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PRICE DURATION WITH TWO-SIDED PRICING RULES

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Abstract

Substantial evidence exists that some firms are lowering prices while other firms are raising prices, even in highly inflationary times. A stochastic process with disparate shocks to individual firms' cost and demand accommodates this phenomenon within the menu cost framework. A Markov chain representation of this process can be related directly to the classical ruin problem in probability theory. If firms follow the optimal price-resetting rules, there is a mapping from data to parameters of the Markov process. Data from a variety of industries in the USA and New Zealand reveal that lower inflation increases the time between price changes whether the changes are up or down and that price duration was less sensitive to a decline in inflation rates in New Zealand than in the USA.

Key words: Menu-costs, *S,s* pricing rules, price duration, inflation.

Journal of economic literature classification: E31

I Introduction *

Prices are said to be "sticky" if they do not respond to new information with sufficient speed to keep supply equal to demand in all markets at all times. When prices are fixed, the most extreme form of price stickiness, any nominal changes, such as an increase in the money supply, will be met by quantity adjustments and possible spillover effects across markets. In fact, with any degree of price stickiness, whether because of delayed price changes or because of partial price adjustments, exogenous nominal changes can have repercussions on quantities. Thus, economists have embarked on numerous theoretical studies to assess why prices should be expected to be less than perfectly flexible and on empirical studies of the actual nature of price stickiness or, much less frequently, to test implications of the theoretical models[1] .

The most well developed theoretical analyses of price stickiness have followed the "menu cost" approach, reflected in the collection of papers in Sheshinski and Weiss (1993). If it is costly to make a price change, firms will delay changes until the private benefits outweigh the private costs. If there is general inflation, a firm's real price will fall in the absence of any changes in its nominal price. Following the inventory literature, theorists have found that it is optimal for a firm to let its real price fall to some lower bound s and with an increase in its nominal price to raise its real price to an upper bound S. These rules have been referred to as S,s pricing rules.

Weiss (1993), in a survey of empirical findings related to these models, comments: "At this early stage of the research, the connections between the data and theory are rather loose." Our purpose here is to help tighten some of those connections.

One ubiquitous fact that has been frequently ignored is evidence that firms lower as well as raise prices, even in periods of high inflation. If these menu-cost models are to predict such behaviour, then they need to develop more fully pricing rules in which price decreases as well as increases can be triggered by

* We thank in particular Lewis Evans and Colin Jeffcoat for their constructive comments and participants at seminars at Victoria University of Wellington and the University of Otago at which earlier versions of this paper were presented.

[1] Fisher and Konieczny (1994) discuss some of these tests.

specific events. Section II looks at one suggested model. Section III derives statistics that would be generated by the theoretical framework, and Section IV explores the link to available empirical statistics.

Sections V is devoted to questions about whether the available evidence is consistent with what would be expected. We find, with data from both the USA and New Zealand, overwhelming support for the hypothesis that higher inflation generates shorter times between price changes whether up or down. We also find evidence that duration increased more as a result of disinflation in the USA than it did in New Zealand. Section VI contains concluding comments.

II Two-Sided *S,s* Pricing Rules

A seminal paper by Sheshinski and Weiss (1977) provides a rigorous analysis of a firm's *S,s* pricing rule when there is a steady inflation in the general price level, the firm's profits depend on its price relative to the general price level, and the firm incurs lump-sum costs to change its nominal price. If firm *i* sets a nominal price P_{it} and the general price level is $\overline{P_t}$, then $Z_{it} = P_{it} / \overline{P_t}$ is the firm's real price. Letting lower case letters denote logs:

$$z_{it} = p_{it} - \overline{p_t}$$

If $\overline{p_t}$ grows steadily as the result of general inflation and the firm's nominal price is unchanged, then z_{it} falls steadily. When z_{it} reaches a lower bound *s*, it is optimal for the firm to raise its nominal price so that z_{it} jumps to *S*. With unchanged parameters, the process repeats itself with regularity. Sheshinski and Weiss prove that a higher general inflation rate will lower *s*, increase *S*, and (subject to a monotonicity condition) increase nominal price duration, which we define as the time between price changes, whether up or down.

In a subsequent paper, Sheshinski and Weiss (1983) carry out the analysis with stochastic inflation. As in their earlier paper, firms never lower prices. Such *S,s* models in which prices change in only one direction involve what are called one-sided *S,s* rules.

Real world evidence shows that, even in inflationary times, some prices are lowered; see for example Lach and Tsiddon (1992), Carlson (1992), and Tommasi (1993). To accommodate this evidence, two-sided pricing rules are clearly needed when there are both price decreases and increases. One example of a two-sided model can be found in Caplin and Leahy (1991). In their model, aggregate demand can fall as well as rise so that at times some firms will lower prices and at other times some firms will raise prices, but it is never the case in their model that, at the same time, some firms are lowering and other firms are raising prices.

