the most recent five years. The first measure had a *positive* and significant coefficient in the duration equation, while the second measure had an insignificant coefficient in the duration equation.

control for inflation uncertainty. When we use the dispersion measure of inflation uncertainty, the expected rate of inflation has no impact on the desired contract duration. However, when we use the ARCH measure of inflation uncertainty, periods of higher expected inflation are associated with shorter contract durations. Adding the aggregate nominal and real uncertainty measures to the ARCH inflation uncertainty measure does not eliminate the effect of the expected rate of inflation on the desired length of contracts. However, when we control for aggregate demand uncertainty using either our proxy for money supply uncertainty or IS uncertainty (as opposed to the composite measure), the expected inflation rate has no additional impact on contract durations.

Contracts which build in flexibility either through indexation or reopening provisions are associated with longer desired contracts. The effect of a reopening provision is twice the magnitude of a COLA clause, though less precisely estimated. The data strongly reject for all specifications the hypothesis that the COLA and reopening clauses are exogenous.

Indexation and reopening clauses may also affect the sensitivity of the desired contract duration to changes in uncertainty. To explore this issue, we interact the COLA and scheduled reopening indicators with the ARCH inflation uncertainty proxy. We find that the negative effect of increased inflation uncertainty on desired contract durations is entirely driven by the non-indexed contracts.[30] Contracts with schedule reopening clauses have the same sensitivity as contracts without COLAs and reopenings. We construct similar interactions with the aggregate demand and supply uncertainty proxies. Here we find that the durations of indexed contracts are less responsive to aggregate demand and supply uncertainty, but we can reject the hypothesis that the duration of indexed contracts does not respond to aggregate uncertainty. In contrast, contracts with scheduled reopenings have durations that are significantly more sensitive to aggregate demand and supply uncertainty.

The industry and union fixed-effects are jointly significant in every specification. In contrast to the decision to adopt a COLA clause, there appears to be no significant scale

[30]Christofides and Wilton (1983) can measure not only whether a contract is indexed, but also the degree of indexation among indexed contracts. They find that the duration response to their inflation uncertainty proxy is larger in absolute magnitude for nonindexed contracts, and that it diminishes with the contract's degree of indexation.

economies in negotiating contracts associated with larger BUs. Holding the size of the BU constant, we also test for whether BUs which have experienced membership declines over the prior contract period prefer longer or shorter contracts. One hypothesis is that BUs faced with membership declines will have a stronger preference for job security protections to be written into the current contract. This gives the union an incentive to write a longer contract in order to extend the life of these job protections. The data suggest that larger membership losses over the prior contract do lengthen the current contract, though the magnitude of this effect is quite small. Following up on Vroman's (1989) idea of controlling for likely strike costs, when we add an indicator for a short strike in the current contract negotiation it has a negative and insignificant effect in all of the specifications.

Desired contract durations also vary with cyclical and trend conditions in the industry and local labor markets. When conditions in the industry are above trend, BUs favor longer contracts. Cyclical conditions in the local labor market have no effect on desired contract durations. This is broadly consistent with Vroman (1989) who finds that tighter aggregate labor market conditions as measured by the aggregate unemployment rate lead to longer contracts. In contrast to the cyclical effects, stronger trend employment growth in the industry or the local labor market is associated with shorter contract durations.

Finally, the data indicate that contract durations were shorter during the Nixon wage and price controls, with the effect being much stronger during Phase II than during the later phases. There is no systematic evidence that the Carter wage and price controls led to shorter contracts. Vroman (1989) reports lower contract durations in her sample for both the Nixon and Carter control periods, with the strongest effects occurring during the Nixon Phase III & IV periods.

Appendix Table A1 explores the sensitivity of the results to the issues of weighting, censoring, the definition of the contract duration, and the sample coverage. Specification (1) repeats the findings from specification (3) of Table 6 as a baseline. Specification (2) of Table A1 mirrors specification (1) except that we weight each contract equally in the estimation. Specification (3) continues to equally weight all contracts and relaxes the assumption that most observed contract durations are censored. Specification (4) changes the definition of the contract duration to be from the effective date of the contract to its planned expiration date.[31] Finally, specification (5) limits the sample to manufacturing contracts negotiated prior to 1981, which restricts our sample to roughly the coverage and time period used by Wallace and Blanco (1991).

What emerges from the exercise is that the weighting, censoring and dating decisions all have meaningful impacts on some of the estimated coefficients. However, the coverage and time period used in the estimation can have a particularly important impact on the inflation uncertainty finding. Restricting the sample to manufacturing contracts negotiated prior to 1981 reduces the magnitude of the real uncertainty effect by almost fifty percent, and the nominal uncertainty measure no longer has a significant negative effect on contract durations.[32] This underscores the important role played by the changes in contract durations that emerged in the mid-1980s in identifying the aggregate uncertainty effects.

V. Conclusions

The behavior of labor contracts in practice is an empirical question. Using different proxies for inflation uncertainty and contract data from different countries, early research built a case for a robust negative relationship between inflation uncertainty and contract durations [Christofides and Wilton (1983), Christofides (1985), Vroman (1989) and Murphy (1992)]. On the other hand, Wallace and Blanco (1991) argued that there was no significant relationship between U.S. contract durations and inflation uncertainty.

In order to resolve this issue, we consider several different measures of inflation uncertainty using a common data set and specification. One such measure is based on the sliding regression technique which has attained considerable popularity in this literature. On closer inspection, however, we suggest that this proxy may be problematic due to the nature of its construct. Moreover, the measure yields no evidence of a significant link between inflation uncertainty and contract durations, and it suggests an inverse relationship between inflation

[31] The expected inflation and ARCH inflation uncertainty in this case are merged in based on the effective date rather than the negotiation date.

[32] When we substitute the Wallace (1999) aggregate uncertainty measures and restrict the sample to negotiations prior to 1981, the nominal uncertainty measure [money supply] remains positive and significant, while the real uncertainty measure [oil prices] becomes negative and insignificant.

uncertainty and the likelihood that a bargaining unit will adopt a COLA and/or a reopening clause. The latter findings seem particularly difficult to reconcile with theory.

