Inflation and Price Setting in a Natural Experiment

Jerzy D. Konieczny
Department of Economics
Wilfrid Laurier University
Waterloo, Ont., Canada, N2L 3C5
jkoniecz@wlu.ca

Andrzej Skrzypacz
Graduate School of Business
Stanford University
Stanford, CA 94305-5015
skrzypacz_andrzej@gsb.stanford.edu

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Abstract

We analyze the behavior of price setters in Poland during the transition from a planned to a market economy, using a large disaggregated data set. The size and frequency of price changes, as well as relative price variability, all increase as inflation rises. The effect of expected inflation on relative price variability is much stronger than the effect of unexpected inflation. Price setters are forward looking and follow a mixture of state- and time-contingent policies. Price changes of heterogeneous goods tend to be staggered. These results are broadly consistent with the menu cost model.

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1. Introduction.

We analyze the behavior of price setters using a unique disaggregated data set from Poland. During the period under analysis the Polish economy underwent dramatic changes. The data start in January 1990, when a big-bang transition to a market economy began. 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]

This paper focuses on the behavior of price setters: store owners and store managers. In the 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 started, they suddenly gained the ability to set prices, choose suppliers and even become owners. Shortages rapidly disappeared and managers were often faced with a competitive market.

This environment created a natural experiment, which allows us to address 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? Is it possible to pin down their expectations?

To answer these questions we analyze the effect of inflation on the size and frequency of price adjustment, on relative price variability and on the time pattern of price changes.

Our results are summarized as follows:

1.The size and frequency of price changes, as well as relative price variability, are positively

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

correlated with the inflation rate.

2. Expected inflation has a greater effect on relative price variability than unexpected inflation.
3. 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 measured as the coefficient of variation of price levels.
4. Price-setting policies are forward-looking.
5. Firms follow a mixture of state- and time-contingent pricing policies.
6. Price changes of heterogeneous goods tend to be staggered.

The first two results are, qualitatively, the same as in Lach and Tsiddon (1992), who analyzed a similar data set from Israel. Their data cover the period of 1978-1984, which follows a decade of very high inflation. 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 the high inflation rate. We treat Lach and Tsiddon's (1992) results as a benchmark and so the bottom line is this: despite the lack of previous experience with market institutions, sellers of sausage, eggs, toothpaste, vacuum cleaners, car-wash services, etc. rapidly learn how to form inflationary expectations and adjust prices in a market environment.

Lach and Tsiddon (1992) point out that the effect of expected inflation (see result 2 above) implies that the relationship between inflation and relative price variability is due to menu costs, rather than Lucas' type aggregate-local confusion. In the presence of menu costs, relative price variability is affected by the expected inflation rate, while the latter implies that the unexpected inflation matters.

Given the qualitative similarity of our results, we devote the rest of the paper to the analysis of this and other predictions of the menu cost model. Quantitatively, result 2 is stronger: the coefficient on expected inflation is twice as large as the coefficient on unexpected inflation in the Polish data and only 20% larger in the Israeli data. Furthermore, both menu costs and Lucas' type models imply that the relationship is symmetric around zero. When we regress relative price variability on the absolute values of expected and unexpected inflation, the coefficient on the absolute value of the expected rate of inflation is eight times larger. As explained below, result 3 also indicates that it is the menu costs that matter.

We also find that pricing decisions are not simply adaptive – price variability depends on expected future inflation, not only on its past realizations. We are able to show that pricing policies are a mixture of time- and state-contingent policies. The state-contingent component is more important when price changes are frequent and departures from the optimal timing of changes are costly. Finally, we look at timing of price changes and find that, in groups consisting of heterogeneous goods, price changes tend to be staggered.

The plan of the paper is as follows. The data are described in section 2. Descriptive analysis of the data is in section 3. 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. In section 7 we analyze the staggering of price changes across goods. The last section concludes.

2. 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 seven-year period from January1990 to December 1996. They 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). From this data we selected all goods which were homogeneous across locations and across time (excluding, for example, "a man's suit"). We required that prices were not regulated and eliminated goods with many missing observations. In the end, the sample consists of 52 goods, in 47 stores each, of which 37 are groceries (19 perishable and 18 storable), 2 are sold in cafeterias/cafes, 10 are non-grocery items and 3 are services. The goods are listed in Table 4. The 47 stores 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 non-existent over the study period.

The frequency of observations varies over time and across goods. Each month, there are between two and four observations per store. Whenever a store is sampled more than once a month, the month is divided into equal intervals and there is one observation per interval.

One problem with the data is that the unique identities of the stores where prices were being sampled are unknown. Price inspectors were expected to visit the same store each time, but this was not enforced and deviations were 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. Information on changes in store ownership in the country as a whole is in Table 1. It probably understates ownership changes, since many privatized stores were small and were less likely to be visited by price inspectors than state or co-operative stores. Also, a store may have altered its pricing strategy due to, management change, for example. Finally, while we took great care to select goods with unchanged characteristics 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 a car wash (good number 49) represents the price for washing a specific make and type of car, but, during 1990-96, this specific car became much less common, and car washes appeared offering different levels of service.