To allow for the sort of disparate pricing behaviour that is apparent in the real world, the models need to incorporate idiosyncratic shocks to individual firms. One way to do this is suggested by Caplin (1993) in a survey of aggregation issues in the face of individual firms following *S,s* pricing rules. He introduces the notion of a firm's "optimal" price. We can interpret this as the price that would maximise a firm's profits if there were no costs of price adjustments. Denote the log of this optimal price by p_{it}^* and assume, as Caplin does, that the change in a firm's optimal price depends on an aggregate component, Δm_t, that is common to all firms and an idiosyncratic component, ε_{it}, with an expected value of zero. Thus:

(1) $$\Delta p_{it}^* = \Delta m_t + \varepsilon_{it}$$

where all variables are measured in logs, so that a first difference can be interpreted as a percentage change.

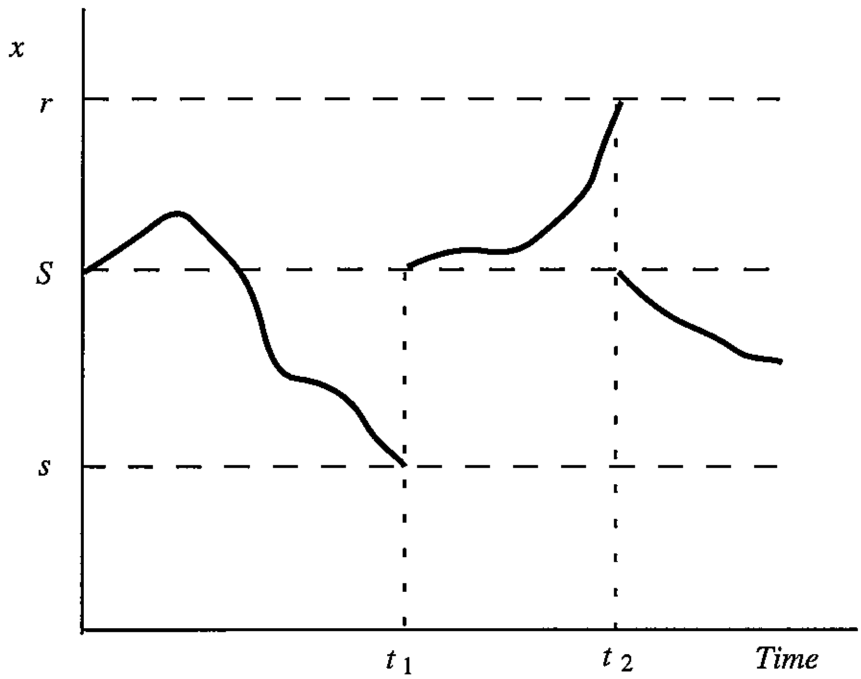
A <u>deviation</u> in a firm's nominal price from its optimal price is denoted by:

(2) $$x_{it} \equiv p_{it} - p_{it}^*$$

Caplin assumes that each firm follows a two-sided pricing rule. Whenever, as a result of cumulative shocks to demand and costs, the deviation x_i reaches a lower or an upper bound, the firm resets its nominal price so that the deviation from the optimal price stays within the upper and lower bounds.

Figure 1 illustrates a possible meandering of these deviations. A similar figure can be found in Blanchard and Fischer (1989, p. 404). Suppose a firm's nominal price is initially set so that $x = S$. If the firm's optimal price rises and its nominal price stays fixed, then x will fall. If x reaches a lower bound s, nominal price is raised so that x jumps back to S as shown in Figure 1 at time t_1. If the firm then gets a series of negative shocks to p^* that raise x to an upper bound r, as shown in Figure 1 at time t_2, the firm lowers its nominal price. Unlike the one-sided *S,s* pricing rule with steady inflation, the time between a firm's price changes can vary and the direction of the nominal price change can be either up or down.

Figure 1: Sample Price Deviations with Two Sided Pricing Rule



An empirical examination of the distribution of deviations of prices from optimal, requires observations of optimal prices. Such variables might be constructed for individual firms from quantities sold, from estimates of the responsiveness of quantity demanded to variations in price and from cost data, but firms' optimal prices are not as a rule observable. Any implications of Caplin's specification about the distribution of x_i will very likely go untested[2] .

On the assumption that the idiosyncratic shocks average out to zero, the average optimal price across all firms grows at the rate Δm_t , which can be interpreted as approximately the general rate of inflation π.

[2] When the x_i are bounded, prices of the same goods in different stores would not as a rule conform to the implications of equations (1) and (2) that nominal prices can over time wander apart by an unlimited amount.

III Expected Price Duration and Probability of Price Increases

In order to link a model such as Caplin's to available statistics about price duration and proportions of firms which raise and lower prices during an interval of time, we utilize a framework analysed by Dixit (1991). First note that in continuous time a state variable x that follows a Brownian motion would be represented by

$$dx = \pi dt + \sigma dz$$

where dz is the increment in a standard Wiener process. This Brownian motion has the property that the change in x over a unit interval of time is a normally distributed variable with mean π and variance σ^2

Dixit recasts Brownian motion in a discrete Markov chain framework. Think of an interval of time being divided into small sub-intervals and let x, the deviation of a firm's price from its optimal price, take on a set of equally spaced discrete values between a lower bound s an upper bound r. From one sub-interval to the next, x can move down a step with probability p and move up a step with probability $1-p$. If x reaches a lower bound s, the firm raises its nomimal price so that x jumps to S. If x reaches an upper bound r, the firm lowers its nominal price so that x jumps down to R. Dixit points out that if there are only lump-sum costs of an impulse control, in our case of changing nominal price, then the optimal reset points S and R coincide. In what follows, we assume that this is the case and S will represent the reset point for x.

