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Relative Price Variability and Inflation in Europe

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The relationship between inflation and the relative variability of prices has been the subject of careful investigation in the United States using data for product groups at the city level. Yet in Europe, where the relationship could have profound effects on the viability of monetary integration, no attempt has been made to study the relationship. This paper fills the gap by examining data disaggregated to the commodity level across 10 EU countries. Evidence is found for logistic smooth transitions in the relative price variability measures within countries and within product groups. When this deterministic component is removed, the stochastic element is not persistent and does not always have the positive relationship with inflation commonly found in US city data.

Introduction

Since the collapse of the exchange rate mechanism and the adoption of Maastricht criteria for monetary union, European nations have focused on the control of inflation. While this is 'business-as-usual' for Germany, where the Bundesbank has officially followed a monetary target, motivated by a commitment to inflation control (Bernanke and Milhov 1997), for other countries this represents a new regime for central banks. Typically, the control of inflation concerns the level of inflation, and a great deal of effort has been given to understanding the causes and behaviour of headline and underlying inflation rates. Little attention has been given to the higher moments of the inflation distribution, in particular to the relative variability of prices around the average level of inflation, for different European countries. This is surprising, since the stated aim and objective of inflation control is to create a stable platform on which to build a monetary union with a single currency. The variability of prices across Europe is an indicator of the degree of convergence in price-setting behaviour: the more variable are relative prices, the less convergent is price-setting behaviour in the countries concerned.

The absence of research into relative price variability in European countries contrasts markedly with a plethora of recent work on inflation variability and uncertainty across cities in the United States.[1] These papers investigate the relative variability of individual commodity prices across American cities and the relative variability of different product groups across each city. Our paper fills the gap by collecting data for 10 countries and 15 different product groups to determine the relative variability in price-setting behaviour in Europe. The results provide empirical information on the extent of variation in relative prices, its persistence over time and its relationship to the level of inflation. We study a panel of data to determine the degree of relative price variability for individual product groups across 10 countries and the degree of price variability within each country across a range of 15 product groups. Examining the relative price variability of prices across Europe, we can determine whether price-setting is convergent in two respects.

First, the information provided by the panel enables us to determine the effectiveness of the single market for goods in driving producers to set prices for a common market place, since our sample corresponds very closely to the period of the EU Single Market programme. Evidence from relative price variation in the price of individual product groups will reveal whether the single market has brought about greater price convergence across European countries. Theoretically, the single market should ensure that there is a single price for homogeneous goods in different countries within Europe, but in practice companies may adopt pricing-to-market rules which create differences in prices for the same product groups across countries. Relative price variability should be low and temporary if the producers of goods really are competing for business in a single market that is not partitioned by national boundaries.

Second, the data for each country allow us to investigate whether it is possible to categorize countries by the degree of relative price variability around the weighted average, giving evidence for sub-groups of countries that have common price variability patterns. Price-setting behaviour that varies across countries likely to participate in a single currency will make inflation control harder for a European monetary authority. Substantial differences in relative price variability may also reflect diversity in monetary policy across Europe and indicate the (in)feasibility of introducing a common monetary policy under a single currency. As things stand, the execution of monetary policy at the national level influences inflation, relative price variability and the relationship between them.[2]

If the data reveal significant relative price variability, its permanence will be an important issue. If price variability is eliminated over a short horizon, then it can safely be regarded as a temporary phenomenon, which need not concern monetary policy-makers whose aim is long-run price stability. If the relative price variability is more persistent, then policy-makers may need to take it into consideration when formulating monetary policy.[3] Our results suggest that shocks to relative price variability are temporary, being almost completely eliminated within a calendar year by market forces. Monetary policy can focus on the single objective of long-term price stability.

The paper proceeds as follows. The next section explains the measurement of relative price variability and its relationship to inflation in previous studies. Section II derives the smooth transition methodology that has been adopted to investigate inflation variability and describes the data-set. Section III reports the results for relative price variability across different product groups within countries, while Section IV reports the relative price variability results for individual product groups across countries. Section V draws the policy conclusions.

I. Measuring Relative Price Variability and its Relationship to Inflation

The empirical evidence suggesting that there may be a positive relationship between inflation and inflation variability was brought to prominence by Okun (1971).[4] His analysis made use of the unconditional standard deviation and the mean of the inflation process from a cross-section of 17 OECD countries over the period 1951–68 to show an empirical relationship between the two variables. The first time-series evidence, presented by Vining and Elwertowski (1976) in a graphical form, confirmed this view, using the standard deviation of changes to relative prices and the inflation rate.

The results of both of these papers can now be questioned in the light of more recent statistical knowledge. Logue and Willett (1976) have shown that the cross-sectional relationship between inflation and the variability of inflation breaks down under disaggregation, while the time series relationship has been shown to be sensitive to the sample period chosen, and to particular shocks to food and energy components of the price series, by Fischer (1981) and Driffill *et al*. (1990). Furthermore, R. Engle (1982), observes that it is the *conditional* standard deviation in inflation and not the *unconditional* standard deviation that matters. Through the development of ARCH and subsequently the derivatives based on GARCH, Engle was able to reopen the debate on the relationship between inflation and the conditional standard deviation of inflation. Papers by Driffill *et al*. (1990) and Joyce (1995) demonstrate that there is a strong positive relationship in time series data betwen inflation and the conditional standard deviation of inflation on US and UK data-sets.

Improvements have also been made to the original cross-sectional measures of variability, which tended to overstate the positive relationship with inflation arising from the effects of aggregation. Aggregation hides the true degree of variability in the data and allows common third causes, such as the oil price shocks of 1973 and 1981, to exert considerable leverage in a regression of inflation variability on the inflation level. For the reasons above, the cross-sectional measures of the unconditional standard deviation of prices at the aggregate level have been called into question.

Parks (1978) and Lach and Tsiddon (1992) introduced new measures of relative price variability using price data disaggregated to the level of the individual product groups used to construct the CPI which were immune to these criticisms. Defining p_{it} to be the price of the ith commodity at time t, inflation is measured as

$$(1) \qquad DP_t = \sum_{i=1}^{n} w_{it}^{*} Dp_{it},$$

where w_{it}^{*} is the expenditure share of commodity i and $Dp_{it} = \ln p_{it} - \ln p_{it-1}$. The relative price variability is defined as

$$(2) \qquad RVP_t = \sum_{i=1}^{n} w_{it}^{*} (Dp_{it} - DP_t)^2.$$

By disaggregating prices, these papers record the true degree of variability without obscuring important information through the aggregation process. The measure they use also corrects for the effect of high leverage on shocks to energy and food prices. Their results support the hypothesis that there is a positive relationship between relative price variability and inflation, but the duration of price quotations falls and RPV is less persistent with higher inflation.

