NOTES D'ÉTUDES

ET DE RECHERCHE

MODELING AND FORECASTING THE FRENCH CONSUMER PRICE INDEX COMPONENTS

Eric Jondeau, Hervé Le Bihan and Franck Sédillot

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Modeling and Forecasting

the French Consumer Price Index Components

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1. Introduction

This paper describes a forecasting model of the French consumer price index. Its purpose is to allow for rapid and detailed analysis of recent inflation developments, as well as frequent forecasts. Its characteristics are therefore the following: a small number of equations, a monthly frequency, and a fairly detailed sectoral breakdown of the price index. A feature of our forecasting procedure is also to systematically supplement the forecast with confidence intervals.

Our strong prior is that the different items within the overall index do not obey the same economic determinants. A disaggregated modelization might improve forecast accuracy. We therefore model and forecast that the various items of the price index according to a least three different methodologies. First the « core » part of the CPI (i.e. food excluding sensitive products, manufactured goods and private sector services) is modeled using reduced-form Phillips-curves.[1] Second, seasonal factors explain most of the fluctuations of the « sensitive » products component, especially fresh products. These may vary in price by as much as 10% from one month to the next. Third, the remainder of the price index is regarded as exogenous: increases in administered prices such as tobacco prices or health care services can be forecast in the short run based on policy announcement. Oil products are linked in en ECM-form to crude oil price, which is also considered exogenous. These products account for only about 5% of the index but may have a significant effect on economic assessments.

Given the main emphasis on forecasting, and the need for an easy-to-handle tool able to be used in junction with a broader forecast or expert judgment, the modelization keeps very simple and parsimonious. Some theoretical motivations for specifications are discussed, but since we aim at tracking short term developments, we do not impose strong theoretical constraints on the estimated equations.

The paper is organized as follows. In section 2 we describe the sectoral breakdown used. Section 3 provides econometric results for foodstuffs, manufactured goods and services. These

[1] This groups accounts for roughly two-thirds of the total index.

are of interest for several reasons. First, they show that the capacity utilization rate is a good predictor of the underlying component of the price index. For manufactured goods, however, this predictive capacity has weakened in the last two years. We show that this is due in particular to agents' lower expectations of inflation since 1995. Surprisingly, it appears that import prices have no effect on the price of manufactured goods. Lastly, wages are the main explanatory variable for prices in services. In section 4 we describe the methods used to forecast oil products and sensitive foodstuffs. For oil products, we find that refiners' mark-up behavior tends to dampen fluctuations in prices per barrel. In section 5 we consider methods for constructing the confidence intervals associated with the forecast, looking at three of them in particular: computation of RMSE, bootstrap simulation, and the approach developed by the Bank of England. Our overall conclusions are contained in section 6.

2. Which price index or indices should we model?

In aggregate macroeconomic models, the inflation rate is generally given by the household consumption deflator. In this context the forecast is based on a theoretical model and a comprehensive set of accounts. Although consistency and the theoretical foundation are the chief advantages of such models, modeling only an overall index is not entirely satisfactory, at least from a short-term standpoint. Indeed the different items within the overall index do not obey the same economic determinants. They may display very dissimilar pattern as can be seen from Figure 1. Furthermore, some components of the index are highly volatile or are not linked to economic determinants such as costs or demand pressures. Thereby, modeling a number of sub-indices seems to be a more promising approach. But what components should the model include?

This question is somewhat related to the issue of underlying inflation. Since Blinder (1982), analysis in terms of core inflation underline the fact that some components show a particularly erratic behavior, or are difficult to explain in terms of economic behavior. The Insee has published since 1996, the monthly growth rates of an index « excluding public tariffs and volatile products adjusted for tax measures and seasonal variations ».[2] We have followed Insee's underlying index spirit by leaving the same items out of the modelization. Altogether, the sectoral model includes the following three components: food excluding sensitive products, private sector manufactured goods and private sector services (including rents and water). Thus, approximately 70% of the index is modeled, the remaining 30% being regarded as exogenous (Table 1)[3].

The different indices are also adjusted for indirect taxation which, for the underlying component, means VAT. The index excluding VAT is evaluated using the following equation:

$$P_{htva} = \frac{P_{ttc}}{1 + r_{vat} / 100} \tag{1}$$

[2] See Insee (1996), Informations rapides, n°179.

[3] The construction of the indices is described in detail in Appendix 1. This Appendix 1 also documents the exogenous pattern of non-modelled indices.

where P_{htva} is the index excluding VAT, P_{ttc} is the index including VAT and r_{vat} is the rate of VAT. For the time being, the adjustment is made using VAT rates supplied by the National Accounts and calculated on the basis of tax receipts.[4]

In addition, we have modeled year-on-year price changes in order to avoid the problem of seasonal variations. Indeed the seasonal pattern of the French CPI has been evolving in recent years, preventing any attempt to model seasonality.[5] Lastly, in order to link changes in price series to economic variables, we use a quarterly frequency (end of quarter). In the course of the forecast, quarterly figures are the interpolated to deliver a monthly forecast. This does not apply for oil products and sensitive products, for which the equations are computed on a monthly basis.

Figure 2 shows the movement of the price index, the « index excluding public tariffs and volatile products adjusted for tax measures and seasonal variations » computed by Insee, and the « adjusted » index that we have reconstructed over a relatively long period from the 265 items in the index for base 1990 (and the 295 items for base 1980).[6] The movement of our « adjusted » index is quite different from that of the NCPI but relatively similar to that of the Insee index. In the forecast process, items are recombined to produce figures consistent with the HICP nomenclature.

Table 1: Breakdown and weighting of the model index

	Weighting
Items modeled (economic factors)	**7081**
of which « underlying » items	**6562**
Food excluding « sensitive » products	712
Private sector manufactured goods	3252
Private sector services (including rents and water)	2598
of which oil products	**519**
Items modeled (seasonality)	**1308**
« Sensitive » food products	1308
Items not modeled (exogenous)	**1611**
Tobacco	200
Electricity, town gas and coal	352
Public tariffs excluding energy	509
Healthcare services	550

Source: Insee, Banque de France calculations.

[4] Except for oil products where TIPP (a special tax on oil products) is included as well as VAT

[5] Especially for manufactured products.

[6] In fact, this involves reconstructing a chained Laspeyres index.

Figure 1
Sectoral inflation rates
(year-on-year change)

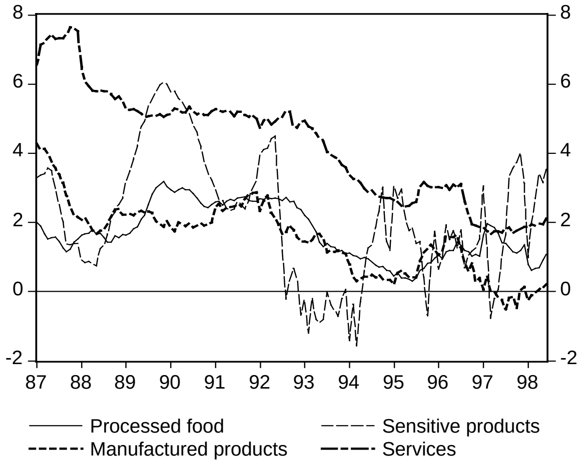
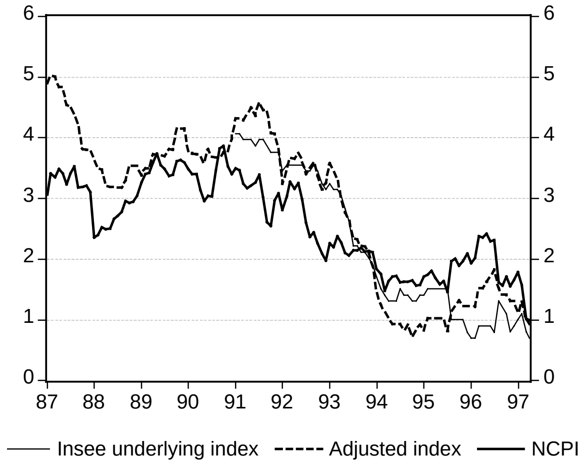


Figure 2
Adjusted inflation and « underlying inflation »
(year-on-year change)



3. Modeling the « underlying » group

Let us now turn to the econometric model of the first three items in Table 1. Three explanatory variables were felt to be relevant for the « underlying » index: the capacity utilization rate (CU), hourly labor costs in the manufacturing sector and hourly labor costs in the service sector[7]. These variables are lagged by one year, allowing for a forecast without extrapolation of exogenous explanatory variables for four quarters.[8] As we will see, the introduction of the capacity utilization rate has a theoretical framework and the model may thus be interpreted as an augmented Phillips curve. However, though hourly labor costs are considered as a natural variable, equations including such a variables have one drawback: they no longer may be interpreted as a reduced form but as a piece of a structural model.

3.1. Theoretical motivation

The variable most naturally suited to inclusion in the price equation is the capacity utilization rate in the manufacturing sector (CU). There are two advantages to using this variable: first, unlike an output gap measure, it is not estimated since it is observed directly from the quarterly survey of manufacturing industry;[9] second, it acts as a pressure on demand and as such is naturally included in an augmented Phillips curve. In many macroeconomic models, the inflation rate is modeled as follows:

$$\pi_t = \pi_t^e + \gamma(y_t - \bar{y}_t) \qquad (2)$$

where π_t is the inflation rate, π_t^e the expected inflation rate and $y_t - \bar{y}_t$ the output gap. If the process of expectations formation is non-adaptive, the Phillips curve is that of the new Keynesians. Using either Taylor's (1979) model of staggered wage setting over two periods, or Calvo's (1983) model, or Rotemberg's (1982) quadratic adjustment cost model, the Phillips curve equation is:

[7] Hourly labour costs are defined as hourly compensation plus employers' contributions.
[8] Except for data revisions.
[9] In addition, only the figures for the last two periods are revised.

$$\pi_t = E_t\pi_{t+1} + \gamma(y_t - \bar{y}_t). \tag{3}$$

However, this approach has a major drawback: if the price level is rigid, inflation adjusts instantaneously to its expected value. In this case a disinflationary policy may, if perfectly expected, be not costly in terms of activity. Phelps (1979), Blanchard (1983), Taylor (1983), Ball (1990) and Fuhrer and Moore (1995) have analyzed this case. However, such a finding has never been verified empirically (Ball, 1993). Fuhrer and Moore (1995), seeking behavior closer to that of adaptive expectations while preserving the rational aspect of expectations, slightly change to Taylor's model by introducing a negotiation on real rather than nominal wages. In this case, the augmented Phillips curve (for a two-period agreement) becomes:

$$\pi_t = \frac{1}{2}(E_t\pi_{t+1} + \pi_{t-1}) + \gamma(y_t - \bar{y}). \tag{4}$$

Lastly, Laxton et al. (1998) derive an augmented Phillips curve from a WS-PS model to which is added a P-bar type equilibrium price adjustment process developed by Mussa (1981). The equation system is written:

$$\begin{aligned} p_t &= w_t + \delta_0 - \delta_1 u_t \\ w_t &= p_t^e + \varphi_0 - \varphi_1 u_t \\ \pi_t &= \lambda(\bar{p} - p_{t-1}) + (1-\lambda)\pi_{t-1} \\ u_t - \bar{u} &= -\gamma(y_t - \bar{y}). \end{aligned} \tag{5}$$

The first equation of system (5) states that price is a mark-up on labor costs. The second equation describes the adjustment of nominal wages to expected prices and labor market pressures. The third equation is the process by which the price adjusts to its long-term target given by the first equation. Lastly, the fourth equation is an Okun law. By solving (5), we obtain the following augmented Phillips curve:

$$\pi_t = \lambda\pi_t^e + (1-\lambda)\pi_{t-1} + \lambda\gamma(\varphi_1 + \delta_1)(y_t - \bar{y}) \tag{6}$$

where $\bar{u} = \frac{\varphi_0 + \delta_0}{\varphi_1 + \delta_1}$ is the equilibrium unemployment rate and $\pi_t^e = p_t^e - p_{t-1}$.

Though we do not intend to estimate a forward-looking model for forecasting purposes, this theoretical background will prove useful in the interpretation of the results.

3.2. Manufactured goods

3.2.1. Using the capacity utilization rate alone

Utilization rates are part of price equation in many French macromodels (Economie et Prevision, 1998). Its inclusion in reduced form Phillips curve has been less often undertook. In the US. Cecchetti (1995) tests the correlation between inflation and seventeen indicators, including variables relating to production (NAPM, hours worked, CU, unemployment), certain prices (oil price, gold and exchange rate) and monetary policy (money supply, short-term interest rates, spread). The capacity utilization rate generally shows a correlation with inflation, though it is not by any means the strongest. However, its ability to predict inflation, carried out in an out-of-sample forecasting exercise, is systematically the best in terms of RMSE for a three-month to two-year time horizon.[10] Stock and Watson (1999) also show that CU helps to forecast inflation over a short term horizon (12 months). Figure 3 shows inflation and capacity utilization in the manufacturing sector (CU_m) lagged by four quarters (this corresponds to the most suitable lag in terms of $\overline{R}^2$). The correlation is plain to see, but has tended to become looser since 1995. The error correction model,[11] estimated over the period 87Q1 to 97Q4, is (**Model 1**):

$$\Delta\Delta_4 p_{m,t} = \underset{(3.10)}{-8.245} - \underset{(3.29)}{0.201}\left(\Delta_4 p_{m,t-1} - \underset{(4.60)}{0.502}\, CU_{m,t-5}\right) + \underset{(4.11)}{0.289}\,\Delta CU_{m,t-4}$$
$$- \underset{(2.65)}{0.713}\, dum901\_911 + \hat{\varepsilon}_t$$
$$\overline{R}^2 = 0.427,\ see = 0.370,\ \text{LM}(4) = 2.262\ (0.08),\ \text{ARCH}(4) = 0.124\ (0.97),$$
$$\text{JB} = 1.642\ (0.44)$$

where the dummy variable dum901_911 is 1 in 90Q1, -1 in 91Q1 and 0 elsewhere[12].