Dixit's setting is exactly the same as what Feller (1950) calls the classical ruin problem, but the variables are given different names. In the classical ruin problem, $r-S$ is the stake of one gambler, $S-s$ is the stake of another gambler, p is the probability that the first gambler wins a bet of one unit, and $1-p$ is the probability that the second gambler wins the one unit. The game ends when one of the gamblers has lost an initial stake (is ruined). Feller derives formulas for the probability that one or the other gambler loses the game and for the expected time until one of them loses.

In the price setting framework, the probability that the firm's price deviation x hits the lower bound s before it hits the upper bound r is analogous to the first gambler winning Feller's game, and the probability that x reaches r before it reaches s is analogous to the second gambler winning the game. Resetting x to S is analogous to restarting the game. If p >.5, then there is a drift toward the lower bound at which nominal prices are raised.

Let T be the expected time between resettings of nominal price. In terms of our price-setting notation, Feller's duration formula (p 286) can be written as follows (letting $\phi = p/(1-p)$ and assuming $p \neq .5$):

(3) $$T = \frac{S-s}{2p-1} - \frac{r-s}{2p-1}\frac{1-\phi^{S-s}}{1-\phi^{r-s}}$$

Denote by $P(up)$ the probability that the price deviation hits s before r and nominal price is raised. Feller's formula (p283) for the probability that the first gambler wins, and in our case that price is raised, becomes:

(4) $$P(up) = \frac{\phi^{r-s} - \phi^{S-s}}{\phi^{r-s} - 1}$$

Suppose a set of firms satisfy the condition necessary to be using a state-contingent two-sided impulse-control rule of the sort postulated by Caplin and Dixit. Suppose also that we observe how frequently these firms change their prices and what proportions of the changes are increases, i.e., we have data for T and $P(up)$. Equations (3) and (4) are then two equations in three unknowns, p, $r-s$ and $S-s$. Is it possible to find a third equation so that these variables can be inferred from values of T and $P(up)$?

For purposes of illustration, we assume here that, contingent on values for $r-s$ and p, the optimal S will have been chosen so that T is as large as possible[3]. This can be found by setting the derivative of the right side of equation (4) with respect to S equal to zero. The result is

[3] A fuller analysis will be presented in another paper. The typical assumption in the literature is that firms choose a strategy that minimizes the expected costs of changing price plus the expected forgone profits while price is not changed. The least-cost combination for the location of the $r-s$ and $S-s$ intervals depends on the relative costs of changing and not changing prices and on the parameter p. Blanchard and Fischer (1989, pp 402-405) work out the cost-minimizing solution when p = .5. In that case, as in Barro (1972), the reset price at the mid point between the upper and lower bounds also maximizes the expected price duration given the upper and lower bounds. With p not equal to .5, maximizing T introduces a slight bias that will be documented in our other paper.

(5) $$S-s=\ln\left\{\frac{\phi^{r-s}-1}{(r-s)\ln(\phi)}\right\}/\ln(\phi)\equiv f(p,r-s)$$

The second derivative of T with respect to S is negative, and as before $\phi=p/(1-p)\equiv\phi(p)$. If $S-s$ from (5) is substituted into equations (3) and (4), then T and $P(up)$ become functions of p and $r-s$.

(6) $$T=\frac{f(p,r-s)}{2p-1}-\frac{r-s}{2p-1}\frac{1-\phi(p)^{f(p,r-s)}}{1-\phi(p)^{r-s}}$$

(7) $$P(up)=\frac{\phi(p)^{r-s}-\phi(p)^{f(p,r-s)}}{\phi(p)^{r-s}-1}$$

To recapitulate: The inflation drift, the variability of the exogenous shocks, and the cost of price changes are assumed to be exogenous to the firm. The firm then chooses the r, s, and S values. Once those values have been chosen, the expected time between price changes will be approximated by equation (6) and the probability that a price change will be an increase is given by equation (7).

IV Estimates of Average Price Increases and Decreases from Qualitative Data

Assuming that firms reset prices optimally, we can use statistics for T and $P(up)$ to reverse the foregoing argument and find values for p, $r-s$, and $S-s$ that would give rise to the available statistics. In Section V we shall present a preliminary analysis of quarterly survey data from both the USA and New Zealand in the light of these theoretical relationships.

To obtain statistics for price duration from qualitative data, let U = percent of firms reporting prices higher in the last three months and D = percent of firms reporting prices lower in the last three months. An average time T in months between price changes is then given by the following:

(8) $$T=3/(U+D)$$

For example, if 75 percent of the firms do not make a price change each quarter and 25 percent do, then a representative firm can expect to go 3/.25 = 12 months before changing price. There are potential biases in this measure. On the one hand, if firms tend to change prices on only some items during the quarter but their average prices are down or up, then $(U+D)$ overestimates the extent of price changing activity and T has a downward bias. On the other hand, if firms change prices more frequently than once a quarter, then T overestimates the average duration between price changes. We suspect that this latter bias becomes increasingly important at higher levels of overall inflation.

