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**INFLATION, RELATIVE PRICES AND NOMINAL RIGIDITIES**

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## Abstract

This paper examines the distribution of Belgian consumer prices and its interaction with aggregate inflation over the period June 1976-September 2000. Given the fat-tailed nature of this distribution, both classical and robust measures of location, scale and skewness are presented. We found a positive short-run impact of the skewness of relative prices on aggregate inflation, irrespective of the average inflation rate. The dispersion of relative prices has also a positive impact on aggregate inflation in the short run and this impact is significantly lower in the sub-sample starting in 1988 than in the pre-1988 sub-sample, suggesting that the prevailing monetary policy regime has a substantial effect on this coefficient. The chronic right skewness of the distribution, revealed by the robust measures, is *positively* cointegrated with aggregate inflation, suggesting that it is largely dependent on the inflationary process itself and would disappear at zero inflation. These results have three important implications for monetary policy.

First, as to the *transmission of monetary policy*, our results are in line with the predictions of menu cost models and therefore suggest that this type of friction can be an important factor behind the short run non-neutrality of monetary policy. Second, as to the *design of robust estimators of core inflation*, economic arguments based on menu cost models tend to highlight the importance of the absence of bias. We have proposed an unbiased estimator by taking the time-varying degree of chronic right skewness explicitly into account. Third, as to the *optimal rate of inflation*, the chronic right skewness found in the data provides no argument against price stability, as it appears as an endogenous response of optimising price setters and would disappear when targeting a zero inflation rate. This conclusion contrasts sharply with the implications of the *exogenously* assumed downward rigidity of Tobin (1972), which would justify targeting a sufficiently positive inflation rate in order to facilitate the adjustment of relative prices. Our empirical findings contradict the latter type of downward rigidity which implies a negative correlation between skewness and inflation. Therefore, the cross-sectional properties of Belgian inflation data do not provide strong arguments against a price stability-oriented monetary policy, such as the one pursued by the Eurosystem.



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## 1. INTRODUCTION

The aim of this paper is to study the cross-sectional properties of Belgian inflation data. The international literature on this issue highlights, broadly speaking, three not mutually exclusive aspects. First, many papers discuss the positive relationship between inflation on the one hand and the dispersion and/or the asymmetry of relative prices on the other. Often these relationships have been interpreted as being symptomatic of nominal rigidities of one form or another. Ball and Mankiw's menu cost model has recently been the focal point of this strand in the literature. Second, in line with the findings of Bryan, Cecchetti and Wiggins (1997), it is often documented that inflation data are fat-tailed and this motivates the stochastic approach to core inflation in which robust estimators of location are proposed. Third, in quite a few countries researchers found not only a considerable degree of frequently switching left and right skewness, but, on average, also a tendency towards right skewness.

The latter characteristic of the data, which is for instance discussed in Roger (2000), raises two additional questions. The first question is whether this chronic right skewness results from an exogenous form of downward nominal rigidity in product markets, such as the one put forward in Tobin (1972), or whether it is endogenously generated as menu cost models suggest. These two sources of asymmetry have different, empirically testable implications for the relation between inflation and skewness: the first source implies a negative relation, whereas the second implies a positive relation. Distinguishing between these two sources of skewness is particularly relevant from a monetary policy perspective, as they have different implications for the optimal rate of inflation. In the case of an exogenous form of downward nominal rigidity, low inflation rates are harmful as they complicate the (downward) adjustment of relative prices. In the other event, low inflation is desirable, as it reduces the costs associated with changing prices, as well as the impact of monetary policy on the distribution of relative prices. The second question raised by the chronic right skewness is how to incorporate it in the construction of robust estimators of core inflation.

This type of analysis often uses the classical characteristics of location, scale, skewness and kurtosis. These measures are, however, very sensitive to outlying values. In this paper we will compare them with robust alternatives, not only for location, as is typically done in the core inflation part of this literature, but also for scale, skewness and tail weight. We will address the question whether robust measures are applicable in this context, because inflation is often strongly influenced by outliers which are both correct and important. Overly strong robustness may downweight outliers too much and thus yield little sensitivity.

Extending the use of robust estimators to the scale, skewness and tail weight of the distribution of inflation data is the main statistical contribution of our paper.

From an economic point of view, one may wonder what justifies downweighting outlying values in inflation data. After all, we all buy goods and services for which prices change dramatically. There is often some assumption about price stickiness underlying this treatment of outliers. In the flexible price benchmark, the classical dichotomy holds: inflation is a purely monetary phenomenon, independent of relative prices which themselves are influenced by real determinants only. In that case it makes no sense to downweight outlying values in order to compute alternative (core) aggregate measures. With prices sticky in the short run, relative prices and inflation are no longer independent and causality goes in both directions. Relative price shocks - large shocks in particular - can affect aggregate inflation in the short run, thus temporarily masking the effect of the factors which are more relevant from a monetary policy perspective, such as the balance of supply and demand in the economy or the impact of inflation expectations. In such a setting, downweighting outliers aims at putting more weight on these factors. This provides the economic rationale for using robust estimators of location as core inflation measures, as Bryan and Cecchetti (1994) do on the basis of the menu cost model of Ball and Mankiw (1995).

In this paper we extend the argument to measures of scale and skewness and find that the robust alternatives are better suited to show the impact of these monetary policy related factors on the distribution of relative prices. Using this type of measures, we found that aggregate inflation had a substantial impact on the dispersion of relative prices in Belgium. Moreover, they revealed endogenous chronic right skewness in the distribution of observed relative prices, which tends to increase with aggregate inflation and which, according to our econometric results, would tend to disappear with zero inflation. This type of asymmetric price adjustment is what the menu cost model of Ball and Mankiw (1994) predicts, and we consider this as our main empirical result.

We came to these results on the basis of readily available inflation data at a relatively high level of aggregation (a 60 items breakdown), whereas it is the standard view that, intra-market rather than inter-market data are the appropriate dimension to uncover a causal relation running from the inflationary process to the distribution of relative prices<sup>1</sup>. We owe this to the fact that we have used the data at the highest frequency available, i.e. 1-monthly price changes, and that the cross-sectional properties were studied on the basis of robust

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<sup>1</sup> See for instance Ball and Mankiw (1994), Lach and Tsiddon (1992) or Cecchetti and Groshen (2000) on this point.



measures of scale and skewness. As to the corresponding classical measures, they allowed us to illustrate the causality in the other direction, i.e. from the distribution of relative price shocks to aggregate inflation.

The remainder of the paper is structured as follows. Section 2 describes the Belgian inflation data and discusses the format in which we will investigate them. In Section 3, robust measures of location, scale and tail weight are proposed and used to detect the presence of outlying values in the data. Section 4 briefly compares the classical and robust measures of location and scale. Section 5 discusses the classical and robust measures of skewness, as well as the chronic right skewness which they reveal. Section 6 presents, for the classical and the robust measures of skewness and scale respectively, their relation with inflation and discusses the empirical findings against the background of menu cost models. Moreover, the results are interpreted in terms of what is known in the literature on the optimal rate of inflation as "grease" and "sand" effects. Section 7 presents a robust unbiased estimator of core inflation. Section 8 sets out the conclusions.

## **2. DESCRIPTION OF THE DATA**

This study is based on Belgian monthly CPI data for the period June 1976<sup>2</sup> up to September 2000, yielding a total of 292 months. For each month, we have aggregated price indices of 60 different product categories. A list of these product categories is given in the Appendix. The index of product category  $i$  ( $i = 1, \dots, 60$ ) at month  $t$  ( $t = 0, \dots, 291$ ) is denoted by  $I_{i,t}$ .

We started by transforming these data into percentage 1-monthly price changes

$\Pi_{i,t}$  ( $t = 1, \dots, 291$ ), defined by

$$\Pi_{i,t} = \frac{I_{i,t}}{I_{i,t-1}} - 1 \quad (2.1)$$

As it was the aim to interpret some of the results obtained in terms of sticky versus flexible price behaviour, the obvious approach was to use data at the highest frequency available. Nevertheless, we also considered 6-monthly price changes. These price changes gave different outcomes, mainly due to the loss of information caused by summarising the data

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<sup>2</sup> It was not possible to start at an earlier point in time, for instance during the low inflation regime of the 60s, for data availability reasons.

over a longer time period. We therefore decided not to consider them subsequently. Also, in Aucremanne (2000) it was found that, for instance, the chronic right skewness was much more pronounced for 1-monthly price changes than for 3-, 6- or 12-monthly price changes.

Note that these price changes can also be seen as the first order approximation of logarithmic price changes

$$\ln\left(\frac{I_{i,t}}{I_{i,t-1}}\right) \quad (2.2)$$

Both approaches did indeed yield approximately the same results, including those related to the presence of chronic right skewness. This suggests that using percentage price changes instead of log price changes is not an important force behind the chronic right skewness, as could have been expected on the basis of Roger (2000). The latter paper inspired us also to test the robustness of the results obtained by removing from the dataset 10 product categories for which prices were obviously adjusting infrequently. These are categories which are mainly regulated by government or law, e.g. bread, tobacco and education. They are marked in the Appendix with an asterisk. Our results - including the endogenous nature of the chronic right skewness - were only slightly influenced by this modification. Therefore, the paper only presents the results for all 60 product categories.

Summarizing, we can represent our final dataset  $\Pi_{i,t}$  with a matrix that consists of 291 rows (indicating the different periods) and 60 columns (for the product categories). A cross-section contains the price changes of one particular month, and thus corresponds with one row of the data matrix. Additionally, we have taken account of the fact that each product category has a time-varying weight  $w_{i,t}$  which is obtained by multiplying for each period  $t$  its fixed Laspeyres-type weight  $w_i$  by the change in its relative price between the base period and period  $t$ . The Laspeyres-type weights  $w_i$  of the CPI reflect the importance of each category in total household consumption expenditure in the base period. Evidently, the weights satisfy  $\sum_{i=1}^{60} w_{i,t} = 1$  for all  $t$ . In so doing, the weighted mean of the 60 product-specific price changes  $\Pi_{i,t}$  corresponds to aggregate inflation.

### 3. DETECTION OF OUTLIERS

In this section we will mainly show that the Belgian price changes contain a substantial amount of outlying values. For this, we will consider the cross-sections and compute robust measures of tail weight for each of them. To show the importance of the tails of univariate data (as the cross-sections are), we can use the classical measure of kurtosis. In general, the kurtosis is said to characterize the fatness of the tails, or equivalently the tail weight, but it also reflects the shape of the density in the centre. Moreover, the classical measure of kurtosis is based on moments of the dataset, and thus it is strongly influenced by outliers. Despite the fact that outliers determine the tail weight, we can construct robust measures to compute it. For this purpose, we first need a robust estimator of *location* and *scale* for univariate data. In general, we denote the sample by  $x_i$  ( $i = 1, \dots, n$ ) and the corresponding weights by  $w_i$ . As usual, the weights sum to 1. Time subscripts are omitted to simplify notations.