We also consider survey-based measures of inflation uncertainty. Among the two measures, we argue that our ARCH uncertainty measure is the preferred proxy, and find that reductions in inflation uncertainty based on this measure are associated with longer labor contracts. The nature and features of this empirical relationship emerge in a much clearer fashion when contracts from the mid-eighties and early nineties are included in the sample.

Because inflation uncertainty reflects the influence of both real and nominal shocks, we also investigated the response of contract durations to changes in real and nominal uncertainty. Risk-sharing models such as Danziger (1988) imply that contract length may increase in response to greater real uncertainty. We provide a direct test of this hypothesis by using structural-based estimates of supply (real) and demand (nominal) uncertainty. Consistent with Gray's (1978) model, we find that both sources of aggregate uncertainty have a significant negative relationship with contract durations. Taken together, this evidence suggests that labor contract durations are endogenous to the economic environment prevailing at the time they are negotiated, but that risk-sharing concerns are not paramount.

References

Blanchard, Oliver, and Danny Quah. "The Dynamic Effects of Aggregate Demand and Supply Disturbances." *American Economic Review* 79 (September 1989): 655-673.

Bloomberg News. "Union Members Ratify Contract." *Bloomberg News*, 28 September 1999.

Canzoneri, Matthew B. "Labor Contracts and Monetary Policy." *Journal of Monetary Economics* 6 (April 1980): 241-255.

Card, David. "An Empirical Model of Wage Indexation Provisions in Union Contracts." *Journal of Political Economy* 94 (June 1986): S144-S175.

------. "Longitudinal Analysis of Strike Activity." *Journal of Labor Economics* 6 (April 1988): 147-176.

Christofides, Louis N. "The Impact of Controls on Wage Contract Durations." *The Economic Journal* 95 (March 1985): 161-168.

------. "The Interaction Between Indexation, Contract Duration and Non-Contingent Wage Adjustment." *Economica* 57 (August 1990): 395-409.

------, and David A. Wilton. "The Determinants of Contract Length." *Journal of Monetary Economics* 12 (August 1983): 309-319.

Cousineau, Jean-Michel, Robert Lacroix, and Banielle Bilodeau. "The Determination of Escalator Clauses in Collective Agreements." *Review of Economics and Statistics* 65 (May 1983): 196-202.

Cramton, Peter, Morley Gunderson, and Joseph Tracy. "The Effect of Collective Bargaining Legislation on Strikes and Wages." *Review of Economics and Statistics*, August 1999.

Cukierman, Alex, and Paul Wachtel. "Differential Inflationary Expectations and the Variability of the Rate of Inflation: Theory and Evidence." *American Economic Review* 69 (September 1979): 595-609.

Danziger, Leif. "Real Shocks, Efficient Risk Sharing, and the Duration of Labor Contracts." *Quarterly Journal of Economics* 103 (May 1988): 435-440.

Ehrenberg, Ronald G., Danziger Leif, and San Gee. "Cost-of-Living Adjustment Clauses in Union Contracts." In *Research in Labor Economics*, edited by Ronald G. Ehrenberg, 1-63. Stamford, CT, JAI Press, 1984.

Engle, Robert F. "Autoregressive Conditional Heteroskedasticity with Estimates of the Variance of United Kingdom Inflation." *Econometrica* 50 (July 1982): 987-1007.

------. "Estimates of the Variance of U.S. Inflation Based Upon the ARCH Model." *Journal of Money, Credit, and Banking* 15 (August 1983): 286-301.

Fischer, Stanley. "Long-Term Contracts, Rational Expectations, and the Optimum Money Supply Rule." *Journal of Political Economy* 85 (February 1977): 191-206.

Friedman, Benjamin M., and Kenneth N. Kuttner. "A Price Target for U.S. Monetary Policy? Lessons from the Experience with Money Growth Targets." *Brookings Papers on Economic Activity* 1 (1996): 77-125.

Gali, Jordi. "How Well Does the IS-LM Model Fit Postwar U.S. Data?" *Quarterly Journal of Economics* 107 (May 1992): 709-738.

Gamboa, Glenn. "Goodyear Union Members in Fayetteville Approve Contract, Return to Work." *Akron Beacon Journal*, 8 May 1997.

Gray, Jo Anna. "Wage Indexation: A Macroeconomic Approach." *Journal of Monetary Economics* 2 (April 1976): 221-35.

------. "On Indexation and Contract Length." *Journal of Political Economy* 86 (February 1978): 1-18.

Heckman, James. "Dummy Endogenous Variables in a Simultaneous Equation System." *Econometrica* 46 (July 1978): 931-959.

Hendricks, Wallace E., and Lawrence M. Kahn. "Cost-of-Living Clauses in Union Contracts: Determinants and Effects." *Industrial and Labor Relations Review* 36 (April 1983): 447-460.

Holland, Steven A. "Wage Indexation and the Effect of Inflation Uncertainty on Unemployment: An Empirical Analysis." *American Economic Review* 76 (March 1986): 235-243.

------. "Uncertain Effects of Money and the Link Between the Inflation Rate and Inflation Uncertainty." *Economic Inquiry* 31 (January 1993): 39-51.

Hyde, Justin. "UAW Gives Ford Right to Spin Off Visteon Plants - Not Workers." *The Associated Press*, October 14 1999.

Kanago, Bryce. "An Examination of the Duration of Wage Contracts." Diss. Iowa City, Iowa, Univeristy of Iowa, 1988.

------. "The Relation Between Contract Duration and Inflation Uncertainty: Further Evidence." *Journal of Macroeconomics* 20 (Fall 1998): 811-819.

Levi, Maurice D., and John H. Makin. "Inflation Uncertainty and the Phillips Curve: Some Empirical Evidence." *American Economic Review* 70 (December 1980): 1022-1027.

Mullineaux, Donald J. "Unemployment, Industrial Production, and Inflation Uncertainty in the United States." *Review of Economics and Statistics* 62 (May 1980): 163-169.

Murphy, Kevin J. "Determinants of Contract Duration in Collective Bargaining Agreements." *Industrial and Labor Relations Review* 45 (January 1992): 352-365.

Pagan, Adrian R., Alistar D. Hall, and Pravin K. Trivedi. "Assessing the Variability of Inflation." *Review of Economic Studies* 50 (October 1983): 585-596.

