NOTES D'ÉTUDES

ET DE RECHERCHE

IMPLEMENTING AND INTERPRETING INDICATORS OF CORE INFLATION: THE FRENCH CASE

Hervé Le Bihan and Franck Sédillot

December 1999

NER # 69

BANQUE DE FRANCE

DIRECTION GÉNÉRALE DES ÉTUDES

DIRECTION GÉNÉRALE DES ÉTUDES
DIRECTION DES ÉTUDES ÉCONOMIQUES ET DE LA RECHERCHE

IMPLEMENTING AND INTERPRETING INDICATORS OF CORE INFLATION: THE FRENCH CASE

Hervé Le Bihan and Franck Sédillot

December 1999

NER # 69

Implementing and interpreting indicators of core inflation: the case of France

Hervé Le Bihan*and Franck Sédillot†

December 1999

Abstract

This paper presents a comparison of alternative indicators of underlying or "core" inflation in the French case. Four broad measures are considered and implemented. The first two are inflation excluding food and energy, and the trimmed inflation indicator. We then implement two methods relying on time-series models: the Dynamic Factor Index and the structural VAR approach. Each indicator stresses on a particular type of shock on the inflation rate. Combining the various indicators conveys valuable information for appraising short term inflation developments. Nevertheless, even in the case of the structural VAR, the theoretical interpretation of core inflation is not straightforward, lacking an explicit representation of monetary policy.

Keywords: Core inflation; Trimmed mean estimators; Dynamic Factor Index.

*Corresponding author. Banque de France, 41-1391 DGEI-DEER, 31 rue Croix des Petits Champs, 75001 PARIS, e-mail herve.lebihan@banque-france.fr. Telephone: +33 1 42 92 49 87.

†Banque de France, 41-1376 DGEI-DEER, e-mail franck.sedillot@banque-france.fr. We are grateful (in alphabetical order) to S. Avouyi-Dovi, O. De Bandt, P. Jacquinot, E. Jondeau and P. Sicsic, as well as participants to a seminar at the Banque de France, for helpful comments. We also thank P. Jacquinot and **Pierre** Sicsic for making their computer programs available to us. Any remaining errors are of course our own responsibility. The views expressed in this paper are not necessarily those of the Banque de France.

1 Introduction

Indicators of underlying or "core" inflation have been increasingly used in order to analyze and implement monetary policy since the beginning of the 90's. Most central banks pursuing an inflation target are currently computing indicators of underlying inflation. The main examples include Canada, Australia, New Zealand, the United Kingdom.[1] Such indicators are also considered by central banks operating under various other monetary arrangements.[2]

Underlying inflation, though, does not have a unique widely accepted theoretical definition. While the general principle is to remove the impact of transitory shocks on the inflation rate, a wide range of empirical indicators are proposed in the literature, based on very contrasting hypothesis and methods.[3] This undoubtedly reflects the fact that the inflation rate may be affected by macroeconomic and sectorial shocks of different sorts.

This paper proposes applications of the main indicators in the case of France. Given the contrasting assumptions, statistical frameworks and datasets on which the alternative measures of core inflation are based, it is difficult to find a single expository framework. We therefore review in section 2 to 5 each of the following indicators in turn. Inflation excluding special factors (such as food and energy), which rely on slightly disaggregated price indices, is presented in section 2. The following section discusses the trimmed mean estimators, which are computed from the cross-sectional distribution of individual prices. The Dynamic Factor Index (DFI) approach, proposed by Bryan and Cecchetti (1993) is implemented in section 4. The DFI relies on a times series unobserved component modelisation of the disaggregate prices indexes and is estimated through the Kalman filter. Last the structural VAR approach relates aggregate inflation to other macroeconomic series in a multivariate framework. In each case we discuss the interpretation of the indicator. Section 6 draws some implications for evaluation of monetary policy and concludes.

2 Exclusion of "special factors"

2.1 Definition

In the first approach, core inflation is computed by systematically excluding some items from the price index. The items excluded are chosen on the ground that their prices show an unpredictable time patterns. Such types of measures are supposed to filter out the -at least first round- effects of transitory shocks originated in the erratic behavior of some particular prices. The computation is usually done at a rather low level of disaggregation of the price index. In the United States, since Blinder (1982), inflation "excluding food and energy" has been a widely used index, particularly in the monetary policy diagnosis. In the case of France, the Insee (1996) releases a

[1] See Bakhshi and Yates (1999), Laflêche (1998), Roger (1998), Quah and Vahey (1995).

[2] See for instance Bryan and Cecchetti (1994) for the US, Alvarez and Matea (1999) in the case of Spain. Wynne (1999) reviews the issue in the context of the European Central Bank.

[3] Detailed surveys are provided by Roger (1998) and Wynne (1999).

similar index which nevertheless excludes more numerous items and is corrected for changes in indirect taxes. Excluding the impact of indirect taxes is motivated by considering that indirect taxes shock only have a one shot effect on the inflation rate. In some countries excluding items is motivated by economic policy issues. In the UK the RPIX index -which excludes mortgage interest rates payments (MIPs), stands as the benchmark index to define the inflation target of the Bank of England. Indeed while a rise in the interest rate is supposed to dampen inflationary pressures, it will actually result, to a certain extent, in pushing upwards the standard RPI measure due to the weight of interest payments by households. Failing to exclude MIPs is doomed to confuse assessment of monetary policy.

2.2 Application

Since the Insee index is unavailable previous to 1990, our application relies on an analogous index we computed over a longer period. This underlying inflation indicator is computed from disaggregated data excluding the following items : fresh food, energy, tobacco and public utilities. The underlying inflation roughly includes industrial manufactured goods, private sector services, and processed food. Figure 1 displays the underlying inflation series “backcast” until 1970. It emerges from Figures 2 that energy and fresh food prices stand out as much more volatile than the other broad aggregate.

One would expect the underlying inflation to significantly diverge from actual inflation during oil shocks. While this is true during the second oil shock and the 1986 counter-shock, core inflation and observed inflation remain close to each other during the first oil shock (Figure 1). This is due to the fact that in 1974 food prices in France grew weakly, which compensate for the spike in energy price with respect to the gap between that particular measure and actual inflation (Chatelain, Odonnat and Sicsic 1996).

Indicators excluding food and energy have obviously several drawbacks. First from a theoretical point of view it is not clear that monetary policy should ignore surges in inflation originated in oil shocks. It should be the case only if such shocks are not persistent. A related issue is how to deal with second round effects of food or energy price shocks operating say through wage pressure. Should they be excluded as well? From a more empirical point of view, using this type of indicators imposes a very strong a priori discrimination between sectors. It amounts to assuming that energy or food price never convey valuable information about core inflation, while conversely that every other price item in any situation does. The trimmed mean indicators are designed to address this last issue.

3 Limited influence estimators : median inflation, 60^{th} percentile, and the trimmed mean

3.1 Definition

Indicators reviewed in this section use limited influence estimators. The basic approach is to remove on a date by date basis the items of the CPI which show the most extreme variation. Such an approach requires an high level of disaggregation of the price index. Its advantage over the previous method is to avoid assuming that food or energy prices never contain relevant information while allowing for correcting for much wider set of sectorial shocks.

This kind of estimators was introduced by Bryan and Pyke (1991) who suggested picking the median item in the cross sectional distribution of price change to capture core inflation. A more general class of estimators is the trimmed mean (Bryan and Cecchetti 1994). This latter indicator is computed by taking the mean of the cross section distribution of price, after removing a part of each of the tails of the distribution. A formal definition of trimming means can be given using the following notations. Let $\{\widetilde{\pi}_1, ..., \widetilde{\pi}_n\}$ be the sectorial inflation rates sorted in ascending order, $\{\omega_1, ..., \omega_n\}$ their respective weights in the total price index, $W_i = \sum_{j=1}^{i} \omega_j$ the cumulated weights and I_α collects the following indices $I_\alpha = \left\{ i \mid \frac{\alpha}{100} < W_i < (1 - \frac{\alpha}{100}) \right\}$. The $\alpha\%$ trimmed mean is then defined as:[4]

$$\overline{\pi}_\alpha = \left(\frac{1}{1-2\frac{\alpha}{100}}\right) \sum_{i \in I_\alpha} \omega_j \widetilde{\pi}_i$$

Estimators in the trimmed mean class include average inflation as a particular case, which is the 0% trimmed mean $\overline{\pi}_{0\%}$, and the median inflation, which corresponds to the 50% trimmed mean $\overline{\pi}_{50\%}$.

3.2 Statistical motivation for trimmed means

A motivation for trimmed mean estimators can be based on the following statistical sampling arguments. Assume that on a given date an individual price observation is a particular draw from a distribution whose (unknown) population mean is the core inflation rate prevailing at that date. Assume furthermore that the distribution has fat-tail, which is consistent with evidence on American data. Then limited influence estimators of the mean, such as the trimmed mean, will have a smaller variance than the usual empirical mean. Bryan, Cecchetti, and Wiggins (1997) illustrate this property performing Monte Carlo experiments. They show that the higher the kurtosis of the distribution, the more the distribution should be trimmed to reach an optimal estimator within this class (see Appendix 1 for further details).

Choosing the optimal trim raises some conceptual problems. First, the (non-normal, leptokurtic) shape of the true distribution of individual prices changes is not

[4] We follow Bryan, Cecchetti and Wiggins (1997) convention that the $\alpha\%$ trimmed mean is obtained by removing $\alpha\%$ from *each* tail of the distribution. Note that this convention does not allow for asymmetrical trims.

known.[5] Bryan, Cecchetti and Wiggins (1997) assume that a moving average of the aggregate inflation is a good approximation for the unknown mean of the distribution. Using such an estimator, they are then able to perform bootstrap experiments. The experiments rely on the deviation between each individual price change and the 36 months moving average, and result in defining the "trimmed mean estimator" with minimal variance. Implicit in such a procedure is nevertheless a core inflation measure based on time series data. If a moving average of inflation is viewed as a good indicator of core inflation, restricting the data set to contemporaneous data for estimating core inflation will probably lead to an inefficient estimator.

In addition for the French case, the distribution of price changes is skewed to the right. The optimal trim need not be symmetrical and defining the optimal trim may be more complex. Indeed the draws of individual price changes with large absolute values will most often come from the right side of the distribution, so minimizing the variance should imply trimming a larger percentage from the right side of the distribution. But such an asymmetric trim mean would then exhibit a systematic deviation from the mean of the distribution, i.e. underlying inflation would on average be lower than actual inflation (see Appendix 1).

Given the difficulty raised by the definition of an optimal trim, in the empirical application we thus set the degree of trimming on *a priori* grounds.

3.3 French data: the need for an asymmetric trim

Following Bryan and Pike (1991), Chatelain, Odonnat and Sicsic (1996) compute a median inflation indicator in the case of France, using a 44 items database covering the 1972-96 period. The authors point out the fact that on French data cross-sectional distribution of price changes is not only leptokurtic, but also skewed to the right (Figure 3).[6]

This result contrasts with the evidence on US data (Bryan and Cecchetti 1994) where the cross sectional distribution of price change is broadly symetrical on average. Chatelain, Odonnat and Sicsic (1996) therefore choose the 55th percentile as an indicator of core inflation so as to obtain a measure of core inflation that on average equals the observed rate of inflation.[7]

We conduct a similar exercise here, although using a more detailed database.[8] This dataset confirms the finding that median inflation is lower than actual inflation. It appears that the percentile corresponding on average to the mean rate of inflation is the 60th (Figure 3). Comparison with the results of Chatelain, Odonnat and Sicsic (1996), suggests the degree of skewness is an increasing function of the degree of disaggregation of the price data.

We also compute a trimmed mean indicator. Because of the skewness of the distribution (see Appendix 1) choosing an symmetric trim would bring out a measure of

[5] Should this be the case, maximum likelihood estimation might apply.

