TIME RIGIDITIES IN THE ADJUSTMENT OF PRICES TO MONETARY SHOCKS: AN ANALYSIS OF MICRO DATA

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A store that increases the nominal price of a given good by more than other stores will tend to wait longer before it increases its nominal price again, but only by as much as 15% of the additional time interval predicted by models which assume fixed menu-type costs for changing nominal prices. This finding is based on large data sets of prices by products and stores during recent inflationary periods in Israel. It suggests that coordination-type costs are important relative to menu-type costs.

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1. INTRODUCTION

How do prices adjust to monetary shocks? This question lies at the heart of monetary economics. Different theories assume different adjustment mechanisms, and the policy implications tend to depend heavily on the particular assumptions made.

Here I focus on the time interval between two consecutive price increases referred to hereafter as "the time interval". The time interval is sometimes treated as a constant, at least in the short run, and sometimes as an endogenous variable.

The first approach may be rationalized by coordination problems. Suppose that a committee is in charge of nominal price changes. The date of the committee meeting must be coordinated among its members, and to facilitate this coordination they may decide to meet at more or less fixed time intervals: At the beginning of each month when average inflation is moderate, or at the beginning of each week when inflation is high.[1] If coordination costs are important, the time interval will depend mainly on aggregate variables which are proxies for the average inflation rate and less on store specific variables. An early attempt to estimate the frequency of price changes form aggregate variables is in Cecchetti (1985).

Store-specific variables are likely to be important if menu-type costs are important. Barro (1972) and Sheshinski and Weiss (1977) show that in a monopolistic environment, with fixed menu-type costs for changing nominal prices and even inflation, it is optimal

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[1] This type of coordination cost is not specific to price setting. For example, the time interval between two consecutive weekends tends to be fixed in the short run and rarely depends on individual factors like fatigue.

for sellers to increase their nominal price by (S - s) whenever the real price hits the lower bound s. See Weiss (1993) for an excellent survey of the more recent literature.

If there is heterogeneity across sellers, due to differences in market power, for example, the size of the band (S-s) is seller-specific. In this case, the time interval between two consecutive price increases should be relatively long for sellers who increase their nominal price by a large percentage and relatively short for sellers who increase their nominal price by a low percentage.

2. THE SEQUENTIAL TRADING MODEL

I now describe a framework that can host both hypotheses about the time interval. I use the sequential trading model in Eden (1994a) which like Prescott (1975) and Butters (1977), derives price dispersion as an equilibrium outcome. This model does not assume monopoly power and ex-ante heterogeneity among sellers. In equilibrium there is a tradeoff between price and the probability of sale, and the expected payoff is the same for all prices which support the equilibrium distribution. As a result, in a certain range, sellers are indifferent to the quoted price.[1]

For our purpose, the main difference between the sequential trading model and the (S,s) model in Sheshinski and Weiss (1977) is in the shape of the equilibrium expected profit function. According to the (S,s) model there is a unique real price which maximizes the

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[1] This idea is also used by Lucas and Woodford (1994), Rotemberg and Summers (1990), Williamson (1993), Eden (1985, 1986, 1990), Landsburg and Eden (1989), Bental and Eden (1993) and Eden and Griliches (1993).

expected profits of the monopoly. In contrast, there is a range of real prices which maximize the expected profits of a seller in the sequential trading model. This is illustrated by Figure 1.

