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Abstract

We analyse the behaviour of price setters in Poland during transition from a planned to a market economy, using a large disaggregated data set. The size of price changes and the frequency of adjustment both fall as the inflation rate declines. Price variability increases with inflation. The effect of expected inflation is much stronger than the effect of unexpected inflation. Price setters are forward looking and follow a mixture of state- and time-contingent policies. These results are broadly consistent with the menu cost model. The bottom line is this: it takes sellers of sausage, eggs, toothpaste, vacuum cleaners, car-wash operators etc. just a few years to figure out how to adjust prices in a market environment.

I. Introduction.

We analyse the behaviour of price setters using a large disaggregated data set for Poland. During the period under analysis the Polish economy undergoes dramatic changes. The data start in January 1990, when a big bang transition to market economy begins. A brief chronology is as follows. In August 1989 the anti-communist opposition unexpectedly won the general election. A new government was formed in September. A package of radical economic reforms was prepared by November and introduced on January 1, 1990. The planned system was abolished, price controls were removed on almost all goods and services, exchange and import controls were relaxed and firms were freed from restrictions on the choice of suppliers.[1]

The paper focuses on the behaviour of price setters: store owners and store managers. In a planned economy prices were set by a central authority or by the producer and were identical in all stores (with the exception of street markets). The role of store managers was to sell, in a noncompetitive environment, goods which were often in short supply. When reforms start, all of a sudden they gain the ability to set prices, choose suppliers and even become owners. Shortages rapidly disappear and managers are faced with a competitive market.

This environment creates a natural experiment, which allows us to answer several interesting questions. How do firms set prices following a regime shift, in an unstable environment, without history on which to base expectations? Would their pricing policies be similar, or different from, those observed elsewhere? How fast do people learn? Is it possible to pin down their expectations?

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[1] See Sachs (1993) for a detailed description and analysis of the reforms.

To answer these questions we analyze the effect of inflation on the size and frequency of price adjustment, on the variability of relative prices across locations and on the time pattern of price changes, using store-level data on prices of 52 products and services.

Our results are summarized as follows:

1. The size and the frequency of price changes are positively correlated with the inflation rate.

2. Relative price variability is positively related to inflation.

3. Expected inflation has a bigger effect on relative price variability than unexpected inflation.

4. The effect of inflation on relative price variability is stronger when variability is measured as the cross-sectional variance of rates of price change than when it is measured as the coefficient of variation of price levels.

5. Price setting policies are forward-looking.

6. Firms follow a mixture of state- and time-contingent pricing policies.

The first three results are the same as in Lach and Tsiddon (1992) who analyze a similar disaggregated data set for Israel. Their data cover the period of 1978-1984, which follows a decade of very rapid inflation. As is well known, the long inflationary period led to many changes in the Israeli economy aimed at reducing its negative effects. In particular, it can be expected that a typical Israeli firm developed optimal, or near optimal, pricing policies to deal with high aggregate inflation rate. We treat Lach and Tsiddon's (1992) results as a benchmark and so the bottom line is this: even in an entirely new economic environment, sellers of sausage, eggs, toothpaste, vacuum cleaners, car-wash operators etc. require just a few years to figure out how to form inflationary expectations and adjust prices in a market environment.

Lach and Tsiddon's (1992) results suggest that the relationship between inflation and relative price variability is due to menu costs (Barro, 1972, Sheshinski and Weiss, 1977). Our results are stronger: while in their study the average coefficient on expected inflation is 20% larger than on unexpected inflation, in our data it is twice larger. When we regress relative price variability on the absolute values of the expected and unexpected inflation, the coefficient on (absolute) expected inflation is eight times larger.

Given these findings the second goal of the paper is to take a look at other predictions of the menu cost model. While the data are not sufficient for a direct test, our results are broadly consistent with the model.

The plan of the paper is as follows. The data are described in section 2. This is followed by nonparametric analysis. In section 4 we analyze the relationship between inflation and relative price variability. Evidence suggesting that agents are forward looking is in section 5. In section 6 we ask whether pricing policies are time- or state-contingent. The last section concludes.

II.   Data.

The data set consists of store level price information on selected products and services in Poland. The data start at the beginning of the big-bang transition and cover the period of seven years from 1990 until 1996. Prices from before 1990 are not comparable as the methodology was completely revamped in January 1990. The data were collected by the Polish Central Statistical Office (GUS) in order to calculate the Consumer Price Index. GUS compiles price information on 1500-1800 products in 307 districts, with one store per district (Bauc et al, 1996, p. 55). Out

of this set we have a balanced sample of data for 52 goods, each in 47 stores (districts). They comprise the complete set for four out of 49 administrative regions in Poland (Voivodships). The prices are the actual transaction prices, as coupons or discounts were very rare or nonexistent over the study period. In addition, we also have data on the national average prices of the 52 goods.

We selected the goods on the basis of several criteria. They had to be well defined, excluding, for example, "a man's suit". Most importantly, we required that they did not change during the period in question. We excluded goods with regulated prices, and with many missing observations. Finally, we wanted to have data on a variety of products. This was difficult as a lot of goods had changed following the collapse of the planned economy. In the end, out of the 52 products, 37 are groceries (19 perishable and 18 storable), 2 are sold in cafeterias/cafes, 10 are nongrocery items and 3 are services. The list of the goods is in Table 4.

The frequency of observations varies over time and across goods. There are between two and four observations per store each month in 1990 and between one and three observations per store each month in 1991-96, depending on the good. Whenever a store is sampled more than once a month, the month is divided into equal intervals and there is one observation per interval. As we would like to take advantage of the variety of products in the data, for a part of our study we use the first observation in each month. For foodstuffs there are 4 observations per store a month in 1990 (except in December 1990, when prices were not collected in the last week) and 3 observations per store a month for the period 1991-96. Hence, for a subsample of goods: the 37

foodstuffs listed in Table 4 plus one perishable and two durable foodstuffs[2], we also use the more frequent (three observations a month) data for the period 1991-96. One problem with the data is that we do not know the identity of the store where a price is being sampled. The procedure is as follows. Each Voivodship is divided into several districts (for example, there are 14 districts in Warsaw)[3]. One store is chosen in each district. Price inspectors are to choose the same store each time, but this is not enforced and changes of stores are not recorded.

Apart from this technical issue, some changes are due to the transformation in the retail sector following the collapse of the planned system. Table 1 provides data on changes in store ownership in the country as a whole. It probably understates ownership changes in our data, since many privatized stores were small and were less likely to be visited by price inspectors than state or cooperative stores. Also, a store may alter its pricing strategy due to, for example, management change. Finally, while we took great care to select goods which did not change during the period in question, the competitive environment was altered both by increased availability of substitutes as well as changes in the retail and service sectors. For example, the price for car wash service (good number 49) is the price for washing a specific make and type but, during the period 1990-96, the car became much less common and there appeared car washes offering different levels of service.

The data set is not complete. The proportion of missing observations varies across goods and time. This proportion decreases almost monotonically over time, from 58% in 1990 to 23%

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[2] Salt, luncheon meat and condensed milk. For these three goods data for 1990 are missing.

[3] We picked the Voivodships with the largest number of districts.

in 1996. Missing observations are not due to market shortages as these disappeared quickly

within the first few months following the big-bang transition, while the proportion of missing

data falls throughout the sample period.