[6] This asymetry is high from 1970 to 1980 but markedly decrease after 1983. Note also that the pattern of the historical skewness is unchanged when oil products are excluded.

[7] Roger (1998) chooses the 57^{th} percentile in the case of New Zealand.

[8] See Appendix 3 for the description of data sources.

core inflation lower on average than actual inflation. We choose an asymmetric trim so as to conform to our conventional identifying assumption that core inflation equals inflation on average. Furthermore we select a small degree of trimming (15% on the whole). This is consistent with evidence from other countries. Bryan, Cecchetti and Wiggins (1997) find that trimming 9% off each tail yields the optimal trim, while the survey in Alvarez and Matea (1997) reports 5% to 20% trimmed means across various countries. Our choice, which is rather arbitrary and mainly aimed at illustrative purpose, is to trim 10% in the left tail of the distribution and 5% in the right tail. Figure 4 shows that the chosen trim mean is quite close to the 60th percentile in general, and that both measures are smoother than the excluding food and energy index.

While Figure 4 reports year on year increases (i.e. a 12 months moving average of the monthly increase), it should be stressed that the trimmed mean indicators of core inflation are also designed to analyze monthly data, and to remove high frequency noise. A recent example in the US case is the monthly increase in the CPI of 0.7% in April 1999, following months of moderate increase of 0.1%. Considering the median inflation indicator which rose only to 0.3 (while inflation excluding food and energy was 0.4%) gives a less alarming picture, pointing to a more transitory pick-up.[9]

In Figure 4, trimmed mean indicator behaves much like excluding food and energy inflation over the first half of the sample. In 1987 an upswing of CPI inflation occurred: this event corresponds to the liberalization of many services prices which were previously administered (car repairs, cafés, hairdressers ...). This upswing is eliminated by the various "limited influence" estimators. This illustrates an attractive feature of limited influence estimators: they allow to remove the impact of noisy disturbances even on items that experience irregular movements only in very rare occurrences. Conversely we can notice that trimmed means closely match the marked decline in observed inflation in 1982. This decline was in fact enforced by prices and wages freeze in 1982. While part of that decrease was transitory, that fact can hardly be detected by a trimmed mean, since roughly every price was affected in the same manner.[10]

3.4 Can the time series dimension be neglected?

Trimmed means are purely static estimators, drawing from the cross-section dimension only. This has some advantages for practical purposes. Since the methodology is implemented on a date by date approach, those measures of core inflation are not affected by the release of new data. This is stressed by Wynne (1999) as a major advantage with respect to the communication of a central bank to the public. Furthermore given the high degree of disaggregation, trimmed means can be implemented without major problems when the nomenclature of prices undergoes a modification, which is not the case for the previous method. At a more fundamental level, trimmed means avoid to specify a time series process for the inflation rate and thus make the

[9] Median inflation is available from the Cleveland FRB Website.

[10] Note that, quite surprisingly, every core inflation measure calculated in this paper yields a decrease in core inflation simultaneous to the price-wage freeze in 1982.

measure robust to structural breaks.

Nevertheless it is not clear that the time series dimension can be escaped. First as discussed above, computation of the optimal trimmed mean relies on a time series moving average of CPI inflation. If we stick to the statistical model motivating trimmed means, the CPI inflation is of no particular interest. Indeed the observations of individual price changes are supposed to have the same population mean - core inflation. CPI inflation is only a particular (inefficient) estimator of core inflation and should not be seen as a reference. In contrast the procedure for computing optimal trimmed means suppose that in the long run average CPI inflation and core inflation coincide, which reflects macroeconomic and time series perspective.

Furthermore in the case of France the identifying assumption that average core inflation should equal average inflation was introduced because we observe a systematic skewness to the right in a time series of data. If we observed a single draw in the distribution of prices, skewness could be considered as a sampling property of a leptokurtic distribution (Bryan, Cecchetti and Wiggins 1997). This underlines that investigating the time series dimension of the data is hardly avoidable.

3.5 Assessing economic motivation for trimmed means

Using sectorial data to evaluate a measure of core inflation (what is done by excluding items or computing trimmed mean inflation or the DFI - see next section) is only to be meaningful if relative price shocks might affect the general price level. That is, if the pure classical model of price level determination does not apply. Furthermore removing these “noises” on the general price level is only relevant as far as they are transitory, so the interpretation relies on short run departures from a classical model of the price level.

The assumption that sectorial prices increase are the sum of two independent terms, core inflation and a relative price shock, might be translated as :

$$\pi_i = \pi_c + \epsilon_i \tag{1}$$

Where π_c, π_i and ϵ_i are respectively core inflation, observed individual price variation and a relative price shock in sector i independent from π_c. For any computation of core inflation to be of interest, it must be the case that aggregate observed inflation (i.e. the average of π_i's) is different from π_c.

Bryan and Cecchetti (1993) draw on the Ball and Mankiw (1992) model, which relies on relative prices nominal rigidities to show that inflation may deviate from its equilibrium value. In that model π_c is expected inflation at the beginning of a period. Equation (1) is interpreted as reflecting the desired price increase of firm i after a temporary zero-mean idiosyncratic shock ϵ_i on the cost or demand of firm i has occurred.[11] Due to menu costs some of the firms (those experiencing of low value of ϵ_i) do not adjust their relative prices and stick to a π_c price change. Those firms for which the cost of not adjusting prices exceed the menu cost will move

[11] The model is discussed in terms of firms. For the empirical analysis we assume that the properties translate to sectoral prices.

their price by π_i. Therefore the aggregate observed inflation will be $E(\pi_i^{actual}) = \pi_c + E\epsilon_i I_{(\epsilon_i > c)}$ (where c is the menu cost and $I_{(.)}$ the dummy variable taking value 1 if $\epsilon_i > c$ and 0 otherwise). Aggregate inflation will still equal core inflation if the distribution of shock ϵ is symmetrical. But it will deviate from core inflation if the distribution is asymmetrical. If for instance the distribution of the shock is skewed to the right, more firms will move their price upwards than downwards, resulting in an inflation rate superior to the core inflation rate π_c. In such a case trimmed mean estimator will, unlike the CPI, deliver a correct extraction of the core inflation rate. The model provides a motivation for trimmed means estimator.[12]

A remaining issue is how is the core inflation rate π_c determined. Bryan and Cecchetti (1994) propose an interpretation in terms of "monetary inflation". Since monetary policy is assumed to determine the inflation rate in the long run it can be stated that :

$$\pi_c = \dot{m} \tag{2}$$

where $\dot{m}$ is money growth, assuming the central bank controls a monetary aggregate.

This interpretation raises some questions. First this bear on the very restrictive hypothesis that only relative price shocks may push observed inflation away from equilibrium inflation.[13,14] Which shock shall be considered as non-monetary is indeed rather arbitrary. Cecchetti (1996) suggests indirect taxes shocks, and even exchange rate shocks should be considered as transitory disturbances to be removed from the monetary inflation. Furthermore, the link between the theoretical determinants of core inflation and the retained measure is rather loose, since monetary aggregates or interest rates are not included in the empirical estimation. Last the notion of monetary inflation requires clarification if the central bank is viewed as reacting to shocks in the economy, and particularly shocks on the inflation rate.

[12] The Ball and Mankiw (1992) also predicts a positive correlation between skewness in the price distribution an the average rate of inflation. The feature is a stylised fact of U.S. data already highlighted by Vining et Elwertowski (1976). Ball and Mankiw (1992) consider it as evidence of the presence of *menu cost.* See Bonnet, Dubois et Fauvet (1998) for an application on French data.

[13] For instance Eckstein (1981) already proposed another decomposition of inflation into several shocks :

$$\pi = \pi_c + \pi_d + \pi_s$$

where π_c is core inflation, π_d the demand component of inflation and π_s the part due to supply shocks (oil price, minimum wage, exchange rates ...).

[14] Bakhshi and Yates (1999) discuss the impact of demand shocks on trimmed means in the context of menu costs.

4 The *Dynamic Factor Index (DFI)*

4.1 Definition

The DFI approach, seemingly less often implemented, consists in extracting the unobserved common factor of several individual price series. It thus makes use both of the cross-sectional and the time-series dimension of prices data. Following Stock and Watson (1991) DFI have traditionally been used to analyze real macroeconomic data in order to build coincident indicators of the business cycle. Bryan and Cecchetti (1993) applied this model to disaggregated price data. The interpretation of the DFI is that each individual item inflation breaks down into a common element termed monetary inflation, and a relative price element.

In the Stock and Watson (1991) notation, the DFI system has the following form:

$$\begin{aligned}
\begin{pmatrix} Y_{1,t} \\ \vdots \\ Y_{n,t} \end{pmatrix} &= \begin{pmatrix} \beta_1 \\ \vdots \\ \beta_n \end{pmatrix} + \begin{pmatrix} \gamma_1 \\ \vdots \\ \gamma_n \end{pmatrix} C_t + \begin{pmatrix} u_{1,t} \\ \vdots \\ u_{n,t} \end{pmatrix} \\
C_t &= \varphi_1 C_{t-1} + \varphi_2 C_{t-2} + \cdots + \varphi_p C_{t-p} + \delta + \eta_t \qquad (3) \\
\begin{pmatrix} u_{1,t} \\ \vdots \\ u_{n,t} \end{pmatrix} &= \begin{pmatrix} \phi_{1,1} u_{1,t-1} + \phi_{1,2} u_{1,t-2} + \phi_{1,k} u_{1,t-k} \\ \vdots \\ \phi_{n,1} u_{n,t-1} + \phi_{n,2} u_{n,t-2} + \phi_{n,k} u_{n,t-k} \end{pmatrix} + \begin{pmatrix} \varepsilon_{1,t} \\ \vdots \\ \varepsilon_{n,t} \end{pmatrix},
\end{aligned}$$

where Y_t is a vector of n observed time series and C_t their - unobserved - common factor. As in Bryan and Cecchetti (1993), the Y_t variables are here the growth rates of the individual items of the CPI. The model is a slightly restricted version of Stock and Watson (1991), for γ is assumed to be unity for every individual component. This warrants a certain degree of homogeneity between the common component and the individual prices. The common component is assumed to follow an $AR(p)$ process,[15] and each idiosyncratic shock $u_{i,t}$ to follow an $AR(k)$.[16] Let us adopt the following notation:

$$\begin{aligned}
C_t^* &= (C_t, C_{t-1}, \cdots, C_{t-p+1})' \\
u_t^* &= (u_{1,t}, \cdots, u_{1,t-k+1}, u_{2,t}, \cdots, u_{2,t-k+1}, \cdots, u_{n,t}, \cdots, u_{n,t-k+1})'.
\end{aligned}$$

The model (3) can then been written in a state-space form and estimated using the Kalman filter. Using the matrix notation detailed in Appendix 3, we get the state and observation equations:

$$\begin{pmatrix} C_t^* \\ u_t^* \end{pmatrix} = \begin{pmatrix} \varphi^*_{(p,p)} & 0_{(p,nk)} \\ 0_{(nk,p)} & \phi^*_{(nk,nk)} \end{pmatrix} \begin{pmatrix} C_{t-1}^* \\ u_{t-1}^* \end{pmatrix} + \begin{pmatrix} \delta \\ 0_{(p+nk-1,1)} \end{pmatrix} + \begin{pmatrix} N_{(p,1)} & 0_{(p,n)} \\ 0_{(nk,1)} & K_{(nk,n)} \end{pmatrix} \begin{pmatrix} \eta_t \\ \varepsilon_t^* \end{pmatrix}$$

[15] Note that the whole model relies on the stationnarity assumption for the inflation process, contrasting for instance with the SVAR approach, examined below.