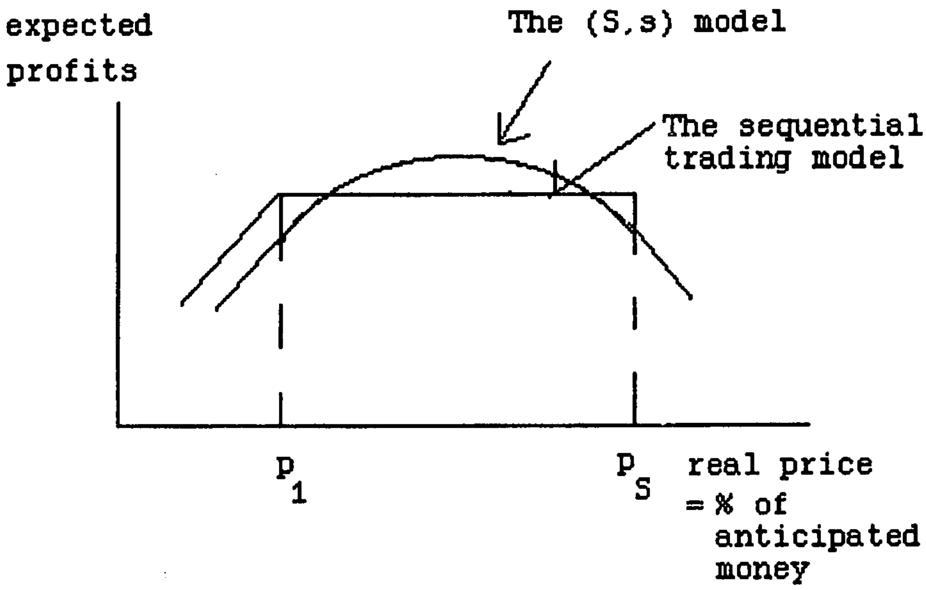

Figure 1

In Figure 1, the equilibrium real price range for the sequential trading model is $p_S - p_1$. When $p_1 \leq p \leq p_S$ expected profits are constant: a lower real price is associated with a higher probability of sale. When $p > p_S$, the probability of sale drops to zero and therefore expected profits are zero. The probability of sale is unity for all $p \leq p_1$, and therefore reducing the real price below p_1 reduces expected profits.

It follows that when inflation erodes the real price, the seller is compensated by an increase in the probability of sale as long as the real price remains in the equilibrium range. In this case there is no reason to change nominal price. From an individual seller's point of view it is perfectly rational to increase his nominal price by $p_S - p_1$ whenever the real price hits the lower bound

p_1: This (S,s) policy does not require fixed menu-type costs for changing nominal price and monopoly power.

The assumption that sellers are ex-ante identical does not imply a lack of variation in their choice of nominal price increases. Unlike some (S,s) models, the equilibrium real price distribution in Eden (1994a) does not depend on the rate of inflation. To maintain the equilibrium real price distribution, sellers will generally have to choose different rates of nominal price increases.

To illustrate, I assume that the equilibrium distribution of real prices (for a given product, across stores) is depicted by the areas A, B and C of Figure 2, and at time t-1 equilibrium prices prevailed. Between time t and t-1, inflation at the rate π erodes the real prices of all sellers, and as a result the distribution at the end of period t-1 shifts to the left: Sellers who were in area A move to C; sellers who were in B move to D and those who were in C move to E (thus by construction A = C = E and B = D).

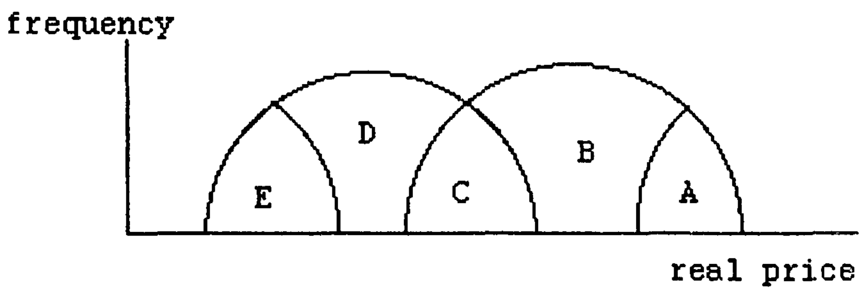

Figure 2

Assume, further, that to restore equilibrium only sellers who are below the lower support of the equilibrium real price distribution adjust nominal prices at the end of period t-1. This can be done as follows: Those who moved from A to C do not change their nominal price; those who moved from B to D move back to B by increasing their nominal price by π; and those who moved from C to E, move to A by increasing their nominal price by 2π.

The percentage change for the three groups are thus: 0, π and 2π. In this example sellers who change their nominal price by 2π will wait two periods before the next nominal price change and those who change it by π will wait one period. Thus the time interval between the changes is proportional to the percentage of change.

The sequential trading model allows for many other possibilities. For example, all sellers may change their nominal price each period by the same percentage π. But we can obtain uniqueness by adding small menu-type costs.[1] In this case, efficiency requires that there are no unnecessary nominal price changes. In the absence of real shocks, this will occur if the individual seller changes his nominal price only when the real price hits the lower bound of the distribution as in Figure 3. This implies a positive relationship between the rate of nominal price increase and the time to the next nominal price increase.

