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# The Relationship between Price Dispersion and Inflation: A Reassessment

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# The Relationship between Price Dispersion and Inflation: A Reassessment

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***Abstract***

The positive relationship between inflation and relative price variability has been taken as a proven case, but recent evidence casts doubt on the verdict. This paper shows that theory can support positive or negative relationships, while past empirical evidence may have been contaminated by aggregation bias and our measures of price dispersion and inflation. We re-examine the empirical evidence for ten years of data over nine commodity groups and seven European countries using a non-parametric approach to estimate the functional form.. Our results suggest that the results offer strong support for the model of Danziger (1987) which predicts a nonlinear functional form for within-commodity across-countries data.

[§] The authors are from University of Leicester and University of Nottingham respectively. This paper was written while the second author was a visiting professor at the European University Institute, Florence. We wish to thank Mike Artis, Anindya Banerjee, Guiseppe Bertola, and Soren Johansen for helpful comments. Any remaining errors are our own.

# 1 Introduction

Supportive evidence for a positive empirical relationship between inflation and price dispersion has been known for some time. Mitchell (1915) discovered a positive association in US data for 1890-1910 and Mills (1927) also found evidence to support this view, using US data for 1892-1926. In Germany, Graham (1930) found an association during the hyperinflation years of 1920-23. Later papers have also supported a positive relationship between price dispersion and inflation, for example, Okun (1971), who used cross-section data for 17 OECD countries to analyse the relationship between the first two moments of the distribution of inflation for the period 1951-68. Vining and Elwertowski (1976) presented graphical evidence to show that time series data of the standard deviation of changes to relative prices and the inflation rate also supports a positive relationship. Lucas' (1973) 'islands' model of price setting behaviour when the general price level is unknown provides theoretical support for these positive results. Consequently, most investigators have concluded that the positive relationship between inflation and relative price variability has been theoretically and empirically proven.

More recently, however, the theory has been questioned, and the evidence is less clear-cut now than it was twenty years ago. A reinterpretation of Lucas' theory by Riensdorf (1994) has shown that the 'islands' model is also consistent with a negative relationship. A significant contribution by Danziger (1987) demonstrates that with costly adjustment of prices, firms will have incentives to set prices only periodically, and under these circumstances the relationship between relative price variability and inflation is not necessarily positive. The study suggests that the relationship between price dispersion and inflation is nonlinear quadratic, the derivative of the function taking positive or negative values for different values of inflation. This goes some way towards explaining why, under certain conditions, a negative rather than a positive relationship has been discovered in US and European data (c.f. Reinsdorf, 1994 and Fielding and Mizen, 2000).

These theoretical developments beg the question why the majority of the early studies found a positive relationship. Hartman (1991) suggests that the degree of time-aggregation and the dependence of both inflation and price dispersion measures on the same underlying random variables can by construction generate a positive sign on inflation in a standard regression of price dispersion on inflation. Common shocks to food and energy components of the price series may have contributed to simultaneous rises in inflation and variability of prices (Fischer, 1981 and Driffill *et al.* 1990). Later studies are less likely to suffer from aggregation effects, because higher frequency time-series data and an increase in the number of commodities covered has facilitated panel estimation with the conditioning on time and group effects (e.g. Parsley, 1986; Debelle and Lamont, 1997). The findings summarised in Table 1 indicate that the number of studies finding an unambiguously positive relationship between relative price variability and inflation has fallen with the use of more methodologies incorporating these effects. Among the annual and quarterly studies of the 1970s and 1980s, all the papers support a positive linear relationship, with the exception of Parks (1978), which was ahead of its time. The monthly studies of the 1980s and 1990s report a larger number of negative correlations between price dispersion and inflation. The increasing heterogeneity of the results is due partly to the application of models of both expected and unexpected inflation, and partly to the use of nonlinear functional forms and higher frequency data. van Hoomissen (1988) captures the nonlinearity of the relationship in a negative coefficient on the square of inflation, while Reinsdorf (1994) reports a negative relationship in US data over the Volcker years. Fielding and Mizen (2000) find a negative relationship around the smooth transition in trend for European data during the single market programme. Most of the studies in the 1990s find linearity is rejected in favour of a nonlinear alternative, although there is little agreement about the functional form or the signs of the coefficients that describe the relationship between price dispersion and inflation. So the conclusion that the relationship is unambiguously positive may be premature.

We re-examine the empirical evidence for ten years of data over nine commodity groups and seven European countries during the period of the single market programme. We take a non-parametric approach to estimation of the relationship in order to allow the greatest possible flexibility with respect to functional form. We also use data collected at monthly and quarterly frequencies for the same commodities and countries to investigate the effects of the measurement frequency (relative to the frequency of price setting) on the nature of the empirically estimated function. Our results suggest that data frequency makes some difference to which functional form best characterises that data. But the greatest contrasts arise as we move between price variability measured across countries for individual commodities and price variability measured across commodities for individual countries. In the first case there is a nonlinear relationship, which appears to be quadratic rather than piecewise linear; in the second there is no significant nonlinearity. This result is very much consistent with the Danziger model of price setting behaviour.

The rest of the paper is organised as follows. Section 2 asks what we know about the theory and measurement of the relationship between price dispersion and inflation. This motivates nonparametric approach that is described in Section 3. The results from this methodology are applied and interpreted in Section 4. Section 5 concludes.

# 2 What do we know about price dispersion and inflation?

In this section we show that there is more dispute in the literature than is commonly supposed with regard to the form of the relationship between price dispersion and inflation. We show that the theoretical literature does not support an unambiguously positive relationship, nor does it offer a unique functional form, but the frequency of observation of prices relative to the frequency of price changes and the measurement of relative price dispersion and inflation are such that a positive relationship is often the most likely outcome.

## 2.1 Theory

Consider the well-known Islands model of imperfect information (Lucas, 1973), which suggests that islanders face a signal extraction problem made more difficult when noise (which is positively correlated with the rate of inflation) increases. Although islanders can directly observe the price on their own island, general prices can only be observed with a lag. When forming expectations about general prices, the weight in placed on the own-island price versus the mean from the prior distribution of general prices declines with the variance of the deviation of the island price from the average (i.e. with relative price variability). This relationship has often been used to motivate a positive relationship between price dispersion and inflation (Lach and Tsiddon, 1992; Debelle and Lamont, 1997), although Hercowitz (1981) demonstrates that it is *unanticipated* monetary policy (inflation) that generates the misperceptions that lead to greater price dispersion i.e. higher relative price variability.

Yet Reinsdorf (1994) demonstrates that unanticipated inflation can create a negative relationship between price dispersion and inflation. He notes that in markets where islanders act as buyers, incomplete information about inflation will mean that buyers do not know whether they have drawn an overpriced seller or whether the good itself has risen in line with the general price level. If buyers have a downwardly biased price distribution because they are unaware of the general inflation, their reservation price will be too low in relation to the true price, and they will engage in more extensive search, so reducing the price dispersion. Yet even this argument is questioned by an earlier literature that proposes that price dispersion might *not* fall if inflation erodes consumers information (Stigler and Kindahl, 1970) or if the desired stock of price information buyers wish to hold falls with inflation (van Hoomissen, 1988). Thus, even the most influential of the models of price setting behaviour can be used to derive both positive and negative relationships between relative price variability and inflation.

The same level of disagreement can be shown using models of price setting in the presence of menu costs. The optimal response to a fixed cost of price adjustment is to set prices discontinuously, such that the price is adjusted sufficiently often to offset the lost profits as the price deviates from the optimal price. So the optimal price represents a target and the maximum deviation that can be tolerated is the threshold in an (S,s) pricing rule (see Sheshinski and Weiss, 1983). This model can generate a positive association between the inflation and price dispersion, provided there is a positive rate of inflation that takes the price to an upper threshold. But if there is the possibility of falling prices and a lower threshold, the model does not generate a positive relationship. Furthermore, 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, 1984); 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.

