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EUROSYSTEM INFLATION PERSISTENCE NETWORK

BREAK IN THE MEAN AND PERSISTENCE OF INFLATION

A SECTORAL ANALYSIS OF FRENCH CPI

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The Eurosystem Inflation Persistence Network

This paper reflects research conducted within the Inflation Persistence Network (IPN), a team of Eurosystem economists undertaking joint research on inflation persistence in the euro area and in its member countries. The research of the IPN combines theoretical and empirical analyses using three data sources: individual consumer and producer prices; surveys on firms' price-setting practices; aggregated sectoral, national and area-wide price indices. Patterns, causes and policy implications of inflation persistence are addressed.

The IPN is chaired by Ignazio Angeloni; Stephen Cecchetti (Brandeis University), Jordi Galí (CREI, Universitat Pompeu Fabra) and Andrew Levin (Board of Governors of the Federal Reserve System) act as external consultants and Michael Ehrmann as Secretary.

The refereeing process is co-ordinated by a team composed of Vítor Gaspar (Chairman), Stephen Cecchetti, Silvia Fabiani, Jordi Galí, Andrew Levin, and Philip Vermeulen. The paper is released in order to make the results of IPN research generally available, in preliminary form, to encourage comments and suggestions prior to final publication. The views expressed in the paper are the author's own and do not necessarily reflect those of the Eurosystem.

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Abstract

This paper uses disaggregated CPI time series to show that a break in the mean of French inflation occurred in the mid-eighties and that the 1983 monetary policy shift mostly accounted for it. CPI average yearly growth declined from nearly 11% before the break date (May 1985) to 2.1% after. No other break in the 1973-2004 sample period can be found. Controlling for this mean break, both aggregate and sectoral inflation persistence are stable and low, with the unit root lying far in the tail of the persistence estimates. However, persistence differs dramatically across sectors. Finally, the duration between two price changes (at the firm level) appears positively related with inflation persistence (at the aggregate level).

Keywords: multiple breaks test, inflation persistence, monetary policy, sectoral prices.

JEL classification: E31, C12, C22.

Non-technical summary

The aim of this paper is to characterize inflation dynamics in France since the seventies. It addresses two questions in particular: has there been any structural change in long-term inflation average (and if yes for which reason), and is inflation a persistent process, meaning that it would slowly come back to its baseline after an external shock?

The paper shows that a single structural change occurred, in the mid-eighties. The break is detected in a very large brand of goods and services, including traded and non-traded items together with energy linked and non-directly linked items, which suggests that the break was driven by domestic factors. Therefore, the monetary policy regime change of March 1983, but also the price and wages freeze in 1982 and the wage bargaining policy from 1983 onwards are good candidates for causing this change.

Turning to the persistence estimates, the paper shows that French inflation is not a highly persistent process, once it is accounted for the structural break in the mean. Measures of persistence at the sectoral level show an aggregation effect, as expected in the theory: the persistence of the aggregate (inflation) is above the average persistence of its sectoral components. In addition, a more puzzling result is found, when these sectoral persistence measures are compared with the average spells between two price changes at the firm level: the longer the duration between two price changes, the more persistent inflation appears to be. If this result were to be confirmed, it would question several usual price-setting models.

On the methodological side, the persistence measures proposed in this paper are rather standard in the literature, whereas the methodology for testing multiple breaks test is more original. It relies on a recent test for which a specific parameterization is proposed in order to limit as much as possible the pitfalls this class of test is generally subject to.

1 Introduction

The dynamics of inflation potentially summarize the key features of an economy. For this reason, an empirical literature has recently paid attention to inflation persistence - defined as the speed with which inflation converges to equilibrium (or baseline) after a shock[1]. The faster this return to equilibrium, the less persistent inflation is. In this framework, inflation dynamics can be characterized with a two-step approach, with the definition of the baseline as the first step and the persistence measure with respect to the baseline as the second.

However, there are several ways to define baseline inflation. Robalo Marques (2004) discusses several models of equilibrium inflation: from filter-based trend components of inflation to discrete changes in the mean. The first group is the most general, since it does not presume any particular pattern for the baseline. Mean inflation is simply considered as a time-varying process. By contrast, a discrete change in the mean can be seen as a restriction of this model. It assumes a stable long run equilibrium level of inflation, which however is allowed to change over time for instance in association with some durable change in monetary policy. The inflation persistence literature[2] has mainly focused on the restricted version of the baseline, i.e. where discrete changes occur in equilibrium inflation.

These studies deduce different break dates in the mean of French inflation. Covering a wide brand of inflation measures on a sample beginning in 1984, Levin and Piger (2004) find no break, except in the GDP deflator inflation in 1993. On similar data, Gadzinski and Orlandi (2004) detect a break in several inflation measures in 1992 or 1993. With a different test and a larger sample, Corvoisier and Mojon (2004) find two breaks, in 1973 and 1985, whereas Benati (2003) finds evidence of possible breaks in 1973 and 1983 but also, depending on the test used, in 1991. Overall, three possible periods of structural change in French inflation emerge: the early seventies, the mid-eighties, and the early nineties. In addition, this literature has shown that it was necessary to take account of the structural break to correctly gauge persistence (Levin and Piger, 2004). Overall, Gadzinski and Orlandi (2004) and Levin and Piger (2004) find a rather moderate degree of persistence in France, as in most developed economies.

However, these studies leave the question of the origins of these shocks unanswered. Since assumptions are made on inflation dynamics, an identification of the structural shock

[1] See for instance Andrews and Chen (1994), Willis (2003) and Robalo Marques (2004).

[2] For instance: Benati (2003), Corvoisier and Mojon (2004), Gadzinski and Orlandi (2004) or Levin and Piger (2004).

is required and other could object either that the break is an artifact or that a more general time-varying model of the mean could be a more appropriate description of the data (Robalo Marques, 2004). In the French case in particular, analysis of the causes of the breaks are still preliminary. This contrast with the discussion of the US break[3].

In addition, the differences between the estimated breaks in the various studies remain striking. It seems that the detected break dates can differ for several reasons - by the sample period, the statistical test implemented, or the inflation measure (CPI or GDP deflator) for instance. No estimation of the relevant importance of these different factors is available so far.

The present paper addresses these two issues.

First, it uses highly disaggregated inflation time series (141 items) over a long period (1972 - 2004). As suggested by Clark (2003) and Cecchetti and Debelle (2004), the use of sectoral prices can strengthen the diagnosis of overall inflation[4]. In particular, we can exploit the sectoral results to reveal the forces driving the structural changes, if any. For instance, if the driving force is external, the structural break should be first detected in the traded goods and services. By contrast, if a monetary policy regime change or some other macroeconomic domestic change is the key factor, the structural break should be homogeneously observed throughout the basket. Moreover, the combined use of a large sample and disaggregated series can help single out general and sectoral shocks among the shocks proposed in the previous studies.

Second, given that at least one of the possible breaks in the mean of French inflation seems to depend on the testing procedure, I investigate the robustness of the results with respect to the parameterization of the Altissimo and Corradi (2003) test, the state of the art of multiple breaks test. In particular, I check its vulnerability to three potential pitfalls: a bias implied by high persistence, a sensitivity to the position of the date break within the sample, and a sensitivity to heteroscedasticity.

[3] An extensive literature investigates the causes of some structural changes (in volatility or average inflation, see Sensier and van Dijk, 2004) that took place in the eighties in the US. To summarize the latter debate as proposed by Ahmed, Levin and Wilson (2004), the structural change can be attributed to "good policies" (essentially monetary policy, as suggested by Clarida, Gali and Gertler, 2000), "good practices" (in the inventories management, McConnell et al., 1999) or "good luck" (as resulting from an external shock, for instance Stock and Watson, 2003). At least in the case of France, the literature on the structural change in inflation has not so far been superseded in this debate.

[4] Lünnemann and Mathä (2004) achieve an important step in the use of highly disaggregated data and find additional evidence of moderate persistence. However the studied sample does not include any of the breaks identified in the literature.

There are four main conclusions.

First, the structural break date estimates among French CPI items are strikingly homogeneous: a single break is detected in aggregate CPI and in 82% of its 141 components, at a date close to the overall CPI break date (May 1985). This confirms Corvoisier and Mojon (2004) who find a break in aggregate CPI inflation in the middle of the eighties (second quarter of 1985) and do not detect a break in the nineties. Average CPI monthly annualized growth declined from 10.9% before May 1985 to 2.1% after that date (see Figure 1). I show that domestic factors - mainly monetary policy - account for this break, as suggested by some previous literature (for instance Trichet, 1992, Blanchard and Muet, 1993, and Bilke, 2004). In addition, I find consistent evidence for a early-nineties break in the services component of the CPI. The results for the overall CPI and for the industrial goods are sensitive to the choice of a sample length: when the sample is shortened to the one used by Levin and Piger (2004) or Gadzinski and Orlandi (2004), i.e. 1984-2003, a second break in mean of these series can be detected in the beginning of the nineties.

