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Time-dependent versus State-dependent Pricing: A Panel Data Approach to the Determinants of Belgian Consumer Price Changes

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Abstract

Using Logistic Normal regressions, we model the price-setting behaviour for a large sample of Belgian consumer prices over the January 1989 - January 2001 period. Our results indicate that time-dependent features are very important, particularly an infinite mixture of Calvo pricing rules and truncation at specific horizons. Truncation is mainly a characteristic of pricing in the service sector where it mostly takes the form of annual Taylor contracts typically renewed at the end of December. Several other variables, including some that can be considered as state variables, are also found to be statistically significant. This is particularly so for accumulated sectoral inflation since the last price change. Once heterogeneity and the role of accumulated inflation are acknowledged, hazard functions become mildly upward-sloping, even in a low inflation regime. The contribution of the state-dependent variables to the pseudo-R^2 of our equations is, however, not particularly important.

JEL code: C23, C25, D40, E31

Keywords: consumer prices, time-dependent pricing, state-dependent pricing, Calvo model, Truncated Calvo model, Taylor contracts.

Non technical summary

This paper presents a panel data approach to the determinants of price changes for Belgian consumer prices. Using a large sample of monthly individual price reports underlying the Belgian CPI, we estimate a model of price adjustment with state-dependent features similar to the one proposed by Cecchetti (1986). However, compared to the specification used by Cecchetti (1986), our model also incorporates variables capturing more time-dependent pricing rules and variables characterising some commercial practices at the retail level. Our specification also allows for unobserved heterogeneity at the product/store level.

The issues we study are closely related to two (partially overlapping) strands of the relevant literature: i) the empirical relevance of state-dependent pricing relative to time-dependent pricing; ii) the slope of the hazard function that characterises the distribution of the prices' durations. Time-dependent pricing rules all have non-decreasing hazard functions (either constant hazards or constant hazards with a spike equal to one at a particular truncation horizon). State-dependent pricing rules tend to produce more smoothly upward sloping hazard functions, at least in steady state and with non-zero inflation, as is shown in Dotsey, King and Wolman (1999).

Although our aggregate (unconditional) hazard is downward sloping, our estimation results show that hazards can increase, once heterogeneity is taken into account and the incentive to adjust prices stemming from sectoral inflation is acknowledged. It is argued in this paper that testing for state-dependence requires this incentive to be explicitly taken into account, rather than just testing whether the probability to observe a price change increases during the period that has elapsed since the previous price change. It is further shown that an appropriate estimation of the role of sectoral inflation on the hazard requires that a distinction is made between price increases and price decreases, as inflation affects both phenomena in the opposite direction.

Our econometric results tend to reject a general pricing rule that is the result of aggregating purely time-dependent theoretical pricing models only, as several other variables, including some that can clearly be considered as state variables, are found to be significant. This is particularly so for sectoral inflation. The probability to observe a price change is, in a significant way, an increasing function of the accumulated sectoral inflation since the last price change. This is particularly so when the role of accumulated inflation is estimated separately for price increases and for price decreases. As a result, we were able to construct upward sloping hazards under certain conditions, once the impact of duration incorporates this effect through accumulated inflation.

This result tends to suggest that sectoral developments condition the relevant "state" for price setting. This does not mean, however, that aggregate inflation is not relevant at all, because movements in aggregate inflation would shift the entire distribution of sectoral inflation rates and, therefore, affect the probabilities to change prices through this channel. In addition, we found that a set of yearly dummies had a significant impact on the price changing probabilities and the coefficients associated with these dummies reflect, to some extent, a pattern which is comparable with the evolution of aggregate (core) inflation over the period considered. We also found that VAT rate changes increase the probability to observe a price change, whereas attractive pricing coincides with a somewhat smaller tendency to adjust prices.