Danziger (1987) provides the clearest case of ambiguity as he explores the implications of a linear quadratic price-setting model based on Rotemberg (1983). In this model, it is assumed that each firm $i$, $i \in [0,1]$ faces a demand function for its product, $Q_{it}$, that depends on the price of the product over the aggregate price level ($P_{it}/P_t$), where $P_t = \exp\left\{\int_0^1 P_{it} di\right\}$, and the real money supply ($M_t/P_t$):

$$(1) \qquad Q_{it} = \left[\left(\frac{P_{it}}{P_t}\right)^{-g}\left(\frac{M_t}{P_t}\right)^{f}\right]$$

If costs are quadratic (so $C = \frac{1}{2} U P_t Q_{it}^2$), then for a constant U > 0 the firm's profit maximisation leads to a desired price in each period of the form:

$$(2) \qquad P_{it}^{*} = \left(\frac{gU}{g-1}\right)^{1/(1+g)}\left(\frac{M_t}{P_t}\right)^{f/(1+g)} P_t$$

Since price adjustment is assumed to be costly, the firm evaluates the fixed cost, $A$, of maintaining the price at $P_{it}$ rather than making continuous adjustments of $P_{it}$ to $P_{it}$*. The instantaneous loss in profits from keeping prices fixed are assumed to be approximately quadratic and are written as $X = B(P_{it} - P_{it}*)^2$, where $B = k(M_t /P_t)^{2f/(1+g)}$ for a constant $k > 0$. Rotemberg shows that if the money stock grows at rate $P$, then this will also be the rate at which $P_{it}$* grows. This means that for firms that are uniformly distributed over the time of their last price change, each firm will adjust prices after an interval of time equal to $T$:

(3) $$T = a(\Pi)^{-2/3}$$

where $a = (6A/B)^{1/3}$, and will set prices at $P_{it} = P^*_{i0}\exp(\Pi T/2)$ starting at $t = 0$.[1] Danziger (1987) then calculates a measure of price dispersion:

(4) $$V = \int_0^1 w_i \left(m_i - m\right)^2 di$$

where $w_i$ is the expenditure share on product $i$ such that $\int_0^1 w_i di = 1$, $m_i$ is the rate of change of prices between observations for the individual firm's product taken at intervals, $b$[2]. The variable $m$ is the change of prices for the aggregate price level, and $m = \ln P_{t+b} - \ln P_t = \Pi b$.

Since equation (3) implies that the length of the adjustment period falls with inflation, it can be shown that there exists a level of inflation for which the adjustment period and the observation period coincide.[3] This is defined as $\overline{\Pi} = (\frac{a}{g})^{3/2}$, so $\Pi \leq \overline{\Pi} \Leftrightarrow b \leq T$ and $\Pi > \overline{\Pi} \Leftrightarrow b > T$. The proportion of firms that adjust their price in the first case is $b/T$ hence the measure of dispersion (assuming that firms that adjust prices do so by $PT$) is written as:

[1] The appendix to Fielding and Mizen (2000) offers linear and quadratic special cases of the Rotemberg model to motivate price-setting behaviour.

[2] In practice the expenditure shares are equal so the weights are uniformly set to unity. $b$ is not necessarily equal to $T$.

[3] Cecchetti (1985), Carlton (1986), Blinder (1991), Lach and Tsiddon (1992), Tomassi (1993) and Kashyap (1994) all show that the duration of price quotations falls during inflationary periods.

$$
\begin{aligned}
V &= (\Pi T - \Pi b)^2 b/T + (0 - \Pi b)^2 (1 - b/T) \\
&= \Pi^2 b(T - b) \qquad (5)\\
&= a^2 (\Pi^{4/3} \overline{\Pi}^{-2/3} - \Pi^2 \overline{\Pi}^{-4/3})
\end{aligned}
$$

This means that V = 0 for $P$ = 0 and $\overline{\Pi}$, and in all other cases between 0 and $\overline{\Pi}$, V > 0, because some firms adjust prices and others do not. Danziger shows that because the relationship in equation (5) is quadratic with a first derivative that is positive for $\Pi \in (0,(2/3)^{3/2}\overline{\Pi})$, zero at $\Pi = (2/3)^{3/2}\overline{\Pi}$ and negative for $\Pi \in ((2/3)^{3/2}\overline{\Pi},\overline{\Pi})$. The empirical estimates of the relationship between V and $P$ will depend on where our observations fall in the range of values for $P$. Since inflation is not independent of the frequency of observation, *b*, relative to the timing of price changes, *T*, it will mean that our observed relationship will be affected by the frequency of the data. It can be shown that $\overline{\Pi}$ depends on *b* in such a way that a higher frequency of observation leads to a higher value for $\overline{\Pi}$, and with it a larger range for which $\Pi \leq \overline{\Pi}$. The higher the frequency of the data the longer the range of inflation values for which the price dispersion measure is rising (because the value of V turns down at a higher level of inflation as the frequency increases). Hence, for a given range of values of inflation, a high frequency measure might detect a uniquely positive slope while a lower frequency of observation would reveal both positive and negative slopes over the same range. In this case $\Pi \in (0,(2/3)^{3/2}\overline{\Pi})$ if frequency of adjustment is slower than 18 months for annual data, 4.5 months for quarterly data or 1.5 months for monthly data, and the detected relationship will be positive, but if price adjustments occur more frequently than that the relationship will be negative because $\Pi \in ((2/3)^{3/2}\overline{\Pi},\overline{\Pi})$.

## 2.2 Measurement

Almost all the papers written in the 1970s, based on annual time-series or cross-sectional measures of variability (usually an unconditional standard deviation), were prone to overstate the positive relationship with inflation due to the effects of aggregation. Aggregation can potentially hide the true degree of variability in the data and allows common third causes, such as global supply shocks, to exert considerable leverage in a regression of inflation variability on the inflation level. The time-series relationship has also been shown to be sensitive to the sample period chosen. For example, Fischer (1981) shows that results are crucially dependent on the impact of particular shocks to the food and energy components of the price series; and Driffill *et al.* (1990) reach a similar conclusion regarding the first oil crisis. Later studies suffer less from aggregation effects because they use higher frequency time-series data; and increases in the number of commodities covered have facilitated panel estimates conditioned on time and group effects (for example Parsley, 1986; Debelle and Lamont, 1997). Steady improvements in data quality and coverage have allowed researchers to use higher frequency data (quarterly and monthly rather than annual data), and greater care has been taken over the measures of inflation variability, which now span commodities, time and cities/countries (Parks, 1978; Lach and Tsiddon, 1992; Parsley, 1996; Debelle and Lamont, 1997; Fielding and Mizen, 2000). They have also facilitated the application of more sophisticated processes to remove trend behaviour in time series (Fielding and Mizen, 2000). With measurement across cities or countries, as well as commodities, the relationship between inflation and price dispersion can be assessed within product groups (across cities/countries) and within countries (across products).

Even though later papers condition on time and group effects, eliminating some of the statistical reasons for a positive bias, there may still be reasons to question whether a positive coefficient estimate is an *economic* phenomenon. Hartman (1991) notes that the variables used in these studies are dependent on the same random disturbances and these can generate sign patterns irrespective of economic behaviour. It may be that these common shocks generate certain functional relationships between measures of actual inflation, expected or unexpected inflation and price dispersion irrespective of the underlying economic behaviour. Hartman considers a proportional rate of change in the price series $\ln(p_{ijt} / p_{ijt-1}) = m_j + S_{jt}$ driven by stationary random variables $S_{1t}$, …, $S_{Nt}$ distributed with a joint normal distribution, with a zero mean $E[S_{it}] = 0$ and a covariance independent of time $E[S_{jt}, S_{kt}] = t_{jk}$ $j,k = 1, \ldots, N$. Since $DP_t = \sum_j w_j \left(\ln\left(\frac{p_{ijt}}{p_{ijt-1}}\right)\right) = m_{DP} + \sum_j w_j S_{jt}$ where $m_{DP}$ is the average price change, and $V_t = \sum_i w_i \left(\ln(\frac{p_{ijt}}{p_{ijt-1}}) - DP_t\right)^2 = \sum_i w_i \left(q_i - \sum_i f_{ij} S_{jt}\right)^2$ both variables are all driven by the same random processes. The only assumptions that are necessary to create a functional relationship between these measures of inflation and inflation variability are that they are drive by the same random variables with stationary properties, jointly normally distributed and with weights, $w_i$, that are constant.

Using the normal equations we can calculate the expected signs of the coefficients for regressions of the type:

(6a) $V_t = a_0 + a_1 DP_t + a_2 DP_t^2$

(6b) $V_t = b_0 + b_1 DP_t + b_2 |DP_t|$

These regressions have expected coefficient values

(7a) $a_2 = \sum_j w_i f_{DPi}^2$

(7b) $b_2 = a_2 v_{DP}$

where $f_{DPi} = \dfrac{\sum_k w_k s_{ik} - s_{DP}^2}{s_{DP}^2}$, $s_{DP}^2 = \sum_j \sum_k w_j w_k s_{jk}$ and

$$v_{DP} = \frac{s_{DP}^2.\text{cov}(DP^2, |DP|) - \text{cov}(DP, DP^2).\text{cov}(DP, |DP|)}{s_{DP}^2.\text{var}(|DP|) - [\text{cov}(Dp, |DP|)]^2}$$

Since $v_{DP} > 0$ we find both $a_2$ and $b_2$ are expected to be positive; although $a_1$ and $b_1$ can take either sign. It is noteworthy that for a linear regression, $V_t = d_0 + d_1 DP_t$, the same calculation implies that the $sgn(d_1) = sgn(cov(V, DP)) = sgn\ (2\sum_i w_i (m_i - m_{DP}) f_{DP} s_{DP}^2)$, which is ambiguous. Thus under conditions of joint normality we cannot sign the coefficient in the linear model even though we can sign higher order terms when a nonlinear functional form is adopted.