More recently, papers by Parsley (1996) and Debelle and Lamont (1997) have sought to provide an additional level of information on price variability and inflation by collecting commodity-level price data over a number of US cities. Parsley's results, based on a regression of the logarithm of relative price variability on the logarithm of inflation across cities and across product groups, find a positive association but little evidence of a long-run relationship. Debelle and Lamont regress city-level relative price variability against city-level inflation (both variables expressed as deviations from the corresponding US national measure), confirming a positive association. They present evidence, however, that the relationship tends to decay as the time horizon over which the variability and inflation are calculated becomes longer. Both papers ignore the possibility that common aggregate shocks could be the cause of a positive relationship between relative price variability and inflation and create a more accurate measure of the degree of variability. They find a positive association between relative price variability and inflation that is robust to city and commodity-level extensions to the data-set.

Not all papers find a positive relationship between relative price variability and inflation, however. Reinsdorf (1994) represents an exception, finding a negative relationship in US data over a period of recession. Other papers by Cecchetti (1985), Blinder (1991), Lach and Tsiddon (1992) and Kashyap (1994) show that the duration of price quotations falls during inflationary periods. While this may raise the frequency of price adjustment, it does not necessarily imply an increase in dispersion of prices. If firms make adjustments to prices towards desired levels (defined by the law of one price) during inflationary periods to avoid (a) the explicit menu costs of continual adjustment (Ball and Mankiw 1994) or (b) the implicit costs through loss of market share (Rotemberg 1982), then price dispersion may fall. In this case, relative price variability will be negatively related to inflation as the firms reduce price dispersion in their own market during periods when the general price level is changing. In the Appendix we present a simple stylized model in which the relationship between the level of inflation and the variability of prices can be positive or negative.

Our paper offers a further extension to the new measures of Parsley–Debelle–Lamont by allowing the panel to extend to prices of product groups across several European countries. Measures of variability across product groups within countries will be equivalent to those reported by Parsley and Debelle and Lamont. Measures of variability for individual product groups across countries will differ from those constructed in the United States because of the need to convert prices reported in different currencies of denomination to a common unit of account. In this case, we will convert all them to Deutschmark equivalent values. To ensure that this does not affect the sign of the relationship between relative price variability and inflation, we condition on an exchange rate variation variable.

The two measures of relative price variability used in this paper are defined below and are based on measures used for a single country case described by Parsley (1996) and Debelle and Lamont (1997). We define p_{ijt} to be the price of the ith commodity in the jth country at time t.

The approach for calculating the relative price variability across product groups within each country is calculated using the average price as the sum across all i prices, p_{ijt}, to give $\bar{p}_{jt}$:

$$\bar{p}_{jt} = \sum_{i=1}^{m} p_{ijt} \Big/ m. \tag{3}$$

Using this average price, we construct the relative price R_{ijt}:

$$R_{ijt} = \ln\left(\frac{p_{ijt}}{\bar{p}_{it}}\right); \tag{4}$$

and, calculating $\bar{R}_{jt}$ as the average over all i, we can construct the relative price variability measure V_{1jt} using the square root of the sum of squares:

$$V_{1jt} = \left(\frac{1}{m-1}\sum_{i=1}^{m}(R_{ijt} - \bar{R}_{jt})\right)^{1/2} \tag{5}$$

The natural logarithm of this measure is used in Section III to compare the relative price variability for different products within each of the countries in our sample with the inflation rate for the relevant country.

We calculate an equivalent measure across countries for the same product group. The average price of the ith commodity over all j countries for each time period and product group is p_{it}, calculated as

$$\bar{p}_{it} = \sum_{j=1}^{n} p_{ijt} \Big/ n. \tag{6}$$

Using $\bar{p}_{it}$, we calculate the relative price, R_{ijt} as

$$R_{ijt} = \ln\left(\frac{p_{ijt}}{\bar{p}_{it}}\right). \tag{7}$$

Again summing over all j countries, we derive the average relative price:

$$\bar{R}_{it} = \sum_{j=1}^{n} R_{ijt} \Big/ n. \tag{8}$$

From the differential between R_{ijt} and $\bar{R}_{it}$, the relative price variability measure V_{2jt} is then calculated as

$$V_{2jt} = \left(\frac{1}{n-1}\sum_{j=1}^{n}(R_{ijt} - \bar{R}_{it})^2\right)^{1/2}. \tag{9}$$

The natural logarithm of this measure is used in Section IV to compare the relationship between relative price variability and average inflation for the same product group across countries.

II. Data and Methodology

The data used in this paper was compiled for the following 10 EU countries using monthly data taken from Eurostat over the sample period 1986(1)–1993(12): Belgium, Denmark, France, Germany, Greece, Italy, Luxembourg, the Netherlands, Spain and the United Kingdom.[5] For each country the prices were recorded for 15 product groups: dairy products, oils and fats; fruit and vegetables; tobacco; non-alcoholic and alcoholic drinks; clothing; footwear; furniture and household textiles; household machinery and appliances; vehicles; public transport; communications; recreational goods; recreational services; books, newspapers and magazines; and hotels, pubs and restaurants.

The sample period and the country set were truncated to produce a panel with an unchanging set of countries and product groups with no missing observations.

In this paper the primary tools of investigation into the relationship between relative price variability within countries and the respective national inflation levels are adopted from time series analysis. The data is de-seasonalized and the stationarity properties of the data are examined using ADF statistics constructed around a deterministic trend. However, there is always the possibility that the deterministic component is more complicated than a linear trend, and the unit root tests are replicated around a logistic trend. The nonlinearity introduced by the logistic terms is a smooth transition process (Leybourne *et al.* 1998). This is particularly appealing in the period covered by our sample, since it corresponds to the period of the EU Single Market programme, when the moments of the price series might be expected to experience transitions.

The suggestion that a smooth transition could be used as a means of representing a structural change arising from deterministic factors was originally proposed by Bacon and Watts (1971) and has been revived more recently by Granger and Terasvirta (1993), Lin and Terasvirta (1994) and Leybourne *et al.* (1998). It has the appealing feature that the transition in the series from one trend path to another is gradual but with limiting cases allowing non-transition or a discrete break in trend. The logistic function used to map the transition process is chosen by a maximum likelihood procedure and is fully described by two parameters: the midpoint, τ, and the speed of transition, γ. There are three possible models:

Model 1: $y_t = \alpha_1 + \alpha_2 S_t(\gamma, \tau) + v_t$,
Model 2: $y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + v_t$,
Model 3: $y_t = \alpha_1 + \beta_1 t + \alpha_2 S_t(\gamma, \tau) + \beta_2 S_t(\gamma, \tau)t + v_t$,

where the logistic transition process can be described as $S_t(\gamma, \tau) = \{1 + \exp[-\gamma(t - \tau T)]\}^{-1}$ and T is the sample size. These models are members of the family of LSTAR models discussed by Michael *et al.* (1997).

The logistic terms provide a mapping from one deterministic regime to another, since $S_{-\infty}(\gamma, \tau) = 0$ and $S_{+\infty}(\gamma, \tau) = 1$. The midpoint τ occurs where $S_{\tau T}(\gamma, \tau) = 0.5$. Thus, in model 1 the logistic describes a transition in the intercept, in model 2 it describes a transition in the intercept around an invariant linear trend, and in model 3 it describes the common transition in intercept and trend.