#### 3.1 *Robust measure of location*

A measure of location should estimate a value that characterises the central position of the data. The best known measure of location is the *weighted mean* or average, defined as:

$$\bar{x} = \sum_{i=1}^n w_i x_i \quad (3.1)$$

Note that with the dataset we analyse, the weighted mean corresponds to observed inflation.

A typical robust measure of location is the *median*. It is defined by first sorting the observations from smallest to largest, and then taking the middle observation (for an even number of observations we take the average of both observations in the middle). Here, we need the weighted variant of the median. In general, all measures based on percentiles can be modified easily to their weighted version by filling in the weighted percentiles. That is the main reason why all robust measures presented in this paper are based on percentiles<sup>3</sup>. For the  $p\%$  weighted percentile, denoted by  $Q_p$ , we first sort the observations from smallest to

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<sup>3</sup> Initial work on other robust measures based on couples and triples of observations was abandoned precisely because it was not straightforward to construct them in weighted terms. Verbrugge (1999) presents a robust measure of skewness based on triples for *unweighted* inflation data. Brys, Hubert and Struyf (2002) also discuss some unweighted robust skewness measures based on percentiles, couples and triples.

largest, and then take the first value with cumulative weight higher than  $p\%$ . The weighted median equals the 50% weighted percentile, or  $Q_{0.50}$ .

### 3.2 Robust measure of scale

Scale characteristics measure how "spread out" the data values are. The classical measure of scale is the *standard deviation*, which is given by:

$$s = \sqrt{\frac{\sum_{i=1}^n w_i (x_i - \bar{x})^2}{1 - \sum_{i=1}^n w_i^2}} \quad (3.2)$$

where  $\bar{x}$  stands for the mean<sup>4</sup>. The *interquartile range* or IQR is a robust measure of scale which, like the median, is based on percentiles and is easy to calculate. The (weighted) interquartile range is the distance between the 75% percentile  $Q_{0.75}$  and the 25% percentile  $Q_{0.25}$ , or

$$\text{IQR} = Q_{0.75} - Q_{0.25} \quad (3.3)$$

### 3.3 Robust measures of tail weight

We propose six alternatives to the classical kurtosis. We will denote them with *taillength* and *tailmass*. The *taillength* measures are based on differences between the outlying observations and the median of the data, while the *tailmass* measures simply count the outliers. To distinguish the outlying values from the regular ones we use the following outlier rules:

1. The *percentile rule*, where an observation is declared an outlier when it lies outside the interval:

$$[Q_p, Q_{1-p}] \quad (3.4)$$

for any  $0 < p < 1$ . Here, we consider two particular choices of  $p$ , namely  $p=12.5\%$  and  $p=25\%$ .

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<sup>4</sup> Formula (3.2) is based on the unbiased sample moments of unequally-weighted frequency distributions presented in Roger (2002).

2. The *general boxplot rejection rule* defines outliers as points outside the interval:

$$\left[ \text{med} - \frac{3}{2} \text{IQR}, \text{med} + \frac{3}{2} \text{IQR} \right] \quad (3.5)$$

This criterion is based on the definition of the whiskers in the univariate boxplot (see Tukey (1977))<sup>5</sup>.

3. The *asymmetric boxplot rule*. This rule is a special case of the bivariate bagplot in Rousseeuw, Ruts and Tukey (1999), which is a bivariate generalisation of the boxplot. Points are considered outlying when they lie outside the interval:

$$\left[ \text{med} - 3(Q_{0.50} - Q_{0.25}), \text{med} + 3(Q_{0.75} - Q_{0.50}) \right] \quad (3.6)$$

Now we can introduce the *taillength* measures. For any of the outlier rules, we define

$$\text{left-taillength} = \sum_{i=1}^n w_i \left( \frac{\text{med} - \hat{x}_i}{\text{IQR}} \right) \quad (3.7)$$

where  $\hat{x}_i$  denote the left outliers found by that specific outlier rule. Note that the median and the IQR are computed on all observations, including the outliers. Using the percentile rule, we so obtain  $\text{left-taillength}(125)$  and  $\text{left-taillength}(250)$ . The general boxplot rule and the asymmetric boxplot rule lead to the taillength measures  $\text{left-taillength}(\text{box})$  and  $\text{left-taillength}(\text{abp})$ . By replacing the left outliers by the right outliers in (3.7), and by using  $(\hat{x}_i - \text{med})$  instead of  $(\text{med} - \hat{x}_i)$  in the numerator, we get  $\text{right-taillength}(125)$ ,  $\text{right-taillength}(250)$ ,  $\text{right-taillength}(\text{box})$  and  $\text{right-taillength}(\text{abp})$ .

Next, we obtain measures of the (left) *tailmass* as the sum of the weights of the (left) outliers, i.e.

$$\text{left-tailmass} = \sum_{i=1}^n w_i \theta_i \quad (3.8)$$

where  $\theta_i$  equals 1 if  $x_i$  is a left outlier and 0 if not. This leads to  $\text{left-tailmass}(\text{box})$  and  $\text{left-tailmass}(\text{abp})$ . Note that it makes no sense to consider  $\text{left-tailmass}(125)$  or

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<sup>5</sup> Note that this interval corresponds approximately to  $[\bar{x} - 2s, \bar{x} + 2s]$  in case of the normal distribution.

left-tailmass(250) since these measures are always equal to 12.5% or 25%, regardless of the data. Considering the right outliers, we obtain right-tailmass(box) and right-tailmass(abp).

### 3.4 Results for Belgian inflation data

All measures defined above were applied to the cross-sections of the Belgian inflation data. In Figures 1 and 2 the resulting time series are plotted, by considering all 291 consecutive cross-sections. On the plots, the scatter points of the different measures are depicted, together with a smoother (solid line). We have chosen to use a *lowess* smoother, which is a robust scatter plot smoother<sup>6</sup>. Lowess uses robust locally linear fits, by placing a window about each point and weighting the points that are inside the window so that nearby points get the most weight.

The tailmass(box) measures show substantial proportions of outliers and these proportions are, moreover, very volatile. This is why we will use robust measures of location, scale and skewness to analyse the data further on. Before doing so, we have a closer look at the tails. On the basis of the smoothed curves, left-tailmass(box) amounted to roughly 5% at the beginning of the sample and increased to approximately 10% at the end of the sample. Right-tailmass(box), in contrast, tends to oscillate around 15% during the whole sample. These observed tail weights are substantially in excess of those for the normal distribution, for which the corresponding tailmass(box) measures amount to 2.15% for each tail. This clearly illustrates the fat-tailed nature of Belgian inflation data, a finding which is in line with the results in the international literature on this issue<sup>7</sup>. It should, however, be stressed that most of the international evidence is based on the classical kurtosis, which may for several reasons underestimate the importance of outliers<sup>8</sup>.

We interpret the fact that the smoothed curve indicates an upward tendency for left-tailmass as a first piece of evidence against the existence of an exogenous form of downward nominal rigidity in Belgian product markets. Indeed, with aggregate inflation substantially lower in the second half of the period, the probability of having nominal price decreases in the left-hand tail increases. If prices are effectively rigid downward- i.e. if they cannot decrease in nominal terms for purely exogenous reasons -, this should reduce left-tailmass at the end of the

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<sup>6</sup> See Cleveland (1979).

<sup>7</sup> See for instance Bryan, Cecchetti and Wiggins (1997), Roger (1997), Kearns (1998), Bakshi and Yates (1999), Meyler (1999 and 2001), Marques and Mota (2000), Aucremanne (2000), Vega and Wynne (2001), Knudsen and Nielsen (2002), Roger (2002).

<sup>8</sup> See for instance Aucremanne (2000) on this so-called masking issue.

period, whereas we observe exactly the opposite. The left-taillength measures do not show a tendency to decrease over time either, thus confirming the conclusion made above.

The fact that  $\text{right-tailmass}(\text{box})$  tends to have larger values than the corresponding left alternative is a first indication of the existence of chronic right skewness, albeit apparently decreasing over time. A possible disadvantage of the general boxplot rejection rule is that it takes the left part of the data into account when defining a right outlier and vice versa. Using the asymmetric boxplot rule overcomes this problem. In so doing, we find that  $\text{left-tailmass}(\text{abp})$  increases and  $\text{right-tailmass}(\text{abp})$  decreases relative to their symmetrical alternatives, which can be seen as another indication of chronic right skewness. These asymmetrically constructed tailmass measures confirm the tendency for the left-hand tail to increase and the more stationary behaviour of the right-hand tail.