Second, using both aggregate and sectoral data, I find that French inflation has fairly moderate persistence once one accounts for the structural break in mean. The null hypothesis of a unit root can be decisively rejected, and the estimated degree of persistence is broadly similar to the results obtained by Levin and Piger (2004) and Gadzinski and Orlandi (2004). Furthermore, as in Cecchetti and Debelle (2004), I find that inflation is well-characterized by a break in mean but not a break in persistence.

Third, the sectoral estimates add some nuances to the Cecchetti and Debelle (2004) observation that the duration between two price changes and inflation persistence are negatively correlated, as expected by the time-dependent price models. To the contrary, in the present dataset it seems that the sectors with a longer duration between two price changes are also the more persistent in price change.

Fourth, on the methodological side, the choice of a proper bandwidth for the computation of robust variance allows to limit the three above mentioned pitfalls for a given sample size. In particular, the bias implied by a high persistence degree can fall to an acceptable level, when compared with the test nominal size.

The remainder of the paper is organized as follows. Section 2 describes the data and briefly introduces the statistical tools used to test for breaks in mean and to measure persistence. Section 3 presents the structural breaks estimates and discusses their determinants.

Section 4 proposes estimates of inflation persistence for France. Section 5 provides a study of the multiple breaks test properties, before Section 5 concludes.

2 Methodology

In this Section, I describe the dataset and briefly exposes the econometric methods.

2.1 Data

The database is an original one and results from Baudry and Tarrieu (2003) retropolation of the base year 1990 CPI and its components on the 1980 base year CPI. The aggregate backdated CPI sample is 1973:1 - 2004:1. The sample of 141 components is 1972:2 - 2004:1 and the sample of 20 other items begins between 1987:2 and 2001:2.

In addition to overall CPI, several sectoral aggregates have been built, following the usual HICP categories: non-processed food (A), processed food (B), non-energy industrial goods (C), energy (D) and services (E).The sample of these Laspeyres chained index aggregates begin in 1973:1. Table 1 proposes additional descriptive statistics for these. All the time series, both at item and aggregate levels, are not seasonally adjusted; the indexes are Laspeyres chained and I use one-month growth rates.

A methodological break in some of the industrial goods price time series may be of particular importance: the French statistical office, INSEE (*Institut National de la Statistique et des Etudes Economiques*), had started accounting for the sales prices in 1992, and then included them progressively until 1998, when they were completely incorporated. The sales mainly concern manufactured goods, more particularly clothing, and the effect of their progressive introduction has not been corrected in the database. I considered that sale prices have to be included since they are the main (if not the only) downward adjustment episode for most of prices. Thus keeping this information and facing a methodological break has been preferred to not using it with seasonally adjusted data. Some previous studies have found evidence that the introduction of sales could have caused a break in average price growth (for instance Cecchetti and Debelle, 2004). It is hence necessary to pay a special attention to this issue.

The break test procedure is performed both at item and aggregate levels, whereas persistence estimates are proposed only for the six aggregates. In order to gauge the robustness of the results to the sample, the break test is also performed over the smaller sample (1984:1 -

2003:3) used by Gadzinski and Orlandi (2004) and Levin and Piger (2004)[5]. The sample size then represents 231 observations with monthly data, instead of 373. Even if this reduction of the sample size may consequently alter the efficiency of the test as shown later on, this is the only way to investigate where the possible differences in the results come from.

2.2 Econometric methods

The overall degree of inflation persistence is evaluated in two steps: testing for the presence of structural breaks, then measuring persistence with respects the possible changes in the baseline. This subsection describes the tools utilized at each step.

2.2.1 Testing for several structural breaks

Given the rather large studied sample, it is certainly necessary to test for the presence of more than one structural change. I have chosen to utilize Altissimo and Corradi (2003) multiple breaks test procedure since it corrects the critical value for the sample size[6]. This test has been performed with an unusual parameter used in the computation of residuals variance, in order to limit as much as possible its sensitivity to a high degree of autocorelation, as will be exposed in detail in Section 5. The minimum allowed interval between two break dates is set to 5% of the number of observations (about 20 months with the large sample). Similarly, no break is allowed to occur in the first 20 months and the last 20 months of the sample.

The dating precision of the procedure has to be known in order to investigate the possible shocks causing the breaks. From simulation exercises presented in Section 5, it results that, in the large sample, a range of plus or minus 20 months around the detected date corresponds to a 96% confidence interval, for a break of a moderate size not to close from the sample beginning or end. It is worth noting that this range is clearly a maximum and it can be substantially reduced when the process is weakly or not at all persistent, which happens to be the case for a large share of the basket.

Last, a careful study of the small sample properties is conducted in Section 5 and its outcome can be quickly summarized in the following way. It is shown that the vulnerability of the procedure to the persistence effect and heteroscedasticity can be limited to a moderate underestimation of the number of breaks for highly persistent processes, which can be raised when a simultaneous change in volatility takes place. A reduction of the sample size in the

[5] A few differences remain however between the two databases, in particular the frequency: we use monthly data whereas Gadzinski and Orlandi (2004) and Levin and Piger (2004) use quarterly time series.

[6] Within the inflation persistence literature, Benati (2003) and Corvoisier and Mojon (2004) also use this procedure.

context of heteroscedasticity and persistence substantially lowers the test efficiency. Last, the test generally overestimates, in a very minor way (below the nominal size of the test, i.e. 0.05), the number of breaks for non persistent series.

2.2.2 Measuring inflation persistence

Following Andrews and Chen (1994) and much of the subsequent literature on inflation persistence (for instance Levin and Piger, 2004 or Clark, 2003), I utilize the sum of autoregressive coefficients as a scalar measure of persistence. This measure requires fitting inflation with an AR(p) model[7], with a lag p implied by Schwarz (1978) criteria. It is well known that OLS estimates of AR parameters suffer from a downward bias when the root is close to unity (Marriott and Pope, 1954), so the sum of the autoregressive coefficients is evaluated using an approximately median unbiased estimator, as proposed by Andrews and Chen (1994)[8].

When a break in the mean has been previously detected, the AR estimates includes two "constants" over the two periods. Each estimated persistence parameter is reported with its 90% confidence band computed as proposed by Andrews and Chen (1994)[9].

3 Evidence on structural breaks

This Section presents and comments the structural changes in French inflation. The impact of several limited shocks like the euro cash changeover and the case of an early nineties break are also discussed.

3.1 Estimated number of breaks and break dates

A single break is detected in overall CPI, in May 1985. Among the 141 items, the test detects a single break in more than 80% of the cases. The occurrence of zero and two breaks is similar (8%), whereas only three items record three breaks.

Table 2 reports this distribution of the estimated number of breaks at the item level, classified by sector. Very few sectoral differences emerge. The service sector reaches the highest number of breaks, with more than 20% of the items recording two breaks or more. On the contrary, a large share of energy items (30%) exhibits zero break (this category can

[7] A basic analysis of the correlogram for inflation time series justifies the AR modelling choice, as autocorrelations decrease very slowly while partial autocorrelations seem to cut off.

[8] Another method for computing the unbiased sum of the autoregressive coefficients - the grid bootstrap - is proposed by Hansen (1999) and is based on a comparable simulation process. See Andrews and Chen (1994) and Robalo Marques (2004) for a discussion of several other measures of persistence.

[9] The simulations are programmed under Gauss, with 1 000 replications at each step of the iteration process.

nonetheless not be fully compared with the others given its small number of items) and a single break is detected in nearly 90% of industrial goods. In the food sectors, a single break is detected in around 80% of the items, whereas the occurrence of two or more breaks is rare. At the sectoral aggregate level (Table 3), the picture is the same: a single break is detected, except in the case of the services aggregate for which two breaks are identified.

Item level break dates appear to be impressively concentrated around the overall CPI single break date (Figure 2): 89% of the items record a break within the three years before or after and, again, there is no significant difference across the sectors. Average price growth falls in a similar way in every sector, except services: from more than 10% annualized to less than 3% (Table 4). The service sector price pattern is a little different, as its price growth remains more vigorous than the other after the mid-eighties break (at nearly 5%) and it falls below 2% after a second break in the early nineties. However as Figure 3 illustrates, the early-nineties are far from being a period of general structural change, since the observed breaks are limited to the service sector. Given the position in the sample of the breaks and their size, we can deduct the dating precision of the estimated breaks: if the process is strongly persistent, the probability of a detected break date to be the exact true break date is 78%, the overall probability that it is within the 3 months before or after is 86% and the probability that it is within the 19 months before or after is 94%, whereas if the process is not persistent at all, the probability for a detected break date to be within the three months before or after is 100%.

The introduction of an unusual bandwidth has no particular influence on overall results at this stage: at the aggregate level, the results are unchanged with the usual parameter. This means that the processes could be only weakly autocorrelated (thus not vulnerable to the persistence bias) and/or that the size of the break is large enough to be properly detected in every case.