We use the test procedure suggested by Leybourne *et al.* (1998), which tests the stationarity of the residuals from models 1, 2 and 3 around a logistic trend against the null of a unit root process. The first step of the test procedure is to compute the nonlinear least squares estimates of deterministic component of the model and derive the residuals:

Model 1: $v_t = y_t - \alpha_1 - \alpha_2 S_t(\gamma, \tau)$,
Model 2: $v_t = y_t - \alpha_1 - \alpha_2 S_t(\gamma, \tau) - \beta_1 t$,
Model 3: $v_t = y_t - \alpha_1 - \alpha_2 S_t(\gamma, \tau) - \beta_1 t - \beta_2 t S_t(\gamma, \tau)$.

Using these residuals, an ADF statistic can be computed as the parameter ρ in

(10) $\Delta v_t = \rho v_{t-1} + \Sigma_i \eta_i \Delta v_{t-i} + \upsilon_t,$

where the last term is a residual. The critical values for the unit root test are tabulated in Leybourne *et al.* (1998).

The NLS estimation of the deterministic process is concentrated with respect to the parameters $\alpha_1, \alpha_2, \beta_1, \beta_2$ since these enter linearly in models 1–3. The minimization of the sum of squares reduces to a problem involving only the speed and midpoint of transition, τ and γ:

(11) $\min\,[SS = \Sigma_t(y_t - \pi' x_t)^2],$

where $\pi = \{a1, \alpha 2, \beta 1, \beta 2\}' = (\Sigma_t x_t x_t')^{-1}(\Sigma_t x_t y_t')$ and $x_t = x(\gamma, \tau) = \{1, t, S_t(), tS_t()\}$.

Having calculated the smooth transition and tested for unit roots, the deterministic component of the measures of relative price variability is removed by subtracting the smooth transition process. The detrended series are used to calculate persistence and the association with the level of inflation, which is also detrended. The equation used is

(12) $v_t = \phi_0 + \Sigma_i \rho_i v_{t-i} + \phi_1 \Delta p_t + \phi_2 (\Delta p_t)^2 + \phi_3 \Delta^2 p_t + \phi_4 \Delta(\Delta p_t)^2,$

where Δp is the average inflation rate corresponding to the variability series, and $(\Delta p)^2$ is this rate squared; the inflation rate is detrended, but in all cases a linear trend suffices.

TABLE 1
STATIONARITY TESTS

Country	Inflation	Lag	Inflation variability	Lag
Belgium	−7.18**	1	−3.57*	6
Denmark	−8.92**	1	−3.56*	12
France	−8.02**	1	−1.58	0
Germany	−5.58**	1	−8.42**	1
Greece	−5.00**	1	−4.25**	7
Italy	−3.92*	1	−2.15	12
Luxembourg	−6.16**	1	−2.66	12
Netherlands	−6.01**	1	−4.66**	12
Spain	−7.05**	1	−1.93	12
UK	−4.20**	1	−4.30**	12

* Significance at the 10% level; ** significance at the 5% level.

Note that in the interpretation of the regressions for each product group the variability measure is based on prices in different countries that have been converted to a single currency using the market exchange rate. If higher inflation provokes exchange rate adjustments within the ERM, then an increase in inflation could reduce the dispersion of prices owing to exchange rate realignments. Goldfajn and Valdes (1998) and Parsley and Popper (1998) suggest that exchange rate realignment is the most important reason for recent findings of reversions to PPP among ERM countries, although Bleaney and

Mizen (1997) find that exchange rate adjustments have only partly accommodated for inflation differentials. We should like to distinguish between the 'pure' effect of the inflation level on its variance and the exchange rate effect. One way of allowing for an exchange rate effect is to include in the regression a term that measures the magnitude of exchange rate movements during the period in question. If the exchange rate realignment effect is at work, then inflation differentials should be lower when there have been large exchange rate adjustments. Any additional negative correlation between the mean and variance of inflation across countries can be (tentatively) ascribed to a 'pure' inflation effect. Therefore, the regressions for variability in individual product group prices across countries include the standard deviation of countries' rates of exchange rate depreciation against the Deutschmark. This variable is intended to capture the magnitude of exchange rate movements.[6]

III. Relative Price Variability within European Countries

In this section we report the estimated relationship between relative price variability, calculated for overall product groups within each of the European countries, and a measure of inflation, recorded as the average change in price in each country in our sample. These correspond to the definitions in Section I that sum over all product groups j to give relative price variability and inflation measures for individual countries i, that is V_{1jt} and p_{it}.

Table 1 reports initial tests for stationarity, around a deterministic trend, in the inflation and relative price variability series. For inflation, the null that the series is non-stationary is rejected at the 5% level for all countries, with the exception of Italy which rejects the null at the 10% level. Relative price variability, on the other hand, presents more mixed results. The series is stationary for Germany, the Netherlands, the United Kingdom and Greece at the 5% level, and is stationary for Belgium and Denmark at the 10% level. The other countries—Luxembourg, Italy and France—fail to reject the null of a unit root, although for France this is quite likely to reflect the fact that there is a clear break in the trend in mid-sample. In order to allow for the possibility that a break in trend is responsible for the failure to reject the null, the unit root tests are then conducted around a smooth transition logistic trend.

The results for the relative price variability measure are reported in Table 2. Examining the ADF statistics around a smooth transition process shows that in only one case—Luxembourg—is there failure to reject the null of non-stationarity. All the other countries have stationary measures of relative price variability around a logistic trend process. The linear trend is estimated by the parameter, β_1, and these are significant in all countries except Belgium and the Netherlands. In all cases where the parameter on the trend process is significant, it is also very small. The parameter on the logistic component, β_2, describing the transition process in trend, $S(\gamma, \tau)$, is significant but again small for Belgium, France, Greece, the Netherlands, Spain and the United Kingdom. The γ parameter measures the speed of adjustment, which in most cases is slow (less than one), when significantly different from zero. In the case of Germany, Greece and Denmark, where the γ parameter is insignificantly different from zero (indicating that no transition occurs), the logistic trend parameter is also insignificant, implying that a deterministic trend is an appropriate

