Delays in Renewal of Labor Contracts: Theory and Evidence

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In many countries, an expired labor contract is automatically extended during the often-protracted delay before the new contract is signed. Our theoretical model focuses on macroeconomic factors in explaining the delay. It emphasizes the importance of the realized nominal and real shocks, and of the levels of nominal and real uncertainty. The model is tested using Israeli collective wage agreements where long delays are frequent. The empirical findings strongly support the theoretical model. Thus, nominal uncertainty is found to increase the delay, and real uncertainty to decrease the delay, but less in the public than in the private sector.

> Delay is preferable to error. (Thomas Jefferson in letter to George Washington, May 16, 1792)

I. Introduction

Contemporary labor contracts are usually signed for a fixed duration. A typical labor contract specifies its expiration date, and it is rare that a contract has no preset duration or includes a clause permitting reopening

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or extension of the contract. Nevertheless, the duration of labor contracts is more flexible than their rigid wording seems to indicate. In particular, it is common practice that the terms of the old contract are automatically extended during the often-protracted holdout period between the stated expiration date of the old contract and the signing of the new one.

In our sample of Israeli labor contracts, 86% of new contracts are signed after the expiration date of the previous contract. The average delay is 213 days, which is 33% of the average stated contract duration. For large contract settlements in U.S. industries, Cramton and Tracy (1992) find that 47% of contract renewals take place 2 or more days after the old contract expired, and, among those contracts, the average holdout from the expiration date of the old contract to the new agreement or the beginning of a strike is 63 days. Among major Canadian collective bargaining contracts, Gu and Kuhn (1998) find a holdout incidence of 81%; the average holdout is 80 days.[1]

Fixed-duration labor contracts play a pivotal role in many models of aggregate fluctuations. Starting with Fischer (1977), Phelps and Taylor (1977), and Taylor (1979, 1980), economists have built macroeconomic models in which staggered multiperiod contracts of fixed duration lead to sluggish adjustment of the aggregate price level and monetary changes to generate real effects.[2] However, variable-duration contracts with state-dependent renewal dates may lead to radically different conclusions; Caplin and Spulber (1987) show that even though prices are changed discretely in response to economic developments, the aggregate price level adjusts immediately to monetary shocks, which therefore have no aggregate consequences.

If the delay in contract renewal is endogenous, the effective duration of the previous contract is state dependent, notwithstanding the formally stated fixed expiration date. Critically, the dependence of the delay on

[1] There are also a few contracts with durations shortened by a new contract negotiated and implemented prior to the expiration date of the old one (4% in our sample, and 12% in Cramton and Tracy's [1992] sample). Contract reopenings are the focus of Danziger (1995). The law in Israel (and similarly in the United States and Canada) posits that the conditions of the old contract govern the employment relationship in the period from the expiration date of the old contract until the signing of the new contract or the beginning of a strike or lockout. This is also true in many other countries, with the major exception of the United Kingdom, where labor contracts are not legally binding. See also Holden (1994).

[2] See Gray (1978), Dye (1985), Harris and Holmstrom (1987), and Danziger (1988, 1992) for theoretical models of the optimal fixed-contract duration. The empirical literature on contract durations includes Christofides and Wilton (1983), Ehrenberg, Danziger, and San (1984), Christofides (1985), Vroman (1989), Wallace and Blanco (1991), Murphy (1992), and Rich and Tracy (2004). Flanagan (1999) surveys the literature concerned with the relationship between collective bargaining systems and macroeconomic performance.

macroeconomic variables may seriously weaken the realism and explanatory power of theories in which fixed-duration labor contracts constitute the propagation mechanism for monetary and other macroeconomic shocks. For example, if an inflationary monetary shock reduces delay in contract renewals, effectively shortening contract durations, the proportion of new contracts increases with the unanticipated inflation. More contracts are concluded after large than after small shocks, and the former are incorporated into new contracts earlier. This mitigates and possibly neutralizes the differential real impacts that the shocks would have if contract durations were truly fixed.

In view of the above, the purpose of this article is to examine how the delay in contract renewal depends on aggregate economic variables. Our approach falls within the implicit-contract paradigm and emphasizes the importance of macroeconomic factors, in particular, the driving forces behind the value of money and the real value of a worker's marginal product in the relevant economic branch. We present a 4-period economy in which a labor contract between a union and a firm expires at the end of period 2, at which time the parties may either immediately conclude a new labor contract for periods 3 and 4 or delay the renewal until the end of period 3. The degree of wage indexation is exogenous to the parties negotiating the individual labor contracts. This reflects the fact that in Israel the degree of wage indexation is determined by law based on an umbrella agreement between the largest union and the largest federation of employers.[3]

The workers are risk averse and have no access to a capital market. The firm is risk neutral, and the discounted expected real payments to a worker in periods 3 and 4 equal the discounted expected real value of a worker's marginal product in these periods. If the contract is renewed immediately, neither the price level nor the real value of the marginal product in period 4 is known at the time of contract renewal. Since the labor contract sets the (partially indexed) nominal wage for period 4, the workers become exposed to the nominal uncertainty (i.e., the uncertainty in the value of money) in period 4 but are fully insured against the real uncertainty (i.e., the uncertainty of the real value of their marginal product). Conversely, if the contract renewal is delayed, both the price level and the real value of the marginal product in period 4 are known at the time of contract renewal. The nominal wage is therefore set so that workers are paid the real value of their marginal product in period 4 (except for the adjustment for under- or overpayment in period 3); the workers are then fully in-

[3] The exogenous wage indexation simplifies the model but is not essential for the results. For models in which the degree of indexation is freely negotiated by the union and the firm, see Ehrenberg et al. (1984), Card (1986), and Danziger (1988, 1992).

demnified for nominal uncertainty but fully exposed to real uncertainty. Hence, for given realizations of the shocks in periods 2 and 3, nominal riskiness favors a delay, while real riskiness favors an immediate renewal.[4]

The realizations of the shocks in periods 2 and 3 determine how much the workers are under- or overpaid given the real value of their marginal product in period 3 if the contract renewal is delayed. A large under- or overpayment would pull toward immediate renewal. Accordingly, the ratio of the real wage in period 3 with a delay to what the real wage would be in a new contract also plays a central role in determining whether contract renewal should be delayed. The optimal timing of contract renewal is shown to follow an (s, S) strategy in this ratio: contract renewal is delayed if the ratio falls between s and S and is immediate if the ratio is either less than s or greater than S.

In order to test our theory, we collected all published collective wage agreements in Israel from 1978 to 1995. This provides us with a sample of 2,103 contracts with a fixed termination date, signed at or after the time the previous contract expired. We can match each contract with the relevant macroeconomic variables, making it possible to base our tests directly on the theoretical model.[5]

The empirical findings provide strong support for the theoretical model. Since our data set includes information not only about whether contract renewal is delayed but also about the length of delay, we can estimate the relationship between the different parameters and the length of delay. The theory predicts differential impacts of real-uncertainty measures for contracts with firms in the private and public sectors. We therefore first examine the relationship in the 1,731 private-sector contracts separately, and then the relationship in the full sample of 2,103 private- and public sector contracts, where we include interaction terms between the public sector and the explanatory variables.

We use a random-effects model to estimate the length of delay in the private sector alone and, by adding interaction terms with the public sector, in the private and public sectors together. All the coefficient estimates have the predicted sign and are significant. Among the implications of the regression results for the private sector are that in the upward-sloping range, a positive one-standard-deviation nominal shock increases

[4] Since a delay shortens the time period during which the provisions of the next contract will be in force, the opposite effects of the two types of riskiness on the decision to delay are similar to the finding in Danziger (1988). There it was shown that with worker risk aversion, nominal uncertainty tends to shorten, and real uncertainty tends to lengthen, contract duration. It is beyond the scope of this article to model the joint nature of the decision to delay and the duration of the next contract.

[5] We therefore obtain a close integration between the assumption of rational behavior and the labormetrics estimation; see Hamermesh (2000).

the average delay by 6 days, while a positive one-standard-deviation real shock decreases the average delay by 16 days. In the downward-sloping range, a positive one-standard-deviation nominal shock decreases the average delay by 26 days, while a positive one-standard-deviation real shock increases the average delay by 69 days. Furthermore, for given realizations of the shocks, a doubling of the variance of the nominal shock would increase the average delay by 22 days, while a doubling of the variance of the real shock would reduce the average delay by 14 days.

Consistent with our theory, the estimates of the coefficients of the interaction terms with the public sector show that real uncertainty shortens the delay less in the public sector than in the private sector, upcoming elections shorten the delay more in the public sector than in the private sector, and unemployment appears to have no significant effect in the public sector in contrast with its negative effect in the private sector.

We also estimate the effects on the likelihood of delay. The evidence is again very clear: whatever increases the delay also increases the likelihood of delay.

With its emphasis on macroeconomic factors, our model of optimal contract delay differs from previous models. Cramton and Tracy (1992, 1994) present and empirically test a bargaining model in which holdouts and strikes are alternative means by which a union can elicit information about a firm's willingness to pay. Holdouts entail a loss of productive efficiency, but since holdouts do not involve work stoppages, they are a less costly form for dispute than strikes. The main focus is to determine the relative importance of holdouts in labor disputes, and Cramton and Tracy show that the frequency and length of holdouts decrease with the uncompensated inflation in the old contract, the local unemployment rate, and the demand for the firm's output.