The dynamic simulation from 1988 Q1 (Figure 4) shows a fair adjustment until the beginning of 1996. Thereafter, however, it shows a growing gap between actual and simulated series. This appears to be due to the fact that there is a substantial gap between inflation and CU series at the end of the period (i.e., from 1995) (cf. Figure 3).

In order to improve the model, we therefore introduced a dummy variable equal to 1 for the period beginning 1995 Q1 and to 0 before. This gives the following error correction model, estimated over the period 87Q1 to 97Q4 (**Model 2**):

$$\Delta\Delta_4 p_{m,t} = \underset{(3.55)}{-9.674} - \underset{(3.70)}{0.285}\left(\Delta_4 p_{m,t-1} - \underset{(5.35)}{0.423} CU_{m,t-5} + \underset{(2.18)}{1.043} D95\_98_{t-1}\right)$$
$$+ \underset{(4.35)}{0.299}\,\Delta CU_{m,t-4} - \underset{(2.61)}{0.687}\, dum901\_911 + \hat{\varepsilon}_t$$
$$\overline{R}^2 = 0.455,\ see = 0.361,\ \text{LM}(4) = 3.104\ (0.03),\ \text{ARCH}(4) = 0.1274\ (0.89),$$
$$\text{JB} = 1.694\ (0.43).$$

The dynamic simulation of Model 2 is much more convincing (Figure 4). What is the justification for such a dummy variable? First, some changes have been made to the price index for manufactured goods in the recent period, notably the inclusion of sales. Second, car prices have fallen significantly in the last three years. The index excluding cars and clothing has an higher trend than the total index, but the overestimation is still apparent. Bearing in mind the underlying theoretical model, this would mean that agents' expectations of inflation changed during the review period.

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[10] Cf. Cecchetti (1995), Table 3, page 11.

[11] Given the uncertainty surrounding the degree of integration of inflation, we did not carry out formal integration tests. Nevertheless, we are aware that if variables are I(1), the OLS method used here provides biased and non-consistent estimates of the long-term parameters. Conversely, if variables are I(0), the equation in level may always be written in difference but the long-run target may not be interpreted as a long-run equilibrium level.

[12] Δ_4 denotes the year-on-year change.

Figure 3
Prices and CU
(year-on-year change and level)

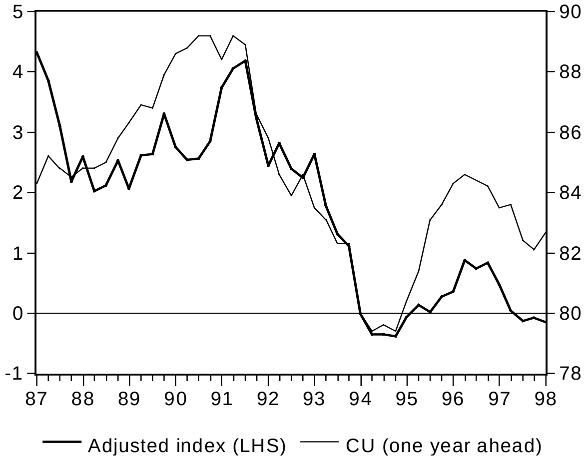
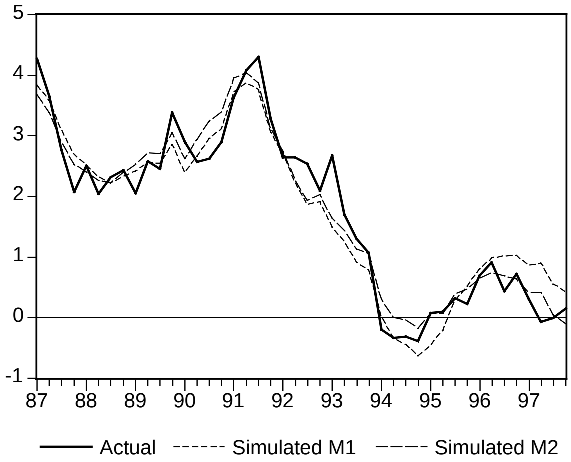


Figure 4
Dynamic simulations (Models 1 and 2)
(year-on-year change)



3.2.2. Taking inflation expectations into account

The first stage was to test the introduction of other theoretical determinants of inflation in the manufacturing sector, namely wages (hourly or per capita), operational labor costs (hourly or per capita) and unit labor costs. The wage variable is introduced on an ad hoc basis and does not form part of the reduced-form theoretical framework we have proposed. Expectations normally already include this variable.

Wage rate or labor cost variables have a considerable capacity to explain long-term inflation trends (level equation). Lagged by two quarters, the year-on-year growth rate of hourly operational labor costs is significant. It eliminates the dummy variable over the period 1995-98 without significantly changing the CU coefficient. How should this effect be interpreted? The sharp fall in inflation from 1992 to 1994 is fully taken into account by the decline of the capacity utilization rate. Wages continued to decline until the summer of 1995, when the minimum wage was raised (cf. Figure 5). This fall balances the rise in the capacity utilization rate from 1994, making the dummy variable unnecessary. However, as Figure 6 shows, the equation tends to slightly overestimate the year-on-year change in prices in 1997. Capacity utilization rates stabilized, whereas wages growth rose from 1.5% to 3%. However, the employers' contributions series used is taken from Quarterly Accounts and therefore does not take reductions in employers' contributions into account. As the regime has been gearing up since 1993, the labor cost to the employer ought to be lower than the series we have used. In fact, this variable is mainly helpful in explaining why prices did not rise in the same proportion as CU in 1994.

The error correction model, estimated from 87Q1 to 97Q4, is (**Model 3**):

$$\Delta\Delta_4 p_{m,t} = \underset{(3.33)}{-8.432} - \underset{(4.07)}{0.255}\left(\Delta_4 p_{m,t-1} - \underset{(4.32)}{0.381} CU_{m,t-5} - \underset{(2.44)}{0.553}\Delta_4 hlc_{m,t-3}\right)$$

$$+ \underset{(4.67)}{0.316}\Delta CU_{m,t-4} - \underset{(2.99)}{0.770} dum901\_911 + \hat{\varepsilon}_t$$

$\overline{R}^2 = 0.484$, $see = 0.351$, LM(4) = 2.868 (0.04), ARCH(4) = 0.968 (0.44), JB = 0.429 (0.89).

The wage-price elasticity is substantially less than 1.

Figure 5
Prices and hourly wage rates
(year-on-year change)

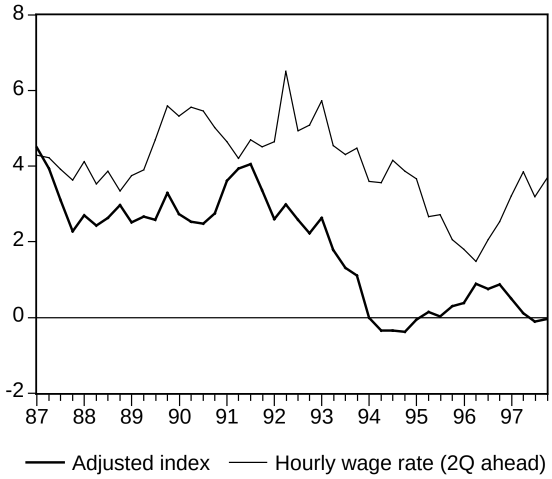
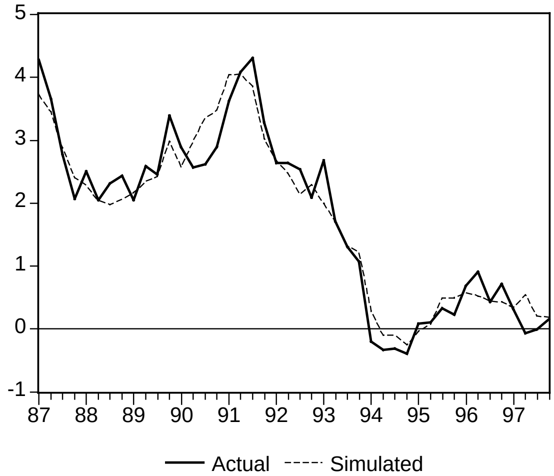


Figure 6
Dynamic simulation of Model 3
(year-on-year change)



The second stage was to estimate the expectation variable. Since this variable is not observable, it is natural to use the Kalman filter. For example, Bomfim and Brayton (1998) adopt this approach to extract US households' expectations of inflation, using a Taylor' rule as a reference equation. Laxton et al. (1998) use the same method to estimate the output gap with the help of a non-linear Phillips curve, though assuming inflation expectations. The tested model is very simple and has the following form:

$$\pi_t^e = \pi_{t-1}^e + \eta_t$$
$$\pi_t = \pi_t^e + \underset{(4.30)}{0.30}\,(CU_{m,t-4} - \overline{CU}_m) + \varepsilon_t$$

with $\sigma_\eta^2 = 0.009$ and $\sigma_\varepsilon^2 = 0.004$.

We also tested a time-varying coefficient for the CU parameter. The model then has the following form:

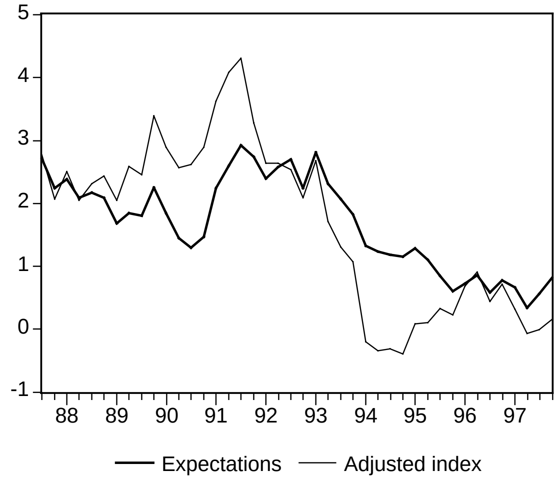
$$\pi_t^e = \pi_{t-1}^e + \eta_t$$
$$\beta_t = \beta_{t-1} + \xi_t$$
$$\pi_t = \pi_t^e + \beta_t(CU_{m,t-4} - \overline{CU}_m) + \varepsilon_t$$

with $\sigma_\eta^2 = 0.009$, $\sigma_\varepsilon^2 = 0.004$ and $\sigma_\xi^2 = 0.000$.

Estimation shows that the coefficient of the deviation of CU from its long-period mean is constant. Figures 7 and 8 show inflation expectations of the Phillips curve and the confidence interval of the expectation variable, computed from a Monte Carlo simulation (cf. Hamilton, 1986 and 1994). The break in expectations from 1995 is indeed significant, tending to validate the introduction of the dummy variable. There is also an uncoupling of inflation expectations and actual inflation at the beginning of the 1990s and, more obviously, between 1994 and 1996. At times of rising prices, expectations rise by less than actual inflation, whereas at times of falling inflation, expectations remain higher than actual inflation.

Figure 7
Inflation expectations
(90% confidence interval)

Figure 8
Expectations and inflation
(year-on-year change)



3.2.3. Taking account of other determinants

The theoretical model we have developed does not include the effects of the exchange rate on inflation. From this standpoint, the price index we have modeled so far would resemble the GDP deflator. Import prices have to be added if the equation is to give a complete description of the retail price index. The exchange rate has two pass-through effects: one is direct, because there is a direct content of imports in consumption, and one is indirect, since import content passes through intermediate consumption. In France, the direct content for end-products is 10%; the indirect content is of the same order of magnitude. For manufactured goods, the total content of imported products is 10%. It is therefore reasonable to expect a coefficient close to 0.1 in the long-term equation.

Import prices, measured by the manufactured goods imports deflator,[13] do not appear to be a significant variable in Model 3, long- or short-term. The objection could be raised, however, that the import price deflator is not constructed in the same way as the price index, since the former is a Paasche index and the latter a chained Laspeyres index. In the recent past, the consumer price deflator in the manufacturing sector calculated by the National Accounts has sometimes differed significantly from the CPI for private sector manufactured goods. We therefore tested the same equations, using the manufactured goods consumption deflator derived from quarterly Accounts rather than consumer prices. The results are identical: import prices are never significant. We also tested the introduction of both nominal and real exchange rates, again without success.

This result may appear counter-intuitive. The same result has been found using US data. For example Tootell (1998) introduces import prices excluding energy and energy prices into a standard Phillips curve and shows that only energy prices have a significant effect on inflation. He explains this result by the fact that, given the volatility of import prices excluding energy, only large variations have an effect on inflation. Such variations are significant when they exceed plus or minus one standard deviation around the mean. In fact, only significant rises

[13] More precisely we have used the deflator of imports of manufactured consumer goods, ie, excluding intermediate goods and part of the "automobile and land transport equipment" item. However, limiting ourselves to consumer goods alone has no effect on the results.

have an impact on US inflation. According to Tootell, the sharp falls in import prices excluding energy observed in the United States do not in fact seem to have affected inflation. Figure 9 allows to understand why this result arises. It shows that the CU trend and a two-year smoothing of import prices coincide, suggesting that importers set their margins according to prices charged by domestic producers and hence French production. Another interpretation lies on the fact there is a common cycle between France and its main commercial competitors. That would explain common changes of import prices and domestic and foreign CUs.

