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ABSTRACT

Delays in Renewal of Labor Contracts: Theory and Evidence[*]

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.

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Delay is preferable to error.

Thomas Je®erson in letter to George Washington, May 16, 1792.

1   Introduction

Contemporary labor contracts are usually signed for a ¯xed duration. A typical labor con-
tract speci¯es its expiration date, and it is rare that a contract has no preset duration or
includes a clause permitting reopening or extension of the contract. Nevertheless, the du-
ration of labor contracts is more °exible than their rigid wording seems to indicate. In
particular, it is common practice that the terms of the old contract are automatically ex-
tended 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 expira-
tion 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) ¯nd that 47% of contract renewals take place two 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 ma-
jor Canadian collective bargaining contracts, Gu and Kuhn (1998) ¯nd a holdout incidence
of 81%, and that the average holdout is 80 days.[1]

Fixed-duration labor contracts play a pivotal role in many models of aggregate °uctu-
ations. Starting with Fischer (1977), Phelps and Taylor (1977), and Taylor (1979, 1980),

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[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 sample).
Contract reopenings may be considered negative delays and are the focus of Danziger (1995).

The law in Israel (and similarly in the U. S. 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/lockout. This is also true in many other countries,
with the major exception of the U. K., where labor contracts are not legally binding. See also Holden (1994).

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, cf. Caplin and Spulber (1987), who 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, dependency of the delay on 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 small shocks, and the former are incorporated into new contracts earlier. This mitigates and possibly neutralizes the differential real impacts the shocks would have if contract durations were truly fixed.

In view of the above, the purpose of this paper 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 four-period economy in which a labor contract between a union and a firm expires at the end of period two, at which time the parties may either immediately conclude a new labor contract for periods three and

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[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 et al. (1984), Christofides (1985), Vroman (1989), Murphy (1992), Wallace and Blanco (1991), and Rich and Tracy (2000). Flanagan (1999) surveys the literature concerned with the relationship between collective bargaining systems and macroeconomic performance.

four, or delay the renewal until the end of period three. The degree of wage indexation is exogenous to the parties negotiating the individual labor contracts. This re°ects 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 ¯rm is risk neutral, and the discounted expected real payments to a worker in periods three and four 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 four are known at the time of contract renewal. Since the labor contract sets the (partially indexed) nominal wage for period four, the workers become exposed to the nominal uncertainty (that is, the uncertainty in the value of money) in period four, but are fully insured against the real uncertainty (that is, 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 four 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 four (except for the adjustment for under- or overpayment in period three); the workers are then fully indemni¯ed for nominal uncertainty, but fully exposed to real uncertainty. Hence, for given realizations of the shocks in periods two and three, nominal riskiness favors a delay, while real riskiness favors an immediate renewal.[4]

The realizations of the shocks in period two and three determine how much the workers are under- or overpaid given the real value of their marginal product in period three if the contract renewal is delayed. A large under- or overpayment would pull toward immediate

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[3] The exogenous wage indexation simpli¯es the model, but is not essential for the results. If the degree of indexation could be freely negotiated, it would depend on the correlation between the price index and the ¯rm's output and input prices. See Ehrenberg et al. (1984), Card (1986), and Danziger (1988, 1992).

[4] Since a delay shortens the time period during which the provisions of the next contract will be in force, the opposite e®ects of the two types of riskiness on the decision to delay are similar to the ¯nding 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 the present paper to model the joint nature of the decision to delay and the duration of the next contract.

renewal. Accordingly, the ratio of the real wage in period three 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 and 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 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

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

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 to 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 obtain 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 paper 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 delay is a nonmonotonic function that first increases and then decreases in this ratio. In the Cramton-Tracy and Gu-Kuhn models the delay always increases in this variable.

2 The Model

Consider a four-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 one 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 $(1 + \iota)(1 + x_i)$, where $\iota > -1$ 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 $[\underline{x}, \overline{x}]$, 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 $(1 + \nu)(1 + y_i)$, where $\nu > -1$ 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 $[\underline{y}, \overline{y}]$, 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 correlated.

In period one, the union and the firm conclude a two-period labor contract covering periods one and two. The contract sets the nominal wage for period one at A_1, and the real

wage is then also A_1. The nominal wage for period two is determined by a base wage b, which is partially indexed to the price level. The degree of indexation is exogenously fixed at μ, $0 < \mu < 1$. Hence, the nominal wage in period two is set at $b\{1 - \mu + \mu[(1 + \gamma)(1 + x_2)]\}$, and the real wage is $b[(1 - \mu)(1 + \gamma)(1 + x_2) + \mu]$. The base wage is set to make the expected real wage equal to the expected real value of the marginal product,[6] that is,

$$b \int_{\underline{x}}^{\overline{x}} [(1 - \mu)(1 + \gamma)(1 + x_2) + \mu] f(x_2) dx_2 = A_1(1 + \omega)$$

$$, \quad b = \frac{A_1(1 + \omega)}{1 + (1 - \mu)\gamma}.$$

The real wage in period two therefore becomes

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

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

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

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[6] It is only for simplicity that the model assumes that the real wage in period one and the expected real wage in period two equals the real value of the marginal product in period one and the expected real value of the marginal product in period two, respectively. 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.