Gu and Kuhn (1998) consider multiple bargaining pairs in an industry. In their model holdouts are also used by unions to elicit information about a firm's willingness to pay, but now indirectly by observing settlements between similar bargaining pairs during the holdout period. Gu and Kuhn do not require holdouts to be associated with a loss of productive efficiency, but they also find that the frequency and length of holdouts decrease with both the erosion of the real wage in the old contract and the firm's profitability. In addition, they show that the incentive to delay increases with the number of similar bargaining pairs and that delays are similar for similar bargaining pairs.

It is beyond the scope of this article to design a test that can distinguish between the different models. However, we note that since in our model the time of contract renewal is determined by an (s, S) strategy in the ratio of the real wage during delay to what the real wage would be in a new contract, the benefit from a delay is a nonmonotonic function that first increases and then decreases in this ratio. In the Cramton-Tracy (1992) and Gu-Kuhn (1998) models, the benefit from a delay always increases in this variable.

II. The Model

Consider a 4-period economy with nominal and real uncertainty. Thus, the future value of money (defined as the inverse of the price level) and the future real value of a worker's marginal product are uncertain.

The value of money in period 1 is unity, and the value of money in period i relative to the value of money in period $i-1$ ($i = 2, 3, 4$) is $(1+\mu)(1+x_i)$, where $\mu > -1$ is the trend in the value of money and $x_i > -1$ is a nominal shock. The nominal shocks have zero mean and are mutually independent and identically distributed with density function $f(x_i)$ on $[\underline{x}, \bar{x}]$, where $\underline{x} > -1$.

Each worker in a union supplies one unit of labor to a firm in each period. The real value of a worker's marginal product in period i is A_i, and the real value of the marginal product in period i relative to that in period $i-1$ is $(1+\xi)(1+y_i)$, where $\xi > -1$ is the trend in the real value of the marginal product and y_i is a real shock. The real shocks have zero mean and are mutually independent and identically distributed with density function $g(y_i)$ on $[\underline{y}, \bar{y}]$, where $\underline{y} > -1$. Since productivity and demand factors may affect both the value of money and the real value of the marginal product, the nominal and real shocks may be dependent.

In period 1, the union and the firm conclude a 2-period labor contract covering periods 1 and 2. The contract sets the nominal wage for period 1 at A_1, and the real wage is then also A_1. The nominal wage for period 2 is determined by a base wage b, which is partially indexed to the price level. The degree of indexation is exogenously fixed at θ, $0 < \theta < 1$. Hence, the nominal wage in period 2 is set at $b\{1 - \theta + \theta/[(1+\mu)(1+x_2)]\}$, and the real wage is $b[(1-\theta)(1+\mu)(1+x_2) + \theta]$. The base wage is set to make the expected real wage equal to the expected real value of the marginal product, that is,[6]

$$b \int_{\underline{x}}^{\bar{x}} [(1-\theta)(1+\mu)(1+x_2) + \theta] f(x_2) dx_2 = A_1(1+\xi)$$

[6] It is only for simplicity that the model assumes that the real wage in period 1 equals the real value of the marginal product in period 1 and the expected real wage in period 2 equals the expected real value of the marginal product in period 2. More generally, it could be assumed that the discounted expected value of the real wages in the two periods equals a fraction (representing the union's relative bargaining strength) of the discounted expected real value of the marginal products from these periods. This would complicate the analysis without changing the central results.

$$\Leftrightarrow b = \frac{A_1(1 + \xi)}{1 + (1 - \theta)\mu}.$$

The real wage in period 2 therefore becomes

$$\frac{A_1(1 + \xi)[(1 - \theta)(1 + \mu)(1 + x_2) + \theta]}{1 + (1 - \theta)\mu} = A_1(1 + \xi)(1 + \gamma x_2),$$

where $\gamma \equiv 1 - \theta/[1 + (1 - \theta)\mu]$ is the fraction of the nominal shock transmitted to the real wage.

When the labor contract covering periods 1 and 2 expires, the union and the firm may choose to conclude a new 2-period contract immediately for periods 3 and 4. The new contract is then similar to the expired contract, except that it sets the nominal wage for period 3 at $A_3/[(1 + \mu)^2(1 + x_2)(1 + x_3)]$ and the base wage for period 4 at $A_3(1 + \xi)/\{[1 + (1 - \theta)\mu][(1 + \mu)^2(1 + x_2)(1 + x_3)]\}$. The real wage in period 3 is therefore A_3, which is the real value of the marginal product in period 3. The realized real wage in period 4 becomes $A_3(1 + \xi)(1 + \gamma x_4)$, and the expected real wage in period 4 is $A_3(1 + \xi)$, which is the expected real value of the marginal product in period 4.

Alternatively, the union and the firm may delay contract renewal to period 4, in which case the provisions of the old contract are automatically extended to cover period 3.[7] The relative change in the base wage between periods 2 and 3 is similar to that between periods 1 and 2, and the base wage is again partially indexed to the price level. Accordingly, the nominal wage in period 3 is set at $A_1(1 + \xi)^2\{1 - \theta + \theta/[(1 + \mu)(1 + x_2)]\}\{1 - \theta + \theta/[(1 + \mu)(1 + x_3)]\}[1 + (1 - \theta)\mu]^{-2}$, and the real wage becomes $A_1(1 + \xi)^2(1 + \gamma x_2)(1 + \gamma x_3)$. Let

$$z \equiv \frac{(1 + \gamma x_2)(1 + \gamma x_3)}{(1 + y_2)(1 + y_3)}$$

denote the ratio of the real wage with a delay to what the real wage would be with a new contract. The real wage in period 3 can then be written as $A_3 z$. The real under- or overpayment to the worker during the extension is $A_3(1 - z)$.

The delayed contract, when it is eventually concluded in period 4, is made retroactive to period 3. As the price level and the real value of the marginal product in period 4 are now known, the wage in period 4 is set so that the real wage is equal to the real value of the marginal product plus a retroactive payment representing the current value of the real under- or

[7] As mentioned in n. 1, this is the legal practice in many countries.

overpayment to a worker during the extension of the previous contract.[8] Let $r > -1$ denote the real interest rate. Accordingly, if the conclusion of the new contract is delayed, the real wage in period 4 is $A_4 + A_3(1 - z)(1 + r) = A_3[(1 + \xi)(1 + y_4) + (1 - z)(1 + r)]$.[9]

A worker cannot borrow or lend. He is risk averse, and his utility in each period is a logarithmic function of his real income during that period. The expected utility from periods 1 and 2 covered by the first contract is independent of the time at which the next contract is concluded. The expected utility from periods 3 and 4, however, depends on when the new contract is concluded. Let $\rho > -1$ denote a worker's discount rate. On the one hand, with an immediate renewal, the discounted expected utility from periods 3 and 4 is

$$\ln A_3 + \frac{1}{1+\rho}\int_{\underline{x}}^{\bar{x}} \ln[A_3(1+\xi)(1+\gamma x_4)]f(x_4)dx_4$$

$$= \ln A_3 + \frac{1}{1+\rho}\left\{\ln[A_3(1+\xi)] + \int_{\underline{x}}^{\bar{x}} \ln(1+\gamma x_4)f(x_4)dx_4\right\}.$$

On the other hand, with a delayed renewal, the discounted expected utility from periods 3 and 4 is

$$\ln(A_3 z) + \frac{1}{1+\rho}\int_{\underline{y}}^{\bar{y}} \ln\{A_3[(1+\xi)(1+y_4) + (1-z)(1+r)]\}g(y_4)dy_4$$

$$= \ln(A_3 z) + \frac{1}{1+\rho}\left\{\ln[A_3(1+\xi)] + \int_{\underline{y}}^{\bar{y}} \ln[1 + y_4 + \frac{(1-z)(1+r)}{1+\xi}]g(y_4)dy_4\right\}.$$

The benefit from a delay depends on z and is obtained by subtracting the discounted expected utility from periods 3 and 4 of a contract that is concluded immediately from the discounted expected utility from periods 3 and 4 of a contract with delayed renewal,

$$B(z) \equiv \ln z + \frac{1}{1+\rho}\int_{\underline{y}}^{\bar{y}} \ln\left[1 + y_4 + \frac{(1-z)(1+r)}{1+\xi}\right]g(y_4)dy_4$$

$$- \frac{1}{1+\rho}\int_{\underline{x}}^{\bar{x}} \ln(1+\gamma x_4)f(x_4)dx_4.$$

Since the utility function is logarithmic, the benefit from a delay is

[8] The model thus captures that new contracts are typically backdated and contain a retroactive payment. Since the discounted expected real profits per worker are $A_3[1 + (1+\xi)/(1+r)]$, whether there is a delay or not, the retroactive payment has the valuable implication that the delay decision is of no concern to the firm.