TABLE 2
STATIONARITY AROUND A LOGISTIC TREND

Country	α_1	β_1	α_2	β_2	γ	ADF(st.)	Lag
Belgium	−0.450	−0.001	−1.431	0.010	0.239	−4.86*	2
	(−0.799)	(−0.197)	(−2.610)	(1.946)	(3.017)		
Denmark	−2.227	0.013	0.008	−0.006	14.664	−5.45**	12
	(−83.332)	(35.026)	(0.085)	(−1.716)	(0.023)		
France	−2.801	0.020	−0.078	0.009	0.170	−4.97*	11
	(−20.650)	(14.073)	(−0.671)	(7.638)	(4.538)		
Germany	−1.926	0.007	−0.154	0.000	22.447	−6.31***	2
	(−19.599)	(8.193)	(−2.460)	(0.000)	(0.255)		
Greece	−1.446	0.012	−0.894	0.032	4.401	−7.21***	10
	(−43.809)	(27.174)	(−9.303)	(8.931)	(0.163)		
Italy	−1.784	0.009	−0.405	−0.005	0.123	−4.92*	12
	(−26.736)	(12.360)	(−3.055)	(−1.472)	(4.441)		
Luxembourg	−1.806	0.007	−0.010	−0.005	0.383	−4.52	0
	(−62.868)	(18.791)	(−1.770)	(−1.453)	(3.331)		
Netherlands	12.696	−0.123	−15.097	0.122	0.095	−5.05*	12
	(1.187)	(−1.196)	(−2.610)	(1.946)	(2.083)		
Spain	−1.500	0.008	0.011	−0.026	0.134	−4.80*	12
	(−11.934)	(6.083)	(0.077)	(−8.245)	(5.972)		
UK	−1.769	0.011	0.000	−0.059	0.049	−4.96*	12
	(−19.009)	(12.540)	(0.000)	(−8.893)	(33.343)		

Notes: The coefficient estimates refer to the parameters of the smooth transition model given as $y_t = \alpha_1 + \beta_1 t + \alpha_2 S(g, t) + \beta_2 S(\gamma, \tau)t + \varepsilon_t$, and the ADF statistic is the stationarity test around the logistic trend.
* Significance at the 10% level; ** significance at the 5% level; *** significance at the 1% level.

model for these countries. The results for other countries indicate that, although the transition is small and slow, allowing for the adjustment in trend over the sample is enough to lead to rejection of the null that the series are I(1).

Table 3 reports the estimated relationship between the relative price variability measure and the average inflation rate for the product group across all of the countries. The relative price variability measure is calculated for each product group across the 10 countries, is detrended by the smooth transition process and is recorded as v_t. The parameters $\rho_1, \rho_2, \ldots, \rho_{12}$ represent the persistence in the stochastic component of relative price variability lagged up to 12 months. The other parameters detect the relationship between relative price variability and inflation. The measure of persistence between current value and the one-period lag is significant at the 5% level in all countries except Spain, where it is significant at the 10% level. Belgium and Greece do not require more lags, but many countries include up to 12 lags of the residuals to remove residual autocorrelation and ARCH. These are jointly significant according to F-tests, despite individual insignificance on t-tests. The diagnostic statistics for the models are acceptable, rejecting ARCH and autocorrelation and misspecification in functional form. Normality is rejected for five countries, but this is due to outliers (and results excluding the outlier observations are very similar to the ones reported). The persistence of a shock to relative price variability is given for each country in the first column of Table 4 in

TABLE 3

INFLATION VARIABILITY AND THE LEVEL OF INFLATION

Country	ϕ_0	ρ_1	ρ_2	ρ_3	ρ_4	ρ_5	ρ_6	ρ_7	ρ_8	ρ_9	ρ_{10}	ρ_{11}	ρ_{12}	ϕ_1	ϕ_2	ϕ_3	ϕ_4
Belgium	0.067	0.689	—	—	—	—	—	—	—	—	—	—	—	–0.581	–2.254	1.153	3.19
	(2.104)	(9.921)												(–1.511)	(–2.604)	(1.114)	(1.72
Denmark	0.002	0.425	–0.051	–0.157	0.013	–0.079	0.084	–0.252	–0.097	0.018	–0.282	–0.017	–0.036	0.017	0.307	–0.073	–0.06
	(0.053)	(3.393)	(–0.337)	(–1.319)	(0.105)	(0.667)	(0.756)	(–2.388)	(0.866)	(0.169)	(–2.601)	(–0.147)	(–0.335)	(0.065)	(0.598)	(–0.144)	(–0.63
France	0.029	1.068	–0.378	0.152	—	—	—	—	—	—	—	—	—	–0.191	–3.433	0.243	6.41
	(0.263)	(10.07)	(–2.645)	(1.512)										(–0.191)	(–2.694)	(0.115)	(2.45
Germany	0.015	0.750	–0.559	—	—	—	—	—	—	—	—	—	—	0.207	–0.623	–0.849	1.90
	(0.519)	(7.763)	(–5.635)											(0.808)	(–0.520)	(–1.165)	(0.87
Greece	–0.022	0.233	—	—	—	—	—	—	—	—	—	—	—	–0.003	–0.474	0.015	0.13
	(–0.122)	(2.281)												(–0.012)	(–0.900)	(0.142)	(0.70
Italy	–0.579	0.814	–0.140	–0.144	—	—	—	—	—	—	—	—	—	2.579	–0.828	–2.816	0.83
	(–2.594)	(7.041)	(–0.988)	(–1.327)										(2.570)	(–0.499)	(–2.543)	(0.46
Netherlands	–0.005	0.527	–0.063	–0.100	–0.0003	0.187	–0.039	0.028	–0.064	–0.02	–0.267	0.235	–0.150	–0.613	0.956	–1.477	–1.89
	(–2.786)	(4.326)	(–0.457)	(–0.759)	(–0.003)	(1.476)	(–0.308)	(0.218)	(–0.502)	(–0.175)	(–2.132)	(1.866)	(–1.144)	(–2.784)	(1.378)	(–2.502)	(–1.42
Spain	0.184	0.186	0.158	–0.138	–0.045	0.107	–0.0778	–0.264	0.084	–0.059	–0.135	—	—	–0.776	0.755	0.788	–0.28
	(0.630)	(1.635)	(1.263)	(–1.059)	(–0.357)	(0.898)	(–0.644)	(–2.185)	(0.707)	(–0.516)	(–1.248)			(–0.594)	(0.522)	(0.552)	(–0.18
UK	–0.028	0.788	0.007	–0.043	–0.043	–0.0784	0.128	–0.225	0.228	0.038	–0.137	0.221	–0.137	0.172	0.013	–0.198	–0.11
	(–1.604)	(5.972)	(0.041)	(–0.258)	–0.273)	(–0.534)	(0.906)	(–1.702)	(1.734)	(0.287)	(–1.036)	(1.656)	(–1.044)	(2.085)	(0.063)	(–2.240)	(–0.58

Notes: The coefficient estimates refer to the parameters of the model given in equation (12). The coefficients ρ_i are reported for up to 12 months. The model for Germany includes term the square of the residual v_t, which are effectively capturing ARCH effects. The coefficients on these three lags are –0.703 (–1.773), 0.163 (0.424) and 0.189 (0.507), respectively.