In order to gain a more precise measurement of the influence of import prices on changes in prices of manufactured goods, we added the smoothed growth rate of import prices to the long-term part of Model 3 with a constrained coefficient of 0.1, representing the import content of manufactured goods in consumption of French manufactured goods.

Figure 9
CU and import prices
(manufacturing sector, year-on-year change)

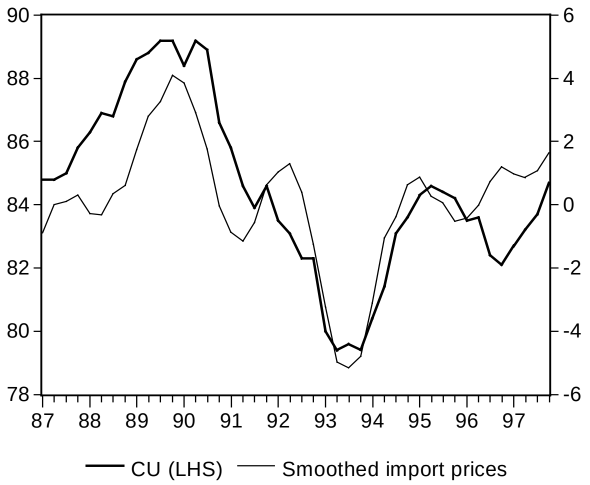
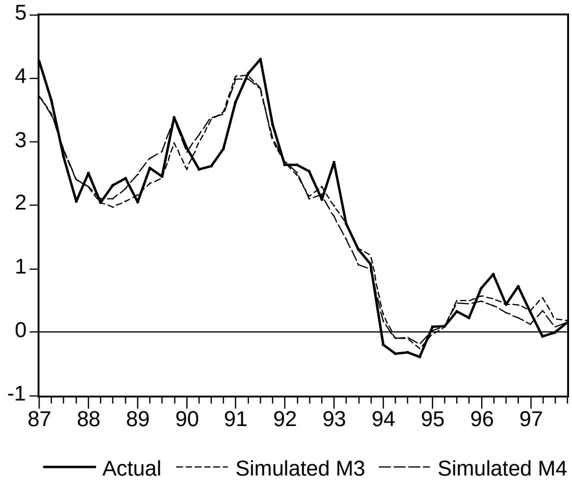


Figure 10
Dynamic simulations (Models 3 and 4)
(year-on-year change)



The error correction model, estimated from 87Q1 to 97Q4, is (**Model 4**):

$$\Delta\Delta_4 p_{m,t} = \underset{(3.47)}{-8.566} - \underset{(4.29)}{0.251}\left(\Delta_4 p_{m,t-1} - \underset{(4.41)}{0.374}\, CU_{m,t-5} - \underset{(2.89)}{0.6120}\, \Delta_4 hlc_{m,t-3} - \underset{(-)}{0.1} * \Delta_4 pimp_{m,t-1} \right)$$

$$+ \underset{(4.48)}{0.299}\, \Delta CU_{mt-4} - \underset{(3.06)}{0.770}\, dum901\_911 + \hat{\varepsilon}_t$$

$\overline{R}^2 = 0.500$, $see = 0.346$, LM(4) = 3.221 (0.02), ARCH(4) = 0.564 (0.69), JB = 2.760 (0.71).

The dynamic simulation is very similar to that of Model 3 (see Figure 10).

3.3. Services

The reduced form is not completely satisfactory. Thereby we add an additional variable, namely the hourly compensation plus employers' contributions -that is labor costs in the services sector (see Figure 11). The variable has been introduced for intuitive reasons, since labor costs may be expected to have a significant impact in sectors where the payroll accounts for the bulk of production costs and producers are in the position of price setters. The theoretical grounds for this equation appear to be looser than for the manufacturing case. The equation estimated may not be interpreted as a reduced form Phillips curve but as a part of a structural model. The introduction of such a variable in the service sector has more theoretical grounds that in the manufactured sector. In the latter sector, as markets are essentially local, companies can more easily pass a wage shock on to prices. The manufacturing sector, being subject to fierce international competition, has much less room for manœuvre in this respect. This finding has also been borne out when tested on American data (cf. Brauer, 1997). Nevertheless, with French data, export prices do not appear to play a significant role in the manufactured price setting (cf. *supra*).

Econometric results partially validate this intuition. As Figure 11 shows, the year-on-year change in service prices excluding VAT ($\Delta_4 p_s$) is correlated to the two-year smoothed year-on-year change in hourly labor costs in the service sector ($\Delta_4 hlc_s$). The long-term equation includes this variable. Adding the capacity utilization rate in the manufacturing sector improves the equation without calling the inclusion of wages into question.[14] The error correction model, estimated from 88Q1 to 98Q1, is:[15]

$$\Delta\Delta_4 p_{s,t} = \underset{(3.31)}{-4.413} - \underset{(2,26)}{0.119}\left(\Delta_4 p_{s,t-1} - \underset{(3.62)}{0.420}\,CU_{m,t-5} - \underset{(3.50)}{1.299}\,\Delta_4 hlc_{s,t-5}\right) - \underset{(7.80)}{1.534}\,dum881 + \hat{\varepsilon}_t$$

$$\overline{R}^2 = 0.698,\ see = 0.163,\ \text{LM}(4) = 0.965\ (0.44),\ \text{ARCH}(4) = 0.641\ (0.64),$$
$$\text{JB} = 1.541\ (0.46).$$

Figure 11
Prices and wages, private sector services
(year-on-year change)

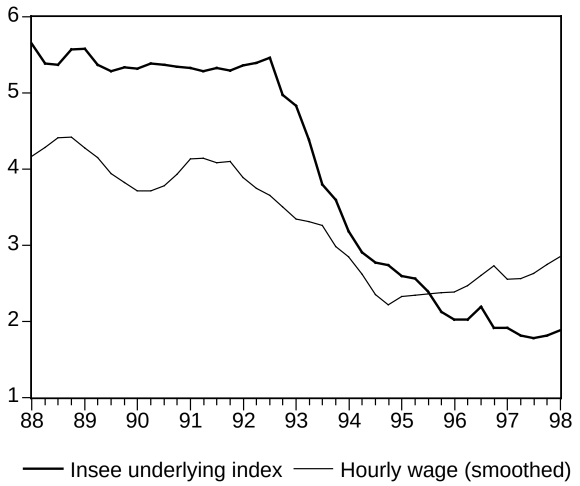
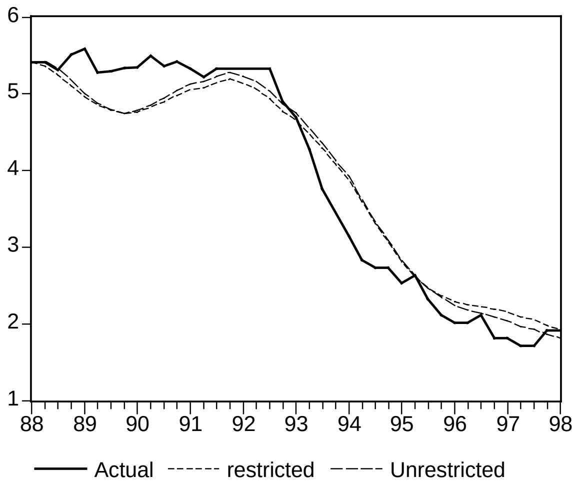


Figure 12
Dynamic simulations
(with restricted wage coefficient or not)



The dynamic simulation since 1988 is satisfactory, as Figure 12 shows. However, the price-wage elasticity is greater than 1. We have therefore restricted it to 1, which leads to the following error correction model, estimated from 88Q1 to 98Q1:

[14] We use this variable as a proxy for utilization rate in the service sector.
[15] The introduction of a dummy variable dum881 is justified by the fact that the index grew strongly in the first quarter of 1987 (as a consequence of the liberalization of prices such as hair dressers, cafés, private vehicle maintenance, ...) and, as a result of the base effect, fell significantly in the first quarter of 1988.

$$\Delta\Delta_4 p_{s,t} = \underset{(3.23)}{-4.027} - \underset{(2.31)}{0.098}\left(\Delta_4 p_{s,t-1} - \underset{(3.78)}{0.481} CU_{m,t-5} - \underset{(-)}{1.0}\Delta_4 hlc_{s,t-5}\right)$$
$$- \underset{(8.79)}{1.463} dum881 + \hat{\varepsilon}_t$$
$$\overline{R}^2 = 0.702,\ see = 0.162,\ \text{LM}(4) = 0.976\ (0.43),\ \text{ARCH}(4) = 0.599\ (0.67),$$
$$\text{JB} = 1.509\ (0.47).$$

Dynamic simulation of this model gives a quarter-by-quarter forecast very close to the preceding one (cf. Figure 12).

3.4. Food excluding "sensitive" products

The disadvantage of excluding the most volatile components is that the adjusted index covers only 35% of the food products index excluding tobacco (i.e. about 8% of the total). We have modeled this index on a quarterly basis (end of quarter). The index shows a correlation with the capacity utilization rate in the manufacturing sector lagged by one year and in the agri-food sector lagged by two quarters (Figure 13). As for the manufacturing case, this equation may be interpreted as an augmented Phillips curve.

Figure 13
Adjusted index and CU
(year-on-year change and level)

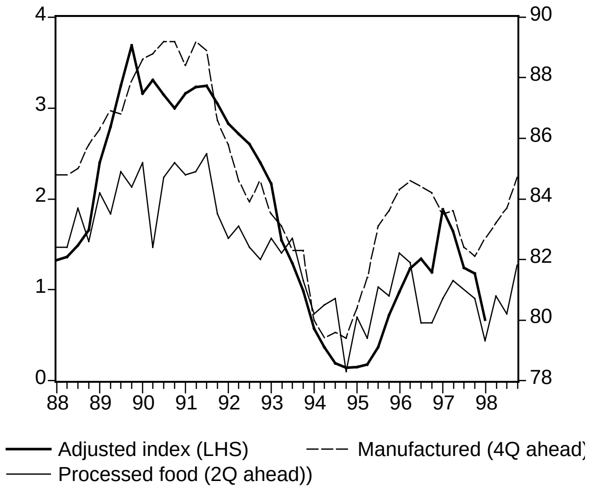
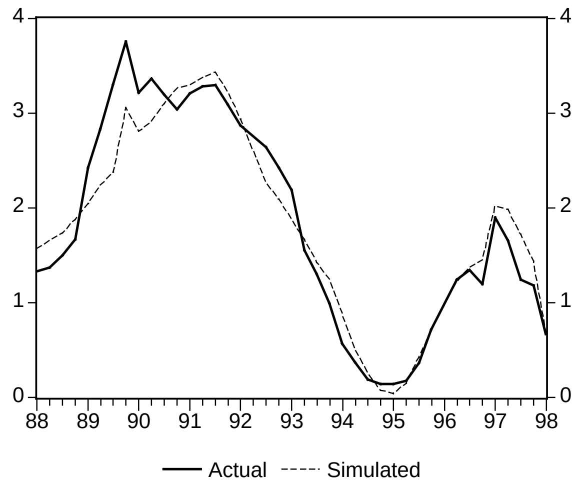


Figure 14
Dynamic simulation of the adjusted index
(year-on-year change)



The equation, estimated quarterly from 87Q3 to 98Q1, has the following form:

$$\Delta\Delta_4 p_{f,t} = -\underset{(2.20)}{7.101} - \underset{(3.18)}{0.217}\left(\Delta_4 p_{f,t-1} - \underset{(7.18)}{0.407} CU_{m,t-6}\right) + \underset{(1.72)}{0.204}\Delta\Delta_4 p_{,ft-1}$$
$$+ \underset{(2.24)}{0.100}\Delta CU_{m,t-4} + \underset{(2.49)}{0.100}\Delta CU_{m,t-5} + \underset{(3.41)}{0.483} dum894\_901 + \underset{(4.13)}{0.590} dum971\_981$$
$$\overline{R}^2 = 0.613,\ see = 0.195,\ \text{LM}(4) = 0.662\ (0.63),\ \text{ARCH}(4) = 0.884\ (0.48),$$
$$\text{JB} = 3.132\ (0.21).$$

where the dummy variable dum971_981 is 1 in 97Q1, -1 in 98Q1 and 0 elsewhere. This is justified by a shock on the adjusted index in the first quarter of 1997 (attributable, according to Insee, to new invoicing rules in the distribution sector), causing a base effect on the year-on-year change which shows up in the first quarter of 1998.

The equation has exactly the same structure as the equation for the manufacturing sector, with the same explanatory variables (CU) as for manufacturing industry. The dynamic simulation is relatively convincing (Figure 14).

4. Modeling the other products

Oil products and sensitive food products constitute a significant proportion of the overall consumer price index (18%). More importantly, they explain much of inflation volatility and as such play a major role in forecasting. Each component has its own specific feature: indirect taxation for oil products, and marked seasonal variations for so-called sensitive products. For energy, the price index, after adjustment for indirect taxation (VAT and TIPP, a special tax on oil products), depends only on the oil price per barrel in French francs. For sensitive products, especially fresh produce, seasonal dummy variables seem to give satisfactory results, at least initially.