[9] It is assumed that $(1+\xi)(1+\underline{y})/(1+r) + 1 > [(1+\gamma\bar{x})/(1+\underline{y})]^2$, so that the real wage is positive.

independent of a possible dependence between nominal and real shocks. Furthermore, the nature of indexation implies that the size of the nominal shocks and the fraction γ of these shocks transmitted to the real wage enter only multiplicatively into the benefit from a delay—as γx_2 and γx_3 in z, and as γx_4. Accordingly, a change in the nominal shocks by a factor of $\lambda > 0$ together with a simultaneous change in the degree of indexation or in the trend in the value of money such that γ changes by a factor of $1/\lambda$ would have no impact on $B(z)$.

The firm is risk neutral, and since its discounted expected real profits per worker are the same with and without a delay, it agrees that the timing of the contract renewal is chosen to maximize a worker's discounted expected utility. Accordingly, the contract renewal is delayed if $B(z) > 0$, the contract is immediately renewed if $B(z) < 0$, and the contract renewal is either delayed or immediate if $B(z) = 0$.

The benefit from a delay is a strictly concave function of z with an internal maximum at $z = z^*$ defined by

$$\int_{\underline{y}}^{\bar{y}} \left\{ \frac{z^*(1+r)}{[(1+\xi)(1+y_4)+(1-z^*)(1+r)](1+\rho)} \right\} g(y_4)dy_4 = 1. \qquad (1)$$

This reflects that if $z < z^*$, larger nominal shocks or smaller real shocks in periods 3 and 4 bring the real wage in period 3 with a delay closer to maximizing the discounted expected utility with a delay; if $z > z^*$, the opposite holds.

It is assumed that the real uncertainty is not so extreme that it would never be optimal to delay the contract renewal, $B(z^*) > 0$. Also, it is assumed that the shocks are sufficiently dispersed that there are shocks for which $z < z^*$, and it is optimal immediately to renew the contract, $B\{[(1+\gamma\underline{x})/(1+\bar{y})]^2\} < 0$, and that there are shocks for which $z > z^*$, and it is optimal immediately to renew the contract, $B\{[(1+\gamma\bar{x})/(1+\underline{y})]^2\} > 0$. The optimal timing of contract renewal can then be described as a two-sided (s, S) strategy in z, where the lower and upper critical values of z, denoted by s and S, are unique and defined by $B(s) = B(S) = 0$, $s < S$: for a given realization of shocks in periods 2 and 3, the renewal is delayed if $z \in (s, S)$, is immediate if $z(s, S)$, and is either delayed or immediate if $z = s$ or $z = S$. Figure 1 shows the benefit from delay and illustrates the choice between delay and immediate renewal. The value of z is measured on the horizontal axis, and $B(z)$ is measured on the vertical axis.[10]

[10] In the special case of no real uncertainty and $\rho = 0$, then $B(z)$ is symmetric around $z^* = 1/2[1 + (1+\xi)/(1+r)]$ and $B(z^*) = \ln\{(2+r+\xi)^2/[4(1+\xi)(1+r)]\} - \int_{\underline{x}}^{\bar{x}} \ln(1+\gamma x_4) f(x_4)dx_4$. A worker is then indifferent between receiving the real wage $z^* - \phi$, $\phi > 0$, in period 3 and the expected real wage, $z^* + \phi$, in period 4, and vice versa. Since the left-hand side of eq. (1) increases with z^* and is a convex function of y_4, real uncertainty leads to a decrease in z^*.

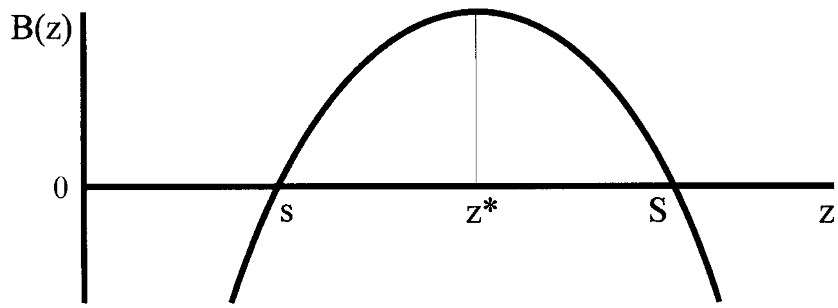

FIG. 1.—The benefit from delay

III. The Benefit from Delay and the Critical Values of z: Comparative Statics

It has been shown that the benefit from delay depends on the realized shocks as manifested in z: the benefit increases with z if $z < z^*$ and decreases with z if $z > z^*$. The benefit therefore increases with the nominal shocks if $z < z^*$ and decreases with the nominal shocks if $z > z^*$, while the real shocks have the opposite effect. We now examine how the parameters affect the benefit for a given z and study the effects of the parameters on the upper and lower critical values of z.

The nominal parameters—the trend in the value of money, the level of the nominal uncertainty, and the degree of indexation—have no effect on the discounted expected utility from a contract with delayed renewal, for a given z. The reason is that a worker in period 3 receives the real wage A_3z, which is given since A_3 and z are given; in period 4 he receives a real wage equaling the real value of the marginal product plus retroactive pay for period 3 and, therefore, independent of the change in the value of money from period 3 to period 4. Thus, delayed renewal protects the worker from any real effects of anticipated and unanticipated changes in the value of money from period 3 to period 4.

However, the same nominal parameters affect the discounted expected utility if the contract is renewed immediately in period 3. In this case, an increase in the nominal uncertainty increases the uncertainty of the real wage in period 4, which reduces the risk-averse worker's discounted expected utility. The magnitude of the impact of the nominal uncertainty depends on the fraction γ of period 4's nominal shock transmitted to the real wage. Since γ increases with μ and decreases with θ, an increase in μ is equivalent to an increase in uncertainty, while an increase in θ is equivalent to a decrease in uncertainty. So for a given z, both the trend in the value of money and the nominal uncertainty lead to a decrease in the discounted expected utility, increasing the benefit from a delay, while the degree of indexation increases the discounted expected utility, decreasing the benefit from a delay. As is clear from figure 1, a larger trend in the value of money and more nominal uncertainty lead to a decrease in s and an increase in S, thereby widening the range of z's for which the contract renewal is delayed; more indexation leads to an increase in s and a decrease in S, thereby narrowing the range of z's for which the contract renewal is delayed.

Turning to the trend in the real value of the marginal product, this has no effect on the third period's wage in either contract for a given A_3 but does affect the fourth period's wage in both contracts.[11] If the contract is renewed in period 3, the fourth period's wage is proportional to $1 + \xi$. If the renewal is delayed and $z = 1$ (so that there is no retroactive pay in period 4), the fourth period's wage is also proportional to $1 + \xi$. In this case the real wage in period 3 is the same irrespective of whether or not the contract renewal is delayed. The benefit from a delay is therefore independent of ξ, and $B(1)$ does not change with ξ.

If the renewal is delayed and $z > 1$, the real wage in period 3 is greater, and the expected real wage in period 4 is less than if the contract is renewed in period 3. Because of the gains from a smoother intertemporal consumption, an increase in ξ is therefore more beneficial if the contract renewal is delayed. If the renewal is delayed and $z < 1$, the opposite is true. Consequently, ξ affects the benefit from a delay positively for $z > 1$ and negatively for $z < 1$. It follows that $ds/d\xi \gtreqless 0$ as $s \lesseqgtr 1$ and that $dS/d\xi \gtreqless 0$ as $S \gtreqless 1$.[12]

Real uncertainty has no effect on the discounted expected utility from a contract concluded in period 3, since the expected real wage in period 4 then equals the expected real value of the worker's marginal product and is independent of the real uncertainty. The worker is fully insured against real uncertainty in period 4. However, uncertainty reduces the expected utility from a delayed contract, since the real wage in period 4 becomes exposed to the real uncertainty. Accordingly, an increase in real uncertainty decreases the benefit from a delay. Hence, s increases and S decreases, leading to a narrower range of z's for which the contract renewal is delayed.

Finally, the effect of the real interest rate works through the value of the retroactive pay in period 4 of a delayed contract. If the retroactive pay is positive (negative), the real wage in period 4 increases (decreases)

[11] We ignore the effects of ξ on the wages through A_3, since the wages in the third and fourth period are proportional to A_3 irrespective of whether the contract renewal is delayed.

[12] The effect on the range of z-values for which the contract renewal is delayed depends on the values of s and S relative to unity: a bigger trend in productivity widens the range of z-values for which the contract renewal is delayed if $1 \leq s < S$, narrows the range if $s < S \leq 1$, and moves the range to the right if $s < 1 < S$.