Few of the empirical studies have found the distributions to be normal, however, so the this argument is considerably weakened. Vining and Elwertowski (1976) take over 1000 goods from the US Bureau of Labour Statistics 1948-74 and reject normality strongly. Mizon *et al.* (1990) examine the normality assumption on monthly observations of 37 categories of goods in the UK 1962-87, showing that price distributions exhibit skewness and kurtosis that lead to overwhelming rejection of the normality assumptions. These studies both show that price data are positively skewed during periods of rising inflation and negatively skewed during disinflationary periods. When prices are rising a few commodities experience price changes considerably above the mean, while the majority lie just below the mean; when prices are falling the reverse is true.

The force of Hartman’s argument lies in the fact that no assumptions about misperceptions of inflation, rigidities or asymmetries are necessary to prove that nonlinear models will be supported provided the data are stationary, normally distributed and the weights are constant. A further prediction of the model is that both quadratic and piecewise-linear models will be supported against linear alternatives under these assumptions, but there is no sense in which one is more probable than another. Evidence in favour of one functional form over another from cross-plots is not particularly illuminating, because since differentiation between alternative nonlinear forms one is likely to be dominated by a small number of outliers. In such a case, the power to discriminate between, say, a quadratic and a piece-wise linear model is likely to be low.

The theory and measurement issues highlight several important conclusions. First, there is not the level of agreement in theory or in empirical studies that is commonly supposed in favour of a positive relationship between relative price variability and inflation. Second, the frequency of the data may influence the nature of the relationship we discover. Third, many of the supportive results can be attributed to aggregation bias and to the construction of the measures themselves as functions of common shocks. These conclusions give some indication of the methodological approach we should take.

First, given the theoretical possibility of either a linear or a nonlinear functional form, we should adopt a methodology that allows the data to choose the functional form rather than impose a form from one of the many available theories. We could then infer which of the theories is supported by the data, since the Lucas-Reinsdorf models imply linear models with positive (negative) coefficients, while the Danziger-Rotemberg model implies a nonlinear model with positive or negative coefficients over different ranges of inflation. Second, if the frequency of observations matters in determining the relationship between price dispersion and inflation, the use of a common data set at different frequencies will determine whether, as Danziger conjectures, the relationship may turn on the relative frequency of observation to the frequency of price setting. Third, we can slice our panel of data in each of two directions, calculating the relative price variability and inflation across countries for the same commodity (in turn), or taking the same measures across commodities for each country sequentially. We can then infer whether the relationship between these variables is consistent, and if not, whether the differences tell us anything about the price setting process. We address these issues in the next section.

## 3. *Methodology*

In order not to make any *a priori* assumptions about the functional form of the relationship between relative price variability and inflation, we will estimate the relationship using the semi-parametric approach described by Robinson (1988) and Härdle (1990, chapter 9.1), using data are from the *Eurostat* database over the period 1983(1)–1994(12). The *Eurostat* data are reported monthly, and this monthly data can be aggregated to create a corresponding quarterly sample. We will be using inflation variability rather than price level variability figures, because we wish to ensure that all of our measures are stationary.[4] The regression equations are the same for both our monthly and our quarterly samples:

(8) $V^1_{it} = b_1 \times V^1_{it-1} + \mathrm{m}^1(p^1_{it}) + u_{it}$

(9) $V^2_{jt} = b_2 \times V^2_{jt-1} + \mathrm{m}^2(p^2_{jt}) + u_{jt}$

where $p^1_{it}$ is a weighted average of deseasonalised inflation rates (in ecus) across the seven countries for the $i^{th}$ product category in month or quarter $t$ and $p^2_{jt}$ is a weighted average of deseasonalised inflation rates (in ecus) across the nine product categories for the $j^{th}$ country in month or quarter $t$.[5] $V^1_{it}$ is a weighted variance of national inflation rates around $p^1_{it}$ for each product category and $V^2_{jt}$ is a weighted variance of product-specific inflation rates around $p^2_{jt}$ for each country. Preliminary regressions of each of these four variables on time dummies and product category or country dummies (time and group effects) are constructed, so that all variables used in the estimation of equations (8-9) are orthogonal to fixed effects. $\mathrm{m}^1(.)$ and $\mathrm{m}^2(.)$ are non-linear functions to be estimated non-parametrically; $u_{it}$ and $u_{jt}$ are residuals. The $b$ parameters allow for the possibility of some persistence in the variability measures we have constructed.

Figures 1-2 illustrate the inflation variability figures. Figure 1 shows histograms representing the frequency of $V^1_{it}$ and $V^2_{jt}$ in both the monthly and the quarterly samples, and Figure 2 shows corresponding QQ plots against a normal distribution. In all cases there are very large tails in the distributions, which are far from normal. This is consistent with the findings of Vining and Elwertowski (1976) and Mizon *et al.* (1990). The same is true of the average inflation rates and, as the scatter-plots in Figure 3 indicate, the outliers are likely to dominate and bias any regression technique that relies on minimising squared residuals without some attempt to modify the influence of the tails of the

[4] Another reason for using inflation rather than price levels is that price levels in our data set are measured in indexed form.

[5] The product categories are: clothes, drink, food, footwear, fuel and rent, furniture and household appliances, recreational goods and services, tobacco and transport and communication; and the countries are: Belgium, Denmark, France, Germany, Italy, the Netherlands and the UK. The weights are based on the level of consumption (in ecus) of each product category in each country in the year in which period $t$ falls.

distribution. The visual evidence in Figure 3 does seem to suggest that there is potentially some nonlinearity in the relationship between inflation and relative price variability within commodities. However, it is impossible to determine from visual inspection whether the relationship is quadratic or piece-wise linear. Nor is it likely that a parametric approach would have great power to discriminate between nonlinear alternatives.

The parameters in equations (8-9) are estimated in several stages. First, the $b$ parameters are estimated by constructing the following regression equations:

(10) $V^1_{it} = b_1 \cdot V^{1\sim}_{it-1} + e^1_{it}$

(11) $V^2_{jt} = b_2 \cdot V^{2\sim}_{jt-1} + e^2_{jt}$

$V^{1\sim}_{it-1}$ and $V^{2\sim}_{jt-1}$ are residuals from non-parametric regressions of $V^1_{it-1}$ and $V^2_{jt-1}$ on $p^1_{it}$ and $p^2_{jt}$ respectively:

(12) $V^1_{it-1} = \mathrm{n}^1(p^1_{it}) + V^{1\sim}_{it-1}$

(13) $V^2_{jt-1} = \mathrm{n}^2(p^2_{jt}) + V^{2\sim}_{jt-1}$

Then the functions $\mathrm{m}^1(.)$ and $\mathrm{m}^2(.)$ are estimated using the residuals $e^1_{it}$ and $e^2_{it}$:

(14) $e^1_{it} = \mathrm{m}^1(p^1_{it}) + u_{it}$

(15) $e^2_{jt} = \mathrm{m}^2(p^2_{jt}) + u_{jt}$

Among the very wide class of kernel density functions that could be used for the non-parametric regressions, we choose the density function proposed and motivated by Fan (1992, 1993). Aside from its other advantages, this is a truncated kernel density function that is robust to the existence of outliers. The point-derivatives of each non-parametric function are estimated for any given value of $p^1$ or $p^2$ ($p^1_0$, $p^2_0$) by a linear weighted least squares regression with weights equal to:

(16) $w_{it} = \frac{15}{16}\left[1 - \left[\frac{p^1_0 - p^1_{it}}{h}\right]^2\right]^2$ if $|p^1_0 - p^1_{it}| = h$ and $w_{it} = 0$ else

$$(17) \quad w_{jt} = \frac{15}{16}\left[1 - \left[\frac{p_0^2 - p_{jt}^2}{h}\right]^2\right]^2 \text{ if } |p_0^2 - p_{jt}^2| = h \text{ and } w_{jt} = 0 \text{ else}$$

where $h$ is a bandwidth parameter which we will select using a mean squared forecast error criterion.[6] Corresponding standard errors will be estimated by bootstrapping.