[16] In the empirical application we also tested ARMA processes for C_t and u_t, but this specification did not provide staisfactory results.

$$(4)$$

$$Y_t = \beta + \left(\gamma N'_{(p,1)} \quad K'_{(nk,p)} \right) \begin{pmatrix} C_t^* \\ u_t^* \end{pmatrix} \tag{5}$$

The system (4) (5) may be written more compactly :

$$\begin{aligned} \alpha_t &= \mu_\alpha + T\alpha_{t-1} + R\zeta_t \\ Y_t &= \beta + Z\alpha_t \end{aligned} \tag{6}$$

$V(R\zeta_t) = R\Sigma R'$and Σ is a diagonal matrix with elements $(\sigma_\eta^2, \sigma_{\varepsilon_1}^2, \cdots, \sigma_{\varepsilon_n}^2)$.[17] It must be noted that in model (6), the parameters β_i and δ (the only non-zero element of μ_α) cannot be jointly identified. As Stock and Watson (1991) we use centered data to estimate the model :

$$\begin{aligned} \widetilde{\alpha}_t &= T\widetilde{\alpha}_{t-1} + R\zeta_t \\ Y_t - \overline{Y} &= Z\widetilde{\alpha}_t \end{aligned} \tag{7}$$

where $\overline{Y}$ is the vector of empirical means of the Y variables.

To compute the constant δ of the unobserved component equation, Stock and Watson (1991) propose the following procedure. They use the property that the extracted component $\widetilde{\alpha}_t$ can be rewritten as a linear combination of current and lagged values of the observed variables Y_t , i.e.:

$$\widetilde{\alpha}_t = W(L)(Y_t - \overline{Y}) \tag{8}$$

where $W(1)$ can be computed from the gain of the Kalman filter. $W(1)$ is a (m, n) matrix, m being the number of state variables, and n the number of observable variables. The first row of this matrix thus gives the implicit weight of each variable in the determination of the common component. Stock and Watson (1991) then assume that the weights given by the $W(1)$ matrix can also be used to compute the growth rate of the common component from the average growth rates of individual series. This means extending (8) to uncentered variables so that $C_t = e'_1 W(L)Y_t$, where the vector e'_1 selects the first row of $W(L)Y_t$. Writing equation the second equation of (3) as $\varphi(L)C_t = \delta + \eta_t$ we get the following estimator for δ :

$$\delta = \varphi(1)W(1)\overline{Y}. \tag{9}$$

This identifying assumption turns out to have important consequences for the interpretation of the results in the case of core inflation. Since Bryan and Cecchetti (1993) use the Stock and Watson (1991) identification restriction, their estimated mean of the common component needs not be equal to the actual mean of the aggregate CPI inflation rate. Indeed their major purpose, rather than estimate a core inflation index, is to compute the bias in price level measurement. Any discrepancy between those two means in their view measures the bias in measurement of the growth rate of the CPI. In our empirical application we use a different identifying assumption, since we impose underlying inflation to have the same mean as actual inflation.

[17] We impose the identifying assumption that $\sigma_\eta^2 = 1$.

4.2 The DFI : estimation results

We estimated the DFI unobserved component model (7) using the following 5 items disaggregation of the Harmonized Index of Consumer Prices : Unprocessed Food, Processed Food, Non energy industrial goods, Energy, Services. The specification retained is an AR(3) model for the common component and an AR(1) model for each idiosyncratic component. A LR test ruled out higher order terms in the autoregressive process. Furthermore, introducing ARMA processes for the common component did not prove fruitful. The empirically estimated system is the following :

$$\begin{aligned}
C_t &= \underset{(7,5)}{0,32}\, C_{t-1} + \underset{(6,8)}{0,22}\, C_{t-2} + \underset{(6,1)}{0,31}\, C_{t-3}. \\
u_{1,t} &= \underset{(1,7)}{0,31}\, u_{1,t-1} \text{ and } \sigma^2_{\varepsilon_1} = \underset{(15,6)}{0,66} \\
u_{2,t} &= \underset{(2,3)}{0,14}\, u_{2,t-1} \text{ and } \sigma^2_{\varepsilon_2} = \underset{(14,5)}{0,35} \\
u_{3,t} &= \underset{(2,1)}{0,21}\, u_{3,t-1} \text{ and } \sigma^2_{\varepsilon_3} = \underset{(14,1)}{0,29} \\
u_{4,t} &= \underset{(6,0)}{0,37}\, u_{4,t-1} \text{ and } \sigma^2_{\varepsilon 4} = \underset{(14,1)}{0,29} \\
u_{5,t} &= \underset{(-1,4)}{-0,11}\, u_{5,t-1} \text{ and } \sigma^2_{\varepsilon_5} = \underset{(11,7)}{0,20}
\end{aligned} \tag{10}$$

The core component of inflation is pictured in Figure 5. It is slightly smoother than the CPI inflation series.[18] The Kalman filter extraction implicitly involves reweighting of the individual components (from equation (8) above). The long run weighting scheme here is 0.04 for Unprocessed food, 0.15 for Processed Food, 0.20 Industrial Goods, 0.01 for Energy and 0.60 for Services.[19] This amounts to giving the Services Item, which is the less variable component, a high weight in the index - in fact twice its weight in the usual CPI. As a consequence, had we applied the Stock and Watson (1991) identifying condition for estimating the mean, the core inflation would have been on average significantly higher than observed inflation, due to the positive trend in relative price of services.

The DFI remains significantly below the actual inflation rate (by 2 points roughly) during the first oil shock. This expresses the fact that while the oil shock rapidly contaminated the open-economy sector, services remained less affected. Actual inflation overshoots core inflation during the 1979 shock as well : the gap is about 1.5 points (year to year increase) in the beginning of 1980. The 1987 upswing of inflation noted above illustrates a weakness of the DFI. This event corresponds to a transitory acceleration of some prices due to the liberalization of many services prices. Since the DFI overweights services, it results in an misleading rise in core inflation. To avoid such unappealing feature a solution would be to apply the DFI to more disaggregated data

[18] The standard deviation of the monthly increase over the period 1975:1 to 1998:12 is 0.39 for the CPI inflation, 0.38 for the excluding food and energy indicator, 0.33 for the trimmed mean, and 0.37 for the DFI.

[19] Note that applying the model to quarterly data, we obtained an even more unbalanced result, with services being affected a weight 0.8.

in the hope to get a more balanced result. But the dimension of the system would make the Kalman Filter estimation unfeasible.[20]

We will not develop an economic interpretation of the DFI. Since the DFI is essentially based on removing sectorial shocks, the same kind of remarks as in the section 3 applies.

5 The structural VAR approach

The SVAR approach is based on applying time series models to the aggregate CPI. Any trend extraction method (moving average, H-P filters etc.) could be selected for that purpose, but the choice among the numerous trend cycle decomposition method is rather arbitrary. A preferred approach is thus to root the decomposition in a multivariate framework and draw on assumed relations between core inflation and other economic variables. The seminal work of Quah and Vahey (1995) uses a structural VAR model to derive a core inflation measure using an economic restriction. The considered VAR models the dynamics of inflation and industrial output.

5.1 The Quah and Vahey (1995) methodology

Quah and Vahey (1995) proceed as follows. They estimate a bivariate inflation-output VAR. Accordingly with the structural VAR methodology, the innovations of the VAR are supposed to be functions of two uncorrelated structural shocks. To identify those shocks the following *a priori* restriction is used: one of the two shocks has no long run effect on the level of output. We term it here the “nominal shock”, while Quah and Vahey used the term “core inflation shock”. Core inflation is then defined as the part of inflation which is attributed to the nominal shock. Jacquinot (1998) shows on French, UK and German data, that since the non core part of inflation is closely correlated to the utilization rate, core inflation measure nearly amounts to correcting inflation for the business cycle.

Technically the Quah-Vahey approach involve three steps. First the following bivariate VAR is estimated :

$$X_t = B(L)X_t + \varepsilon_t. \tag{11}$$

where $X_t = \begin{pmatrix} \Delta y \\ \Delta \pi \end{pmatrix}$ and Δy is the growth rate of output and $\Delta \pi$ the variation of the inflation rate. Furthermore $\Sigma = E(\varepsilon_t \varepsilon_t')$ is the covariance matrix of the innovations. In the second step the structural shocks u_t are computed as a linear combination of the VAR residuals :

$$\varepsilon_t = Pu_t. \tag{12}$$

The matrix P is to be estimated. The structural shocks are assumed to be uncorrelated, so that $V(u_t) = I$. From (12) it is also true that $\Sigma = E(\varepsilon\varepsilon_t') = E(Pu_t u_t' P') =$

[20] Alternatively we may think of using the dynamic generalizations of principal component analysis presented by Stock and Watson (1999b) and Forni *et alii* (1999).

PP'. This condition provides three equations, while P has four parameters. Another hypothesis is required to complete identification of the VAR. In the Quah and Vahey approach the following restriction is used : the long run impact of the "nominal" shock on output is zero. In the bivariate case, P can be computed very easily. From (11) and (12) the moving average representation of the VAR is :

$$X_t = (I - B(L))^{-1} P u_t. \tag{13}$$

The identifying assumption is that the $u_{2,t}$ shock has no long run impact on the variation of output. This implies that the long run response matrix $H = (I - B(1))^{-1}P$ is lower triangular. H can then be obtained taking the Cholesky decomposition of M, where M is defined as $(I - B(1))^{-1}\Sigma(I - B(1))^{-1\prime} = HH'$. M can thus be recovered from the estimated parameters. From H the matrix $P = H(I - B(1))$ is computed and the shocks u_t can be recovered. The last step consists in computing the historical decomposition of the series into *core* and *non core* components. It is achieved using equation (13) where the shocks u_{1t} and u_{2t} are independent. Of particular interest is the inflation series generated using only the nominal shock u_{2t} : it turns out to be the core inflation series.

5.2 Estimation

We estimate a bivariate VAR model identical to that presented in Jacquinot (1998), based on the specification of Quah and Vahey (1995). The endogenous variables are the monthly growth rate in industrial production and the monthly variation of the inflation rate. Seasonal dummies are included in the equations. The estimation period is here 1973:6 to 1998:12 in order to provide basis for comparison with other indicators. Structural shocks are estimated according to the methodology sketched above.

Figure 6 presents the estimated core inflation, assuming core inflation equals inflation at the beginning of the sample period. We obtain results close to Jacquinot (1998).[21] Note that core inflation seems to deviate from actual inflation only in the second part of the sample.

The striking divergence between inflation and the SVAR indicator at the time of the 1993 recession is one of the most noticeable feature of the indicator. Since this measure amounts to correct inflation for the business cycle, the measures points to an acceleration in core inflation up to 3.5% at the end of 1993, when real activity declined. Symmetrically, during the subsequent recovery, core inflation decreases.

5.3 Interpreting the SVAR results

The Quah-Vahey approach relies on a macroeconomic identification assumption, based on the hypothesis of vertical long run Phillips curve. The observed dynamics of inflation and output is characterized as generated by two independent fundamental shocks, one only of which has a long-run impact on output. It seems natural in such a framework to view the shock with no long run impact on output as a demand shock,

[21] The VAR model was also estimated with oil price as an exogenous variable without altering the results.

and the other one as a supply shock. Quah and Vahey nevertheless "prefer to be agnostic on the exact determinants of core inflation", and define the two underlying shocks respectively as a "core shock" and a "non-core-shock".