3.2 Interpretation of the structural change

The main finding is that there occurred a single period of general change in the price growth in the mid-eighties. France is not the only country where structural changes occurred at that time and there is, for instance, an extensive literature on this issue concerning the United States. Some people argue that this change in the US was mainly driven by external factors (Stock and Watson, 2003), whereas the domestic factors, such as monetary policy, were the determining factors for the others (for instance Clarida, Gali and Gertler, 2000, but also Ahmed, Levin and Wilson, 2004, for the fall in average inflation). In the case of

France, a monetary policy rule can hardly be computed to test for a structural change in the conduct of monetary policy. Indeed, the interest rates were not a representative instrument for monetary before 1984 (Bilke, 2004). For this reason, the use of highly disaggregated time series can give some useful insight on the causes of a structural break in inflation. In particular, I propose to investigate the respective roles played by some external factors like the exchange rates, the oil prices or the degree of openness of the economy, and some other domestic and policy related factors. In the following, I show that disaggregated price sectoral break dates gives support for the domestic factors.

A change in the overall exchange rate regime is sometimes considered as the driving factor in the eighties structural change in the US. However in France, an exchange rate structural shock could be partly linked with a monetary policy change of which the external face was the "franc fort" policy. But let us consider the case of a purely externally driven exchange rate regime change. In that situation, the prices of internationally traded goods should be affected first, before spillover to the non-traded goods and services sectors takes place. Evidence from sectoral aggregate and item level break dates in the eighties does not support this view. The services (traditionally less internationally traded than the goods) experienced a break in the eighties before the other four sectors, in September 1983 (Table 6). At the item level, many non traded service items experienced a break early in the eighties, for instance (see Table 9): the medical, dental and paramedical services (overall 4.5% of CPI, in April or June 1983), the cultural services (1983), the services for the maintenance of the dwellings (April 1984), water supply (May 1982), the education services (November 1983), the restaurants (April 1983) or hairdressing (October 1983). Overall, no systematic difference between the traded and non-traded items can be found.

The 1986 counter oil price shock is another external candidate explanation. The main oil price decrease occurred during the first quarter of 1986. Given the break dating precision, all the breaks before June 1984 would be out of the 96% confidence interval around the oil shock[10]. Except for the service aggregate, the sectoral break dates in the eighties are posterior, thus they could be related with the oil price shock. However, item level break dates offer some crucial additional information: the break dates of 45% of the individual items for which a break is detected around the mid-eighties are between January 1982 and May 1984, outside the 1986 oil shock confidence band. In addition, among the pre-May 1984 breaking item prices, we even find some energy goods (electricity and solid fuels),

[10]For highly correlated processes only, otherwise the interval around the oil price decrease would be far smaller as shown in Section 5.

whereas some other energy goods do not experience any break in the mid-eighties (liquefied hydrocarbons, liquefied fuels or lubricants). Given these observations, I can clearly reject the hypothesis that the mid-eighties shock is mainly caused by the 1986 oil price drop.

The two previous arguments can also be used to reject the explanation of a greater exposure to external competition, linked for instance with the European market integration. First because, as previously stated, there is no significant difference between traded and non traded items. And second, because the growing exposure to external competition mainly occurred at the very end of the eighties, outside the confidence intervals around the mid eighties detected break dates.

As a consequence, we should likely look for explanations among the domestic factors. Among them, several policy changes are serious candidates for justifying a structural break: the prices and wages freeze in 1982, the collective wage agreement policy starting in 1983 and the monetary policy tightening together with the beginning of the "franc fort" policy in March 1983 (see for instance Bilke, 2004, for a description of these three shocks). The 1982 freeze episode had certainly a cooling down effect on inflation which then fell from its two digit levels, whereas the following collective wage agreement policy could have heavily weighed on the expectations (Blanchard and Sevestre, 1989). Overall, Blanchard and Muet (1993) consider that the combination of these policies had a major effect on the disinflation process and Trichet (1992) has shown how the competitiveness through the disinflation goal has been efficient since then. Given the confidence intervals of our break test procedure and the usual transmission lag, the findings on the break of the mid-eighties fully support the policy-oriented explanation. The average inflation of the post 1985 period (2.1%) is also consistent with the implicit or explicit reference that would have governed monetary policy during that regime.

3.3 Transitory shocks

Our test results can also be useful to gauge the impact of some specific shocks on inflation.

3.3.1 Euro cash changeover

So close to the end of the sample, the expected power of the test should be low, especially for highly persistent processes. However, it is worth noting that within the 141 item price series and 6 aggregates series, no break was detected in the two years before or after the changeover (January 2002). Thus, given the large number of considered time series, I would suggest that the changeover did not have a structural impact on inflation. This finding is, of course, compatible with a possible durable effect of the changeover on the price level of the level price.

3.3.2 Value added tax rate changes

The mid-eighties break does not coincide with a value added tax (VAT) rate decrease like the one in the end-eighties/early-nineties (see average VAT rate proposed by Bilke, 2004). Regarding the August 1995 VAT rate rise, a single break in mean occurred in October 1995 and concerned the other services for the maintenance of dwellings. Besides, no change occurred around the April 2000 main rate decrease. The same observation can be formulated regarding the 1977 main rate change: nearly no break occurred nearby. Thus, the two last general value added tax (VAT) rate changes did not have a permanent effect on inflation.

Regarding the specific VAT rate changes, the picture is a little more balanced. For instance, the July 1982 decrease in the VAT rate for food and publishing did not entail any break for the former but could have caused a break for the latter (a break has been detected for the newspapers in April 1982). However around the decrease in the rate of the new cars and non-alcoholic beverages in September 1987, no break is recorded in the relevant items. Similarly no break happened after the January 1990 fall in the rate of drugs and publishing goods.

Overall, the general VAT rate changes do not seem to have an impact on average inflation, whereas, in some circumstances, specific rate changes may have had one.

3.3.3 Regulated prices

An additional aggregate has been built in order to identify the regulated and formerly regulated sectors[11]. As shown by Bilke (2004), these prices have offered a significant contribution to the curbing of inflation (0.2 percentage point of yearly price growth from 1985 to 2003). However, their structural features remain broadly similar to the rest of the economy since a single break has been detected, in May 1984.

3.4 The case of the early-nineties break

The same test has been performed over a smaller sample, similar to the one used by Levin and Piger (2004) and Gadzinski and Orlandi (2004), in order to gauge the relevant effect

[11] This aggregate includes the former monopolies (airways and telecommunication) and industries for which the public authorities set prices (electricity, gas and taxis) and sectors which combines both situations (railway, combined transport, postal and TV and radio fees). See Bilke (2004) for further discussion on this sector.

of the sample length and the break test. Both work rely on another approach to test for structural change than the one followed in this paper, in Corvoisier and Mojon (2004) or in Benati (2003) for instance. While the latter rely on the use of a multiple breaks test procedure, the former estimate an autoregressive process, and then gauge the possibility of a break. Gadzinski and Orlandi (2004) find a break in most inflation series in 1992-1993 and Levin and Piger (2004) find a break at the same time in one out of four inflation measures (GDP deflator).

Multiple breaks test procedures are often judged to have low power. If a special attention is paid here to minimize this risk (Section 5 is largely devoted to this issue), the previously reported estimates are however computed again over the reduced sample with Altissimo and Corradi (2003) procedure, in order to account for the effect of the reduction of the sample length (for this test). The results over the reduced sample are shown in Table 6 and can be compared with the results reported in Table 3. First, it is worth noting that despite its proximity with the beginning of the sample, the mid-eighties break is still detected for CPI, industrial goods and services, while it was not in Gadzinski and Orlandi (2004) or Levin and Piger (2004). Second, the early nineties break is then also detected, in the same three time series, while it was not in CPI and industrial goods over a larger sample (Table 3). A possible explanation could be the following: the sample length reduction increases the test sensitiveness to a volatility change, and a change in the volatility of the series very likely happened in the early nineties -with the introduction of sales prices. Section 5 will show in particular that for the most persistent processes (as will appear to be the 3 relevant aggregates), the reduction of the sample size can make the test spuriously detecting a mean break when there is a volatility change. This would be consistent with the evidence pointed by Cecchetti and Debelle (2004) that the introduction of sales prices can coincide with the detection of a break in the mean. Overall, while its aim is not to compare the relative power of parametric and non-parametric break test, the present exercise shows that the differences between the present results and Levin and Piger (2004) or Gadzinski and Orlandi (2004) can be attributed to the break test and the sample length in the case of the mid-eighties break, and more likely to the sample in the case of the nineties break.

4 Evidence on persistence

This Section outlines estimates of inflation persistence in France after taking account of structural changes and also discuss the implication of some sectoral differences.

4.1 Overall CPI persistence

To gauge the effect of introducing structural changes, I first estimate inflation persistence by deliberately using the incorrect hypothesis of a stable mean. Table 7 reports these estimates. Inflation persistence appears to be strong and the hypothesis of a unit root can hardly be rejected for overall CPI, industrial goods and services. These estimates are larger than Levin and Piger (2004) estimate on overall CPI (0.77).