III. Preliminary Analysis.

We begin the analysis by taking a cursory look at the annual data. They are illustrated in

Figure 1. We plot the inflation rate, and the proportions of increased, unchanged and decreased

prices as well as the ratio of decreases to increases. We obtain these proportions by computing

the ratios of increases, unchanged prices and decreases to the number of cases with two

consecutive observations.

The annual inflation rate is the 12-month price increase December to December. The rate

in 1990 was 249% (due to the price shock following the freeing of most prices in January 1990)

and is omitted from the graph for clarity. The 12-month inflation rate falls systematically over

time.[4] Throughout the sample period it increases only in March 1993, Oct-Nov 1994 and Feb-

May 1995. Monthly inflation is, however, much more erratic and increases in the monthly

inflation rate are frequent.

All series follow a similar, nearly monotonic, pattern. As the inflation rate declines over

time, the proportion of price increases and of price decreases falls, while the proportion of prices

which are left unchanged rises. Price decreases become relatively less common than price

increases, except for 1991. The fact that some price observations are missing creates a bias in the

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[4] The general picture of a near monotonic decline in 12-month inflation rates does not depend on the choice of the
month.

duration of price changes, computed as the average length of the period during which the nominal price is held constant. The more frequent are gaps in the data, the less likely are long, uninterrupted price observations and so the less likely are we to observe long price durations. The proportion of missing data is positively correlated with inflation. This creates a spurious inverse relationship between duration and inflation. To avoid this problem, instead of price duration we show the probability of a price change in a single observation. It is calculated as the ratio of the number of price changes to the number of observations in which a change could occur, i.e. the number of two consecutive price observations. Using these data we compute the probability of price change per month. In a complete data set this probability would be monotonically related to the duration of prices. The average values are in table 2. The monthly probability of a price change decreases with inflation. The same pattern is observed at the level of individual goods, but with some exceptions. The probability is the highest for perishable foodstuffs, somewhat lower for durable foodstuffs and much lower for manufactured goods and services. Between 1990 and 1996 it declines the most for services and the least for durable foodstuffs. The extreme values are 0.94 for eggs in 1990 and 0.06 for ECG test in 1996.

Information on the size of price changes is in Table 2. The average size of price increases falls as the inflation rate declines over time, except in 1995. It varies from over 30% in 1990 to about 10% in 1994-96, after the inflation rate has fallen from the high levels which followed the initial price shock. The bias discussed above means that the data in Table 2 most likely underestimate the changes in adjustment size. For individual goods the average size of changes varies from 7.1% for one of the meat products to 36% for EKG test. Price increases are

the smallest for perishable foodstuffs (in particular meats) and the largest for manufactured products.

The average size of price decreases falls from 12% in 1990 to 7% in 1993 and increases somewhat in the last three years. The size of price decreases varies much less over time than the size of price increases. While in 1990 the average decrease is equal to about 40% of average increase, in the last three years they are of similar size. Decreases are the smallest for foodstuffs and the largest for services and manufactured products.

The recorded proportion of price decreases varies between 27% in 1990 and 15% in 1996 of all price <u>changes</u>. These numbers appear to be large, judging by earlier results.[5] In principle, the large proportion of price decreases may be the result of price inspectors not always sampling the same store. It is not likely that this factor dominates pricing behaviour in our data. If price changes were due to changes in sampled stores (or in pricing policies) their distributions would be generated by random sampling from the contemporaneous distribution of price levels. In Konieczny and Skrzypacz (2000), however, we find that the distribution of price levels rapidly approaches long-run values. The probability of finding a price which is over 10% (20%, 30%) away from the geometric mean of prices of the given good is 52% (28%, 16%, respectively) in January 1990, falls to 30% (14%, 6%, respectively) by January 1991 and varies little throughout

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[5] Cross-country evidence suggests that price decreases become more frequent as the rate of inflation rises. The intuition is that, as the rate of inflation is high, the behaviour of prices is more erratic and mistakes or price experimentation on the part of price setters happen more often. In Cecchetti (1986) the inflation rate is 0.2-12% per year and there are no price decreases; in Sheshinski, Tishler and Weiss (1979) inflation is 0.25-2.9% per month and 2% of changes are decreases; in Lach and Tsiddon (1992) inflation is over 4% per month and 15% of changes are decreases; in Tommasi (1993) the inflation rate is in the range -5% to +38% per week and there are 36% price decreases. One exception is Dahlby (1992); in his data inflation is 8-12% per year and there are 7% price cuts.

the rest of the sample. In contrast, the average size of price increases falls continuously until 1995, and the average size of price decreases falls continuously until 1993. Finally, as we show below, price changes are related to various measures of inflation; this relationship is unlikely to be generated by store changes.

To sum up, as inflation falls price changes become less frequent. This is consistent with earlier findings (Sheshinski, Tishler and Weiss, 1981, Cecchetti, 1986, Danziger, 1987, Dahlby, 1991, Lach and Tsiddon, 1992, Tommasi, 1993, Kashyap, 1995 and Fisher and Konieczny, 1999). The finding that the size of price changes falls as inflation declines is less common: in Sheshinski, Tishler and Weiss (1981) and in Cecchetti (1986) there is little effect of inflation on adjustment size, while Lach and Tsiddon (1992) and Kashyap (1995) find several instances when price changes become larger as inflation falls.

IV. Inflation and Relative Price Variability.

There is substantial empirical literature on the relationship between inflation and relative price variability (see, for example, Mills (1927), Vinning and Elwertowski (1976), Parks (1978), Fisher (1981), Domberger (1987), Van Hoomissen (1988), Lach and Tsiddon (1992), Parsley (1996) and Debelle and Lamont (1997)). The general conclusion of this literature is that various measures of relative price variability are positively related to inflation.

The two main explanations of this relationship are based on the menu cost and on the incomplete information approaches. When price changing is costly, changes in the inflation rate affect the size and the frequency of price changes (Sheshinski and Weiss, 1977, 1983). Variability of relative prices increases with inflation if price changes are not perfectly

synchronized. The incomplete information model of Lucas (1973) suggests that the reason is the inability of firms to distinguish between aggregate and local shocks. Relative price variability increases with inflation if the history or persistence of local shocks and/or supply and demand elasticities differ across markets. We discuss the implications of both approaches below.

Lach and Tsiddon (1992), henceforth L-T, point out that the menu cost approach implies that relative price variability is affected by <u>expected</u> inflation while the incomplete information approach implies a relationship with <u>unexpected</u> inflation. They analyse a disaggregated data set on prices of foodstuffs in Israel during 1978-1984. The effect of expected inflation on relative price variability is stronger than the effect of unexpected inflation.

We treat the Lach and Tsiddon's (1992) (L-T) results as a benchmark. They study the Israeli economy after it had gone through more than 10 years of rapid inflation. It was among the most developed countries which experienced substantial and prolonged inflation. Israeli price setters can be expected to use optimal, or near optimal, pricing policies. The comparison of the price behaviour of Polish firms, therefore, sheds light on their ability, and the speed of, adaptation to inflationary environment.