4.1. Oil products

The oil products sub-index is both the largest and the most volatile part of the energy component of the CPI. Movements of the index are attributable to two factors: oil prices and taxation. We shall pay particular attention here to the second factor, since the taxation of oil products is both specific and significant. A special tax on oil products, TIPP, and other taxes of minor importance are added to ex-refinery prices. Taken together, these taxes constitute an excise duty corresponding to a fixed amount per unit of product sold. VAT is then levied on the ex-refinery price plus excise. The following equation gives the retail price of a litre of super, for example:

$$P_{ttc} = (P_{refinery} + TIPP) * (1 + r_{VAT}) . \quad (7)$$

Taxation is thus exogenous, since tax rates are set in the Finance Act (disclosed at the end of September) and on occasion in interim budgets (e.g., the increases in the summer of 1993).

France does import all its oil. Intuitively, ex-refinery prices depend on oil prices in French francs and the refiner's margin, as shown in Figures 15 to 19.[16] In particular, we can see that between 1995 and 1996, refiners did not pass on the entire rise in Brent prices (in FRF) to prices excluding taxes. Their margin therefore shrank. Since mid-1997, however, the opposite occurred. Each month, the hydrocarbons department at the Industry Ministry provides the pump price of super, unleaded super, diesel and fuel, together with ex-refinery prices for the same products.[17] We can also compute the TIPP and other excise taxes levied on each of these products. Figure 19 shows the oil product price index computed from a dynamic simulation of the four equations.

Figure 15
Price of super (excl. taxes, log.)

Figure 16
Price of unleaded super (excl. taxes, log.)

[16] The equations are given in Appendix 3.
[17] The data are available one month before the release of the inflation figure.

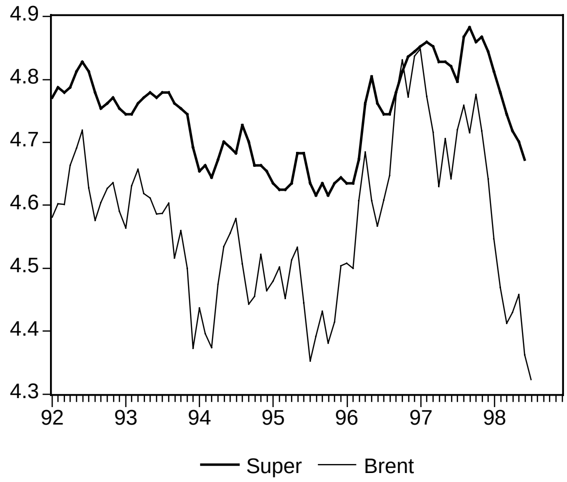


Figure 17
Price of diesel (excl. taxes, log.)

Figure 18
Price of fuel oil (excl. taxes, log.)

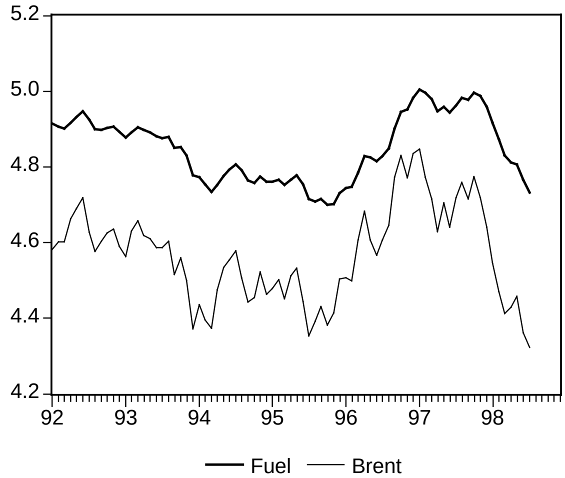
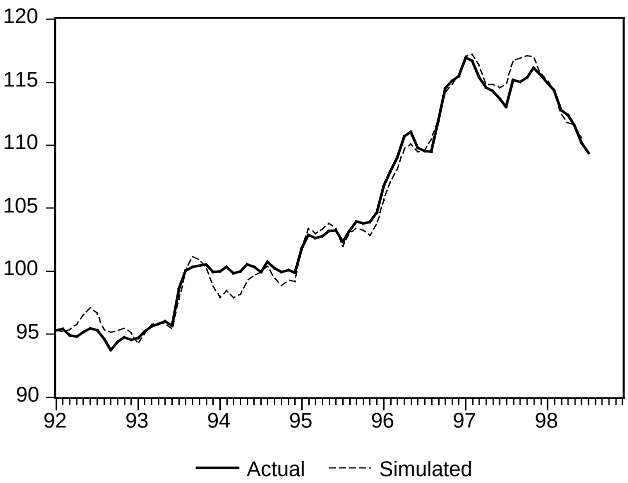


Figure 19
Dynamic simulation of the price of oil products
(level)



4.2. Sensitive food products

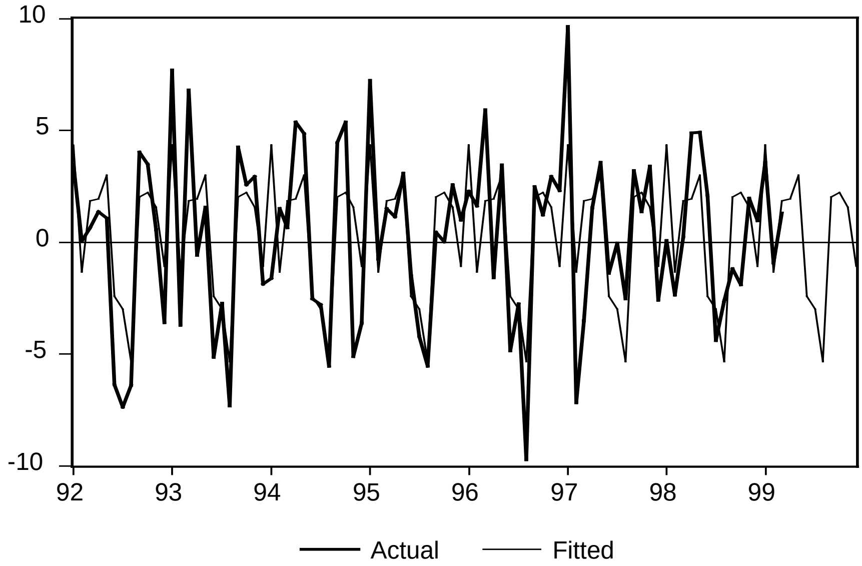
There are four main elements in this item: fresh produce and meat; tea, cocoa and chocolate; dairy products, cheese and eggs; wines and spirits. Tea, cocoa and chocolate are modeled using the import price of these raw materials, published monthly. Fresh produce and dairy products, eggs and cheese are subject to marked seasonal variations. They are modeled using seasonal dummy variables whose explanatory capacity appears to be satisfactory. Given a zero growth rate since 1992, movements are due solely to unforeseeable factors such as frost, etc., since the coefficients of seasonal variations seem to be stable. In these cases, the year-on-year change may show considerable fluctuations (sharp rises followed by sharp drops) attributable to previous shocks. The following equation provides an example for the fresh food and is estimated from 91M1 to 97M12:

$$\Delta p_{ff,t} = \sum_{i=1}^{12} \alpha_i dm_{i,t} + \varepsilon_t \text{ where } dm_i \text{ is a seasonal dummy for month } i$$

$$\overline{R}^2 = 0.55,\ see = 2.90,\ \ \text{DW} = 2.04.$$

Figure 20 displays a simulation of the previous equation. It fits the data rather well. When the forecast is implemented, the pattern of the year-on-year change may be stumbled if the historical value were subject to a change not predicted by the seasonality.

Figure 20
Simulation of the price of fresh food
(monthly growth rate)



4.3. Other products

Other products are mainly modeled using a judgmental approach. For instance, the tobacco component is relatively easy to predict, since the price of tobacco is left to the government's discretion (see Figure A1.6 in Appendix 1). Price rises generally take effect in January of each year but are announced in the previous September when the Finance Bill is presented. Sometimes, as in 1995, rises may be introduced in an interim budget.

5. Confidence intervals and distribution of the forecast

It seems valuable to associate the delivery of a forecast with some measure of the uncertainty surrounding it. The confidence interval of a forecast may be constructed in several ways. We shall consider here three cases: in-sample dynamic simulations, bootstrap and the method proposed by the Bank of England.

5.1. In-sample dynamic simulations method

First, it may be computed from errors in the context of an in-sample dynamic simulation. This method measures the error resulting directly from the use of the equations for forecasting purposes. The interval usually represented, of plus or minus two standard errors, corresponds to an interval in which the forecast will lie with a probability of 95%, assuming that error distribution is normal. Figure 21 gives an example for April 1999. We have added nine confidence intervals around this mean forecast, covering 10% to 90% of the probability distribution. It will be seen that the interval widens slightly with the time horizon of the forecast. In addition, as the residuals follow a normal distribution whose parameters are known, the probability density can be easily described (Figure 22).

Figure 21
Fan chart – under normality assumption

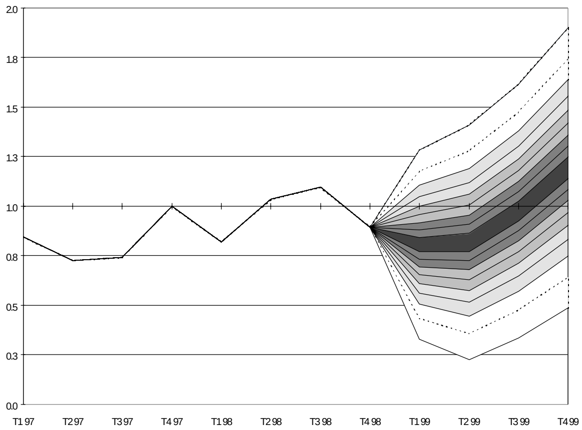
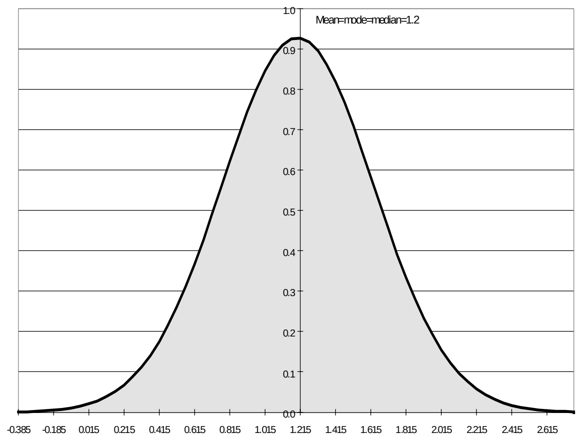


Figure 22
Probability distribution in December 1999

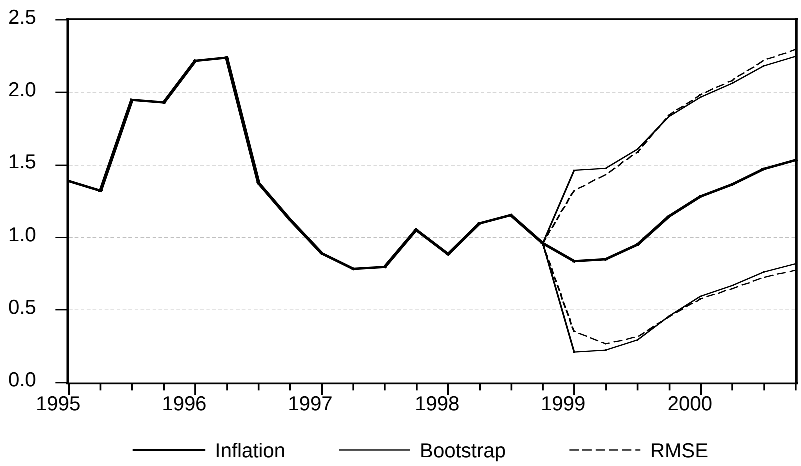


x axis = modeled inflation

5.2 Bootstrap method

Second, the bootstrap method can be used. This method, unlike the first, does not assume normal error distribution. It works as follows. Once the equation has been estimated (this equation is called the reference equation), we have T residuals from the estimation period. These T residuals are reordered according to a uniform distribution and a new inflation series is computed using the reference equation. The equation is re-estimated (with the same specifications as the reference equation) using the new inflation series and a forecast is made for one to eight quarters. This operation is repeated N times. This then gives N forecast series of one to eight quarters, i.e., a complete distribution for each quarter of the forecast. A confidence interval is then deduced. Figure 23 shows that the confidence intervals obtained from the root mean square error (RMSE) and the bootstrap are very close.

Figure 23
90% confidence intervals
Standard deviations using the RMSE and bootstrap methods



It will be seen that the confidence interval widens rather rapidly (two quarters) then stabilizes, whereas uncertainty might be expected to increase steadily with the time horizon. This is due to the fact that the random element of the exogenous variables is not taken into account. In this case, as the equations are similar to a stationary autoregressive model, the forecasting error at *k* quarters approaches a constant. Lastly, the standard deviations are derived from an in-sample simulation. In view of the limited number of observations, it was not possible to conduct this exercise outside the estimation period. An out-of-sample evaluation would probably give a forecasting interval that widens over time.