Phillips, David, and Peronet Despeignes. "Costs May Outweigh UAW Gains: Automakers Wait for Results of Four-Year Contracts With Union." *The Detroit News*, 27 October 1999.

Rich, Robert W., Jennie E. Raymond, and J.S. Butler. "The Relationship Between Forecast Dispersion and Forecast Uncertainty: Evidence From a Survey-Data ARCH Model." *Journal of Applied Econometrics* 7 (April 1992): 131-148.

Schnell, John F., and Cynthia L. Gramm. "Learning by Striking: Estimates of the Teetotaler Effect." *Journal of Labor Economics* 5 (April 1987): 221-241.

Taylor, John B. "Aggregate Dynamics and Staggered Contracts." *Journal of Political Economy* 88 (February 1980): 1-23.

Thomson, Linda. "Kennecott Workers Ratify Pact." *The Desert News*, 25 October 1996.

Vroman, Susan. "Inflation Uncertainty and Contract Duration." *Review of Economics and Statistics* 71 (November 1989): 677-681.

Wallace, Frederick H. "The Effects of Shock Size and Type on Labor Contract Duration." Working Paper. Appalachian State University, April, 1999.

------, and Hermino Blanco. "The Effects of Real and Nominal Shocks on Union-Firm Contract Duration." *Journal of Monetary Economics* 27 (June 1991): 361-380.

Zarnowitz, Victor, and Louis A. Lambros. "Consensus and Uncertainty in Economic Predictions." *Journal of Political Economy* 95 (June 1987): 591-621.

Appendix A: Econometric Specification Issues

In this appendix we provide additional details regarding the econometric specification we estimate. A natural framework is to allow for contract durations to depend on indexation/reopening clauses and vice versa. Our specification only allows the duration decision to depend on the presence or absence of indexation and/or reopening clauses. There is no feedback between the desired contract duration and the decisions to include these clauses in the contract. The rationale for this restriction is detailed below.

Heckman (1978) discusses the following general framework.

$$y_{1i}^* = x_{1i}\alpha_1 + d_i\beta_1 + y_{2i}^*\gamma_1 + u_{1i}$$
$$y_{2i}^* = x_{2i}\alpha_2 + d_i\beta_2 + y_{1i}^*\gamma_2 + u_{2i}$$
$$d_i = \begin{matrix} 1 & \text{iff } y_{2i}^* > 0 \\ 0 & \text{otherwise} \end{matrix}$$

where the *y* variables are potentially unobserved latent variables, and the *x* variables are assumed to be exogenous. Heckman shows that this is a valid statistical framework if and only if $\gamma_2\beta_1 + \beta_2 = 0$.

Consider a simplified version of our model. Ignoring the issue of censoring of some contract durations and the presence of reopening provisions, we can write our basic specification into Heckman's framework.

$$D_{it} = X_{1it}\alpha_1 + C_{it}\beta_1 + C_{it}^*\gamma_1 + u_{1i}$$
$$C_{it}^* = X_{2it}\alpha_2 + C_{it}\beta_2 + D_{it}\gamma_2 + u_{2i}$$
$$C_{it} = \begin{matrix} 1 & \text{if } C_{it}^* > 0 \\ 0 & \text{otherwise} \end{matrix}$$

Here we assume that the contract duration is observed.

Consider first the contract duration equation. In the general framework, both the presence of a COLA clause as well as the underlying latent variable affecting the indexation decision are allowed to influence the contract duration decision. In this specific application, it is reasonable to allow the presence of a COLA to affect the length of a contract since it provides additional flexibility ($\beta_1 \neq 0$). However, there is little reason to believe that variation in the latent variable C^* would have any independent influence on the contract duration ($\gamma_1 = 0$).

Turn now to the specification of the latent variable governing the indexation decision. There is little justification in this context to assume that a COLA clause would shift the latent variable C^*. This suggests that we set β_2 to zero. Heckman's restriction is that $\gamma_2\beta_1 + \beta_2 = 0$. For this restriction to be satisfied, we drop *D* from the equation (set $\gamma_2 = 0$).

Data Appendix B

Data Definitions and Sources	*Summary Statistics* Mean	Standard Deviation
Sliding Regression Technique. Inflation is measured as the quarterly growth (at an annual rate) in the consumer price series *PCU*. The sliding regressions are initially estimated using a sample covering the period 1966:Q2-1969:Q3.		
Expected Inflation The expected inflation series is computed according to equation (5), where the conditional mean of the inflation process is specified as an AR(4) process.	5.79	2.96
Inflation Uncertainty The inflation uncertainty measure is computed according to equation (5) as the square of the standard error of the estimate.	3.86	0.46
The Survey Research Center (SRC) Expected Price Change Series. The SRC survey is conducted at the University of Michigan and actually begins in 1948, but has undergone changes in both the sampling interval and in the form of the survey questionnaire. We restrict our attention to post-1966:Q2 data because this corresponds to a transition in the survey to complete closed-end quantitative data. In addition, the survey made a transition from quarterly data to monthly data in 1977. Therefore, we examine data corresponding to the months of February, May, August and November to maintain consistency in the dates of the surveys. Further details on the SRC survey series are provided in Juster and Comment (1980) and Rich, Raymond and Butler (1992).		
Expected Inflation Measures of expected inflation are forecasts of CPI inflation over the next year.	6.08	2.13
Inflation Uncertainty - Forecast Dispersion Measure The forecast dispersion measure is computed according to equation (6) as the cross-sectional variance of the inflation forecasts.	56.86	23.64

Data Definitions and Sources	*Summary Statistics*	
	Mean	Standard Deviation
Inflation Uncertainty - ARCH Measure The inflation uncertainty measure is computed according to equation (7). The forecast error is constructed as the difference between the annual rate of change in the consumer price series *PCU* and the consensus inflation forecast. The ARCH specification includes the fourth lagged squared forecast error and is estimated using data from the 1968:Q2-1995:Q3 surveys.	3.42	2.07
3. *Gali (1992) Model*. All data are quarterly and the estimation is conducted over the sample period 1959:Q1-1995:Q4.		
Output The output series *GDPH* is measured as real gross domestic product in chain-weighted 1992 dollars.		
Prices The price data are a quarterly average of the monthly consumer price series *PCU* for all urban consumers.		
Interest Rates The interest rate data are a quarterly average of the monthly series *FTBS3* which represents the yield on three-month Treasury Bills.		
Money Stock The data on the money stock are for M1 and are a quarterly average of the series *FM1*. The measures of M1 prior to 1959 are taken from the *Federal Reserve Bulletin*.		
Aggregate Real and Nominal Uncertainty The measure of real uncertainty is computed as a rolling variance of the aggregate supply shock with a five-year window with equal weights. The measure of nominal uncertainty is computed as a rolling variance of a composite aggregate demand shock with a five-year window with equal weights.		
Aggregate Supply	0.98	0.48
Aggregate Demand	3.76	1.80
IS Shock	1.40	0.67
Money Demand Shock	1.40	0.83
Money Supply Shock	1.19	0.45