[1] The expected profit in the sequential trading model is still the same for all prices in the equilibrium range, but now the equilibrium expected revenue from lower prices must be somewhat higher than the expected revenue from higher prices, compensating for the need to change a low nominal price sooner.

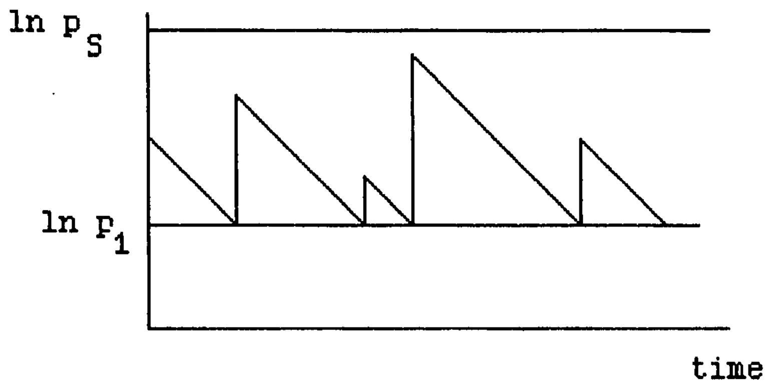

Figure 3

Figure 4, illustrates the alternative hypothesis in which the time interval is fixed.

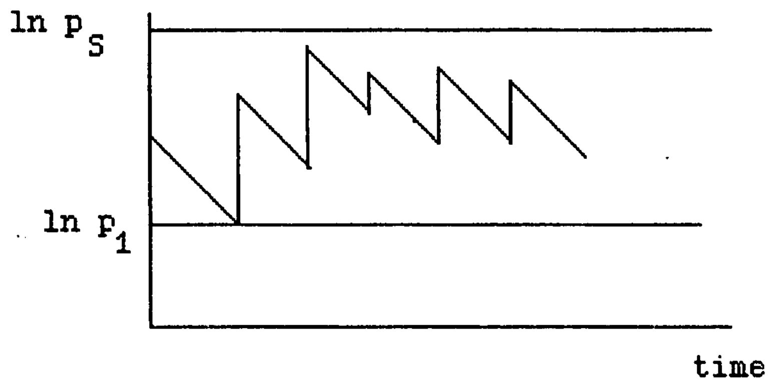

Figure 4

Assuming that the realized inflation rate was constant in the recent past allows us to derive stronger predictions. I consider two stores which changed the nominal price of the same product at the same time, say t = 10 in Figure 5. It is assumed that this change occurred after their real price hit the lower support of the equilibrium real price distribution, which is 10 in the example used. Assuming that at the last nominal price change episode, store 1 increased its nominal price by 10% (to 11) and store 2 by 5% (to 10.5), what can be said about the time interval between the two consecutive nominal price increases? If both stores experienced the same constant rate of inflation, then by similarity of triangles, the time interval for store 1 must be twice the time interval for store 2. In the example, the time interval for store 1 is 4 and the time interval for store 2 is 2. Thus, the ratio of the initial percentage increase is the same as the ratio of the time interval between changes: 10/5 = 4/2.

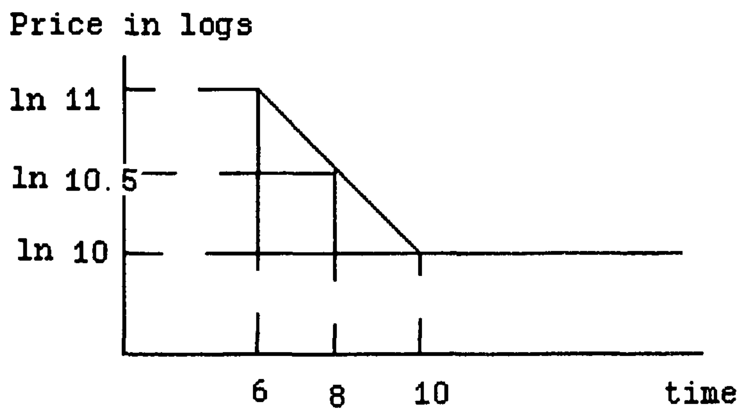

Figure 1

Nominal price reductions ("Sales") may introduce noise to the relationship between the percentage increase in nominal price and the time interval between changes. A Sale may occur when the seller spots a temporary disequilibrium that implies an expected profit-making opportunity. Sales may also occur when the seller has made a mistake and increased his nominal price above the upper support of the equilibrium distribution, or when the product has been on the shelf for a relatively long period and no longer looks new. In this case, the next price increase will be larger than average, reflecting in part the change in the perceived quality of the product. I therefore restrict attention to histories of positive nominal price changes. In particular, I test the following hypothesis.