Although our estimation results clearly reject a purely time-dependent representation of price-setting in a statistically significant way and therefore favour a more heterogeneous representation that mixes both time-dependent and state-dependent features, it is important to assess the contribution of each type of pricing in more quantitative terms. This has been done by evaluating the contribution of the time- and state-dependent features to the pseudo-R^2 of the estimated models.

In this respect, our results point out that the time-dependent characteristics are far more important than the state-dependent characteristics. For the core CPI, the contribution of an infinite mixture of Calvo rules to the pseudo-R^2 attains somewhat more than 50 p.c. for price increases (but 82 p.c. for price decreases). For price increases, seasonality and truncation make a contribution of nearly 14 p.c., a level comparable to the contribution of accumulated inflation. For price decreases, the contribution of the two latter is less than half their contribution for price increases. The role of accumulated inflation is most important for processed food (somewhat more than 20 p.c. of the pseudo R^2 for price increases in that sector). Seasonality and truncation are very widespread in services. Their contribution to the pseudo R^2 for price increases is close to 50 p.c. in this sector.

As to the truncation phenomena, we were able to disentangle those that occur on a regular basis (Taylor pricing) from those that occur irregularly, for instance as a result of so-called truncated Calvo pricing. It turned out that Taylor pricing is estimated significantly, particularly at the 6 and 12 months horizons. However, it seems that this type of Taylor pricing is only relatively important for services, where price spells tend to come to an end at fixed durations of 12 months in a fairly synchronised way at the end of the year (new prices observed in January).

All in all, although our econometric results clearly reject a purely time-dependent representation of price-setting in a statistically significant way, it turns out that the importance of state-dependent features in price-setting is relatively limited. Indeed, at most 20 p.c. of the pseudo-R^2 of the estimated models for the core components of the CPI is attributable to the role of accumulated inflation.

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1. INTRODUCTION

This paper presents a panel data approach to the determinants of price changes for a large sample of Belgian consumer prices. We use mixed Logit models (logistic normal regression), which can be considered as an alternative to duration models with unobserved heterogeneity and time-dependent covariates. We consider a model of price adjustment with state-dependent features similar to the one proposed by Cecchetti (1986) as our starting point. The aim is to analyse the probability to observe that the price of one specific product sold in store i, belonging to product category j, is changed after T periods. These probabilities - conditional on the time period that has elapsed since the previous price change - are often referred to as hazards in duration models. We use a large sample of monthly individual price reports underlying the Belgian CPI observed during the period January 1989 - January 2001. In doing so, we extend the framework used by Cecchetti in several respects.

First of all, whereas Cecchetti (1986) only analysed one type of product, our sample covers 529 product categories and, therefore, has two cross-sectional dimensions, leading to heterogeneity both across and within product categories. Taking advantage of the heterogeneity across product categories is similar to what is done by Konieczny and Skrzypacz (2003) for Polish data. However, they do not analyse the heterogeneity within product categories. Combining both dimensions seemed only natural to us, as in a more descriptive analysis of the frequency of price changes for Belgian consumer prices (Aucremanne and Dhyne (2004)) we found a huge amount of heterogeneity in the observed durations of price spells, not only across product categories but also within relatively narrowly defined product categories.

Moreover, compared to the specification used by Cecchetti (1986), our model incorporates not only variables, such as the accumulated inflation rate since the last price change, that indicate the use of state-dependent pricing rules but also variables capturing more time-dependent pricing rules and variables characterising some commercial practices at the retail level. Combining different types of price-setting behaviour is motivated by the fact that our above-mentioned analysis showed both state-dependent and time-dependent pricing. Indeed, periods characterised by higher inflation were also characterised by a higher aggregate frequency of price changes, while at the same time significant seasonal patterns in the frequency of price changes were identified. A similar mixture of both types of pricing practices is found in Aucremanne and Druant (2004). The time-dependent component of our model comprises, besides monthly seasonal dummies, variables that describe price-setting models with random durations (as in Calvo (1983)) and fixed durations (as in Taylor (1980)), as well as the so-called truncated Calvo model. We particularly incorporate a variable in the estimated model that allows distinguishing between spikes in the hazard function due to Taylor contracts, on the one hand, and spikes due to truncated Calvo, on the other hand.