Table 4
Model Diagnostics

Country	Half-life (months)	R^2	Equation standard error	F-statistic (p-value)	ARCH (p-value)	*RESET* $F(1, 80)$	Normality (p-value)	Error auto-correlation (p-value)
Belgium	1–2	0.586	0.0629	0.000	0.473	0.203	0.000	0.291
Denmark	0–1	0.352	0.0376	0.017	0.981	0.063	0.703	0.708
France	2–3	0.748	0.0351	0.000	0.641	0.132	0.003	0.447
Germany	not unique	0.475	0.104	0.000	0.521	0.771	0.000	0.356
Greece	0–1	0.119	0.0556	0.053	0.249	0.511	0.162	0.523
Italy	3–4	0.625	0.0266	0.000	0.817	0.166	0.611	0.198
Netherlands	1–2	0.558	0.0575	0.000	0.124	0.052	0.002	0.228
Spain	0–1	0.402	0.0466	0.001	0.118	0.227	0.003	0.701
UK	2–3	0.758	0.0228	0.000	0.122	0.506	0.923	0.966

terms of the two calendar months between which the half-life falls. The greatest persistence is seen in France, Italy and the United Kingdom, where the half-life is around three months; all the other countries have a half-life of less than two months and in three cases it is less than one month. This indicates that the relative price variability series around a logistic trend are short memory processes, whose shocks decay quickly.

The relationship between relative price variability and inflation, inflation-squared, the change in inflation and the change-in-inflation-squared is more mixed. For Italy, the Netherlands and the United Kingdom, the level of inflation and the level-of-inflation-squared are significant. When evaluated at their mean values, the net effect is positive for Italy, Spain and the Netherlands; for Belgium, Germany and France there is a net negative effect. The pattern between relative price variability and inflation shows a marked contrast to the results from US cities reported by Debelle and Lamont (1997) and Parsley (1996), where all the findings confirmed a positive relationship. This illustrates that the relationships differ across countries in our data-set. (Our parameter estimates are not constrained to be equal across the countries.) Rather than a systematic common relationship between countries within Europe, there is considerable diversity. This implies that it is not possible to provide simple categorizations of countries into groups, and this will make the control of inflation harder for a European monetary authority.

Although there is not a systematic relationship between inflation and relative price variability across Europe, there is some similarity in the measures of the persistence of shocks to relative price variability. The persistence measure demonstrates that almost the entire shock to relative price variability is eliminated over a four-month horizon. This accords with the findings of Blinder (1991), who found that, among 72 large companies surveyed in the United States, prices were changed once per year on average and that adjustment to a shock to prices had a mean lag of three to four months. The lack of persistence in shocks to relative price variability beyond one calendar year implies that any asymmetric shocks to prices disappear over a very short horizon. Thus, the actions of monetary authorities can be dedicated towards the longer-term objective of inflation control rather than the reduction, in any systematic way, of short-term relative price variability within European countries.

IV. Relative Price Variability within Product Groups

In this section we examine the behaviour of relative price variability in each of the 15 product groups across European countries. The methodology is similar to the previous section. These measures correspond to the definitions in Section I that sum over all countries i, to give relative price variability and inflation measures for individual product groups, j, that is V_{2jt} and p_{jt}.

Table 5
Stationarity Tests

Commodity	Inflation	Lag	Inflation variability	Lag
Dairy produce	−10.49***	0	−6.05***	0
Fruit and vegetables	−6.39***	0	−4.01**	0
Tobacco	−9.72***	0	−2.18	13
Drinks	−8.14***	0	−1.27	11
Clothing	−8.88***	0	−2.10	12
Footwear	−10.53***	0	−2.66	12
Furniture	−9.09***	0	−2.47	11
Appliances	−8.52***	0	−2.59	11
Vehicles	−7.44***	0	−2.39	9
Public transport	−9.49***	0	−2.42	13
Communications	−4.57***	3	−1.10	9
Recreational goods	−9.78***	0	−2.22	0
Recreational services	−9.57***	0	−2.90	11
Books	−8.65***	0	−4.35**	11
Hotels	−8.45***	0	−1.65	9

* Significance at the 10% level; ** significance at the 5% level; *** significance at the 1% level.

Table 5 reports the stationarity tests around a deterministic trend and a constant for the relative price variability measure for each commodity across the European countries and for inflation in each commodity. It is possible to reject the null of nonstationarity for all of the inflation series, but for only 3 out of 15 relative price variability measures. Following the procedure outlined in the last section, we examine the hypothesis that the series is stationary around a smooth transition, described by a logistic trend.

Table 6 reports the results of smooth transition analysis. For 13 of the 15 product groups there are significant parameters on the smooth transition process in trend and intercept. These product groups reject the nonstationary null once the deterministic component describes a nonlinear transition: four at the 10% level, seven at the 5% level and one at the 1% level. Three series—tobacco, appliances and recreational services—fail to reject nonstationarity even allowing for a smooth transition process. Two others—fruit and vegetables and books—for which smooth transition processes are not detected, were among the series that were found to reject the null of nonstationarity under a linear trend. For these product groups a linear trend process is an adequate representation of the deterministic process. The significant γ coefficients for all the other series are small, indicating slow transition processes; the absence of γ coefficients for fruit and vegetables and books reflects the fact that no transition has occurred. The relative price variability for the majority of individual

TABLE 6
STATIONARITY AROUND A LOGISTIC TREND

Commodity	α_1	β_1	α_2	β_2	γ	ADF(st.)	Lag
Dairy produce	−3.234 (−11.584)	0.009 (3.335)	1.340 (4.776)	−0.011 (−4.348)	−5.158 (−0.112)	−5.29**	15
Fruit and vegetables	−0.707	—	—	—	—	−4.01**	0
Tobacco	−1.667 (−24.786)	−0.011 (−8.935)	8.024 (5.332)	−0.057 (−4.520)	0.128 (5.399)	−4.54	17
Drinks	−2.084 (−216.58)	—	−0.226 (−6.375)	—	2.532 (1.068)	−4.04*	3
Clothing	−1.642 (−64.603)	−0.009 (−13.066)	−0.855 (−4.809)	0.014 (6.664)	0.390 (1.658)	−4.82*	12
Footwear	−2.431 (−16.932)	0.013 (3.014)	−0.029 (−0.216)	−0.008 (−2.016)	0.233 (3.565)	−5.24**	13
Furniture	−1.877 (−30.984)	−0.008 (6.509)	0.912 (9.663)	0.016 (10.819)	0.720 (1.189)	−4.85*	6
Appliances	−3.359 (−15.086)	0.041 (5.707)	−0.980 (−5.001)	−0.038 (−5.613)	0.164 (2.714)	−4.55	12
Vehicles	−2.145 (−50.70)	0.0015 (1.923)	0.474 (3.992)	−0.0014 (−1.085)	0.473 (3.280)	−6.21***	17
Public transport	−7.789 (−13.113)	0.028 (7.210)	6.236 (7.513)	0.104 (14.736)	−0.023 (−21.05)	−4.62*	0
Communications	−2.870 (−56.619)	0.020 (13.046)	−2.719 (−11.545)	0.012 (9.874)	−0.098 (−11.898)	−5.01**	1
Recreational goods	−3.598 (−29.173)	0.012 (9.948)	0.449 (1.823)	0.025 (4.469)	−0.127 (−7.428)	−5.03**	1
Recreational services	−1.549 (−30.096)	−0.010 (−9.530)	0.178 (1.645)	0.003 (1.994)	0.974 (1.843)	−4.46	0
Books	−0.539 (−4.210)	0.001 (3.849)	—	—	—	−4.35**	11
Hotels	−1.806 (−47.820)	−0.007 (−9.115)	−5.631 (−8.408)	−0.044 (−8.078)	0.108 (6.639)	−5.06**	17

Notes: see notes to Table 2. The ADF statistics are taken from Leybourne *et al*. (1998, table 1). The exception is the ADF statistic for fruit and vegetables, which has standard critical values (Dickey and Fuller 1976).

product groups across the 10 selected European countries appear to be described accurately by a smooth transition process.