Figure 1 – Left-tailweight

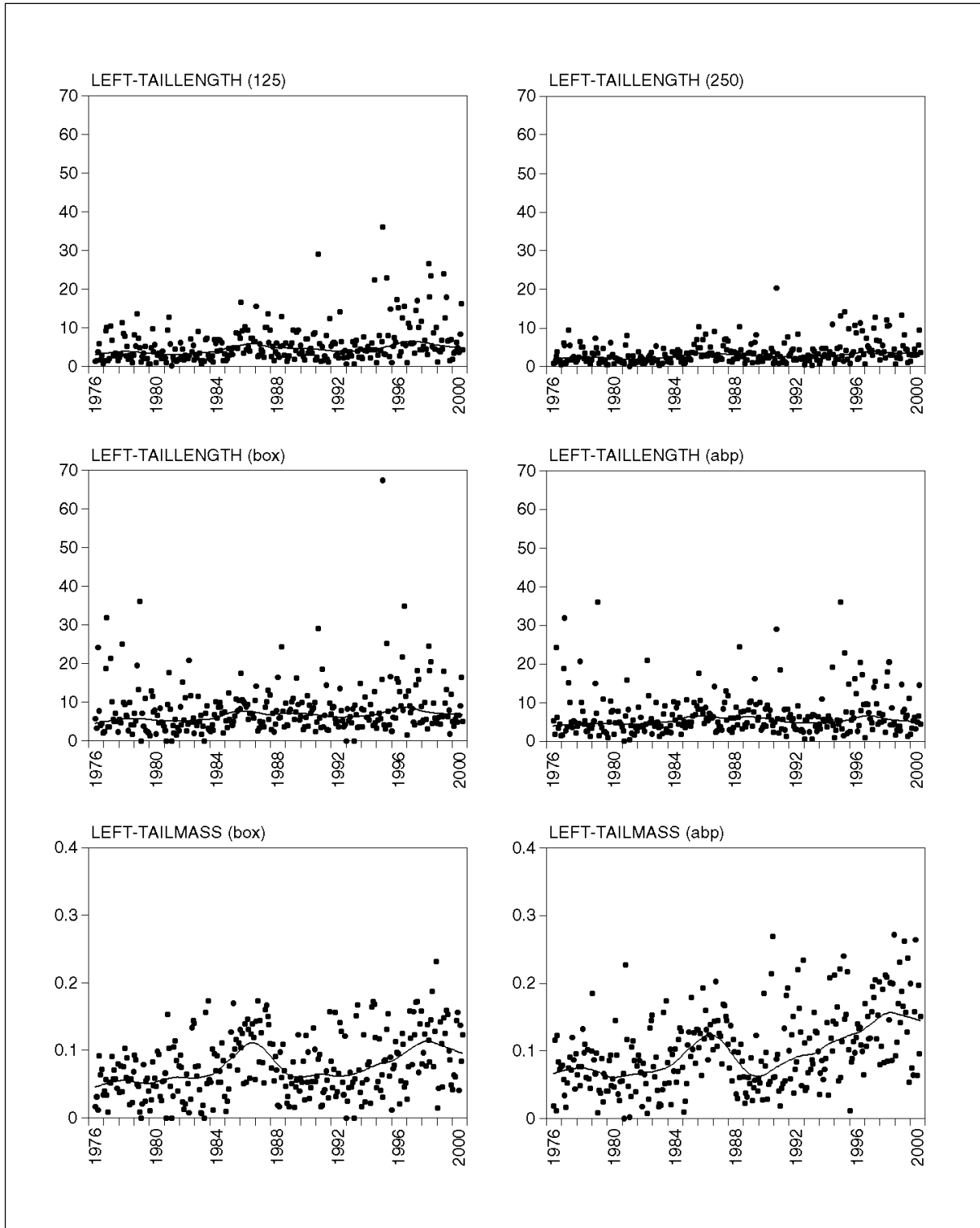
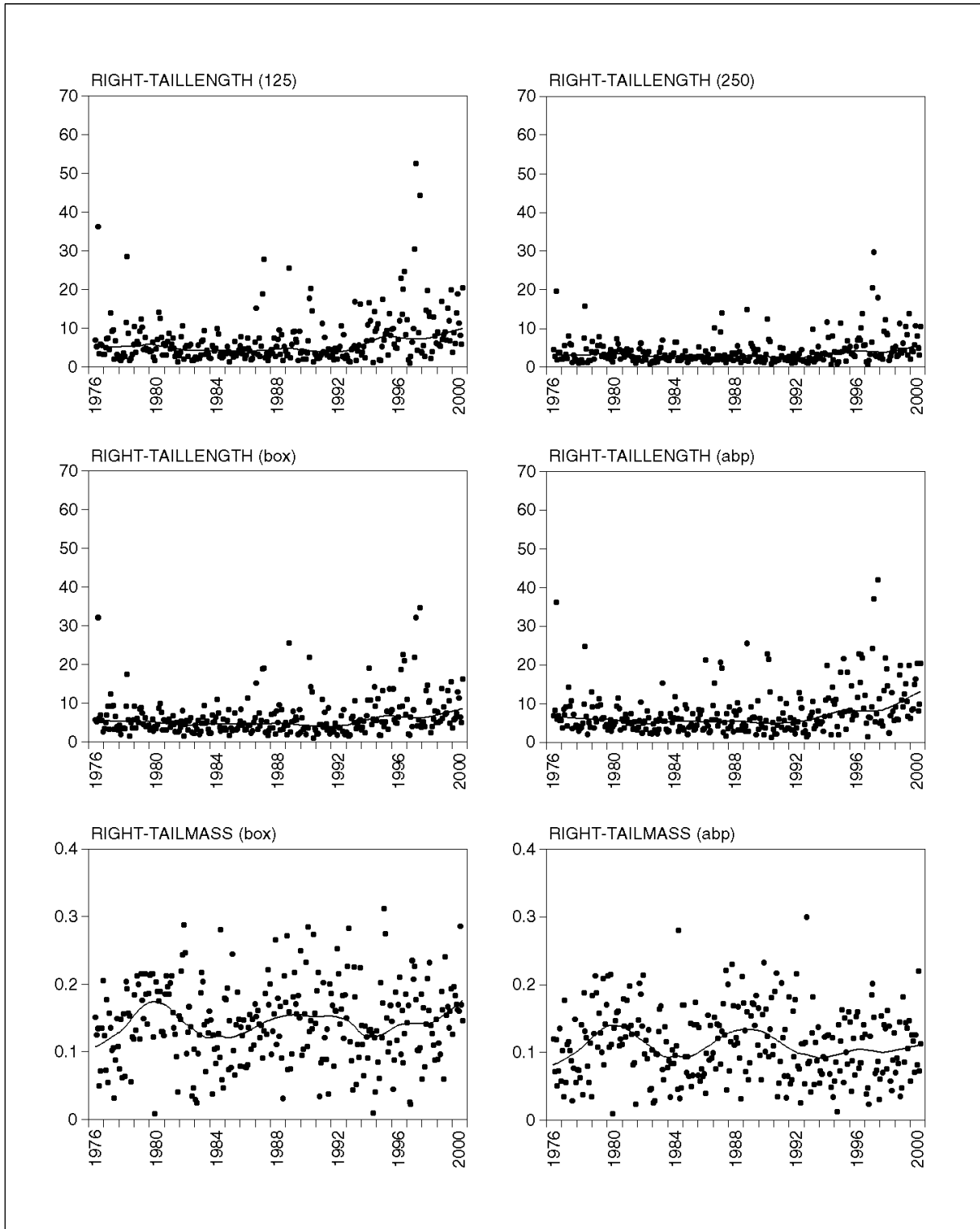




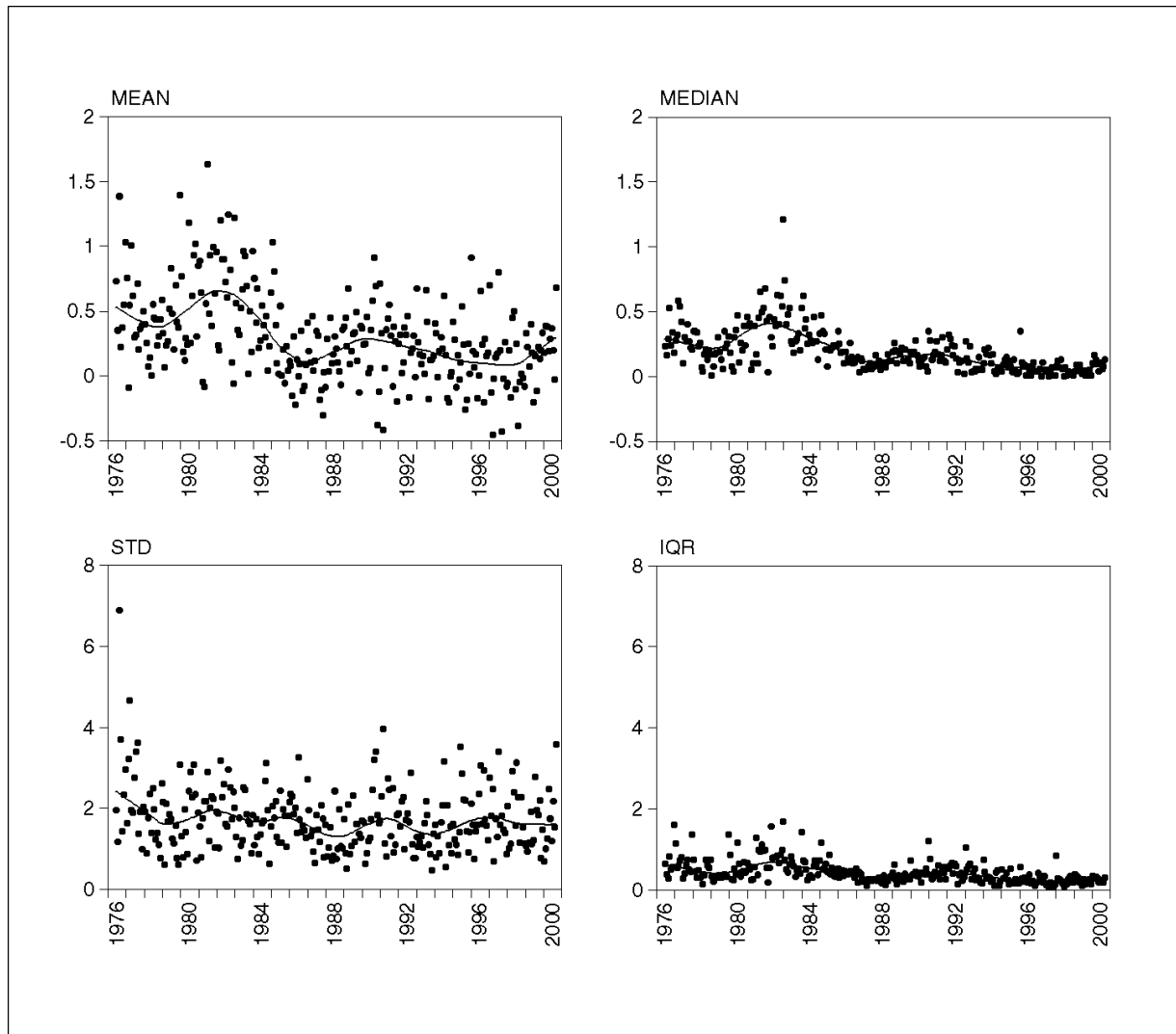
Figure 2 – Right-tailweight



#### 4. CLASSICAL AND ROBUST MEASURES OF LOCATION AND SCALE

The classical and robust measures of location and scale are plotted in Figure 3. The robust alternatives were already used in the previous section to construct robust measures of tail weight. In this section we will compare them to their classical counterparts.

**Figure 3 - Classical and robust measures of location and scale**



## 4.1 Location

As the use of robust measures of location for inflation data is well-developed in the core inflation literature, we only give some brief comments when comparing the *mean* to the *median*. The interested reader is referred to other papers in this field for a deeper analysis<sup>9</sup>. As the ADF-tests reported in Table 1 fail to reject the null hypothesis of a unit root, both time series contain a stochastic trend. The smoothed curves in Figure 3 seem to confirm this conclusion. The median is substantially less volatile around its trend than the mean, thus confirming that using a robust estimator of location yields a substantial gain in efficiency in the case of a fat-tailed distribution. This could motivate the use of the median as a measure of core inflation statistically. The first robust core inflation measure, proposed by Bryan and Pike (1991), was indeed the median. Subsequently, a wide range of alternative robust estimators of core inflation have been proposed (see footnote 9 for references).

**Table 1 - Unit root tests for the classical and robust measures of location, scale and skewness**

(Augmented Dickey Fuller test statistic)

	With a constant in the test equation <sup>1</sup>	Significance of the constant - P-value <sup>2</sup>	Lag number <sup>3</sup>
Location			
Mean	-1.74		11
Median	-1.58		11
Scale			
Standard deviation	-3.97	0.00	11
IQR	-1.74		11
Skewness			
Skew(class)	-3.82	0.12	11
Dskew(class)	-5.80	0.02	11
Dskew(125)	-2.55		11
Dskew(250)	-3.44	0.00	11
Meme	-2.59		11

<sup>1</sup> The 95 p.c. critical value for rejection of the null hypothesis of a unit root is -2.87. Inclusion of a trend in the test equations for those variables for which the null of a unit root was not rejected did not change the results. In other words, they were not trend stationary either.