Tests of the significance of deviations in the data from linear, piecewise-linear and quadratic functional forms will be conducted by estimating nonparametric regressions of the form:

$$(18) \quad v^1_{it} = \mathrm{m}^1(p^1_{it}) + u_{it}$$

$$(19) \quad v^2_{jt} = \mathrm{m}^2(p^2_{jt}) + u_{jt}$$

where $v^1_{it}$ ($v^2_{jt}$) represents the series of residuals from linear (or piecewise-linear, or quadratic) regressions of $V^1_{it}$ ($V^2_{jt}$) on $p^1_{it}$ ($p^2_{jt}$). Under the null of no deviations, fitted values of $v^1_{it}$ ($v^2_{jt}$) should equal zero at every observation. As a test of the null, we will test whether the sum of the fitted absolute or squared values of $v^1_{it}$ ($v^2_{jt}$) is significantly different from zero, again using a bootstrap.

## 4 Results

Table 2 reports several statistics relating to estimation of the relationship between price dispersion and inflation for our two variability measures ($V^1_{it}$, $V^2_{jt}$) at both monthly and quarterly frequencies. (For $V^1$ there are 1,395 observations in the monthly data and 459 observations in the quarterly data; for $V^2$ there are 1,085 and 357 observations respectively.) For each of the four models, we report mean squared forecast errors for a selection of bandwidth values around the one that minimises the forecast error, and corresponding values of estimated $b$. The estimates of $b$ are not sensitive to the value of $h$ chosen; they indicate significant persistence at the monthly frequency, but not at the quarterly frequency.

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[6] The criterion is constructed as follows. The model is estimated over (n – 1) years of the sample n times over (each time omitting a different year), and forecasts of the n$^{th}$ year are made on the basis of these estimates. The mean value of all the forecast errors for every year is then calculated, ignoring observations where $p$ is more than two standard deviations from its mean (so that outliers do not bias the shape of the function). The value of $h$ is chosen so as to minimise this criterion.

For each of the four models, Table 3 reports tests of the null that the model is linear. These are constructed in the way described in the previous section, using 1,000 bootstrap samples drawn from the distribution of the estimated residuals $u_{it}$ ($u_{jt}$) in equations (18-19) at both monthly and quarterly frequency. In these tests, the bandwidth chosen is the one selected according to the forecast error criterion reported in Table 2. For $V^1$, the null of linearity can be rejected at the 1% level at both frequencies, whether the test is based on fitted absolute or fitted squared residuals. This is an important result because it supports the theoretical model that suggests the behaviour of price setters for the same commodity will generate a quadratic relationship between relative price variability and inflation. The results for each product across countries are consistent with the prediction of the Danziger-Rotemberg model when firms set prices at intervals, $T$, and incur losses from allowing prices to deviate from the optimal value for a short duration[7]. It is perhaps more likely that within product data will conform to this model, since the relevant decision for a producer of a commodity is the position that his commodity price occupies in the distribution relative to the average price. Adjustment will depend on the cost of adjusting the price relative to the gain in profits to be made from so doing, as Danziger (1987) and Rotemberg (1983) have described.

For $V^2$, the null of linearity can be rejected at neither frequency at the 5% level. This is not entirely surprising: it is not obvious that the producers of different products within the same country will behave as Danziger and Rotemberg suggest, because the within-country distribution reflects the average price of different products with different adjustment speeds. Since there is only one price for each product in each country, a price setter cannot simultaneously set a price with reference to the within-product distribution and at the same time refer to the within-country distribution. Indeed, the strong result in favour of a nonlinear Danziger-Rotemberg model using the within-commodities data makes it less likely that adjustment within countries will lead to a nonlinear relationship.

We turn now to the question of the exact functional form of the relationship between relative price variability within-commodity across-countries. Since the null of linearity can be rejected for $V^1$, we test the null that the relationship with inflation is quadratic, and the null that it is piecewise linear[8]. In the latter case, we allow for a linear slope that changes its value discretely at some endogenously determined value of inflation; the breakpoint is

[7] The exact shape of the nonlinear function is considered below and conforms to the quadratic shape suggested by Danziger (1987).

[8] We do not consider the question for $V^2$ since the null of linearity could not be rejected.

the one that minimises the squared residuals. The quadratic and piecewise linear test statistics are also reported in Table 3. Using the quarterly data, neither piecewise linear nor quadratic functional forms can be rejected; however, the small sample size means that these tests have lower power than those based on the monthly data. Using monthly data, the piecewise linear model can be rejected at the 10% level, but the quadratic model cannot. There is therefore some evidence in favour of a quadratic functional form.

The piecewise-linear model is generated by the response of relative price variability to the absolute value of inflation, creating some V-shaped function, while the quadratic model is the response of relative price variability to squared inflation terms, which gives a U-shaped distribution. There is not a great deal of difference between the two. Both imply symmetric responses to rising and falling inflation on the part of the relative price variability measure, in the first case the weight on rising and falling inflation is constant, while in the latter the weight varies proportionally with inflation itself. Debelle and Lamont (1997) explain that such a result might be found within countries (in their case it was city-by-city data) if a positive (negative) demand shock caused non-traded goods prices to rise (fall) while traded goods prices remain at the international level, determined by arbitrage. The increase in within country dispersion for both positive and negative demand shocks would create the V-shaped relationship. In our case, however, the argument does not apply because we are considering the relationship between relative price variability and inflation within products across countries. While we can think of the demand shock applying to the specific product, we cannot consider the commodity in one country as more or less tradable than the same commodity in another country. Rather we suggest that our evidence offers further support for the Danziger-Rotemberg argument that firms set prices in the context of a cost of adjustment, creating a quadratic relationship between relative price variability and inflation within commodities. In fact, the case that corresponds to the Debelle-Lamont example does not support a non-linear model, since both the monthly and the quarterly versions of the model, estimated $\partial \mathrm{m}^2/\partial p^2{}_0$ insignificantly different from zero at all values of $p^2{}_0$.

In the final stage of the analysis we consider the estimates of the nonparametric functions. This is a redundant concept in the case of the $V^2$ model because there is no significant nonlinearly, (so we do not discuss the shape of the nonparametrically estimated $\mathrm{m}^2(.)$ function) but Figures 4-7 illustrate the estimates of the $\mathrm{m}^1(.)$ functions in the monthly and quarterly versions of the $V^1$ model. The figures show the derivatives of the functions for different values of $p^1{}_0$ from two standard deviations below the mean value of $p^1$ to two standard deviations above.

Figure 4 shows different estimates of the derivative of the function using the monthly data with alternative bandwidths around the one that minimises the mean squared forecast error ($h$ = 0.01). Changing the value of $h$ does not affect the general shape of the function, which is downward-sloping and close to zero at two standard deviations above the mean inflation rate; but smaller bandwidths make the derivate of the function steeper (in other words, they increase the curvature of the function). The figure indicates that the variability of inflation rates across countries is a positive function of the mean inflation rate in each product category, but that the effect is decreasing in the inflation rate, so that at very high values of inflation an incremental increase in the rate has no significant effect on the variability. Figure 5 provides more information on the level of significance of the point estimates of $\partial m^1/\partial p^1_0$. It shows the estimate of the function with $h$ = 0.01 along with a line depicting points two standard errors above the point estimate and another depicting points two standard errors below. The error bars are based on 1,000 bootstrap replications drawing from the distribution of the estimated $u_{i\ t}$. $\partial m^1/\partial p^1_0$ is significantly positive up to an inflation rate about 1.75 standard deviations above the sample mean.

Figures 6-7 illustrate the results using the quarterly data, for which the forecast error criterion is minimised at $h$ = 0.005. Figure 6 shows that the overall shape of the function estimated on quarterly data is not affected by the choice of $h$, but that for very small bandwidths there are erratic changes in its slope. The most notable feature of the quarterly model, indicated by Figure 7, is that the derivative is significantly positive when inflation is more than about 1.25 standard deviations below its sample mean and significantly negative for values of inflation above this mean. In other words, at very low values of average inflation a marginal increase in the average inflation rate corresponds to an increase in the variability of inflation across countries; but at higher rates a marginal increase is associated with lower variability. This would be also consistent with a quadratic model of the kind suggested by Danziger (1987) since there is predicted to be a range of inflation rates which leads to increasing variation in relative prices and a range of higher values where variability falls. The implicit level of inflation at which the turning point is located varies across commodities, because the estimated model is based on de-meaned data. Using the monthly data, the monthly inflation rates at the turning point of the function (two standard deviations above the mean) range from 1.12% (food) to 1.27% (tobacco). Using the quarterly data, the range of quarterly inflation rates at the turning point (1.25 standard deviations below the mean) is from 0.40% (food) to 0.85% (tobacco).