Alternatively, the union and the ¯rm may delay contract renewal to period four, in which case the provisions of the old contract are automatically extended to cover period three.[7] The relative change in the base wage between periods two and three is similar to that between periods one and two, and the base wage is again partially indexed to the price level. Accordingly, the nominal wage in period three is set at $A_1(1 + »)^2 f1 ¡ \mu + \mu=[(1 + {}^1)(1 + x_2)]g f1 ¡ \mu + \mu=[(1 + {}^1)(1 + x_3)]g[1 + (1 ¡ \mu)^1]^{i\ 2}$, and the real wage becomes $A_1(1 + »)^2(1 + °x_2)(1 + °x_3)$. Let

$$z \, ´ \, \frac{(1 + °x_2)(1 + °x_3)}{(1 + y_2)(1 + y_3)}$$

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 three can then be written as A_3z. 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 four, is made retroactive to period three. As the price level and the real value of the marginal product in period four are now known, the wage in period four 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 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 four is $A_4 + A_3(1 ¡ z)(1 + r) = A_3[(1 + »)(1 + y_4) + (1 ¡ z)(1 + r)]$.[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 one and two covered by the ¯rst contract is independent of the time at which the

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

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

[9] It is assumed that $(1 + »)(1 + \underline{y})=(1 + r) + 1 > [(1 + °\overline{x})=(1 + \underline{y})]^2$, so that the real wage is positive.

next contract is concluded. The expected utility from periods three and four, however, depends on when the new contract is concluded. Let $\frac{1}{2} > \i 1$ denote a worker's discount rate. On the one hand, with an immediate renewal, the discounted expected utility from periods three and four is

$$\ln A_3 + \frac{1}{1 + \frac{1}{2}} \int_{\underline{x}}^{\overline{x}} \ln[A_3(1 + \gg)(1 + {}^\circ x_4)] f(x_4) dx_4$$
$$= \ln A_3 + \frac{1}{1 + \frac{1}{2}} \left(\ln[A_3(1 + \gg)] + \int_{\underline{x}}^{\overline{x}} \ln(1 + {}^\circ x_4) f(x_4) dx_4 \right) :$$

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

$$\ln(A_3 z) + \frac{1}{1 + \frac{1}{2}} \int_{\underline{y}}^{\overline{y}} \ln f A_3[(1 + \gg)(1 + y_4) + (1 \i z)(1 + r)] gg(y_4) dy_4$$
$$= \ln(A_3 z) + \frac{1}{1 + \frac{1}{2}} \left(\ln[A_3(1 + \gg)] + \int_{\underline{y}}^{\overline{y}} \ln \left[1 + y_4 + \frac{(1 \i z)(1 + r)}{1 + \gg} \right]^{\#} g(y_4) dy_4 \right) :$$

The bene¯t from a delay depends on z and is obtained by subtracting the discounted expected utility from periods three and four of a contract which is concluded immediately from the discounted expected utility from periods three and four of a contract with delayed renewal,

$$B(z) \acute{} \ln z + \frac{1}{1 + \frac{1}{2}} \int_{\underline{y}}^{\overline{y}} \ln \left[1 + y_4 + \frac{(1 \i z)(1 + r)}{1 + \gg} \right]^{\#} g(y_4) dy_4 \i \frac{1}{1 + \frac{1}{2}} \int_{\underline{x}}^{\overline{x}} \ln(1 + {}^\circ x_4) f(x_4) dx_4 :$$

Since the utility function is logarithmic, the bene¯t from a delay is independent of a possible correlation between nominal and real shocks. Furthermore, the nature of indexation implies that the size of the nominal shocks and the fraction $^\circ$ of these shocks transmitted to the real wage enter only multiplicatively into the bene¯t from a delay { as $^\circ x_2$ and $^\circ x_3$ in z, and as $^\circ x_4$. Accordingly, a change in the nominal shocks by a factor of $\_ > 0$ together with a simultaneous change in the degree of indexation or in the trend in the value of money such that $^\circ$ changes by a factor of $1=\_$ would have no impact on $B(z)$.

The ¯rm is risk neutral and since its discounted expected real pro¯ts per worker are the same with and without a delay, it agrees that the timing of the contract renewal is chosen

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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 $z = z^{\alpha}$ deﬁned by

$$\int_{\underline{y}}^{\bar{y}} \left(\frac{z^{\alpha}(1 + r)}{[(1 + \omega)(1 + y_4) + (1 - z^{\alpha})(1 + r)](1 + \frac{1}{2})} \right) g(y_4) dy_4 = 1. \tag{1}$$

This reﬂects that if $z < z^{\alpha}$, larger nominal shocks or smaller real shocks in periods three and four bring the real wage in period three with a delay closer to maximizing the discounted expected utility with a delay; if $z > z^{\alpha}$, 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, $B(z^{\alpha}) > 0$. Also, it is assumed that the shocks are suﬃciently dispersed that there are shocks for which $z < z^{\alpha}$ and it is optimal to immediately renew the contract, $Bf[(1 + °\underline{x})=(1 + \bar{y})]^2 g < 0$, as well as shocks for which $z > z^{\alpha}$ and it is optimal to immediately renew the contract, $Bf[(1 + °\bar{x})=(1 + \underline{y})]^2 g > 0$. 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 two and three, the renewal is delayed if $z \in (s; S)$, is immediate if $z \notin (s; S)$, 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]

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[10] In the special case of no real uncertainty and $\frac{1}{2} = 0$, then $B(z)$ is symmetric around $z^{\alpha} = \frac{1}{2}[1 + (1 + \omega)=(1 + r)]$ and $B(z^{\alpha}) = \ln f(2 + r + \omega)^2 = [4(1 + \omega)(1 + r)]g - \int_{\underline{x}}^{\bar{x}} \ln(1 + °x_4)f(x_4)dx_4$. A worker is then indi®erent between receiving the real wage $z^{\alpha} - \acute{A}$, $\acute{A} > 0$, in period three and the expected real wage $z^{\alpha} + \acute{A}$ in period four, and vice versa. Since the left-hand side of eq. (1) increases with z^{α} and is a convex function of y_4, real uncertainty leads to a decrease in z^{α}.

3 The Beneᶠt from a Delay and the Critical Values of z: Comparative Statics

It has been shown that the beneᶠt from delay depends on the realized shocks as manifested in z: the beneᶠt increases with z if $z < z^{\mu}$ and decreases with z if $z > z^{\mu}$. The beneᶠt therefore increases with the nominal shocks if $z < z^{\mu}$ and decreases with the nominal shocks if $z > z^{\mu}$, while the real shocks have the opposite e®ect. We now examine how the parameters a®ect the beneᶠt for a given z, and also study the e®ects 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 e®ect on the discounted expected utility from a contract with delayed renewal, for a given z. The reason is that a worker in period three receives the real wage $A_3 z$, which is given since A_3 and z are given; in period four he receives a real wage equaling the real value of the marginal product plus retroactive pay for period three, and therefore independent of the change in the value of money from period three to four. Thus, delayed renewal protects the worker from any real e®ects of anticipated and unanticipated changes in the value of money from period three to four.