Data Definitions and Sources	*Summary Statistics* Mean	Standard Deviation
Contract Duration The uncensored duration is measured as the number of months between the negotiation date of the contract and the expected expiration date. Source: BLS Major Agreement file.	32.74	9.11
Cost-of-Living Clause Indicator taking a value of one if the current contract contains a cost-of-living adjustment clasue. Source: BLS Major Agreement file.	0.35	0.48
Reopening Clause Indicator taking a value of one if the current contract contains either one or more scheduled reopening clauses or a cost-of-living reopening clause. Source: BLS Major Agreement file.	0.04	0.21
Bargaining Unit Size Log of the number of workers covered by the labor contract. Source: BLS Major Agreement file.	7.94	0.93
Percent Reduction in Bargaining Unit Size Maximum of zero and the percent change in the size of the bargaining unit between the current contract and the prior contract. Source: BLS Major Agreement file.	9.22	15.26
State Employment Residual Residual from a regression of quarterly log state private sector nonagricultural employment on a quadratic trend and quarterly seasonal indicators. See equation (10) in text. Source: BLS *Employment and Earnings: States & Areas* data.	−0.002	0.028
State Employment Trend Trend growth rate of quarterly state employment implied from equation (10) in text. Source: BLS *Employment and Earnings: States & Areas* data.	0.004	0.002

Data Definitions and Sources	*Summary Statistics*	
	Mean	Standard Deviation
Industry Employment Residual Residual from a regression of quarterly log 2-digit industry employment on a quadratic trend and quarterly seasonal indicators. See equation (10) in text. Source: BLS *Employment and Earnings: National* data.	0.002	0.005
Industry Employment Trend Trend growth rate of quarterly 2-digit industry employment implied from equation (10) in text. Source: BLS *Employment and Earnings: National* data.	−0.002	0.066
Wage & Price Controls:		
Nixon Phase II Indicator taking a value of one for contracts negotiated between 1 August 1971 and 31 December 1972.	0.00	0.02
Nixon Phase III & IV Indicator taking a value of one for contracts negotiated between 1 January 1973 and 30 April 1974.	0.02	0.15
Carter Phase I Indicator taking a value of one for contracts negotiated between 1 November 1978 and 31 December 1979.	0.05	0.22
Carter Phase II Indicator taking a value of one for contracts negotiated between 1 January 1980 and 31 August 1980.	0.05	0.21

Table 1. Comparison of Timing Between Negotiation, Effective, and Lag Expiration Dates

	Difference in Quarters								
Reference Date	−4+	−3	−2	−1	0	1	2	3	4+
Lag Expiration (Dur_1)	4.8	2.0	4.3	22.6	54.7	4.2	2.0	1.3	4.1
Effective (Dur_2)	1.6	1.2	3.1	13.9	74.8	4.3	0.6	0.3	0.2

Notes: Figures represent the percent of the sample where the negotiation date and the reference date differ by the indicated number of calendar quarters.

Table 2. Distribution of COLA and Reopener Clauses

		Reopener Clause	
		No	Yes
COLA Clause	No	61.3	4.7
	Yes	33.3	0.7

Notes: Numbers indicate the percent of contracts which have the indicated combination of COLA and reopener clauses.

Table 3. Panel Lengths and Contract Durations

Panel Length	% of Sample	% of BUs	Median Duration	Mean Duration
1	1.37	7.67	35	37.8
2	1.75	4.89	35	33.7
3	5.30	9.87	35	35.2
4	9.90	13.83	36	34.3
5	8.61	9.61	35	33.5
6	13.05	12.14	35	33.4
7	24.63	19.65	35	35.3
8	19.81	13.83	35	33.2
9	5.98	3.71	26.5	28.6
10	4.53	2.53	23	26.1
11	2.49	1.26	24	23.8
12+	2.58	1.01	12	16.0

Notes: Sample restricted to the set of contracts used in the estimation.

Table 4. Bivariate Probit - COLA and Reopener Clause, Expected Inflation and Inflation Uncertainty

Variable	COLA	Reopener	COLA	Reopener	COLA	Reopener
Lag COLA	2.093**	−0.132	2.029**	−0.160	2.040**	−0.144
	(0.060)	(0.100)	(0.058)	(0.103)	(0.058)	(0.101)
Lag reopener	0.040	1.117**	0.047	1.111**	0.043	1.123**
	(0.128)	(0.104)	(0.128)	(0.105)	(0.129)	(0.105)
Log BU Size	0.098**	−0.006	0.106**	−0.002	0.105**	−0.003
	(0.024)	(0.039)	(0.024)	(0.039)	(0.024)	(0.039)
Expected inflation:						
• Regression forecast	0.219**	0.003				
	(0.041)	(0.056)				
• SRC survey concensus			0.277**	−0.051	0.302**	0.044
			(0.077)	(0.091)	(0.039)	(0.051)
Inflation uncertainty:						
• Regression SEE	−0.240**	−0.087				
	(0.035)	(0.050)				
• SRC - dispersion of forecasts			0.064	0.150*		
			(0.067)	(0.083)		
• SRC - ARCH					0.074**	0.080**
					(0.025)	(0.029)
State employment residual	−0.022	−0.114**	−0.013	−0.112**	−0.010	−0.113**
	(0.025)	(0.032)	(0.025)	(0.031)	(0.025)	(0.032)
State employment trend	0.037	0.077**	0.042	0.077**	0.038	0.075**
	(0.026)	(0.033)	(0.027)	(0.033)	(0.027)	(0.033)
Industry employment residual	0.080**	−0.020	0.100**	−0.020	0.105**	−0.016
	(0.024)	(0.031)	(0.024)	(0.031)	(0.024)	(0.030)
Industry employment trend	0.060	0.155	0.289**	0.201	0.259**	0.165
	(0.116)	(0.160)	(0.120)	(0.165)	(0.120)	(0.165)
Nixon phase II	−0.086	0.835	0.973	1.284*	0.982	1.258
	(0.678)	(0.774)	(0.673)	(0.763)	(0.664)	(0.769)
Nixon phase III & IV	−0.528**	0.129	−0.109	0.233	−0.189	0.156
	(0.191)	(0.208)	(0.165)	(0.185)	(0.167)	(0.185)
Carter phase I	−0.515**	0.242	−0.792**	0.082	−0.728**	0.181
	(0.133)	(0.168)	(0.139)	(0.175)	(0.140)	(0.175)
Carter phase II	−0.089	−0.073	−0.146	−0.203	−0.099	−0.137
	(0.145)	(0.213)	(0.125)	(0.176)	(0.127)	(0.178)
Rho	−0.307**		−0.295**		−0.306**	
	(0.062)		(0.062)		(0.062)	