When there is a price change, the estimated probability of raising price is given by:

(9) $$P(up) = U / (U + D)$$

For example, if 20 percent of the firms raise prices and 5 percent lower prices each quarter, $P(up)$ would be estimated to be .20/.25 = .80, or an 80 percent chance of reaching s before r and raising rather than lowering prices.

One more piece of information, the actual rate of inflation among the firms being sampled, enables us to estimate how large price changes have been on average. As noted earlier, values of T and $P(up)$ can be used in equations (5), (6) and (7) to obtain values for p, $r-s$ and $S-s$, and hence of $r-S$. With U percent of the firms raising prices by an average of $S-s$ steps and D percent of the firms lowering prices by $r-S$ steps, the size of each step denoted by δ can be determined from the following equation:

(10) $$\delta\{(S-s)U-(r-S)D\} = \pi$$

where π is the general rate of inflation among the firms for which the survey data have been obtained. The estimated average price increase is given by $\delta(S-s)$ and the estimated average price decrease is $\delta(r-S)$.

To illustrate the potential information that can be gleaned from these formulas, we examine data reported by Lach and Tsiddon (1992), who make a careful study of monthly observations of prices of retail food and wine products in Israel. Table 1 reproduces a few of the statistics.

Table 1: Statistics taken from Lach and Tsiddon (1992)

	1978-79	1982
Inflation rate (percent/month) (π)	3.9	7.3
Reported price duration in months	2.21	1.48
Reported average price increase (in percent)	12.3	12.9
Reported *S,s* bands (in percent)	9.1	11.5
Reported percent up each month (U)	34	57
Reported percent down each month (D)	5	4
Inferred price duration (T)	2.56	1.64
Inferred $P(up)$	0.872	0.934
Inferred average price increase $\delta(S-s)$	12.1	13.0
Inferred average price decrease $\delta(r-S)$	4.8	2.8
Inferred total band $\delta(r-s)$	16.9	15.8

They report that the average price duration fell from 2.21 months in 1978-79 when inflation was 3.9 percent per month to 1.48 months in 1982 when inflation was 7.3 percent per month. They also report that when prices were increased, they were raised by an average of 12.3 percent in 1978-79 and by an average of 12.9 percent in 1982[4] . We interpret each of these latter two figures as an estimate of a $\delta(S-s)$ interval.

What is puzzling about their statistics is how to interpret an *S,s* band which is smaller than the average price increase. They report that these bands rose from 9.1 to 11.5 percent between 1978-79 and 1982. This might refer to a "representative" firm that makes a net change in price, with its decreases on some products offsetting its increase on other products. We prefer to think of the total band width as encompassing a range in which prices are neither increased nor decreased, i.e., the $r-s$ range in Figure 1.

[4] The Lach and Tsiddon estimates of the duration and size of price change can be compared to those from an earlier study by Liebermann and Zilberfarb (1985) who analysed price adjustments by Israeli manufacturers over an eleven month period from October 1978 through August 1979. They estimated that the average duration of prices was 1.8 months (55 days) when average monthly inflation was 5.4 percent, while the average rate of price increase was 11.7 percent.

Now consider what can be inferred if all we know is the fraction of the items on which prices are higher each month and the fraction which are lower. These are recorded in Table 1 as 34 percent up and 5 percent down in 1978-79 and 57 percent up and 4 percent down in 1982. Since these are monthly instead of quarterly data, we need to modify (8) to

$$T = 1/(U+D)$$

The duration implied by this formula, based only on qualitative monthly data, is 2.56 months [1/(.34+.05)=2.56] in 1978-79 and is 1.64 months [1/(.57+.04) = 1.64] in 1982. These are slightly higher than the durations calculated by Lach and Tsiddon but note that with either set of figures average duration in the latter period was about two-thirds of what it was in the earlier period[5].

The qualitative U and D statistics can also be used in equation (9) to calculate $P(up)$, the proportion of firms that raised price each month among those that changed price. This proportion rose from 87 to 93 percent between the two periods.

The numbers in the last three lines of Table 1 were calculated using equations (5), (6) and (7) with the T and $P(up)$ statistics to get values for p, $r-s$, $S-s$, and hence for $r-S$, and then using equation (10) with the π statistics to get δ. We find that the estimates for the average price increases are 12.1 percent in 1978-79 and 13.0 percent in 1982[6]. These are remarkably close to the figures of 12.3 and 12.9 reported by Lach and Tsiddon.

The inferred size of the average price decrease fell from 4.8 percent in 1978-79 to 2.8 percent in 1982. As a result, the total band width inferred from the qualitative data is smaller in the higher inflation period.

Before turning to survey data from the USA and New Zealand, we want to state hypotheses about the effects of higher inflation on the size of price changes and on price duration if price changes are generated by a process of the sort

[5] With estimates from the qualitative data 1.64/2.56 = 0.64 and with Lach and Tsiddon's figures 1.48/2.21 = 0.67.