Most DSGE macro models assume time-dependent pricing rules, often à la Calvo. In these models, the timing of price adjustment is exogenous, irrespective of economic conditions. In his survey on staggered price- and wage-setting in macroeconomics, Taylor (1999) points out that precisely this exogeneity is one of the most criticised assumptions of sticky price models. State-dependent pricing seems to be a more realistic representation in this respect, as it implies that the timing of the adjustment is affected by economic conditions. However, these pricing rules are more difficult to model and are therefore very rarely used in dynamic macro models, Dotsey, King and Wolman (1999) being an important exception. Being able to discriminate between these different representations of the pricing behaviour is nonetheless of great interest, as Wolman (1999) shows that, for a given marginal cost shock, inflation dynamics depend on the pricing rule used. The Calvo specification yields a relatively small, but fairly persistent response of inflation, whereas inflation reacts much more strongly with a state-dependent pricing rule.

The issues we study are closely related to two (partially overlapping) strands of the relevant literature. First of all, the empirical relevance of state-dependent pricing relative to time-dependent pricing is also analysed in Klenow and Kryvtsov (2004), be it from a different methodological angle. In their breakdown of the variance of US CPI inflation, they find that only roughly 10 p.c. of the variance is stemming from state-dependent factors. With such a limited fraction of state-dependent behaviour, they find that the dynamics of a mixed model do not differ much from that of a time-dependent model. The results of ad hoc surveys on price-setting behaviour in several euro area countries, reveal however a higher fraction of state-dependent pricing (see for instance Aucremanne and Druant (2004) for Belgium and Fabiani et al. (2004) for a summary of euro area results in this respect).

Second, our paper fits quite well in with the literature that addresses the issue of the slope of the hazard function. The above-mentioned time-dependent pricing rules all have non-decreasing hazard functions (either constant hazards or constant hazards with a spike equal to one at a particular truncation horizon). State-dependent pricing rules tend to produce more smoothly upward sloping hazard functions, at least in steady state and with non-zero inflation, as is shown in Dotsey, King and Wolman (1999). That is why Mash (2004) calibrates a model with an increasing hazard and he presents this approach as "*a stepping stone between the current workhorse models based on a rather simple time dependent rule (Calvo pricing) and future models based on full state dependence*"[1]. While theoretical price-setting models have either non-decreasing or increasing hazards, empirical studies often find downward sloping hazards[2]. Álvarez et al. (2004) show, as to price-setting models, that aggregation of heterogeneous non-decreasing hazards can lead to downward sloping aggregate hazards. This aggregation effect is also known for instance in the

1 Mash (2004), page 6.
2 For empirical evidence on this issue, see for instance Álvarez et al. (2004), Campbell and Eden (2004), Veronese et al. (2004), Fougère et al. (2004), Jonker et al. (2004).

literature on unemployment[3] and investment[4]. Although our aggregate (unconditional) hazard is downward sloping, our estimation results show that hazards can increase, once heterogeneity is taken into account and the incentive to adjust prices stemming from sectoral inflation is acknowledged. It is argued in this paper that testing for state-dependence requires this incentive to be explicitly taken into account, rather than just testing whether the probability to observe a price change increases during the period that has elapsed since the previous price change. It is further shown that an appropriate estimation of the role of sectoral inflation on the hazard requires that a distinction is made between price increases and price decreases, as inflation affects both phenomena in the opposite direction.