The results obtained using the Qhah and Vahey approach raise an interpretation issue. The impulse response functions (Figure 7) imply that price and output have a positive response to the non-core inflation shock.[22] This feature usually characterizes demand shocks. Therefore, the shock with a long run impact on output seems to contain a demand-led component. This results from the characteristics of the empirically estimated VAR. The estimated system is the following :

$$\begin{aligned}
\Delta y_t &= B_{yy}(L)\Delta y_t + B_{yp}(L)\Delta \pi_t + \varepsilon_t^y \\
\Delta \pi_t &= B_{y\pi}(L)\Delta y_t + B_{\pi\pi}(L)\Delta \pi_t + \varepsilon_t^\pi
\end{aligned}$$

Concerning British and French data two important empirical restrictions emerge.[23] First the equation residuals are uncorrelated (the estimated Σ is diagonal), and furthermore the absence of lagged impact of prices on output (i.e.$B_{y\pi}(L) = 0$) cannot be rejected. Thus, even before the identification stage, the estimated system is nearly recursive and structural.

The innovation of the price equation has an impact on output neither in the short run (since residuals are uncorrelated) nor in the long run (since all terms in $B_{y\pi}(L)$ are zero). This innovation is therefore close to the structural core inflation shock, while the output equation innovation closely resembles the non core inflation shock.[24]

In order to propose an interpretation for the structural shocks retrieved from the SVAR it may be useful to put forward a simple macroeconomic model which explicitly includes a long run vertical Phillips curve. The vertical Phillips curve is indeed the theoretical motivation underlying the identifying assumption. Consider the following simple aggregate model:

$$\begin{aligned}
y_t - y_t^* &= \phi\left(y_{t-1} - y_{t-1}^*\right) - \upsilon\left(i_t - \pi_t - \rho\right) + \varepsilon_t^d && (14) \\
i_t &= \rho + \pi_t + \beta\left(\pi_t - \overline{\pi}_t^m\right) && (15) \\
\pi_t &= \pi_{t-1} + \alpha\left(y_t - y_t^*\right) + \varepsilon_t^p && (16)
\end{aligned}$$

(14) is a standard backward looking IS equation where $y_t - y_t^*$ is the output gap and ρ is the equilibrium real interest rate, assumed to be known and constant. ε_t^d is a private or government demand shock. Equation (15) is the reaction function of the monetary authorities with $\overline{\pi}_t^m$ the inflation target. No discretionary element is introduced in monetary policy, which is thus completely defined by the (potentially time-varying) inflation target and the parameters of the reaction function. (16) is a reduced-form backward-looking Phillips curve, which includes a price shock ε_t^p that

[22] The pattern are similar to those obtained on UK data. Note that Blix (1995) obtains very different pattern, using a similar dataset to Quah and Vahey, when estimating the VAR with the level of inflation rather than with its rate of change.

[23] Those restrictions are accepted using separate hypotheses tests, see Jacquinot (1998).

[24] Jacquinot (1998) thus uses the term «real shock» for the latter.

might be attributed to cost-push phenomena. Four exogenous shocks thus drive fluctuations: two demand shocks (ε_t^d and the monetary policy shock $\overline{\pi}_t^m$), a supply shock which drives y_t^* and a price shock ε_t^p. The shocks are assumed to be independent:[25] (14), (15), (16) is the model that a structural econometric estimation should attempt to unveil. We can put this model in an autoregressive reduced form so as to compare it with the estimated VAR. An aggregate demand curve is drawn from (14) and (15):

$$y_t - y_t^* = \phi\left(y_{t-1} - y_{t-1}^*\right) - \upsilon\beta\left(\pi_t - \overline{\pi}_t^m\right) + \varepsilon_t^d. \tag{17}$$

Using (16) and (17) the model can be expressed in an autoregressive form :

$$\begin{aligned} y_t - y_t^* &= \left(\frac{1}{1+\alpha\beta\upsilon}\right)\left[\phi\left(y_{t-1} - y_{t-1}^*\right) - \upsilon\beta\pi_{t-1} - \overline{\pi}_t^m - \upsilon\beta\varepsilon_t^p + \varepsilon_t^d\right] & (18) \\ \pi_t &= \left(\frac{1}{1+\alpha\beta\upsilon}\right)\left[\alpha\phi\left(y_{t-1} - y_{t-1}^*\right) + \pi_{t-1} + \alpha\upsilon\beta\overline{\pi}_t^m + \varepsilon_t^p + \alpha\varepsilon_t^d\right] & (19) \end{aligned}$$

The long run of the model is straightforward: output is equal to potential, demand shocks have no long run effect; inflation is equal to the central bank target. It seems therefore sensible to consider that in the model the underlying inflation rate is $\overline{\pi}_t^m$. Demand and supply shocks, as well as price shocks, have only a temporary effect on the inflation rate.

Equations (18) and (19) may then be compared to the estimated VAR. The potential output of (18) is not empirically observed. The empirical model is formulated in terms of growth rate rather than in terms of gap, and potential output implicitly follows a process with a unit root.[26] The output equation of the VAR is thus a mix of demand curve and potential output. Therefore, given the nearly structural characteristics of the estimated model in both British and French cases, the "non-core" (or real) shock is a mix of demand shocks and supply shocks. Disentangling the two shocks would certainly require incorporating additional variables (Blanchard and Quah (1989) use the unemployment rate for this purpose).[27]

In the empirical SVAR for the two mentioned countries the core inflation shock is the price equation shock ε_t^p. This contradicts the interpretation of the model sketched above, in which core inflation is the central bank's inflation target. Consider a disinflation experiment in which the central bank reduces its inflation target. This create a transitory recession and leads to a permanent decrease in the inflation rate. The shock will be captured by the estimated SVAR as a real shock (or non core shock), and the calculated core inflation will not decrease. Identifying a fully-fledged monetary policy shock would probably require introducing monetary policy instrument in the VAR, as in Gali (1992).

[25]Althought ε_t^p and y_t^*could be considered as correlated on the grounds that for instance a shock in pay-roll taxes, or other element of the wage wedge, has a permanent effect on output. Here ε_t^p interpreted as a short term supply shock (for instance a schock on oil price, on the wage wedge,etc).

[26]If potential output is driven by a deterministic trend, no shock has long run impact on GDP.

[27]Alternatively, imposing a constraint on the relation between potential output and the price-cost push shock might be considered.

Theoretical interpretation of empirical indicators of core inflation is not straightforward. In the above sketched small model, core inflation would rather be defined as the inflation target.

6 Conclusion

This paper has implemented and compared in the case of France four methods for computing indicators of underlying inflation: inflation excluding food and energy, trimmed mean, the structural VAR approach and the DFI, *i.e.* an unobserved component model. These indicators rely on very different assumptions, statistical frameworks and datasets.

Because they stress on different types of shocks, the various indicators might provide contrasting estimates at a given moment. At the end of 1998 one may consider actual French inflation to be above structural inflation if looking at the SVAR indicator, or conversely to be below it if looking at inflation excluding food and energy index (Figure 8). This is because 1998 is characterized by a fairly strong output growth together with a historically low rate of inflation, explained in part by low commodity prices. For the SVAR indicator the conjunction of output recovery and very low inflation indicates that core inflation has been reduced to zero. "Limited influence and exclusion measures" suggest that core inflation (1.1% roughly) is above actual inflation, due to transitory shocks in commodity prices. Each of the various indicators might convey valuable information for appraising short term inflation developments. Indeed they present alternative summaries of the data, helping to track the source of any shock in the inflation rate.

Nevertheless the theoretical interpretation of the indicators remains somewhat unwarranted, lacking an explicit representation of monetary policy and of the long term determination of the inflation rate. A way to provide a satisfactory foundation for core inflation might build on the inflation targeting literature (for instance Svensson 1997). In Svensson's model, some shocks (which might be oil price shocks as well as demand shocks) have effects on inflation and output, that can not be stabilized by monetary policy because they are unpredictable and because monetary policy affects output and inflation with some significant lags. Assessment of monetary policy should abstract from the effects of those shocks, and therefore an inflation indicator able to exclude the impact of such shocks should be useful to evaluate monetary policy. Core inflation might be in this respect related to an inflation forecast (conditional on the information available at the time of the monetary policy decision). Deriving an operational measure of core inflation along those lines would require not only to figure out what kind of shocks affect inflation, but also to examine the transmission lags of monetary policy.

References

[1] Bakhshi, H., and T. Yates (1999) To trim or not to trim? An application of a trimmed mean inflation estimator to the United Kingdom. *Bank of England Working Paper Series* N°97, July.

[2] Ball, L. and N.G. Mankiw (1992) Relative-Price Changes as Aggregate Supply Shocks. *NBER Working Paper* N°4168.

[3] Ball, L. and N.G. Mankiw (1999) Interpreting the Correlation between Inflation and the Skewness of Relative prices: a comment on Bryan and Cecchetti. *Review of Economics and Statistics* 81(2): 197-198.

[4] Blanchard, O.J. and D. Quah (1989) The Dynamic Effects of Aggregate Demand and Supply Disturbances. *American Economic Review* 79(4): 655-673.

[5] Blinder, A. S (1982) Anatomy of Double-Digit Inflation. In *Inflation: Causes and Effects (*R.E Hall Ed.) University of Chicago Press, pp 261-282.

[6] Blinder, A. S (1997) Commentary, *Federal Reserve Bank of Saint Louis Review,*79(3) 157-160.

[7] Blix, M. (1995) Underlying Inflation - A common Trends Approach. *Sveriges Riksbank Working Paper* n°23.

[8] Bonnet, X., E. Dubois and L. Fauvet (1998) Asymétrie des inflations relatives and *menus costs* : tests sur l'inflation française. *Revue Française d'Economie, 50(3): 547-556.*

[9] Bryan, M.F. and C.J. Pike (1991) Median Prices Changes: an Alternative Approach to Measuring Current Monetary Inflation. *Economic Commentary*, Federal Reserve Bank of Cleveland, December.

[10] Bryan, M.F. and S.G. Cecchetti (1993) The CPI as measure of Inflation. *Economic Review*, Federal Reserve Bank of Cleveland, 4^{th}quarter.

[11] Bryan, M.F. and S.G. Cecchetti (1994) Measuring Core Inflation. In *Monetary Policy*, (G. Mankiw Ed). NBER Studies in Business Cycle, vol 29.

[12] Bryan, M.F. and S.G. Cecchetti (1996) Inflation and the Distribution of Prices Changes. *NBER Working Paper* N°5783.

[13] Bryan, M.F. and S.G. Cecchetti (1999) Inflation and the Distribution of Prices Changes. *Review of Economics and Statistics* 81(2):188-196.

[14] Bryan, M.F., S.G. Cecchetti and R.J. Wiggins II (1997) Efficient Inflation Estimation. *NBER Working Paper* N°6183.

[15] Cecchetti, S.G. (1996) Measuring Short Run Inflation for Central Bankers. *NBER Working Paper,* N°5786.

[16] Chatelain, J.B., I. Odonnat and P. Sicsic (1996) Construction sur longue période d'un indicateur d'inflation sous-jacente.Mimeo, Banque de France.

[17] Eckstein, O. (1981) *Core Inflation*, Englenwood Cliffs, Prentice-Hall, Inc.

[18] Forni M., M. Hallin, M. Lippi and L. Reichlin (1999) The generalized dynamic factor model : identification and estimation. Mimeo presented at the NBER/NSF Forecasting Seminar .

[19] Gali, J. (1992) How Well Does the IS-LM Model Fit the Postwar US Data?. *Quaterly Journal of Economics*, CVII(2): 709-738.

[20] Hamilton, J.D. (1994) *Time Series Analysis*, Princeton University Press.

[21] Huber, P.J. (1981) *Robust Statistics,* John Wiley and Co.

[22] Insee (1996) Indice corrigé des variations saisonnières and indice hors tarifs publics and produist à prix volatils corrigé des mesures fiscales and des variations saisonnières. *Informations Rapides* 179.

[23] Jacquinot, P. (1998). L'inflation sous-jacente à partir d'une approche structurelle des VAR : une application à la France, l'Allemagne and le Royaume-Uni. *Notes d'Études and de Recherche de la Banque de France* N°51.