Hypothesis: The time interval between two consecutive nominal price increases is proportional to the first rate of nominal price increase.

To test this hypothesis I introduce the following notations. I use i to index a product, j to index a store (seller) and t to index time (month).

N = the number of stores that have changed nominal price;

dp = the rate of the current nominal price change;

dpl (dp last) = the last nominal price change (the first lag of dp which is not zero);

dpll = the second lag of dp which is not zero;

MO (mean other) = the mean of dp across all other stores in the subsample used: $MO_{ijt} = \Sigma_{j'\neq j}\ dp_{ij't}/(N_{it} - 1)$;

MOL (mean other last) = the mean of dpl across all other stores in the subsample used: $MOL_{ijt} = \Sigma_{j'\neq j}\ dpl_{ij't}/(N_{ijt} - 1)$;

td = the time difference between the current nominal price change and the last time that such a change was made;

TO = the average td across all other stores in the subsample used: $TO_{ijt} = \Sigma_{j'\neq j}\ td_{ij't}/(N_{ijt} - 1)$;

Y = the percentage difference between the td and TO: $Y_{ijt} = \ln td_{ijt} - \ln TO_{ijt}$;

Z = the percentage difference between dpl and MOL: $Z_{ijt} = \ln dpl_{ijt} - \ln MOL_{ijt}$.

I look at all stores that have a history of positive nominal price changes. Specifically, I eliminate all observations which do not satisfy the following condition: $dp_{ijt} > 0$ and $dpl_{ijt} > 0$ and $dpll_{ijt} > 0$. The hypothesis requires two consecutive nominal price increases. The reason for requiring dpll > 0, is that dpll < 0, tends to increase dpl to correct for the "Sale".

The hypothesis implies:

$$Y_{ijt} = Z_{ijt} + \varepsilon_{ijt}. \tag{1}$$

The error term in (1) can result from human error: The manager was sick and did not change the price when it hit the floor. Or the manager was planning to go on a trip and changed the nominal price before it hit the floor. The manager may also have changed the nominal price because he spotted a temporary disequilibrium and an expected profit-making opportunity. Finally, there are measurement errors in Y: We observe only the month and not the day of the nominal price changes. This means that the error term is not normally distributed and therefore the reported standard errors should be interpreted with care.

The relationship (1) requires that the realized inflation rate was roughly constant before the current nominal price increase. It also requires that the same product-specific inflation rate was experienced by all stores. Later I relax this assumption and allow for "store-specific" inflation rates.

3. DATA

The data were collected by Israel's Central Bureau of Statistics as inputs for the computation of the consumer price index (CPI), and consist of monthly data from three periods: 1978 - 1979, 1981-1982 and 1991-1992. For the first two periods the data comprise prices of 26 food products (mostly meat and wine). These data were used by Lach and Tsiddon and are described in their 1992 article.

The data for 1991-92 which are new and are similar to the Lach and Tsiddon data. They consists of 119,748 monthly observations of prices by stores and products. These observations were collected from 458 stores selling 390 different products (each store sold only a subset of the products). I eliminated all products whose prices are controlled by the government. The definition of a product is rather narrow, for example, there are 10 different kinds of bread, two kinds of Coca-Cola and three kinds of olives.

The average monthly inflation rate was 4.1% in the 1978-79 sample and 6.2% in the 1981-82 sample. The 1991-92 sample was divided into three categories. The nominal price of all products increased by an average of 0.7%. The nominal price of all food products increased by an average of 0.8% as did "defined food", in which a weight is specified in the description of the product (see the Appendix in Eden (1994b) for a list of the products in each category).

The sample of strictly positive histories of nominal price changes (dp > 0 and dpl > 0 and dpll > 0) is used in this paper.[1] Table 1 gives summary statistics for these samples.

[1] An observation was included in this sample if it satisfies the condition dp > 0, dpl > 0 and dpll > 0 and there was no missing value for the price between the time at which dpll occurred and the time at which dp occurred. If there was a missing value for the price in this time interval, but the price after the missing value was the same as the price before the missing value, the observation was included under the assumption that the price did not change during the month in which there was no value for the price. Whenever the price before the missing value was different than the price after the missing value, we assumed, for the purpose of computing td, that the change in price occurred after the missing value.