Notes: Standard errors are given in parentheses and have been adjusted for any nonindependence across contracts in the same BU. Sample size is 6,524. Continuous RHS variables have been standardized to have a zero mean and unit standard deviation. All specifications control for 2-digit industry fixed effects and 43 union fixed effects.

** significant at the 5% level

* significant at the 10% level

Table 5. Bivariate Probit - COLA and Reopener Clause, Aggregate Demand and Supply Uncertainty

Variable	COLA	Reopener	COLA	Reopener
Expected inflation:				
• SRC survey consensus	0.081*	0.076	0.070	0.045
	(0.045)	(0.058)	(0.045)	(0.058)
Uncertainty:				
• Aggregate demand	0.148**	0.093*	0.145**	0.087
	(0.040)	(0.053)	(0.040)	(0.053)
• Aggregate supply	0.527**	0.109	0.518**	0.087
	(0.054)	(0.072)	(0.054)	(0.073)
• Inflation, SRC - ARCH			0.031	0.078**
			(0.026)	(0.029)

Notes: Standard errors are given in parentheses and have been adjusted for any nonindependence across contracts in the same BU. Sample size is 6,524. Continuous RHS variables have been standardized to have a zero mean and unit standard deviation. All specifications control for 2-digit industry fixed effects and 43 union fixed effects.

** significant at the 5% level

* significant at the 10% level

Table 6. Contract Durations - Inflation, Aggregate Demand & Supply Uncertainty

Variable	(1)	(2)	(3)	(4)	(5)
COLA[1]	1.17*	1.43**	1.31**	1.65**	1.65**
	(0.062)	(0.63)	(0.62)	(0.60)	(0.60)
Reopener[1]	3.23	3.23	3.81	3.99*	4.18*
	(2.43)	(2.44)	(2.36)	(2.35)	(2.37)
Log BU Size	0.06	0.06	0.07	0.06	0.06
	(0.19)	(0.19)	(0.18)	(0.18)	(0.19)
Percent change in BU size	0.18	0.17	0.16	0.22*	0.22*
(negative changes)	(0.13)	(0.13)	(0.13)	(0.13)	(0.13)
Expected inflation:					
• Regression forecast	−1.15**				
	(0.29)				
• SRC survey consensus		−0.03	−1.11**	−0.61*	−0.59*
		(0.39)	(0.29)	(0.30)	(0.30)
Inflation uncertainty:					
• Regression SEE	−0.03				
	(0.19)				
• SRC - dispersion of forecasts		−1.25**			
		(0.40)			
• SRC - ARCH			−0.33**		−0.12
			(0.11)		(0.12)
Aggregate uncertainty:					
• Aggregate demand				−2.05**	−2.03**
				(0.34)	(0.34)
• Aggregate supply				−2.66**	−2.62**
				(0.35)	(0.36)

Table 6. Continued

Variable	(1)	(2)	(3)	(4)	(5)
State employment residual	−0.13	−0.22	−0.15	−0.01	−0.02
	(0.13)	(0.14)	(0.13)	(0.13)	(0.13)
State employment trend	−0.74**	−0.72**	−0.72**	−0.58**	−0.58**
	(0.20)	(0.19)	(0.19)	(0.19)	(0.19)
Industry employment residual	0.74**	0.69**	0.70**	0.65**	0.64**
	(0.22)	(0.22)	(0.22)	(0.21)	(0.21)
Industry employment trend	−3.74**	−3.36**	−3.23**	−0.58	−0.58
	(1.12)	(1.20)	(1.21)	(1.39)	(1.39)
Nixon phase II	−10.94**	−11.78**	−11.72**	−11.76**	−11.91**
	(4.82)	(4.79)	(4.79)	(4.82)	(4.82)
Nixon phase III & IV	−1.54*	−1.43*	−1.04	−1.01	−0.88
	(0.92)	(0.93)	(0.95)	(0.94)	(0.95)
Carter phase I	0.39	1.09*	0.88	0.02	−0.03
	(0.52)	(0.58)	(0.57)	(0.57)	(0.57)
Carter phase II	1.22**	0.34	0.26	−1.09**	−1.12**
	(0.60)	(0.38)	(0.38)	(0.38)	(0.38)

Notes: Standard errors are given in parentheses and have been adjusted for any nonindependence across contracts in the same BU. Sample size is 6,524. Continuous RHS variables have been standardized to have a zero mean and unit standard deviation. All specifications control for 2-digit industry fixed effects and 43 union fixed effects.

[1] Predicted values used

** significant at the 5% level

* significant at the 10% level

Table 7. Decomposition of Aggregate Uncertainty

Variable	(1)	(2)	(3)	(4)
Expected inflation - SRC survey consensus	−0.61*	−0.74**	−0.05	−0.07
	(0.30)	(0.31)	(0.27)	(0.28)
Aggregate demand uncertainty				
• Composite	−2.05**			
	(0.34)			
• Money supply uncertainty		−1.47**		
		(0.19)		
• Money demand uncertainty			−1.45**	
			(0.18)	
• IS uncertainty				−1.33**
				(0.16)
Aggregate supply uncertainty	−2.66**	−1.80**	−1.88**	−1.75**
	(0.35)	(0.22)	(0.24)	(0.23)

Notes: Standard errors are given in parentheses and have been adjusted for any nonindependence across contracts in the same BU. Sample size is 6,524. Continuous RHS variables have been standardized to have a zero mean and unit standard deviation. All specifications control for 2-digit industry fixed effects and 43 union fixed effects.