[24] Laflêche, T. (1997) Mesure du taux d'inflation tendanciel. *Document de Travail de la Banque du Canada,* N°97-9.

[25] Quah, D. and S.P. Vahey (1995) Measuring Core Inflation *Economic Journal* 105: 1130-1144.

[26] Roger, S. (1998) Core Inflation: Concepts, Uses and Measurement. *Reserve Bank of New Zealand Discussion Paper* N°G98/9.

[27] Stock, J.H and M.W. Watson (1991) A Probability Model of the Coincident Economic Indicator. In *Leading Economic Indicators New Approaches and Forecasting Records* (K. Lahiri and G. Moore Eds.) Cambridge University Press, pp 63-89.

[28] Stock, J.H and M.W. Watson (1999a) Forecasting Inflation. *Journal of Monetary Economics*, 44: 293-335.

[29] Stock, J.H and M.W. Watson (1999b) Diffusion indexes. Mimeo presented at the NBER/NSF Forecasting Seminar (revised version of *NBER Working Paper* N°6702, 1998).

[30] Svensson, L.E.O (1997) Inflation Forecast Targeting: Implementing and Monitoring Inflation Targets. *European Economic Review*, 41: 1111-1146.

[31] Vining, D. R. and T.C Elwertowski (1976) The Relationship between Relative Prices and the General Price Level. *American Economic Review.* 66: 699-708.

[32] Wynne, M.A. (1997) Commentary. *Federal Reserve Bank of Saint Louis Review,* 79(3):161-167.

[33] Wynne, M.A. (1999) Core Inflation: a Review of Some Conceptual Issues. ECB Working Paper N°5.

Appendix 1
Optimal trim in an asymmetric distribution

Bryan, Cecchetti and Wiggins (1997) illustrate through Monte Carlo simulations the choice of an optimal trim, i.e. the degree of trimming which yields an estimator with minimum variance. Their example is a mixture of two normal distributions with different variances. Indeed a standard result in robust statistics (Huber 1981) is that for such distributions the empirical mean does not yield an efficient estimation of the theoretical mean of the distribution. The mixture distribution is the following:

$$\begin{aligned} z &= sy_1 + (1-s)y_2 \\ y_1 &\rightarrow N(0,1) \\ y_2 &\rightarrow N(0,\sigma) \\ P(s &= 1) = p \end{aligned}$$

Its first moments are $E(z) = 0$ and $V(z) = p^2 + (1-p)^2 \sigma^2$. The kurtosis of this distribution depends on p and σ. It is worth $K(p,\sigma) = \frac{3(p+(1-p)\sigma^2}{p^2+(1-p)^2\sigma^2}$. For $\sigma = 1$ (z is normal) the kurtosis is 3, while for large values of σ, kurtosis tends to $\frac{3}{1-p}$.

If draws are taken from the first distribution only (i.e. $p = 1$) the variance of the empirical mean is $V(\overline{Y}_1) = \frac{1}{n}$ while that of the empirical median $V(median(Y_1)) = (\frac{\pi}{2})\frac{1}{n}$. In such a case the empirical mean is the most efficient estimator. But in the case $1-p > 0$, that is if drawn can be taken from the second distribution, this is no longer the case. When for instance σ goes to infinity, the variance of the empirical mean is unbounded while the variance of the empirical median is bounded. In such a case trimmed means show a good performance. In the above example, the optimal trim increases with the kurtosis and reaches a plateau at 15% (figure 9).[28] The trimmed mean estimator has the property to be unbiased.

We now consider the case when the population mean of the second distribution differs from that of the first one, that is :

$$y_2 \rightarrow N(\mu,\sigma), \quad \mu > 0.$$

The probability distribution of the resulting z variable is then asymmetrical. This feature is consistent with stylized facts of the cross-sectional distribution in the French case. When drawing from the z distribution, large values will most of the times be drawn from the right side of the distribution (if the mean μ is positive and large enough). An subsequent intuition is that the optimal trim should not be symmetrical, but that more should be trimmed on the right side of the distribution. Another feature is that the trimmed mean will be systematically biased. Indeed, any symmetrical trim will remove observations on the right side of the distribution, whose weight is heavier

[28]The graph is obtained by bootstrap experiments for each value of the kurtosis. One replication consists in 100 draws in the above distribution (supposed to be embody 100 items of the price index). A number of trimmed mean is computed from these observations. 1000 replications are then performed, allowing to compute the variance of the trimmed mean estimators.

than that of the observations removed on the left side of the distribution. As a result the trimmed mean will be inferior to the mean. An illustration of this is the fact that for an asymmetric distribution, the mean is inferior to the median. An alternative approach to implement trimmed mean in this context is to choose an asymmetric trim that yields a trimmed mean on average close to the empirical mean. But this bears a cost in terms of variability.

Table 1 reports the results of computing trimmed means according to various criteria, when the distribution of z has a kurtosis 12 and for values of the μ comprised between 0 and 10. The first columns reports the trimmed mean that yields the closer result to the empirical mean. The second column reports the trim leading to the minimal variance of the trimmed mean estimator. The third column reports the trim obtained when the objective function is a weighted average between mean and variance criteria.[29] It is apparent from column 1 that the more the asymmetry increases, the more the trim has to be weighted on the left side of the distribution in order to yield an unbiased estimator of the (time series) mean of CPI-inflation. Conversely, for highly skewed distributions, the trim is more important on the right side of the distribution, yielding an underlying inflation series systematically inferior to the aggregate inflation. Unsurprisingly, the fourth column yields intermediate results.

Of course the relevance of the particular distribution to represent disaggregated price data is questionable. It is nevertheless very particular to suppose that every item has the same distribution, which happens to be a mixture of two distributions. It might be considered that each item in the price index has its own distribution.

Table 1

Optimal trim - distribution of z ($p = 0,9$ $\sigma = 4$)

Mean of y_2 (1)	Criteria : Minimum Bias (2)	Variance (3)	Bias and variance (4)
μ =0	(5,5)[30]	(10,10)	(10,10)
μ =1	(30,25)	(15,15)	(15,15)
μ =2	(35,25)	(10,15)	(15,20)
μ =3	(40,25)	(15,15)	(10,15)
μ =4	(40,20)	(15,20)	(10,15)
μ =5	(45,20)	(15,25)	(25,25)
μ =6	(45,15)	(10,20)	(15,15)
μ =7	(50,15)	(10,20)	(20,15)
μ =8	(55,15)	(10,20)	(25,15)
μ =9	(60,15)	(10,20)	(30,15)
μ =10	(65,15)	(10,10)	(50,25)

As a tentative experience to implement optimal trim on real data, we performed bootstrap experiments on French inflation from 1993:1 to 1998:12, under the assumption that the average inflation for the whole period could be considered as the core

[29] An alternative approach would be to minimize the variance conditionnal on the trimmed mean being unbiased.

[30] Trim in % in each side between brackets. Determination of optimal trims was conducted by grid search, with 5% steps.

inflation for each period (such an assumption is necessary in order to recover "residuals" , i.e. deviations from core inflation necessary for the bootstrap).

Results are reported in Table 2. The optimal trim for the variance criteria yields a narrow interval around the 40th percentile. The time-series average of that trimmed mean is in that case close to 0% while the monthly average inflation over the period is close to 0.1%.

Table 2
Optimal trim - French inflation data

Estimator of core inflation	Trim	Mean of estimator	Variance
Average inflation	(0,0)	0.0950	24.6
Symmetrical trimmed mean	(15,15)	0.0784	2.9
Minimum Variance trimmed mean	(35,55)	0.0057	1.1
Minimum bias trimmed mean	(15,35)	0.0939	1.7
Combination bias and variance criteria	(35,30)	0.0059	1.1
Median	(50,50)	0.0657	1.6
$55^{th} percentile$	(55,45)	0.1042	2.2
$60^{th} percentile$	(60,40)	0.1481	2.4

Note : trims in the left-side and right side of the distribution between brackets.

Appendix 2
Notations used in the state-space form of the DFI model

The notations used to write the DFI model in the state-space form (4),(5) are :

$$\varphi^*_{p,p} = \begin{pmatrix} \varphi_1 & \cdots & \varphi_{p-1} & \varphi_p \\ & I_{p-1} & & 0 \end{pmatrix}$$

$$\phi^*_{nk,nk} = \begin{pmatrix} \phi^*_1 & 0_{k,k} & \cdots & 0_{k,k} \\ 0_{k,k} & \ddots & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0_{k,k} \\ 0_{k,k} & \cdots & 0_{k,k} & \phi^*_n \end{pmatrix},$$

with $\phi^*_j = \begin{pmatrix} \phi_{j,1} & \cdots & \phi_{j,k-1} & \phi_{j,k} \\ & I_{k-1} & & 0 \end{pmatrix}$ for $j \in [1, n]$.

Lastly, $N_{(p,1)} = (1, 0, \cdots, 0)'$ and $K_{(nk,n)} = I_{(n,n)} \otimes e'_{(1,k)}$ with $e_{(1,k)} = (1, 0 \cdots 0)$.

Formula (6) is obtained simply by writing :

$\alpha_t = \begin{pmatrix} C^*_t \\ u^*_t \end{pmatrix}$,

$T = \begin{pmatrix} \varphi^*_{(p,p)} & 0_{(p,nk)} \\ 0_{(nk,p)} & \phi^*_{(nk,nk)} \end{pmatrix}$,

$\mu_\alpha = \begin{pmatrix} \delta \\ 0_{(p+nk-1,1)} \end{pmatrix}$,

$R = \begin{pmatrix} N_{(p,1)} & 0_{(p,n)} \\ 0_{(nk,1)} & K_{(nk,n)} \end{pmatrix}$,

$Z = \begin{pmatrix} \gamma N'_{(p,1)} & K'_{(nk,n)} \end{pmatrix}$.

Appendix 3: Data Sources

Disaggregated price data are taken from the INSEE Consumer Price Indices Database. Two databases have been merged. The first one (base year 1980), extending from 1970 to 1992, has 296 items. The second database (base year 1990) runs from 1993 to December 1998 and has 265 items. From January 1999, onwards a new base year (1995) and a new nomenclature of the CPI have been introduced, which are not considered here.

The five broad sectorial indices are computed aggregating detailed sub-components drawn from the above described databases.

The industrial production index used in the VAR model is taken from Datastream.