The rest of this paper is structured as follows. In section 2 we present the data set and some basic facts characterising the price-setting process in the Belgian economy, in particular the downward sloping non-parametric estimates of the aggregate hazard. Next, section 3 presents a Logit representation of the different pricing rules (Calvo, Taylor or truncated Calvo for the time-dependent rules and for state-dependent pricing à la Cecchetti) and proposes a baseline representation of our price-setting model, both for price changes and for price increases and price decreases separately. We also propose an alternative specification that distinguishes between spikes in the hazard function due to Taylor contracts and spikes due to truncated Calvo. Section 4 presents the estimation results of the baseline and the alternative models, based on the entire sample and various sub-samples. Section 5 tries to assess the empirical importance of state-dependent pricing with respect to that of time-dependent pricing. Finally, our main conclusions are summarised in section 6.

2. THE DATA SET

As mentioned above, our empirical analysis is based on a sample of price reports underlying the computation of the Belgian CPI. Those prices are collected on a monthly basis by the Federal Public Services "Economy, SMEs, Self-employed and Energy" (former Ministry of Economic Affairs). For reasons set out in Aucremanne and Dhyne (2004), the period covered by this analysis starts in January 1989 and ends in January 2001. Our raw data set considers 583 product categories and consists of 18,910,857 price records.

This data set covers only the product categories for which the prices are recorded in a decentralised way, i.e. 68.1 p.c. of the Belgian CPI in December 2000. The remaining 31.9 p.c. pertain to product categories that are monitored centrally by the Federal Public Services, such as housing rents, electricity, gas, telecommunications, most health care services, cars, newspapers, education services, and insurance services. It should be stressed that food and non-alcoholic beverages are

[3] For instance Lancaster and Nickell (1980).
[4] For instance Cooper et al. (1999).

over-represented in terms of observations in our data set. In fact, these products represent 57.4 p.c. of the price reports in December 2000, whereas their actual weight in the CPI amounts to only 21.4 p.c.

For each particular individual price report in the data set, we have information on the outlet where the price report has been carried out (a store code identifying 26,809 stores located in 65 cities[5]) and on the product surveyed (price, packaging and brand). An important characteristic of our data set is that the prices reported account for all types of discounts and temporary promotions, except for those related to the winter and summer sales periods, which typically take place in January and July. During these sales periods, the prices reported refer to the prices quoted without the discounts; they are not the prices paid by the customers.

Using the information characterising the individual products and the individual outlets enabled us to follow the price of a specific product (that means an individual product sold in a particular store) over time.

We define P_{ijt} as the price at time t of the individual product sold in store i belonging to the product category j.

As to the price reports associated with such an individual product, the following notations can also be introduced:

- $\left(P_{ij\tilde{t}}, \tilde{t}, T\right)$ represents a price spell associated with the individual product sold in store i and belonging to the product category j. This price spell starts in time $\tilde{t}$ and is characterised by a duration of T periods. During the period $\left(\tilde{t}, \tilde{t}+T\right)$, the price of the individual product under review remains unchanged at the level $P_{ij\tilde{t}}$;
- $\left\{\left(P_{ijt_1}, t_1, T_1\right), \left(P_{ijt_1+T_1}, t_1+T_1, T_2\right), \ldots, \left(P_{ijt_1+T_1+T_2+\ldots+T_{n-1}}, t_1+T_1+T_2, \ldots+T_{n-1}, T_n\right)\right\}$ characterises the price trajectory of the individual product sold in store i and belonging to the product category j. t_1 is the first month that individual product is reviewed, while $t_1+T_1+ \ldots +T_{n-1}+T_n$ is the last month.

The first price spell of a price trajectory is left-censored because the starting date of that price spell, t_1, is not necessarily associated with the month during which the price P_{ijt_1} was set. Similarly, the

5 Because the average observation period for one particular store is 51.4 months, the total number of stores observed in our data set has to be considered taking into account the fact that, on average, 3 different stores in a row should be surveyed in order to be able to follow continuously the price of one particular product category in one particular location over the January 1989 - January 2001 period.

last price spell of a price trajectory is right-censored because the end of that price spell, $t_1+T_1+ \ldots +T_{n-1}+T_n$, does not necessarily coincide with a price change. The other price spells of a price trajectory are uncensored.