The data set is not complete. The proportion of missing prices decreases over time, from 58% in 1990 to 23% 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.

3. Preliminary Analysis.

We begin the analysis by taking a cursory look at the annual data. 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.[2] The annual inflation rate shown in Figure 1 is the rate of price increase over 12 months since the previous 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

[2] We cannot calculate the duration of prices, as some price observations are missing. When gaps in data are frequent, we are less likely to observe long periods with unchanged prices. The proportion of missing data is positively correlated with inflation. This creates a spurious inverse relationship between duration and inflation. To avoid this problem, we calculate the probability of price change. It is obtained as the ratio of the number of price changes to the number of two consecutive observations. In a complete data set this probability would be monotonically related to the duration of prices.

12-month inflation rate fell systematically over time.[3] Monthly inflation was, however, much more erratic and increases in the monthly inflation rate were frequent.

All series follow a similar, nearly monotonic, pattern. As the inflation rate declines over time, both price increases and price decreases become less frequent; also, price decreases become relatively less common than price increases, except for 1991. The same pattern is observed at the level of individual goods, but with some exceptions. The monthly probability of price change is the highest for perishable foodstuffs, somewhat lower for durable foodstuffs and much lower for manufactured goods and services. Between 1990 and 1996 it declined the most for services and the least in durable foodstuffs. The extreme values are 0.94 for eggs in 1990 and 0.06 for an electrocardiogram (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. For individual goods the average size of changes varies from 7.1% for one of the meat products to 36% for the ECG 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 a little in the last three years. The size of price decreases varies less over time than the size of price increases. While in 1990 the average decrease is equal to about 40% of the average increase, in the last three years they are of similar size. Decreases are the smallest for foodstuffs and the largest for

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[3] The general picture of a near monotonic decline in 12-month inflation rate does not depend on the choice of the month. Throughout the sample period it increases only in March 1993, Oct-Nov 1994 and Feb-May 1995.

services and manufactured products.

The recorded proportion of price decreases varies between 27% in 1990 and 15% in 1996 of all price changes[4]. In principle, the large proportion of price decreases may be the result of changes in sampled stores. It is not likely that this factor dominates pricing behavior in our data. If price changes were due to random 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 approaches long-run values by January 1991, while the average size of price increases falls continuously until 1995, and the average size of price decreases falls continuously until 1993.

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.

4. Inflation and Relative Price Variability.

[4] These numbers appear to be large, judging by earlier results. Cross-country evidence suggests that price decreases become more frequent as the rate of inflation rises. The rationale 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 12% 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 from this pattern is Dahlby (1992); in his data inflation is 8-12% per year and there are 7% price cuts.

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 principal explanations of this relationship are based on the menu cost and on incomplete information approaches. When price changing is costly, inflation affects the size and 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) implies 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 (Hercovitz, 1981). Lach and Tsiddon (1992), henceforth L-T, point out that the menu cost approach implies that relative price variability is affected by expected inflation while the incomplete information approach implies a relationship with unexpected inflation. They analyze a disaggregated data set on prices of foodstuffs in Israel during 1978-1984 and find that the effect of expected inflation on relative price variability is stronger than the effect of unexpected inflation.

L-T study the Israeli economy after it had gone through more than ten years of rapid inflation. It was among the most developed countries to experience substantial and prolonged inflation. Israeli price setters can be expected to use optimal, or near optimal, pricing policies. Therefore we treat their results as a benchmark and start our analysis by checking if the same relationship holds in the Polish data. Following L-T, we measure relative price variability by the cross-sectional variance of the rates of price changes. Denote the price of good i in store j at time t by P_{ijt}.[5] 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) $$SDP_{it} = \left[\frac{1}{N_{it} - 1} \sum_j (DP_{ijt} - DP_{it})^2 \right]^{1/2}$$

where N_{it} is the number of observations in which price change could be observed (i.e. the number of two consecutive non-missing observations) and $DP_{it} \equiv (1/N_{it})\, \Sigma_j\, DP_{ijt}$ is the in-sample rate of inflation of good i at time t.

The division of inflation into its expected and unexpected parts 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,[6] its nature was quite different from the subsequent inflationary process. Most prices were regulated and price increases required the approval of planning authorities. In all stores (with the exception of street markets) prices were identical. Inflation in the 1980s was a result of the planners' attempts to reduce rampant shortages. The rate of inflation was determined by the whim of bureaucrats, rather than by observable phenomena like the money supply; even the degree of

[5] In this section we use the first observation each month. This allows us to use the entire data set.

[6] Until 1970 the inflation rate was low, in the range of 0-3% per year. In 1970s 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.

shortages for individual goods did not play a crucial 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 mechanisms or rules-of-thumb developed prior to 1990 were useless for the period following the jump to a market economy. An additional difficulty for the Lucas' approach is that past history of local shocks is an unreliable predictor of the current shock structure.