[6] These estimates used monthly intervals. Similar results would emerge, however, if we used shorter intervals, such as weeks or even days. With shorter intervals the number of steps between r and s would be more numerous and p per step would be smaller but the transitions would be assumed to occur more frequently.

discussed by Dixit and utilized in deriving our formulas. *Ceteris paribus,* with two sided pricing rules, our hypotheses are that higher inflation will:

(a) Raise the $S-s$ band (larger price increases),
(b) Lower the $r-S$ band (smaller price decreases), and
(c) Decrease the duration between price changes.

Intuitively, we would generally expect higher inflation to raise the total $r-s$ band within which prices are neither raised nor lowered but, as seems to have occurred with the Lach and Tsiddon data, the decrease in $r-S$ might at times outweigh the increase in $S-s$[7].

V Price Duration Across Countries, Industries and Inflation Rates

Similar data in both the USA and New Zealand make possible inter-country comparisons of the effects of inflation on price duration and how those effects vary across types of industry.

In the USA since October 1973, the National Federation of Independent Business (NFIB) has been conducting quarterly surveys (in January, April, July and October) of its more than 500,000 members. Each quarter, questionnaires are mailed to a random sample of about 7,000 firms and responses are received from about one-third of those sampled[8]. The NFIB members are generally small businesses. However, "small businesses" in the USA consitutute about 40 percent of GNP and 50 percent of total employment, so their reported patterns of behaviour reflect an important part of overall economic activity. See Carlson and Dunkelberg (1989) for an analysis of some of the data in these surveys.

Since 1978:4 respondents have been asked the following question about their actual price changes: "How are your average selling prices now compared to three months ago?" The choices for answers are: "Lower now, No difference, Higher now, Don't know." We will use the data from the 59 surveys from 1978:4 to 1993:2 and ignore the don't know and no answer responses. The NFIB surveys also ask firms to indicate an industry category. Four categories - manufacturing, construction, retail, and services (other than professional or

[7] These hypothese will be examined sytematically in the paper mentioned in footnote 3.
[8] A copy of the questionnaire can be obtained from William C. Dunkelberg, Chief Economist for the NFIB, School of Business and Management, Temple University, Philadelphia, PA 19122.

financial) - will be used because these are closely matched by available industry groupings in the New Zealand data.

In New Zealand the Quarterly Survey of Business Opinion (QSBO), compiled by the New Zealand Institute of Economic Research, has responses to similar questions about changes in prices during the last three months by manufacturers, builders, merchants and non-financial services. See Bowie and Easton (1987) for details about these surveys. We will use data from the 75 quarters 1975:1 to 1993:3.

Table 2: Average Tradeoffs Between Inflation and Price Duration

USA	Higher inflation (1978:4 to 1981:4)	Lower inflation (1982:1 to 1993:2)	Difference
Average quarterly inflation rate (%)	2.86	0.94	-1.92
Average duration of prices (months)			
All sectors	4.66	8.51	+3.85
Manufacturing	5.57	10.02	+4.45
Construction	4.81	8.50	+3.69
Retail	4.03	7.40	+3.37
Services	6.26	10.94	+4.68
New Zealand	Higher inflation (1975:1 to 1982:3)	Lower inflation (1990:1 to 1993:3)	Difference
Average quarterly inflation rate (%)	3.59	0.55	-3.04
Average duration of prices (months)			
All sectors	4.67	7.39	+2.72
Manufacturers	4.70	8.48	+3.78
Builders	4.69	6.50	+1.81
Merchants	3.55	6.74	+3.19
Services	5.74	7.85	+2.11

Table 2 provides evidence of the effects of lower inflation on price duration, which was calculated using equation (8). In the USA, quarterly CPI inflation averaged 2.86 percent for the three years from the end of 1978 to the end of 1981. Beginning in 1982 and continuing to 1993, inflation has been about one-third as high, averaging 0.94 percent per quarter. For all four industries conbined, price duration in the USA rose from 4.66 months in 1979-81 to 8.51 months in 1982-93. Thus, each one percentage point drop in quarterly inflation apparently added almost two months to price duration on average. None of the four industry groupings separately exhibits patterns markedly different from the changes in price duration for all groups combined.

Inflation in New Zealand varied substantially over the years 1975 to 1993. It stayed higher for longer and in the last three years has come down lower than in the USA. In the period 1975:1 to 1982:3, CPI inflation averaged 3.59 percent per quarter. In the period 1990:1 to 1993:3, average inflation has been 0.55 percent per quarter. There has been a marked increase in price duration between the two periods for all four industry groupings. The changes, however, have been less than in the USA. Each one percentage point drop in quarterly CPI inflation in New Zealand has added about one month to price duration on average.