It should be pointed out that, within a price trajectory, a price spell can end because of a true price change or because of a product replacement. Based on arguments presented in Aucremanne and Dhyne (2004), both events are considered as equivalent in the empirical analysis.

Using the above definition of a price spell and considering the 3,278,526 price spells that were identified in the entire data set, Aucremanne and Dhyne (2004) analysed the unconditional frequency of price changes that characterised the Belgian consumer prices during the January 1989 - January 2001 period.

Several basic facts characterising the frequency of price changes have been put forward:

1. price setting is very heterogeneous across product categories;
2. the pricing behaviour is also characterised by a substantial degree of heterogeneity within relatively homogeneous product categories;
3. the frequency of price changes is the highest for energy products and unprocessed food; processed food holds an intermediate position, while the frequency of price changes is the lowest for non-energy industrial goods and services;
4. each month, nearly 17 p.c. of consumer prices are changed on average;
5. price decreases are not uncommon, except for services;
6. the frequency of price changes is positively affected by important shocks such as VAT rate changes. Moreover, for the core components of the CPI, the frequency of price changes seems to be positively correlated with aggregate (core) inflation. These findings point towards state-dependent pricing practices;
7. however, there seems to be evidence showing time-dependent aspects in the price setting behaviour as well, as price changes are more likely at the beginning of the year and in September and October, while they are less frequent in the summer and in December.

As has been mentioned in the introduction, the purpose of this article is to try to explain this heterogeneous behaviour by incorporating both elements of state- and time-dependence in the pricing rules applied by firms. For empirical tractability purposes, this has been conducted using a sub-sample of the entire data set. This sample was selected using the following procedure.

Firstly, we disregarded price reports associated with seasonal products, in other words product categories that are not reviewed during the entire year.

Secondly, we disregarded all the price spells that were affected by left-censoring. This selection was imposed by our empirical specification of the probability to observe a price change after T periods, which considers the accumulated inflation during those T periods as an explanatory variable. As the starting date of left-censored price spells is unknown, a simple measure of the accumulated inflation since the starting date of those price spells was not available. It should be pointed out that this selection does not seem to lead to a severe selection bias because of the repeated nature of the price spells (see Fougère et al. (2004) for a discussion on this issue).

Finally, we randomly drew 5 p.c. of the remaining 244,993 price trajectories. This last sample of 12,250 price trajectories or 95,656 price spells, associated with 558,579 price reports, is our main sample for this analysis. It has been used for the empirical analysis presented in section 4 and is referred to as the "Full CPI" sample, since the 5 main components of the CPI (unprocessed food, processed food, energy, non-energy industrial goods and services) are covered by this sample. In order to check whether our sample could be considered as representative of the whole data set, we conducted simple representativity tests. Chi-square tests confirm that the structure of our sample in terms of product categories[6] or in terms of the length of the price trajectories[7] was not significantly different from the one of the entire data set.

As our sample could be considered as representative, we used it to compute the non-parametric estimate of the probability to observe a price change conditional on the time period during which the price has remained constant and we compared it with the similar estimation based on the entire sample. Both estimations are presented in Figure 1.

This figure shows a strong decline in the hazard rates after 1 period. The large peak observed after 1 month is associated with both flexible pricing and commercial practices (temporary promotions). This is also related to the over-representation of short price spells in our data, as price trajectories associated with products characterised by price flexibility are made up of more price spells (see Dias, Robalo Marques and Santos Silva (2004) for a discussion on that issue).