Table 7 reports the results of estimating the relationship between price variability and inflation. (Results for the three variability series for which the null of nonstationarity was not rejected are also reported, since in all cases the stationarity test statistics are close to the 10% critical value; the reported t-ratios in these three cases should be treated with extreme caution.) The table shows first of all the estimates of the coefficients on the lagged dependent variable (the ρs), which measure persistence. There is generally high persistence, but this falls short of and is significantly different from one; the sum of the ρ_i coefficients lie on the range 0.591–0.905, although 11 of the 15 values are greater than 0.7.[7] As in the cross-commodity model, the persistence estimates are conditioned on detrended inflation, inflation-squared, change in inflation and change-in-inflation-squared, which are all stationary series. There is an extra variable in this regression, however: the (detrended) cross-country

TABLE 7

INFLATION VARIABILITY AND THE LEVEL OF INFLATION

Commodity	ρ_1	ρ_2	ρ_3	ϕ_1	ϕ_2	ϕ_3	ϕ_4	ϕ_5
Dairy produce	0.569	0.395	−0.147	−3.703	1.903	−13.205	−21.696	0.002
	(5.303)	(2.985)	(−1.348)	(−4.368)	(3.692)	(−0.859)	(−1.387)	(0.238)
Fruit and vegetables	0.700	−0.013	−0.095	−3.925	0.412	−57.129	39.604	0.016
	(5.995)	(−0.101)	(−0.735)	(−3.542)	(0.403)	(−1.086)	(1.080)	(0.614)
Tobacco	0.877	—	—	−1.464	0.666	19.158	−0.634	0.016
	(15.860)			(−1.291)	(0.886)	(1.081)	(−0.057)	(0.875)
Drinks	0.774	—	—	−1.004	1.897	−7.348	20.480	0.023
	(11.570)			(−0.986)	(2.588)	(−0.189)	(0.740)	(2.132)
Clothing	0.762	—	—	−1.978	1.232	−109.67	65.044	0.028
	(10.449)			(−1.375)	(1.293)	(−2.670)	(2.343)	(2.418)
Footwear	0.717	—	—	−2.151	0.815	−17.681	53.368	0.028
	(10.267)			(−2.114)	(1.302)	(−0.858)	(3.134)	(2.505)
Furniture	0.770	—	—	−3.784	0.4051	−116.80	28.982	0.021
	(13.680)			(−3.053)	(0.476)	(−2.005)	(0.671)	(2.207)
Appliances	1.047	−0.178	—	−2.545	−1.360	−63.921	−8.779	0.020
	(9.412)	(−1.683)		(−2.200)	(−1.563)	(−1.183)	(−0.217)	(2.282)
Vehicles	0.875	—	—	−1.472	0.693	11.571	−0.613	−0.006
	(16.490)			(−1.547)	(1.104)	(0.778)	(−0.061)	(−0.363)
Public transport	0.667	—	—	−0.985	0.443	−15.952	20.097	0.006
	(7.784)			(−1.293)	(0.827)	(−0.805)	(1.401)	(0.466)
Communications	0.628	—	—	−2.648	0.214	−43.825	8.467	0.002
	(7.167)			(−3.366)	(0.369)	(−1.203)	(0.325)	(0.198)
Recreational goods	0.577	—	—	2.870	−1.621	270.36	−64.589	0.023
	(6.499)			(1.378)	(−1.190)	(2.110)	(−0.674)	(1.136)
Recreational services	0.492	0.265	0.127	0.221	−0.540	−173.60	59.273	−0.015
	(4.200)	(2.116)	(1.056)	(0.243)	(−0.887)	(−3.972)	(1.753)	(−1.424)
Books	0.993	−0.166	—	−2.238	0.418	−26.986	−11.885	0.019
	(9.264)	(−1.618)		(−2.931)	(0.734)	(−0.576)	(−0.353)	(2.425)
Hotels	0.806	—	—	−1.365	0.689	−72.554	14.462	0.019
	(13.107)			(−1.445)	(1.021)	(−1.411)	(0.404)	(1.948)

Notes: see notes to Table 3. Drinks includes lags of the dependent variable squared up to order 3. The parameters on these terms are 0.717 (0.833), −3.889 (−4.689) and 3.017 (4.298), respectively. 'Fruit and vegetables' includes lags of the level up to order 7. The higher-order lags parameters are −0.190 (−1.466), 0.003 (0.022), 0.165 (1.291) and −0.213 (−1.992), respectively. 'Recreational services' includes lags of the level up to order 12. The higher-order lags are 0.093 (0.751), −0.174 (−1.210), 0.033 (0.222), 0.104 (0.745), −0.057 (−0.412), −0.165 (−1.201), −0.054 (−0.391), 0.376 (2.755) and −0.328 (−2.911), respectively.

standard deviation of exchange rate depreciation against the Deutschmark. This measure of the magnitude of exchange rate movements is intended to capture exchange rate realignment effects. The coefficient on the exchange rate term is ϕ_5.

The half-lives of the shocks are recorded in the first column of Table 8. These results indicate that the shocks to relative variability within product groups take notably longer to decay than the shocks to relative price variability within countries. There are no product groups for which the half-life is less than a calendar month, and there are several product groups for which it is greater than four months. This demonstrates that persistence is higher in relative price variability within product groups than the equivalent measure within countries, and the higher persistence is often associated with product groups that are imperfectly substitutable (such as books, hotels, recreational services). Carlton (1986, 1989) has shown that prices are more rigid (implying that relative price variability would be more persistent after a shock) for more product groups from concentrated industries. All the product groups, with the exception of dairy products, show evidence that shocks decay away completely within one year.

The effect of an increase in the level of inflation on relative price variability is predominantly negative, but the short-run effect is typically smaller than the long-run effect; this is indicated in the table as a negative value of ϕ_1 and a positive value of ϕ_2. The negative ϕ_1 term is significantly negative for seven series, while the ϕ_2 term is significantly positive for just two. The coefficients on inflation-squared are usually significantly negative, indicating some nonlinearity; and in all cases (except recreational goods) the net effect of higher inflation, calculated at its mean, is negative; i.e., as inflation falls relative price variability rises. The negative net effect is significant in 12 cases. The implication of these results is that, as inflation in the price of product groups increases, so the relative variability between the prices in each of the countries gets smaller, but if prices accelerate there is some effect in the opposite direction.