Figures (1 to 9)

Figure 1. Inflation excluding food, energy and public services

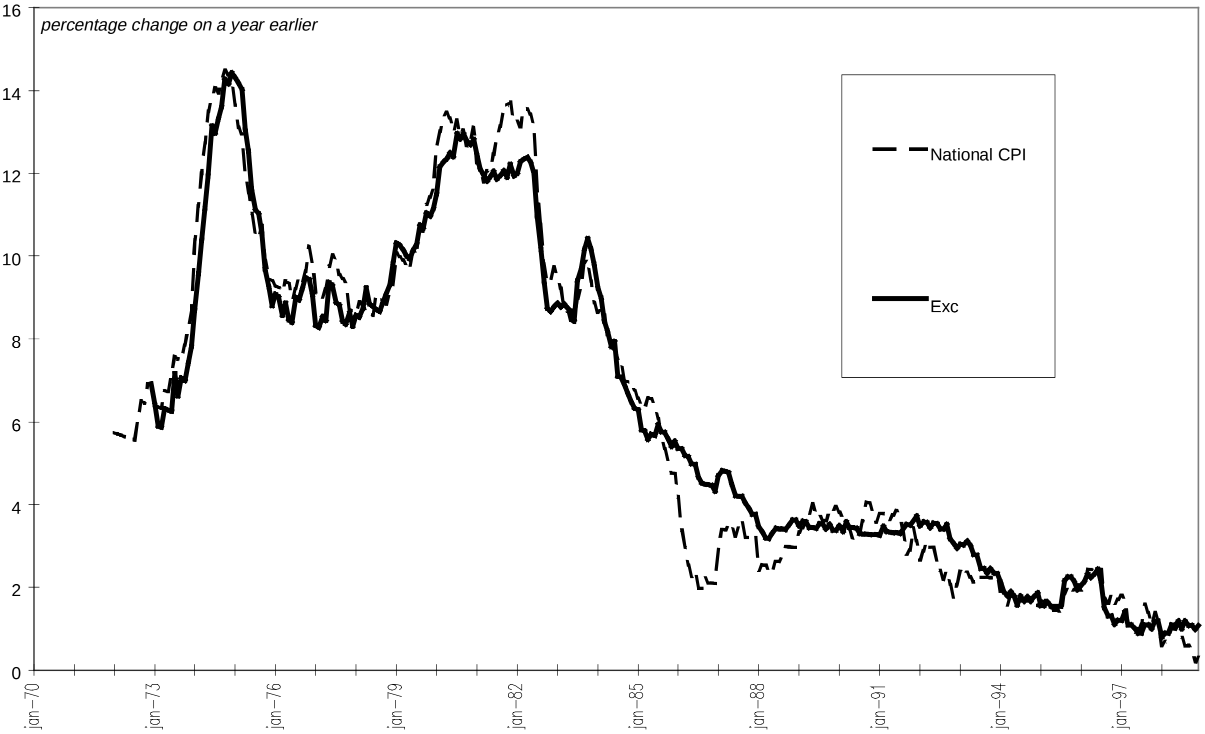


Figure 2. Sectoral inflation

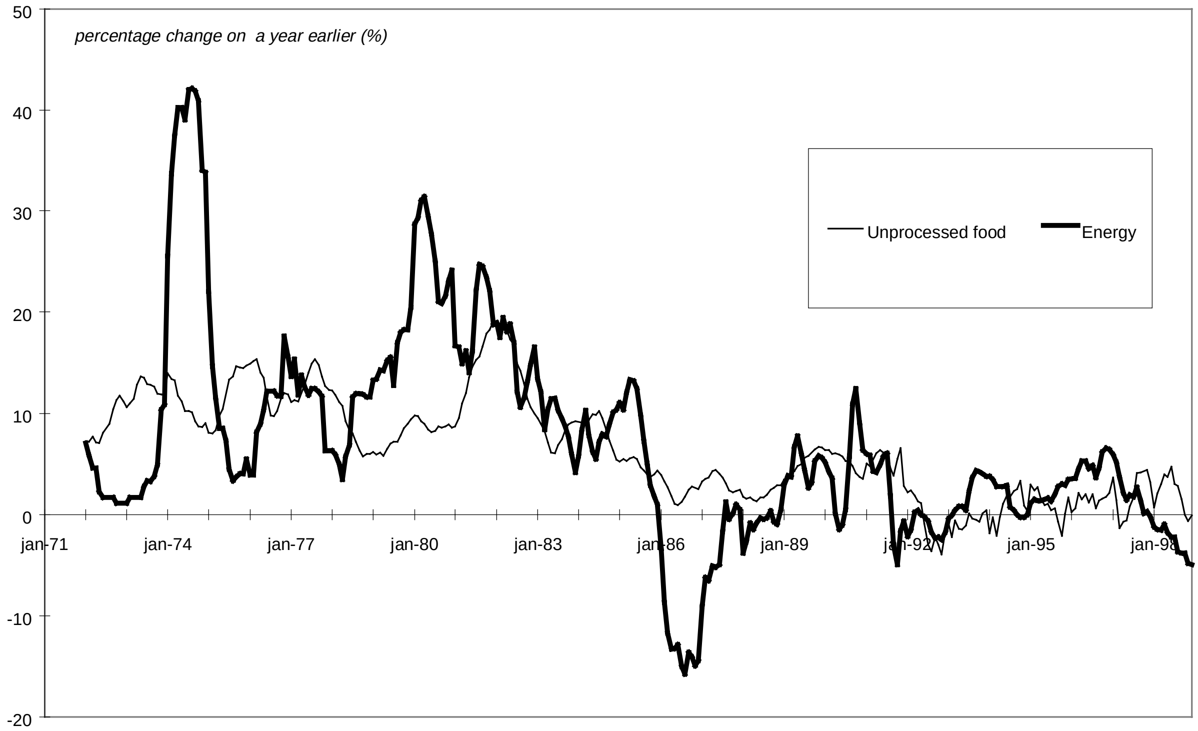
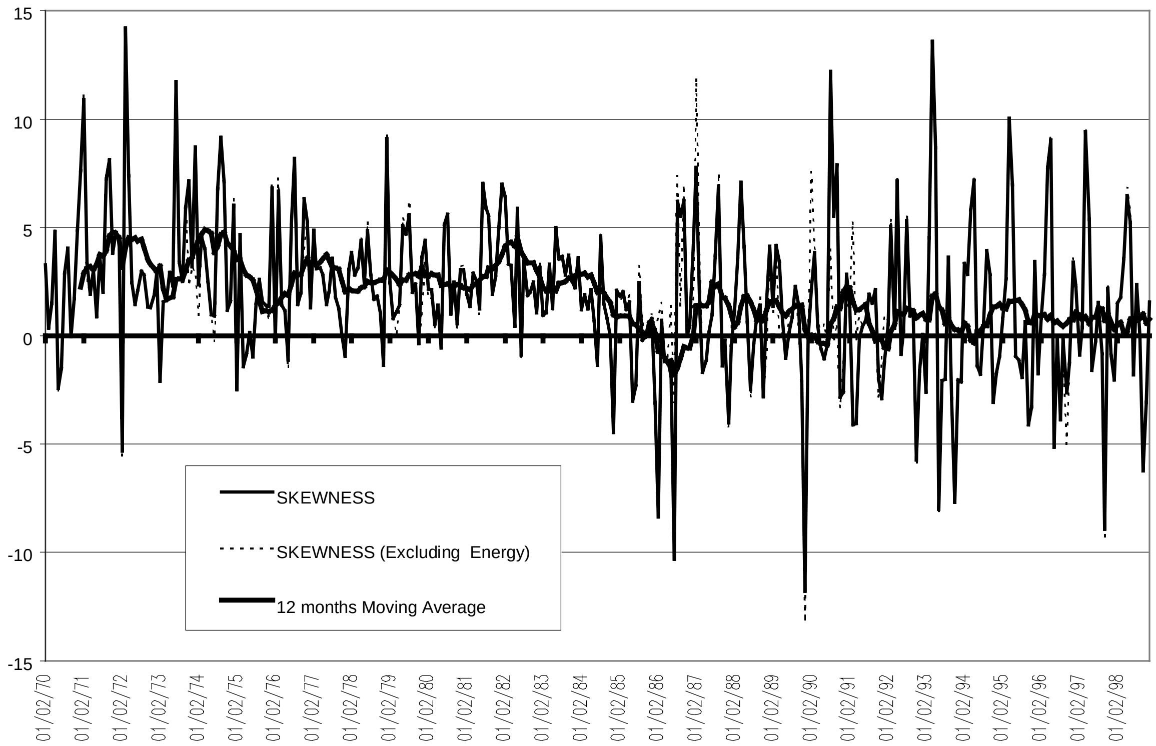


Figure 3. Skewness of the cross-sectional distribution in prices changes, 1970-1998.



Note : Unweighted Skewness of growth rates of the CPI items.