In order to avoid such an over-representation of short spells in our sample, we might have used a different sampling strategy. Instead of the random drawing of 5 p.c. of the price trajectories observed, we might have randomly drawn one price spell per trajectory. This strategy would have reduced the over-representation of short spells and therefore the negative slope of the Kaplan-Meier estimates of the hazard function. However, by checking for unobserved heterogeneity in the estimations presented in section 4, this over-representation should no longer have a strong effect on the shape of our hazard functions. Moreover, the use of all the uncensored or right-censored spells of a particular price trajectory (rather than one randomly drawn spell per trajectory) can

6 Chi2 representativity test in terms of product category : Chi2-stat : 559.15 < Chi2(528, 0.05) : 582.56.
7 Chi2 representativity test in terms of duration : Chi2-stat : 129.43 < Chi2(143, 0.05) : 171.09.

improve our estimations by a better characterisation of the firm's pricing behaviour (for instance through the unobserved heterogeneity parameter).

The potential impact of the over-representation of short spells, which are mostly observed in unprocessed food and in oil products, on our results has also been reduced by the use of more homogeneous sub-samples. For this purpose, we considered the sub-sample of the price trajectories of product categories belonging to the core part of the CPI (excluding energy and unprocessed food) and the sub-samples associated with each of the 3 components of the core CPI (processed food, non-energy industrial goods and services).

At first sight, the declining hazard function observed in Figure 1 is counterintuitive, compared to what is implied by the theoretical pricing rules considered in macroeconomic models. Álvarez et al. (2004) present a characterisation of the hazard functions associated to standard time-dependent pricing rules, which are all non-decreasing. However, the decline in the aggregate hazard rates could reflect a composition effect in the sample of price spells. Álvarez et al. (2004) show indeed that mixing non declining hazards, such as those associated to Calvo or Taylor pricing rules, can produce a declining hazard at the aggregate level. A declining hazard function contrasts even more with state-dependent pricing rules, as these tend to produce increasing hazards, at least in steady state and with non-zero inflation, as is shown in Dotsey, King and Wolman (1999). It is argued later on that testing for state-dependence requires that the incentive to adjust prices stemming from inflation is explicitly taken into account, rather than just testing whether the probability to observe a price change increases in the course of the period that has elapsed since the previous price change. We further put that appropriately estimating the impact (sectoral) inflation has on the hazard requires that a distinction is made between price increases and price decreases, as inflation affects both phenomena in the opposite direction. Our estimation results tend to confirm this and allow us to conclude that the decreasing shape of the aggregate hazard could even be compatible with the existence of a statistically significant fraction of state-dependent pricing.

Another interesting element provided by Figure 1 is the fact that some truncated Calvo or Taylor phenomena seem to take place after 6, 12, 18 and 24 months, as local modes in the hazard function are observed for those specific horizons. This phenomenon is particularly pronounced at the 12 months duration.

Finally, Figure 1 also confirms the good representativity of our sub-sample, as the hazard functions estimated using our sample or the entire data set are very close, both in shape and magnitude. This has been confirmed by a Chi-square test[8].

8 Chi2-stat : 37.18 < Chi2(72, 0.05) : 92.80.

To summarise, the analysis of the non-parametric estimation of the hazard function associated with price changes clearly stresses the need to account for (unobserved) heterogeneity at the store-product level and to allow for truncations at specific durations when analysing price spells durations. Moreover, we will acknowledge the effect inflation has on the probability to change a price in an explicit way, rather than just testing whether the hazard increases during the time span that has elapsed since the previous price change. The baseline model presented in the following section is constructed in order to meet these requirements.

3. MODELLING THE PROBABILITY TO OBSERVE A PRICE CHANGE

In this section, we present alternative specifications of a Logit equation modelling the probability to observe a price change after T periods under several assumptions.

To that end, we define Y_{ijt} as a binary variable indicating that the price of the individual product sold by firm i that belongs to product category j is changed at the end of period t or:

$$Y_{ijt} = \begin{cases} 1 & \text{if } P_{ijt} \neq P_{ij,t+1} \\ 0 & \text{otherwise} \end{cases} \tag{1}$$

We first present the specifications associated with Calvo, truncated Calvo and state-dependent pricing rules à la Cecchetti. Finally, we present a general specification that combines both state- and time-dependent features. This general specification encompasses the simpler specifications as special cases.