Two measures of relative price variability have been used in the literature. Parks (1978) proposed that it be evaluated by the cross-sectional variance of rates of change of prices. This measure has been used by, for example, Domberger (1987), Van Hoomissen (1988) and L-T. Define the price of good i in store j at time t as P_{ijt}. Whenever we have two consecutive observations in a given store we can calculate its rate of change between t-1 and t:

$DP_{ijt} \equiv \ln P_{ijt} - \ln P_{ijt-1}$. Relative price variability is defined as the standard deviation of DP_{ijt}

across stores, SDP_{it}:

$$(1) \qquad SDP_{it} = \left[\frac{1}{N_{it} - 1} \sum_j (DP_{ijt} - DP_{it})^2 \right]^{1/2}$$

where N_{it} is the number of observations in which price change could be observed (i.e. the

number of two consecutive nonmissing observations) and $DP_{it} \equiv (1/N_{it}) \sum_j DP_{ijt}$ is the in-sample

rate of inflation of good i at time t.

An alternative measure of relative price variability is the coefficient of variation of price

levels across stores, CV_{ij} :

$$(2) \qquad CV_{it} \equiv \left[\frac{1}{N_{it} - 1} \sum_j \left(\frac{P_{ijt} - P_{it}}{P_{it}} \right)^2 \right]^{1/2}$$

where $P_{it} \equiv (1/N_{it}) \sum_j P_{ijt}$ is the average price of good i across stores at time t.

The benefit of using both measures of relative price variability is that they allow

distinguishing between the menu cost and the incomplete information explanations of the

relationship between variability and inflation. When the relationship is the consequence of menu

cost considerations, results may differ depending on which measure is used. On the other hand,

when the relationship is the consequence of incomplete information, it is the same for both

measures.

In the menu cost approach, a firm which faces fixed costs of price adjustment sets two

bounds for its real price and keeps its nominal price constant as long as the real price is between

the bounds. Whenever inflation reduces the real price to the lower bound, the firm increases the nominal price so as to reset its real value at the upper bound. As the expected inflation rate rises, the size of price changes increases (Sheshinski and Weiss, 1977, 1983). The effect on the frequency of price changes is ambiguous; a sufficient condition for it to increase is that the real profit function be strictly concave in the log of the real price.

The fact that the choice of the variability measure matters is easiest to demonstrate in a situation in which identical firms stagger price changes uniformly over time, as in Caplin and Spulber (1987). Assume that, as inflation increases, the frequency of changes rises but the size is unaffected. Consider the ordering of firms by the real price. Price changes affect the position of firms in the ordering, but do not affect the distribution of real prices. A firm which just changed its price simply goes from having the lowest real price to the highest. Hence, as long as the size of price adjustment does not change, the *CV* measure of relative price variability, which is based on price levels, is not affected. On the other hand, the *SDP* measure of variability, which is based on the rates of price change, increases.[6] [7]

When the relationship between inflation and relative price variability is the consequence of incomplete information, both the *SDP* and the *CV* measures of variability increase. The size of an individual firm's price response to an unexpected increase in inflation depends on its idiosyncratic history of shocks and its market conditions. As a result, both the sizes of price

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[6] As long as a minority of firms change prices between consecutive observations. Danziger (1987) shows that, for the positive correlation to hold, observations must be at least twice more frequent than price changes. In our data this requirement is met by all goods except good 9 - fresh eggs.

[7] If higher inflation results in larger price changes but the frequency of adjustment is not affected, both the *SDP* and the *CV* measures of relative price variability increase, as long as price changes are not perfectly synchronized.

changes, as well as the observed price levels, differ across firms.

Empirical studies find that the frequency of price changes increases with inflation, but price changes sometimes become smaller (Lach and Tsiddon, 1992, Kashyap, 1995) or are not affected by inflation (Sheshinski, Tishler and Weiss, 1981, Cecchetti, 1986). Price changes appear to be staggered (Lach and Tsiddon, 1992 and 1996, Tommasi, 1993, and Fisher and Konieczny, 2000).

In our data both the frequency and the size of price changes increase with inflation. This implies that, if menu costs are the reason for the positive relationship between inflation and relative price variability, we should find that expected, rather than unexpected, inflation matters and the effect should be stronger when variability is measured by *SDP*. If the relationship is the result of incomplete information, we should find that unexpected inflation matters and the effect should be similar for both measures of variability.

The division of inflation into its expected and unexpected part is difficult in the "natural experiment" economy. We simply do not know how people form expectations following a dramatic regime change. While there was, at times, significant inflation in Poland prior to the "big bang" transition,[8] its nature was quite different than the subsequent inflationary process. Most prices (with the exception of fruits, vegetables as well as prices in street markets) were regulated; price increases required approval of planning authorities. In all stores (with the exception of street markets) prices were identical. Inflation in 1980's was a result of the

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[8] Until 1970 the inflation rate was low, in the range of 0-3% per year. In 1970's it varied between 0 and 10%. In 1981-88 there was unprecedented (in a planned economy) inflation, which varied between 11% (1985) and 103% (1982). In the fall of 1989 many prices were freed and inflation reached over 200% in that year.

planners' attempt to reduce rampant shortages. Its rate was determined by the whim of bureaucrats, rather than by observable things like the money supply; even the degree of shortages for individual goods did not play much role. Moreover, the big-bang reforms in January 1990 changed completely the organization of the economy (for example, the exchange rate policy), altering the relationship between inflation and aggregate variables. Hence any expectation mechanism or "rules-of-thumb" developed prior to 1990 were useless for the period following the jump to a market economy. Finally, an additional difficulty for the Lucas' approach is that past history of local real shocks is an unreliable predictor of the current shock structure.

There are several estimates for the expected CPI rate of inflation in Poland. They are produced by various research institutes; Reuters collect survey data. Unfortunately, they do not cover the whole period of study. Hence we construct our own measure of expected CPI inflation. Also, for each good, we construct a measure of expected own inflation (i.e. the inflation rate for that particular good). The reason is that CPI is not a good proxy for price pressures in a given market. The price structure inherited from the planned economy was artificial and, as a result, there are large relative price changes as prices adjust to market clearing levels.

We construct expected CPI inflation by regressing inflation on its past values, time, time squared and monthly dummies. Expected own inflation is obtained in a similar fashion, except that we include past values of own inflation as well as of CPI inflation. To make things as simple as possible, we chose a model with three lags of the dependent variable (and three lags of CPI in the regressions for own inflation). Time and time squared are included to control for transition-induced changes in the economy structure; monthly dummies control for seasonal

effects. We take the explained part of inflation to be the expected inflation and the residual to be the unexpected one.