The density of the forecast associated with the two methods described above is based solely on the processing of previous errors. In the case of RMSE, the distribution is symmetrical by construction. With a bootstrap method, while asymmetry is possible, the equations residuals in our example are normal, and the distribution resulting from the resampling process has little chance of being asymmetrical. Yet it is often the case that when assessing a given situation, experts point to an asymmetry in the balance of risks. We therefore consider the Bank of England methodology, explicitly designed to tackle such an issue.

5.3. The Bank of England's approach

5.3.1. Density forecast for a specific quarter

The method applied by the Bank of England (Britton, Fisher and Whitley, 1997) differs from the two approaches described above in that it assumes that the density probability underlying the forecast is not related to the estimation process.[18] In addition, the density used is not the normal distribution but a two-piece normal distribution.[19] This means that two normal distributions with identical means but different standard deviations are used on each side of the mode (see Appendix 4). It is fully defined by three parameters: the mode (in this case, the means of the two original distributions), and the two standard deviations (σ_1 and σ_2). The main advantage of this distribution is that it is asymmetrical although it has only a limited number of parameters (three).

5.3.2. Introduction of asymmetry and uncertainty into the forecast

[18] This approach is more closely related to a bayesian one.

[19] The distribution presented in Appendix 4 is that of the Bank of England, but written with another parametrization (see Wallis, 1999).

Each parameter must now be determined. Mode, i.e. the most likely inflation scenario, is defined as being the central inflation forecast established by the central bank. The distribution variance can easily be derived either by simulating a model or, more simply, from the RMSE of past forecasting errors. The Bank of England has this series dating back to 1985. Evaluating the deviation between mode and mean is a bit more complex. This parameter reflects the risk asymmetry and plays a key role in establishing the forecast. The computation calls for expert judgment as well as the use of economic models (Britton, Cunningham and Whitley, 1997). An example of such an evaluation in the case of France based on April 1999 is given in Table 2 below. A description of its construction is given in the Appendix 5. To summarize, the deviation between the mean and the mode is the sum of the effects on inflation of shocks on different variables, weighted by the probability of their occurrence and likely magnitude.

Table 2: Determination of the mean-mode deviation over four quarters

	« High » risk (%) (1)	« Low » risk (%) (2)	Risk balance (%) (3)=(1)-(2)	Impact of a shock of one std deviation (4)	Uncertainty coefficient (5)	Expected impact on inflation (Q4 1999) (6)=(3)x(4)x(5)
CU	25	75	-50	0.20	1.10	-0.11
Wages	40	60	-20	0.38	1.00	-0.08
Prices	40	60	-20	0.34	1.25	-0.08
Total impact (mean-mode) :						-0.27

Now that we have the empirical variance and the difference between the mode and the mean, the method of moments allows us to determine σ_1 and σ_2 using the following two equations which provide the first two moments of the two-piece normal distribution:

$$E(X) = \mu + \sqrt{2/\pi}(\sigma_2 - \sigma_1)$$
$$V(X) = \left(1 - \frac{2}{\pi}\right)(\sigma_2 - \sigma_1)^2 + \sigma_1\sigma_2. \tag{7}$$

5.3.3. The fan chart

Once all parameters have been estimated, the confidence intervals associated with a given forecast can easily be computed. The Bank of England uses the following method. Two equally probable points are selected on each side of the mode, so that 10% of the distribution is covered. Two other points with the same probability are then selected in order to cover 20% of the distribution. This process is repeated until 90% of the distribution is covered. Each of these bands is then shaded, ranging from very dark for the narrowest band to very light for the largest band. In this illustration of density, the mode is always in the darkest band, but this is not necessarily the case for the median or the mean, especially if the distribution is asymmetrical. Figure 24 shows an example using the mean-mode deviation derived from Table 2. A comparison of Figures 21 and 24 clearly shows the downside asymmetry of the distribution.

Figure 21 (for comparison)
Fan chart-assuming normality
Inflation forecast for France - symmetry

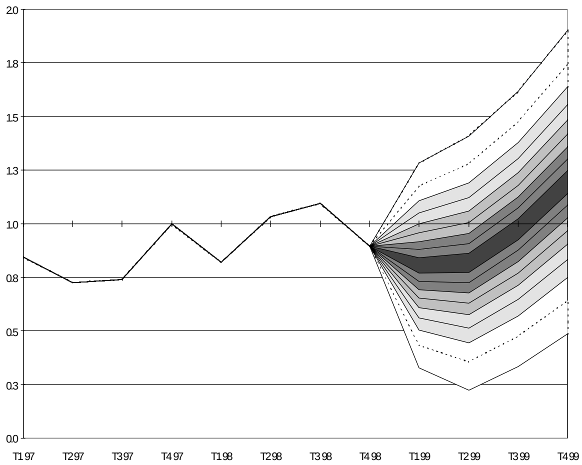
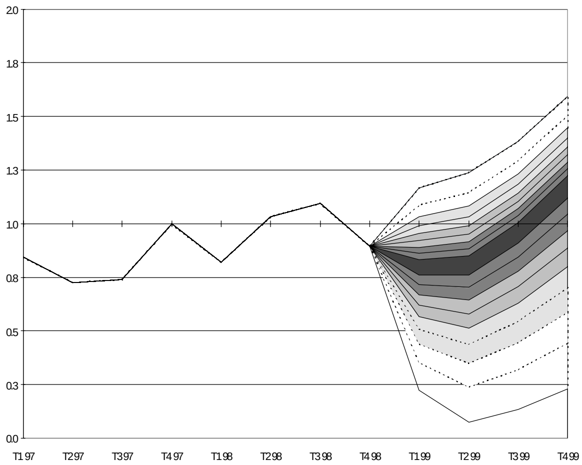


Figure 24
Fan chart - mode-centered intervals
Inflation forecast for France - asymmetry



5.3.4. An illustration of Wallis's criticism of the fan chart

The presentation adopted by the Bank of England deserves to be regarded cautiously (Wallis, 1999). Firstly, the concept of confidence intervals does not match with the one traditionally used. When one is presented with an interval covering 90% of the distribution, one tends to assume that this excludes 5% to the right and 5% to the left. The probability of falling outside of the band is therefore equal on the right and the left. However, this is not exactly the case in the preceding method. There is a greater probability that inflation will overshoot the upper bound of the interval when there is a positive asymmetry. This is because the interval is computed on the basis of the mode, which in this specific case lies to the left of the median, rather than on the median itself. Wallis recommends the use of a centered interval. This interval is wider than the preceding one and, if the asymmetry coefficient is positive, it will be skewed to the upper side of the preceding interval, thus seeming to reflect a higher price increase corresponding roughly to the difference between the median and the mode.

Figures 24 and 25 compare the impact of the different presentations on our inflation forecast for France. It should be remembered that we introduced here a downward asymmetry. In Figure 25, the darkest band is centered on the median: at the end of the period this band includes the value 1%. By contrast, in Figure 24 the same band is centered on the mode, which is higher, and significantly diverges from 1%, giving an impression of higher inflation. These results are detailed in Table 3, which shows the confidence intervals computed with the two methods for the forecast with a time horizon of four quarters. For example, given a 10% threshold, a presentation similar to that of the Bank of England results in an interval of [1.12%, 1.22%]. There is a 63.5% probability that inflation will overshoot the band and a 26.5% probability that it will undershoot the band. With Wallis's method, a forecast interval of [0.92%, 1.03%] is obtained. There is thus an equal probability of inflation being higher or lower than the band (45% for a threshold of 10%).

Figure 24 (for comparison)
Fan chart - mode-centered intervals
Inflation forecast for France - asymmetry

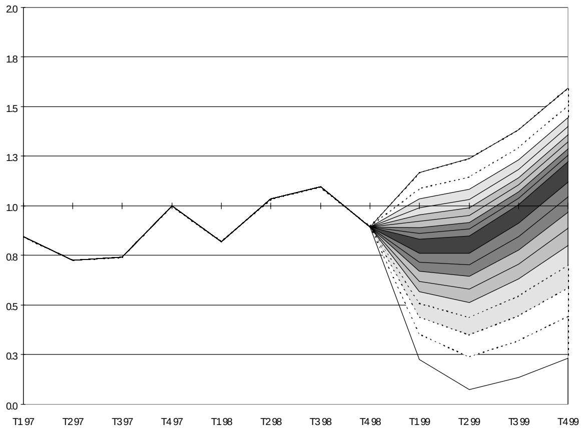


Figure 25
Fan chart- median-centered intervals
Inflation forecast for France - asymmetry

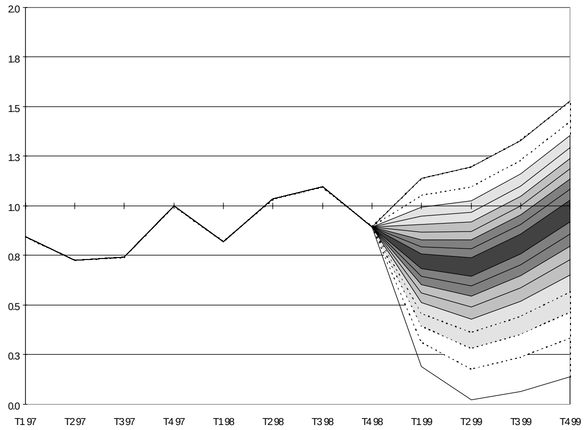


Table 3 : France - forecast intervals computed according to two methods (forecast horizon: December 1999)

	Bank of England approach				Wallis approach			
Threshold %	% < Lower bound	Lower bound	Upper bound	% > Upper bound	% < Lower bound	Lower bound	Upper bound	% > Upper bound
10	63.5	1.12	1.22	26.5	45	0.92	1.03	45
20	56.5	1.04	1.25	23.6	40	0.86	1.08	40
30	49.4	0.97	1.29	20.6	35	0.79	1.14	35
40	42.3	0.89	1.32	17.7	30	0.73	1.19	30
50	35.3	0.80	1.36	14.7	25	0.65	1.24	25
60	28.2	0.70	1.40	11.8	20	0.57	1.29	20
70	21.2	0.59	1.45	8.8	15	0.46	1.35	15
80	14.1	0.44	1.51	5.9	10	0.33	1.43	10
90	7.1	0.23	1.59	3.0	5	0.14	1.53	5

The two presentations have the same underlying distribution and the choice between them is, to a certain extent, arbitrary. However, Wallis contests the use of the mode of distribution rather than the mean or the median as a central scenario on statistical grounds. This approach consists in selecting the most likely inflation figure, regardless of the distribution. The loss function implicit in this forecast considers all forecast errors to be equally wrong, irrespective of their magnitude.[20] The mean, however, minimizes the mean square error. Wallis (1999) recommends using the median, which in fact minimizes the mean forecast error in absolute terms.

[20] In fact, this method results in the smallest distance between the upper and lower bounds for the probability threshold chosen. This is illustrated in Table 3. The values covered by the 90% interval is 1.36% for the mode-centred interval and 1.39% for the median-centred interval.

6. Conclusion

This sectoral model of the price index has several advantages. From a practical standpoint, forecasts can be carried out fairly quickly. However, the method is not designed to replace a comprehensive, complete forecast conducted using a macroeconomic model. From an economic standpoint, a relatively detailed level of aggregation makes it possible to identify shocks which have affected or might affect certain sub-items of the index, and thus to distinguish between some very transitory disturbances from more persistent shocks.

Even if not enough time has passed for any valid retrospective assessment to be made, the results of the first three forecasting exercises point to a certain number of features. The behavior of the equations for the agri-food and services sector is satisfactory. However, the two equations that can be used for manufactured goods do not track recent developments correctly. They tend to overestimate the rise in prices in the manufacturing sector, even if the introduction of a break in the constant has significantly improved the results.

Anyway, further extensions of this work still remain ahead. Ongoing research focus on the use of survey data, comparison with results obtained using multivariate methods (such as VAR or BVAR) and the estimation of some equations on a monthly frequency which would allow to carry out more formal tests on the out-of-sample accuracy of the model.

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Appendix 1: Composition and tax rates of the indices

The tables in this appendix present the detailed composition of each group of the price index, together with rates of VAT. In our model rates of VAT for each sector are calculated using apparent rates from quarterly National Accounts, weighted for the components of each sector consumption as a share of total purchases of each sector good and services. This approximation is rather crude, since the quarterly rates provided by the National Accounts are calculated on the basis of receipts rather than tax rates. A more accurate calculation would require the real VAT rates for each of the elementary components of the CPI. The Appendix 2 shows the weight of VAT in 1998 for the various components: 83.4% of the items in the sub-index are subject to VAT at 20.6%, 3.8% to VAT at 5.5% and 12.8% to VAT at 2.1%. Since tax receipts might be affected by random fluctuations, differences between the apparent rate and the official rate might occur and bias the forecast. This Appendix tries to appraise this bias comparing our excluding tax index with the INSEE underlying index which uses official rate at a detailed level.

1. Price index in the manufacturing sector

In 1998, the price index for private sector manufactured goods represented 32.52% of the national consumer price index (NCPI). It is seasonal, to take account of twice yearly clothing and footwear sales. A number of methodological changes have been made since 1995, such as the inclusion of sales, tax breaks for new vehicle purchases, etc., that have amplified the fluctuations of monthly growth rates. The composition of the index is given in Table A1.1.