Our results in this section contrast with the mixed results from countries, and the strong positive relationship detected in US city data by Parsley (1996) and Debelle and Lamont (1997). It is consistent with the empirical finding that the law of one price tends to hold more strongly with higher inflation. Studies that test for PPP on disaggregated data, such as Rogers and Jenkins (1995) and C. Engel and Rogers (1996), support the hypothesis that there are frictions to the price-setting process, justifying a negative relationship between price variability and inflation. Our results show that relative price variability within commodity groups is persistent but that the effect of inflation brings down the relative price variation, and accords with the menu cost argument proposed by Ball and Mankiw (1994). Essentially, the argument implies that, when there are menu costs associated with changing prices, commodity prices are not adjusted very readily and the differences in prices for the same commodity across countries tend to persist. When inflation does occur, it creates the need to change menu prices to account for the rising average price level, and while these changes are taking place adjustments are made to reduce the relative price variability in the commodity price across countries. This reduces the implicit costs of price changes that result in lost market share (Rotemberg

TABLE 8
MODEL DIAGNOSTICS

Commodity	Half-life (months)	R^2	Equation standard error	F-statistic (p-value)	ARCH (p-value)	RESET (p-value)	Normality (p-value)	Error auto-correlation (p-value)
Dairy produce	2–3	0.782	0.041	0.000	0.939	0.777	0.515	0.307
Fruit and vegetables	1–2	0.803	0.103	0.000	0.229	0.167	0.711	0.724
Tobacco	5–6	0.797	0.077	0.000	0.905	0.963	0.001	0.201
Drinks	not unique	0.808	0.044	0.000	0.799	0.499	0.803	0.808
Clothing	2–3	0.797	0.049	0.000	0.671	0.301	0.174	0.349
Footwear	2–3	0.811	0.046	0.000	0.263	0.369	0.665	0.721
Furniture	2–3	0.843	0.041	0.000	0.982	0.377	0.618	0.496
Appliances	5–6	0.899	0.037	0.000	0.299	0.651	0.800	0.434
Vehicles	5–6	0.817	0.039	0.000	0.631	0.309	0.002	0.707
Public transport	1–2	0.536	0.500	0.000	0.134	0.465	0.886	0.316
Communications	1–2	0.599	0.039	0.000	0.617	0.118	0.024	0.501
Recreational goods	1–2	0.554	0.087	0.000	0.300	0.125	0.000	0.399
Recreational services	not unique	0.875	0.038	0.000	0.565	0.985	0.197	0.901
Books	4–5	0.847	0.033	0.000	0.124	0.897	0.075	0.178
Hotels	3–4	0.773	0.042	0.000	0.820	0.752	0.273	0.624

1982). Dynamic models of menu costs predict an increase in the frequency of price changes with a rising inflation rate, reducing persistence. Evidence from Lach and Tsiddon (1992) on Israeli foodstuffs implies that the duration of the price quotes falls with inflation, and this result has been confirmed on other data-sets by Kashyap (1994). This may explain why we find a net negative relationship between inflation and relative price variability.

V. Conclusions

The paper presents new information on inflation and relative price variability across Europe. Previous papers have disaggregated national data of non-European countries to the commodity level by focusing on evidence from cities within a single country. The evidence reported here has taken 10 countries and 15 different product groups to examine the diversity in relative price variability and inflation in Europe, where significant differences in behaviour will have important implications for the viability of monetary integration.

The measures of relative price variability employed are similar to those reported by Parsley (1996) and Debelle and Lamont (1997). The data on prices are used to calculate relative price variability measures across product groups within the same country and across countries for the same product group. The time series behaviour of these measures is investigated, and for 9 out of 10 countries and 13 out of 15 product groups the null of non-stationarity is rejected around a smooth transition in the deterministic trend. In all cases, across both countries and product groups, the smooth transitions are very gradual, indicating that although there is a significant transition the adjustment is slow.

These findings of stationarity are confirmed by tests of persistence in the relative price variability measure which, while showing some persistence for many of the 10 countries and all 15 product groups, reject the null of a unit root. The memory of the series shows that the decay to the relative price variability measure within countries is rapid—with reported half-lives of less than four months—and is highest for the relative price variability measures within product groups. As a result, shocks to relative price variability are all but eliminated after 12 months.

The evidence for a relationship between relative price variability and inflation is not as systematic in European countries as it is in the United States. It does not show the pronounced positive influence found in data from cities within the United States reported by Parsley (1996) and Debelle and Lamont (1997), for example, and this makes the task of a monetary authority in Europe all the more difficult. Within product groups, there is more uniform evidence for a significant negative relationship. A negative relationship is consistent, on the demand side, with the empirical evidence on the law of one price, which tends to show that mean reversion is stronger when the price level is changing. It also shows the importance of menu costs to price-setting on the supply side, since the results show that price-setters typically eliminate relative differences in prices across countries when they need to make changes to prices to account for inflation. So there is more persistence in relative price variability when inflation is low, but when inflation occurs relative price variability falls.

Two points stand out from these results. First, for data within countries and within product groups, shocks to relative price variability are virtually eliminated over a 12-month horizon. The half-lives of the stochastic components around logistic trends are extremely short, and in no case does the half-life exceed six calendar months. Second, there is an absence of a systematic strong and positive relationship to inflation that suggests that intervention to control inflation could not be relied upon to control relative price variability. This is most evident in the relative price variability within countries, and it implies that attempts to find a policy acceptable to all the European nations would be frustrated by the lack of any systematic relationship between inflation and relative price variability. However, since the relative price variability series are short memory processes which decay quickly, this is less problematic than it first appears. Even within product groups, although relative price variability is negatively related to inflation, the effect of shocks is eliminated within one year. The speed of the decay demonstrates that shocks to variability in prices do not persist for long. In policy terms, this suggests that monetary policy can be dedicated to the long-term task of reducing inflation, while the harmonization of short-term relative price variability can be left to market forces.

APPENDIX: A MODEL OF MENU COSTS AND PRICE DISPERSION

Here we present a simple model relating to the price-setting behaviour of a firm. We do not attempt to bridge the gap between the individual firm and the behaviour observed within countries or within commodity groupings. If firms behave as we suggest, the relationship between cross-sectional price dispersion and inflation can be positive or negative. In this model a negative relationship arises when a higher inflation rate is associated with lower 'menu costs', and a positive result arises if the opposite is the case. The term 'menu costs' covers a range of possible costs associated with the act of raising prices—for example administrative costs, or the possibility that brand loyalty is decreasing in the number of price changes. Their existence can account for positive results found by Debelle and Lamont (1997) and Parsley (1997) or for negative results reported by Reinsdorf (1994) and our paper based on the response of price-setting behaviour by firms to inflation. Other authors (e.g. Caballero and Engel 1994) present menu cost models with a much more complex structure (but with an exogenous bandwidth corresponding to our endogenous variable g), in which a negative relationship can arise for slightly different reasons.