## 5 Conclusions

This paper has shown that the theoretical and empirical support in the literature for a positive linear relationship between relative price variability and inflation is less clear-cut than is commonly supposed. The theoretical literature provides supporting arguments for *both* positive and negative linear relationships, *and* for linear and nonlinear functional forms. The empirical evidence is also inconclusive since aggregation and measurement can bias the findings in favour of a positive correlation between the two variables.

We reinvestigate the issue using a non-parametric methodology that does not specify the nature of the functional relationship in advance, allowing us to discriminate between the alternatives without imposing a particular result in advance. We find strong support for a quadratic model against linear and piecewise linear alternatives within commodities across countries and a linear relationship in the corresponding relationship across commodities within countries. Given that there is a single price for each commodity per country at any one point in time, we can only support a single theory of price setting, despite the fact that in our empirical work we slice the data in two directions in order to consider two alternatives. The alternatives are that *either* the firm sets its price in relation to the distribution of the prices of the same commodity in the other countries, *or* it sets prices with reference to the distribution of prices within its own country (i.e. taking account of the distribution of the general price made up of many different commodity prices). Our results are supportive of the former. For European data at monthly and quarterly frequencies over the period of the Single Market programme (1983 – 1994), prices are set with respect to the distribution of prices of the same commodity across countries rather than with reference to the distribution of prices of within a country. The findings are supportive of a number of the theoretical predictions made by Danziger (1987), since the preferred relationship is quadratic (not linear or piecewise linear) and it is sensitive to the observation frequency of the data, in the sense that the relationship between price dispersion and inflation shifts with the frequency of observation. These results confirm Dabziger's theoretical point that it is possible that the dispute in the empirical literature may have arisen as a result of the choice of data frequency and the range of values of inflation within the sample used by different studies.

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## Table 1: Previous Studies of Relative Price Variability and Inflation

*(i) Annual Studies*

| *Author* | *Variables* | *Functional Form* | *Data Sample and Coverage* | *Methodology* | *Findings* |
|---|---|---|---|---|---|
| Okun (1971) | RPV1, DP | None | 1951-1968<br>17 OECD countries | Comparison of sample moments | Positive association |
| Logue and Willett (1976) | RPV1, DP | Linear: RPV, average inflation for class | 1950 – 1970<br>41 industrialised and less developed countries | Pooled regressions based on classified data e.g. industrialised, Latin American and other countries | Positive coefficient on inflation for whole sample, differs between classifications. |
| Vining and Elwertowski (1976) | RPV1, DP | None | Over 1000 goods from Bureau of Labour Statistics,<br>US 1948 – 74 | Comparison of sample moments | Positive association |
| Parks (1978) | RPV4, DP | Nonlinear: RPV, absolute value of DP, DP squared, positive/negative values of DP, DP squared, expected and unexpected DP | US 1929 – 1975, 12 consumer goods<br><br>Netherlands 1921- 1963 16 consumer goods | Time series, OLS. | Positive coefficient on squared inflation and squared unexpected inflation. Relationship stronger for declining prices than for rising prices. |
| Blejer and Leiderman (1980) | RPV4, DP | Linear: RPV and DP | Mexico, 1951 – 1976 | Time series, OLS | Weak positive coefficient on unanticipated inflation |
| Ball and Mankiw (1994) | DP, RPV1, RPV3 | Linear DP and RPV | US 1949 – 1989 | Time series, OLS | Positive, some evidence of asymmetry. Note inflation is regressed on RPV, lagged inflation etc not the other way around. |

*(ii) Quarterly Studies*

| *Author* | *Variables* | *Functional Form* | *Data Sample and Coverage* | *Methodology* | *Findings* |
|---|---|---|---|---|---|
| Domberger (1987) | RPV1, DP | Nonlinear: RPV, DP squared | UK 1974(1) – 1984(4) 80 goods categories from SIC | Time series, SURE, OLS and GLS | Positive coefficient on squared term |
| Fischer (1982) | | | Germany 1969 – 1980<br>US 1948 – 1980 | | |
| Stockton (1988) | RPV2, DP | Linear: DP , RPV | US 1949(3) – 1980(3)<br>91 sectors | Time series, OLS | Positive coefficient when DP regressed on RPV. |
| Hartman (1991) | RPV2, DP | Nonlinear: RPV, DP, square and absolute values of DP, expected and unexpected DP | US data 1973(1) – 1990(1) split into two subsamples 1973(1) –1981(4) and 1982(1) – 1990(1). Data for durables, nondurables and services | Time series, OLS | Negative sign on linear term, DP, but positive sign on square and absolute value of DP. This result was robust to expected and unexpected measures of inflation and the change in inflation. |

*(iii) Monthly Studies*

| *Author* | *Variables* | *Functional Form* | *Data Sample and Coverage* | *Methodology* | *Findings* |
|---|---|---|---|---|---|
| Hercowitz (1981) | RPV1, money growth, variance of money growth | Linear: RPV, money growth, variance of money growth | Germany, 1921(2) – 1923 (7) 60 commodities | Time series, OLS, Cochrane-Orcutt | Positive coefficient on money growth but no significant coefficient on the variance of money growth. |
| Ashley (1981) | RPV1, DP | Linear: RPV and DP | US, 1953(1) –1975(6) 20 commodity groups | Time Series: Box-Jenkins approach | Evidence exists for a positive (negative) relationship at lag 2 (12) |
| Cukierman and Leidermann (1984) | RPV4, unexpected money growth | Linear: RPV and unexpected money growth | Israel, 1966(10) – 1980(12) 23 components of CPI | Time series, maximum likelihood estimation | Positive effect of unanticipated monetary growth on RPV. Some prices are controlled during the sample. |
| Danziger (1987) | RPV1, DP | Specific nonlinear functional form: $RPV1 = a + bDP^{4/3} + cDP^{2}$ | Israel 1968 – 1983100g of Kosher Salami | Time series, OLS | Positive coefficient on the inflation term and negative on the quadratic term. The net effect depends on the level of inflation |
| Van Hoomissen (1988) | RPV1, DP | Nonlinear RPV and Square of DP | Israel, 1971(1) – 1984(10) 13 food items | Time series, SURE Joint GLS, OLS | Positive coefficient on inflation in the level, negative on the squared term (expect for books and fridges). |
| Mizon et al. (1990) | RPV1, DP | None | UK 1962 – 1983 37 goods | none | Assessment of the moments of the distributions of each category of goods |
| Mizon (1991) | RPV3, DP | Linear: RPV, DP | UK 1965 – 1987 | Time series, OLS | Positive coefficient on the level of inflation. Efforts to remove the effects of oil shocks, on RPV and DP. |
| Lach and Tsiddon (1992) | RPV1, expected and unexpected DP | Nonlinear: RPV, level of inflation, absolute value of unexpected inflation | Israel 1978(1) – 1984 (9) 26 food products | Time series, SURE | Positive association |

| | | | | | |
|---|---|---|---|---|---|
| Reinsdorf (1994) | CV, DP | Nonlinear: expected and unexpected DP, absolute value of DP, square of DP | US 1980(1) – 1982(12)<br>9 cities<br>65 food products | Time series, Swamy random coefficient estimator | Negative in inflation during the Volcker disinflation. Positive in absolute value of inflation, and negative in inflation squared. Declines follow unexpected inflation, but expected inflation has a positive effect on RPV |
| Parsley (1996) | RPV1, RPV2, RPV3, RPV4, DP | Linear: RPV and DP | US 1975(1) – 1992(12)<br>48 cities 32 goods | Panel estimation with time and fixed effects | Positive coefficient on inflation for all measures |
| Debelle and Lamont (1997) | RPV1, DP (both with US average subtracted) | Nonlinear: RPV on DP, absolute value of DP, square of DP | Panel 1 US 1954 – 1986, 19 cities, 14 goods<br><br>Panel 2 US 1977- 1986, 24 cities, 18 goods | Panel estimation with time and fixed effects | Positive coefficient on inflation, absolute and quadratic terms are insignificant when both included, but absolute value has a positive effect when included without the squared term. |
| Fielding and Mizen (1999) | RPV1, RPV2, DP after removing smooth transition in the trends | Nonlinear: RPV on DP, square of DP, change in DP and change in square of DP | 10 EU countries in Single Market<br>15 products<br>1986(1) – 1993(12) | Time series: maximum likelihood estimation of smooth transition and nonlinear relationship across products and countries separately | Negative association after the smooth transition in the trend has been removed within product groups. Fast decay in response to shocks around the logistic trend. |