<sup>2</sup> Only reported when the null hypothesis of a unit root is rejected.

<sup>3</sup> Given the monthly frequency of the data, 11 lags were necessary to produce residuals without autocorrelation.

<sup>9</sup> See for instance Bryan, Cecchetti and Wiggins (1997), Roger(1997), Kearns (1998), Alvarez and de los Llanos Matea (1999), BIS (1999), Bakshi and Yates (1999), Le Bihan and Sédillot (1999), Meyler (1999 and 2001), Roger (2000), Aucremanne (2000), Landau (2000), Vega and Wynne (2001), Knudsen and Nielsen (2002), Roger (2002).

As can be seen from Table 2, the median is cointegrated with the mean, but with a coefficient (0.5934) that is substantially less than one. In Figure 3 it can indeed be verified that the median tends to be lower than the mean. This is another indication of chronic right skewness in the Belgian inflation data and raises the question of constructing a core inflation measure which incorporates this characteristic of the data (see Section 7). As such, the median violates the first condition put forward in Marques, Neves and Sarmento (2000) for an ideal core inflation measure. It is said to be a *biased* estimator. Note also that the constant in the cointegration relation is very close to zero, suggesting that the chronic asymmetry would disappear completely at zero inflation (see also Section 6).

**Table 2 - Cointegration of the robust measure of location with actual inflation (mean)**

	Median
Lag number	11
Cointegration equation (CE) <sup>1</sup>	
Constant	0.0002 (0.01)
Mean	-0.5934 (-10.54)
Trace statistic <sup>2</sup>	
No CE	35.09
At most 1 CE	4.70

<sup>1</sup> Normalised cointegration coefficients, with the coefficient of the variable on top of the column normalised to 1.

<sup>2</sup> The 95 p.c. critical value is 19.96 for rejection of the null hypothesis of no cointegration equation and 9.24 for rejection of the null hypothesis of at most 1 cointegration equation. The cointegration rank and the specification of the deterministic components of the model were determined jointly on the basis of the general test procedure discussed in Johansen (1992). However, models having trends in the levels of the variables (constants in the VAR) were not considered.

## 4.2 Scale

Comparing the classical standard deviation with IQR indicates that relying on a robust alternative dramatically alters the characteristics of the measure of scale. According to the ADF-tests in Table 1, IQR contains a unit root whereas the classical measure appears as a stationary variable. Moreover, Figure 3 shows that the variability of IQR around its smoothed curve is substantially lower than the variability of the classical standard deviation. We owe this to the fact that the classical measure is dominated by the impact of outliers or, in economic terms, by large relative price shocks. In the absence of outliers, IQR would in contrast tend to be larger than the standard deviation (in the case of the normal distribution

IQR  $\cong 1.35s$ ). Evidently, the impact of outliers on the classical measure of scale is magnified by squaring their deviations from the mean, as can be seen from formula (3.2). A more detailed analysis of the measures of scale and their relation with aggregate inflation is presented in Section 6.

## 5. CLASSICAL AND ROBUST MEASURES OF SKEWNESS

Skewness reflects the shape of the distribution. A symmetric distribution has zero skewness, a distribution which is asymmetric with the largest tail to the right has a positive skewness, and a distribution with a longer left tail has a negative skewness.

### 5.1 *Classical measures of skewness*

Normally, the standardized, unitless measures of asymmetry (skewness) are used. The best known measure of skewness is the *classical skewness*<sup>10</sup> which is defined as:

$$\text{skew(class)} = \frac{\sum_{i=1}^n w_i (x_i - \bar{x})^3}{\left(1 - 3 \sum_{i=1}^n w_i^2 + w_i^3\right) s^3} \quad (5.1)$$

However, we will consider destandardized measures of asymmetry as well. In this way skewness becomes dependent on scale, but this correlation does not bother us. Indeed, the economic models we rely on suggest that scale and skewness interact (see also Section 6). Ball and Mankiw (1995) show that scale magnifies the effect of skewness on inflation, while the asymmetric range of inaction in Ball and Mankiw (1994), which is a possible source of right skewness in the data, is also unscaled. All destandardized measures of asymmetry are expressed in some units, and the robust measures have the same units as the data and the measures of location and scale. The latter will facilitate their use in the construction of an unbiased core inflation estimator in Section 7. The destandardized *classical skewness* is defined as:

$$\text{dskew(class)} = \frac{\sum_{i=1}^n w_i (x_i - \bar{x})^3}{1 - 3 \sum_{i=1}^n w_i^2 + 2 \sum_{i=1}^n w_i^3} \quad (5.2)$$

<sup>10</sup> Formulae (5.1) and (5.2) are based on the unbiased sample moments of unequally-weighted frequency distributions presented in Roger (2002).

## 5.2 Robust measures of skewness

A disadvantage of the classical measures of skewness is that they change dramatically when we introduce an outlier. The impact of outliers is even stronger than in the case of the mean and the standard deviation. On the other hand, skewness is supposed to measure the asymmetry of the observations, so we may not downweight outliers too heavily. The first robust alternative we propose is  $dskew(125)$ , which is given by:

$$dskew(125) = (Q_{0.875} - Q_{0.50}) - (Q_{0.50} - Q_{0.125}) \quad (5.3)$$

with  $Q_p$  again the  $p\%$  (weighted) percentile. The following measure is very similar, but it is based on other percentiles. We compute  $dskew(250)$  as:

$$dskew(250) = (Q_{0.75} - Q_{0.50}) - (Q_{0.50} - Q_{0.25}) \quad (5.4)$$

Another measure of skewness can be obtained by standardizing the data as:

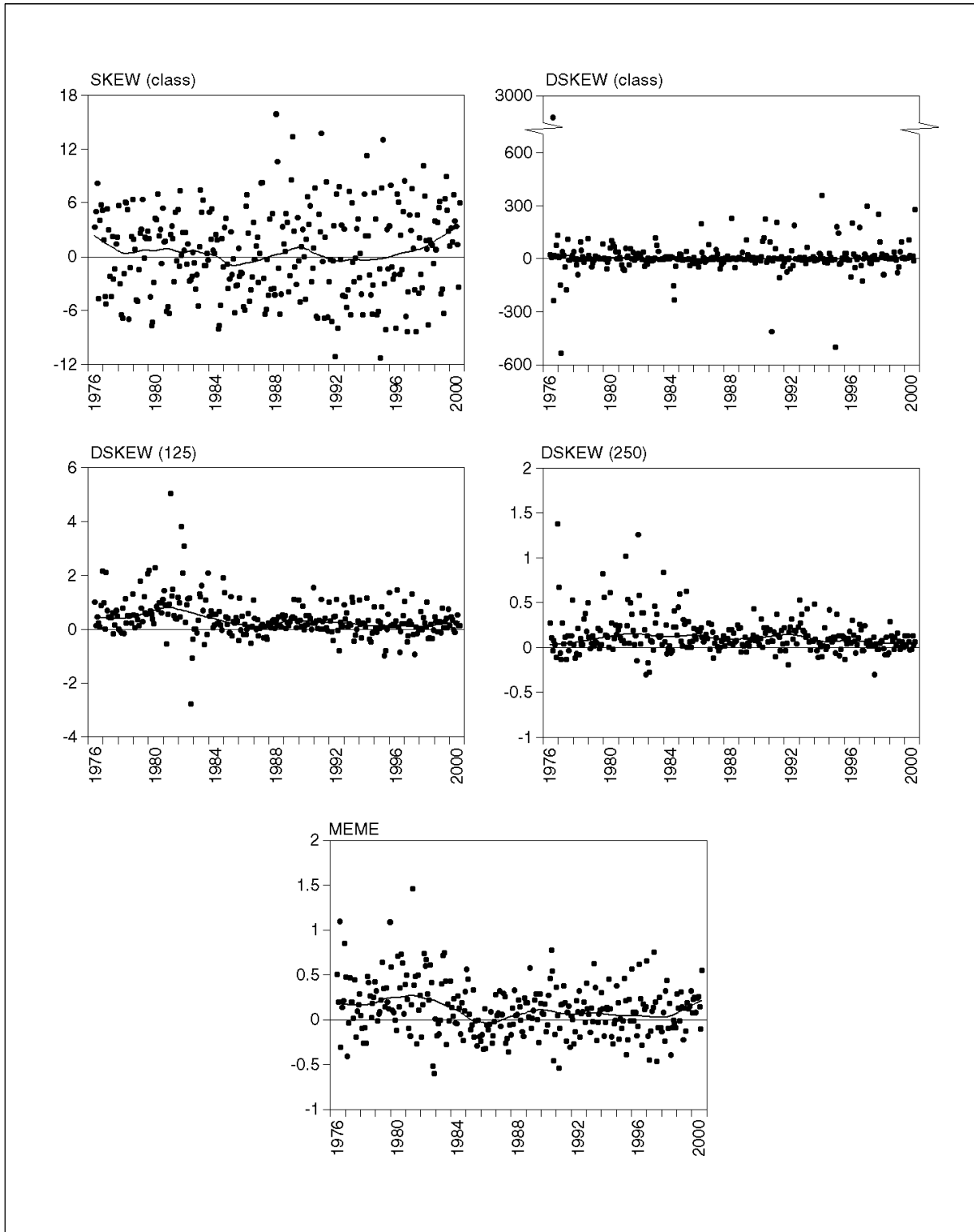
$$z_i = x_i - \text{med} \quad (5.5)$$

and then taking the first moment of these standardized observations which corresponds to the *mean-median difference*, yielding:

$$\text{meme} = \bar{z} = \sum_{i=1}^n w_i z_i \quad (5.6)$$

Strictly speaking this measure is not robust, as it uses the mean. However, we consider it here as an alternative measure of skewness, as it has, compared to the classical measures, the advantage that the influence of what happens in the tails is far less accentuated.

Figure 4 - Classical and robust measures of skewness



### 5.3 Chronic right skewness

We have applied both types of measures of skewness on the cross-sections of the Belgian inflation data. The resulting time series are plotted in Figure 4. The classical measures of skewness show a substantial degree of frequently switching left and right skewness, which is a typical result in the literature in regard to the analysis of inflation data<sup>11</sup>. However, they do not show a clear tendency towards positive values. As for scale, the time series properties of the robust measures of skewness are fundamentally different from those of the classical ones. As the ADF tests in Table 1 indicate, the null hypothesis of a unit root in the classical measures was rejected at the 99% significance level. Moreover, the constant term in the test equation was not different from zero at the conventional significance levels for *skew(class)*. For *dskew(class)* the constant was significant at the 95% level, but not at the 99% level. Also, in this case the estimated value of the constant (11.14) is small relative to the variability observed in the series. Hence, the classical measures hardly reveal any form of chronic right skewness in a significant way, notwithstanding the fact that we provided evidence of a tendency towards right skewness in Sections 3 and 4.