We chose this simple approach for two reasons. First, given the degree of disaggregation in our data, collecting additional market-level information was not practical. Second, we want to compare the results to those obtained by L-T. There is an advantage of using a simple approach in a comparison like this, as it avoids atheoretical "massaging" the data. In the end, our approach is similar to that in L-T.[9]

To estimate the effect of inflation on relative price variability we ran OLS regressions, separately for each good, with various measures of own inflation and aggregate inflation, as well as time, time squared and monthly dummies as explanatory variables. A typical regression is:

$$(3) \qquad Variability_{it} = \beta_{io} + \beta_{i1}INFE_{it} + \beta_{i2}INFU_{it} + \beta_{i3}\,t + \beta_{i4}\,t^2 + \boldsymbol{\beta d_{md}} + e_{it}$$

where *Variability* is a measure of variability: *SDP* or *CV*, as defined in equations (1) and (2), INF_{it} denotes own inflation of good i at time t, $INFE_{it}$ and $INFU_{it}$ are its expected and unexpected parts, and $\boldsymbol{d_{md}}$ is the vector of monthly dummies. Own inflation is better than CPI as it reflects changes in demand/supply conditions in a given market, which affect the price bounds in the menu cost model (Cecchetti, 1986) and local response in the incomplete information approach. The data used for own inflation are national averages for the given good, rather than in-sample averages. In order to compute the rate of inflation for a particular good, GUS first calculates the

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[9] They construct the measures of expected own inflation by regressing inflation on past values of own and of CPI inflation and various time-related dummies. The lags are chosen on the basis of F-test. They select three lags of the CPI for all goods, and three lags of the dependent variable for 80% of goods. In our case CPI inflation has little effect, due to large changes in relative prices.

price level in each voivodship and then computes the national price level as the unweighted average of the 49 voivodship values. Our data cover 4 out of the 49 voivodships, or about 8% of GUS's sample. Time is included as a proxy for structural change; we expect the change to be fast initially and slow down over time, as the economy approaches the new equilibrium. To account for this we also include time squared. Finally, monthly dummies are included as we have many seasonal goods.

The inflation rate, as well as both measures of relative price variability, were very high at the beginning of 1990 and so the initial observations are outliers. We estimate the relationships using data for the period May 1990 to December 1996. The exclusion of the February-April 1990 data (we cannot calculate price changes for January 1990) allows to avoid the results being dominated by the three outliers. If these months are included, the positive relationship between inflation and relative price variability is much stronger. Eliminating the subsequent months has little effect on the results. The additional benefit is that all regressions cover the same period (the measure of expected inflation can be calculated from April 1990 on).

We found that the stochastic components have nonspherical distribution. In particular, we detected significant autocorrelations, correlations across goods and heteroscedasticity.

Serial correlation was first detected using Durbin-Watson statistics. It was confirmed when we tried to estimate the model with the Maximum Likelihood method using different AR(p) specifications for the disturbances: the lagged values of disturbances turned out to be statistically significant.

One reason for the heteroscedasticity may be that our sample is unbalanced: not every

good is quoted in every location every time. For some locations and months we are not able to

calculate the monthly inflation rates and so the number of observations over which we calculate

SDP or *CV* varies over time and over goods.

We tried to find a way to estimate the models more efficiently than by OLS. Experiments

with AR(p) specifications showed that the distribution of the disturbances is not simple. For

example, when we tried an AR(12) model, different lags turned out to be significant for different

goods, without any visible pattern. Besides, in several cases the ML estimation could not be

conducted because, at the initial estimates, the distribution was nonstationary under the

hypothesis of AR(12).

In the end we decided to use simple OLS estimation with consistent non-parametric

estimation of standard errors. Despite the loss of efficiency, this approach is used because, first,

the results are still significant and, second, the complex form of the distribution of disturbances

could cause loss of consistency if the AR(p) model is misspecified. To estimate standard errors

we employ the methods described in Newey and West (1994) and Andrews and Monahan

(1992)[10].

We also tried to estimate the model as a seemingly unrelated regression system. The

coefficients were quite similar and they were significant in similar instances as those in OLS

estimation. As we don't know the true form of the covariance matrix of the disturbances,

however, we cannot be sure that these estimates are more efficient than the OLS ones. Also, in

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[10] Following recommendations in this literature based on Monte-Carlo results, we use a quadratic-spectral kernel
(Andrews (1991) and Andrews and Monahan (1992)), prewhitening procedure (Andrews and Monahan (1992)) and
automatic lag selection (Newey and West (1992)).

our model, (rational) expectations play a crucial role and we don't know how exactly these are formed on the basis of observed variables and available information. Under those circumstances SUR estimators may be inconsistent. We decided not to draw any conclusions from these estimates and they are not reported here.

Regression results are in tables 3-5. In tables 3 and 5 we report the average, maximum and minimum values of the coefficients in the regressions for individual goods as well as the number of significant coefficients at the 5% level, using a two-sided alternative; in table 4 we report the results of selected regressions for individual goods.

We begin by replicating the analysis in L-T and then conduct additional tests. In column 1 of table 3 the summarized regression is:

(4) $$SDP_{it} = \alpha_{io} + \alpha_{i1}INF_{it} + \alpha_{i2}\,t + \alpha_{i3}\,t^2 + \boldsymbol{\alpha d_{md}} + u_{it}$$

where INF_{it} is own inflation of good i at time t and $\boldsymbol{d_{md}}$ is the vector of monthly dummies. Relative price variability, as measured by the standard deviation of the rates of price change, is positively related to the inflation rate. The coefficient on inflation is positive for most goods; it is positive and significant for 38 out of the 52 goods (in all cases significance is at the 5% level and the tests use two-sided alternatives). The coefficient is never negative and significant.

In columns 3 and 4 of table 1 we summarize the results when own inflation is split into its expected and unexpected parts:

(5) $$SDP_{it} = \beta_{io} + \beta_{i1}INFE_{it} + \beta_{i2}INFU_{it} + \beta_{i3}\,t + \beta_{i4}\,t^2 + \boldsymbol{\beta d_{md}} + e_{it}$$

These are, essentially, the same regression as in L-T (their equations (2') and (2),

respectively); the differences are that we include time and time squared. The motivation for including the additional variables is discussed above.

It is clear that the effect of expected inflation is stronger than the effect of unexpected inflation: the average coefficient on expected inflation is twice bigger and it is significant for more goods. It is negative and significant for one good (40 - citric acid). The effect of time is as predicted: variability falls with time but at a decreasing rate.[11] Monthly dummies are jointly significant. Results for individual goods are in table 4. The coefficient on expected inflation is larger than on unexpected inflation for 40 out of 52 goods; the difference is significant for 13 goods. In one case (good 26 - apple juice) the difference is negative and significant.

Our results are qualitatively identical to those in L-T, despite the fact that the method of calculating expected inflation we use is, if anything, overly simplified. Quantitatively, our results are quite similar but the difference between the effect of expected and unexpected inflation is much larger. All goods in their study are foodstuffs sold in a store, so in the comparison below we look only at goods 1-37; the remaining goods are either industrial products or services. The values from Lach and Tsiddon are in brackets. The average coefficient on own inflation in regression (4) is 0.39 (0.41); the average coefficient on expected inflation in regression (5) is 0.57 (0.43); the average coefficient on unexpected inflation is 0.30 (0.36).[12] The coefficient on

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[11] The average coefficient on time (i.e. the marginal effect of time, equal to $\hat{\alpha}_{i2} + 2\hat{\alpha}_{i3}t$, evaluated at sample mean, where $\hat{\alpha}$ is the estimated coefficient in equation (5); and similarly for equation 4) is negative for 50 goods in regression (5) and for 52 goods in regression (4).