Table 1
The Benefit from Delay and the Critical Values of *z*

	$B(z)$	s	S
z	$\gtreqless 0$ as $z \lesseqgtr z^*$		
μ	>0	<0	>0
Nominal uncertainty	>0	<0	>0
θ	<0	>0	<0
ξ	$\gtreqless 0$ as $z \gtreqless 1$	$\gtreqless 0$ as $s \lesseqgtr 1$	$\gtreqless 0$ as $S \gtreqless 1$
Real uncertainty	<0	>0	<0
r	$\lesseqgtr 0$ as $z \gtreqless 1$	$\lesseqgtr 0$ as $s \lesseqgtr 1$	$\lesseqgtr 0$ as $S \gtreqless 1$

with the real interest rate. It follows that r has the opposite effect of ξ on the benefit of a delay. So, if $z = 1$, the benefit from a delay is independent of r; if $z > 1$, it decreases with r; and if $z < 1$, it increases with r. Accordingly, $ds/dr \lesseqgtr 0$ as $s \lesseqgtr 1$, and $dS/dr \lesseqgtr 0$ as $S \gtreqless 1$.[13] The comparative-static relationships are summarized in table 1.

IV. The Israeli Economy, 1978–95

The development of some of the major macroeconomic variables in Israel during 1978–95 is summarized in table 2. During this time span, the Israeli economy experienced two distinct periods of inflation: the annual inflation rate climbed from 48.1% in 1978 to about 800% in May 1985 (with a highly variable monthly inflation); thereafter, the stabilization program, enacted in May 1985, reduced the inflation rate to 19.7% in 1986 and further to 8.1% in 1995. The average annual inflation rate for the entire period was around 60%. This is illustrated in figure 2.

The growth rate of per capita gross domestic product (GDP) exhibits no clear trend. The growth rate varied between an annual minimum of −0.56% in 1989 and an annual maximum of 10.6% in 1991 (caused by mass immigration that started in September 1989 and the Oslo accords). This is illustrated in figure 3. The annual real interest rate exhibits a hump-shaped pattern, similar to the pattern of the inflation rate. The real interest rate is on average 19.9% and varies between a minimum of −11.4% in 1979 and a maximum of 90.6% in 1985. The unemployment rate rose from an annual average of about 3% in 1978–79 to a peak of 11.2% in 1992 (because of the mass immigration during 1989–92). The trend then changed, and unemployment decreased during 1993–95. The number of work days lost because of strikes shows no apparent trend. The average is 574 days per thousand workers, ranging from a minimum of 63 days in 1991 to a maximum of 1,552 days in 1982.

[13] A higher real interest rate narrows the range of z's for which the contract renewal is delayed if $1 \leq s < S$, widens the range if $s < S \leq 1$, and moves the range to the left if $s < 1 < S$.

Table 2
Macro Economic Background Data (Annual)

Year	Inflation Rate (%)	Growth Rate of Per Capita GDP (%)	Real Interest Rate (%)	Unemployment Rate (%)	Work Days Lost (due to Strikes) per 1,000 Employees
1978	48.1	1.93	1.1	3.4	1,054.46
1979	111.4	2.05	−11.4	2.9	488.50
1980	132.9	.67	18.5	4.6	195.54
1981	101.5	2.88	34.3	5.1	684.85
1982	131.5	−.46	3.8	5.0	1,552.30
1983	190.7	.68	−3.1	4.5	818.02
1984	444.9	.15	59.8	5.9	835.15
1985	185.2	2.72	90.6	6.7	450.94
1986	19.7	2.91	33.4	7.1	339.85
1987	16.1	4.61	38.6	6.1	804.29
1988	16.4	1.91	25.6	6.4	410.49
1989	20.7	−.56	11.3	8.9	185.14
1990	17.6	2.82	10.2	9.6	750.88
1991	18.0	10.60	10.0	10.6	63.40
1992	9.4	3.20	11.5	11.2	234.54
1993	11.2	.68	6.2	10.0	925.52
1994	14.5	4.18	4.7	7.8	413.36
1995	8.1	4.02	13.3	6.9	126.43

Sources.—Bank of Israel 1978–95; Bank of Israel, Research Department 1978–95; Central Bureau of Statistics 1978–95a, 1978–95b.
Note.—GDP = gross domestic product.

V. The Sample of Labor Contracts

Israeli employers are required to report all collective bargaining agreements to the Ministry of Labour and Social Affairs. The main characteristics of the labor contracts are subsequently published in the ministry's monthly bulletin, which identifies the parties to the contract, the contract's starting date (assumed to be the expiration date of the previous one), the signing date, the termination date, the economic branch, and important contract provisions, such as wages, vacations, and pensions. Our sample includes all published contracts dealing with wage provisions signed between 1978 and 1995 with a fixed termination date.[14] All the contracts, whether in the private or public sector, resulted from negotiation by the parties and not from legislative intervention. There are 92 (4% of total) new contracts that were signed before the previous contracts terminated and, therefore, exhibited a negative delay. These contracts are excluded from the empirical study, leaving a sample of 2,103 contracts. They stem from 711 different firms (including public-sector employers), since 325

[14] In the beginning of the 1980s, the unionization rate was about 85%. In 1995, health insurance was separated from union membership, causing a sharp fall in the unionization rate to about 50% as well as radical changes in labor relations (Cohen et al. 2001). We therefore decided not to include contracts signed after 1995 in the sample.

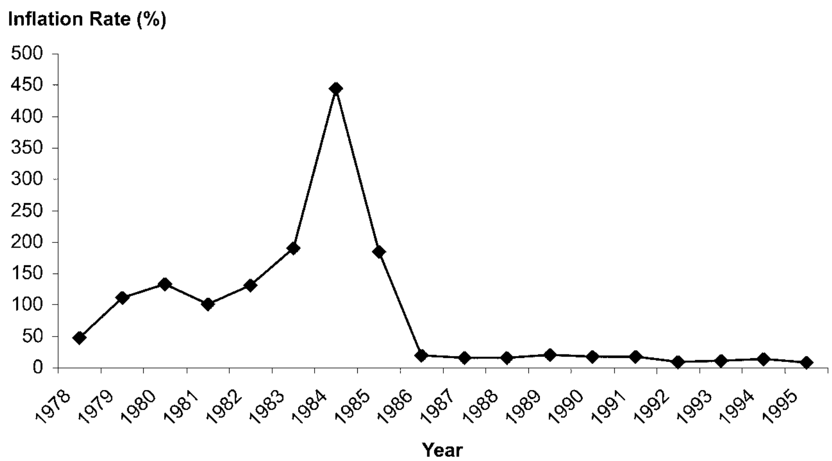

FIG. 2.—Annual inflation rate in Israel, 1978–95

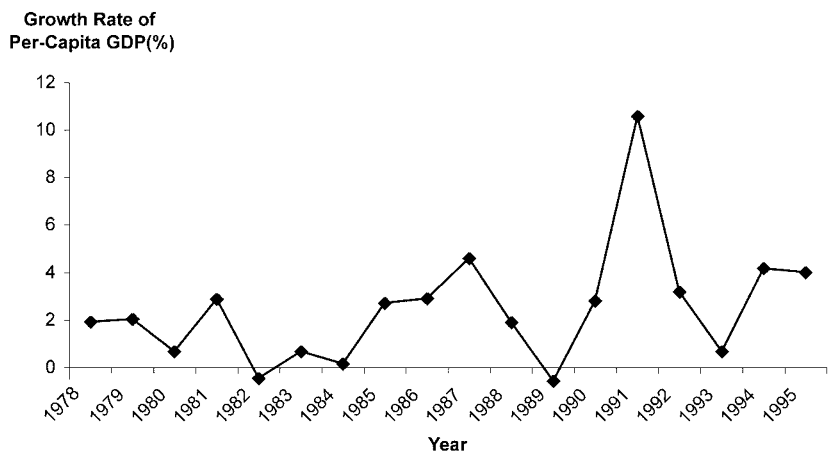

FIG. 3.—Annual growth rate of per capita gross domestic product in Israel, 1978–95

firms enter the sample more than once as they conclude several agreements over the years. As a result, our sample is an unbalanced panel data set. We distinguish between contracts in the private and public sectors.[15] Table 3 shows the distribution of the contracts by sector and economic branch.[16]

[15] The public sector includes municipalities, universities, and most hospitals, among others.

[16] Firms in the public-services branch are not necessarily in the public sector. For example, private schools and private hospitals belong to the public-services branch but not to the public sector.

Table 3
Distribution of Contracts by Sector and by Economic Branch, Israeli Labor Market, 1978–95

Sector/Economic Branch	Share (%)
Sector:	
Private	82.31
Public	17.69
Economic branch:	
Private services	27.29
Public services	9.18
Manufacturing	56.97
Commerce	4.37
Banking institutions	2.19

SOURCE.—Authors' calculations based on Ministry of Labour and Social Affairs (1978–95).
NOTE.—2,103 contracts.

About 82% of the contracts are in the private sector (1,731 contracts), and 18% are in the public sector (372 contracts). Concerning the economic branches, over one-half of the contracts are in manufacturing, over one-quarter are in private services, and about 9% are in public services, while only 4.4% are in commerce and 2.2% are in banking institutions.

The delay in contract renewal is the difference between the signing date of the new contract and the expiry date of the old one. Table 4 provides summary statistics of the delays in contract renewals and contract durations. Only 10% of the contracts are signed on time, and only another 1.9% are signed with a delay of at most 1 week. The average delay is 213 days. The delay varies greatly, from 0 to 1,529 days, with a standard deviation of 207 days. Figure 4 shows the distribution of delays. The private sector has an average delay of 194 days, while the public sector has a longer average delay of 304 days. In the public sector the delays are also more dispersed.