There are several estimates for the expected CPI rate of inflation in Poland but 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). Expected CPI inflation is obtained 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 economic 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: our measure of inflation differs across goods and, even if we were able to collect additional data to estimate expectations, comparisons across goods would not be straightforward. 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 a theoretical "massaging" of the data. In the end, our approach is similar to that of L-T.[7]

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:

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

where 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 optimal price bounds in the menu cost model (Cecchetti, 1986) and local response in the incomplete information approach. Also, there are large relative price changes in our data as prices adjust to market clearing levels from the artificial price structure inherited from the planned economy. The data used for own inflation are national averages for the given good, rather than in-sample averages.[8] Time and time squared are 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. 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

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[7] L-T construct 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.

[8] In order to compute the rate of inflation for a particular good, GUS first calculates the 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.

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) prevents the results from being dominated by the three outliers. If these months are included, the positive relationship between inflation and relative price variability is much stronger. The elimination of the subsequent months has little effect on the results.

While estimating equation (2) we found that the stochastic components have nonspherical distribution. In particular, we detected significant autocorrelations, correlations across goods and heteroscedasticity.

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* and CV (defined below) varies over time and over goods, and is correlated with the independent variables.

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 standard non-parametric methods described in Newey and West (1994) and Andrews and Monahan (1992)[9].

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 do not 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 our model, (rational) expectations play a crucial role and we do not 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 so they are not reported here.

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:

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

where INF_{it} is own inflation of good i at time t and $\boldsymbol{d_{md}}$ is the vector of monthly dummies. Relative

[9] Following recommendations in this literature based on Monte-Carlo results, we use a quadratic-spectral kernel (Andrews, 1991 and Andrews and Monahan, 1992), pre-whitening procedure (Andrews and Monahan, 1992) and automatic lag selection (Newey and West, 1994).

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 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 3 we summarize the results when own inflation is split into its expected and unexpected parts:

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

These are, essentially, the same regression as in L-T (their equations (2') and (2), respectively).[10] 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 more often positive and significant. 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.[12]

Our results are, qualitatively, identical to those in L-T, despite the fact that our method of calculating expected inflation is, if anything, overly simplified. The similarity of these results with those of L-T suggest that there is nothing special about the behavior of price setters in a transition

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[10] The differences are that we include time and time squared.

[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 equations (3) and (4), is negative for 50 goods in regression (4) and for 52 goods in regression (3).

[12] In one case (good 26 - apple juice) the difference is negative and significant.

economy. Despite the lack of previous experience with market institutions, sellers of sausage, eggs, toothpaste, vacuum cleaners, car-wash services, etc. rapidly learn how to adjust prices in a market environment.

The qualitative similarity of our findings to those in L-T leads us to believe that we can treat the data as we would have treated data coming from an established market economy and analyze further the nature of price setting policies. In the remainder of the paper we concentrate on this and on various other implications of the menu costs model.

One problem with the results in L-T is that they are not very strong: the average coefficient on expected inflation is 20% higher than the coefficient on unexpected inflation. The difference between the effect of expected and unexpected inflation is much larger in the Polish data. All goods in L-T 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 L-T are in brackets. The average coefficient on own inflation in regression (3) is 0.39 (0.41); the average coefficient on expected inflation in regression (4) is 0.57 (0.43); the average coefficient on unexpected inflation is 0.30 (0.36).[13] The coefficient on expected inflation is greater than the coefficient on unexpected inflation for 30 out of 37 goods (16 out of 26). The difference is significant in 9 (5) cases.[14]

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

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[13] The difference between the median coefficients on expected and unexpected inflation in Lach and Tsiddon's data compared to our own 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).

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

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.[15]

Equations (3) and (4) are misspecified, as 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 figure 2, which shows the scatter plot of own inflation and the *SDP* measure of variability. Therefore we replace inflation with its absolute value in regression (3).[16] 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 (4) the coefficients on absolute expected inflation becomes eight times larger than the coefficient on absolute unexpected inflation.[17]

The effect of CPI inflation, summarized in columns 7-9 of Table 3, presents a similar picture to those obtained using own inflation, but the explanatory power is lower: while some coefficients

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

[16] 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.

[17] 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.

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.

An alternative measure of relative price variability is the coefficient of variation of price levels across stores, CV_{ij}:

$$CV_{it} \equiv \left[\frac{1}{N_{it}-1} \sum_j \left(\frac{P_{ijt} - P_{it}}{P_{it}} \right)^2 \right]^{1/2} \tag{5}$$

where $P_{it} \equiv (1/N_{it})\ \Sigma_j\ P_{ijt}$ 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 to distinguish 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.

The choice of the variability measure matters in the menu cost case, a fact which 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 that 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. [18] On the other hand, 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.[19] 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 changes, as well as the observed price levels, differ across firms.