In contrast, the robust measures *dskew(125)* and *dskew(250)* show a more pronounced tendency towards positive values. Hence, they better reveal the chronic right skewness of the data. The mean-median difference *meme* also shows such a tendency, although it takes negative values more frequently than the two other robust measures. We owe the latter to the fact that *meme* is in fact not robust. The null hypothesis of a unit root is rejected for *dskew(250)* at the 95 p.c. significance level, suggesting mean reversion around a significant *positive* constant. In contrast, for *dskew(125)* and *meme*, the null of a unit root is not rejected, suggesting a time-varying degree of chronic right skewness. Particularly in the second half of the sample the tendency towards right skewness seems less pronounced, thus confirming the findings of Sections 3 and 4. A chronic tendency towards right skewness is also found in several other countries<sup>12</sup> and some of these studies also mention that this tendency is less pronounced in the most recent period, when inflation is lower.

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<sup>11</sup> This property of the data is documented, for several other countries, in for instance Bryan, Cecchetti and Wiggins (1997), Roger (1997), Kearns (1998), Alvarez and de los Llanos Matea (1999), Bakhshi and Yates (1999), Le Bihan and Sédillot (1999), Marques and Mota (2000), Aucremanne (2000), Meyler (2001), Vega and Wynne (2001), Knudsen and Nielsen (2002), Roger (2002).

<sup>12</sup> See for instance Chatelain, Odonnat and Sicsic (1996) and Le Bihan and Sédillot (1999) for France, Roger (1997) and Roger (2002) for New Zealand, Kearns (1998) for Australia, Marques and Mota (2000) for Portugal, Meyler (1999) and Meyler (2001) for Ireland, Vega and Wynne (2001) for the euro area as a whole and for participating countries. In contrast, in the US the distribution is symmetric (Cecchetti and Groshen (2000)).



## 6. INFLATION AND RELATIVE PRICES

In this section we report our empirical results regarding the relationship between aggregate inflation and the classical and robust measures of scale and skewness, and interpret them against the background of menu cost models, as these have recently become the focal point in this literature. We start with an intuitive discussion of the main implications of these models.

### 6.1 *Menu cost models*

Menu cost models assume that changing prices involves a cost, which is fixed in the sense that it does not depend on the magnitude of the price change. Examples of this type of costs are changes to price lists, catalogues, advertising, etc. Simplifying the economy to one without trend inflation, where only zero mean relative price shocks occur, Ball and Mankiw (1995) show that the menu cost produces a *symmetric* range of inaction. Firms react only to large relative price shocks - responding equally to positive and negative shocks -, for which the advantage of changing prices outweighs the menu cost.

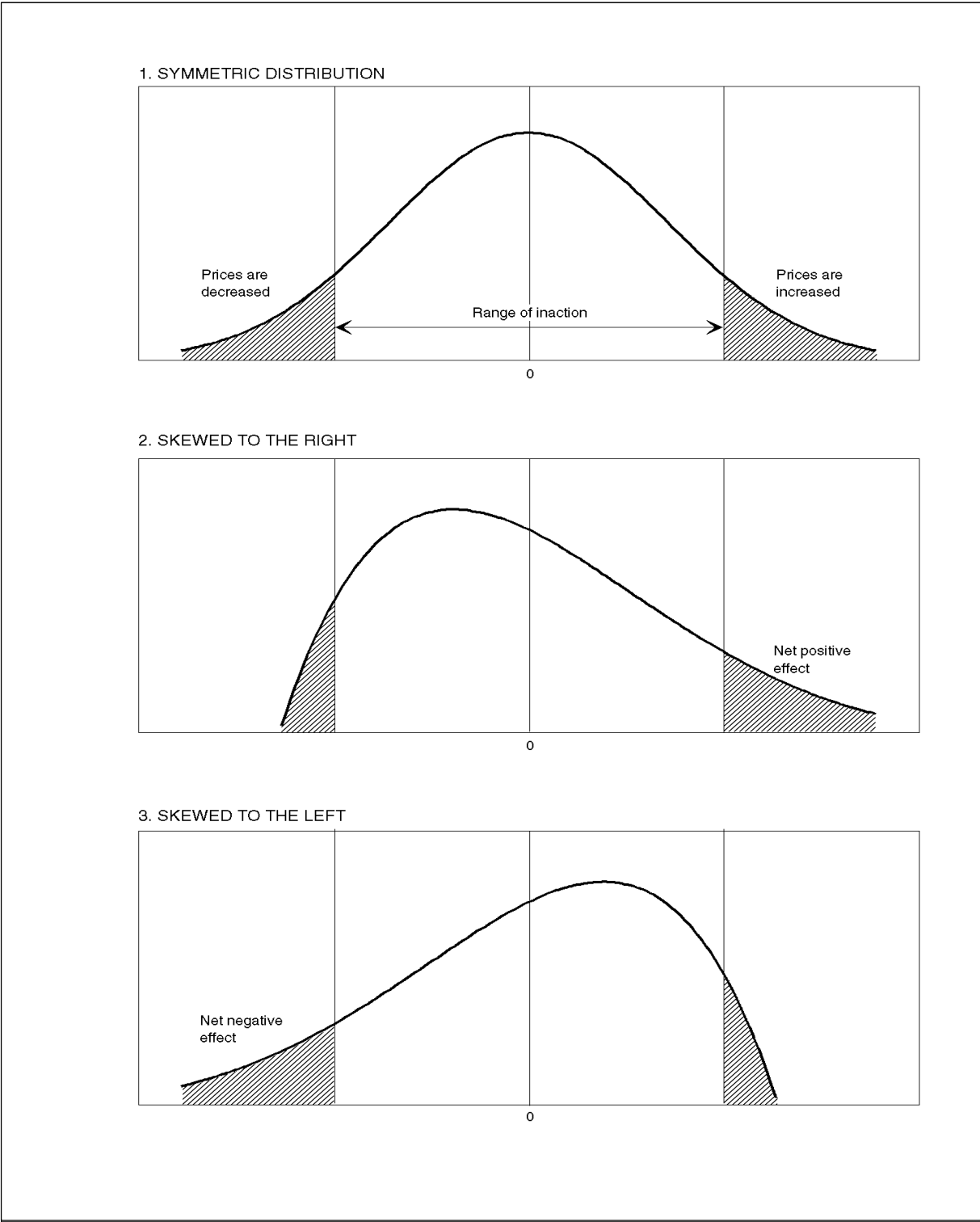
As Figure 5 below illustrates<sup>13</sup>, symmetric distributions of relative price shocks do not affect aggregate inflation, as (large) price increases are exactly compensated by (large) price decreases. It is easy to verify that, in this case, increasing the variance of the (symmetric) distribution of relative price shocks has no effect on aggregate inflation either. However, things change when the distribution of relative price shocks is asymmetric. When these shocks are skewed to the right (left), large relative price increases (decreases) will no longer be compensated totally by large relative price changes in the opposite direction, and aggregate inflation will increase (decrease) in the short run. In other words, the model predicts that inflation is positively related to the skewness of the relative price shocks *in the short run* and, by extension, also to the skewness of observed relative price changes. Increasing the variance of the skewed distributions - a phenomenon which can be measured by using destandardised measures of skewness - magnifies the impact of skewed relative price shocks on inflation. Ball and Mankiw have interpreted the positive correlation between inflation and the skewness of the distribution of relative prices as evidence in favour of menu cost models.

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<sup>13</sup> The Figure is based on Ball and Mankiw (1995).

**Figure 5 - The impact of the shape of the distribution of relative price shocks in a menu cost model without trend inflation**

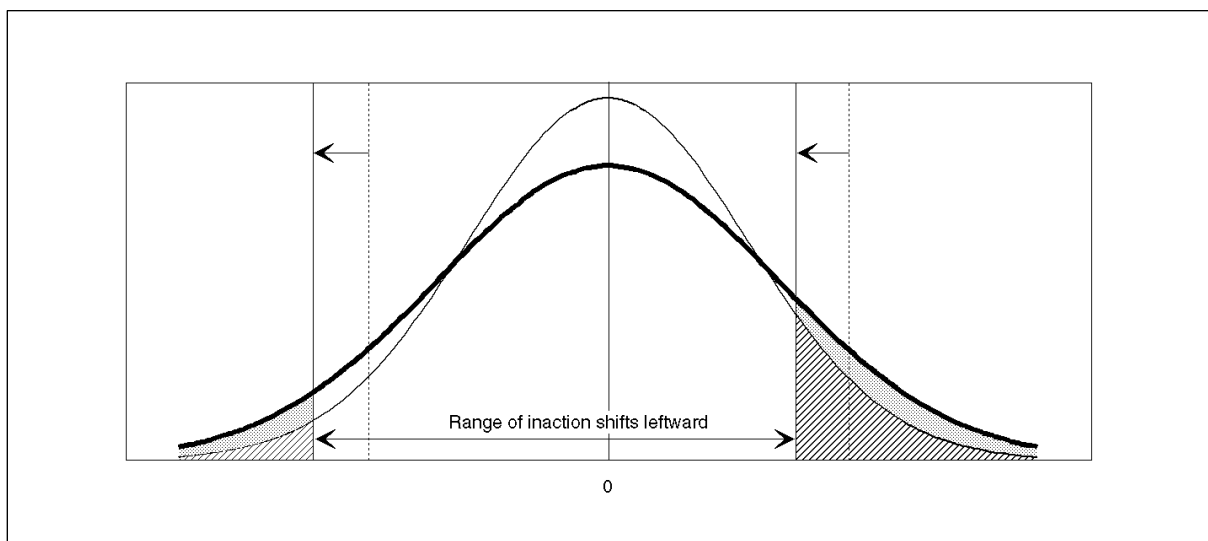
(density functions of zero mean relative price shocks)



When trend inflation is introduced in an otherwise similar model, the optimising behaviour of price setters leads to an *asymmetric* range of inaction, as is illustrated in Figure 6. Trend inflation shifts the range of inaction *endogenously* leftwards: for a given absolute value of a relative price shock, an upward adjustment is more likely than a downward adjustment. The intensity of this shift depends on trend inflation<sup>14</sup>. Intuitively, this asymmetry comes from the fact that trend inflation can lead to the desired relative price decreases even if nominal prices are not changed, whereas it magnifies the incentive to change nominal prices when a relative price increase is desirable.

**Figure 6 - Introducing trend inflation in a menu cost model**

(density functions of zero mean relative price shocks)



This has three implications for the relation between inflation and relative prices: (i) as is illustrated in Figure 6, increasing the variance of a symmetric distribution of relative price shocks now has a positive impact on observed inflation in the short run, and this impact increases with the intensity of the leftward shift of the range of inaction or, equivalently, with trend inflation; (ii) the impact of asymmetric price shocks on inflation in the short run is not fundamentally altered by the introduction of trend inflation in the model; and (iii) the model generates a tendency towards right skewness in the distribution of observed price changes, as price adjustment is more likely for positive relative price shocks than for negative shocks and this effect, too, increases with trend inflation. Clearly, for the latter effect, causality runs from trend inflation to the distribution of observed relative prices and this impact would

<sup>14</sup> For a formal derivation of this result we refer to Ball and Mankiw (1994) or to Andersen (2000).

disappear when targeting a zero inflation rate, in which case the range of inaction is again symmetric as in the first model.