Figure 4. Underlying inflation : trimmed mean indicators

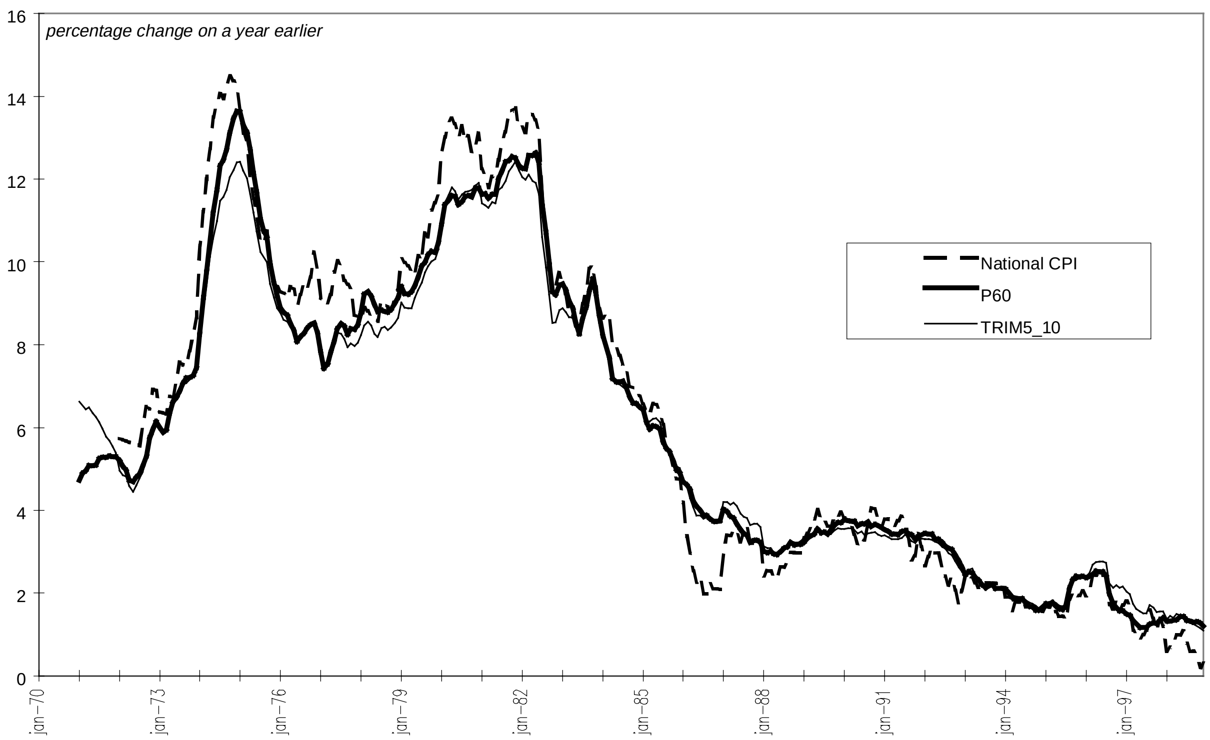


Figure 5. Core inflation : DFI estimate Kalman filter

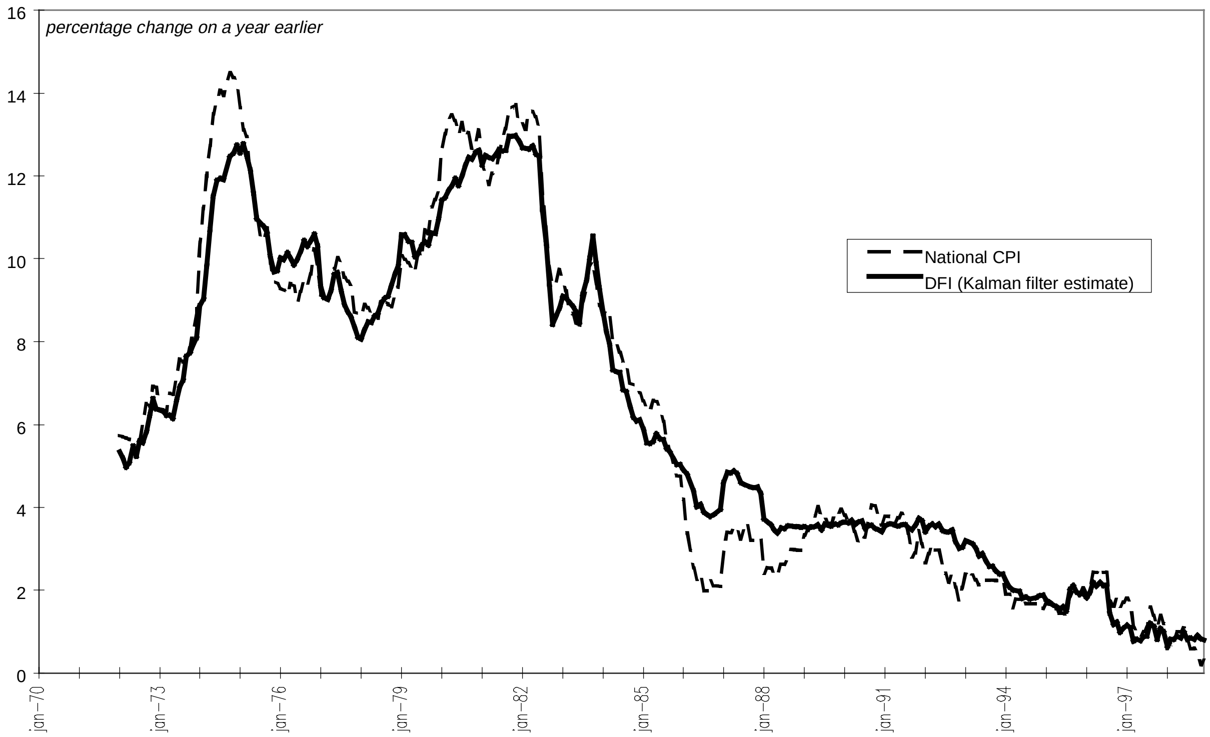


Figure 6. Core inflation : SVAR measure

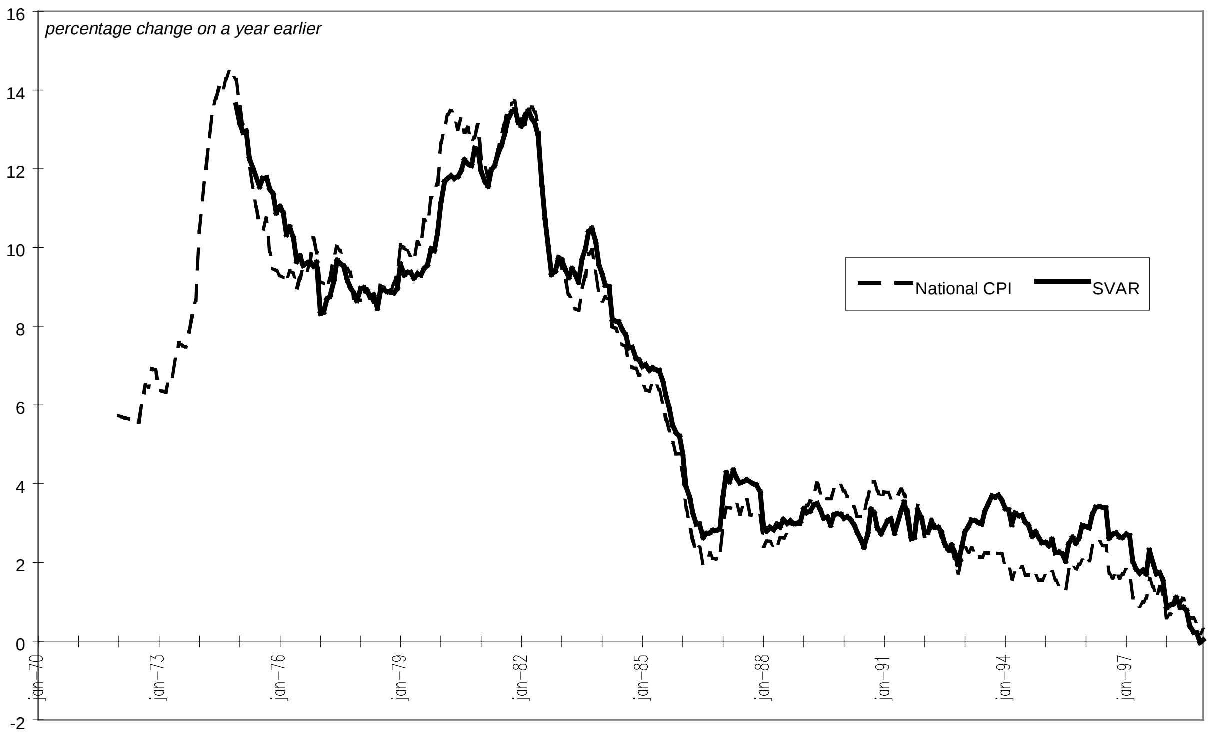


Figure 7. SVAR impulse response functions

Fig. 7a. Output response to real shock

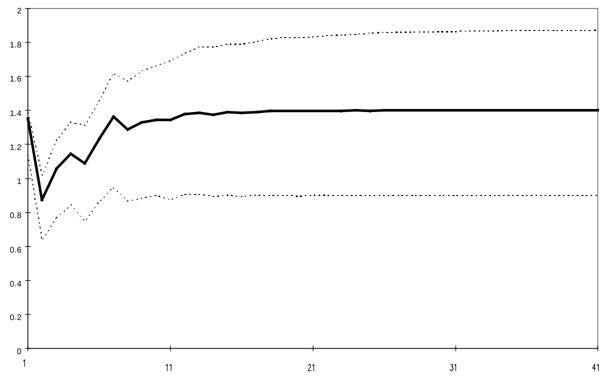


Fig. 7b. Output response to nominal shock

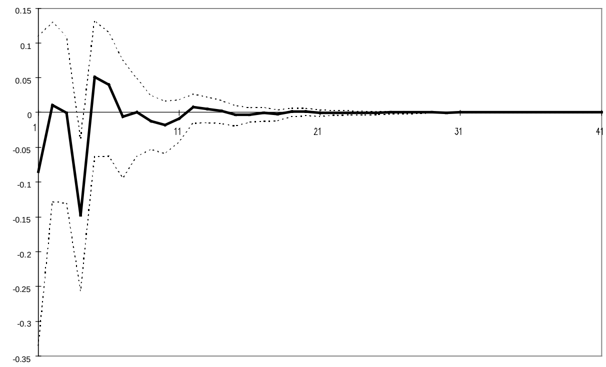


Fig. 7c. Inflation response to real shock

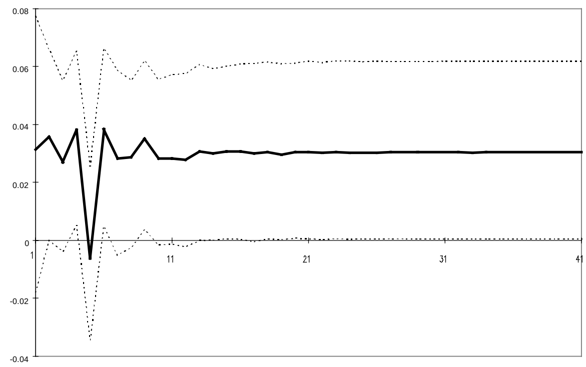
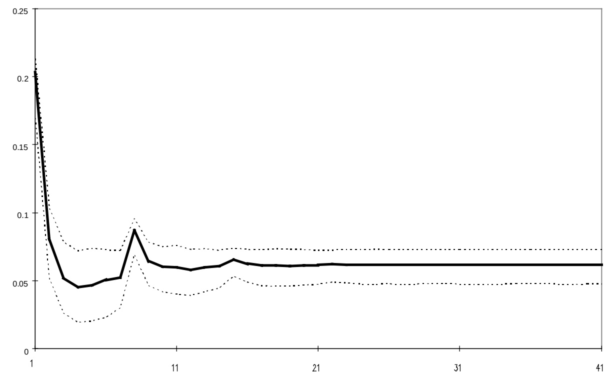


Fig. 7d. Inflation response to nominal shock



Note : The dashed lines indicate a 5% confidence interval. Impulse response standard errors were calculated using the *bootstrap* (1000 replications).

Fig. 8. Selected underlying inflation indicators, 1991-98.

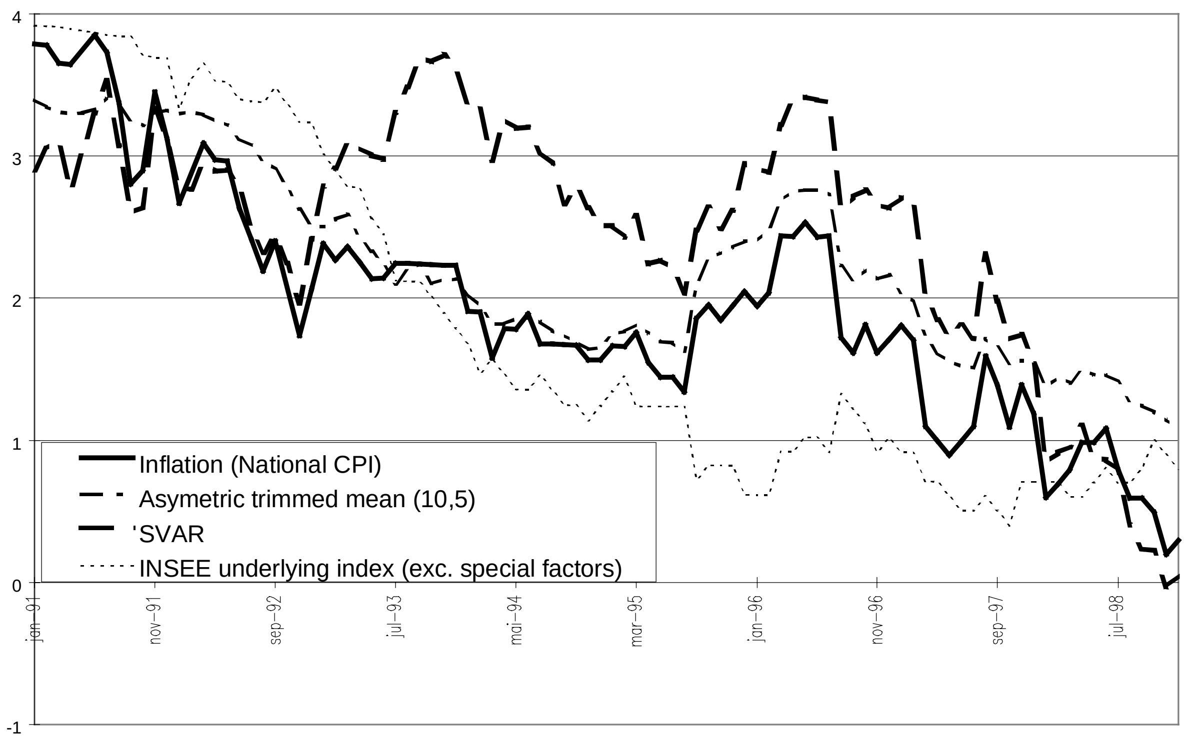
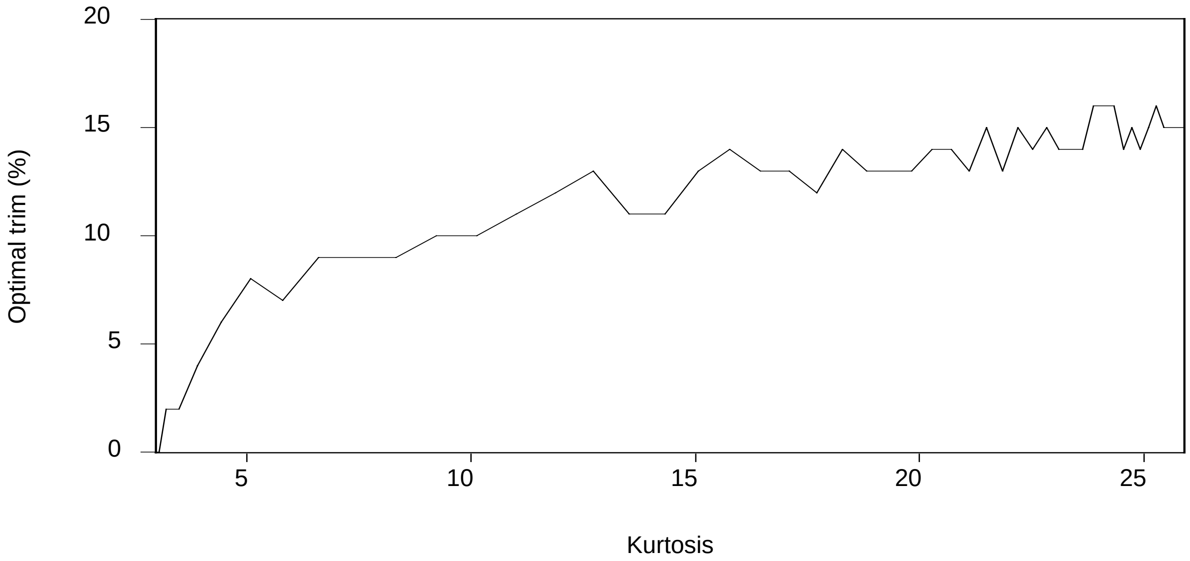


Fig. 9. The optimal trim as a function of the kurtosis (mixture of two normal distributions)



Note : Results from Monte-Carlo experiments, with 1000 replication of 100 draws in the mixture distribution for each value of the kurtosis. Optimal trim search was performed by 5% steps, so that 10 trimmed means are computed for each replication.