Table 1: Summary statistics; sample dp > 0, dpl > 0 and dpll > 0.

	obs.	mean	standard deviation	(Min; Max)
		Variable = dp		
1978-79	2,299	0.10	0.095	(0.001; 2.513)
1981-82	4,413	0.12	0.084	(0.001; 1.001)
1991-92; all	3,850	0.067	0.051	(0.000; 0.486)
1991-92; food	2,381	0.070	0.052	(0.001; 0.486)
91-92; defined food	915	0.074	0.053	(0.001; 0.440)
		Variable = td		
1978-79	2,299	1.9	1.4	(1; 12)
1981-82	4,413	1.6	1.0	(1; 11)
1991-92; all	3,850	3.7	3.0	(1; 21)
1991-92; food	2,381	3.6	2.8	(1; 21)
91-92; defined food	915	3.0	2.5	(1; 20)

Clearly, the time interval td is not a constant. It varies inversely with the average rate of inflation. The question is whether the time interval varies in the short run according to store-specific variables. To test the hypothesis we now turn to regression (1).

4. RESULTS

The result of OLS regressions (1) for the 5 samples are given in Table 2.[1]

Table 2: dependent variable = Y ; sample dp > 0, dpl > 0 and dpll > 0.

	obs.	coefficient of Z	standard error	Adj. R^2
1978-79	2,234	0.16	0.01	0.05
1981-82	4,388	0.16	0.01	0.06
1991-92; all	2,577	0.18	0.02	0.04
1991-92; food	1,728	0.16	0.02	0.03
1991-92; defined food	767	0.18	0.03	0.05

The coefficients of Z are similar across samples and all are close to 0.17. This suggests a rejection of the hypothesis that the coefficient of Z is unity.

One possible reason for a deviation of the coefficient of Z from unity is variations in the rate of inflation in the relevant time interval. But the coefficient of Z is similar across very different samples: Inflation rose in 1978-82 and fell in 1991-92.

Store-specific inflation rate is another possible reason for the large deviation of the coefficient of Z from unity. In Eden (1994a) changes in the money supply affect all stores in the same way. This model may be extended to the case in which stores are

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[1] The estimated probability of heteroskedasticity is small and the asymptotic standard errors are almost identical to the standard errors from the simple OLS regressions.

spatially separated and the rate of change in the money supply differs across locations. In this more general case, the rate of inflation varies across products and stores.

To see if store-specific inflation can explain the above finding, I use π_{ijt} to denote the rate of inflation of product i experienced by store j in the relevant period before time t. Under the hypothesis that stores change their nominal price when their real price hits a store-specific lower bound we should have: $td_{ijt} = dpl_{ijt}/\pi_{ijt}$. This says that a store which increased its nominal price by 10% will wait two months if it experiences a monthly rate of inflation of 5%. As before, I assume that there is a random component in the manager's behavior (due to sickness, forgetfulness, and other random constraints on his behavior) which is captured by an i.i.d. shock ε. This leads to:

(2) $$\ln td_{ijt} = \ln dpl_{ijt} - \ln\pi_{ijt} + \varepsilon_{ijt}.$$

Under the hypothesis that all stores change their nominal price when the real price hits a (store-specific) lower bound, the variable MOL_{ijt}/TO_{ijt} is the rate of product i inflation experienced by other stores which changed the nominal price of product i at the same time as store j. We can therefore use MOL_{ijt}/TO_{ijt} as a proxy for π_{ijt}.

I assume that at the time of the last nominal price change store j's prediction of its own rate of inflation conditional on MOL and TO can be written as:

(3) $$\ln\pi_{ijt} = \ln MOL_{ijt} - \ln TO_{ijt} + \upsilon_{ijt} ,$$

where υ_{ijt} is an i.i.d. shock with zero mean which is independent of the variables $\ln MOL_{ijt}$ and $\ln TO_{ijt}$. Since υ_{ijt} is not known at the time that dpl is chosen, υ_{ijt} is also independent of $\ln dpl_{ijt}$. Substituting (3) into (2) leads to:

$$(4) \qquad \ln td_{ijt} = \ln dpl_{ijt} + \ln TOL_{ijt} - \ln MOL_{ijt} + \upsilon_{ijt} + \varepsilon_{ijt}.$$

Note that (4) is a less restricted form of (1). Under the assumptions made here the missing variable υ_{ijt} is not correlated with the explanatory variables, and therefore the OLS estimates of the coefficients are not biased.