Our findings with respect to the time series properties of the classical and the robust measures of scale and skewness respectively are not without consequences for their relation with aggregate inflation. They indicate that it is not appropriate to relate the classical measures of scale and skewness to inflation. Rather they should be related to the first difference of inflation. This relation will be discussed in Section 6.2. With our data, relating the robust measures of scale and skewness to inflation must, in contrast, be done in a cointegration framework. The latter relations will be discussed in Sections 6.3 and 6.4.

## **6.2 Using classical measures: relative prices affecting inflation**

Table 3 reports our regression results linking the first difference of observed monthly inflation with the classical measures of scale and skewness. Following De Abreu Lourenco and Gruen (1995), the sample was split into two sub-periods to illustrate the role of trend inflation. According to the menu cost theory, the prevailing inflationary regime should particularly affect the estimated coefficient for scale. The first sub-period, running from 1976.09 to 1987.12, was characterised by a substantially higher average inflation rate than the second sub-period, running from 1988.01 to 2000.09. For each sub-period a first regression is reported using the contemporaneous standard deviation, as well as two lags of this variable. The next equation adds the classical skewness and two lags as explanatory variables. Adding *dskew(class)* in a similar way did not yield significant coefficients for this variable<sup>15</sup>. Table 3 gives the estimated coefficients, as well as the corresponding t-statistics. As the Durbin Watson statistics of the estimated regressions reveal the presence of autocorrelation, the t-statistics were calculated using the heteroskedasticity and autocorrelation consistent (HAC) covariance matrix estimator of Newey and West (1987).

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<sup>15</sup> Results of this third type of regression are therefore not reported in Table 3.

**Table 3 - Regression results relating the change in inflation to the classical measures of scale and skewness**

	Constant	Standard deviation			Skewness (class)			Adjusted R <sup>2</sup>	DW	Coefficient restrictions <sup>1</sup>	
		P-value		P-value		C(2)=-C(3), C(4)=0, C(5)=-C(6), C(7)=0	Ho: C(2)=0, C(3)=0, C(4)=0,				
		t	t-1	t-2	t					t-1	t-2
C(1)	C(2)	C(3)	C(4)	C(5)	C(6)	C(7)					
<b>Period</b>											
<b>1976.09-1987.12</b>	-0.073 (-0.98)	0.197 (5.39)	-0.186 (-3.38)	0.026 (0.62)				0.200	2.58	0.58	0.00
<b>Average annual inflation: 5.3%</b>	-0.061 (-1.02)	0.189 (4.88)	-0.164 (-3.77)	0.008 (0.26)	0.039 (4.94)	-0.043 (-4.08)	-0.001 (-0.08)	0.378	2.71	0.89	0.00
<b>Period</b>											
<b>1988.01-2000.09</b>	0.109 (1.36)	0.079 (1.53)	-0.088 (-1.99)	-0.056 (-1.39)				0.045	2.60	0.27	0.12
<b>Average annual inflation: 2.8%</b>	0.098 (1.70)	0.044 (1.24)	-0.060 (-1.66)	0.041 (-1.38)	0.024 (6.65)	-0.032 (-7.20)	0.002 (0.53)	0.383	2.70	0.32	0.20

<sup>1</sup> Wald test.

All the estimated coefficients for the contemporaneous standard deviation (C(2)) and the contemporaneous skew(class) (C(5)) are positive. The estimated coefficients for the first lag of these variables (C(3) and C(6) respectively) have more or less the same magnitude in absolute value as the coefficient of the corresponding contemporaneous variable, but have systematically the opposite sign. All the coefficients for the second lag are statistically non-significant (C(4) and C(7) respectively). The null hypothesis which imposes these features as restrictions on the coefficients was not rejected by the first Wald test presented in Table 3. This clearly illustrates the short-run nature of the estimated equations: while having a clear impact on the *change* in inflation, the standard deviation and the skewness do not have a *permanent* effect on inflation. This suggests that these relations illustrate the impact of relative price shocks on inflation, which according to economic theory is short-lived. This would imply that, in the estimated regressions, causality is running from the standard deviation and skew(class) to aggregate inflation. Overall, these results are in line with the predictions of menu cost models, outlined in the previous section. But do they really allow us to discriminate sticky from flexible price behaviour?

Finding a positive coefficient for contemporaneous skewness is not particularly relevant in this respect, as the traditional approach which interprets the classical mean-skewness correlation as evidence in favour of menu costs has recently been criticised on two grounds. In economic terms, Balke and Wynne (2000) showed that a flexible price general equilibrium model can generate this correlation as well. Statistically, Bryan and Cecchetti (1999) have attributed this result to the small sample bias characterising this correlation in the case of fat-tailed distributions<sup>16</sup>. However, the impact of the inflation regime on the estimated coefficients seems more relevant in this respect. Comparing the regression results for the two sub-periods shows that the impact of the prevailing inflation regime is moderate for the coefficients of skewness. The inflation regime matters much more for the impact of scale, as the coefficients for the standard deviation are substantially reduced from the first to the second sub-sample, to the extent that they are no longer significant. The second Wald-test presented in Table 3 clearly rejects the null hypothesis that all the coefficients of the standard deviation are zero in the first sub-sample, but not in the second sub-sample. Also, this result is in line with the predictions of menu cost models and suggests an important role for trend inflation, which in principle is not crucial in flexible price models.

### **6.3 Using robust measures: inflation affecting relative prices**

As is shown in Table 4, using the robust measure  $dskev(125)$  reveals a clear positive cointegration relationship between asymmetry and inflation. Each percentage point of additional (monthly) inflation tends to increase  $dskev(125)$  by 1.5 percentage points. Moreover, the constant term in the cointegration relation is very small both in statistical and in economic terms, suggesting that the chronic right skewness would disappear completely at zero inflation<sup>17</sup>. As such, the long term behaviour of the chronic right skewness measured by  $dskev(125)$  seems to be an integral part of the inflationary process itself.

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<sup>16</sup> The idea is that when samples are drawn from a fat-tailed distribution, they will tend to contain outliers (i.e. observations drawn from the left- or the right-hand tail), which influence both the mean and the classical skewness in the same direction. As a consequence, even for symmetric, but fat-tailed distributions it is very likely that skewed samples will be drawn. In that event the sampling may be an important source of the observed mean-skewness sample correlation, which does not necessarily characterise the distribution. Interesting in this respect is the observation that the coefficient of skewness is, for both periods, no longer significant when the regressions are estimated with the median instead of the mean as the inflation measure. Having no significant impact of  $skew(class)$  on the change in inflation beyond the first lag also illustrates the simultaneity in the mean-skewness correlation. See also the comments of Ball and Mankiw (1999) on this small sample bias.

<sup>17</sup> The sign of the constant term even suggests that the distribution would become slightly skewed to the left at zero inflation.

**Table 4 - Cointegration of the robust measures of skewness and scale with actual inflation (mean)**

	Dskew(125)	Dskew(250)	IQR
Lag number	11	11	11
Cointegration equation (CE) <sup>1</sup>			
Constant	0.0412 (0.73)	-0.0534 (-2.55)	-0.1790 (-5.39)
Mean	-1.5037 (-9.42)	-0.2579 (-4.38)	-0.8088 (-8.65)
Trace statistic <sup>2</sup>			
No CE	30.86	24.88	28.21
At most 1 CE	3.41	3.39	4.90

<sup>1</sup> Normalised cointegration coefficients, with the coefficient of the variable on top of each column normalised to 1.

<sup>2</sup> The 95 p.c. critical value is 19.96 for rejection of the null hypothesis of no cointegration equation and 9.24 for rejection of the null hypothesis of at most 1 cointegration equation. The cointegration rank and the specification of the deterministic components of the model were determined jointly on the basis of the general test procedure discussed in Johansen (1992). However, models having trends in the levels of the variables (constants in the VAR) were not considered.

Also for *dskew(250)*, the Johansen cointegration test reveals the existence of a positive cointegration relation with inflation, notwithstanding the fact that the ADF-test reported in Table 1 suggests that *dskew(250)* is stationary<sup>18</sup>. This measure of asymmetry tends to increase by 0.26 percentage points for each percentage of (monthly) inflation. The constant in the cointegration relation for *dskew(250)* suggests the existence of some form of chronic right skewness which is independent of the inflationary process. The impact of this constant is, however, relatively small compared to the impact of inflation. Indeed, each 0.2% of additional monthly inflation (approximately 2.4% on a yearly basis) increases *dskew(250)* by the same amount as that resulting from the constant. This suggests that, relative to what has been observed historically, a substantial part of the chronic asymmetry *measured by dskew(250)* disappears at zero inflation.

Summarising these results, for both robust measures of skewness there exists a *positive* long-run relation with aggregate inflation. Moreover, all of the asymmetry measured by *dskew(125)* and a substantial part of the asymmetry measured by *dskew(250)* would disappear at zero inflation. These findings confirm the result found for the mean-median relationship documented in Section 4. By uncovering this type of endogenous chronic right skewness, these results tend to provide more evidence in favour of menu cost models than

that found on the basis of the classical measures only. Moreover, these results are qualitatively different from those of the previous subsection on the impact of skewness. Relating robust measures of skewness to aggregate inflation in a cointegration framework overcomes, by the very long-run nature of such a relation, the small-sample bias, which typically characterises the contemporaneous correlation between inflation and skewness. The long-run nature of the relationship also suggests that monetary policy related factors, such as the balance of supply and demand in the economy and inflation expectations, rather than relative price shocks, are the driving force<sup>19</sup>. This also corroborates our presumption, put forward in the introduction, that downweighting outliers gives more weight to these monetary policy related factors. The finding that causality runs from aggregate inflation to relative prices fits less well in a flexible price story.

It is also important to note that this type of endogenous right skewness only reveals *apparent* downward nominal rigidity, as it would largely disappear at zero inflation. This differs substantially from the *exogenously* assumed downward nominal rigidity of Tobin (1972), which implies a negative, rather than a positive relation between aggregate inflation and asymmetry, essentially on the basis of the same argument as the one that we used in Section 3 to motivate a reduction of the left-hand tail. This difference is essential when the implications of our results for the optimal rate of inflation are examined, as is done in the next subsection.

#### **6.4 Grease and sand effects**

Tobin (1972) used the existence of exogenous downward nominal rigidity in labour markets to justify the statement that "*higher prices or faster inflation can diminish involuntary, disequilibrium unemployment, even though voluntary, equilibrium labour supply is entirely free of money illusion*" (see also Akerlof, Dickens and Perry (1996) for a more recent treatment). The idea is that inflation relaxes the constraint that downward rigidity imposes on the adjustment of relative wages. According to this argument, inflation "greases" the wheels of the labour market. Extrapolated to product markets, the "grease" argument essentially states that, due to downward rigidities, there is *insufficient* relative price variability at low inflation rates, and that increasing inflation therefore leads to a more flexible economy and hence to a better economic performance. Maintaining the metaphor, inflation does, however, also have *sand* effects. The sand effect emphasises the fact that higher inflation can lead to

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<sup>18</sup> In the case of conflicting results, the Johansen-test dominates the ADF-test, implying that *dskew(250)* also contains a unit root.



unnecessary relative price variability which disrupts the allocative role that relative prices play in a market economy. These sand effects can result either from the staggered nature of price changes in the case of price stickiness or from a Lucas-type of imperfect information, whereby higher and more variable inflation complicates the breakdown of observed nominal price changes in an aggregate and a relative component (Lucas (1972) and Lucas (1973)). Sand effects emphasise the excess of relative price variability at higher inflation rates and provide arguments in favour of a low inflation target.