Figure A1.1
Price of manufactured goods
(Year-on-year change)

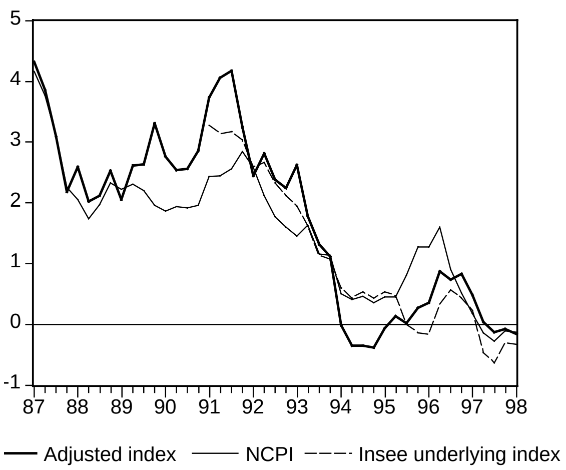


Figure A1.2
Adjusted manufactured goods index

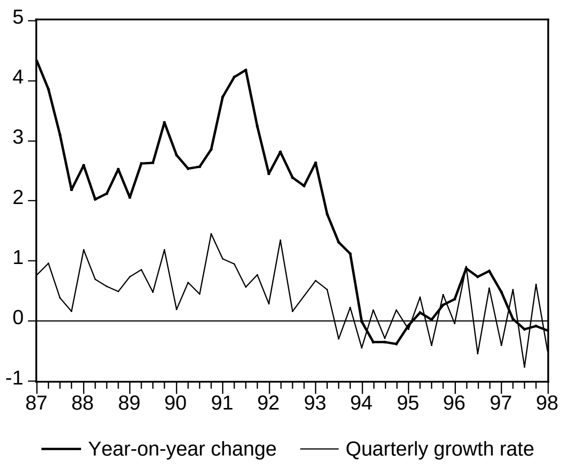


Table A1.1: Composition of the price index for manufactured goods

Item	Weight	BMS N°	Indirect taxation
Manufactured goods	**3252**		
Sub-total clothing and footwear[1]	*637*		
- Clothing	513	21	VAT at 20,6%
- Footwear	124	22	VAT at 20,6%
Sub-total other manufactured goods	*2615*		
- Furniture, carpets and floor coverings	187	41	VAT at 20,6%
- Household textile goods, other furnishings	87	42	VAT at 20,6%
- Domestic appliances	107	43	VAT at 20,6%
- Glassware, crockery, household utensils	91	44	VAT at 20,6%
- Non-durable household goods	130	451	VAT at 20,6%
- Pharmaceutical products	342	51	
Specialty pharmaceutical products	321	5101	VAT at 2,1% or 5,5%[2]
Parapharmaceutical products	21	5102	VAT at 20,6%
- Therapeutic devices	50	52	VAT at 5,5% or 20,6%[3]
- Tyres	22	6211	VAT at 20,6%
- Vehicle spare parts	296	6212	VAT at 20,6%
- Vehicle purchases	404	61	VAT at 20,6%
- Lubricants	28	6221	VAT at 20,6%
- Radio, television	70	711	VAT at 20,6%
- Photography, music, other durable goods	34	712	VAT at 20,6%
- Other recreational items	253	713	VAT at 20,6%
- Books, periodicals and newspapers	168	73	VAT at 5,5% or 2,1%[4]
- Grooming products	190	812	VAT at 20,6%
- Other personal items	156	82	VAT at 20,6%

Source: Insee, BMS.
Notes: [1] Excluding repairs. [2] The 2.1% rate applies to prescriptions reimbursed by Social Security. [3] The 20.6% rate applies to medical optics (spectacles, etc.) and the 5.5% rate to prosthetic devices. [4] The 2.1% rate applies to press products (except in certain special cases) and the 5.5% rate to other printed and published products.

Figure A1.1 shows the growth of the underlying index calculated by Insee (it is adjusted for tax measures and has the same scope as the index of manufactured goods with two exceptions)[21] and our price of private sector manufactured goods excluding VAT. Changes are similar, though a scissors pattern is in evidence from the beginning of 1994 to mid-1995, when the adjusted index (excluding VAT) passes below the INSEE underlying index. This difference is probably due to the non-inclusion of certain government measures, such as incentives to buy new vehicles (catalytic exhausts, premium for scrapping old vehicles, etc.), which had an isolated effect on the adjusted index (excluding VAT).

[21] These are "flowers and plants" and "specialty pharmaceutical products", the weight of which is not negligible (3.21% of the total index but 9.9% of the "private sector manufactured goods" sub-index).

2. The price index for private sector services

The price index for private sector services represents slightly less than 20% of the CPI. If rents and water are added, this proportion rises to 26% (Table A1.2). The only indirect tax on services is VAT. VAT is levied on private sector services at three rates: 7.4% of products are exempt, 17.9% are subject to VAT at 5.5% and 74.7% at the standard 20.6% rate.

Figure A1.3 shows our VAT-adjusted index and the Insee underlying index. They are similar except when the standard rate of VAT was raised in 1995, resulting in a discrepancy on the year-on-year change over 12 months. Given the similarity of the two indexes and in order to avoid introducing a dummy variable in 1995-96, we have modeled the Insee underlying index. It has been backcast with our adjusted index for the period prior to the first quarter of 1991.

Table A1.2: Composition of price index for private sector services

Item	Weight	BMS N°	Indirect taxation
Services	**2598**		
Sub-total private sector services	*1916*		
- Clothing and footwear repairs	7	23	VAT at 20,6%
- Rural rents	32	3112	Exempt
- Dwelling services	206	452	VAT at 20,6%
- Private vehicle repairs	234	6213	VAT at 20,6%
- Private vehicle maintenance	25	6214	VAT at 20,6%
- Driving schools	18	6232	VAT at 20,6%
- Private vehicle rental	12	6233	VAT at 20,6%
- Transport services (excluding rail and urban transport)	138	6303+6304 +6305+6306	VAT at 5,5%
- Repairs of recreational items	30	714	VAT at 20,6%
- Leisure and culture excl. license fees	150	72 excl. 7201	VAT at 5,5% or 20,6%[1]
- Teaching	47	74	Exempt
- Hairdressing, beauty care	99	811	VAT at 20,6%
- Restaurants, cafés, hotels (excl. school or university canteens)	772	83 excl. 8313	
Restaurants and cafés	589	831 excl. 813	VAT at 20,6%
Hotels and accommodation services	183	832	VAT at 5,5%
- Package tours	13	84	VAT at 20,6%
- Financial services	63	85	Exempt
- Other household services	70	86	VAT at 20,6%
Sub-total rents and water	*682*		
- Rent of principal residence	576	3111	Exempt
- Water and urban heating	106	312	VAT at 5,5%

Source: Insee, BMS.
Note: [1] Only the "cinema" and "theaters and concerts" items are subject to the 5.5% rate.

Figure A1.3
Price of private sector services
(year-on-year change)

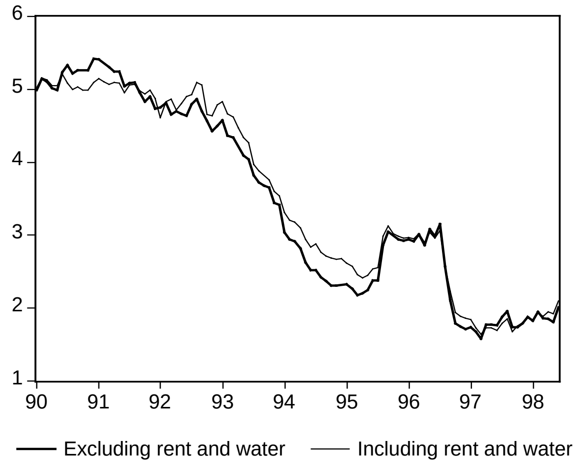
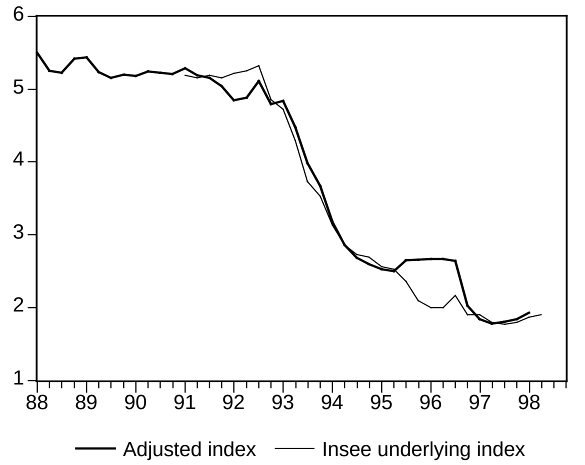


Figure A1.4
Adjusted and underlying indexes for private sector services (year-on-year change)



3. Index of food prices

The index of food prices represented 22.8% of the CPI in 1998. The high volatility (Figure A1.5) is attributable to two factors: tobacco and so-called sensitive products For sensitive products, we have included items excluded by Insee for the construction of its food underlying index. Thus, the index excludes fresh produce, which fluctuates by as much as 10% from month to month (equivalent to 0.3 of a point of the monthly growth rate for the index, see Figure A1.7), and other products like meat and coffee.

Table A1.3: Composition of the food price index

Item	Weight	BMS N°	Indirect taxation
Food and tobacco	**2220**		
Sub-total sensitive products	*1308*		
- Fresh produce	275		
Fresh fish	52	1131	VAT at 5,5%
Fresh shellfish	15	1133	VAT at 5,5%
Fresh vegetables	111	1161	VAT at 5,5%
Fresh fruit	97	1171	VAT at 5,5%
- Meat	532	112	VAT at 5,5%
- Milk, cheese, eggs	250	114	VAT at 5,5%
- Coffee, tea and cocoa	52	119	VAT at 5,5%
- Chocolate bars	29	11A2	VAT at 5,5%
- Chocolate-based confectionery	24	11A3	VAT at 20,6%
- Everyday wines	24	1301	VAT at 20,6%+ excise
- Superior wines	88	1302	VAT at 20,6%+excise
- Champagne and sparkling wines	34	1303	VAT at 20,6%+ excise
Sub-total other products	*712*		VAT at 5,5%
Sub-total tobacco	*200*	*14*	VAT at 20,6%+ excise

Source: Insee, BMS.

The exclusion of tobacco and sensitive products gives the adjusted food index. The pattern of the adjusted index and Insee underlying index are similar (Figure A1.8). The only differentials are due to changes of excise duties (affecting alcohol), which are not included in our calculation. This difference explains the discrepancy between the two indexes from 1993 to 1995. A detailed breakdown of the index is given in Table A1.3. Excise duties are levied on alcoholic beverages and tobacco products in addition to VAT.

Figure A1.5
Food prices excluding tobacco
(monthly growth rate)

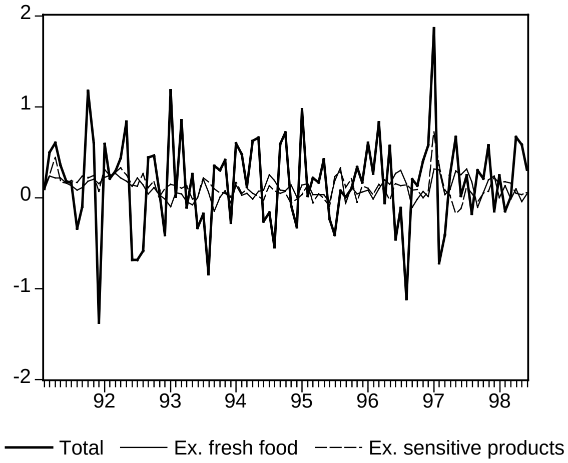


Figure A1.6
Tobacco price
(level)

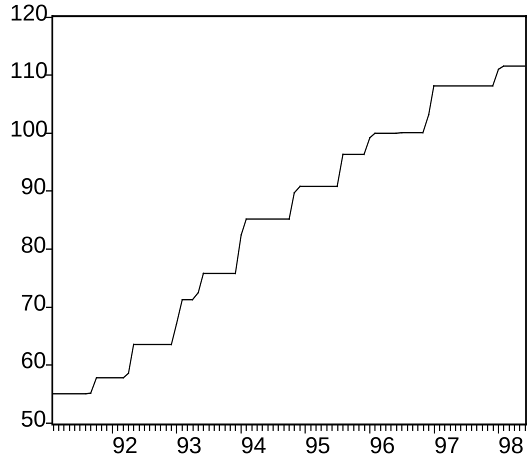


Figure A1.7
Volatile sub-indexes
(monthly growth rate)

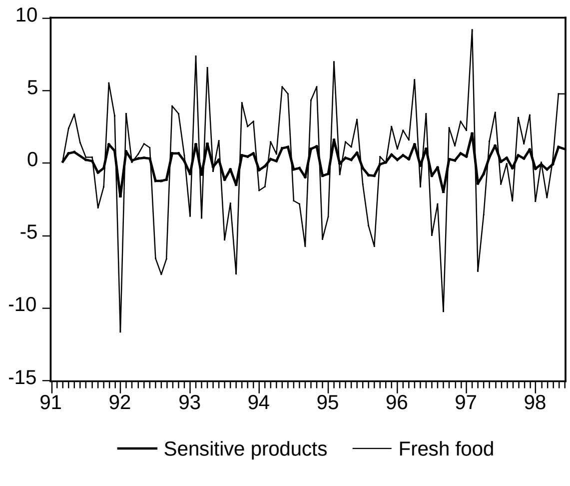
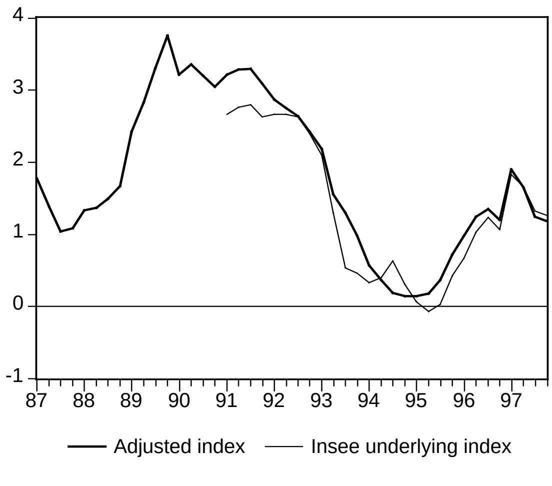


Figure A1.8
Adjusted and underlying indexes
(year-on-year change)



4. Index of energy prices

The energy price index represented 8.71% of the CPI in 1998. It includes two sub-indexes: oil products (5.19%) and other products (3.52%). Table A1.4 gives a detailed breakdown of the energy component of the CPI. Figures A1.9 and A1.10 show the prices of the oil products sub-index and the elementary electricity index. Two remarks may be made. First, not unexpectedly, the « oil products » component is the most volatile. Second, the « other products » sub-index consists principally of electricity, the price of which changes stepwise because tariff increases or reductions are generally decided once a year. The model of this item therefore concerns the first sub-index only.[22]

Table A1.4: Composition of the energy price index

Item	Weight	BMS N°	Indirect taxation
Energy	**871**		
Sub-total oil products	*519*		
- Petrol (super, unleaded super and diesel)	418	6222	VAT at 20,6% and TIPP
- Domestic fuel oil	70	323	VAT at 20,6% and TIPP
- LPG	31	3223	VAT at 20,6% and TIPP
Sub-total other products	*352*		
- Electricity	256	321	VAT at 20,6%
- Town gas	90	3222	VAT at 20,6%
- Coal and other solid fuels	6	324	VAT at 20,6%

Source: Insee, BMS.