We have no information about whether a contract settlement is reached following the onset of a strike, or when such a strike begins. Our measure of delay therefore includes the length of any strike prior to the signing of the new contract. However, such strikes are rare since the average yearly number of strikes in all of the Israeli economy during 1983–92 is 127, and only 13% of all strikes are caused by the inability of the parties to reach a new contract (Bar-Zuri 1994).

The duration of a new contract includes the delay and is therefore calculated as the difference between the termination dates of the new and old contracts. The average duration of a labor contract is 649 days, which is about three times the average delay. There are two cluster points: at 1 year (21% of the contracts) and at 2 years (52% of the contracts). Only 10% of the contracts exceed 2 years, and a mere 3% exceed 3 years. Contracts in the public sector are longer than ones in the private sector:

Table 4
Delays in Contract Renewals and Contract Durations by Sector, Israeli Labor Market, 1978–95

	Total	Private Sector	Public Sector
Negative delay (%)	3.97	4.36	2.11
No delay (%)	10.13	11.21	5.11
1–7 days delay (%)	1.90	1.85	2.15
Average delay (days)	213.29	193.79	304.02
	(207.20)	(188.78)	(258.92)
Relative delay (delay divided by duration)	.3337	.3151	.4200
Average duration (days)	649.49	628.57	746.83
	(264.57)	(232.09)	(365.86)
Duration of 1 year (%)	20.73	21.43	17.47
Duration of 2 years (%)	51.55	51.70	50.81
Duration of more than 2 years (%)	10.27	8.43	18.82
Duration of more than 3 years (%)	3.28	1.62	11.02
Sample size	2,103	1,731	372

Source.—Authors' calculations based on Ministry of Labour and Social Affairs (1978–95).
Note.—Figures in parentheses are standard deviations. The 92 contracts with negative delay are not included in the sample size.

in the private sector the average duration is 629 days, while in the public sector the average duration is 747 days. Among the private-sector contracts, 8% are signed for more than 2 years and 2% for more than 3 years. The corresponding percentages for the public sector are more than double: 19% for more than 2 years and 11% for more than 3 years.

VI. The Empirical Implementation

In order to test the model's predictions for the delay in contract renewals, we take a second-order Taylor approximation of the benefit $B(z)$ from delaying the new contract at $(x_4, y_4) = (0, 0)$. This yields

$$\ln z + \frac{1}{1+\rho}\ln\left[1 + \frac{(1-z)(1+r)}{1+\xi}\right] - \frac{\sigma_y^2}{2(1+\rho)[1+(1-z)(1+r)/(1+\xi)]^2} + \frac{\gamma^2\sigma_x^2}{2(1+\rho)},$$

where σ_x^2 is the variance of the nominal shock and σ_y^2 is the variance of the real shock. Since $(1-z)(1+r)/(1+\xi)$ is small, we further approximate the benefit from a delay by

$$\bar{B}(z) \equiv \ln z + \frac{1}{1+\rho}\ln\left[1 + \frac{(1-z)(1+r)}{1+\xi}\right] - \frac{\sigma_y^2}{2(1+\rho)} + \frac{\gamma^2\sigma_x^2}{2(1+\rho)}.$$

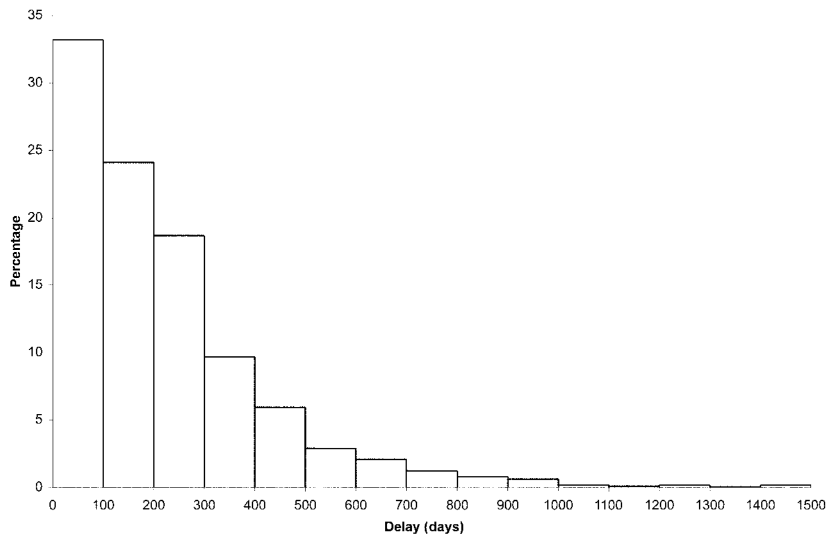

FIG. 4.—Delay (days) in contract renewal in Israel, 1978–95

In the empirical model, we interpret $\bar{B}(z)$ first as the delay and afterward as the likelihood of delay.

All variables are measured at the time the old contract expires, except the degree of indexation. The logic is that the decision about a possible delay in contract renewal is based on the data available at the point at which the old contract expires.

We now define and describe each of the variables used in the empirical analysis. Table 5 presents the means and standard deviations together with their minimum and maximum values.[17]

The rate of change in the actual value of money from month $m - 1$ to month m is $\mu_m \equiv (p_{m-1}/p_m) - 1$, where p_{m-1} and p_m are the consumer price indexes in months $m - 1$ and m, respectively.[18] We use AR(6) to estimate the trend in the value of money from month $m - 1$ to month m, $\hat{\mu}_m$.[19] The estimated average $\hat{\mu}_m$ is -0.0385, with a standard deviation of 0.0384.

[17] The values are calculated over the 2,103 contracts. The values of the macrovariables calculated over the years 1978–95 are similar.

[18] The monthly consumer price index is published in the *Monthly Bulletin of Statistics* (Central Bureau of Statistics 1978–95a).

[19] AR = autoregression. We also tried to estimate the trend in the value of money by 6-month moving arithmetic or geometric averages, which were very similar. The estimations were repeated using AR(12). While the estimates are somewhat different, they lead to similar conclusions in the regression analyses. In addition, we experimented with 6- and 12-month partly backward- and partly forward-looking estimates. These perform less well in the regression analyses.

Table 5
Characteristics of Regression Variables, Israeli Economy, 1978–95

Variable	Mean	Standard Deviation	Minimum Value	Maximum Value
Delay (days)	213.29	207.20	0	1,529
Value of money:				
Trend in the value of money	−.0385	.0384	−.1533	−.0041
Nominal shock	−.0159	.0238	−.1185	.1385
Expected variance of nominal shock	6.8*E*−4	1.2*E*−3	2.9*E*−6	7.5*E*−3
GDP (per capita):				
All economic branches				
Trend in GDP	.0032	.0140	−.0224	.0527
Real shock	.0038	.0263	−.0306	.0780
Expected variance of real shock	2.3*E*−4	5.5*E*−4	5.9*E*−9	2.6*E*−3
Private services (574 contracts):				
Trend in GDP	.0042	.0191	−.0177	.0527
Real shock	.0055	.0297	−.0306	.0780
Expected variance of real shock	2.4*E*−4	6.1*E*−4	1.5*E*−6	2.6*E*−3
Public services (193 contracts):				
Trend in GDP	.0029	.0083	−.0096	.0238
Real shock	.0012	.0140	−.0191	.0339
Expected variance of real shock	6.5*E*−5	1.4*E*−4	9.5*E*−8	7.0*E*−4
Manufacturing (1,198 contracts):				
Trend in GDP	.0024	.0120	−.0224	.0252
Real shock	.0048	.0278	−.0230	.0751
Expected variance of real shock	2.7*E*−4	5.6*E*−4	2.9*E*−7	2.4*E*−3
Commerce (92 contracts):				
Trend in GDP	.0011	.0100	−.0095	.0274
Real shock	.0024	.0206	−.0265	.0510
Expected variance of real shock	1.4*E*−4	1.7*E*−4	2.2*E*−6	5.8*E*−4
Banking institutions (46 contracts):				
Trend in GDP	.0021	.0124	−.0135	.0331
Real shock	.0034	.0216	−.0149	.0631
Expected variance of real shock	1.2*E*−4	2.4*E*−4	5.9*E*−9	8.4*E*−4
Degree of indexation	.5888	.1756	.0336	1.0000
γ	.4035	.1775	0	.9663
z	.9916	.0254	.9031	1.0696
Real interest rate	.0153	.0129	−.0101	.0537
Election years	.2240			
Unemployment rate	.0705	.0232	.0288	.1118

Sources.—Authors' calculations based on Bank of Israel (1978–95), Bank of Israel, Research Department (1978–95), Central Bureau of Statistics (1978–95a, 1978–95b), and Ministry of Labour and Social Affairs (1978–95).

Note.—2,103 contracts. GDP = gross domestic product. Trend, shock, and real interest rate are per month. The value of z (the ratio of the real wage with delay to the real wage without delay) is measured in percent.