Against this background, it is clear that the interpretation of the relation between inflation and relative price variability, which will be discussed below, is not straightforward, as a positive effect of inflation on relative price variability can be symptomatic of both grease and sand effects and, depending on which effect is dominant, leads to completely different policy implications as to the optimal rate of inflation. The complex nature of the link between inflation and relative price variability was already put forward in Fischer (1981).

Using the robust measure of scale IQR, we find indeed a strong positive cointegration relation between inflation and the dispersion of relative prices (see Table 4). The estimated impact of aggregate inflation on the dispersion of relative prices is again substantial, as a monthly inflation rate of 0.2 p.c. nearly doubles the dispersion that would prevail in the absence of inflation.

The literature on grease and sand effects<sup>20</sup> emphasises the fact that inter-market data, i.e. the dimension of inflation data we are examining, are particularly relevant for finding grease effects, while sand effects should even occur among prices of identical goods and hence should *also* be present in intra-market data. As to sand effects, the dimension of the data is also important when trying to distinguish between the two possible sources of sand effects. Inter-market data are the most relevant for imperfect information induced sand effects, while staggering should *also* affect intra-market data.

Cecchetti and Groshen (2000) also point to the fact that sand effects should be symmetrical, while grease effects may have an asymmetric impact, as they facilitate downward adjustment in particular. With this in mind, our result that the chronic right skewness is largely endogenous, rather than resulting from presumed downward rigidities, considerably reduces the scope for substantial grease effects. Hence, we interpret the inflation-IQR relation as

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<sup>19</sup> In sections 6.2 and 6.3 the causal direction was inferred from the short- and long run nature of the respective relationships. This approach differs from the one followed in Hall and Yates (1998), where the concept of Granger causality was used.

<sup>20</sup> See for instance Cecchetti and Groshen (2000).

sand rather than as grease, and we conclude that the cross-sectional properties of Belgian inflation data do not provide strong arguments against a price stability-oriented monetary policy, which targets a low inflation rate. Theoretically, this conclusion is in line with the implications of menu cost models. In Ball and Mankiw (1994), the endogenous asymmetry erases grease effects completely so that these authors find that, in their model, the optimal rate of inflation is zero, as it allows (i) a maximal reduction of sand effects and (ii) a reduction of the resources spent on changing prices. Empirically, our results appear in line with those reported in Yates (1999) where the absence of a negative relation between inflation and skewness is interpreted as evidence against the existence of specific downward rigidities. Suvanto and Hukkinen (2002) found that declining prices were rather frequent in Finland during the period 1995-2001 and conclude from this that the assumption of downward rigidity of nominal prices cannot be generalised to all situations.

## **7. AN UNBIASED MEASURE OF CORE INFLATION**

In the *statistical* approach to core inflation, robust estimators of location are proposed on the basis of the fat-tailed nature of the distribution of price changes. Section 3 documented this characteristic of the data for Belgian 1-monthly price changes. On the basis of the results presented in Sections 5 and 6 we found, moreover, that the distribution is not only fat-tailed but also skewed to the right. Given the chronic right skewness, symmetrically constructed robust estimators tend to underestimate aggregate inflation, as has been documented in Section 4 for the median. Several asymmetrically constructed robust estimators have been proposed to overcome this problem<sup>21</sup>. They all have in common that the robust estimator is centred around a higher percentile than the 50% percentile and that this central percentile is essentially obtained by averaging over time some indicator of skewness. In so doing, this approach assumes a degree of chronic right skewness which is stable over time. However, by using robust measures we were able to uncover the time-varying and endogenous nature of the chronic right skewness in Belgian 1-monthly price changes. This characteristic of the data presumably explains why some instability in the central percentile was found in Aucremanne (2000), when it was calculated for different periods, each of which was characterised by a different average inflation rate.

The use of robust estimators of location as measures of core inflation has also been motivated in *economic terms* by Bryan and Cecchetti (1994) on the basis of a variant of the

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<sup>21</sup> See for instance Chatellain, Odonnat and Sicsic (1996), Roger (1997), Kearns (1998), Aucremanne (2000), Marques and Mota (2000), Meyler (1999 and 2001).

(symmetric) menu cost model of Ball and Mankiw (1995). Since then, the economic rationale of robust estimators has, however, been criticised by several authors, mainly for two reasons.

First, as was emphasised above, the short run mean-skewness correlation was found not to provide sufficient evidence in favour of menu costs. Second, even if menu costs or some other form of price stickiness indeed exist, one should be careful in downweighting outliers. In that event, not only relative price shocks, but also the inflationary process itself affects the distribution of relative prices. Hence, observing an outlier can be either the result of a pure relative price shock, in which case downweighting this observation seems justified. Alternatively, it can also correspond to a price change induced by aggregate demand fluctuations, which are more relevant from a monetary policy perspective. In a context of infrequent adjustment of prices also this type of price change will typically be large. This point was made in several papers, such as Zeldes (1994), Bakshi and Yates (1999) and more recently in Aoki (2001) and in Mankikar and Paisley (2001). Finding an endogenous degree of chronic right skewness illustrates that this problem is not a pure theoretical issue, but has a clear practical relevance as well.

Against this background, it is our view that the chronic asymmetry found in the Belgian inflation data provides additional evidence in favour of menu cost models, on top of the evidence obtained on the basis of the classical mean-skewness correlation only. As such, this reinforces the economic rationale for the use of robust estimators of location, but at the same time it complicates their construction as the endogenous nature of the chronic right skewness should be taken into account. These economic arguments tend to put a lot of weight on the absence of bias in the estimator, whereas initially the focus was much more on the efficiency gain. In the case of skewed distributions there is indeed a trade-off between efficiency and absence of bias. This point is illustrated, for instance, in Le Bihan and Sédillot (1999). Below we present a robust estimator of location based on the median as a measure of core inflation, which is by construction unbiased and is nevertheless more efficient than the mean.

Our starting point is the first condition of Marques, Neves and Sarmento (2000) for an ideal core inflation measure, which provides an operational definition for an unbiased estimator. According to their terminology, a core inflation estimator is considered unbiased if cointegration *with unitary coefficient* is found between actual and core inflation. In the previous sections, we provided evidence on two occasions that this was not the case for the median. First, in Section 4 we found a coefficient which is substantially less than one (0.5934,

to be precise) in the cointegration equation linking the median and the mean. Second, in Section 5 it was found that the mean-median difference *meme* is a non-stationary variable, even in a test equation incorporating a constant term. The latter finding illustrates clearly that fixing the bias is a more complicated issue than merely correcting for a constant term, as it is sometimes suggested.

Our strategy is to link *meme* in a cointegration relation to one of the robust measures of (chronic right) skewness and to exploit this cointegration relation subsequently to construct an unbiased estimator of core inflation. In so doing, we found that *meme* was cointegrated with *dskew(125)* in the following way:

$$meme - 0.3057 * dskew(125) \cong I(0)$$

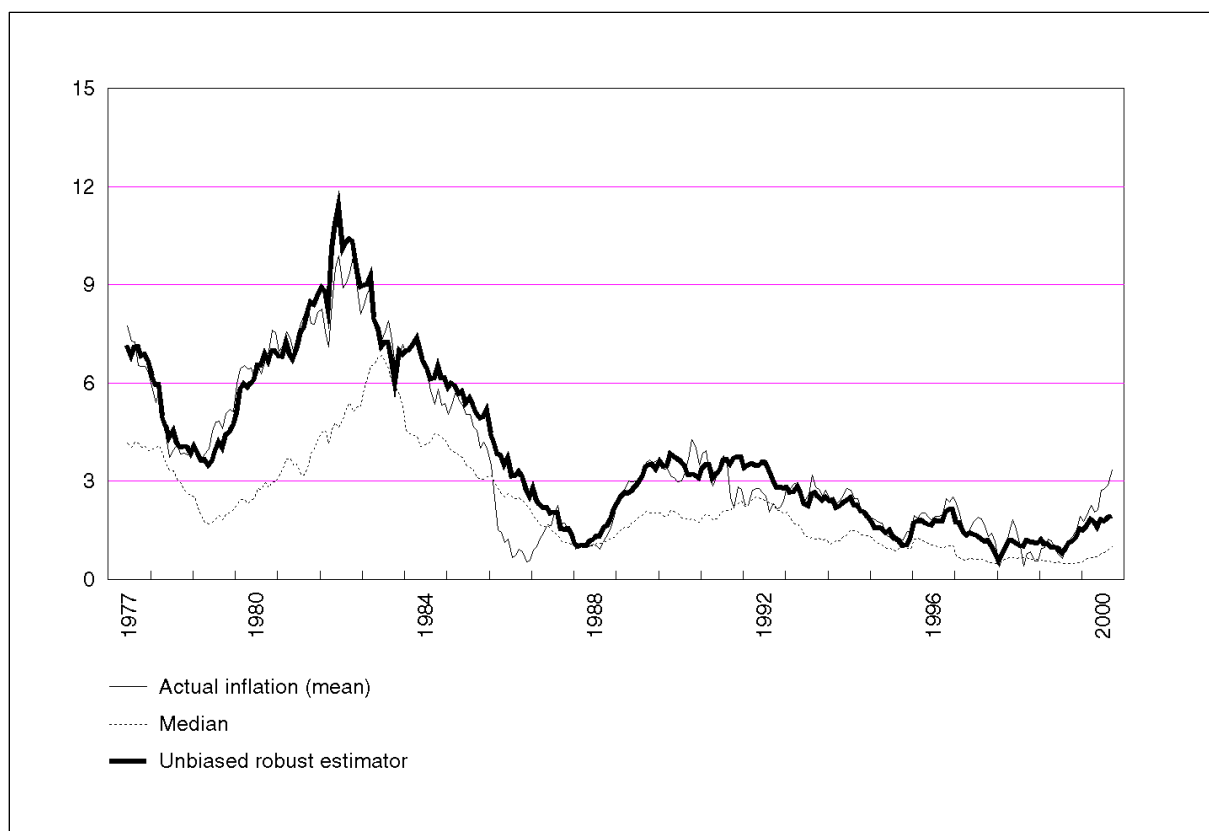
By construction, this implies that the variable  $median + 0.3057 * dskew(125)$  is cointegrated with a unitary coefficient with the mean (or, equivalently, with actual inflation). In consequence, this variable can be considered as an unbiased estimator of core inflation. The rationale of this approach is that, by incorporating *dskew(125)*, the median is corrected upwards in order to take account of the chronic right skewness of the distribution, which is, as was emphasised above, time-varying and largely part of the inflationary process itself. Hence, the fraction of the mean-median difference which is attributable to the chronic right skewness is incorporated in the core inflation measure, while the latter ignores only the fraction of the mean-median difference which is influenced by what is happening far down in the tails.