Notes to Table 1

$$\text{RPV1} = \left[ \frac{1}{n-1} \left( \sum_j DP_{ijt} - DP_{i.t} \right)^2 \right]^{1/2}$$

where $DP_{ijt} = \ln p_{ijt} - \ln p_{ijt-1}$ and $DP_{i.t} = \frac{\sum_j DP_{ijt}}{n}$

$$\text{RPV2} = \left[ \frac{1}{n-1} \left( \sum_j R_{ijt} - R_{i.t} \right)^2 \right]^{1/2}$$

where $R_{ijt} = \ln \frac{p_{ijt}}{p_{i.t}}$, $R_{i.t} = \frac{\sum_j R_{ijt}}{n}$ and $p_{i.t} = \frac{\sum_j p_{ijt}}{n}$

$$\text{RPV3} = \left[ w_j \left( \sum_j R_{ijt} - R_{i.t} \right)^2 \right]^{1/2}$$

where $R_{ijt} = \ln \frac{p_{ijt}}{p_{.jt}}$, $R_{i.t} = \frac{\sum_i R_{ijt}}{m}$ and $p_{.jt} = \frac{\sum_i p_{ijt}}{m}$

$$\text{RPV4} = \left[ w_j \left( \sum_j DP_{ijt} - DP_{.jt} \right)^2 \right]^{1/2}$$

where $DP_{ijt} = \ln p_{ijt} - \ln p_{ijt-1}$ and $DP_{.jt} = \frac{\sum_i DP_{ijt}}{m}$

$$\text{CV} = \frac{1}{p_{ijt}} \left[ \frac{1}{n-1} \left( \sum_j p_{ijt} - p_{i.t} \right)^2 \right]^{1/2}$$

where $p_{i.t} = \frac{\sum_j p_{ijt}}{n}$.

## Table 2: Selected Statistics from Non-parametric Regression Estimates

Mean squared forecast errors and estimated *b*s for the $v1_{it}$ model (monthly)

| *h* | *m.s.f.e* (´ $10^{-6}$) | *b (s.e.)* |
|---|---|---|
| 0.005 | 7.8078 | 0.150218 (0.026613) |
| 0.006 | 7.7912 | 0.126091 (0.026756) |
| 0.007 | 7.7785 | 0.114914 (0.026793) |
| 0.008 | 7.7667 | 0.113383 (0.026786) |
| 0.009 | 7.7563 | 0.117005 (0.026767) |
| **0.010** | **7.7470** | **0.121516 (0.026746)** |
| 0.020 | 7.8123 | 0.126357 (0.026721) |
| 0.030 | 7.8328 | 0.121577 (0.026739) |
| 0.040 | 7.8484 | 0.122284 (0.026723) |
| 0.050 | 7.8635 | 0.122997 (0.026708) |
| 0.060 | 7.8513 | 0.124211 (0.026693) |
| 0.070 | 7.8871 | 0.125759 (0.026681) |
| 0.080 | 7.9370 | 0.127067 (0.026672) |
| 0.090 | 7.9886 | 0.128072 (0.026666) |
| 0.100 | 8.0353 | 0.128836 (0.026662) |

Mean squared forecast errors and estimated *b*s for the $v2_{jt}$ model (monthly)

| *h* | *m.s.f.e* (´ $10^{-6}$) | *b(s.e.)* |
|---|---|---|
| 0.010 | 9.2499 | 0.104371 (0.030309) |
| 0.020 | 9.2075 | 0.108296 (0.030370) |
| **0.030** | **9.2004** | **0.108841 (0.030351)** |
| 0.040 | 9.2016 | 0.109026 (0.030349) |
| 0.050 | 9.2030 | 0.109013 (0.030347) |
| 0.060 | 9.2030 | 0.108929 (0.030345) |
| 0.070 | 9.2026 | 0.108829 (0.030343) |
| 0.080 | 9.2024 | 0.108746 (0.030340) |
| 0.090 | 9.2026 | 0.108681 (0.030339) |
| 0.100 | 9.2028 | 0.108629 (0.030337) |

**Table 2: (continued)**

Mean squared forecast errors and estimated *b*s for the $v1_{it}$ model (Quarterly)

| *h* | *m.s.f.e* (´ $10^{-6}$) | *b(s.e.)* |
|---|---|---|
| 0.001 | 7.8779 | 0.063771 (0.047218) |
| 0.002 | 7.8623 | 0.079216 (0.046347) |
| 0.003 | 7.7979 | 0.051759 (0.047509) |
| 0.004 | 7.7829 | 0.031546 (0.047820) |
| **0.005** | **7.7793** | **0.019871 (0.047925)** |
| 0.006 | 7.7823 | 0.007694 (0.048066) |
| 0.007 | 7.7882 | -0.003197 (0.048081) |
| 0.008 | 7.7932 | -0.009987 (0.048027) |
| 0.009 | 7.7994 | -0.012711 (0.047979) |
| 0.010 | 7.8023 | -0.013410 (0.047949) |

Mean squared forecast errors and estimated *b*s for the $v2_{jt}$ model (Quarterly)

| *h* | *m.s.f.e* (´ $10^{-6}$) | *b (s.e.)* |
|---|---|---|
| 0.001 | 2.8250 | 0.006332 (.003862) |
| 0.002 | 2.4306 | 0.014385 (0.008546) |
| 0.003 | 2.4027 | 0.060513 (0.039500) |
| **0.004** | **2.3948** | **0.039695 (0.051840)** |
| 0.005 | 2.3953 | 0.029261 (0.052069) |
| 0.006 | 2.3957 | 0.019303 (0.051863) |
| 0.007 | 2.3976 | 0.017309 (0.051985) |
| 0.008 | 2.4005 | 0.017245 (0.052226) |
| 0.009 | 2.4026 | 0.016137 (0.052356) |
| 0.010 | 2.4036 | 0.014722 (0.052397) |

**Table 3: Tests of the Nonparametric $V^1$ and $V^2$ Models Against Linear, Quadratic and Piecewise-Linear Models**

*Figures indicate p-values corresponding to the null that the model takes a particular functional form.*

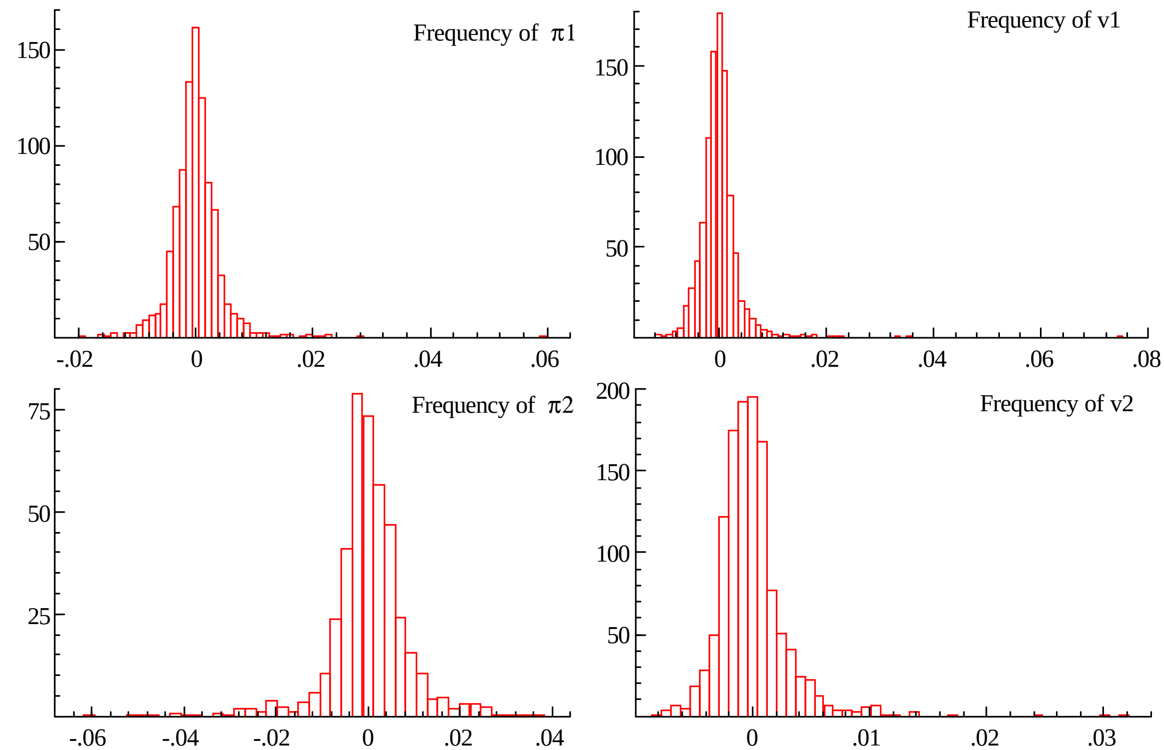
*(i) Using Fitted Squared Residuals*

| *Model* | Linear | Quadratic | Piecewise Linear |
|---|---|---|---|
| $V^1$ (monthly data) | 0.000 | 0.547 | 0.057 |
| $V^1$ (quarterly data) | 0.003 | 0.746 | 0.785 |
| $V^2$ (monthly data) | 0.784 | | |
| $V^2$ (quarterly data) | 0.068 | | |