However, the same nominal parameters a®ect the discounted expected utility if the contract is renewed immediately in period three. In this case, an increase in the nominal uncertainty increases the uncertainty of the real wage in period four, which reduces the risk-averse worker's discounted expected utility. The magnitude of the impact of the nominal uncertainty depends on the fraction ° of period four's nominal shock transmitted to the real wage. Since ° increases with ¹ and decreases with µ, an increase in ¹ is equivalent to an increase in uncertainty, while an increase in µ 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 beneᶠt from a delay, while the degree of indexation increases the discounted expected utility, decreasing the beneᶠt 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,[11] but does affect the fourth period's wage in both contracts. If the contract is renewed in period three, the fourth period's wage is proportional to $1 + $ ». If the renewal is delayed and $z = 1$ (so that there is no retroactive pay in period four), the fourth period's wage is also proportional to $1 + $ ». In this case the real wage in period three is the same irrespective of whether or not the contract renewal is delayed. The benefit from a delay is therefore independent of », and $B(1)$ does not change with ».

If the renewal is delayed and $z > 1$, the real wage in the third period is greater and the expected real wage in the fourth period is less than if the contract is renewed in period three. Due to the gains from a smoother intertemporal consumption, an increase in » is therefore more beneficial if the contract renewal is delayed. If the renewal is delayed and $z < 1$, the opposite is true. Consequently, » affects the benefit from a delay positively for $z > 1$ and negatively for $z < 1$. It follows that ds/d» $\gtrless 0$ as $s \lessgtr 1$ and dS/d» $\gtrless 0$ as $S \lessgtr 1$.[12]

Real uncertainty has no effect on the discounted expected utility from a contract concluded in period three, since the expected real wage in period four then equals the expected real value of worker's marginal product and is independent of the real uncertainty. The worker is fully insured against real uncertainty in period four. However, uncertainty reduces the expected utility from a delayed contract, since the real wage in period four becomes exposed to the real uncertainty. Accordingly, an increase in real uncertainty decreases the

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[11] We ignore the effects of » 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 $1 \cdot s < S$; narrows the range if $s < S \cdot 1$; and moves the range to the right if $s < 1 < S$.

bene⁻t 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 e®ect of the real interest rate works through the value of the retroactive pay in period four of a delayed contract. If the retroactive pay is positive (negative), the real wage in period four increases (decreases) with the real interest rate. It follows that r has the opposite e®ect of » on the bene⁻t of a delay. So, if z = 1, the bene⁻t 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 ⪋ 0 as s ⪋ 1 and dS=dr ⪋ 0 as S ⪌ 1.[13]

The comparative-static relationships are summarized in Table 1.

4 The Israeli Economy 1978-1995

The development of some of the major macroeconomic variables in Israel during 1978-1995 is summarized in Table 2. During this time span, the Israeli economy experiences two distinct periods of in°ation: The annual in°ation rate climbs from 48.1% in 1978 to about 800% in May 1985 (with a highly variable monthly in°ation); thereafter, the stabilization program, enacted in May 1985, reduces the in°ation rate to 19.7% in 1986 and further to 8.1% in 1995. The average annual in°ation rate for the entire period is around 60%. This is illustrated in Chart 1.

The growth rate of per-capita GDP exhibits no clear trend. The growth rate varies 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 Chart 2. The annual real interest rate exhibits a hump-shaped pattern, similar to the pattern of the in°ation 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 rises from an annual average of about 3% in 1978-79 to a peak of 11.2% in 1992 (due

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

to the mass immigration during 1989-1992). The trend then changes and unemployment decreases during 1993-1995. The number of work days lost due to 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.

5   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, result from negotiation by the parties and not from legislative intervention. There are 92 (4% of total) new contracts signed before the previous contracts terminated and therefore exhibiting a negative delay. These contracts are excluded, leaving a sample of 2,103 contracts which are used in the empirical study. They stem from 711 different firms (including public-sector employers), since 325 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]   About 82% of the contracts are in the private sector (1,731 contracts) and 18%

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[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.

[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.

are in the public sector (372 contracts). Concerning the economic branches, over half of the contracts are in Manufacturing, over a quarter in Private Services, 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 di®erence 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 one week. The average delay is 213 days. The delay varies greatly, from zero to 1,529 days with a standard deviation of 207 days. Chart 3 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. So the public sector is characterized by longer and more dispersed delays in contract renewals.

We have no information about whether a contract settlement is reached following the onset of a strike, or when such 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 1983-1992 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 di®erence 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 one year (21% of the contracts) and at two years (52% of the contracts). Only 10% of the contracts exceed two years, and a mere 3% exceed three years. Contracts in the public sector are longer than in the private sector: 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 two years and 2% for more than three years. The corresponding percentages for the public sector are more than double: 19% for

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more than two years and 11% for more than three years.

6 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

$$\ln z + \frac{1}{1+\frac{1}{2}} \ln \left[1 + \frac{(1 - z)(1 + r)}{1 + \gg} \right] - \frac{\sigma_y^2}{2(1+\frac{1}{2})[1 + (1 - z)(1 + r)/(1 + \gg)]^2} + \frac{\circ 2\sigma_x^2}{2(1+\frac{1}{2})};$$

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

$$\tilde{B}(z) \approx \ln z + \frac{1}{1+\frac{1}{2}} \ln \left[1 + \frac{(1 - z)(1 + r)}{1 + \gg} \right] - \frac{\sigma_y^2}{2(1+\frac{1}{2})} + \frac{\circ 2\sigma_x^2}{2(1+\frac{1}{2})};$$

In the empirical model, we interpret $\tilde{B}(z)$ first as the delay and afterwards 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 $\iota_m \approx (p_{m-1}/p_m) - 1$, where p_{m-1} and p_m are the consumer price indexes in months $m - 1$ and m.[18] We use AR(6) to estimate the trend in the value of money from month $m - 1$ to month m,

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[17] The values are calculated over the 2,103 contracts. The values of the macro variables calculated over the years 1978-1995 are similar.