As inflation data are often expressed as year-on-year growth rates, we decided to express the core inflation measure considered as 12-monthly price changes as well, by simply compounding the original 1-monthly price changes of the last 12 months. A similar approach - robust estimation on the basis of 1-monthly price changes and subsequent smoothing by compounding it over the last twelve months - is applied in the Bank of England (see Mankikar and Paisley (2001)) and was also followed in Aucremanne (2000). All in all, the proposed core inflation measure of period  $t$ , expressed in terms of 12-monthly price changes, corresponds to:

$$\prod_{s=0}^{11} [1 + (\text{median}_{t-s} + 0.3057 * dskew(125)_{t-s})] - 1 \quad (7.1)$$

**Figure 7 - Unbiased robust estimator of core inflation**

(compounded results, percentage changes)



In Figure 7 the proposed estimator of core inflation is plotted, together with the 12-monthly inflation rate and the median, which has been compounded in a similar way to obtain 12-monthly price changes as well. A comparison of the proposed core inflation measure with the median clearly illustrates the upward, but *time-varying*, correction made to the median, which makes the new estimator unbiased. But it also shows that the proposed core inflation measure is less smooth over time than the median, thus clearly illustrating the trade-off between absence of bias and efficiency. Despite this loss in efficiency relative to the median, the proposed core inflation measure remains smoother than actual inflation<sup>22</sup>. Of course, smoothness is a desirable property of a core inflation measure. However, as the smoothness of the shocks underlying the (core) inflationary process, as well as the smoothness of their transmission into prices, is unknown, it is not clear a priori exactly how smooth the core inflation measure should be. The loss of smoothness compared to the median is therefore not necessarily a drawback. Indeed, it seems that it is compensated for by the fact that the proposed core inflation measure is more deeply embedded in economic theory, by trying to

abstract only from relative price changes which are due to pure relative price shocks, but not from those which, in a context of infrequent price adjustment, are the result of the ongoing inflationary process itself. The proposed core inflation measure deviates particularly from observed inflation in 1976, when oil prices dropped sharply and in 2000 when oil prices increased substantially.

## 8. CONCLUSIONS

In this paper we have studied the properties of the distribution of Belgian consumer prices for the period June 1976 - September 2000, by using both classical and robust measures of location, scale, skewness and tail weight. The robust measures of tail weight showed clearly (i) the fat-tailed nature of the distribution, (ii) the short run volatility of the tail weights, (iii) the tendency for the right-hand tail to dominate the left-hand tail, which points towards chronic right skewness and (iv) the increase over time of the left-hand tail and the more stationary behaviour of the right-hand tail. The presence of outliers makes the classical measures of location, scale and skewness very volatile relative to the robust alternatives which have been used in the paper. Moreover, in the case of scale and skewness, the influence of the outliers on the classical measures erased any long run time profile, whereas for their robust alternatives a stochastic trend was discovered, as was the case for the inflationary process itself. Finally, the classical measures of skewness hardly revealed any form of chronic right skewness, whereas the robust measures clearly bring this characteristic of the data, as well as its time-varying nature, into the picture.

In the econometric part of the paper we discussed the relation between the distribution of relative prices and aggregate inflation. According to models with nominal rigidities - sticky price models in general, or more particularly, menu cost models -, the causality of this relation can run in two directions and both directions have been discussed in the paper. In particular, we were able to describe (i) the short run impact of relative price shocks on aggregate inflation by using the classical measures of scale and skewness and (ii) to reveal the impact of the inflationary process on the distribution of observed relative prices by relating robust measures of scale and skewness to aggregate inflation in a cointegration framework.

Our main empirical results in this respect are threefold. First of all, we found a short run impact of the skewness of observed relative prices on aggregate inflation, in line with Ball and

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<sup>22</sup> As in Aucremanne (2000), smoothness has been measured here as the standard deviation of the *first difference* of the estimator considered.

Mankiw (1995). Second, it was found that the short run impact of the dispersion of observed relative prices on inflation was substantially lower in the sub-sample starting in 1988 than in the pre-1988 sub-sample, suggesting that the prevailing monetary policy regime has a substantial effect on this coefficient. This result is in line with the model of asymmetric price adjustment of Ball and Mankiw (1994). Third, our further empirical results highlight the endogeneity of the asymmetry in the distribution of Belgian consumer prices. Indeed it was found that the chronic right skewness was *positively* cointegrated with aggregate inflation. Moreover, we found little evidence of chronic right skewness not dependent on the inflationary process, suggesting that the asymmetry would disappear at zero inflation. Overall, these results are in line with the predictions of menu cost models and, importantly, they go beyond the contemporaneous mean skewness correlation. The latter has been criticised recently for not providing *sufficient* evidence in favour of menu costs on its own. The results presented are symptomatic of nominal rigidities, in the sense that prices are adjusted infrequently. However, they do not point in the direction of specific downward rigidities, other than those endogenously generated by the prevailing inflation rate.

These results have three important policy implications.

First, as to the *transmission of monetary policy*, they suggest that menu costs can be an important factor behind the short run non-neutrality of monetary policy and the observed output and inflation dynamics. It should be taken into account that these models generate asymmetric output effects in consequence of the asymmetric nature of price adjustment to which they lead endogenously. As prices tend to adjust more slowly in the case of a negative shock than in the case of a positive shock, output is affected more when monetary policy is tightened than in the opposite event. According to menu cost models, this asymmetry depends on the prevailing inflation rate: it is more pronounced at higher rates of inflation and disappears when inflation is equal to zero. In this respect it is interesting to remember that Peersman and Smets (2001) found larger monetary policy effects on output in recessions than in booms, not only in the euro area as a whole, but also in Belgium in particular. They did not attempt to distinguish between the various theories explaining this asymmetry. According to our results, menu costs could be a part of the story, at least for Belgium.

Second, as to the *design of core inflation measures*, providing evidence in favour of menu cost models reinforces the economic rationale for the use of robust estimators of location, but at the same time it complicates their construction as the endogenous right skewness of the distribution of relative prices should be taken into account. These economic arguments tend to put a lot of weight on the absence of bias in the estimator. We have proposed an estimator which has this property, as the estimator takes the time-varying degree of chronic right skewness explicitly into account.

Third, as to the *optimal rate of inflation*, the chronic right skewness found in the data provides no argument against price stability, as it reveals only *apparent* downward nominal rigidity in product markets, which is an endogenous response of optimising price setters and which would disappear when targeting a zero inflation rate. The policy implication of this result contrasts sharply with the *exogenously* assumed downward rigidity of Tobin (1972), which would justify targeting a positive inflation rate in order to facilitate the adjustment of relative prices. In the absence of *specific downward* nominal rigidities, finding that aggregate inflation affects the distribution of relative prices rather provides an argument in favour of price stability, as it allows the impact of monetary policy on relative prices to be minimised. Clearly, this impact disrupts the allocative role prices play in a market economy. In the menu cost model of Ball and Mankiw (1994), the optimal rate of inflation is indeed equal to zero. Evidently the analysis presented here does not take into account other factors which could justify a low but positive inflation target, such as downward nominal rigidities in labour markets, the zero bound for nominal interest rates and the presumably positive bias in aggregate inflation measurement. In any case, the cross-sectional properties of Belgian inflation data do not provide strong arguments against a price stability-oriented monetary policy, such as the one pursued by the Eurosystem.

For future empirical work, it seems interesting to test whether the implications of menu cost models are also verified for detailed datasets which have both an inter-market and an intra-market dimension. The results of this paper were indeed obtained on the basis of readily available inflation data which, in the cross-section, only have an inter-market (or sectoral) dimension. It is, however, the standard view that intra-market data are particularly relevant for uncovering the impact of monetary policy on the distribution of relative prices.



**Description of the 60 items breakdown of the CPI used**

<b>Product name</b>	<b>Product code</b>
Bread and cereals	1.1.1 (*)
Meat	1.1.2
Fish	1.1.3
Milk, cheese and eggs	1.1.4
Oils and fats	1.1.5
Fresh fruit	1.1.6 part
Fresh vegetables, excluding potatoes and other tubers	1.1.7 part
Potatoes and other tubers	1.1.7 part
Tinned fruit and vegetables	1.1.6/7 part
Sugar	1.1.8 part
Jam, honey, syrups, chocolate and confectionery	1.1.8 part
Food products n.e.c.	1.1.9
Coffee, tea and cocoa	1.2.1
Mineral waters, soft drinks and juices	1.2.2
Alcoholic beverages	2.1
Tobacco	2.2 (*)
Clothing	3.1
Footwear, excluding repairs	3.2 part
Footwear, repairs	3.2 part
Actual rentals for housing	4.1
Regular maintenance and repair of the dwelling	4.3
Other services relating to the dwelling	4.4 (*)
Electricity	4.5.1
Gas	4.5.2
Liquid fuels	4.5.3
Solid fuels	4.5.4
Furniture, furnishings and decorations, carpets and other floor coverings and repairs	5.1
Household textiles	5.2
Major household appliances whether electric or not and small electric household appliances	5.3.1/2
Repair of household appliances	5.3.3
Glassware, tableware and household utensils	5.4
Tools and equipment for house and garden	5.5
Non-durable household goods	5.6.1
Domestic services	5.6.2 part
Home care services	5.6.2 part
Medical and pharmaceutical products	6 part
Therapeutic appliances and equipment	6 part
Medical services	6 part (*)
Hospital services	6 part
Purchase of vehicles	7.1

<b>Product name</b>	<b>Product code</b>
Spare parts and accessories	7.2.1
Fuels and lubricants	7.2.2
Maintenance and repairs	7.2.3
Other services in respect of personal transport equipment	7.2.4 (*)
Transport services	7.3 (*)
Communications	8 (*)
Equipment for the reception, recording and reproduction of sound and pictures; Photographic and cinematographic equipment and optical instruments; Data processing equipment	9.1.1/2/3
Other major durables for recreation and culture; Games, toys and hobbies; Equipment for sport, camping and open-air recreation; Recording media for pictures and sound; Gardening;	
Pets	9.1.4/5/6/7/8
Repair of equipment and accessories for recreation and culture	9.1.9
Recreational and cultural services	9.2
Newspapers, books and stationery	9.3
Package holidays	9.4
Education	10 (*)
Hotels, cafes and restaurants	11
Hairdressing salons and personal grooming establishments	12.1.1
Appliances, articles and products for personal care	12.1.2
Personal effects n.e.c.	12.2
Insurance	12.4 (*)
Banking services n.e.c.	12.5 (*)
Other services n.e.c;	12.6

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