We also ran regressions to look more closely at the shorter run tradeoff between inflation and price duration. Before reporting some of those results, we would like to discuss plots of the relationship between inflation and price duration for manufacturers and for builders in USA (Figure 2) and in the New Zealand (Figure 3). Several points emerge from an examination of these plots:

1. When general inflation is very small, the duration of prices depends primarily on the variability of sector specific shocks. This can be seen most strongly in the New Zealand data as a wide scatter of points for duration when the inflation rate is close to zero. If inflation actually stayed at zero indefinitely, price duration could become very large for many firms.
2. The frequency of data collection imposes a lower bound on the estimated duration. With quarterly data, the lower bound is three months, which arises if every firm changes prices at least once every quarter. As noted earlier, there is a bias in the duration measure when firms change prices more frequently than every quarter at higher levels of inflation. Thus, the flattening out of the duration statistics at higher levels of inflation is to some extent due to the fact that the data are collected quarterly rather than more frequently.

3. Outliers may be attributable to special circumstances. New Zealand has had a variety of price controls, described by Boston (1984). Another outlier is in 1986:4 when New Zealand instituted a 12.5 percent Goods and Services Tax (GST). That caused a large one-time jump in the CPI.

Regressions, with both the USA and NZ data, were run for each of the four groups of firms (manufacturers, builders, merchants and non-financial services) separately and for a pooled data set across industries. A more detailed analysis, including dummy variables for special circumstances, is given in an Appendix. In Table 3, we report only the coefficients on quarterly inflation in the CPI. In all of the regressions, price duration is significantly lower when general inflation is higher. As suggested by changes in the averages shown in Table 2, the regression coefficients are almost twice as large for USA firms.

We do not know why changes in price duration appear to be more responsive to lower inflation in the USA than in New Zealand. Several possibilities suggest themselves[9] . Aside from the 1982-1984 price freeze, New Zealand has only recently moved into a low inflation environment. Perhaps, if New Zealand can maintain another decade of inflation of no more than two percent per year, we will see the average price duration among New Zealand firms rise to figures closer to and perhaps higher than those in the USA.

However, a greater proportion of New Zealand firms trade in foreign markets, especially since the post-1985 liberalisation programme, which is described in Bollard and Buckle (1987). Within the menu-cost framework price duration is influenced by the variability of the shocks hitting individual firms, the sensitivity of demand and hence profits to price deviations from optimal, the costs of price changes, and general inflation. It may be that firms in a small open economy, such as New Zealand, experience greater demand sensitivity to relative prices and relatively greater shocks to its demand and costs compared to firms in a somewhat more insulated economy, such as the USA. In addition, firms exposed to greater foreign competition are more sensitive to variations in foreign prices and hence less sensitive to domestic inflation than their counterparts in a more insulated economy.

[9] There is a difference in sampling procedures that is unlikely to make much difference. The NFIB surveys in the USA take a new sample every quarter. The QSBO surveys in New Zealand are a panel that stay in the sample from one quarter to the next. The sample has been revised just twice during the period covered by this study.

Figure 2: Inflation and the Duration of USA prices (1978:4 - 1993:2)

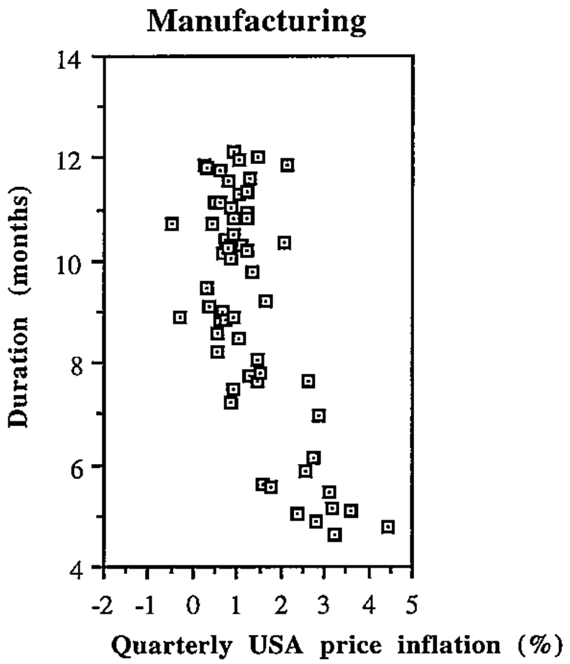


Figure 3: Inflation and the Duration of New Zealand prices (1975:1 - 1993:3)