[18] The monthly consumer price index is published in the Monthly Bulletin of Statistics.

$\uparrow_m$.[19]   The estimated average $\uparrow_m$ is $\lnot 0{:}0385$, with a standard deviation of 0.0384.

The estimated nominal shock is the di®erence between the rate of change in the actual value of money and the estimated trend in the value of money, $x_m \, ´ \, {}^1m \, \lnot \, \uparrow_m$. 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 six 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,[20] and we calculate the rate of change in the actual per-capita GDP in a branch for each month m in year a as $»_m = [(GDP_a{=}GDP_{a\lnot 1}) \, \lnot \, 1]^{1{=}12}$, where GDP_a and $GDP_{a\lnot 1}$ are the per-capita GDP in the branch in year $a \, \lnot \, 1$ and a. We use AR(2) to estimate the trend in branch per-capita GDP from year $a \, \lnot \, 1$ to year a, $\hat{»}_a$,[21] and calculate the trend in the branch per-capita GDP for each month m in year a as $\hat{»}_m \, ´ \, \hat{»}_a^{1{=}12}$. The average branch-speci¯c 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, $y_m \, ´ \, »_m \, \lnot \, \hat{»}_m$. As an estimate of the expected variance of the shock in each month of the year, $¾^2_{ym}$, we use the (moving-average) variance of the shocks in the months of the previous

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[19] We also tried to estimate the trend in the value of money by six-month moving arithmetic or geometric averages, which were very similar. The estimations were repeated using AR(12). While the estimates are somewhat di®erent, they lead to similar conclusions in the regression analyses.

 In addition, we experimented with six- and twelve-months partly backward- and partly forward-looking estimates. These perform less well in the regression analyses.

[20] The annual branch GDP is published in the Monthly Bulletin of Statistics.

[21] Similarly 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 signi¯cance of the coe±cients in the delay regressions.

two 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, $\hat{»}_m$, ranges from {2.24% to 5.27%, with an average monthly trend for 1978-1995 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) extends 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 in°ation rate. Empirically, the degree of indexation is measured as the rate of change in the wage during the contract period due to 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 due to in°ation.

The annual real interest rate is published by the Bank of Israel.[24] 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%.

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[22] We have also experimented with using economy-wide GDP productivity measures, which have the advantage that the GDP ¯gures 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.

[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] See Bank of Israel Annual Report.

We calculate the fraction of a nominal shock which is transmitted to the real wage as $\gamma_m = (1 - \mu)(1 + \kappa_m)[1 + (1 - \mu)\kappa_m]$ for a contract with starting date in month m. The average γ_m is 0.4035, and it varies from zero (when $\mu = 1$) 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 $z^\ast$ we use the value of z which maximizes $\overset{.}{B}(z)$, $z^\ast = (2 + \hat{x}_m + r_m)/[2(1 + r_m)]$.[25] The average $z^\ast$ is 0.9916, and 48% of the contracts have $z < z^\ast$ with the remaining 52% having $z > z^\ast$.

Upcoming elections 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 one year before elections, and equal to zero otherwise. In our sample, 22.4% of the previous contracts expire less than one year before elections. Similarly 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.[26] This is a traditional measure of labor-market tightness.[27] 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%.

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[25] For simplicity, we set $\psi = 0$.

[26] The annual unemployment rate, published in the Annual Statistical Abstracts of Israel, is the average of the estimated quarterly unemployment rates based on the Labor Force Surveys.

[27] See, for example, Vroman (1989) and Murphy (1992).

To account for di®erences in the constraints of the economic environment and the labor-relations culture, we include dummy variables to capture the e®ects of the ¯ve economic branches in the economy.[28]

We separate between the private and public sectors as we expect systematic di®erences. The public sector, being less exposed to the vicissitudes of market forces than the private sector, should react less to real uncertainty. Similarly, it should be less a®ected by the unemployment rate. At the same time, upcoming elections likely indicates more real uncertainty for public- than private-sector workers, and we therefore surmise that the e®ect may be more pronounced in the public sector.[29]

7   Econometric Estimation

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

$$DELAY_{tj} = ® + ^-X_t + ±Y_j + ÁW_{tj} + ²_{tj};$$

where $DELAY_{tj}$ is the delay for a contract starting at time t for ¯rm j, ® 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 $²_{tj}$ is a disturbance term. The nature of the distribution of $²_{tj}$ determines the choice of the estimation model, which potentially could be either random-e®ects, ¯xed-e®ects, or ordinary-least-squares. Statistical tests indicate the superiority of the random-e®ects model for our data, which is therefore

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[28] We also experimented with a variable for days lost due to strikes in the whole economy during the year the previous contract expired, but found no signi¯cant e®ect. We do not use the year of the contract as an explanatory variable, since it might obscure some of the e®ects of the macroeconomic variables. However, when we did try to include the year as a variable, we obtained a signi¯cant negative estimate of its coe±cient, similar estimates of the other coe±cients, and a higher R^2.

[29] 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, e®ect may also be found there. These arguments also imply that elections should have a negative e®ect, and that the e®ect might be more pronounced in the public sector.

our preferred estimation method.

The random-effects estimation does not require independence of the ε_{tj} disturbance terms. The method assumes that $\varepsilon_{tj} = u_j + e_{tj}$, where u_j and e_{tj} are classical disturbance terms. The disturbance term u_j is a firm-specific constant that is randomly distributed across firms, and independent of e_{tj} and time. It is also assumed that u_j is uncorrelated with all explanatory variables.[30] We use the Hausman (1978) test for the null hypothesis that there is no correlation between u_j and the observed explanatory variables, a requirement of the random-effects model (Nerlove 2002, p. 38). Since our panel data set is unbalanced, we use the \feasible" version of generalized-least-squares to estimate the random-effects model.[31]

In contrast to the random-effects model, the fixed-effects model assumes that u_j is a firm-specific constant, i.e., 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 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

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

[31] 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.