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 a stronger effect 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.

In Table 5 we summarize results of regressions in which the left hand side variable is the coefficient of variation of price levels, *CV*. The results are similar to those obtained for the *SDP* measure of variability, but much less significant. This is consistent with an environment in which menu costs matter.

5. Are Price Setters Forward-looking?

While the results in the previous section are consistent with the menu cost model, the modeling of expectations is unsatisfactory. The model assumes that agents are forward-looking: the

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[18] 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.

[19] 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 as frequent as price changes. In our data all goods meet the requirement, except for good 9 - fresh eggs.

optimal price bounds are set on the basis of the expected inflation until the next price change. The measure of expected inflation used in section 4, 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.

Modeling expectations in the transition economy is, as argued above, extremely difficult. Given the data limitations we tried a different approach: regressing the dispersion on the realized past, current and future own inflation. This method has two problems. First, all these variables are at best noisy proxies of the expected future inflation and so the estimated coefficients are smaller than those we would obtain using true expected inflation. Second, these three variables are highly correlated with each other, hence the significance tests have low power.

When we add in regression (3) 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 ten 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,[20] 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

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

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. Assume expectations are backward looking. Then, as time progresses within a month and new information is acquired, the effect on future inflation should not change. On the other hand, if expectations are forward-looking, new information should be included in expectation formation.

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:

$$(6) \quad SDP_{it}^{k} = \gamma_{i0} + \gamma_{i1} INF_{i(t-1)} + \gamma_{i2} INF_{it} + \gamma_{i3} INF_{i(t+1)} + \gamma_{i4} t + \gamma_{i5} t^{2} + \boldsymbol{\gamma d}_{\boldsymbol{MD}} + \nu_{it} \qquad k=1,2,3$$

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 *k*th 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* in month *t*-1, *t* and *t*+1, respectively.

Results, summarized in Table 6, suggest that expectations are forward-looking. 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 half of the average coefficient on past month's inflation, $INF_{i(t-1)}$; with the data from the middle ten days of each month it is four times larger. When the data are from the last ten days of each month, the average coefficient on last month's inflation is actually negative, while the average 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.

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

Price setting policies are 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 whether 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.[21] In the absence of priors, we restrict our attention to checking for temporal clustering at regular time intervals: the beginning of a 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

[21] 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.

September the proportion varies between 12.9% in January and 11.3% in February.[22] 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 us to take a closer look at clustering within months. Table 7 summarizes the information on the proportion of price changes by each ten day period. Price changes are concentrated in the first ten days of each month: with the exception of price increases in 1991, more than half of all changes take place in the first ten days of each month.[23]

What determines the time-contingent behavior of changing prices at the beginning of the month? Both time series, and cross-sectional evidence indicate that it is affected by the frequency of price changes. As can be seen in Table 7, in the early years, when prices are changed often, the proportion of changes in the first ten 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 price changes are for a given good, 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

[22] 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).

[23] 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. Dutta et al (1999) document that a U.S. grocery chain and a pharmacy chain use weekly pricing rules.

proportion of changes in the first ten 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 ten days of each month (68%). The correlation coefficient across goods between the monthly probability of price increases and the proportion of all increases which take place in the first ten days of the 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 itself only when price changes are infrequent. In that case the change in timing required to adjust at the beginning of the month, rather than at the optimally chosen time, is relatively insignificant (i.e. it has little effect on profits). When price changes are frequent, the change in timing from optimal to the beginning-of-the-month strategy does matter, 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.

7. Synchronization of Price Changes within Markets and Staggering across Markets (?)

The timing of price changes across price setters is important in the menu cost literature, since the only time-invariant distribution of the time of price changes is uniform (Caplin and Spulber, 1987; Benabou, 1988). This is often seen as a condition for consistent aggregation: if the optimal pricing rules of identical firms are based on the expectation of a constant inflation rate, the distribution of price changes has to be uniform for the expectations to be met. On the other hand, if price changes are synchronized, the price level is discontinuous.

The aggregation of individual pricing policies in the presence of menu costs has been analyzed by Rotemberg (1983), Caplin and Spulber (1987), Caballero and Engel (1991), Tsiddon (1991, 1993) and Caplin and Leahy (1991, 1997). Generally speaking, differences across firms and idiosyncratic shocks lead to staggering, while aggregate shocks promote synchronization. Sheshinski and Weiss (1992) analyze the timing of price changes by a multi-product monopolist. They show that positive interactions between prices in the firm's profit function, as well as the costs of adjustment of the "menu" type (i.e. independent of the number of goods changing prices) lead to synchronization of price changes, while negative interactions or the costs of the "management" type (which are increasing in the number of changes) promote staggering.[24]

Empirical studies concentrate on within-store synchronization. Tomassi (1993) finds that price changes of different groceries sold by a given store are staggered. In contrast, Lach and Tsiddon (1996) find that price changes of different groceries within a given store are synchronized, while price changes of a given product in different stores are staggered. Dutta et al (1999) report that a large proportion of price changes takes place at certain times (early in the week for grocery chains and on Fridays for a drugstore chain). Fisher and Konieczny (2000) find that prices of newspapers published by the same owner tend to change together, while prices of independent newspapers do not. Chakrabarti and Scholnick (2001) report within-store synchronization by internet booksellers. Finally, casual observation provides many examples of within-market synchronization (for example

[24] A unique analysis of various aspects of the price changing costs in Zbaracki et al (2000) shows the importance of the "management" type costs.

list car prices or gasoline prices).