Figure A1.9
Price of oil products (level)

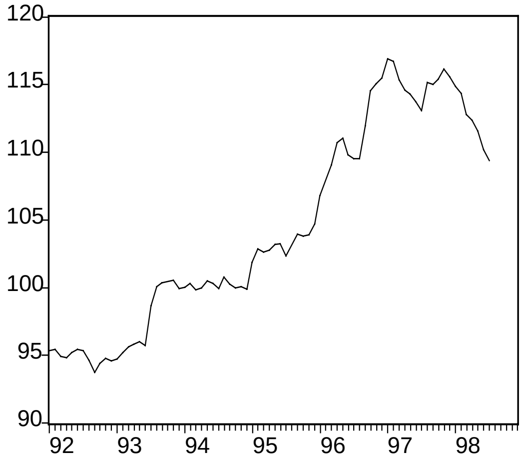
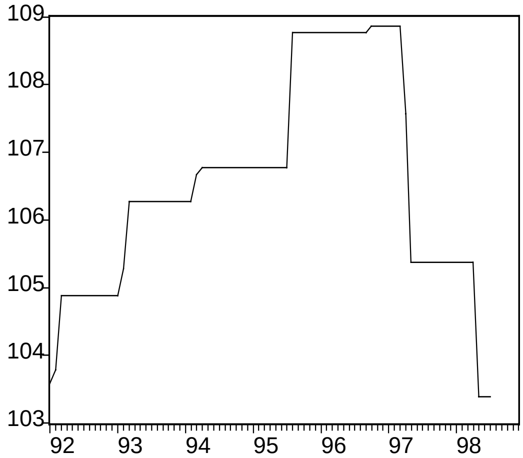


Figure A1.10
Electricity price (level)



5. Other prices

A certain number of items in the CPI in addition to sensitive food products may be regarded as exogenous. They include public tariffs excluding energy (Figures A1.11 to A1.14), tobacco (which is grouped together with food) and healthcare services, the trend of which is closely linked to agreements between the government and healthcare professionals. The variations in the different sub-indexes comprising the "healthcare services" item (Figures A1.15 to A1.18) provide ample evidence of this. The breakdown of the exogenous price index is given in Table A1.5. The total weighting of public services excluding energy is slightly greater than the weighting used in the "Informations rapides" publication because we cannot entirely isolate road tolls and the TV license. Private sector services (excluding rents and water) will clearly be reduced by the same amount.

Table A1.5: Composition of exogenous prices

[22] All the items in this sub-index are excluded from the price index "excluding public tariffs and volatile prices adjusted for tax measures" published monthly by Insee.

Item	Weight	BMS N°	Indirect taxation
Public services	**509**		
- Road tolls and parking	61	6231	VAT at 20,6%
- Rail transport	67	6301	VAT at 5,5% or 20,6%[1]
- Urban and suburban passenger transport	60	6302	VAT at 5,5%
- Communication	182	64	
Telecommunication services	155	6401	VAT at 20,6%
Postal services	27	6402	Exempt
- TV license and subscriptions	60	7201	2,1% or 5,5%[2]
- Meals in school or university canteens	79	8313	Exempt
Healthcare services	**550**	53	
- Medical services	288	5301	Exempt
- Laboratory analyses	50	5302	Exempt
- Medical auxiliaries	79	5303	Exempt
- Dentists	115	5304	Exempt
- Ambulance services	18	5305	Exempt

Source: Insee, BMS.

Notes: [1] Goods transport services are subject to VAT at 20.6%. [2] 2.1% for licenses and rights of use and 5.5% for subscriptions to television services.

Figure A1.11
Price of public services (level)

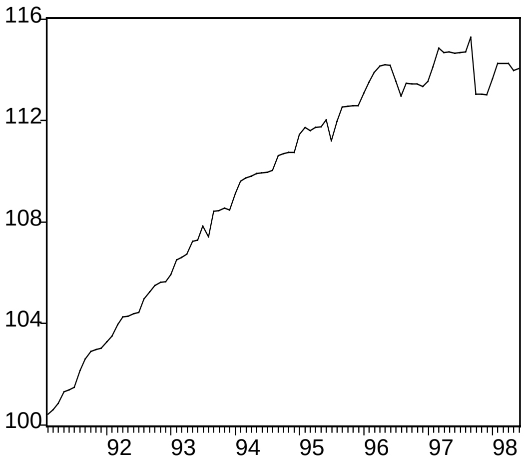


Figure A1.12
Price of passenger transport (level)

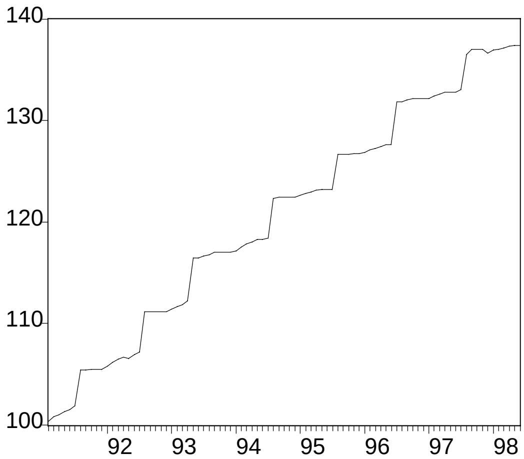


Figure A1.13
Price of communications (level)

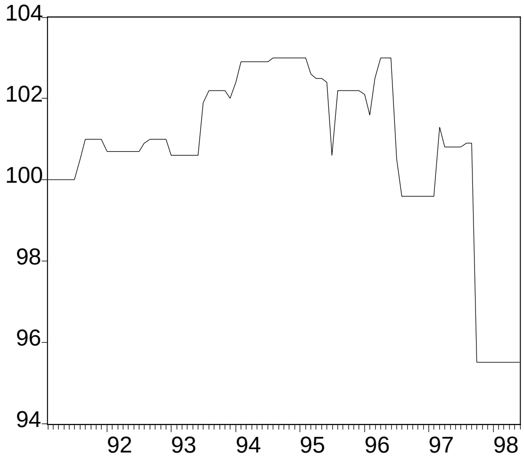


Figure A1.14
Price of TV license and subscriptions (level)

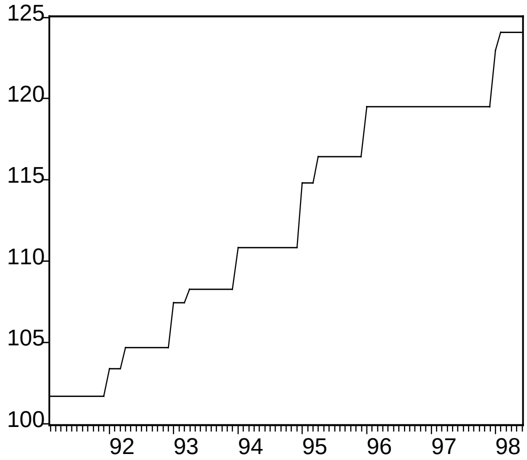


Figure A1.15
Price of healthcare services (level)

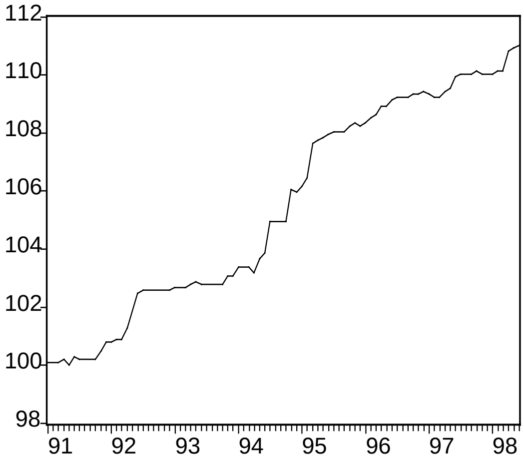


Figure A1.16
Price of medical services (level)

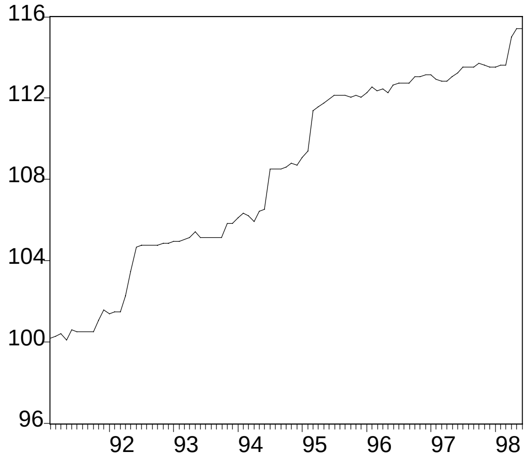


Figure A1.17
Price of laboratory analyses (level)

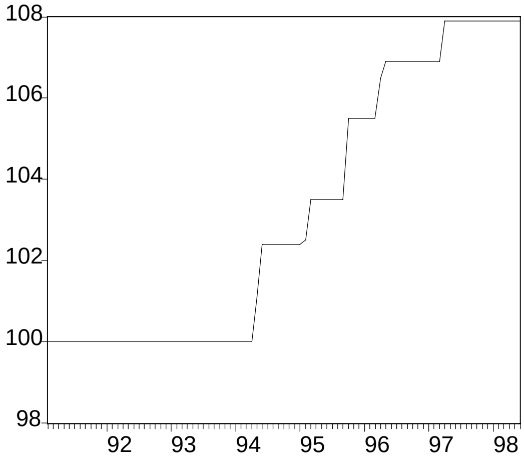
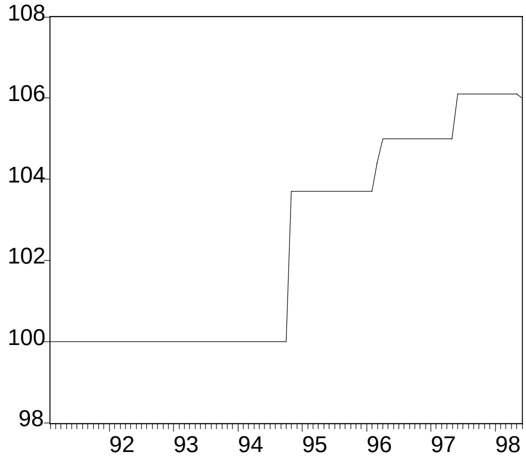


Figure A1.18
Price of dentists (level)



Appendix 2: Breakdown of NCPI by rate of VAT

In 1998, 61.6% of the NCPI was subject to VAT at the standard rate, 24.7% at the reduced rate and 4.9% at 2.1%. 8.8% of the items in the index were exempt.

Table 6: Breakdown of the NCPI by rate of VAT[23]

% of items

VAT rate (%)	0.0	2.1	5.5	20.6
Total	**8.8**	**4.9**	**24.7**	**61.6**
Food			85.0	15.0
Private sector manufactured goods	0.0	12.8	3.8	83.4
Private sector services	7.4	0.0	17.9	74.7
Energy				100.0
Public tariffs excluding energy	20.8	11.8	24.9	57.1
Healthcare services	100.0	0.0	0.0	0.0

Source : Banque de France computations.

[23] Percentages cannot be measured with absolute precision because the products or services that comprise certain elementary indexes are subject to VAT at different rates.

Appendix 3: Modeling oil products prices

The following four equations model ex-refinery prices of super, unleaded super, diesel and fuel oil. They are estimated monthly from January 1992 to December 1997.