I then add the structural changes to the persistence evaluation. As expected from Perron (1990), and also observed by Levin and Piger (2004) for several other countries, inflation persistence dramatically decreases in every case (Table 8). The most significant decreases occur in the two prices for food aggregates. The median unbiased estimate of overall CPI persistence falls from 1.01 to 0.76 after accounting for the structural change. Overall inflation persistence is similar to Levin and Piger (2004) one[12]. However, it remains above Gadzinski and Orlandi (2004) estimates which rely on detecting a structural break in the early-nineties[13]. Impulse response functions offer another view of these persistence measures (Figure 4), and reveal that half-lives are below two months in every case, except for price of industrial goods.

Taylor (1998) suggests that a structural change in monetary policy can cause a change in the persistence of inflation itself. I investigate the possibility of a change in the persistence parameter by computing rolling regressions over 10-year samples. From a methodological point of view, I follow Pivetta and Reis (2003) and O'Reilly and Whelan (2004). Both conclude there is stability in the persistence parameter, respectively in the US and the euro area. Figure 5 reports the estimated persistence parameter for the French CPI, with its 90% confidence interval band. The latter appears to provide important information, since the bands have enlarged over the past few years. The persistence of overall inflation could have decreased in the beginning and the middle of the nineties, but we cannot conclude that this observation remains valid either at the end of the nineties or the beginning of the following decade.

[12] The estimated persistence measure in Levin and Piger (2004) is 0.77 for CPI, 0.78 for GDP price, and 0.79 for PCE price. Two opposite statitical effects play a role in this comparison: the present estimates are based on monthly data which rises the persistence when compared with quarterly estimates, but the additional noise introduced in monthly frequency (sales effect for instance) reduces the persistence measure.

[13] The CPI persistence measure is 0.5 in Gadzinski and Orlandi (2004).

4.2 Persistence at the sector level

At the sector level, I investigate two issues: the aggregation effect and the link between price persistence and price duration.

From Granger (1980) and the literature on long memory, it is well known that we can expect an aggregate to exhibit a stronger autocorrelation than the average autocorrelation of its constituent series. Our measures confirm this expected theoretical aggregation effect on persistence: in Tables 7 and 8, CPI persistence is either the largest or one of the largest of the six aggregates.

A promising and emerging literature aims to measure the duration between two price changes at the firm level (Baudry, Le Bihan, Sevestre and Tarrieu, 2004, for France). From a theoretical viewpoint, linking the empirical observation of the time between two *price level* changes (persistence in level) and aggregate *inflation* persistence (persistence in price changes) is difficult. Time-dependent pricing models, such as the Taylor (1980) contract model and the Calvo (1983) price model, imply persistence in *price level.* As exposed by Cecchetti and Debelle (2004) and demonstrated in the case of Taylor contracts by Whelan (2004), persistence in price changes should be negatively correlated with persistence in level. In other words, the longer the length of time between two price changes, the smaller the inflation persistence should be. In variants of the Calvo model that relax the assumption of forward lookingness (Fuhrer and Moore, 1995), inflation persistence can be positively correlated with the degree to which pricing decisions look backward. In this framework, some previous work found empirical evidence supporting the time-dependent theories of price setting. Bils and Klenow (2002) and Cecchetti and Debelle (2004) observe that the longer the time between two price level changes, the lower the persistence in price change.

Our findings on French CPI inflation are different. Baudry, Le Bihan, Sevestre and Tarrieu (2004) find strong sector differences in the duration between two price changes: service sector prices are changed more rarely than prices of industrial goods, which in turn are revised more rarely than food or energy prices[14]. In the sectoral inflation persistence measures reported in Table 8, the sector aggregates can be split into two groups: services and industrial goods appear to be more persistent, while food and energy price changes appear less persistent. Thus, in our dataset, the longer the duration between two price changes, the

[14]Some studies have found a similar hierarchy for other countries, for instance: the United States (Bils and Klenow, 2002), Belgium (Aucremanne and Dhyne, 2004) or Portugal (Dias, Dias and Neves, 2004). See also Loupias and Ricart (2004) for a survey-based study on French data.

more persistent the price change appears to be. To be consistent with the time-dependent model, our findings should imply that the sectors with more persistence in price changes should also be the sectors with the largest component that looks backward.

5 Insights from simulation experiments

Implementation of multiple breaks test requires some caution in finite sample. This Section investigates to which extent the Altissimo and Corradi (2003) multiple breaks test is affected by three potential pitfalls: a bias implied by a strong persistence of the process, a bias implied by the position within the sample of the true break date, and the incidence of heteroscedasticity.

5.1 Size and power with persistent processes

Multiple breaks test results are supposed to converge asymptotically towards the true number of breaks and the relevant break dates in a time series. In the case of the Altissimo and Corradi (2003) test, the convergence towards the true number of breaks is regarded as "perfect", in the sense that no asymptotic error is expected to occur. However, the asymptotic theory for these tests is not fully consistent with the presence of a high autocorrelation degree in the time series (see Kiefer and Vogelsang, 2002). More precisely, as shown by Vogelsang (1999), the size effect and the power of CUSUM tests depend on the degree of autocorrelation of the initial process. The Altissimo and Corradi (2003) multiple breaks test statistics is a ratio of cumulative distance to a mean on the variance of the residuals of the model estimated under the null hypothesis. The null hypothesis is zero change in a first stage, then a single change, then two changes, and so on until it cannot be rejected. The variance of the residuals is heteroscedasticity and autocorrelation consistent (HAC) which implies the use of a kernel over a given bandwidth. Two kinds of truncation lags are usually proposed: fixed ones which only depend on the sample size (usually $\left[T^{1/3}\right]$ with the Bartlett kernel) and data dependent ones as proposed by Andrews (1991).

Studies of the CUSUM tests in finite sample (Vogelsang, 1999 and Kiefer and Vogelsang, 2002) have shown that a small bandwidth tends to under-estimate the variance of the residuals, thus leading to an upward bias in the test statistic and an over-rejection of H0 (type I error, i.e. an overestimation of the number of breaks) in the case of highly correlated variables, while a large bandwidth may lead to an under-rejection of H0 (type II error or low power, i.e. an underestimation of the number of breaks). The latter can even be so strong that the test could exhibit non-monotonic power. In the case of a multiple breaks test procedure, it means that the ability of the test to detect a break would be negatively related with the size of the break. In a small sample, it could also be impossible to find a bandwidth leading to both acceptable type I and II errors, in the case of highly autocorrelated variables. In this regards, Kiefer and Vogelsang (2002) have shown that automatic bandwidth selection procedure provides worse results than fixed bandwidth. Multiple breaks tests are generally suspected to have low power, which is a materialization of this trade-off between size effect and power.

Monte-Carlo simulations allow for an analyze of this trade-off for highly autocorrelated processes with the Altissimo and Corradi (2003) multiple breaks test for a given sample size. I used this test with a Newey and West (1987) estimation of the variance with fixed truncation lags[15] and an implementation of the finite sample correction of the critical values proposed by Altissimo and Corradi (2003)[16] for a nominal size of the test of 5%. The following DGP is simulated to determine the extent to which the actual size is close to the nominal size:

$$X_t = \mu_1 + 1\{k, T\} \times i + \varepsilon_t \tag{1}$$

$$\text{with } \varepsilon_t = \begin{cases} u_t & \text{DGP NIID } u_t \sim N(0,1) \\ \rho\varepsilon_{t-1} + \nu_t & \text{DGP AR(1) } \nu_t \sim N(0, 1-\rho^2) \end{cases}$$

$1\{k, T\}$ means that 1 is the average between date k and T. In the simulations[17], I set μ to 1, ρ to 0.9, and the sample size (T) to 380, close to that of our case study in the previous Sections. When there is a break in the process, i is set to 1.5 thus the mean shifts from 1 to 2.5, a one and a half times change in the standard deviation of the innovation. This can then be considered as a rather small break. The break date is placed at one third of the sample size ($k = 1/3$). Overall, this framework allows to gauge the finite sample properties of the test and to find the best parameterization for the bandwidth with our sample size (380)[18]. In practise, the procedure needs to combine two features: a high ability to detect

[15] Note than Altissimo and Corradi (2003) have proposed a local mean correction, but it is not fully applicable here. Their procedure relies on a parameter h which governs the number of observations in the neighborhood of the t-th observation which are left out in the computation of the local mean. But their proper h depends on the DGP: small in the NIID case and large in the AR one.

[16] Our simulations lead to the following critical values: 0.702 and 0.736 for 380 and 120 observations at the 95% level.

[17] In the whole Section, each Monte-Carlo exercise relies on 5 000 simulations, under the Gauss program.

[18] Our case study is restricted to NIID and AR cases, with no consideration of MA and ARMA cases. Actually, in their simulations, Kiefer and Vogelsang (2002) observe that the MA case is similar to the NIID one and that the ARMA case is a combination of AR and NIID cases.

the true number of breaks and, when an error is committed, the absence of a strong bias towards underestimation or overestimation. I propose an additional set of simulation with a smaller sample size in order to gauge if our findings can be generalized or not.