First, consider an imperfectly competitive firm which is maximizing profits over a finite time horizon, $t=[0,N]$. In any one period (the log of) the optimal output price set by the firm, in the absence of menu costs, is equal to $b(t)$; the losses arising from any deviation of (the log of) the actual output price ($p(t)$) from the optimum are some function of the difference between the two, say $c_1(t)=f(p(t)-b(t))$. Impose the normalization $b(0)=0$, and let $\mathrm{d}b/\mathrm{d}t=\pi$ (π is some positive function of aggregate price inflation). The menu costs of adjusting the output price k times in the N periods are proportional to k: $c_2=ak$.

Case 1 (*linear costs*): $c_1(t)=p(t)-b(t)$

In order to illustrate the problem, consider Figure A1. The optimal price level in each period (disregarding menu costs) is given by the line labelled πt. In the figure the firm has chosen $k=3$: there are three price levels, evenly spaced over the range πN, and the total cost of price deviations is the sum of the area of the six triangles Δ_3. In general,

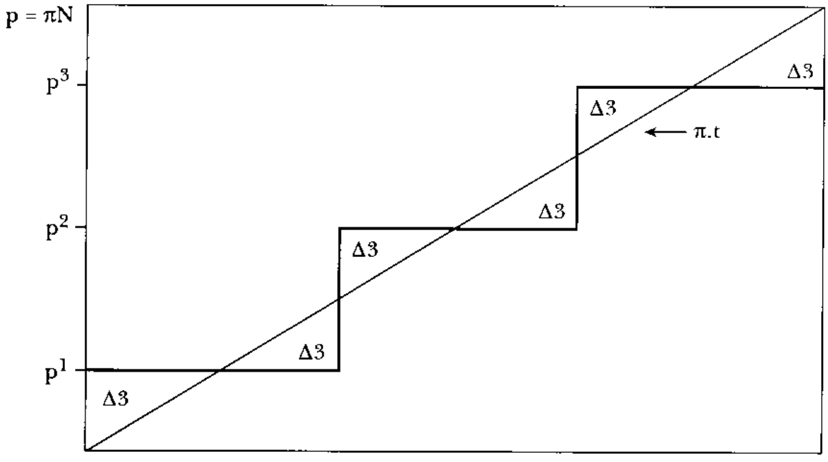

FIGURE A1

the total cost of price deviations is

$$\text{(A1)}\qquad \int_{t=0}^{t=N} c_1(t)\,\mathrm{d}t = 2k\left(\int_{=0}^{t=N/2k} \pi t\,\mathrm{d}t\right) = \pi N^2/4k.$$

The firm chooses k to minimize the cost of price deviations plus menu costs, $\pi N^2/4k + ak$, so that k is the integer closest to

$$\text{(A2)}\qquad k^* = N\sqrt{(\pi/4a)},$$

and the gap between prices ($\pi N/k^*$) is

$$\text{(A3)}\qquad g^* = \sqrt{(4a\pi)}.$$

If menu costs are independent of π, then g^* will be increasing in π. In an equilibrium with a constant π and a constant number of identical firms coming into existence in every instant, the distribution of prices set will be uniform on an interval of width g^*. A higher π entails a higher g^*, and so an increase in the dispersion of prices set. This means a positive relationship between inflation and cross-sectional price variability.

However, if menu costs are decreasing in the inflation rate π, then this result no longer necessarily holds. Menu costs might be decreasing in the inflation rate because, for example, higher inflation means that consumers loyal to a particular brand are less likely to commence a search for a better price when the price of that brand rises. (In the presence of high general inflation, consumers are less likely to interpret a rise in the price of 'their' brand as a fall in competitiveness.) Suppose that marginal menu costs are given by $a = h\pi^{-\beta}$, with $\beta > 0$; this means that the optimal g is equal to

$$\text{(A4)}\qquad g^* = \sqrt{(4h\pi^{1-\beta})},$$

which is decreasing in π for $\beta > 1$.

Case 2 (*quadratic costs*): $c_1(t) = (p(t) - b(t))^2$
This is like case 1, except that we now have

$$\text{(A5)}\qquad \int_{t=0}^{t=N} c_1(t)\,\mathrm{d}t = 2k\left(\int_{=0}^{t=N/2k} (\pi t)^2\,\mathrm{d}t\right) = \pi^2 N^3/8k^2,$$

which entails

(A6) $k^* = N \cdot (\sqrt[3]{(\pi^2/4a)}),$

(A7) $g^* = \sqrt[3]{(4a\pi)},$

and the same general observations are true as in Case 1.

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NOTES

1. The exceptions are the papers by R. Engle (1982) and Joyce (1995) for the UK and Van Hoomissen (1988) for Israel. Section I gives further details of the papers that have studied inflation variability and inflation uncertainty.
2. This makes our study different from Parsley (1996) and Debelle and Lamont (1997), who studied cities within the same country with a common monetary policy.
3. If the common finding of a positive association between inflation variability and the level of inflation revealed in studies in the United States is confirmed within Europe, then the monetary policy may not be quite so complicated. If there is a positive relationship between inflation and relative price variability, then macroeconomic control of inflation will imply control of relative price variability. Only if there is an absence of evidence of a positive relationship will policies aimed at the microeconomic level may be required to eliminate relative price variability.
4. In this paper we consider the empirical measurement and methodology of inflation variability and do not discuss the theoretical models that generate such results. We refer the reader to papers by Lucas (1973), Barro (1976), Seshinski and Weiss (1977), Friedman (1977), Cukierman (1982, 1984), Benabou (1988), Van Hoomissen (1988), Cecchetti (1985) and Ball and Mankiw (1994, 1995), which illustrate the link between inflation and relative price variability using imperfect information, menu costs and search costs. Friedman's Nobel Lecture makes the connection between inflation in the level, its variability and the uncertainty surrounding forecasts of inflation.
5. Although we might like to include other European countries, especially those that are likely to participate in the first phase of the single currency, some countries, such as Eire and Portugal, were excluded because of constraints on the availability of data. The countries in our sample are the only countries that report disaggregated data on a comparable basis for the full sample period.
6. A warning ought to be included here. If some large exchange rate movements are adjustments of the Parsley and Popper type and some are not, then our exchange rate variable (intended to capture just the former) will contain some measurement error. However, since the main reason for adjustments within the ERM has been real exchange rate misalignment, this error is likely to be very small.
7. These levels of persistence explain why the ADF results, after accounting for a nonlinear transition in trend, generally reject at the 5% or 10% level. For two of the series—appliances and books—the parameter on the first lag is close to unity and, given the low power of ADF tests in this context, explains why the smooth transition models for these series could not reject a unit root in the residual. The sum of the first two ρ_i coefficients is also very close to unity for dairy products. When we include the effect of other significant lags, the combined effect of the ρ_i coefficients gives a memory for the series that is less than infinite, rejecting a unit root. These results are confirmed by calculating the half-lives.

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