A simple way to get an idea of the degree of synchronization of price changes is to look at the standard deviation of the proportion of price changes. We look at the entire sample period, as well as 1992-96. The reason for including the shorter period is that the proportion of price changes falls rapidly over the first three years (see Table 2), increasing the standard deviation independently of the degree of price change synchronization. Theoretically, the standard deviation is in the interval [0, SD_{max}], where the lower bound would obtain under perfect staggering and the upper bound under perfect synchronization.[25] The value of SD_{max} depends on the frequency of price changes in the group. Under perfect synchronization, the proportion of price changes is either one or zero (for example, if prices change in a quarter of the observations, the proportion of price changes is 1 in 25% of the cases and zero in 75% of the cases). The most meaningful statistic is the ratio of the sample standard deviation to SD_{max}. In 1992-96 it varies between 0.17 (good 52: radiator coolant) to 0.48 (good 7: sausage "Torunska"), with a median value of 0.30. For the entire sample period the median value is 0.39; the ratio exceeds 0.5 for only two goods. This indicates that the behavior of price changes is, if anything, closer to perfect staggering or to independence than to perfect synchronization.

We can test the hypothesis that the sample values of the variances are equal to the corresponding values under the assumption of independence and perfect synchronization. The null hypothesis is that the variances are equal and the alternative is a one-sided test (sample variance

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[25] An alternative for the lower bound is when price changes are independent. As mentioned above, perfect staggering is one of the possible implications of Sheshinski and Weiss (1992). All results below are the same whether the lower bound corresponds to perfect staggering or to independence.

larger than under independence, and sample variance smaller than under perfect synchronization). For all groups, and for almost all individual goods, we reject the null at the 5% level using a one-sided chi squared test.

It should be noted, however, that aggregation does not require the staggering of price changes in every market. Assume, for example, that there are m markets with n identical firms in each market[26], which change prices once every T periods. Even with perfect within-market synchronization, the observed price level increases smoothly if the timing of price changes is staggered across markets (i.e. prices change each period in m/T markets), or if the period of observation is a multiple of T.

As we do not know the identity of the stores, our data do not permit us to address the issue of within-store synchronization of price changes. They allow, however, an analysis of synchronization across goods and markets. To do this, we compare the behavior of the proportion of price changes for individual goods and for groups of goods.

In Figure 3 we plot the proportion of price changes for all goods as well as for various subgroups. As before, we compute the probability of price change as the ratio of price changes to the number of two consecutive observations, separately for each good and for each group of goods. It is clear that the degree of staggering increases with the heterogeneity of the group. The proportion varies the most for homogeneous groups (breads, meats and sausages) and the least for heterogeneous groups (all goods, industrial goods).

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[26] More generally, a market may consist of more than one good. If there are k goods in a market, assume there are n/k identical firms in that market for the result to hold.

While the average proportion of price changes varies quite a bit over time, most of the variations are due to two factors: many price changes take place each January, and prices of most meats and sausages change in late summer/early fall, usually in September. The first factor is specific to Poland, where prices of some basic commodities like heating and electricity were regulated and changed on January 1 each year. The second is a seasonal factor; its role in our data is exaggerated as meats and sausages constitute almost 20% of the sample.

Figure 3b illustrates the synchronization within markets and staggering across markets. The proportion of price changes for breads and for meats and sausages varies greatly over time, but the peaks rarely, if ever, coincide. Hence the proportion of price changes for both groups taken together would vary less than for each group separately.

There are two ways to look at the hypothesis more formally. First, we test whether the variance of the proportion of price changes in a given group is higher than the median variance of the proportion for goods in the group. The null hypothesis is that the two variances are equal to each other and the (one-sided) alternative is that the group variance is smaller than the median one. Under the null, the ratio of the two variances (both calculated using the same number of monthly observations, n) has an F distribution with (n-1, n-1) degrees of freedom. The data and the results are in Table 8. In column 4 we report the values of the ratios of the median variance for goods in a given group to total group variance, and indicate which are statistically significantly higher than 1 at a 1% level.[27] We reject the null hypothesis for all heterogeneous groups, but do not for the two homogeneous groups.

[27] Whenever we could not reject the null at the 1% significance level, the test failed to do so also at the 5% level.

