

Inflation and price setting in a natural experiment*

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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. Despite the unusual economic environment, the results are qualitatively identical, and quantitatively stronger, than those in Lach and Tsiddon (1992).

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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 faced with a competitive market.

This new environment created a natural experiment, in which the obvious question is whether pricing policies are similar or different from those observed elsewhere. To answer it we analyze the effect of inflation and its components on relative price variability.

We find that (a) the size and frequency of price changes, as well as relative price variability, are positively correlated with the inflation rate (b) expected inflation has a greater effect on relative price variability than unexpected inflation, and (c) 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 by the coefficient of variation of price levels.

The first two results are, qualitatively, the same as in Lach and Tsiddon (1992), who

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

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.

Quantitatively, result (b) is stronger: the coefficient on expected inflation is twice as large as the coefficient on unexpected inflation in the Polish data while it is 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 (c) also indicates that it is the menu costs that matter.

The plan of the paper is as follows. The data are described in section 2. In section 3 we analyze the relationship between inflation and relative price variability. 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 January 1990 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 47 stores comprise the complete set for 4 out of 49 administrative regions in Poland (Voivodships). The goods are listed in the Appendix.

The prices are the actual transaction prices, as coupons or discounts were very rare or non-existent over the study period. As retail prices, they were not subject to quantity discounts. Temporary sales and promotional packaging (e.g. 120 gram for the price of 100 gram) were uncommon.

One problem with the data is that the unique identities of the stores where prices were being sampled are unknown. Price inspectors were instructed to visit the same store each time, but this was not enforced and deviations were not recorded.

The frequency of observations differed across goods and time, from 1 to 4 each month in each store. To make the analysis comparable across goods we use the first observation each month. The data set is not complete. The proportion of missing prices decreases over time, from 41% in 1990 to 18% in 1995. 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.

In Table 1 we summarize the information on the size and probability of price changes.² The annual inflation rate is the rate of CPI change December to December. It falls systematically over time.³ As the inflation rate declines, both price increases and price decreases become less frequent and smaller. With some exceptions, the same pattern is observed at the level of individual goods.⁴

INSERT TABLE 1 HERE

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. The extreme values are 0.94 for eggs in 1990 and 0.06 for an electrocardiogram (ECG) test in 1996.

Price increases are the smallest for perishable foodstuffs (in particular meats) and the largest for manufactured products. The average size of price changes varies from 7.1% for one of the meat products to 36% for the ECG test.

3. 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; Domberger, 1987; Van Hoomissen, 1988; Lach and Tsiddon, 1992; and Debelle and Lamont,

² We cannot calculate the duration of prices, as some price observations are missing. Instead, 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.

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. Monthly inflation was, however, much more erratic, and increases in the monthly inflation rate were frequent.

⁴ The finding that the probability of price changes falls with inflation is consistent with earlier findings. 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.

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). 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 for relative price variability is the inability of firms to distinguish between aggregate and local shocks. 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 check 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 in month t by P_{ijt} . Whenever we have two consecutive price 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} :

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

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 $DP_{it} \equiv (1/N_{it}) \sum_j DP_{ijt}$ 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,⁵ 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 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

⁵ Until 1970 the inflation rate was low, in the range of 0-3% per year. In the 1970s it varied between 0 and 10%. From 1981 to 1988 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.

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 the 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 atheoretical "massaging" of the data. In the end, our approach is similar to that of L-T.⁶

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

⁶ They 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 own inflation for 80% of goods.

time, time squared and monthly dummies as explanatory variables. Own inflation is superior to 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.⁷ Time and time squared are included as a proxy for structural change; we expect the change to be fast initially and to slow down over time, as the economy approaches the new equilibrium. Monthly dummies are included as we have many seasonal goods.

The inflation rates, 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) 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.

The model is estimated using simple OLS with consistent non-parametric estimation of standard errors.⁸ To estimate standard errors we employ standard non-parametric methods described

⁷ 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.

⁸ We use nonparametric estimates of the standard errors as we have detected both autocorrelation and heteroscedasticity. We also estimated the models as a seemingly unrelated regression system. The results are similar, but as the coefficient estimates may be inconsistent, we report the OLS results only.

in Newey and West (1994) and Andrews and Monahan (1992).⁹

In Table 2 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.¹⁰ In columns 1 and 2 we summarize the results when own inflation is divided into its expected and unexpected parts:

$$SDP_{it} = \beta_{i0} + \beta_{i1}INFE_{it} + \beta_{i2}INFU_{it} + \beta_{i3}t + \beta_{i4}t^2 + \beta d_{md} + e_{it} \quad (2)$$

This is, essentially, the same regression as in L-T (their equation 2).¹¹ 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 coefficient on expected inflation is larger than on unexpected inflation for 40 out of 52 goods; the difference is significant for 13 goods.¹²

INSERT TABLE 2 HERE

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 suggests that there is nothing special about the behavior of price setters in a transition

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).

¹⁰ Results for individual goods are in the working paper version, Konieczny and Skrzypacz (2001).

¹¹ The difference is that we include time and time squared and monthly dummies.

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.