[12] The difference between the median coefficients on expected and unexpected inflation in our and Lach and Tsiddon's data is even larger: the median coefficient on expected inflation is 0.6 (0.46); the median coefficient on unexpected inflation is 0.23 (0.37).

expected inflation is greater than the coefficient on unexpected inflation for 30 out of 37 goods (16 out of 26). The difference is positive and significant in 9 (5) cases.[13]

What is responsible for the stronger effect of expected inflation on relative price variability in the Polish data? We can only speculate that, despite the fluid economic environment, inflation was easier to predict. Its rate was lower than in Israel and fell systematically over time. While the monthly inflation rate was erratic, the 12-month inflation rate fell, compared with the analogous period in the previous year, in 64 out of 72 months. From 1992 on, the standard deviation of the monthly inflation rate in the previous 12 months was quite stable, varying between 1 and 1.5%. With the exception of 1994, the inflation rate was predictably high in January, largely due to increases in regulated prices. Hence, at least in the later years, it was relatively easy to forecast.[14]

The similarity of these results with those of Lach and Tsiddon (1992) establishes the main result of the paper. There is nothing special about the behaviour of price setters in a transition economy. Despite the lack of previous experience with market institutions, it takes sellers of sausage, eggs, toothpaste, vacuum cleaners, car-wash operators etc. just a few years to figure out how to adjust prices in a market environment.

This finding leads us to believe that we can treat the data as we would have treated data coming from an established market economy and analyse further the nature of price setting policies. In the remainder of the paper we concentrate on various implications of the menu costs

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[13] In our data it is negative and significant in one case.

[14] Note, however, that these arguments apply only to CPI inflation while evidence discussed above is based on own inflation rates.

model.

Both the menu cost as well as the incomplete information approaches imply that the relationship between inflation and relative price variability is symmetric around zero. This is evident in our data. Figure 2 shows the scatter plot of own inflation and the *SDP* measure of variability. As the rate of inflation increases above zero or decreases below zero, *SDP* rises.

This suggests that equations (4) and (5) are misspecified. Therefore we replace inflation with its absolute value in regression (4).[15] The average coefficient on (absolute) inflation increases by about 15%. When we replace the expected and unexpected components of inflation with their absolute values in regression (5) the coefficients on absolute expected inflation becomes eight times larger than the coefficient on absolute unexpected inflation.[16]

The effect of CPI inflation, summarized in columns 7-9 of table 3, presents a similar picture to those obtained using own inflation, but their explanatory power is lower: while some coefficients are larger, so are standard errors and so results are significant only for about 1/4 of the goods. This is not surprising, as there were large relative price changes during transition from the artificial relative price structure imposed by planners to relative prices dictated by the market.

Finally, in Table 5 we summarize results of regressions in which the left hand side

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[15] This regression may still be misspecified if either the effects of negative and positive inflation are not symmetric or the relationship is nonlinear and negative inflation is less frequent in the data (see Bomberger (1999)). To check this we ran regressions allowing the coefficients on positive and negative inflations to differ and tested the null that they add up to zero. We rejected the null in 17 (25) cases when the independent was *INF* (*INFE* and *INFU*). To be consistent across goods and to save space we report only the results of regressions with absolute values of inflation.

[16] Among all 52 goods the coefficient on expected inflation is larger for 43 goods, significantly so for 16 goods. Among the 37 foodstuffs the numbers are 31 and 13. The difference is never negative and significant.

variable is the coefficient of variation of price levels, *CV*. The results are similar to those obtained for the *SDP* measure of variability, but less significant. As argued above, this is consistent with an environment in which menu costs matter.

Overall, our findings are consistent with the menu cost explanation of the relationship between inflation and relative price variability.

V. Are Price Setters Forward Looking?

While the results in the previous section are consistent with the menu cost model, the modelling of expectations is unsatisfactory. The model assumes that agents are forward-looking: the optimal price bounds are set on the basis of the expected inflation until the next price change. The measure of inflation used in section IV, however, is simply a weighted sum of the inflation rates in the previous three months. Hence it is not clear whether price setters' expectations are, indeed, forward-looking.

Modelling expectations in the transition economy is, as argued above, extremely difficult. In any case, our data are insufficient for this as, apart from prices, we do not have any other information on individual markets. The behaviour of prices differs across goods and so relying only on aggregate variables would not be adequate.

Given the data limitations we tried a different approach: regressing the dispersion on the realized past, current and future inflation. This method has two problems. First, all these variables are correlated with the expected future inflation and so, most probably, are correlated with dispersion if the expected inflation is. Second, these three variables are highly correlated and hence the significance tests have low power.

When we add in regression (4) the values of previous month and next month inflation rates as explanatory variables, the results indicate that the most important is current inflation, followed by past inflation and future inflation. The average coefficient on future inflation is much smaller than on the other two inflation rates. This suggests backward-looking expectations. It is important to note, however, these regressions use prices recorded in the first observation each month; for most goods, these observations are made in the first 10 days of the month. While the price setters may have a good idea of the previous month's inflation rate for the product they sell,[17] neither the value of current own inflation nor, of course, next month's inflation is known in that period.

To explore further the hypothesis that the price setters make decisions based on expected future inflation, we propose a second approach which uses the high frequency data. From 1991 to 1996, for 40 foodstuffs, price observations in each store are made at regular intervals three times a month. These include goods 1-37 plus three additional foodstuffs (salt, luncheon meat and condensed milk for which data for 1990 are missing). The idea is as follows. If expectations are backward looking, as time progresses within a month and new information is acquired, the effect future inflation should not change. To test this hypothesis we regress the *SDP* measure of relative price variability, separately for each 10-day period, on last month's, current and next month's inflation:

$$(5) \quad SDP_{it}^{k} = \gamma_{i0} + \gamma_{i1}INF_{i(t-1)} + \gamma_{i2}INF_{it} + \gamma_{i3}INF_{i(t+1)} + \gamma_{i4}t + \gamma_{i5}t^{2} + \gamma d_{MD} + v_{it} \qquad k=1,2,3$$

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[17] The value of inflation for individual goods is published by GUS in April or May of the following year.

Here SPD_{it}^{k} is the standard deviation of $DP_{ijt}^{k} \equiv \ln P_{ijt}^{k} - \ln P_{ijt-1}^{k}$, where P_{ijt}^{k} is the kth

observation in month t of the price of good i in store j. $INF_{i(t-1)}$, INF_{it} and $INF_{i(t+1)}$ are the rates of

own inflation of good i at time t-1, t and t+1, respectively.

Results summarized in Table 6 present a consistent picture. Within each month, as time

progresses, the explanatory power of next month's inflation increases, at the expense of

explanatory power of both the current and previous month's inflation. When SDP is computed

using data from the first ten days of each month, the average coefficient on next month's

inflation, $INF_{i(t+1)}$, is a half of the coefficient on past month's inflation, $INF_{i(t-1)}$. When the data

are from the middle ten days of each month it becomes four times larger. When the data are from

the last ten days of each month the coefficient on the next month's inflation is positive and

significant for almost 40% of the goods in the sample; it is never negative and significant. In

contrast, the explanatory power of past month's inflation falls as time progresses within each

month. In regressions in which the dependent variable is computed using data from the last 10

days in each month the average coefficient is actually negative. Overall, the results suggest that

price setters have forward-looking expectations.