** significant at the 5% level

* significant at the 10% level

Table A1. Contract Durations - Sensitivity Analysis

Variable	(1) Baseline	(2) Unweighted	(3) No Censoring[2]	(4) Effective Date[3]	(5) Mfg, 1970-1980[4]
COLA[1]	2.19**	2.78**	2.79**	2.95**	5.73**
	(0.58)	(0.59)	(0.59)	(0.59)	(0.99)
Reopener[1]	5.19**	5.66**	5.60**	5.93**	5.07
	(2.41)	(2.41)	(2.41)	(2.33)	(3.65)
Log BU Size	0.05	0.18	0.18	0.17	−0.23
	(0.18)	(0.15)	(0.15)	(0.14)	(0.21)
Percent change in BU size (negative changes)	0.27**	0.21**	0.21**	0.25**	0.33
	(0.13)	(0.10)	(0.10)	(0.10)	(0.22)
Expected inflation:	−0.63**	−0.37**	−0.33**	−0.22	0.74**
• SRC survey consensus	(0.32)	(0.20)	(0.20)	(0.19)	(0.35)
Aggregate uncertainty:	−2.14**	−1.74**	−1.69**	−1.68**	−0.63
• Demand (nominal)	(0.35)	(0.21)	(0.20)	(0.24)	(1.11)
• Supply (real)	−2.77**	−2.55**	−2.49**	−2.61**	−1.16**
	(0.37)	(0.25)	(0.24)	(0.24)	(0.45)
State employment residual	−0.00	−0.06	−0.06	−0.07	−0.17
	(0.13)	(0.13)	(0.12)	(0.12)	(0.28)
State employment trend	−0.65**	−0.64**	−0.62**	−0.53**	0.21
	(0.19)	(0.19)	(0.19)	(0.19)	(0.25)
Industry employment residual	0.66**	0.65**	0.64**	0.69**	0.04
	(0.23)	(0.12)	(0.12)	(0.12)	(0.31)
Industry employment trend	−0.26	−1.38	−1.56*	−1.71*	0.77
	(1.43)	(0.88)	(0.86)	(0.82)	(2.95)
Nixon phase II	−13.02**	−10.62*	−10.47*	−9.69*	−8.54*
	(5.15)	(5.50)	(5.51)	(5.22)	(4.72)
Nixon phase III & IV	−1.24	−1.41*	−1.15	−1.87	−1.64**
	(0.94)	(0.79)	(0.77)	(0.82)	(0.92)
Carter phase I	0.12	−0.45	−0.48	−0.44	−0.77
	(0.58)	(0.50)	(0.49)	(0.50)	(0.84)
Carter phase II	−1.15**	−1.40**	−1.41**	−1.44**	0.00
	(0.38)	(0.41)	(0.40)	(0.39)	(0.54)

Notes: Baseline specification is the same as specification (4) from Table 6 with the union fixed-effects omitted. Standard errors are given in parentheses and have been adjusted for any nonindependence across contracts in the same BU. Sample size is 6,524 for specifications (1) - (4), and 1,404 for specification (5). Continuous RHS variables have been standardized to have a zero mean and unit standard deviation. All specifications control for 2-digit industry fixed effects and 43 union fixed effects.

[1] Predicted values used

[2] Maintain equal weighting of all contract negotiations

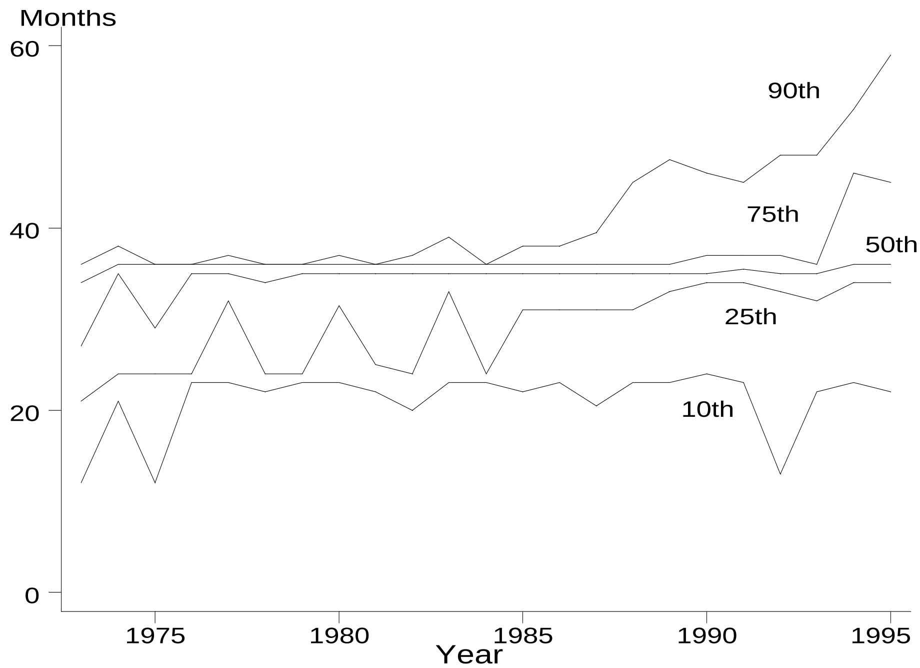
[3] Maintain equal weighting of all contract negotiations and assume no censoring.

[4] Maintain equal weighting of all contract negotiations, assume no censoring and use effective date.

** significant at the 5% level

* significant at the 10% level

Figure 1. Evolution of Contract Durations



Notes: Private sector labor contracts covering at least 1,000 workers. Duration is measured from the effective date to the expiration date.