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, $D(z - z^\pi)$, $\sigma^2\sigma_x^2$, σ_y^2, ELEC, and UNEMP (for simplicity, we omit the time subscripts). The delay is longest for $z = z^\pi$, 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 which equals unity if $z > z^\pi$ and equals zero otherwise. In the regression analysis we enter the variable $D(z - z^\pi)$ in addition to z in order to distinguish between the effects of z for $z < z^\pi$ and for $z > z^\pi$. The estimate of the nominal uncertainty that impacts the real wage with an immediate renewal is $\sigma^2\sigma_x^2$, which is σ^2 times the variance of the shock to the value of money. The estimate of the real uncertainty that impacts the real wage with a delayed renewal is σ_y^2, which is the variance of the real shock. The variable ELEC is a dummy variables for elections and UNEMP is the unemployment rate. 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:

$$
\begin{aligned}
\text{DELAY} \;=\; & \alpha + \hat{\beta}_1 z + \hat{\beta}_2 D(z - z^\pi) + \hat{\beta}_3 \sigma^2\sigma_x^2 + \hat{\beta}_4 \sigma_y^2 + \hat{\beta}_5 \text{ELEC} \\
& + \hat{\beta}_6 \text{UNEMP} + \hat{\gamma}\,\text{BRNCH}.
\end{aligned}
$$

Our model yields the following predictions:

- $\hat{\beta}_1 > 0$. This is the positive effect of z on the delay if $z < z^\pi$.

- $\hat{\beta}_1 + \hat{\beta}_2 < 0$. This is the negative effect of z on the delay if $z > z^\pi$.

- $\hat{\beta}_3 > 0$. 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.

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² $\hat{\beta}_5 < 0$. Elections is another indicator of real uncertainty as the economic policy may change if/when a new government is elected. We therefore expect a negative coefficient.

² $\hat{\beta}_6 < 0$. A high unemployment rate is associated with more real uncertainty and therefore expected to have a negative effect on the delay.

The model provides no a priori predictions about $\hat{\beta}_0$ 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.[32] The model predicts:

² 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.

² The coefficient of the interaction between the public sector and elections should be negative: elections shorten the delay in the public sector more than in the private sector.

² 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.

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

7.1 The Delay Regressions

In Table 6, regressions (1) and (2) present the random-e®ects coe±cient estimates of the length-of-delay regressions, using the ¯rm for the cross-section index[33] and assuming that ½ = 0. Regression (1) is based on the 1,731 contracts with ¯rms in the private sector only. Regression (2) is based on the pooled sample of 2,103 contracts with ¯rms in the private and public sectors, and it includes interaction terms between the public sector and each explanatory variable. As the coe±cients for the private sector are very similar in regressions (1) and (2), and there is more information in regression (2), we discuss only the results from this regression in detail.[34]

The Hausman test shows that the key assumption in the random-e®ects model, namely that the u_j disturbance term is uncorrelated with the explanatory variables, is satis¯ed.[35] The Breusch-Pagan test indicates that the random-e®ects model is preferable to the ordinary-least-squares model. The very large values of $Â^2$ is a sign that a basic assumption of the ordinary-least-squares model, namely that $var(u_j) = 0$, is violated. Accordingly, taken together, the Breusch-Pagan and the Hausman tests show that the random-e®ects model is the correct speci¯cation of the delay function.

The coe±cient estimates provide strong support for the theoretical model. They all have the predicted signs and are highly signi¯cant. This is true for both the private and the public sectors, the only di®erence being in the magnitude of the e®ect of some of the explanatory variables. The uninteracted variables refer to the private sector, and we start by discussing their estimated e®ects.[36]

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[33] We have also tried to specify the economic branch for the cross-section index. This leads to basically similar results and generally higher signi¯cance levels for the coe±cients of the key variables.

[34] 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 a week of delay (88%). The results are again similar. The same holds true for the likelihood-of-delay estimation.

[35] (Prob> $Â^2$) = 0:1365 in regression (1), and (Prob> $Â^2$) = 0:0973 in regression (2).

[36] For the numerical illustrations we use the means of the macro variables from Table 5 which are based

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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 $z = z^{\alpha}$. The regression results support both the increasing and the decreasing portion of this relationship. This is because $\hat{\ }_1 + \hat{\ }_2 = {}_{\textrm{¡}} 26{:}274$, and an $\hat{A}^2$-test yields $\hat{A}^2(1) = 30{:}35$, which is significant at less than the 0.001 level. Accordingly, for $z < z^{\alpha}$ an increase in z by 0.01 causes the delay to be lengthened by 5.998 days, while for $z > z^{\alpha}$ 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 $^{\circ}$ (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 $z < z^{\alpha}$ $(z > z^{\alpha})$ 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 $1_{\textrm{¡}} {}^{\circ}$. In the face of a positive one-standard-deviation nominal shock, for $z < z^{\alpha}$ wage indexation decreases the delay by $599{:}8 \times 0{:}0238 \times 0{:}5965 = 8{:}5152$ days through its effect on z, while for $z > z^{\alpha}$ wage indexation increases the delay by $2627{:}4 \times 0{:}0238 \times 0{:}5965 = 37{:}3$ days through its effect on z.

Since $^{\circ}$ 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 $z < z^{\alpha}$ the trend in the value of money on average increases the delay by $599{:}8 \times 0{:}0238 \times (0{:}5888 {}_{\textrm{¡}} 0{:}4035) = 2{:}6452$ days through its effect on z, while for $z > z^{\alpha}$ the trend in the value of money on average decreases the delay by $2627{:}4 \times 0{:}0238 \times (0{:}5888 {}_{\textrm{¡}} 0{:}4035) = 11{:}587$ days

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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.