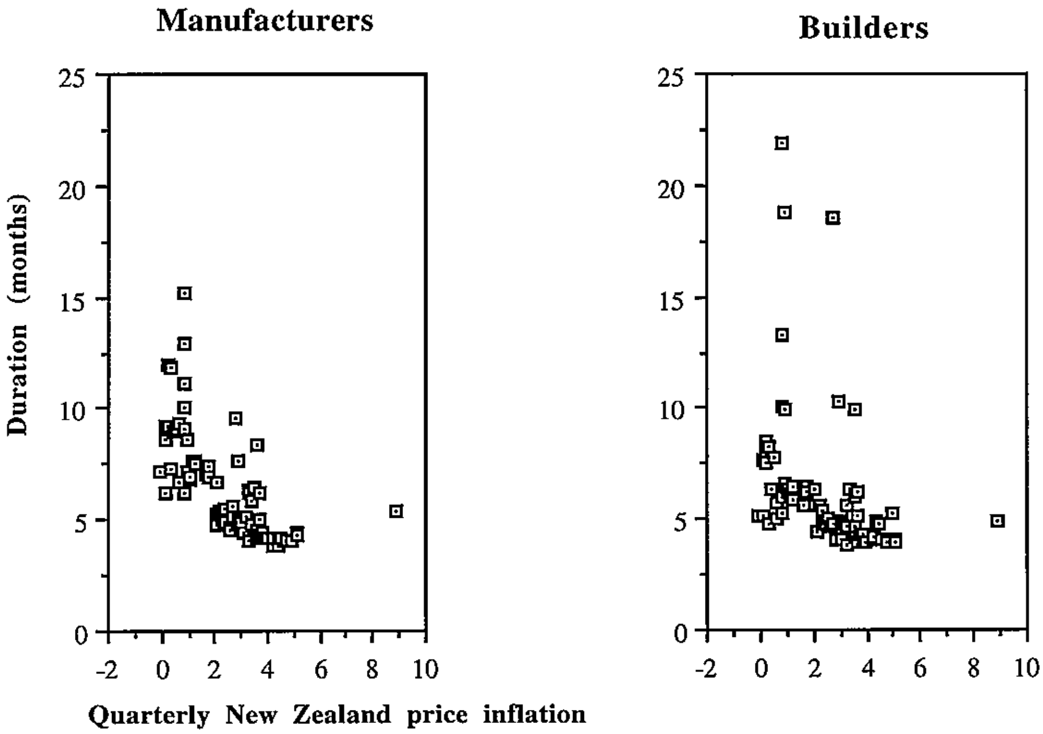


Table 3: Coefficients on Quarterly CPI Inflation in Regressions to Explain Price Duration

	USA 1978:4-1993:2	New Zealand 1975:1-1993:3
Manufacturers	-1.58	-1.08
	(7.14)	(11.1)
Builders	-1.40	-0.54
	(7.47)	(3.98)
Merchants	-1.22	-0.80
	(7.49)	(12.5)
Services	-1.86	-0.47
	(7.94)	(3.07)
Combined	-1.52	-0.78
	(15.03)	(10.7)

Note: t ratios are in parentheses

Despite these intercountry differences, the evidence from both the averages and regression coefficients is overwhelming supportive of the hypothesis that higher inflation results in shorter duration when prices may be lowered as well as raised

VI Concluding Remarks

We have shown with data from both the USA and New Zealand that price duration is markedly higher when general inflation is lower. This is true across a variety of industry groupings in both countries. What is unique about our estimates of duration is that we consider the time until a price is changed, whether down or up, and not just the average time between price increases in a period of general inflation.

In making comparisons across countries, we find that price duration among firms in New Zealand is less sensitive to changes in domestic inflation than among firms in the USA. There are several possible explanations that merit further study in the future. One involves the credibility of sustained low inflation. New Zealand has had a longer history of high and variable inflation than the USA and has had fewer recent years of low inflation. New Zealand is also a smaller, more open, economy. As a result, firms in New Zealand may be more sensitive to changes in foreign prices and less sensitive to changes in domestic inflation in deciding when to change prices than their USA counterparts.

Formal models that incorporate price changes in either direction have not as yet progressed very far, despite the well-established fact that prices decrease even in highly inflationary times. We have considered one such model. When coupled with the classical ruin problem from probability theory, the model can be linked to readily observable data on the frequency of price changes and the proportions that are up or down. Numerous surveys around the world contain such information and a reference to many of these data sources is CIRET (1990). This paper provides an important step toward formal tests of the model. At this stage, however, we are not satisfied that the aggregate data we have are adequate to test the model, which is essentially microeconomic in nature. We are continuing to analyse this model in an attempt to develop stronger links between theory and evidence.

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Appendix: Regressions of Price Duration on Quarterly CPI Inflation

In the regressions reported in Table A1, the dependent variable is price duration calculated in accordance with equation (8) in the text. The key right side variable, denoted Qinf, is the percentage change in the CPI from the prior quarter.

In New Zealand, there were a variety of incomes policies that might affect the estimated tradeoff between price duration and inflation. To check on these potential effects, we constructed dummy variables to capture different episodes:

a. Stabilisation of Prices Regulations from July 1974 to April 1979. Since the size of price increases were limited under these regulations, firms may have increased the frequency of price increases and hence lowered the duration. (D74 = 1 until 1979:1 and 0 thereafter).

b. A price and rent freeze was imposed from 17 August 1976 and expired 31 December 1976. This would have induced relatively fewer price increases and increased the implied duration between price changes. (D76 = 1 in 1976:4, and 0 elsewhere).

c. A wage price freeze from June 1982 until March 1984. This would have also increased the measured duration between price changes. (D82 = 1 from 1982:3 to 1984:1, and 0 elsewhere).

The introduction of a 12.5 percent Goods and Services Tax (GST) in October 1986 meant that the CPI took an unusually large jump in that quarter. (GST = 1 in 1986:4, and 0 elsewhere).