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 in our sample are either industrial products or services. The values from L-T are in brackets. The average coefficient on expected inflation in regression (2) 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 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

¹³ The difference between the median coefficients in our results compared to L-T 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.

increases in regulated prices. Hence, at least in the later years, it was relatively easy to forecast.¹⁵

Equation 2 is 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 1, which shows the scatter plot of own inflation and the *SDP* measure of variability. Therefore we replace the expected and unexpected components of inflation with their absolute values in regression (2). The coefficient on absolute expected inflation becomes eight times larger than the coefficient on absolute unexpected inflation (see columns 3 and 4 of Table 2).¹⁶

INSERT FIGURE 1 HERE

The effect of CPI inflation presents a similar picture to those obtained using own inflation, but the 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.¹⁷ 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} :

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

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.

¹⁷ For brevity, we do not report these results here. They are in the working paper version (Konieczny and Skrzypacz, 2001).

$$CV_{it} \equiv \left[\frac{1}{N_{it} - 1} \sum_j \left(\frac{P_{ijt} - P_{it}}{P_{it}} \right)^2 \right]^{1/2} \quad (3)$$

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 us 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 having the highest real price. 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

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.

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 columns 5-8 of Table 2 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.

4. 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).

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 adapt their behavior 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 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, assured that households which have no experience with a market economy will respond correctly to market incentives.

Appendix.

List of goods:

1. Back bacon "Sopocka", 1kg; 2. Sausage "Krakowska sucha", 1 kg; 3. Sausage "Mysliwska sucha", 1 kg; 4. Sausage "Krakowska parzona", 1kg; 5. Sausage "Zwyczajna", 1kg; 6. Pork wieners, 1 kg; 7. Sausage "Torunska", 1 kg; 8. Sausage "Zywiecka", 1 kg; 9. Eggs, each; 10. Herring, salted, 1 kg; 11. Sprats, smoked, 1 kg; 12. Powdered baby milk, 500g; 13. Cheese "Gouda", 1 kg; 14. Cheese "Edamski", 1 kg; 15. Butter, 82.5% fat, 250g; 16. Flour "Tortowa", 1 kg; 17. Flour "Krupczatka", 1 kg; 18. Flour "Poznanska", 1 kg; 19. Pearl barley "Mazurska", 1 kg; 20. Rye bread, 1 kg; 21. Bread "Baltonowski", 1 kg; 22. Bread "Wiejski", 1 kg; 23. Sugar, 1 kg; 24. Plum butter, 460g jar; 25. Jam, blackcurrant, 460g jar; 26. Apple juice, 1 l box; 27. Pickled cucumbers, 900g jar; 28. Margarine "Palma", 250g; 29. Veggie butter, 250g tub; 30. Candy "Krowka", 1 kg; 31. Cookies "Delicje szampanskie", 1kg; 32. Cookies "Petit Beurre" type, 100g; 33. Pretzel sticks, 100g; 34. Halvah, 1 kg; 35. Vinegar, 10%, 0.5l bottle; 36. Citric acid, 10g bag; 37. Tea "Madras", packed domestically, 100g; 38. Mineral water in a cafeteria, 0.33l bottle; 39. Pastry "W-Z" in a café, each; 40. Razor blade "Polsilver", each; 41. Vacuum cleaner, type 338,5; 42. Kitchen mixer, type 175,5; 43. Folding bicycle "Wigry-3"; 44. Radio receiver "Ania" 45. Paint thinner, 0.5l; 46. Toothpaste "Pollena", 98g; 47. Shaving cream; 48. Sanitary pads "Donna", box of 20; 49. Car-wash, of car: "FSO 1500"; 50. Varnishing of hardwood floor, twice, 1m²; 51. EKG test; 52. Radiator coolant "Borygo" or "Petrygo", 1 l.

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Table 1**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: authors' calculations

Table 2

Effect of inflation on price variability

Dependent Variable		SDP				CV			
Column		1	2	3	4	5	6	7	8
Regression Specification		INFE t	INFU t	INFE t	INFU t	INFE t	INFU t	INFE t	INFU t
average R ²		0.60		0.60		0.44		0.44	
max R ²		0.83		0.83		0.76		0.79	
min R ²		0.25		0.20		0.08		0.07	
Measure of inflation	coeff.	0.659	0.322	0.784	0.098	0.404	0.165	0.403	0.187
	std	0.262	0.160	0.274	0.299	0.351	0.160	0.335	0.297
	max	3.04	1.28	4.17	0.95	7.20	0.82	6.17	3.33
	min	-0.86	-0.33	-1.05	-1.08	-2.32	-0.97	-2.42	-1.73
# of significant coeff. (5% level)	ve+	37	31	41	13	21	19	22	12
	ve-	1	0	1	3	2	0	2	2
time	coeff.	-0.201		-0.201		-0.101		-0.099	
	std	0.074		0.074		0.122		0.112	
	max	0.14		0.16		0.86		0.88	
	min	-0.68		-0.86		-2.31		-2.17	
# of significant coeff. (5% level)	ve+	0		0		7		6	
	ve-	36		31		12		14	
time ²	coeff.	0.0016		0.0016		0.0012		0.0012	
	std	0.0008		0.0008		0.0014		0.0013	
	max	0.0064		0.0085		0.0301		0.0285	
	min	-0.0020		-0.0017		-0.0104		-0.0106	
# of significant coeff. (5% level)	ve+	29		26		14		14	
	ve-	2		1		7		6	

Figure 1**Inflation and the SDP Measure of Relative Price Variability**