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 $194840 \times 0.4035^2 \times 6.8E{-}4 = 21.571$ days. Without wage indexation, the delay would have increased by $194840 \times 6.8E{-}4 = 132.49$ 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 $61220 \times 2.3E{-}4 = 14.081$ days.

If the previous contract expires during the year before an election, the delay is shorter by 17.619 days. An increase in the unemployment rate of one percentage point reduces the delay by 30 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.[37] 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

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[37] 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).

between real uncertainty and delay, while still negative and signi¯cant, is weaker than in the private sector. A doubling of the variance of the real shock would decrease the delay by only $(61220 ¡ 43410) \times 2:3E{4 = 4:3263$ days in the public sector versus 14.081 days in the private sector. Upcoming elections reduces the delay in the public sector by an additional 68.930 days, and the e®ect of elections on delay is more than three times as strong as in the private sector. The signi¯cant positive coe±cient of 18.309 for the interaction term Public sector×Unemployment cancels the signi¯cant negative coe±cient in the private sector, so that unemployment appears to have no signi¯cant e®ect on the delay in the public sector.[38]

In regressions (1) and (2) the delay is the dependent variable, and the regressions therefore require information about the length of the delay.[39] 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 regression (3) the results of a probit random-e®ects model are presented. The dependent variable is a dummy variable which 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 regression (2), and therefore include public-sector interactions. The hypotheses are identical to those speci¯ed in section 7.1, except that the dependent variable is now dichotomous. The value of $\hat{A}^2$ for the likelihood-ratio test rejects the hypothesis that the panel estimator is the same as the pooled standard probit estimator.

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[38] A $\hat{A}^2$-test shows that the sum of the two coe±cients is not signi¯cantly di®erent from zero.

[39] We have also tried to use relative delay (´ delay/duration) as the dependent variable, rather than delay itself. The sign and signi¯cance of the coe±cients are comparable to what is obtained with delay as the dependent variable, except that the coe±cient of z is insigni¯cant. The implied e®ects 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 °, 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 to 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 to a decrease in the delay by $6:3$ days.

The estimates of the non-interacted variables are generally similar to those in regression (2) in terms of sign and signi¯cance. The main di®erence is that the coe±cient of z is insigni¯cant (z = 0:19). Additionally, the likelihood of delay is highest in manufacturing, while there is no signi¯cant di®erence between the other branches. All interaction terms are insigni¯cant, implying that the likelihood of delay is similar in the two sectors.

8 Conclusion

In many countries, the typical labor contract has a ¯xed duration. Often, however, the ¯xed 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 a (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 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 contract signed from 1978 to 1995. The renewal is delayed for 86% of these contracts, with the average delay being 213 days. The empirical ¯ndings strongly support the theory. The coe±cient estimates all have the predicted signs and are signi¯cant. 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 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 increases the real uncertainty, it also reduces 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 ¯rms in both the private and public sectors is used to gauge the di®erent impact of the explanatory variables in the two sectors. The coe±cient estimates of the interaction terms with the public sector provides 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 private sector.

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

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TABLE 1

The Benefit from Delay and the Critical Values of z

	$B(z)$	s	S
z	$\gtrless 0$ as $z \lessgtr z^{*}$		
[1]	> 0	< 0	> 0
Nominal Uncertainty	> 0	< 0	> 0
μ	< 0	> 0	< 0
ν	$\gtrless 0$ as $z \gtrless 1$	$\gtrless 0$ as $s \lessgtr 1$	$\gtrless 0$ as $S \gtrless 1$
Real Uncertainty	< 0	> 0	< 0
r	$\lessgtr 0$ as $z \gtrless 1$	$\lessgtr 0$ as $s \lessgtr 1$	$\lessgtr 0$ as $S \gtrless 1$

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 Thousand Employees
1978	48.1	1.93	1.1	3.4	1054.46
1979	111.4	2.05	-11.4	2.9	488.50
1980	132.9	0.67	18.5	4.6	195.54
1981	101.5	2.88	34.3	5.1	684.85
1982	131.5	-0.46	3.8	5.0	1552.30
1983	190.7	0.68	-3.1	4.5	818.02
1984	444.9	0.15	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	-0.56	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	0.68	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: *Bank of Israel Annual Report*, various issues.
Bank of Israel, Research Department: *Recent Economic Development*s, various issues.
Central Bureau of Statistics: *Monthly Bulletin of Statistics* and *Statistical Abstract of Israel*, various issues.

TABLE 3

Distribution of Contracts by Sector and by Economic Branch
Israeli Labor Market, 1978-1995 (2,103 Contracts)

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: *Labour, Social Affairs and National Insurance* (Monthly Bulletin), various issues.

TABLE 4

Delays in Contract Renewals and Contract Durations by Sector
Israeli Labor Market, 1978-1995

	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 (207.20)	193.79 (188.78)	304.02 (258.92)
Relative delay (delay/duration)	0.3337	0.3151	0.4200
Average duration (days)	649.49 (264.57)	628.57 (232.09)	746.83 (365.86)
Duration of one year (%)	20.73	21.43	17.47
Duration of two years (%)	51.55	51.70	50.81
Duration of more than two years (%)	10.27	8.43	18.82
Duration of more than three 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: *Labour, Social Affairs and National Insurance* (Monthly Bulletin), various issues.

Note: Figures in parentheses are standard deviations.