The estimated nominal shock is the difference between the rate of change in the actual value of money and the estimated trend in the value of money, $x_m \equiv \mu_m - \hat{\mu}_m$. The standard deviation of the shock is 0.0238, and the shock ranges from −0.1185 to 0.1385. We use the (moving-average) variance of the shocks in the previous 6 months as an estimate of the expected variance of the shock in month m. The expected variance is, on average, $6.8E-4$ and ranges from $2.9E-6$ to $7.5E-3$.

Since there is no published data on the real value of the marginal product, we instead use the per capita GDP in the empirical analysis on the assumption that the real value of the marginal product is proportional to the per capita GDP. The GDP is published annually for each economic branch, and we calculate the rate of change in the actual per capita GDP in a branch for each month m in year a as $\xi_m = [(\mathrm{GDP}_a/\mathrm{GDP}_{a-1}) - 1]^{1/12}$, where GDP_a and GDP_{a-1} are the per capita GDP in the branch in year $a - 1$ and a, respectively.[20] We use AR(2) to estimate the trend in branch per capita GDP from year $a - 1$ to year a, $\hat{\xi}_a$, and calculate the trend in the branch per capita GDP for each month m in year a as $\hat{\xi}_m \equiv \hat{\xi}_a^{1/12}$.[21] The average branch-specific per capita monthly GDP trend ranges from 0.11% in commerce to 0.42% in private services.

The estimated real shock in a branch in a month in year a is the deviation of the estimated trend in the branch per capita GDP from the rate of change in actual branch per capita GDP, $y_m \equiv \xi_m - \hat{\xi}_m$. As an estimate of the expected variance of the shock in each month of the year, σ^2_{ym}, we use the (moving-average) variance of the shocks in the months of the previous 2 years. The ranking of the branches in terms of increasing uncertainty of their real shocks is public services, banking institutions, commerce, private services, and manufacturing.

Taking an average over all 2,103 contracts, the monthly estimate of the trend in the branch per capita GDP, $\hat{\xi}_m$, ranges from -2.24% to 5.27%, with an average monthly trend for 1978–95 of 0.32%. The standard deviation of the monthly real shocks is 0.0263, and the shock size ranges from -0.0306 to 0.0780. The expected variance of the monthly real shocks is on average $2.3E - 4$ and ranges from $5.9E - 9$ to $2.6E - 3$.[22]

The wage indexation rules in Israel are determined by agreements negotiated between the Histadrut (the largest union) and the Coordinating Bureau of Economic Organizations (the largest federation of employers). Subsequently, the Knesset (the Israeli parliament) extended the rules to cover all workers in the economy, and the wage indexation is therefore exogenous to the parties concluding a labor contract. The typical indexation agreement is complicated since the degree of indexation depends on the inflation rate. Empirically, the degree of indexation is measured as the

[20] The annual branch GDP is published in the *Monthly Bulletin of Statistics* (Central Bureau of Statistics 1978–95a).

[21] Similar to estimating the trend in the value of money, we also tried moving arithmetic and geometric versions, which leads to similar estimates. Using AR(3) to estimate the productivity trend does not change the signs and significance of the coefficients in the delay regressions.

[22] We have also experimented with using economy-wide GDP productivity measures, which have the advantage that the GDP figures are available on a quarterly basis. Since branch-productivity measures are more relevant for the delay decision in a given branch, the regression results using economy-wide GDP productivity measures are less satisfactory.

rate of change in the wage during the contract period resulting from indexation divided by the rate of change in the consumer price index over the same period.[23] The average degree of indexation is 0.5888, implying that indexation on average compensates the workers for 58.9% of the decrease in their real wages resulting from inflation.

The annual real interest rate is published in the *Bank of Israel Annual Report* (1978–95). If r_a denotes the real interest rate in year a, we calculate the monthly real interest rate for each month in year a as $r_m = (1 + r_a)^{1/12} - 1$. The average monthly real interest rate is 1.53% and ranges from -1% to 5.4%.

We calculate the fraction of a nominal shock that is transmitted to the real wage as $\gamma_m = (1 - \theta)(1 + \hat{\mu}_m)[1 + (1 - \theta)\hat{\mu}_m]$ for a contract with starting date in month m. The average γ_m is 0.4035, and it varies from zero (when $\theta = 1$) to 0.9663. The value of z for a contract with starting date in month m is obtained as $z = (1 + \gamma_m x_m)/(1 + y_m)$. The average z is 0.9916, with a standard deviation of 0.0254, indicating that the distribution of z is very concentrated. As the empirical measure of z^*, we use the value of z, which maximizes $\bar{B}(z)$, $z^* = (2 + \hat{\xi}_m + r_m)/[2(1 + r_m)]$.[24] The average z^* is 0.9916, and 48% of the contracts have $z < z^*$, with the remaining 52% having $z > z^*$.

An upcoming election to the Knesset is an additional indicator of increased real uncertainty as the election of a new government may presage important economic and political changes. This is of particular importance for workers employed in the public sector, where budgets are determined by a political process; upcoming elections are therefore accompanied by more real uncertainty for workers in the public sector than for workers in the private sector.

We use a dummy variable equal to one if the previous contract expires less than 1 year before elections, and equal to zero otherwise. In our sample, 22.4% of the previous contracts expire less than 1 year before elections. Similar to the variance of the real shocks, we expect upcoming elections to have a negative effect on the delay.

We also use the annual unemployment rate when the previous contract expired as an explanatory variable.[25] This is a traditional measure of labor-

[23] The degree of indexation in a contract is therefore measured on average, which is the appropriate measure for determining the real wage during a delay. The marginal degree of indexation is preferable for determining the impact of the uncertain future value of money, but we assume that the degree of indexation in a contract is constant so that the average and marginal degrees of indexation are identical.

[24] For simplicity, we set $\rho = 0$.

[25] The annual unemployment rate, published in the *Statistical Abstract of Israel* (1978–95b), is the average of the estimated quarterly unemployment rates based on the Labor Force Surveys.

market tightness (Vroman 1989; Murphy 1992). Since a high unemployment rate is likely to increase a worker's risk of being fired and might also lead to policy interventions, it is associated with more real uncertainty. In our sample, the average unemployment rate when the previous contract expired is 7.05%, with a minimum of 2.88% and a maximum of 11.8%.

To account for differences in the constraints of the economic environment and the labor-relations culture, we include dummy variables to capture the effects of the five economic branches in the economy.[26] We separate between the private and public sectors as we expect systematic differences. The public sector, being less exposed to the vicissitudes of market forces than the private sector, should react less to real uncertainty stemming from productivity shocks. Similarly, it should be less affected by the unemployment rate. At the same time, an upcoming election likely indicates more real uncertainty stemming from political changes for public- than private-sector workers, and we therefore surmise that the effect may be more pronounced in the public sector.[27]

VII. Econometric Estimation

To estimate our theoretical model, we run cross-sectional time-series regressions of the following type:

$$\mathrm{DELAY}_{tj} = \alpha + \beta X_t + \delta Y_j + \phi W_{tj} + \epsilon_{tj},$$

where DELAY_{tj} is the delay for a contract starting at time t for firm j, α is a constant, X_t is a row vector of time-varying regressors, Y_j is a row vector of time-invariant regressors, W_{tj} is a row vector of interactions between X_t and Y_j variables, and ϵ_{tj} is a disturbance term. The nature of the distribution of ϵ_{tj} determines the choice of the estimation model, which potentially could be either random effects, fixed effects, or ordinary least squares. Statistical tests indicate the superiority of the random-effects model for our data, which is therefore our preferred estimation method.

The random-effects estimation does not require independence of the ϵ_{tj} disturbance terms. The method assumes that $\epsilon_{tj} = u_j + e_{tj}$, where u_j and

[26] We also experimented with a variable for days lost because of strikes in the whole economy during the year the previous contract expired, but we found no significant effect. We do not use the year of the contract as an explanatory variable, since it might obscure some of the effects of the macroeconomic variables. However, when we did try to include the year as a variable, we obtained a significant negative estimate of its coefficient, similar estimates of the other coefficients, and a higher R^2.

[27] Elections may make the government more inclined to grant wage rises to employees in the public sector, which may shorten the negotiations and, hence, the delays. Since employers in the private sector, at least to some extent, have to match the wages in the public sector, a similar, but possibly smaller, effect may also be found there. These arguments also imply that elections should have a negative effect and that the effect might be more pronounced in the public sector.

e_{tj} are classical disturbance terms. The disturbance term u_j is a firm-specific constant that is randomly distributed across firms and is independent of e_{tj} and time. It is also assumed that u_j is uncorrelated with all explanatory variables.[28] We use the Hausman (1978) test for the null hypothesis that there is no correlation between u_j and the observed explanatory variables, which is a requirement of the random-effects model (Nerlove 2002, 38). Since our panel data set is unbalanced, we use the "feasible" version of generalized least squares to estimate the random-effects model.[29]

In contrast to the random-effects model, the fixed-effects model assumes that u_j is a firm-specific constant, that is, that the intercept term differs among firms. The simplest estimation of models with fixed effects includes a dummy variable for each firm in the sample, which is identical to taking deviations from firm means and then estimating an ordinary least squares regression. In our case, the random-effects model is more satisfactory than the fixed-effects model because the latter would exclude all firms that have one contract only (and therefore lose much information) and cannot estimate the coefficients of the time-invariant variables. The fixed-effects model also assumes that all firms are represented, which is not satisfied by our sample, which in turn is drawn from a large population of firms.