The OLS regression results in Table 3 do not support the hypothesis about the coefficients. The coefficient of dpl is again similar across samples but now it is even lower: Being in the neighborhood of 0.13.

Table 3*: dependent variable = lntd ; sample dp > 0, dpl > 0, dpll > 0.

	lndpl	lnTO	lnMOL	Adj. R^2
1978-79	0.13 (0.01)	0.70 (0.03)	-0.16 (0.02)	0.53
1981-82	0.13 (0.01)	0.60 (0.02)	-0.18 (0.01)	0.45
1991-92; all	0.11 (0.02)	0.59 (0.02)	-0.23 (0.02)	0.71
1991-92; food	0.11 (0.02)	0.59 (0.02)	-0.23 (0.02)	0.72
1991-92; defined food	0.14 (0.03)	0.67 (0.04)	-0.23 (0.03)	0.72

* Standard errors in parentheses.

The alternative hypothesis is that the time interval is roughly constant in the short run and tends to change slowly and uniformly across sellers. This suggests a coefficient of unity for lnTO and a zero coefficient for lnMOL and lndpl. This hypothesis is also clearly rejected by the data. The coefficient of lnTO is about 0.65 (again, surprisingly similar across samples) and the coefficients of lnMOL and lndpl are small but significantly different from zero. The results suggests that the "truth" is somewhere between the hypotheses considered.

5. CONCLUDING REMARKS

I used large data sets on prices disaggregated by products and stores. These data sets are from three periods in the recent Israeli history: 1978-79 with a monthly inflation rate of about 4%, 1981-82 with a monthly inflation rate of about 6% and 1991-92 with a monthly inflation rate of about 1%. I use a subsample in which positive nominal price changes were preceeded by two other positive nominal price changes.

The main finding is that when holding a proxy for the rate of inflation constant, the elasticity of the time interval between two consecutive nominal price-increases with respect to the first nominal price-increase is positive but much less than unity. Surprisingly the estimated elasticity is similar across periods with very different inflation rates. It is about 0.16 when a specification that does not allow for store-specific inflation rate is used (Table 2) and about 0.13 if we allow for store-specific inflation rate (Table 3). This can be interpreted as a rejection of the (S,s) model in Sheshinski and Weiss (1977). I doubt whether more sophisticated (S,s) models would yield drastically different predictions about this elasticity. We can also reject the hypothesis that stores change their nominal prices only when their real price hits the lower bound of the equilibrium real price distribution in Eden (1994a).

One possible interpretation is that coordination-type costs are important relative to menu-type costs.[1] Take the example in which a

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[1] Note that both coordination-type costs and menu-type costs serve only as tie-breakers in Eden (1994a) and in this sense, both types

committee meeting is required for changing nominal prices, and to facilitate coordination problems they tend to meet at time intervals which are relatively fixed in the short run. Assume that there are other items on the agenda of this meeting and nominal price changes are not always discussed. The priority of the item "nominal price changes" gets higher when the real price gets closer to the lower bound of the equilibrium real price distribution. This implies a negative relationship between the time interval and the realized inflation rate as measured by MOL/TO and a positive relationship between the time interval and dpl. But there is no clear prediction about the elasticities.

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of costs are unimportant. The statement is therefore on the relative importance only.

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R. Melnick and E. Yashiv - The Macroeconomic Effects of Financial Innovation: The Case of Israel. 94.05

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94.10 א. ארנון, ג'. וינבלט, - פוטנציאל הסחר בין ישראל, הפלסטינים וירדן.

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94.12 ק. פלוג, נ. (קלינר) קסיר - הציות לחוק שכר המינימום בסקטור העסקי.

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94.14 א. ספיבק, - משבר קרנות הפנסיה בישראל: מסגרת מושגית ובחינת ההמלצות לפתרון.

94.15 ל. מרידור, ש. פסח - שער החליפין הריאלי בישראל: פרספקטיבה של שלושה עשורים.

B. Eden - Time Rigidities in The Adjustment of Prices to Monetary Shocks: An Analysis of Micro Data. 94.16