TABLE 5

Characteristics of Regression Variables, Israeli Economy, 1978-1995 (2,103 Contracts)

Variable	Mean	Standard Deviation	Minimum Value	Maximum Value
Delay (days)	213.29	207.20	0	1529
Value of money				
- Trend in the value of money	-0.0385	0.0384	-0.1533	-0.0041
- Nominal shock	-0.0159	0.0238	-0.1185	0.1385
- Expected variance of nominal shock	6.8E-4	1.2E-3	2.9E-6	7.5E-3
GDP (per-capita)				
All economic branches				
- Trend in GDP	0.0032	0.0140	-0.0224	0.0527
- Real shock	0.0038	0.0263	-0.0306	0.0780
- Expected variance of real shock	2.3E-4	5.5E-4	5.9E-9	2.6E-3
Private services (574 contracts)				
- Trend in GDP	0.0042	0.0191	-0.0177	0.0527
- Real shock	0.0055	0.0297	-0.0306	0.0780
- Expected variance of real shock	2.4E-4	6.1E-4	1.5E-6	2.6E-3
Public services (193 contracts)				
- Trend in GDP	0.0029	0.0083	-0.0096	0.0238
- Real shock	0.0012	0.0140	-0.0191	0.0339
- Expected variance of real shock	6.5E-5	1.4E-4	9.5E-8	7.0E-4
Manufacturing (1,198 contracts)				
- Trend in GDP	0.0024	0.0120	-0.0224	0.0252
- Real shock	0.0048	0.0278	-0.0230	0.0751
- Expected variance of real shock	2.7E-4	5.6E-4	2.9E-7	2.4E-3
Commerce (92 contracts)				
- Trend in GDP	0.0011	0.0100	-0.0095	0.0274
- Real shock	0.0024	0.0206	-0.0265	0.0510
- Expected variance of real shock	1.4E-4	1.7E-4	2.2E-6	5.8E-4
Banking Institutions (46 contracts)				
- Trend in GDP	0.0021	0.0124	-0.0135	0.0331
- Real shock	0.0034	0.0216	-0.0149	0.0631
- Expected variance of real shock	1.2E-4	2.4E-4	5.9E-9	8.4E-4
Degree of indexation	0.5888	0.1756	0.0336	1.0000
γ	0.4035	0.1775	0	0.9663
z	0.9916	0.0254	0.9031	1.0696
Real interest rate	0.0153	0.0129	-0.0101	0.0537
Election years	0.2240	-	-	-
Unemployment rate	0.0705	0.0232	0.0288	0.1118

Sources: Authors' calculations based on:

     Bank of Israel: *Bank of Israel Annual Report*, various issues.

     Bank of Israel, Research Department: *Recent Economic Development*s, various issues.

     Central Bureau of Statistics: *Monthly Bulletin of Statistics* and *Statistical Abstract of Israel*, various issues.

     Ministry of Labour and Social Affairs: *Labour, Social Affairs and National Insurance* (Monthly Bulletin), various issues.

Notes:

1. Trend, shock, and real interest rate are per month.
2. The value of z (= the ratio of the real wage with delay to the real wage without delay) is measured in percent.

TABLE 6

Random-Effects Delay Regressions
Israeli Labor Market, 1978-1995

Independent variables	(1) Length-of-Delay in Private Sector Coefficient (z–value)	(2) Length-of-Delay in Private and Public Sector, With Public-Sector Interactions Coefficient (z-value)	(3) Likelihood-of-Delay in Private and Public Sector, With Public-Sector Interactions Coefficient (z-value)
z	5.471 (2.66)	5.998 (2.60)	-0.010 (0.19)
$D(z-z^*)$	-30.910 (5.79)	-32.272 (5.35)	-0.362 (2.85)
$\gamma^2 *$ exp. variance of nominal shock	19.031 (6.07)	19.484 (5.51)	0.263 (2.88)
Exp. variance of real shock	-6.029 (7.85)	-6.122 (7.11)	-0.065 (4.01)
Elections	-28.685 (2.87)	-30.003 (2.67)	-0.287 (1.22)
Unemployment rate (%)	-16.827 (7.73)	-17.619 (7.28)	-0.230 (4.39)
Economic Branch			
Public Services	-85.512 (1.92)	-82.821 (1.69)	1.080 (0.81)
Manufacturing	-21.460 (1.95)	-19.477 (1.45)	0.777 (2.97)
Commerce	-15.504 (0.62)	-8.514 (0.32)	0.641 (1.04)
Banking Institutions	60.518 (1.38)	82.421 (1.91)	1.322 (1.29)
Public Sector	-	-12.301 (0.36)	-1.625 (0.09)
Public Sector $* z$	-	-2.793 (0.49)	0.026 (0.14)
Public Sector $* D(z-z^*)$	-	-32.311 (1.97)	-0.489 (1.31)
Public Sector $* \gamma^2 *$ exp. variance of nominal shock	-	14.540 (1.49)	0.092 (0.24)
Public Sector $*$ Exp. variance of real shock	-	4.341 (1.94)	-0.023 (0.47)
Public Sector $*$ Elections	-	-68.930 (2.33)	0.752 (0.68)
Public Sector $*$ Unemployment rate (%)	-	18.309 (3.59)	0.104 (0.61)
Constant	-193.222 (0.98)	-241.018 (1.09)	5.254 (0.98)
R^2	0.0958	0.1312	
χ^2 for Hausman test	14.88	28.54	
χ^2 for Breusch–Pagan test	268.97	203.35	
χ^2 for likelihood-ratio test			76.47
Sample size	1,731	2,103	2,103