VI. Do Price Setters Follow Time- or State-Contingent Pricing Policies?

The nature of price setting policies is of crucial importance in macroeconomics. If price

setters change prices on a time-contingent basis (for example once a quarter), monetary shocks

have large effects on output as prices are fixed in the short run. These effects are persistent when

individual decisions are staggered (Fischer, 1977 and Taylor, 1980). On the other hand, if pricing

policies are state-contingent, monetary policy is less effective (Caplin and Spulber, 1987, provide an example of an economy in which most individual nominal prices are fixed but monetary policy is ineffective; see, however, Caplin and Leahy, 1992).

It is clear from table 2 that pricing policies are, at least in part, state-contingent: as the rate of inflation falls, price changes become less frequent. Verifying that policies are time-contingent is, however, difficult. One approach would be to estimate a probit model. Our data are insufficient for this task as we do not have any other market specific information and some observations are missing. This leads us to look at temporal clustering of price changes.[18] In the absence of priors, we restrict our attention to checking for temporal clustering at regular time intervals: the beginning of year, quarter, month, or in specific months.

There is little tendency for price changes to cluster in particular months. The proportion of price changes varies between 15.8% in September and 10.2% in June. Excluding June, August and September the proportion varies between 12.9% in January and 11.3% in February.[19] There is also little clustering across months in quarters. The proportion of price changes varies between 12.6% in the third month of each quarter and 12.2% in the first and the second month of each quarter. Price increases, as well as price decreases, also show little tendency to cluster in particular months or quarters.

The high-frequency data (3 observations a month) for 40 foodstuffs over 1991-96 allow

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[18] It is important to note that time-contingent policies need not imply clustering. For example, firms may change prices once a quarter, with a third changing prices each month.

[19] The large value in September is due to the fact that meats, which constitute almost a quarter of the sample, all have the highest proportion of changes in that month (excluding meats, the proportion varies between 13.1% in September and 10.4% in November).

us to take a closer look at clustering within months. Table 7 summarizes the information on the

proportion of price changes by each 10-day period. Price changes are concentrated in the first 10

days of each month: with the exception of price increases in 1991 more than half of all changes

take place in the first 10 days of each month.[20]

What determines the time-contingent behaviour of changing prices at the beginning of the

month? Table 7 indicates that the tendency to cluster price changes is inversely related to the

frequency of price changes. In the early years, when prices are changed often, the proportion of

changes in the first 10 days of each month is relatively low. The proportion rises as price

changes become less frequent. This is supported by cross-sectional evidence. The less frequent

are price changes, the more they cluster at the beginning of the month. Over the whole period, the

good with the highest frequency of price changes (good 9 - eggs) has the lowest proportion of

changes in the first 10 days of each month (45%), and the good with the lowest frequency of

price changes (good 36 - citric acid) has the highest proportion of changes in the first 10 days of

each month (68%). The correlation coefficient across goods between, on one hand, the

proportion of price increases (among nonmissing observations) each month and, on the other

hand, the proportion of all increases which take place in the first 10 days of each month, is -0.79;

the corresponding number for decreases is -0.64.

These findings suggest that firms follow a mixture of state- and time-contingent pricing

policies. They prefer to change prices at the beginning of each month. This tendency manifests

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[20]   It is possible these changes take place at the beginning of each month (first working day), but we have no data to
support, or disprove, this hypothesis.

itself only when price changes are relatively rare, and so the change in timing required to adjust at the beginning of the month, rather than at the optimally chosen time, is relatively insignificant. When price changes are frequent, the change in timing from optimal to the beginning-of-the-month strategy matters and stores are more likely to change prices during the month. Our findings are consistent with the popular assumption in macroeconomics that, when inflation is low (and price changes infrequent) firms tend to follow simple, time-contingent policies. As the rate of inflation rises and the frequency of price changes increases, they switch to state-contingent policies.

VII. Conclusions.

In this paper we analyse the relationship between inflation and relative price variability using a disaggregated data set for Poland. The period covered starts at the beginning of the big-bang transition of the Polish economy from planning to market. Even though the economy is undergoing changes unlike anything encountered in earlier studies of the relationship, the results are remarkably similar to those obtained by Lach and Tsiddon (1992). We also find that behaviour of price setters is broadly consistent with the predictions of the menu cost model.

Most importantly, in our view, the results indicate an astonishing degree of rationality among price setters. Together with the evidence in the companion paper (Konieczny and Skrzypacz, 2000) where we find that agents learn rapidly, search for the best price and arbitrage price differences between markets, the results suggest that the learning curve is really steep.

These findings have clear implications for policy making in transition economies and, similarly, in developing countries. Despite the lack of experience with market-driven allocations,

at the individual level agents behave precisely as theory, and evidence from advanced market economies, suggest. This means that policymakers should not be afraid that households will not respond correctly to market incentives and should concentrate on reforming institutions.

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Table 1

Changes in Store Ownership

Year	Total	State Owned	Cooperative
1990	237425	14312	68454
1991	310966	9440	51044
1992	352502	9613	42448
1993	380582	8620	36187
1994	415449	7533	32369
1995	425600	6287	29372
1996	405563	5399	26316

Sources:
Rynek wewnetrzny w 1993 r. (Domestic Market in 1993), GUS, Warsaw, 1994
Notatka informacyjna dotyczaca publikacji "Rynek Wewnetrzny w 1996r."
  (Information Note on the Publication "Domestic Market in 1996"),
  GUS, Warsaw, July 1997

Table 2

**The Duration of Prices
and the Size of Price Changes**

Year	Inflation rate in % per year	Average probability of price change per month	Average price increase in %	Average price decrease in %
1990-96	54.23	0.37	14.42	-9.45
1990	249.3	0.62	32.92	-12.39
1991	60.4	0.44	15.42	-9.94
1992	44.3	0.39	13.19	-8.95
1993	37.6	0.34	11.09	-7.28
1994	29.5	0.32	9.35	-8.24
1995	21.6	0.31	9.87	-7.54
1996	18.5	0.29	9.05	-9.42

Source: GUS

Table 3

Effect of inflation on price variability
(measured by coefficient SDP)

Column		1	2	3	4	5	6	7	8	9
Regression Specification		INF t	|INF t|	INFE t	INFU t	|INFE t|	|INFU t|	CPI t	CPIE t	CPIU t
average R²		0.57	0.59	0.60		0.60		0.47	0.49	
max R²		0.80	0.80	0.83		0.83		0.78	0.78	
min R²		0.25	0.28	0.25		0.20		0.17	0.18	
Measure of inflation	coeff.	0.428	0.483	0.659	0.322	0.784	0.098	0.318	0.954	0.255
	std	0.147	0.155	0.262	0.160	0.274	0.299	0.292	0.802	0.287
	max	1.40	1.43	3.04	1.28	4.17	0.95	2.95	7.56	2.50
	min	-0.04	-0.12	-0.86	-0.33	-1.05	-1.08	-0.54	-1.94	-0.63
# of significant coeff. (5% level)	ve+	38	40	37	31	41	13	12	15	11
	ve-	0	0	1	0	1	3	1	0	1
time	coeff.	-0.217	-0.204	-0.201		-0.201		-0.209	-0.154	
	std	0.067	0.068	0.074		0.074		0.074	0.103	
	max	0.04	0.02	0.14		0.16		0.05	0.12	
	min	-1.03	-0.95	-0.68		-0.86		-1.18	-0.83	
# of significant coeff. (5% level)	ve+	0	0	0		0		0	0	
	ve-	37	37	36		31		33	16	
time²	coeff.	0.0016	0.0015	0.0016		0.0016		0.0016	0.0013	
	std	0.0007	0.0007	0.0008		0.0008		0.0008	0.0009	
	max	0.0093	0.0084	0.0064		0.0085		0.0112	0.0088	
	min	-0.0009	-0.0006	-0.0020		-0.0017		-0.0008	-0.0013	
# of significant coeff. (5% level)	ve+	27	26	29		26		22	13	
	ve-	0	0	2		1		0	0	