Another approach is to look at the behavior of the size of price changes. The more staggered price changes are, the lower is the variance of the average value of the monthly price change. Again, we compare the median variance of price change for goods in a given group with the variance in the group (see Table 8). For 1992-96 the results are the same as before.[28] Overall, our results indicate a greater degree of staggering of price changes in heterogeneous groups of goods than for homogeneous groups or for individual goods.

8. Conclusions.

In this paper we analyze 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 a planned to a market economy. 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 behavior 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 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

[28] For the entire sample period we cannot reject the null hypothesis. The reason is that, at the beginning of transition, price changes are very large, a fact which dominates the variances.

the individual level agents behave precisely as theory, as well as evidence from advanced market economies, suggest. This means that policymakers can introduce market institutions without being afraid that households, who have no experience with market economy, will not respond correctly to market incentives.

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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:
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(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	0.59	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	0.35	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	0.30	9.05	-9.42

Source: GUS

Table 3

Effect of inflation on price variability (measured by 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 R^2		0.57	0.59	0.60		0.60		0.47	0.49	
max R^2		0.80	0.80	0.83		0.83		0.78	0.78	
min R^2		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	ve+	38	40	37	31	41	13	12	15	11
coeff. (5% level)	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	ve+	0	0	0		0		0	0	
coeff. (5% level)	ve-	37	37	36		31		33	16	
$time^2$	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	ve+	27	26	29		26		22	13	
coeff. (5% level)	ve-	0	0	2		1		0	0	

Table 4

Effect of inflation on price variability
(measured by 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	0.26 **	0.45 **	0.23 **	0.60 **	0.23 **
2	0.21 **	0.65 **	0.08	1.00 **	0.23
3	0.05	1.07 **	-0.17	1.26 **	-0.03
4	0.30 **	0.67 **	0.18 **	0.89 **	0.07
5	0.26 **	0.54 **	0.18 **	0.77 **	-0.11
6	0.21 **	0.98 **	-0.02	1.23 **	-0.11
7	0.19 **	0.51 **	0.09	1.00 **	0.18 *
8	0.30 **	0.60 **	0.22 **	0.68 **	0.22 *
9	-0.04	-0.07	-0.04	0.26 **	0.05
10	0.14	0.33 **	-0.03	0.41 **	-0.79 **
11	0.09	-0.14	0.23 **	-0.17	-0.22
12	0.50 **	0.64 **	0.39 **	0.53 **	0.40 **
13	0.42 **	0.44 **	0.37 **	0.47 **	0.13
14	0.41 **	0.41 **	0.39 **	0.36 **	0.46
15	0.11 **	0.15 **	0.07	0.15 **	0.25 **
16	0.52 **	0.74 **	0.34 **	0.72 **	0.42 **
17	0.59 **	0.75 **	0.38 **	0.68 **	0.59 **
18	0.36 **	0.66 **	0.20 **	0.66 **	0.24 **
19	0.65 **	0.78 **	0.37 **	0.87 **	0.19
20	0.70 **	1.12 **	0.53 **	0.99 **	0.33
21	0.35 **	0.49 **	0.27 **	0.44 **	0.26 **
22	0.36 **	0.51 **	0.27 **	0.47 **	0.12
23	0.12 **	-0.41	0.13 **	0.29	0.26 **
24	1.06 **	1.49 **	0.69 **	1.53 **	0.95 **
25	1.40 **	1.54 **	1.28 **	1.62 **	-0.33
26	0.27	-0.21	1.08 **	-0.23	0.03
27	0.31 **	0.36 **	0.12	0.37 **	0.29
28	0.57 **	0.74 **	0.45 **	0.70 **	0.39
29	0.51 **	0.38 **	0.59 **	0.43 **	0.77 **
30	0.67 **	0.92 **	0.45 **	0.95 **	-0.42
31	0.41 **	1.83 **	0.27 **	1.88 **	0.10
32	0.30	0.97 *	0.15	1.12 **	0.07
33	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	0.05	0.05	0.04	0.06	0.30
35	0.50 **	0.81 **	0.26 *	1.05 **	-0.62 **
36	-0.04	-0.86 **	0.09	0.24	-0.17
37	0.58 **	0.92 **	0.09	0.98 **	-0.30
38	0.59 **	0.88 *	0.54	0.94 **	-0.67
39	-0.02	-0.32	0.02	-0.11	-0.27
40	0.86 **	2.89 **	0.56 *	4.18 **	-0.61
41	0.51 **	0.60 **	0.42 **	0.94 **	0.41 *
42	0.57 **	0.93 **	0.38 **	0.98 **	0.64 **
43	0.60 **	0.61 **	0.59 **	0.63 **	0.37
44	0.68 **	0.94 **	0.54 **	0.83 **	0.76 **
45	0.61 **	2.09 **	0.06	2.06 **	-0.39
46	0.68 *	1.71 **	-0.33	1.81 **	0.18
47	0.87 **	3.04 **	0.55	2.85 **	-1.08 **
48	0.78 *	0.70	0.88 **	1.50 **	-0.07
49	0.11	-0.60	0.21	-0.54	0.08
50	0.02	-0.50	0.17	-1.05 **	-0.04
51	0.42 *	-0.69	0.55 **	-0.60	0.32
52	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 R^2		0.43	0.44	0.44		0.44		0.37	0.38	
max R^2		0.76	0.80	0.76		0.79		0.75	0.75	
min R^2		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
	std	0.170	0.182	0.351	0.160	0.335	0.297	0.275	0.701	0.241
	max	1.52	1.47	7.20	0.82	6.17	3.33	2.29	4.46	2.43
	min	-0.54	-0.54	-2.32	-0.97	-2.42	-1.73	-1.00	-2.14	-1.12
# of significant	ve+	19	24	21	19	22	12	6	2	6
coeff. (5% level)	ve-	1	0	2	0	2	2	3	1	3
time	coeff.	-0.082	-0.077	-0.101		-0.099		-0.095	-0.096	
	std	0.111	0.113	0.122		0.112		0.122	0.136	
	max	0.90	0.93	0.86		0.88		0.94	0.88	
	min	-1.09	-1.07	-2.31		-2.17		-1.13	-1.10	
# of significant	ve+	7	7	7		6		8	7	
coeff. (5% level)	ve-	12	11	12		14		13	11	
time^2	coeff.	0.0010	0.0009	0.0012		0.0012		0.0011	0.0011	
	std	0.0013	0.0013	0.0014		0.0013		0.0014	0.0014	
	max	0.0164	0.0162	0.0301		0.0285		0.0166	0.0144	
	min	-0.0107	-0.0110	-0.0104		-0.0106		-0.0110	-0.0106	
# of significant	ve+	14	15	14		14		13	13	
coeff. (5% level)	ve-	7	6	7		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 during the month	Proportion of all 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