The ordinary least squares model assumes independence of ϵ_{tj}. This does not appear realistic given that we have repeated observations for the same firm. Breusch and Pagan (1980) have devised a Lagrangian multiplier test for the ordinary least squares model versus the random-effects model based on the ordinary least squares residuals. The null hypothesis is that the variance of the u_j's vanishes, and rejection of the null hypothesis means that there is evidence in favor of the error structure of the random-effects model.

In order to test the hypotheses derived from our model, we specify the following time-variant variables (X_t): z, $D(z - z^*)$, $\gamma^2\sigma_x^2$, σ_y^2, ELEC (a dummy variables for elections), and UNEMP (the unemployment rate; for simplicity, we omit the time subscripts). The delay is longest for $z = z^*$ and is first increasing and then decreasing in z. To estimate the relationship we use a piece-wise linear approximation. Let D be a dummy variable that equals unity if $z > z^*$ and equals zero otherwise. In the regression analysis we enter the variable $D(z - z^*)$ in addition to z in

[28] If this is not the case, there is an omitted-variables problem, and estimates would be biased.

[29] The generalized least squares estimation weighs the observations in inverse relationship to their variances. Since the variances of the disturbance terms are unknown, a two-stage estimation procedure is used to accomplish the weighing. In the first stage, ordinary least squares is run, and the residuals are then used to calculate estimates of the variances. These variance estimates are used in the second stage to obtain the generalized least squares parameter estimates.

order to distinguish between the effects of z for $z < z^*$ and for $z > z^*$. The estimate of the nominal uncertainty that affects the real wage with an immediate renewal is $\gamma^2\sigma_x^2$, which is γ^2 times the variance of the shock to the value of money. The estimate of the real uncertainty that affects the real wage with a delayed renewal is σ_y^2, which is the variance of the real shock. In addition, we specify the cross-section (time-invariant) variable (Y_j): BRNCH for the firm's economic branch (for simplicity, we omit the firm subscripts). If the private and public sectors are considered separately, we then have the following estimation equation for each sector:

$$\text{DELAY} = \hat{\alpha} + \hat{\beta}_1 z + \hat{\beta}_2 D(z - z^*) + \hat{\beta}_3\gamma^2\sigma_x^2 + \hat{\beta}_4\sigma_y^2 + \hat{\beta}_5\text{ELEC} + \hat{\beta}_6\text{UNEMP} + \hat{\delta}'\text{BRNCH}.$$

Our model yields the following predictions:

$\hat{\beta}_1 > 0$. This is the positive effect of z on the delay if $z < z^*$.
$\hat{\beta}_1 + \hat{\beta}_2 < 0$. This is the negative effect of z on the delay if $z > z^*$.
$\hat{\beta}_3 > 0$. The nominal uncertainty affects the delay positively, and a higher gamma is equivalent to larger absolute values of the shocks to the value of money and, hence, to more nominal uncertainty.
$\hat{\beta}_4 < 0$. The real uncertainty affects the delay negatively.
$\hat{\beta}_5 < 0$. Elections are another indicator of real uncertainty as the economic policy may change if or when a new government is elected. We therefore expect a negative coefficient.
$\hat{\beta}_6 < 0$. A high unemployment rate is associated with more real uncertainty and is therefore expected to have a negative effect on the delay.

The model provides no a priori predictions about $\hat{\delta}'$ for the economic branches, which serve as controls.

To examine the empirical validity of the predictions, we first run a regression for the length of delay in contract renewal based on the private-sector sample. To check for differential effects in the private and public sectors, we then run an extended regression of the pooled sample in which we include interaction terms between the public sector and each of the explanatory variables. A significant interaction term indicates that the interacted explanatory variable affects the delay differently in the two sectors.[30] The model predicts:

The coefficient of the interaction between the public sector and real uncertainty should be positive: real uncertainty shortens the delay in the public sector, but less than in the private sector.
The coefficient of the interaction between the public sector and elections

[30] Alternatively, we could run separate regressions for the two sectors and test for significant coefficient differences.

should be negative: elections shorten the delay in the public sector more than in the private sector.

The coefficient of the interaction between the public sector and the unemployment rate is positive: unemployment shortens the delay in the public sector less than in the private sector.

A. The Length-of-Delay Regressions

In table 6, regressions 1 and 2, we present the random-effects coefficient estimates of the length-of-delay regressions, using the firm for the cross-section index and assuming that $\rho = 0$.[31] Regression 1 is based on the 1,731 contracts with firms in the private sector only. Regression 2 is based on the pooled sample of 2,103 contracts with firms in the private and public sectors, and it includes interaction terms between the public sector and each explanatory variable. As the coefficients for the private sector are very similar in regressions 1 and 2, and there is more information in regression 2, we discuss the results from this regression only in detail.[32]

The Hausman test shows that the key assumption in the random-effects model, namely, that the u_j disturbance term is uncorrelated with the explanatory variables, is satisfied.[33] The Breusch-Pagan (1980) test indicates that the random-effects model is preferable to the ordinary least squares model. The large values of chi-square is a sign that a basic assumption of the ordinary least squares model, namely, that $\mathrm{Var}(u_j) = 0$, is violated. Accordingly, taken together, the Hausman and Breusch-Pagan tests show that the random-effects model is the correct specification of the delay function.

The coefficient estimates provide strong support for the theoretical model. They all have the predicted signs and are highly significant. This is true for both the private and the public sectors, the only difference being in the magnitude of the effect of some of the explanatory variables. The uninteracted variables refer to the private sector, and we start by discussing their estimated effects.[34]

[31] We have also tried to specify the economic branch for the cross-section index. This leads to basically similar results and generally higher significance levels for the coefficients of the key variables.

[32] The regressions are based on all 2,103 contracts with zero and positive delay. Using only the 1,890 contracts with positive delay (90%) leads to similar results. We also estimated the delay regressions using only the 1,850 contracts with more than 1 week of delay (88%). The results are again similar. The same holds true for the likelihood-of-delay estimation.

[33] $(\mathrm{Prob} > \chi^2) = 0.1365$ in table 6, regression 1, and $(\mathrm{Prob} > \chi^2) = 0.0973$ in table 6, regression 2.

[34] For the numerical illustrations we use the means of the macrovariables from table 5, which are based on the pooled sample. The means in the private and public sector are very similar since the contracts in the two sectors are over the same years.

Table 6
Random-Effects Delay Regressions Israeli Labor Market, 1978–95

	Coefficient (z-Value)		
Independent Variables	Length of Delay in Private Sector (1)	Length of Delay in Private and Public Sector, with Public-Sector Interactions (2)	Likelihood of Delay in Private and Public Sector, with Public-Sector Interactions (3)
z	5.471 (2.66)	5.998 (2.60)	−.010 (.19)
$D(z-z^*)$	−30.910 (5.79)	−32.272 (5.35)	−.362 (2.85)
γ^2 × expected variance of nominal shock	19.031 (6.07)	19.484 (5.51)	.263 (2.88)
Expected variance of real shock	−6.029 (7.85)	−6.122 (7.11)	−.065 (4.01)
Elections	−28.685 (2.87)	−30.003 (2.67)	−.287 (1.22)
Unemployment rate (%)	−16.827 (7.73)	−17.619 (7.28)	−.230 (4.39)
Economic branch:			
Public services	−85.512 (1.92)	−82.821 (1.69)	1.080 (.81)
Manufacturing	−21.460 (1.95)	−19.477 (1.45)	.777 (2.97)
Commerce	−15.504 (.62)	−8.514 (.32)	.641 (1.04)
Banking institutions	60.518 (1.38)	82.421 (1.91)	1.322 (1.29)
Public sector		−12.301 (.36)	−1.625 (.09)
Public sector × z		−2.793 (.49)	.026 (.14)
Public sector × $D(z-z^*)$		−32.311 (1.97)	−.489 (1.31)
Public sector × γ^2 × expected variance of nominal shock		14.540 (1.49)	.092 (.24)
Public sector × expected variance of real shock		4.341 (1.94)	−.023 (.47)
Public sector × elections		−68.930 (2.33)	.752 (.68)
Public sector × unemployment rate (%)		18.309 (3.59)	.104 (.61)
Constant	−193.222 (.98)	−241.018 (1.09)	5.254 (.98)
R^2	.0958	.1312	
χ^2 for Hausman test	14.88	28.54	
χ^2 for Breusch-Pagan (1980) test	268.97	203.35	
χ^2 for likelihood-ratio test			76.47
Sample size	1,731	2,103	2,103

Note.—STATA 7.0 is used for estimation. The values of z and of $D(z-z^*)$ are measured in percent. The expected variances of the nominal and real shocks are measured in (percent)^2. The firm is used as the cross-section identifier. The reference group for economic branch is private services. The null hypothesis of the Hausman test is that there is no correlation between u_j and the explanatory variables, in which case the random-effects specification is correct. The null hypothesis of the Breusch-Pagan (1980) test is that var(u_j) = 0, in which case ordinary least squares estimates would be better than random-effects estimates. The null hypothesis of the likelihood-ratio test is that the panel estimator is the same as the pooled standard probit estimator. Regressions 2 and 3 also include interactions between the public sector and the economic branches. These are not reported as none of their coefficients is significant.