Table 4

**Effect of inflation on price variability
(measured by coefficient SDP)
Results for individual goods**

	Column	1	3	4	5	6
	Regression Specification					
Good		**INF t**	**INFE t**	**INFU t**	**|INFE t|**	**|INFU t|**
1	Back bacon "Sopocka"	0.26 **	0.45 **	0.23 **	0.60 **	0.23 **
2	Sausage "Krakowska sucha"	0.21 **	0.65 **	0.08	1.00 **	0.23
3	Sausage "Mysliwska sucha"	0.05	1.07 **	-0.17	1.26 **	-0.03
4	Sausage "Krakowska parzona"	0.30 **	0.67 **	0.18 **	0.89 **	0.07
5	Sausage "Zwyczajna"	0.26 **	0.54 **	0.18 **	0.77 **	-0.11
6	Pork wieners	0.21 **	0.98 **	-0.02	1.23 **	-0.11
7	Sausage "Torunska"	0.19 **	0.51 **	0.09	1.00 **	0.18 *
8	Sausage "Zywiecka"	0.30 **	0.60 **	0.22 **	0.68 **	0.22 *
9	Eggs, fresh	-0.04	-0.07	-0.04	0.26 **	0.05
10	Herring, salted	0.14	0.33 **	-0.03	0.41 **	-0.79 **
11	Sprats, smoked	0.09	-0.14	0.23 **	-0.17	-0.22
12	Powdered baby milk	0.50 **	0.64 **	0.39 **	0.53 **	0.40 **
13	Cheese "Gouda"	0.42 **	0.44 **	0.37 **	0.47 **	0.13
14	Cheese "Edamski"	0.41 **	0.41 **	0.39 **	0.36 **	0.46
15	Butter, 82.5% fat	0.11 **	0.15 **	0.07	0.15 **	0.25 **
16	Flour "Tortowa"	0.52 **	0.74 **	0.34 **	0.72 **	0.42 **
17	Flour "Krupczatka"	0.59 **	0.75 **	0.38 **	0.68 **	0.59 **
18	Flour "Poznanska"	0.36 **	0.66 **	0.20 **	0.66 **	0.24 **
19	Pearl barley "Mazurska"	0.65 **	0.78 **	0.37 **	0.87 **	0.19
20	Rye bread	0.70 **	1.12 **	0.53 **	0.99 **	0.33
21	Bread "Baltonowski"	0.35 **	0.49 **	0.27 **	0.44 **	0.26 **
22	Bread "Wiejski"	0.36 **	0.51 **	0.27 **	0.47 **	0.12
23	Sugar	0.12 **	-0.41	0.13 **	0.29	0.26 **
24	Plum butter	1.06 **	1.49 **	0.69 **	1.53 **	0.95 **
25	Jam, blackcurrant	1.40 **	1.54 **	1.28 **	1.62 **	-0.33
26	Apple juice	0.27	-0.21	1.08 **	-0.23	0.03
27	Pickled cucumbers	0.31 **	0.36 **	0.12	0.37 **	0.29
28	Margarine "Palma"	0.57 **	0.74 **	0.45 **	0.70 **	0.39
29	Vegetable spread	0.51 **	0.38 **	0.59 **	0.43 **	0.77 **
30	Candy "Krowka"	0.67 **	0.92 **	0.45 **	0.95 **	-0.42
31	Cookies "Delicje szampanskie"	0.41 **	1.83 **	0.27 **	1.88 **	0.10
32	Cookies "Petit Beurre"	0.30	0.97 *	0.15	1.12 **	0.07
33	Pretzel sticks	0.75 **	0.37	0.85 **	0.35	0.86 **

Note: "*" denotes significance at 10% level and "**" at 5%

Table 4
continued

	Column	1	3	4	5	6
	Regression Specification	**INF t**	**INFE t**	**INFU t**	**|INFE t|**	**|INFU t|**
Good						
34 Halvah		0.05	0.05	0.04	0.06	0.30
35 Vinegar, 10%		0.50 **	0.81 **	0.26 *	1.05 **	-0.62 **
36 Citric acid		-0.04	-0.86 **	0.09	0.24	-0.17
37 Tea "Madras"		0.58 **	0.92 **	0.09	0.98 **	-0.30
38 Mineral water in a cafeteria		0.59 **	0.88 *	0.54	0.94 **	-0.67
39 Pastry "W-Z" in a café		-0.02	-0.32	0.02	-0.11	-0.27
40 Razor blade "Polsilver"		0.86 **	2.89 **	0.56 *	4.18 **	-0.61
41 Vacuum cleaner, type 338		0.51 **	0.60 **	0.42 **	0.94 **	0.41 *
42 Kitchen mixer, type 175,5		0.57 **	0.93 **	0.38 **	0.98 **	0.64 **
43 Folding bicycle "Wigry-3"		0.60 **	0.61 **	0.59 **	0.63 **	0.37
44 Radio receiver "Ania"		0.68 **	0.94 **	0.54 **	0.83 **	0.76 **
45 Paint thinner		0.61 **	2.09 **	0.06	2.06 **	-0.39
46 Toothpaste "Pollena"		0.68 *	1.71 **	-0.33	1.81 **	0.18
47 Shaving cream		0.87 **	3.04 **	0.55	2.85 **	-1.08 **
48 Sanitary pads "Donna"		0.78 *	0.70	0.88 **	1.50 **	-0.07
49 Car-wash, of car: "FSO 1500"		0.11	-0.60	0.21	-0.54	0.08
50 Varnishing of hardwood floor		0.02	-0.50	0.17	-1.05 **	-0.04
51 ECG test		0.42 *	-0.69	0.55 **	-0.60	0.32
52 Radiator coolant "Borygo" or "Petrygo"		0.57 **	0.79 *	0.53 **	0.77	0.21

Table 5

Effect of inflation on price variability
(measured by coefficient of variation CV)