For super:

$$\Delta\log(su) = \underset{(2.22)}{0.176}\Delta\log(su_{-1}) - \underset{(-1.94)}{0.138}\Delta\log(su_{-2}) - \underset{(-3.06)}{0.107}\left(\log(su_{-1}) - \log(brent_{-1})\right)$$
$$+\underset{(6.57)}{0.206}\Delta\log(brent) + \underset{(3.78)}{0.155}\Delta\log(brent_{-1}) + \underset{(3.12)}{0.047}\,dum978 + \underset{(3.51)}{0.052}\,dum948 + \underset{(2.68)}{0.016}$$
$$\overline{R}^2 = 0.709,\ see = 0.014,\ \text{LM(12)} = 0.665\ (0.78),\ \text{ARCH(12)} = 1.309\ (0.43),$$
$$\text{JB} = 7.094\ (0.03).$$

For unleaded super:

$$\Delta\log(su98) = \underset{(3.49)}{0.244}\Delta\log(su98_{-1}) - \underset{(-1.83)}{0.129}\Delta\log(su98_{-2}) - \underset{(-2.61)}{0.065}\left(\log(su98_{-1}) - \log(brent_{-1})\right)$$
$$+\underset{(6.23)}{0.176}\Delta\log(brent) + \underset{(4.66)}{0.162}\Delta\log(brent_{-1}) + \underset{(4.22)}{0.057}\,dum978 + \underset{(4.03)}{0.057}\,dum948 + \underset{(2.35)}{0.017}$$
$$\overline{R}^2 = 0.759,\ see = 0.013,\ \text{LM(12)} = 0.453\ (0.17),\ \text{ARCH(12)} = 0.897\ (0.56),$$
$$\text{JB} = 2.227\ (0.33).$$

For diesel:

$$\Delta\log(go) = \underset{(2.96)}{0.262}\Delta\log(go_{-1}) - \underset{(-2.25)}{0.095}\left(\log(go_{-1}) - \log(brent_{-1})\right)$$
$$+\underset{(7.19)}{0{,}304}\Delta\log(brent) + \underset{(2.85)}{0.145}\Delta\log(brent_{-1}) + \underset{(2.19)}{0.020}$$
$$\overline{R}^2 = 0.590,\ see = 0.020,\ \text{LM(12)} = 1.049\ (0.42),\ \text{ARCH(12)} = 0.813\ (0.66),$$
$$\text{JB} = 13.908\ (0.00).$$

For fuel oil:

$$\Delta\log(fi) = \underset{(3{,}08)}{0.263}\Delta\log(fi_{-1}) - \underset{(-2.52)}{0.087}\left(\log(fi_{-1}) - \log(brent_{-1})\right)$$
$$+\underset{(3.08)}{0.263}\Delta\log(brent) + \underset{(2.41)}{0.023}$$
$$\overline{R}^2 = 0.458,\ see = 0.022,\ \text{LM(12)} = 0.681\ (0.76),\ \text{ARCH(12)} = 1.023\ (0.44),$$
$$\text{JB} = 0.171\ (0.92).$$

Figures A3.1 to A3.4 show dynamic simulations of these equations beginning in January 1992.[24] They are of good quality and give an accurate forecast for the first half of 1998, which is outside the sample.

From these prices inclusive of tax, we recomposed the "petrol" sub-index, weighting each price for the weight of the component (super, unleaded super and diesel) in total petrol consumption. Then, with the fuel oil weighting, we recomposed an approximate index inclusive of tax for "oil products" (approximate because we did not model LPG). A simple regression of the "oil products" component of the CPI (*pp*) on our approximate index (*pttc*) gives a simulated index:

$$\log(pp) = \underset{(113.92)}{1.008}\log(pttc) - \underset{(30.85)}{1.693} + \hat{u}_t$$

$$\overline{R}^2 = 0.994,\ see = 0.005 \text{ and } DW = 0.458.$$

Figure A3.1
Super

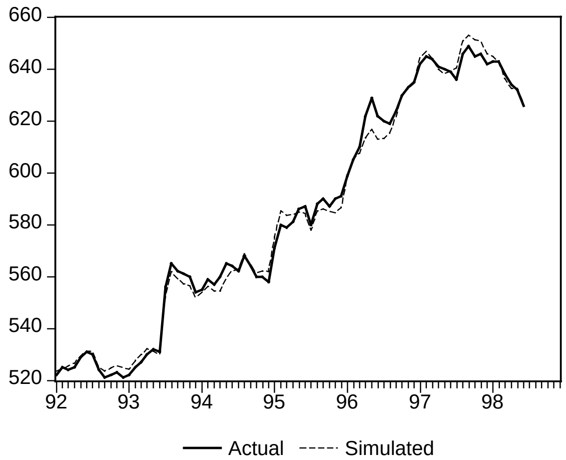


Figure A3.2
Unleaded super 98

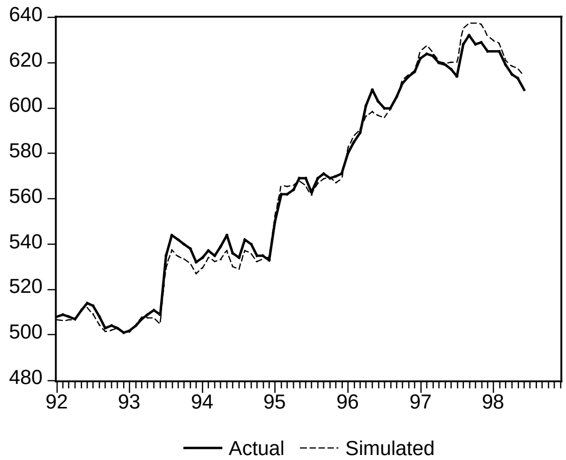


Figure A3.4
Diesel

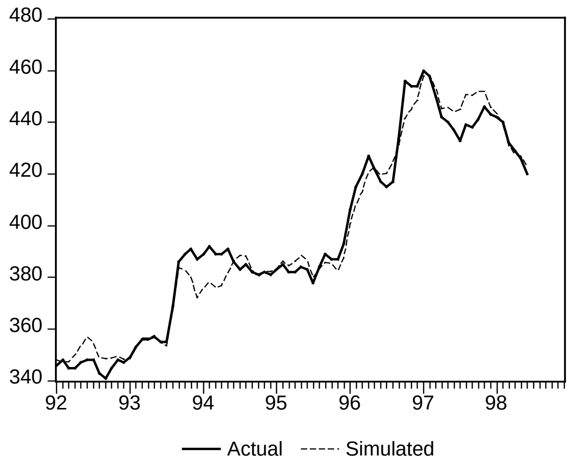
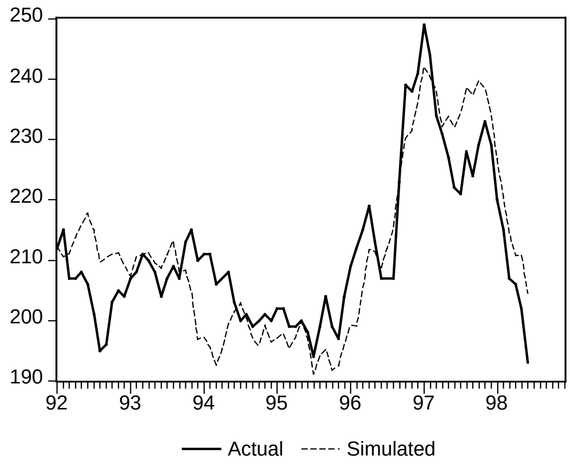


Figure A3.3
Domestic fuel oil



[24] The charts show pump prices, adding the various taxes to ex-refinery prices.

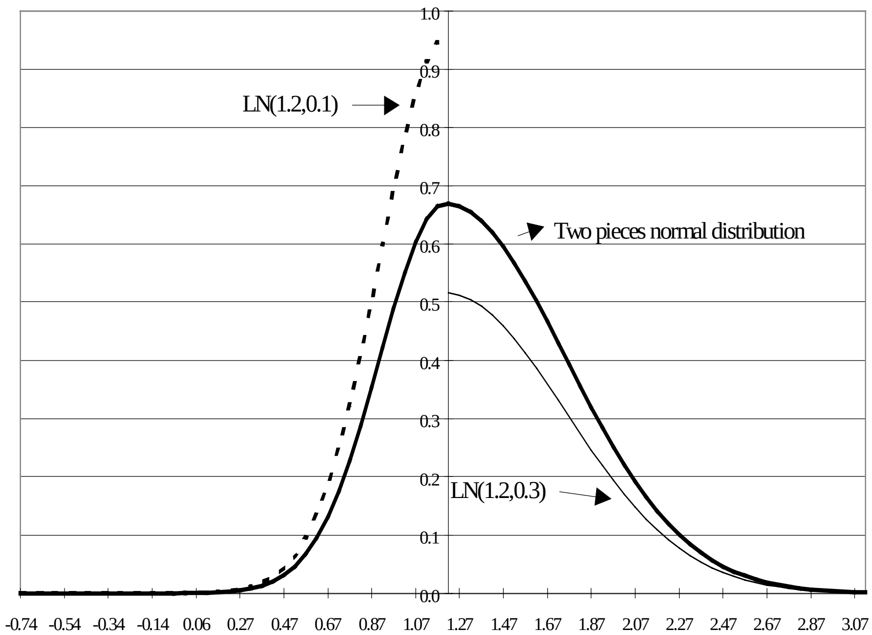
Appendix 4: The two-piece normal distribution

The density of a two-piece normal distribution is written as follows:

$$f(x) = \begin{cases} A\exp\left(-0.5\left(\dfrac{x-\mu}{\sigma_1}\right)^2\right) & x \le \mu \\ A\exp\left(-0.5\left(\dfrac{x-\mu}{\sigma_2}\right)^2\right) & x \ge \mu \end{cases}$$

The constant $A = \left(\sqrt{2\pi}\left(\dfrac{\sigma_1+\sigma_2}{2}\right)\right)^{-1}$ normalizes the density so that its integral is equal to the unity (Figure A4.1). This ensures that the function is continuously differentiable. The distribution is slightly asymmetrical in this example, since the mode is 1.2% whereas the mean is 1.3%.

Figure A4.1.
Density of the two-piece normal distribution



The distribution mode is μ. Where $\sigma_2 > \sigma_1$, the distribution is skewed to the right. In this case, the median is higher than the mode and the mean is higher than the median.

Appendix 5: Computation of the difference between the mode and the mean

One begins by estimating the upside and downside risks relating to the various macroeconomic indicators. The judgment of experts is supposed to contribute information on the sign and the probability of macroeconomic shocks. The impact of a shock on the inflation rate must then be measured in order to deduce the impact on mean inflation. This calibration is effected using the properties (multipliers and standard deviations) of macroeconomic models. To illustrate our case in the French example,[25] we estimated a VAR model with three variables: inflation, growth in the hourly wage rate and the capacity utilization (CU) rate using quarterly data with six lags over the period 1977-1997. The year-on-year changes of prices and wages are the endogenous variables. The standard deviations of the innovations are 0.38, 0.43 and 0.75 respectively. They give an order of magnitude for the uncertainty regarding future shocks. The VAR estimate provides a response function for the shocks. Here we orthogonalize the residuals by ordering the VAR as follows: first the CU rate, then wages and prices. The CU rate is contemporaneously affected only by activity shocks. At four quarters, the impact on inflation of orthogonalized shocks of one standard deviation on the variables of CU rate, wages and prices is respectively: 0.20, 0.38 and 0.34.[26] These multipliers give the expected impact on inflation of shocks of one standard deviation. Lastly, in the Bank of England's approach, each of these multipliers is adjusted according to the subjective assessment of uncertainty regarding the amplitude of the shocks, from opinion of experts.

Table 2 illustrates the above procedure applied to the French case. Columns 1 and 2 contain the quantified probability of shocks on the three VAR variables. For the CU rate, there is a 25% upside risk and a 75% downside risk. We also assume that there is a 60% downside risk and a 40% upside risk for both wages and the intrinsic price shock. Column 3 shows the difference between columns 1 and 2, i.e. -50% for the CU rate and -20% for the two other variables. We then postulate that this net value represents the probability of a negative shock occurring in this specific case (the shock would be positive if the upside risk were greater than the downside risk). In Table 2, the risks described are more downside than upside. This is because we introduced an asymmetry to the left rather than to the right in order to approximate the context of the inflation forecast of April 1999. There are several reasons for believing that in the present case there are more downside risks: the impact of the crisis in the emerging countries on activity, price deregulation, etc. Column 4 simply multiplies the impact on inflation of a shock of one standard deviation in the variable over four quarters. As for column 5, it represents the uncertainty regarding the magnitude of a shock on CU rates, wages or prices. The usual magnitude is one standard deviation, but it may be slightly higher. For example, if the uncertainty regarding changes in the CU rate is high, the variability of the variable is 10% higher than that recorded in the past (1.1). All this information is then used to establish the mean inflation forecast (as a deviation to the modal scenario), which is the sum of the figures listed in column 6. This sum is obtained as follows. For each variable, we multiply the figure in column 3 with that in column 4. The result is then multiplied by the value in column 5. For example, the expected shock on the CU rate will have an impact of -0.5*0.20*1.1=0.11. The total impact of all these potential shocks, i.e. the mean-mode deviation, is -0.27 inflation points.

[25] The Bank of England uses various models (Bank of England, 1999, Fisher and Whitley, 1998). Different models may be used from one forecasting exercise to the other. These models include traditional econometric models, a VAR model, a reduced form of *output gap*/Phillips curve, etc.

[26] When residuals are normal, one standard deviation covers 68% of possible shocks.

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.

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