Table 8

Proportion of Price Changes	Value of the proportion of changes	Variance in group	Median variance for goods in group	Ratio of median variance to total group variance
1992-96				
All goods	0.33	0.005	0.019	3.64 *
Foodstuffs	0.37	0.007	0.023	3.29 *
Perishable foodstuffs	0.40	0.013	0.034	2.55 *
Meats and sausages	0.39	0.041	0.046	1.13
Breads	0.31	0.036	0.034	0.93
Industrial goods	0.24	0.005	0.016	3.13 *
1990-96				
All goods	0.37	0.016	0.032	2.00 *
Foodstuffs	0.41	0.017	0.036	2.08 *
Perishable foodstuffs	0.45	0.025	0.046	1.84 *
Meats and sausages	0.45	0.053	0.058	1.10
Breads	0.33	0.037	0.038	1.01
Industrial goods	0.28	0.017	0.028	1.65 *

Average Monthly Price Change	Value of the average monthly price change	Variance in group	Median variance for goods in group	Ratio of median variance to total group variance
1992-96				
All goods	2.0%	0.000	0.001	4.25 *
Foodstuffs	2.0%	0.000	0.001	2.90 *
Perishable foodstuffs	1.8%	0.000	0.001	2.45 *
Meats and sausages	1.7%	0.001	0.001	1.03
Breads	2.6%	0.001	0.001	1.10
Industrial goods	2.0%	0.000	0.000	3.73 *
1990-96				
All goods	2.8%	0.003	0.0035	1.37
Foodstuffs	2.7%	0.003	0.0037	1.38
Perishable foodstuffs	2.6%	0.003	0.0038	1.36
Meats and sausages	2.5%	0.003	0.0027	0.98
Breads	3.1%	0.007	0.0070	0.98
Industrial goods	3.3%	0.004	0.0046	1.30