Table 9 reports the number of detected breaks, first when there is no break in the initial process then when there is one, each time with both NIID variables and highly autocorrelated variables. With 380 observations and a standard truncation lag ($\left[T^{1/3}\right]$, i.e. 7 observations), the results are correct in the NIID case (the error is below the nominal size of 0.05), but the test clearly over-estimates the number of breaks in highly autocorrelated variables. The type I error (overestimation of the number of breaks) when there is no break then reaches 0.44. When there is a single break, the test detects it only with a probability of 0.54, overestimates it with a probability of 0.43, and underestimates it with a probability of 0.02. This implies that the bandwidth may be too small. The outcome with a larger bandwidth ($\left[T^{1/2}\right]$, i.e. 19 observations) is more convincing. Type I error in the NIID case remains below 0.05. The overestimation of the number of breaks appears more limited with the highly auto-correlated process, as expected, whereas the underestimation rises without becoming unacceptable. The type I error when there is no break now reaches 0.12. When there is a single break, the test detects it with a probability of 0.75, overestimates it with a probability of 0.09, and underestimates it with a probability of 0.15. Thus, despite a type I error probability around 0.12 in the absence of break in highly persistent processes, the overall test accuracy is significantly improved when there is a break in highly persistent processes because the error is then slightly balanced towards an underestimation of the number of breaks. In unreported simulations, I have checked that this configuration does not lead to non-monotonic power for this sample size.

Over a smaller sample size (120 observations, Table 1), two findings can be highlighted: the test then exhibits an unacceptable low power when the process is persistent and the more suitable bandwidth with 380 observations ($\left[T^{1/2}\right]$) is no longer the more efficient any more.

Overall, this small sample study of Altissimo and Corradi (2003) leads us to select a larger bandwidth than usual, in order to limit the persistence effect with a given sample size. The size effect is then satisfactory, even in the case of highly persistent processes, whereas the traditional expected low power of the multiple breaks test procedures is present but rather moderate for highly persistent processes (underestimation of 0.15 versus an overestimation of 0.09). The combination of these two features provides a balanced picture of this test which can not be suspected of an unconditional bias in one direction or the other. However, a persistence effect could remain with a smaller sample size: the test would then clearly underestimate the true number of breaks, leading in turn to a likely overestimation of persistence of the most persistent processes. The bandwidth choice can thus not be generalized, which reflects the absence of an asymptotic foundation for it.

5.2 Dating the break

This issue is rarely documented, however it may be of particular interest for the implementation of multiple breaks test. The position within the sample of the true break can influence both the ability of the test to detect a break and its ability to find the true date. I propose to investigate these two questions, again in the case of the Altissimo and Corradi (2003) procedure[19]. The DGP proposed in Equation 1 are again simulated, with $\left[T^{1/2}\right]$ and $k = 1/2, 2/3, 4/5$.

With a highly persistent process and a small break ($i = 1.5$), the ability of the test to detect the break clearly decreases after $k = 2/3$ (see Table 10). However, it is worth noting that the dating precision is not reduced, thus the test does not detect a break at a wrong date even when the break is close to the end of the sample. The autocorrelation effect is here again perceptible since no decrease in the ability of the test to detect a break can be emphasized with DGP NIID.

The distortion depends on the size of the break. When the size of the break increases ($i = 3$), the distortion with the highly autocorrelated process is reduced since a break at the end of the sample is even more easily detected than a break in the middle. However, it is worth noting that this improvement has a cost, a rise in the probability to overestimate the true number of breaks.

Overall, three findings have to be highlighted: (1) small breaks in highly persistent time series are more difficult to detect when they occur near the sample borders, (2) the dating precision is not affected by the position of the break in the sample, but (3) it is affected by the degree of autocorrelation and the size of the break. In this paper, Table 10 has been used to estimate the confidence intervals around the estimated breaks, conditionally on the position of the break within the sample.

[19]The detected number of breaks is specific to the Altissimo and Corradi (2003) procedure, whereas the dating of the breaks (the precision) in their test is based on the usual criteria of minimization of the residuals as proposed by Bai (1997).

5.3 Volatility change

I conduct a third set of simulations in order to gauge the sensitivity of the test to a volatility change. The sample is divided in two: in its first part, the DGP is similar to Equation 1; whereas, in the second part, the variance of the residuals doubles: $(1-p^2)*\sqrt{2}$. The data are first generated without a break in the mean then with a break at the same date than the volatility change, still with $i = 1.5$. It is worth noting that, in our simulation, the volatility change is of a magnitude significantly larger than the break in the mean, thus leading to a particularly unfavorable situation for the test. The study is for two sample sizes (380 and 120) and two bandwidths, to ensure the parameter previously proposed ($\left[T^{1/2}\right]$) is again the more accurate. The results are reported in Table 11.

The introduction of heteroscedasticity does not have a real impact when there is no break in the true process, with the large sample and the $\left[T^{1/2}\right]$ bandwidth (top of Table 11 compared to top of Table 9). In other words, in these circumstances, the test does not spuriously detect a volatility change as a mean break. On the opposite, when there is a break in the real process, a simultaneous volatility change can hide it in presence of autocorrelation. Like in the first set of simulations, persistence decreases the power of the test.

The reduction of the sample size in the presence of heteroscedasticity is also of particular interest. In the presence of heteroscedasticity and for a persistent process, the test is less efficient in all directions: when there is no break, one is spuriously detected and when there is one, it is less easily detected. Again, this effect does not apply to the NIID case.

Overall, with 380 observations and a larger bandwidth than usual, the vulnerability of the procedure to the persistence effect and heteroscedasticity can be limited to the following: a minor overestimation of the number of break for the non persistent series (below the nominal size of the test) and a mild underestimation of the number of breaks for highly persistent processes which can be raised when a simultaneous change in volatility takes place. In other situations, the test properties are rather satisfactory.

6 Concluding remarks

Thanks to the use of highly disaggregated time series, the dynamics of French inflation is clearly defined. First, during the past thirty years, a single structural change occurred in the mid-eighties. This change is broadly diffused across the entire CPI basket and can be directly linked to a major monetary policy change, among several policy related shocks.

Second, inflation persistence in France is moderate, once this structural change is accounted for. It is not possible to highlight any evidence of a structural decrease in the persistence measure. Particular caution is required here as it seems uncertainty around the parameter measure has increased recently.

Finally, some work remains to reconcile the evidence found at the firm and sector levels. To test if the time-dependent model is still compatible with our findings, an investigation of how the degree of backward lookingness varies among the sectors may be a follow-up of the present work.

From a methodological point of view, note that even the test considered to be the state of the art of multiple breaks tests (Altissimo and Corradi, 2003) requires a careful treatment of autocorrelation. Therefore, if not accounting for a break can lead to misestimate persistence, I have shown that the reverse is also true, that not accounting for persistence can lead to spurious estimates of structural breaks. A possible way to explore is certainly to relate this test with the new asymptotic developments proposed by Kiefer and Vogelsang (2002).

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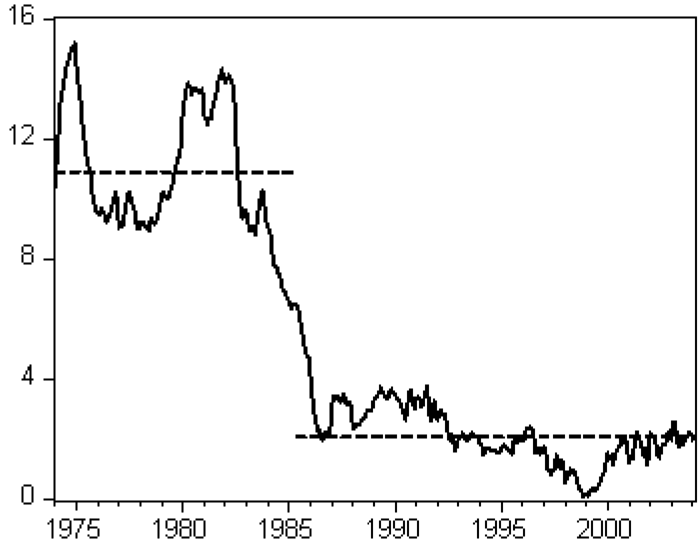
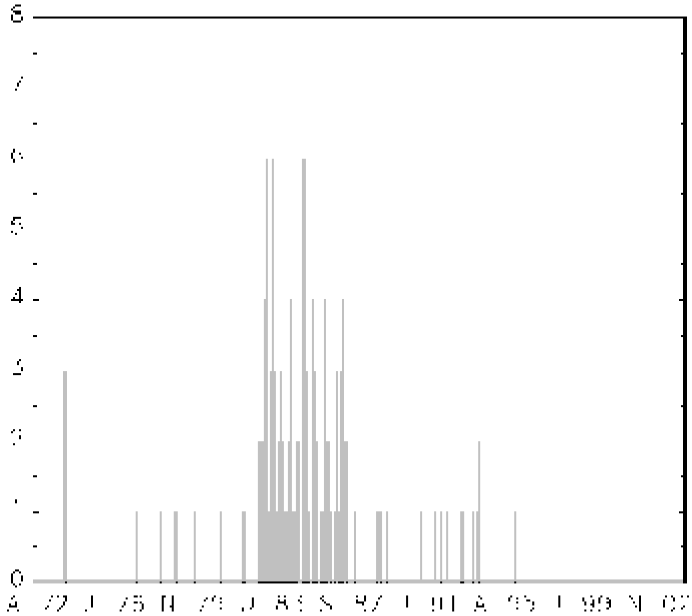
FIGURES