If a dummy variable was insignificant in any regression, we dropped the variable and reestimated the regression. OLS results for New Zealand are shown in the top part of Table A1. The Qinf slope coefficients ranged from -1.08 for manufacturers to -0.47 for services. The 1974 price regulation had a significant effect only among merchants and had the expected effect of reducing price duration. The 1976 freeze had a significant positive effect in three out of four industries. The 1982 wage-price freeze showed up as significant in all four industries, having the most notable impact in increasing price duration for builders and services. The introduction of the GST dummy did not affect the services regression significantly but did show up in regressions for the other industries.

When the data were pooled and we tested for significant differences in intercepts and Qinf slope coefficients among the four industries, we did not find that the slope coefficients were significantly different. When the equation was reestimated constraining the slopes to be the same, we found that the intercept for merchants was significantly below the intercept for the other three industries. This is shown under Dmer in Table A1.

The USA did not have comparable on-and-off incomes policies during the period considered and did not institute anything like New Zealand's GST. Consequently, we ran OLS regressions of price duration only on quarterly CPI inflation, with results shown in the bottom part of Table A1.

Table A1: OLS Regression Estimates

New Zealand	**(1975:1 - 1993:3)**						
	Const	Qinf	D74	D76	D82	GST	Dmer
Manufacturers	8.68	-1.08		2.19	3.87	6.34	
	(30.5)	(11.1)		(1.82)	(8.03)	(4.70)	
Builders	6.62	- 0.54		5.26	8.86	3.07	
	(16.8)	(3.98)		(3.16)	(13.2)	(1.64)	
Merchants	6.74	- 0.80	- 0.46		0.89	3.93	
	(37.8)	(12.5)	(2.11)		(2.94)	(4.53)	
Services	7.45	- 0.47		4.93	6.93		
	(15.7)	(3.07)		(2.32)	(8.10)		
Pooled	7.97	- 0.78			5.05	4.21	-1.77
	(36.3)	(10.7)			(13.9)	(4.14)	(7.42)

United States	**(1978:4 - 1993:2)**				
	Const	Qinf	Dmfg	Dretl	Dserv
Manufacturers	11.14	-1.58			
	(30.5)	(7.14)			
Construction	9.94	-1.40			
	(31.0)	(7.48)			
Retail	8.29	-1.22			
	(33.6)	(8.18)			
Services	12.38	-1.86			
	(32.2)	(7.95)			
Pooled	9.70	-1.52	1.34	-1.03	2.22
	(40.7)	(15.03)	(4.82)	(3.70)	(7.97)

Note: t ratios are in parentheses

For each industry classification, the coefficients show that USA firms adjusted price duration more in response to changing inflation than did firms in New Zealand. We pooled the data and introduced separate dummy variables for the intercepts and slopes for manufacturing, retail, and services to see if they were significantly different than for construction. The intercept dummy variables were all significant. We did not find any significant differences in the slopes. The highest t-ratio was -1.62 for services. However, the slope coefficient for services was significantly different from the slope coefficient for retail.

When we constrained the slopes to be the same and reran the regressions, the intercepts were again significantly different. This is reported in Table A1 as a pooled regression, along with the coefficients on the intercept dummy variables (denoted Dmfg, Dretl, and Dserv for manufacturing, retail and services, respectively). Retail firms tended to have a significantly lower price duration than the firms in other industries, as was true for merchants in New Zealand. The coefficients also indicate that manufacturing and service firms tended to have a higher duration than construction firms in the USA.

There was some evidence of first-order autocorrelation in the residuals for some of the industry regressions in both countries. For services, the Durbin-Watson statistic (DW) was about 1 87 in both countries and it was 1.72 for New Zealand merchants. All of the other DW statistics were significantly different from 2.00 and so we reestimated those equations after making autoregressive transformations. The results are reported in Table A2.

The Durbin-Watson statistic of 2.67 for builders in New Zealand suggests negative autocorrelation in the residuals. These arise primarily because of large residuals of alternating signs over 6 quarters. Otherwise, the pattern of errors looks more like one of positive first order correlation. In all the industries for which the DW statistic indicated positive autocorrelation, reestimating with an autoregressive transformation substantially lowered the estimated responsiveness of duration to quarterly inflation.

We interpret these new coefficients, which still have significant t-ratios, as indicative of the short-run (quarter to quarter) responsiveness of price duration to changes in the rate of general inflation. The longer run responsiveness to different sustained levels of inflation are better reflected in a comparison of averages over different inflation episodes, as shown in Table 2 in the text.

Table A2: Effects of Autoregressive Transformations on the Inflation Coefficient

	ols	dw	auto	rho
New Zealand				
Manufacturers	-1.08	1.39	-0.49	0.688
	(11.1)		(3.48)	
Builders	-0.54	2.67	-0.59	-0.402
	(3.98)		(6.58)	
United States				
Manufacturing	-1.58	1.46	-1.18	0.237
	(7.14)		(4.90)	
Construction	-1.40	1.41	-1.07	0.258
	(7.48)		(5.27)	
Retail	-1.22	1.16	-0.72	0.386
	(8.18)		(4.73)	

Notes: t ratios are in parentheses; ols denotes ordinary least squares estimates of the Qinf coefficient from Table A1; dw denotes Durbin-Watson statistics in the OLS regressions; auto denotes new slope coefficient on Qinf after an autoregressive transformation; rho denotes autoregressive coefficient.

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