* significant at 99% level

Figure 1

General Characteristics of the Data

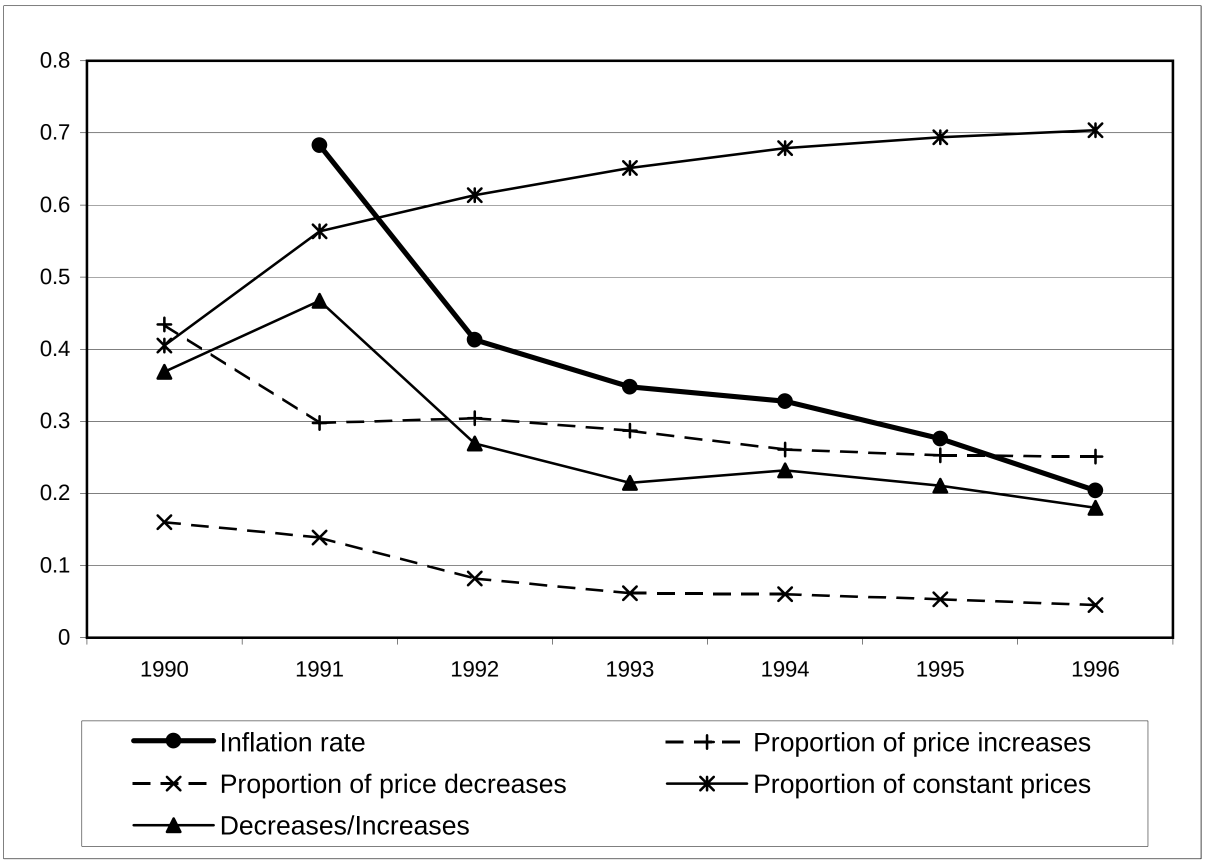

Figure 2

Inflation and the SDP Measure of Relative Price Variability

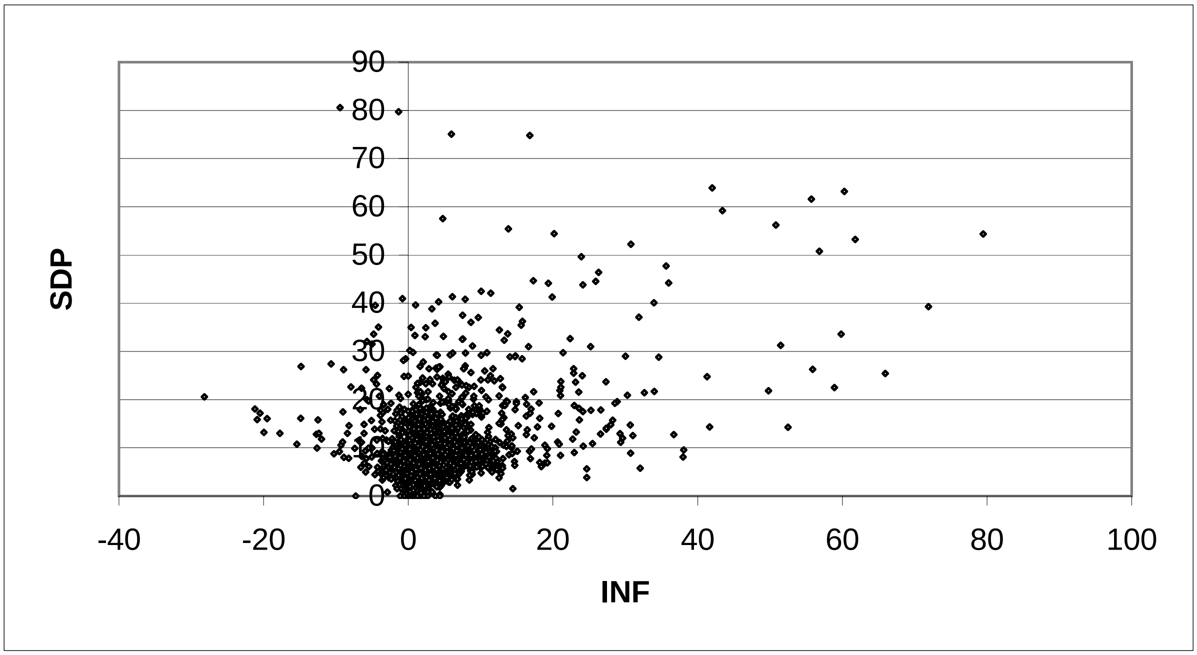

Figure 3

Proportion of price changes, 1992-96