Figure 1: CPI, 12 months growth rate %



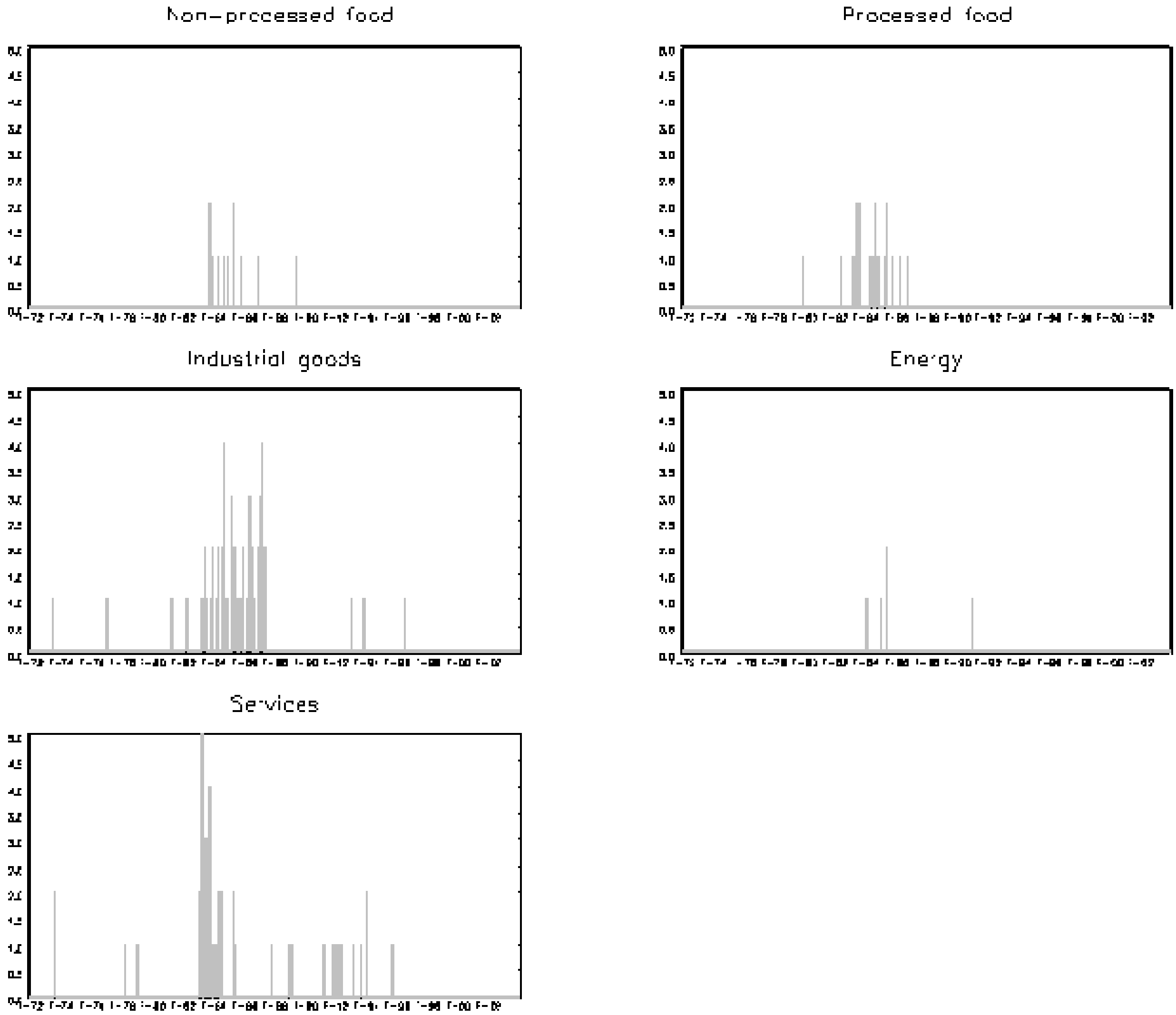
Notes: the CPI is as retropolated by Baudry and Tarrieu (2003). The dot lines are the two historical averages: 10.9% before 1985:5 and 2.1% after.

Figure 2: break dates distribution, item level



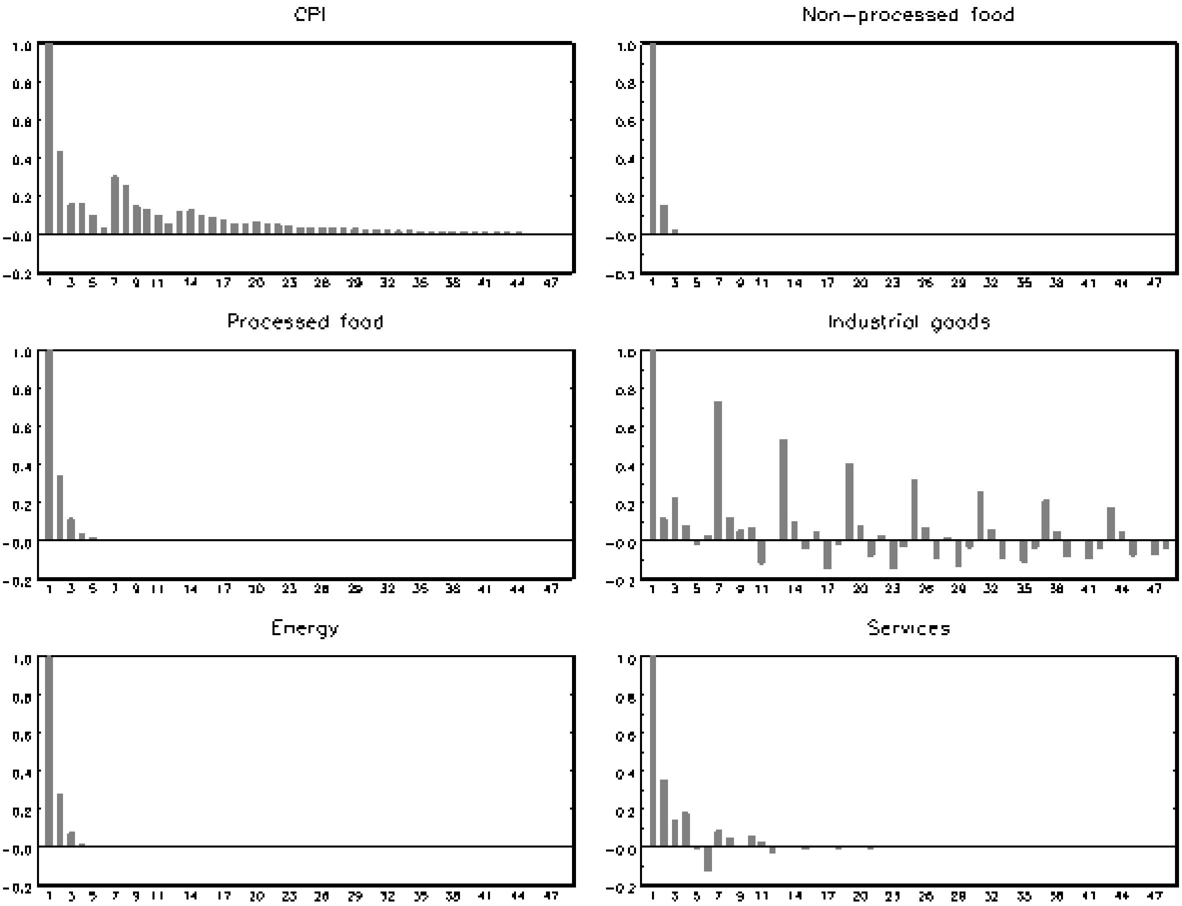
Notes: the chart report the number of breaks detected by the multiple breaks test procedure in the 141 items, at each month.