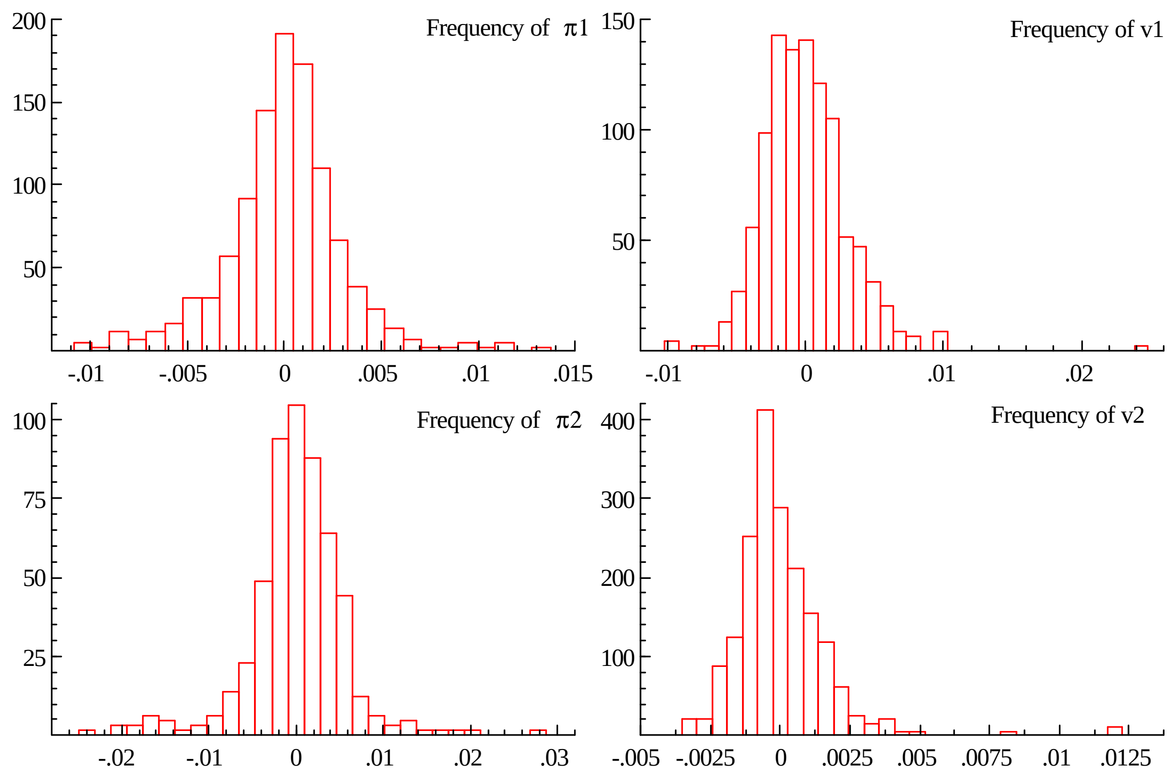
*(ii) Using Fitted Absolute Residuals*

| *Model* | Linear | Quadratic | Piecewise Linear |
|---|---|---|---|
| $V^1$ (monthly data) | 0.000 | 0.653 | 0.097 |
| $V^1$ (quarterly data) | 0.009 | 0.645 | 0.735 |
| $V^2$ (monthly data) | 0.858 | | |
| $V^2$ (quarterly data) | 0.415 | | |

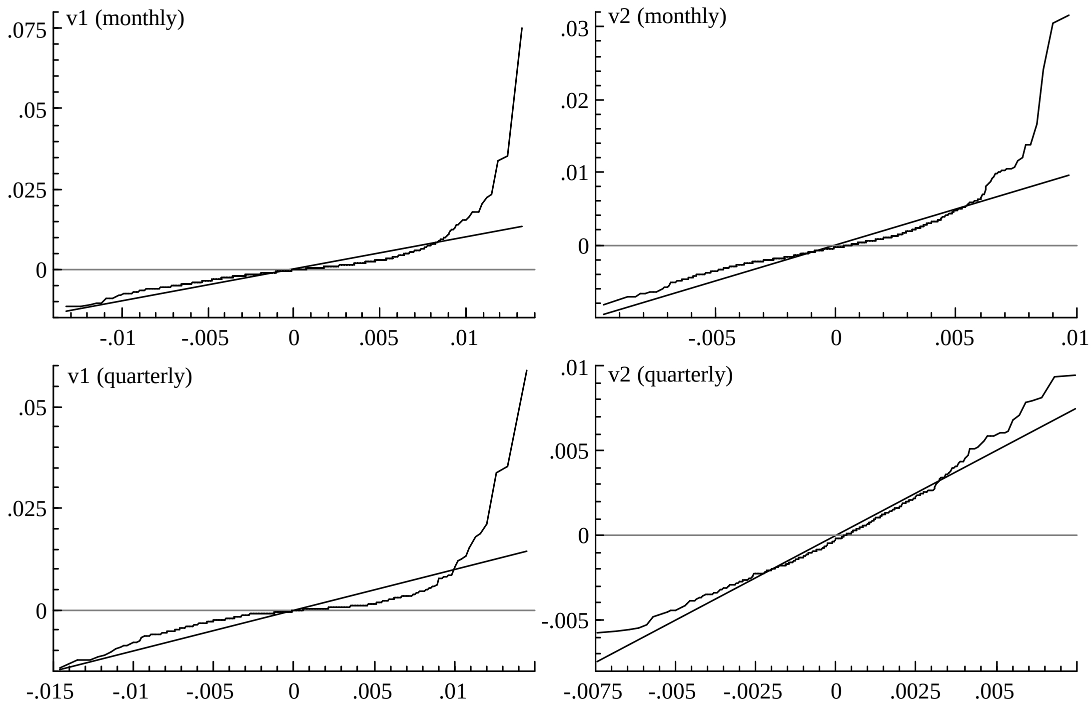
## Figure 1(a): Frequency distributions of the monthly variables