Recall that z is the ratio between the real wage with delay and what the real wage would be without delay; the delay first increases and then decreases in z, with a peak when the shocks are such that $z = z^*$. The regression results support both the increasing and the decreasing portion of this relationship. This is because $\hat{\beta}_1 + \hat{\beta}_2 = -26.274$, and a chi-square test yields $\chi^2(1) = 30.35$, which is significant at less than the 0.001 level. Accordingly, for $z < z^*$, an increase in z by 0.01 causes the delay to be lengthened by 5.998 days, while for $z > z^*$ an increase in z by 0.01 causes the delay to be shortened by 26.274 days.

The value of z depends on the realized shocks. A positive nominal shock equal to one standard deviation (0.0238) at the average γ (0.4035) leads to an increase in the real wage by $0.0238 \times 0.4035 = 0.96\%$, or approximately 1%, if the contract renewal is delayed. A positive real shock equal to one standard deviation (0.0263) leads to an increase in the per capita GDP by 2.63% and, hence, to this increase in the real wage if the contract is renewed. Consequently, for $z < z^*(z > z^*)$, a positive one-standard-deviation nominal shock increases (decreases) the delay in the average contract by about 6 days (26 days), while a positive one-standard-deviation real shock decreases (increases) the delay by about 16 days (69 days).

The effect of wage indexation on z is the same as if the absolute value of the nominal shock were reduced by a factor of $1 - \gamma$. In the face of a positive one-standard-deviation nominal shock, for $z < z^*$ wage indexation decreases the delay by $599.8 \times 0.0238 \times 0.5965 = 8.5152$ days through its effect on z, while for $z > z^*$ wage indexation increases the delay by $2{,}627.4 \times 0.0238 \times 0.5965 = 37.3$ days through its effect on z.

Since γ increases with the trend in the value of money, a higher value of the latter is equivalent to an increase in the absolute value of the nominal shock. At a positive one-standard-deviation nominal shock and the average degree of wage indexation, for $z < z^*$ the trend in the value of money on average increases the delay by $599.8 \times 0.0238 \times (0.5888 - 0.4035) = 2.6452$ days through its effect on z, while for $z > z^*$ the trend in the value of money on average decreases the delay by $2{,}627.4 \times 0.0238 \times (0.5888 - 0.4035) = 11.587$ days through its effect on z.

Turning to the nominal and real uncertainty, the regression coefficients show, as predicted by the theory, that the nominal uncertainty increases the delay, while the real uncertainty decreases the delay. At the average value of γ, a doubling of the variance of the nominal shock would increase the delay by $194{,}840 \times 0.4035^2 \times 6.8E - 4 = 21.571$ days. Without wage indexation, the delay would have increased by $194{,}840 \times 6.8E - 4 = 132.49$ days. So wage indexation, through its effect on γ, reduces the average impact of the variance of the nominal shock to only 16% of what it would have been in the absence of indexation. A doubling of the variance of the real shock would decrease the delay by $61{,}220 \times 2.3E - 4 = 14.081$ days.

If the previous contract expires during the year before an election, the delay is shorter by 30 days. An increase in the unemployment rate of one percentage point reduces the delay by 17.619 days.

Using the private-services branch as the reference group, the dummy variables of the economic branches show that the delay is longest for banking institutions, which has a coefficient of 82.421. The delay for banking institutions is therefore 39% longer than the average delay of 213.29 days in all economic branches.[35] Public services, manufacturing, and commerce are not significantly different from private services.

The interactions between the public sector and the independent variables indicate the differences between the two sectors. The only significant differences are in the effect of z on the downward-sloping portion of the relationship between z and the delay, of the variance of the real shock, of elections, and of unemployment. While we have no theoretical explanation of why the effect of z on the delay on the downward-sloping portion of the relationship between z and the delay is more negative in the public sector, the latter three differences are as predicted by the theory. In the public sector, the negative relationship between real uncertainty and delay, while still negative and significant, is weaker than in the private sector. A doubling of the variance of the real shock would decrease the delay by only $(61{,}220 - 43{,}410) \times 2.3E - 4 = 4.3263$ days in the public sector versus 14.081 days in the private sector. Upcoming elections reduce the delay in the public sector by an additional 68.930 days, and the effect of elections on delay is more than three times as strong as in the private sector. The significant positive coefficient of 18.309 for the interaction term public sector × unemployment cancels the significant negative coefficient in the private sector, so that unemployment appears to have no significant effect on the delay in the public sector.[36]

B. The Likelihood-of-Delay Regressions

In table 6, regressions 1 and 2, the delay is the dependent variable, and the regressions therefore require information about the length of the de-

[35] Negotiations in banking institutions take place among several similar bargaining pairs. The workers, therefore, have an additional incentive to delay the new contract in order to obtain better information about their own institution's ability to pay by observing the outcomes reached with the other institutions. See Gu and Kuhn (1998).

[36] A chi-square test shows that the sum of the two coefficients is not significantly different from zero.

lay.[37] We now estimate the likelihood of delay. This approach has the disadvantage that it disregards the available information about the length of delay given that there is a delay, but it is nevertheless instructive and provides a closer link to the theoretical model.

In table 6, regression 3, the results of a probit random-effects model are presented. The dependent variable is a dummy variable that takes the value of unity if the new contract is delayed and the value of zero if there is no delay. The independent variables are the same as in table 6, regression 2, and, therefore, include public-sector interactions. The hypotheses are identical to those specified in Section VII.A, except that the dependent variable is now dichotomous. The value of chi-square for the likelihood-ratio test rejects the hypothesis that the panel estimator is the same as the pooled standard probit estimator.

The estimates of the noninteracted variables are generally similar to those in table 6, regression 2, in terms of sign and significance. The main difference is that the coefficient of z is insignificant ($z = 0.19$). Additionally, the likelihood of delay is highest in manufacturing, while there is no significant difference between the other branches. All interaction terms are insignificant, implying that the likelihood of delay is similar in the two sectors.

VIII. Conclusion

In many countries, the typical labor contract has a fixed duration. Often, however, the fixed duration is not nearly as sacrosanct as it appears. In reality, there may be lengthy delays before the next contract is concluded, and during these delays the provisions of the "expired" contract remain in force. We present a theoretical model that focuses on macroeconomic factors in explaining delays in contract renewal. In particular, we emphasize the importance of the realized nominal and real shocks, as well as of the levels of nominal and real uncertainty. We show that whether the contract renewal will take place on time or be delayed can be described by an (s, S) strategy in the ratio of the real wage with a delay to what the real wage would be with a new contract. We also demonstrate that

[37] We have also tried to use relative delay ($\equiv$ delay/duration) as the dependent variable, rather than delay itself. The sign and significance of the coefficients are comparable with what is obtained with delay as the dependent variable, except that the coefficient of z is insignificant. The implied effects of the variables tend to be smaller. For example, in the pooled sample of contracts in the private and public sectors, at the average value of γ, a doubling of the variance of the nominal shock would increase the relative delay in the private sector by 0.008, which at the average contract duration corresponds with an increase in the delay by 5.0 days. A doubling of the variance of the real shock would decrease the relative delay in the private sector by 0.010, which at the average contract duration corresponds with a decrease in the delay by 6.3 days.

nominal uncertainty tends to favor delay, while real uncertainty tends to favor immediate contract renewal.

The model is tested using data from all published Israeli labor contracts signed from 1978 to 1995. The renewal is delayed for 86% of these contracts, with the average delay being 213 days. The empirical findings strongly support the theory. The coefficient estimates all have the predicted signs and are significant. In the private sector, increasing the real wage during a delay by 1% of the real wage in a new contract would lengthen the average delay by 6 days in the increasing range and would shorten the average delay by 26 days in the decreasing range. A doubling of the level of nominal uncertainty would increase the delay by 22 days, while a doubling of the level of real uncertainty would reduce the delay by 14 days. Since upcoming elections increase the real uncertainty, they also reduce the delay: if the previous contract expires during the year before an election, the delay is reduced by 29 days.

The full sample of contracts with firms in both the private and public sectors is used to gauge the different impact of the explanatory variables in the two sectors. The coefficient estimates of the interaction terms with the public sector provide further support for the theoretical model. For example, real uncertainty reduces the delay by less in the public sector, while upcoming elections reduce the delay by more than in the private sector.

The empirical variables are constructed so that they correspond closely with the variables in the theoretical model, and the period under consideration includes subperiods with extreme differences in inflation and growth rates. This enables us to get robust estimates of the model parameters. In fact, separate regressions for different subperiods yield coefficient estimates that are similar even though the economic environments are very different. Needless to say, however, it would be desirable also to test the model with data from other countries.

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