Column	1	2	3	4	5	6	7	8	9
Regression Specification	INF t	|INF t|	INFE t	INFU t	|INFE t|	|INFU t|	CPI t	CPIE t	CPIU t
average R²	0.43	0.44	0.44		0.44		0.37	0.38	
max R²	0.76	0.80	0.76		0.79		0.75	0.75	
min R²	0.05	0.06	0.08		0.07		0.02	0.02	
INF — coeff.	0.230	0.266	0.404	0.165	0.403	0.187	0.098	0.082	0.099
INF — std	0.170	0.182	0.351	0.160	0.335	0.297	0.275	0.701	0.241
INF — max	1.52	1.47	7.20	0.82	6.17	3.33	2.29	4.46	2.43
INF — min	-0.54	-0.54	-2.32	-0.97	-2.42	-1.73	-1.00	-2.14	-1.12
# of significant coeff. (5% level) — ve+	19	24	21	19	22	12	6	2	6
ve-	1	0	2	0	2	2	3	1	3
time — coeff.	-0.082	-0.077	-0.101		-0.099		-0.095	-0.096	
time — std	0.111	0.113	0.122		0.112		0.122	0.136	
time — max	0.90	0.93	0.86		0.88		0.94	0.88	
time — min	-1.09	-1.07	-2.31		-2.17		-1.13	-1.10	
# of significant coeff. (5% level) — ve+	7	7	7		6		8	7	
ve-	12	11	12		14		13	11	
time² — coeff.	0.0010	0.0009	0.0012	0.0012	0.0012	0.0012	0.0011	0.0011	
time² — std	0.0013	0.0013	0.0014	0.0014	0.0013	0.0013	0.0014	0.0014	
time² — max	0.0164	0.0162	0.0301	0.0301	0.0285	0.0285	0.0166	0.0144	
time² — min	-0.0107	-0.0110	-0.0104	-0.0104	-0.0106	-0.0106	-0.0110	-0.0106	
# of significant coeff. (5% level) — ve+	14	15	14	14	14	14	13	13	
ve-	7	6	7	7	6	6	7	6	

Table 6

Effect of inflation on price variability
(measured by SDP) 10-DAY DATA

Column		1	2	3	4	5	6	7	8	9
		First 10 days			Second ten days			Third ten days		
Regression Specification		INF t-1	INF t	INF t+1	INF t-1	INF t	INF t+1	INF t-1	INF t	INF t+1
average R^2			0.58			0.49			0.48	
max R^2			0.89			0.87			0.91	
min R^2			0.32			0.20			0.12	
INF	coeff.	0.026	0.379	0.013	0.013	0.229	0.058	-0.038	0.228	0.097
	std	0.108	0.132	0.120	0.089	0.108	0.100	0.088	0.114	0.104
	max	0.33	1.13	0.44	0.37	0.54	0.58	0.23	0.78	0.61
	min	-0.54	-0.28	-0.52	-0.33	-0.53	-0.28	-0.30	-0.16	-0.36
# of significant	ve+	7	31	3	5	28	4	2	21	15
coeff. (5% level)	ve-	4	0	4	4	0	1	6	0	0
time	coeff.		-0.105			-0.047			-0.038	
	std		0.047			0.032			0.035	
	max		0.09			0.07			0.05	
	min		-0.32			-0.15			-0.15	
# of significant	ve+		0			1			0	
coeff. (5% level)	ve-		25			16			10	
$time^2$	coeff.		0.0009			0.0004			0.0003	
	std		0.0006			0.0004			0.0004	
	max		0.0031			0.0015			0.0016	
	min		-0.0016			-0.0010			-0.0010	
# of significant	ve+		18			12			6	
coeff. (5% level)	ve-		0			1			1	

Table 7

**Proportion of Price Changes
by 10-day Periods**

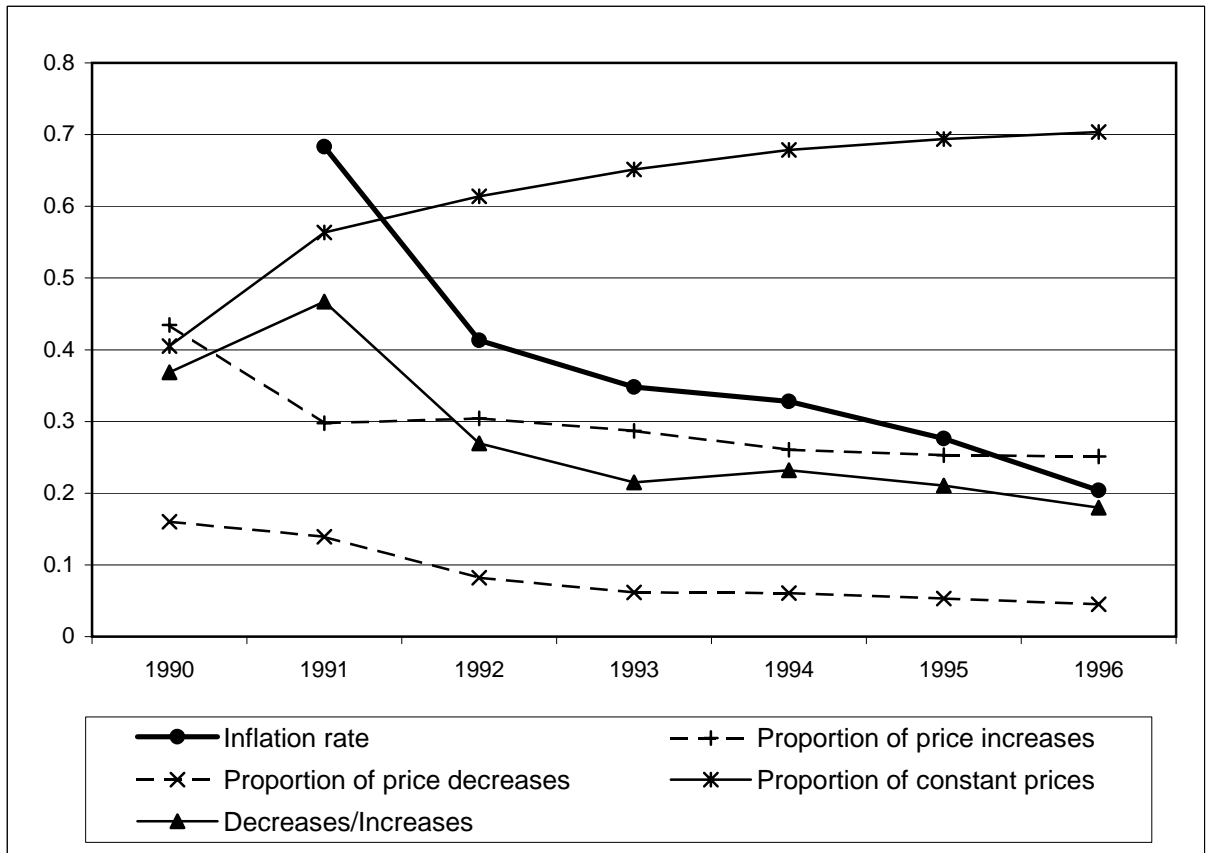| Year	Observation in each month	Proportion of all	
		increases	decreases
		during the month	
1991	1	0.49	0.59
	2	0.29	0.20
	3	0.23	0.21
1992	1	0.56	0.64
	2	0.24	0.18
	3	0.20	0.18
1993	1	0.52	0.60
	2	0.24	0.20
	3	0.24	0.20
1994	1	0.59	0.63
	2	0.20	0.18
	3	0.21	0.18
1995	1	0.61	0.66
	2	0.21	0.18
	3	0.18	0.17
1996	1	0.58	0.61
	2	0.23	0.20
	3	0.19	0.19

Figure 1

General Characteristics of the Data

Figure 2

Inflation and the SDP Measure of Relative Price Variability