Notes:
1. STATA 7.0 is used for estimation.
2. 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.
3. The firm is used as the cross-section identifier.
4. The reference group for Economic Branch is: Private Services.
5. 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 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.
6. 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.
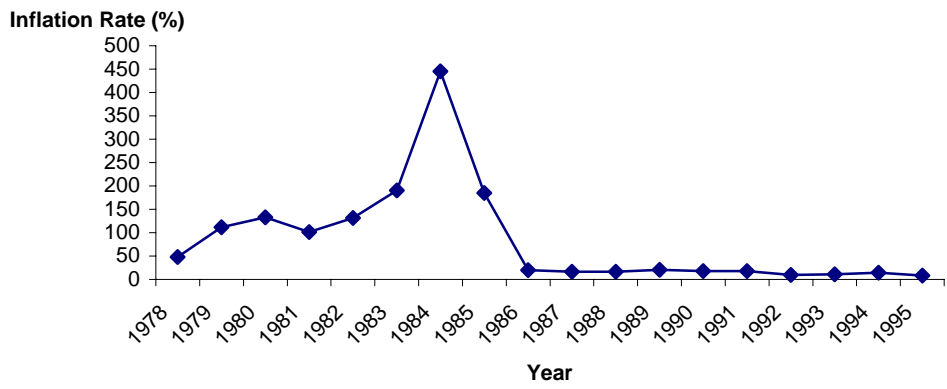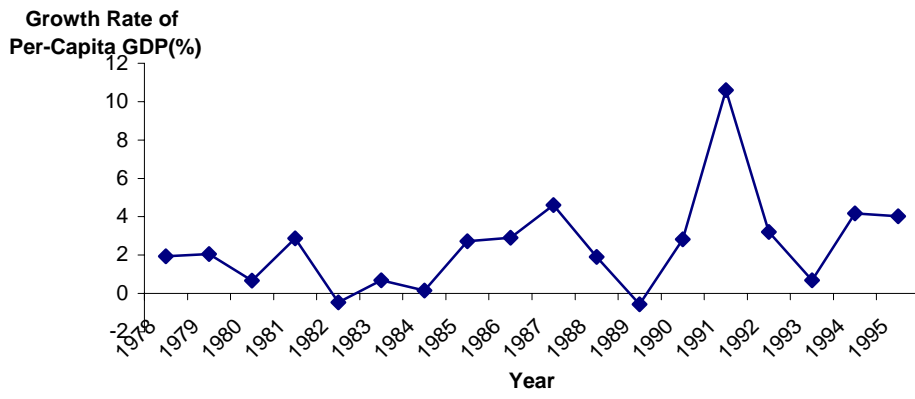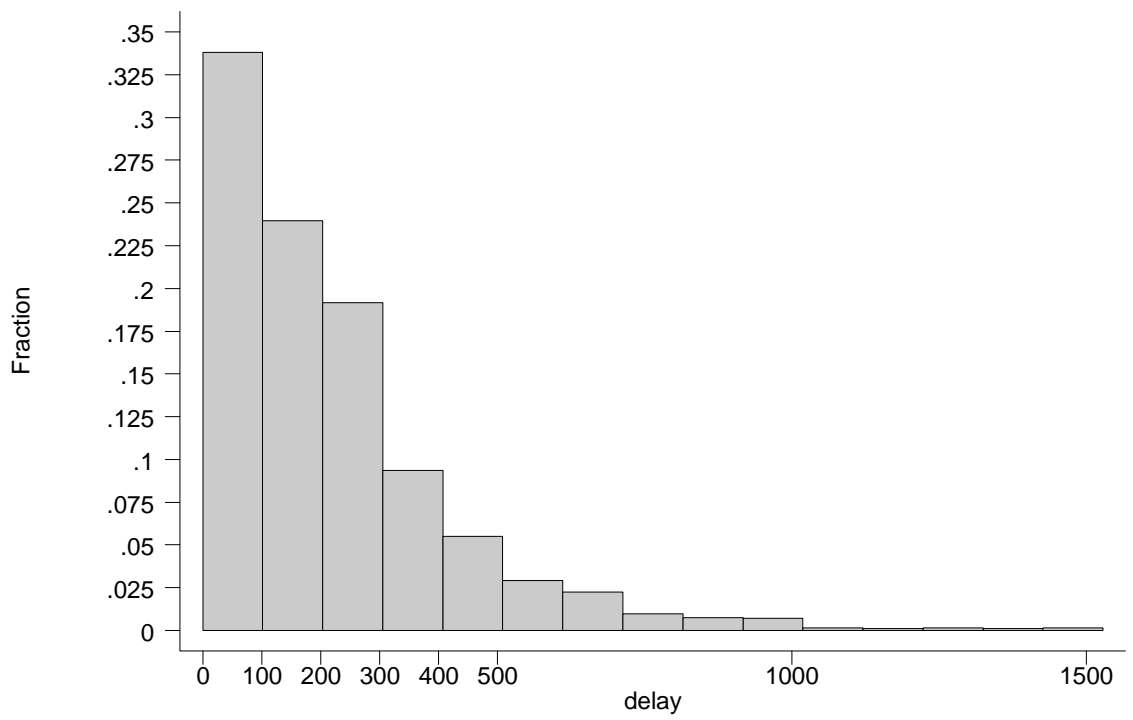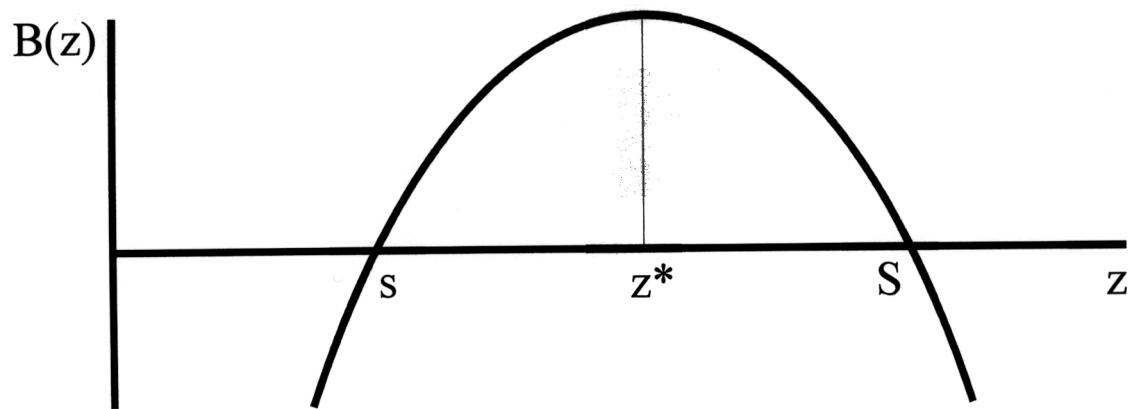
Chart 1

Inflation Rate (%)

Year

Chart 2

**Growth Rate of
Per-Capita GDP(%)**

Year

FIGURE

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