Figure 3: item level break dates, distribution by sector



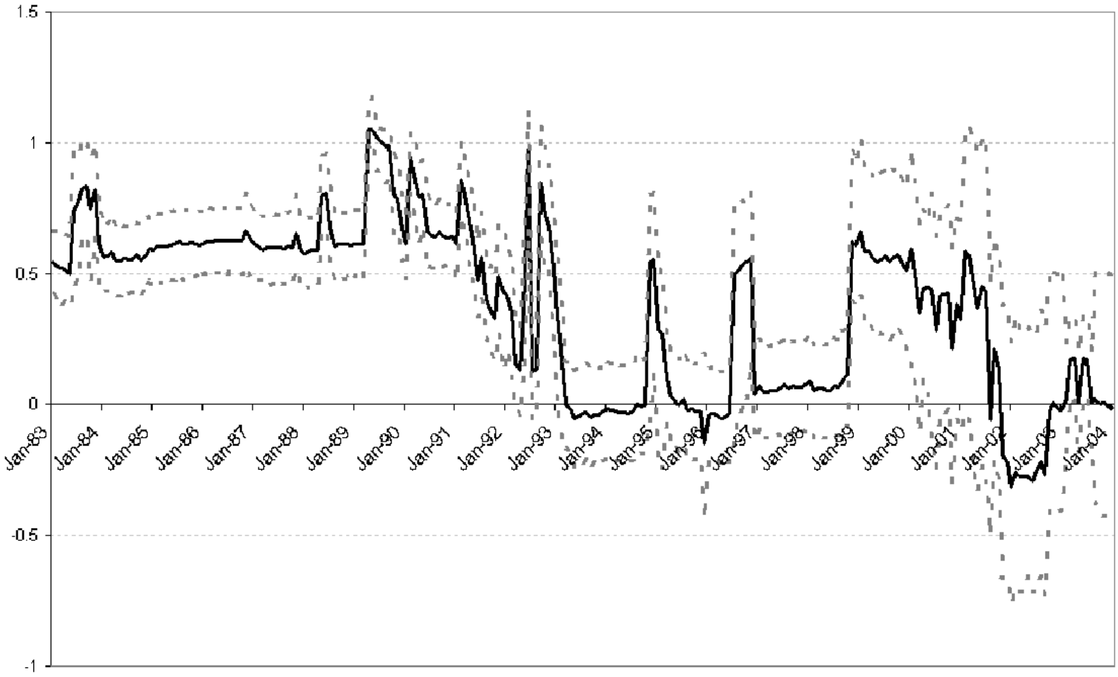
Notes: Figure 3 is the sectoral breakdown of Figure 2.

Figure 4: impulse response functions of the fitted process, once accounting for the structural break



Notes: the charts report the effect of an unit single deviation of the innovation on the prices, as obtained by numerical simulation. The constant coefficients are forced to be nulled, so that the IRF account for the structural breaks but are not representative of what happens in the immediate neighborhood of the breaks.

Figure 5: CPI persistence, 10 years rolling regressions



Notes: the chart report estimates of CPI persistence over rolling periods of 10 years, together with their 90% confidence interval band.

TABLES

Table 1: sectoral aggregates descriptive statistics

Group name	Number of items (2004:1)	Weight (/10 000) Average 1972-2003
Non processed food (A)	12	1 126
Processed food (B)	27	1 382
Industrial goods (C)	61	3 462
Energy (D)	8	850
Services (E)	53	3 180

Notes: this table reports the sectoral composition of the French CPI at the 161 items level, following the standard HICP disaggregation.

Table 2: distribution of the estimated number of breaks, item level

number of breaks=	0	1	2	≥ 3
CPI aggregate	.00	1	.00	.00
141 components	.08	.82	.08	.02
in which				
Non-processed food	.17	.83	.00	.00
Processed food	.19	.78	.04	.00
Industrial goods	.02	.89	.09	.00
Energy	.29	.71	.00	.00
Services	.03	.76	.13	.08

Notes: the table reports the distribution of the estimated number of breaks for CPI and the 141 item level prices, classified by sector. For instance, in the case of non-processed foods, no break has been detected for 17% of the items. Detailed item level results are proposed in Table 5.

Table 3: estimated break dates, aggregate level

	1st date	2nd date
CPI	1985:5	-
Non-processed food	1984:7	-
Processed food	1984:5	-
Industrial goods	1985:7	-
Energy	1985:4	-
Services	1983:9	1993:2

Note: the sample is 1973:1 - 2004:1.

Table 4: change in the mean after the breaks, aggregate level

Mean, monthly growth rate (annualized)	before 1st break	after 1st break	after 2nd break
CPI	.86 (10.9)	.17 (2.1)	-
Non-processed food	.86 (10.8)	.19 (2.3)	-
Processed food	.84 (10.6)	.25 (3.0)	-
Industrial goods	.76 (9.5)	.11 (1.3)	-
Energy	1.16 (14.8)	.06 (0.7)	-
Services	.90 (11.4)	.39 (4.8)	.16 (1.9)

Notes: the table reports the average price growth at the aggregate level, once taking into account the structural changes. The break dates are as reported in Table 3. The figures into brackets are annualized monthly growth rates, the others are non-annualized monthly growth rates.

Table 5: estimated break dates, item level

Item	INSEE id.	Weight	Type	Break	Dates
CPI		10 000	-	1	1985:5
Bread	i01111	97	B	1	1983:9
Pasta products	i01112	86	B	1	1984:9
Pastry-cook products	i01113	49	B	1	1985:11
Cereals	i01114	47	B	1	1985:10
Fresh, chilled or frozen meat of bovine animals	i01121	184	A	1	1983:10
Fresh, chilled or frozen meat of swine	i01122	54	A	1	1989:6
Fresh, chilled or frozen meat of sheep and goat	i01123	52	A	1	1984:5
Pork meat and cooked pork meat	i01124	239	A	1	1984:10
Fresh, chilled or frozen meat of poultry	i01125	82	A	1	1985:1
Other preserved or processed meat and meat preparations	i01126	87	A	1	1983:12
Fresh, chilled or frozen fish and seafood	i01131	72	A	1	1986:12
Dried, smoked or salted fish and seafood	i01132	49	A	1	1985:11
Milk and cream	i01141	75	B	1	1984:7
Yogurt and milk based dessert	i01142	40	B	1	1984:11
Cheese	i01143	151	B	1	1985:10
Eggs	i01144	32	B	0	-
Butter	i01151	53	B	2	1983:9 1979:12
Oil and margarine	i01152	37	B	0	-
Fresh fruits	i01161	118	A	0	-
Frozen fruits	i01162	11	A	1	1985:5
Fresh vegetables	i01171	138	A	0	-
Cooked vegetables	i01172	49	A	1	1983:10
Sugar based products	i01181	59	B	1	1986:5

Item	INSEE id.	Weight	Type	Break	Dates
Chocolate based products	i01182	44	B	1	1986:11
Ice creams	i01183	28	B	1	1984:6
Condiments and sauces	i01191	19	B	1	1985:10
Processed cook for children and dietetical products	i01192	11	B	1	1983:5
Other food products n.e.c.	i01193	9	B	1	1984:9
Cocoa and powdered chocolate	i01211	10	B	0	-
Coffee	i01212	47	B	0	-
Tea and infusion	i01213	4	B	1	1985:4
Mineral or spring water	i01221	32	B	1	1985:5
Soft drink	i01222	34	B	1	1984:10
Aperitif	i02111	26	B	1	1983:7
Brandy and liquor	i02112	28	B	1	1983:7
Wine	i02121	146	B	1	1983:4
Champagne, sparkling wine and cider	i02122	23	B	1	1982:6
Beer	i02131	31	B	1	1984:5
Tobacco	i02211	171	B	0	-
Clothing materials	i03111	16	C	1	1987:4
Garments for men	i03121	94	C	1	1986:12
Garments for women	i03122	135	C	1	1985:11
Garments for children	i03123	47	C	1	1986:5
Sport clothes	i03124	23	C	1	1986:5
Underwear for men	i03125	78	C	1	1985:12
Underwear for women	i03126	111	C	1	1987:2
Underwear for children	i03127	54	C	1	1986:5
Other articles of clothing and clothing accessories	i03131	67	C	2	1986:4 1993:2
Cleaning, repair and hire of clothing	i03141	18	E	3	1983:2 1994:1 1973:8
Footwear	i03211	113	C	1	1985:12
Other footwear including repair	i03212	47	C	1	1986:12
Actual rentals paid by tenants	i04111	554	E	1	1988:12
Actual rentals for holidays	i04112	12	E	1	1984:8

Item	INSEE id.	Weight	Type	Break	Dates
Materials for the maintenance and repair of the dwelling	i04311	28	C	1	1985:10
Floor covering and wall repair services	i04321	51	E	1	1984:4
Other services for the maintenance of the dwelling	i04322	72	E	1	1984:4
Water supply	i04411	83	C	2	1982:5 1996:7
Other services related with the dwelling n.e.c.	i04414	12	E	2	1983:7 1995:10
Electricity	i04511	203	D	1	1984:3
Gas for domestic use	i04521	89	D	1	1985:1
Repair of personal transport equipment	i07232	174	E	1	1983:4
Toll and carparks	i07241	36	E	1	1992:2
Other services for personal vehicles	i07242	27	E	1	1983:10
Passenger transport by railway	i07311	67	E	2	1985:5 1978:4
Passenger transport by road	i07321	55	E	1	1984:2
Taxis	i07322	14	E	1	1983:3
Combined passenger transport	i07351	47	E	1	1983:8
Other purchased transport services	i07361	12	E	1	1983:3
Postal services	i08111	31	E	1	1983:6
Telecommunication services	i08122	110	E	1	1984:8
Equipment for the reception and recording of sound and pictures	i09111	100	C	1	1987:4
Photographic and cinema equipment, optical instruments	i09121	22	C	2	1987:6 1993:11
Recording media for pictures and sound	i09141	44	C	1	1983:12
Other major durable for recreation and culture	i09211	23	C	1	1983:6
Games, toys hobbies	i09311	68	C	1	1986:6
Equipment for sport, camping and open-air recreation	i09321	31	C	1	1986:6
Flowers and plants	i09331	38	C	0	-
Seeding an seeds	i09332	28	C	1	1983:12