Figure 2. Distribution of Contract Durations - Timely Settlements

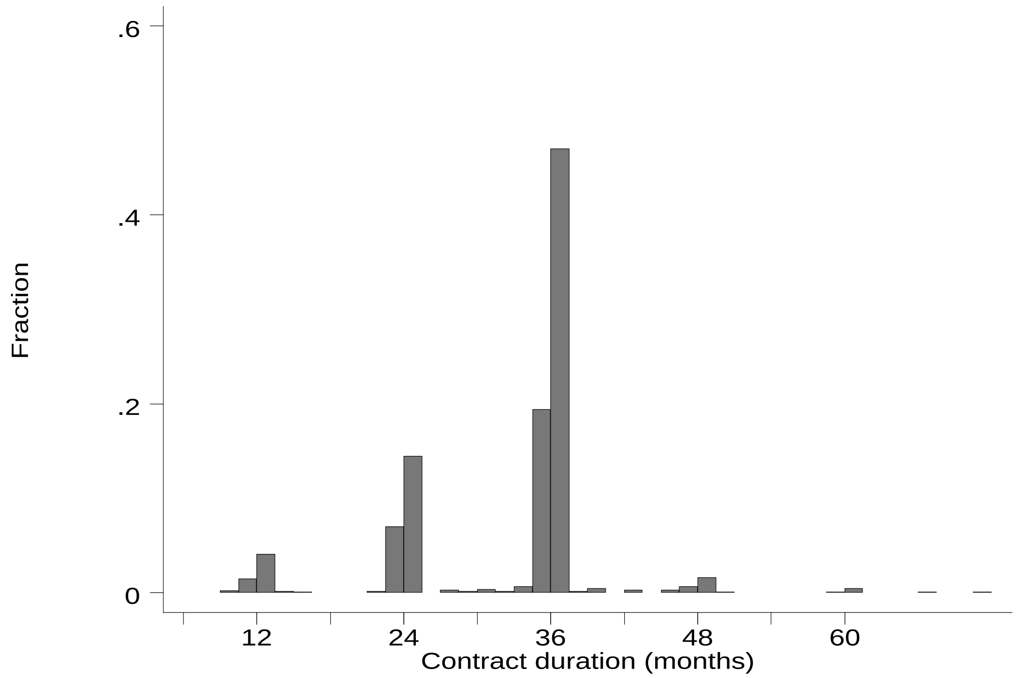


Figure 3. Distribution of the Change in Contract Durations - Timely Settlements

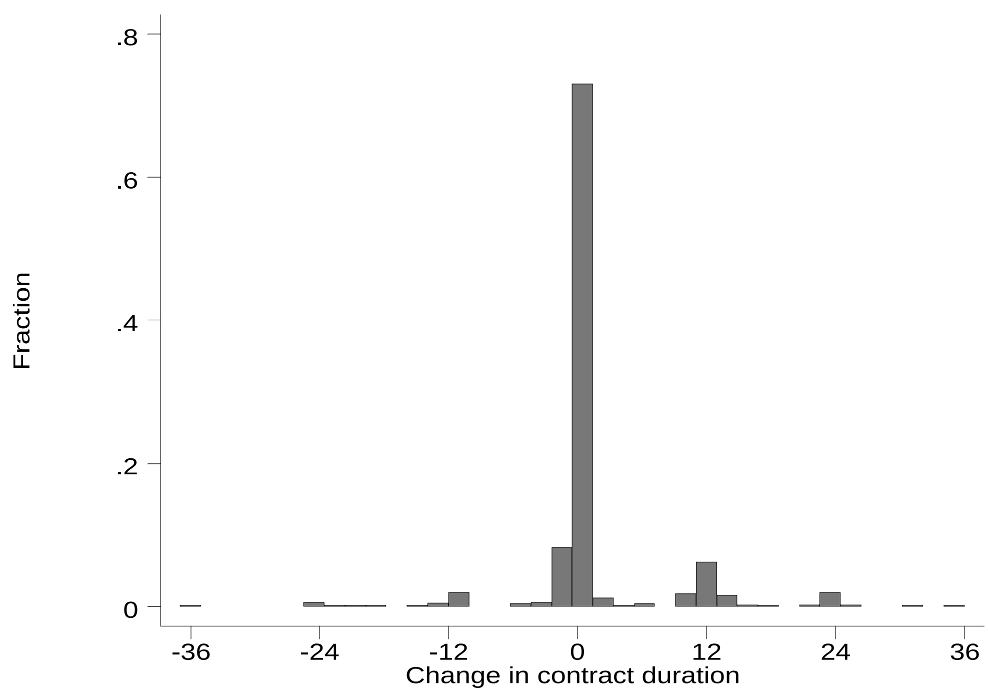
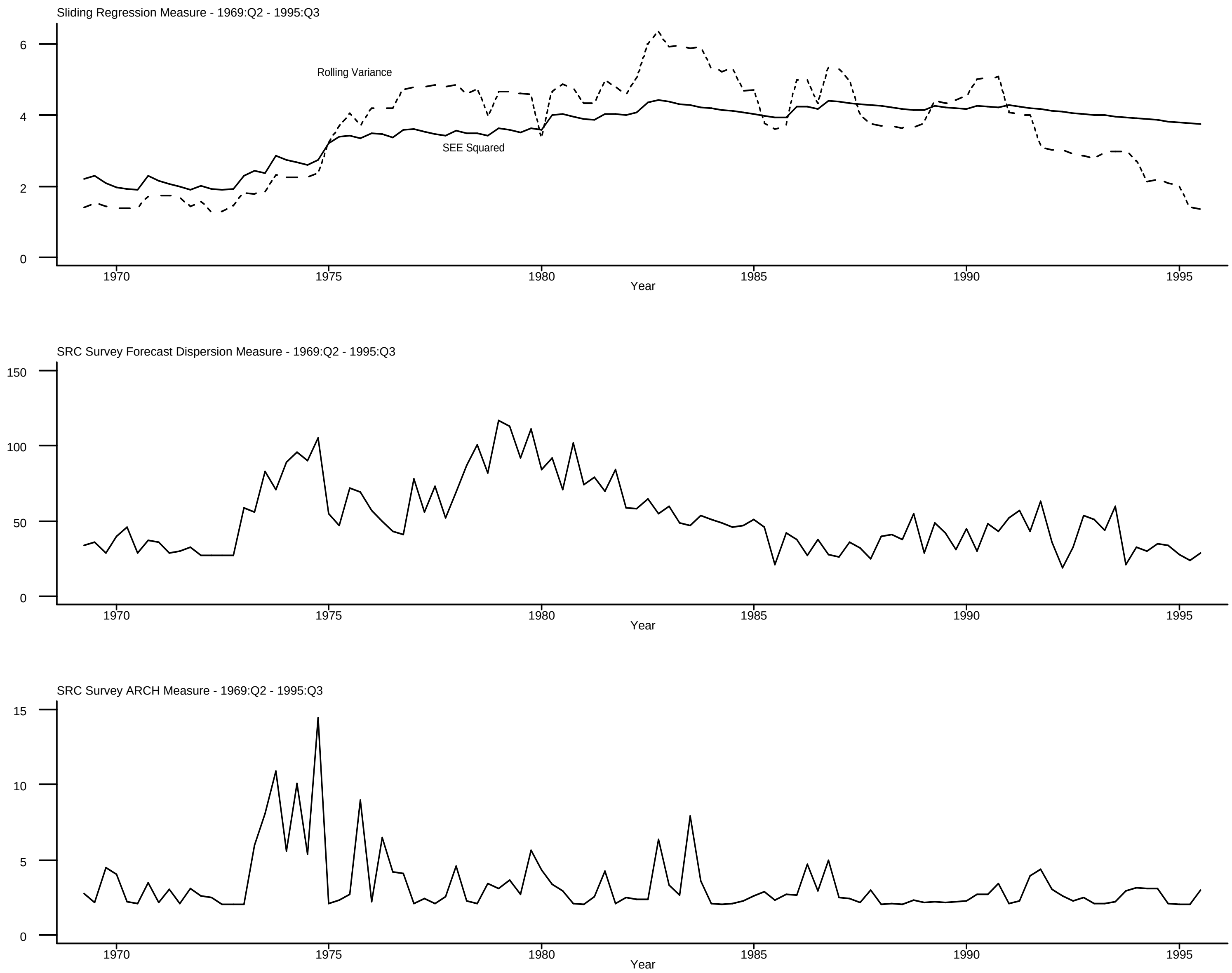
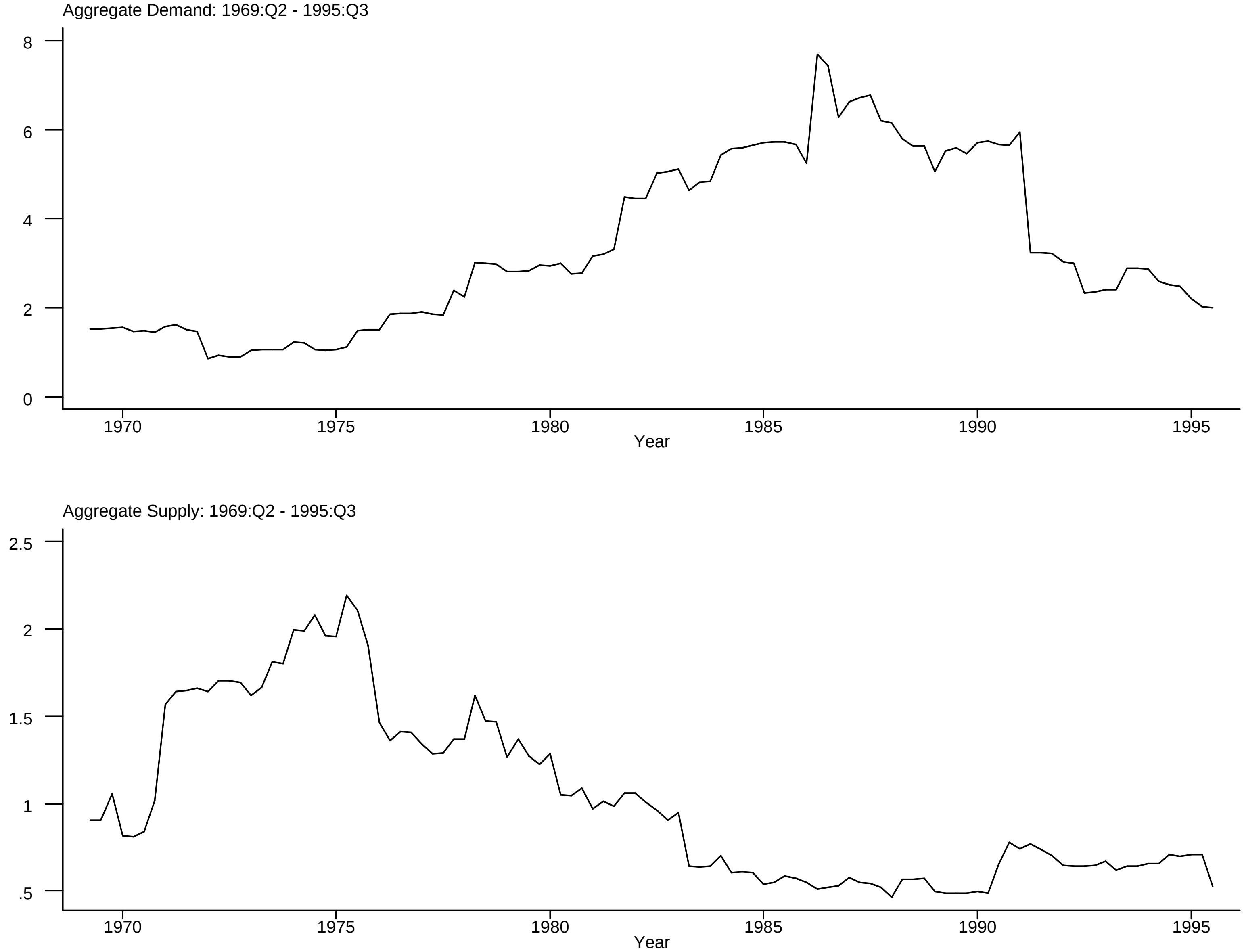


Figure 4. Alternative Measures of Inflation Uncertainty



Note: Dates on the horizontal axis correspond to the end of the five-year window in the top panel.

Figure 5. Variances of Structural Shocks − Gali Model



Note: Dates on the horizontal axis correspond to the end of the five-year window.