Notes d'Études et de Recherche

1. C. Huang and H. Pagès, “Optimal Consumption and Portfolio Policies with an Infinite Horizon: Existence and Convergence,” May 1990.

2. C. Bordes, « Variabilité de la vitesse et volatilité de la croissance monétaire : le cas français », février 1989.

3. C. Bordes, M. Driscoll and A. Sauviat, “Interpreting the Money-Output Correlation: Money-Real or Real-Real?,” May 1989.

4. C. Bordes, D. Goyeau et A. Sauviat, « Taux d'intérêt, marge et rentabilité bancaires : le cas des pays de l'OCDE », mai 1989.

5. B. Bensaid, S. Federbusch et R. Gary-Bobo, « Sur quelques propriétés stratégiques de l’intéressement des salariés dans l'industrie », juin 1989.

6. O. De Bandt, « L'identification des chocs monétaires et financiers en France : une étude empirique », juin 1990.

7. M. Boutillier et S. Dérangère, « Le taux de crédit accordé aux entreprises françaises : coûts opératoires des banques et prime de risque de défaut », juin 1990.

8. M. Boutillier and B. Cabrillac, “Foreign Exchange Markets: Efficiency and Hierarchy,” October 1990.

9. O. De Bandt et P. Jacquinot, « Les choix de financement des entreprises en France : une modélisation économétrique », octobre 1990 (English version also available on request).

10. B. Bensaid and R. Gary-Bobo, “On Renegotiation of Profit-Sharing Contracts in Industry,” July 1989 (English version of NER n° 5).

11. P. G. Garella and Y. Richelle, “Cartel Formation and the Selection of Firms,” December 1990.

12. H. Pagès and H. He, “Consumption and Portfolio Decisions with Labor Income and Borrowing Constraints,” August 1990.

13. P. Sicsic, « Le franc Poincaré a-t-il été délibérément sous-évalué ? », octobre 1991.

14. B. Bensaid and R. Gary-Bobo, “On the Commitment Value of Contracts under Renegotiation Constraints,” January 1990 revised November 1990.

15. B. Bensaid, J.-P. Lesne, H. Pagès and J. Scheinkman, “Derivative Asset Pricing with Transaction Costs,” May 1991 revised November 1991.

16. C. Monticelli and M.-O. Strauss-Kahn, “European Integration and the Demand for Broad Money,” December 1991.

17. J. Henry and M. Phelipot, “The High and Low-Risk Asset Demand of French Households: A Multivariate Analysis,” November 1991 revised June 1992.

18. B. Bensaid and P. Garella, “Financing Takeovers under Asymetric Information,” September 1992.

19. A. de Palma and M. Uctum, "Financial Intermediation under Financial Integration and Deregulation," September 1992.

20. A. de Palma, L. Leruth and P. Régibeau, "Partial Compatibility with Network Externalities and Double Purchase," August 1992.

21. A. Frachot, D. Janci and V. Lacoste, "Factor Analysis of the Term Structure: a Probabilistic Approach," November 1992.

22. P. Sicsic et B. Villeneuve, « L'Afflux d'or en France de 1928 à 1934 », janvier 1993.

23. M. Jeanblanc-Picqué and R. Avesani, "Impulse Control Method and Exchange Rate," September 1993.

24. A. Frachot and J.-P. Lesne, "Expectations Hypothesis and Stochastic Volatilities," July 1993 revised September 1993.

25. B. Bensaid and A. de Palma, "Spatial Multiproduct Oligopoly," February 1993 revised October 1994.

26. A. de Palma and R. Gary-Bobo, "Credit Contraction in a Model of the Banking Industry," October 1994.

27. P. Jacquinot et F. Mihoubi, « Dynamique et hétérogénéité de l'emploi en déséquilibre », septembre 1995.

28. G. Salmat, « Le retournement conjoncturel de 1992 et 1993 en France : une modélisation V.A.R. », octobre 1994.

29. J. Henry and J. Weidmann, "Asymmetry in the EMS Revisited: Evidence from the Causality Analysis of Daily Eurorates," February 1994 revised October 1994.

30 O. De Bandt, "Competition Among Financial Intermediaries and the Risk of Contagious Failures," September 1994 revised January 1995.

31. B. Bensaid et A. de Palma, « Politique monétaire et concurrence bancaire », janvier 1994 révisé en septembre 1995.

32. F. Rosenwald, « Coût du crédit et montant des prêts : une interprétation en terme de canal large du crédit », septembre 1995.

33. G. Cette et S. Mahfouz, « Le partage primaire du revenu : constat descriptif sur longue période », décembre 1995.

34. H. Pagès, "Is there a Premium for Currencies Correlated with Volatility ? Some Evidence from Risk Reversals," January 1996.

35. E. Jondeau and R. Ricart, "The Expectations Theory : Tests on French, German and American Euro-rates," June 1996.

36. B. Bensaid et O. De Bandt, « Les stratégies "stop-loss" : théorie et application au Contrat Notionnel du Matif », juin 1996.

37. C. Martin et F. Rosenwald, « Le marché des certificats de dépôts. Écarts de taux à l'émission : l'influence de la relation émetteurs-souscripteurs initiaux », avril 1996.

38. Banque de France - CEPREMAP - Direction de la Prévision - Erasme - INSEE - OFCE, « Structures et propriétés de cinq modèles macroéconomiques français », juin 1996.

39. F. Rosenwald, « L'influence des montants émis sur le taux des certificats de dépôts », octobre 1996.

40. L. Baumel, « Les crédits mis en place par les banques AFB de 1978 à 1992 : une évaluation des montants et des durées initiales », novembre 1996.

41. G. Cette et E. Kremp, « Le passage à une assiette valeur ajoutée pour les cotisations sociales : Une caractérisation des entreprises non financières “gagnantes” et “perdantes” », novembre 1996.

42. S. Avouyi-Dovi, E. Jondeau et C. Lai Tong, « Effets “volume”, volatilité et transmissions internationales sur les marchés boursiers dans le G5 », avril 1997.

43. E. Jondeau et R. Ricart, « Le contenu en information de la pente des taux : Application au cas des titres publics français », juin 1997.

44. B. Bensaid et M. Boutillier, « Le contrat notionnel : Efficience et efficacité », juillet 1997.

45. E. Jondeau et R. Ricart, « La théorie des anticipations de la structure par terme : test à partir des titres publics français », septembre 1997.

46. E. Jondeau, « Représentation VAR et test de la théorie des anticipations de la structure par terme », septembre 1997.

47. E. Jondeau et M. Rockinger, « Estimation et interprétation des densités neutres au risque : Une comparaison de méthodes », octobre 1997.

48. L. Baumel et P. Sevestre, « La relation entre le taux de crédits et le coût des ressources bancaires. Modélisation et estimation sur données individuelles de banques », octobre 1997.

49. P. Sevestre, “On the Use of Banks Balance Sheet Data in Loan Market Studies : A Note,” October 1997.

50. P.-C. Hautcoeur et P. Sicsic, “Threat of a Capital Levy, Expected Devaluation and Interest Rates in France during the Interwar Period,” January 1998.

51. P. Jacquinot, « L’inflation sous-jacente à partir d’une approche structurelle des VAR : une application à la France, à l’Allemagne et au Royaume-Uni », janvier 1998.

52. C. Bruneau et O. De Bandt, « La modélisation VAR structurel : application à la politique monétaire en France », janvier 1998.

53. C. Bruneau et E. Jondeau, “Long-Run Causality, with an Application to International Links between Long-Term Interest Rates,” June 1998.

54. S. Coutant, E. Jondeau et M. Rockinger, “Reading Interest Rate and Bond Futures Options’ Smiles: How PIBOR and Notional Operators Appreciated the 1997 French Snap Election,” June 1998.

55. E. Jondeau et F. Sédillot, « La prévision des taux longs français et allemands à partir d'un modèle à anticipations rationnelles », juin 1998.

56. E. Jondeau et M. Rockinger, "Estimating Gram-Charlier Expansions with Positivity Constraints," January 1999.

57. S. Avouyi-Dovi et E. Jondeau, "Interest Rate Transmission and Volatility Transmission along the Yield Curve," January 1999.

58. S. Avouyi-Dovi et E. Jondeau, « La modélisation de la volitilité des bourses asiatiques », janvier 1999.

59. E. Jondeau, « La mesure du ratio rendement-risque à partir du marché des euro-devises », janvier 1999.

60. C. Bruneau and O. De Bandt, "Fiscal policy in the transition to monetary union: a structural VAR model," January 1999.

61. E. Jondeau and R. Ricart, "The Information Content of the French and German Government Bond Yield Curves: Why Such Differences?" February 1999.

62. J.-B. Chatelain et P. Sevestre, « Coûts et bénéfices du passage d'une faible inflation à la stabilité des prix », février 1999.

63. D. Irac et P. Jacquinot, « L'investissement en France depuis le début des années 1980 », Avril 1999.

64. F. Mihoubi, « Le partage de la valeur ajoutée en France et en Allemagne », mars 1999.

65. S. Avouyi-Dovi and E. Jondeau, "Modelling the French Swap Spread," April 1999.

66. E. Jondeau and M. Rockinger, "The Tail Behavior of Stock Returns: Emerging Versus Mature Markets," June 1999.

67. F. Sédillot, « La pente des taux contient-elle de l'information sur l'activité économique future ? », juin 1999.

68. E. Jondeau, H. Le Bihan et F. Sédillot, « Modélisation et prévision des indices de prix sectoriels », septembre 1999.

69. H. Le Bihan and F. Sédillot, "Implementing and Interpreting Indicators of Core Inflation : The French Case", September 1999.

Pour tous commentaires ou demandes sur les Notes d'Études et de Recherche, contacter la bibliothèque du Centre de recherche à l'adresse suivante :

For any comment or enquiries on the Notes d'Études et de Recherche, contact the library of the Centre de recherche at the following address:

BANQUE DE FRANCE
41.1391 - Centre de recherche
75 049 Paris CEDEX
tél : 01 42 92 49 55
fax :01 42 92 62 92
email : thierry.demoulin@banque-france.fr