Item	INSEE id.	Weight	Type	Break	Dates
Recreational services	i09411	45	E	1	1987:11
Cinemas	i09421	20	E	2	1983:4 1991:4
Museums, zoological gardens	i09422	24	E	1	1989:2
Television and radio taxes and hire of equipment	i09423	42	E	1	1984:1
Other cultural services	i09424	42	E	1	1983:11
Books	i09511	45	C	1	1983:11
Newspapers	i09521	46	C	1	1982:4
Magazines	i09522	68	C	1	1986:3
Miscellaneous printed matter	i09531	32	C	1	1985:7
Other office accessories	i09532	14	C	1	1985:7
Package holidays	i09611	24	E	1	1985:7
Education services	i10111	26	E	1	1983:11
Restaurants	i11111	324	E	2	1983:4 1991:11
Cafés, bars and the like	i11112	201	E	3	1983:9 1993:9 1979:1
School or university canteen	i11121	70	E	1	1983:10
Professional canteen	i11122	88	E	2	1984:5 1993:3
Hotel	i11211	96	E	1	1983:7
School or university pension	i11212	16	E	1	1984:3
Holiday accommodation	i11213	25	E	0	-
Hairdressing	i12111	82	E	1	1983:10
Other aesthetic services	i12112	8	E	1	1992:6
Perfumes and beauty Products	i12131	75	C	1	1984:5
Personal care products	i12132	55	C	1	1985:6
Other toilet articles and equipment	i12133	37	C	1	1985:8
Jewelry	i12311	122	C	1	1981:4
Leather working and travel goods	i12321	41	C	1	1986:9
Other personal effects, incl. repair	i12322	36	C	1	1987:6

Notes: this table reports the estimated number of breaks and the break dates for the 141 items, over the sample 1972:2 - 2004:1. The weights sum to 10,000. The "type" stands for the aggregate the item belongs to: A for non-processed food, B for processed food, C for non-energy industrial goods, D for energy goods and E for services. The "break" column reports the number of breaks detected in the time series. "INSEE id" is the code used by INSEE to identify the time series. For more details on the backward retropolated time series, see Baudry and Tarrieu (2003).

Table 6: estimated break dates with the reduced sample, aggregate level

	1st date	2nd date
CPI	1985:7	1991:11
Non-processed food	-	-
Processed food	-	-
Industrial goods	1987:4	1993:3
Energy	-	-
Services	1985:6*	1993:2

Notes: the table reports the break dates detected in the 6 aggregate time series over the sample 1984:1 - 2003:3 used by Gadzinski and Orlandi (2004) and Levin and Piger (2004). * indicates that the break is not robust to a change in the bandwidth. This table can be compared with Table 3 which covers a larger sample.

Table 7: persistence without structural break

	α	90% CI	p
CPI	.98	[.92,1.01]	6
Non processed food	.54	[.40,.68]	6
Processed food	.80	[.73,.90]	6
Industrial goods	.97	[.89,1.01]	8
Energy	.37	[.29,.45]	1
Services	1.00	[.95,1.03]	11

Notes: persistence at the aggregate level is estimated under the wrong hypothesis that the mean is constant over time.

Table 8: persistence with structural breaks

	α	90% CI	p
CPI	.76	[.64, .88]	6
Non processed food	.15	[.07, .23]	1
Processed food	.34	[.26, .41]	1
Industrial goods	.72	[.58, .84]	8
Energy	.28	[.19, .36]	1
Services	.44	[.23, .60]	6

Notes: persistence at the aggregate level is estimated once taking account of the structural breaks as reported in Table 3.

Table 9: estimated number of breaks and persistence

			0	1	2	≥3
T=380	$\left[T^{1/2}\right]$	AR, p=.9				
		No break	.88	.11	.01	.00
		1 break, i=1.5	.15	.75	.08	.01
		NIID				
		No break	.96	.04	.00	.00
		1 break, i=1.5	.00	.96	.04	.01
	$\left[T^{1/3}\right]$	AR, p=.9				
		No break	.56	.26	.13	.05
		1 break, i=1.5	.02	.54	.26	.17
		NIID				
		No break	.96	.04	.00	.00
		1 break, i=1.5	.00	.95	.04	.00
T=120	$\left[T^{1/2}\right]$	AR, p=.9				
		No break	.85	.14	.01	.00
		1 break, i=1.5	.74	.24	.02	.00
		NIID				
		No break	.97	.03	.00	.00
		1 break, i=1.5	.28	.68	.03	.01
	$\left[T^{1/3}\right]$	AR, p=.9				
		No break	.43	.32	.19	.06
		1 break, i=1.5	.33	.37	.20	.10
		NIID				
		No break	.96	.04	.00	.00
		1 break, i=1.5	.11	.84	.04	.00

Notes: this table reports the results of Monte-Carlo simulations of Equation 1, for two data generating process (AR and NIID) without break and with a break (i=1.5). T is the sample size. The Table shows the estimated number of breaks with two choices of bandwidth for the computation of HAC robust variance, $\left[T^{1/2}\right]$ i.e. 19 observations and the standard $\left[T^{1/3}\right]$ i.e. 7 observations. The nominal size of the test is 0.05.

Table 10: power of the test and dating precision, as a function of the break position within the sample

		k	Number of detected breaks 0	1	≥ 2	Dating precision Exact	$\pm.01$	$\pm.05$	$\pm.10$	$> \pm.10$
					T=380					
i=1.5	DGP AR	1/2	.08	.79	.13	.28	.12	.27	.15	.18
		2/3	.14	.77	.09	.28	.13	.28	.15	.17
		4/5	.47	.48	.06	.27	.11	.28	.13	.22
	DGP NIID	1/2	.00	.91	.09	.47	.44	.09	.00	.00
		2/3	.00	.95	.05	.47	.43	.10	.00	.00
		4/5	.00	.95	.05	.48	.42	.10	.00	.00
i=3.0	DGP AR	1/2	.00	.82	.18	.80	.07	.10	.03	.01
		2/3	.00	.87	.13	.78	.08	.10	.03	.01
		4/5	.01	.86	.12	.78	.07	.11	.03	.01
	DGP NIID	1/2	.00	.91	.08	.86	.14	.00	.00	.00
		2/3	.00	.96	.04	.85	.15	.00	.00	.00
		4/5	.00	.96	.04	.86	.14	.00	.00	.00
					T=120					
i=3.0	DGP AR	1/2	.02	.78	.20	.83	.03	.06	.04	.04
		2/3	.05	.81	.14	.83	.03	.06	.04	.04
		4/5	.42	.49	.09	.84	.02	.04	.03	.07

Notes: k stands for the position of the break within the sample (k=1/2 means that the break in the simulated process occurs at observation 190 out of 380). The dating precision columns report the probability that the detected break is exactly at the true date ("exact"), is in more or less 1% of the sample around the true date but not at the true date ("$\pm.01$"), and so on until it is outside more or less 10% around the true date ("$> \pm.10$"). The nominal size of the test is 0.05.

Table 11: estimated number of breaks when there is a change in volatility

			0	1	2	≥3
T=380	$[T^{1/2}]$	AR, p=.9				
		No break	.90	.10	.01	.00
		1 break, i=1.5	.24	.68	.08	.00
		NIID				
		No break	.98	.02	.00	.00
		1 break, i=1.5	.00	.93	.07	.00
	$[T^{1/3}]$	AR, p=.9				
		No break	.60	.23	.12	.04
		1 break, i=1.5	.05	.52	.32	.11
		NIID				
		No break	.97	.03	.00	.00
		1 break, i=1.5	.00	.94	.06	.00
T=120	$[T^{1/2}]$	AR, p=.9				
		No break	.82	.16	.02	.00
		1 break, i=1.5	.45	.47	.08	.00
		NIID				
		No break	.97	.03	.00	.00
		1 break, i=1.5	.00	.90	.09	.01
	$[T^{1/3}]$	AR, p=.9				
		No break	.42	.31	.21	.06
		1 break, i=1.5	.16	.39	.32	.13
		NIID				
		No break	.96	.04	.00	.00
		1 break, i=1.5	.00	.92	.08	.00

Notes: the volatility of the innovation doubles between the first half and the second half of the sample. When there is a mean break ("1 break, i=1.5"), the volatility change and the mean break are simultaneous. The nominal size of the test is 0.05.

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