## Figure 1(b): Frequency distributions of the quarterly variables



## Figure 2: QQ plots of the distributions of $V^1$ and $V^2$ against a standard normal distribution




## Figure 3: Scatterplots of inflation variability (vertical axes) against average inflation (horizontal axes)

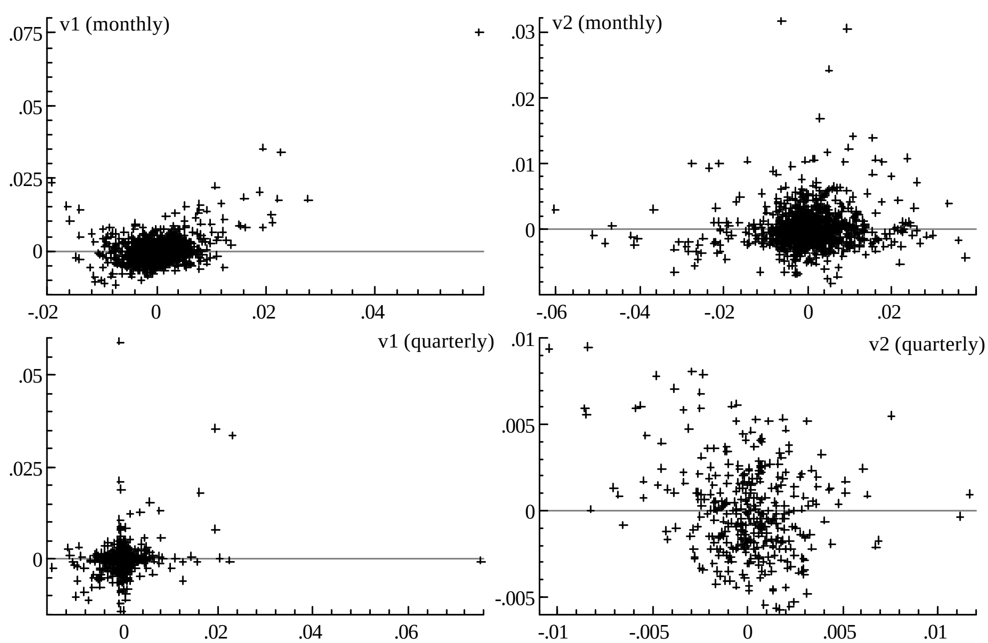


**Figure 4: Derivative of the $m^1$(.) function for different values of *h*, monthly data**

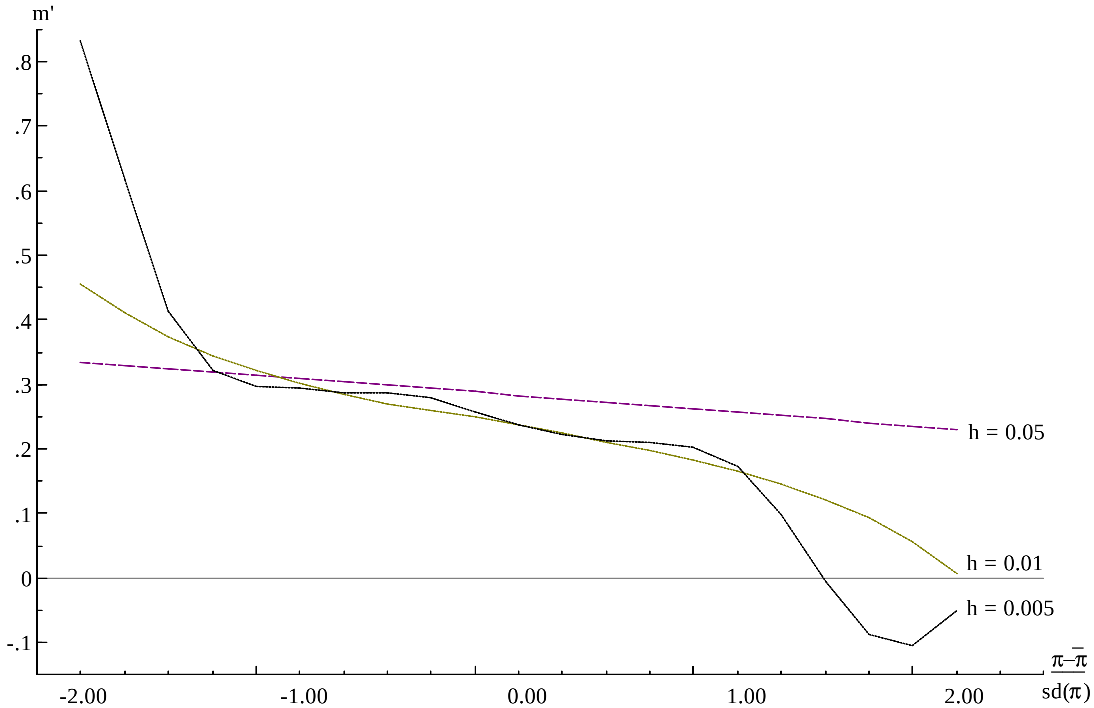



**Figure 5: Derivative of the $m^1$(.) function ± 2s.e. for *h* = 0.01, monthly data**

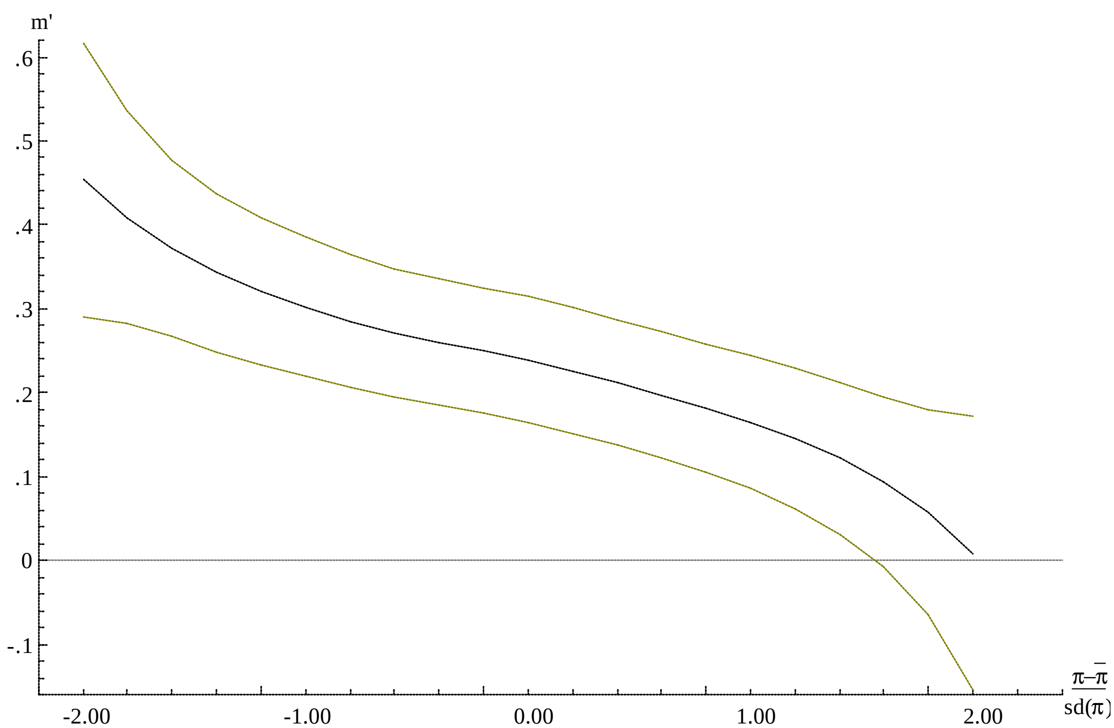


**Figure 6: Derivative of the $m^1$(.) function for different values of *h*, quarterly data**

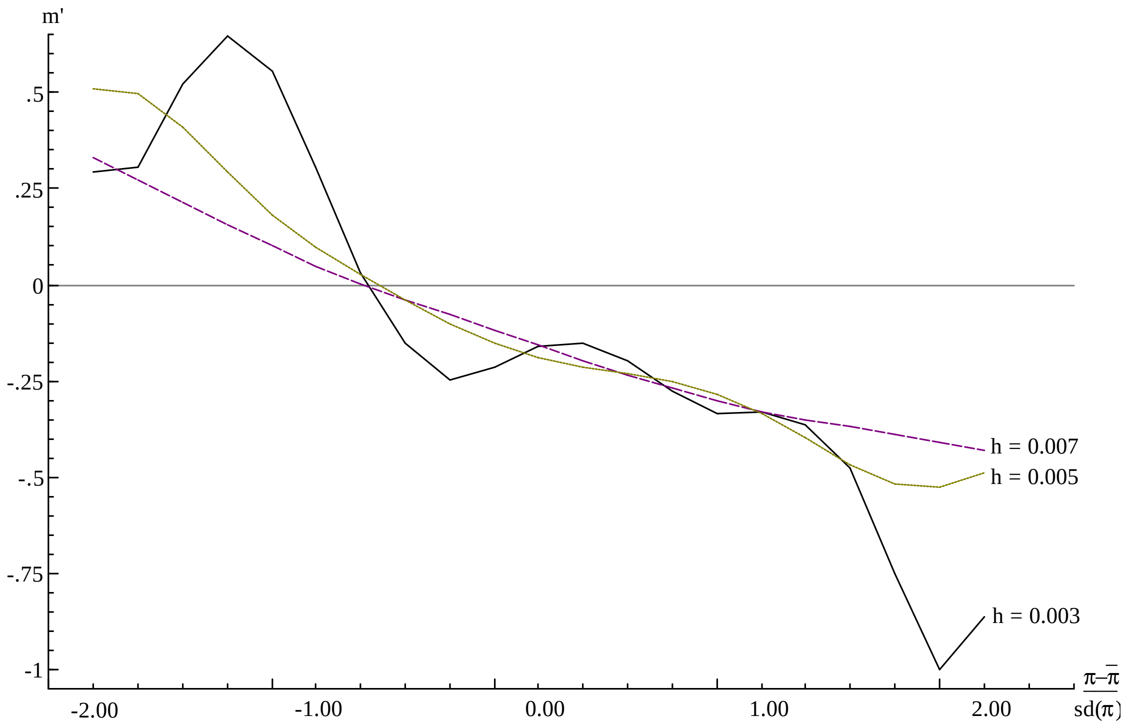



**Figure 7: Derivative of the $m^1$(.) function ± 2s.e. for *h* = 0.01, quarterly data**