3.1 The Calvo pricing rule

On the assumption that firms apply a Calvo pricing rule, the Logit representation of the probability that firm i changes the price of the product belonging to product category j at the end of period t, taking into account that its price was set T periods before, is provided by:

$$P\left[Y_{ijt} = 1\right] = \frac{\exp(\beta_0)}{1 + \exp(\beta_0)} \tag{2}$$

As the hazard rate is constant under a Calvo pricing rule, this probability does not depend on the number T of periods having elapsed since the price P_{ijt} was set. The smaller β_0 is, the less frequently prices change.

3.2 The Truncated Calvo pricing rule

Assuming that the probability to observe a price change is constant (as in Calvo) for the first horizons and that some truncation occurs after n_{ij} periods, the Logit representation of the probability that firm i changes the price of the product belonging to product category j at the end of period t, taking into account that its price was set T periods before, is provided by equation (3):

$$P\left[Y_{ijt} = 1\right] = \begin{cases} \dfrac{\exp(\beta_0)}{1+\exp(\beta_0)} & \text{if } T < n_{ij} \\ 1 & \text{if } T = n_{ij} \end{cases} \tag{3}$$

For this specification, the maximal length of a price spell is specific to the firm i and the product category j. However, equal maximal length for all firms and/or product categories can be imposed.

This truncated Calvo model can be seen as a general model that nests two well-known time-dependent pricing rules as special cases, namely the original Calvo rule and the so-called Taylor contracts[9]. Indeed, when n_{ij} tends to infinity, the truncation component of equation (3) will disappear, leading to the Calvo rule of equation (2). When β_0 tends to minus infinity, the first term which yields random durations will tend to zero and, hence, fixed durations of length n_{ij} are obtained, as is the case for Taylor contracts. The more negative β_0 is, the more the pricing rule tends to a pricing rule à la Taylor.

3.3 State-dependent pricing rule

Let us now consider firms that reset their prices using a state-dependent pricing rule. According to Cecchetti (1986), firm i that faces fixed price adjustment costs will change the price of the product belonging to product category j after T periods if the difference between its actual price P_{ijt} and its desired price P^*_{ijt} exceeds a certain threshold, or

$$P\left[Y_{ijt} = 1\right] = P\left[\log\left(\frac{P^*_{ijt}}{P_{ijt}}\right) \geq h^c_{ij}\right] \tag{4}$$

where h^c_{ij} is the largest difference between the log of the frictionless optimal price level and the log of the actual price level tolerated by firm i for the product belonging to product category j.

9 For more details, see Burriel et al. (2004).

Referring to Iwai (1981), Cecchetti (1986) has shown that this probability can be expressed in terms of (i) $\pi_{acc,t-T,t}$, the accumulated overall inflation recorded since the price level P_{ijt} was set; (ii) T, the time that has elapsed since the last price change; (iii) LDP_{ijt}, the relative size of the previous price change (in percentage points), and (iv) X_{ijt}, changes in demand conditions.

Under state-dependent price setting, the Logit representation of the probability that firm i changes the price of the product belonging to product category j at the end of period t, taking into account that its price was set T periods before, becomes:

$$P[Y_{ijt} = 1] = \frac{\exp(\beta_0 + \beta_1 \pi_{acc,t-T,t} + \beta_2 \log T + \beta_3 LDP_{ijt} + \beta_4 X_{ijt})}{1 + \exp(\beta_0 + \beta_1 \pi_{acc,t-T,t} + \beta_2 \log T + \beta_3 LDP_{ijt} + \beta_4 X_{ijt})} \quad (5)$$

In equation (5) the estimates of the different β_i will capture both the impact of the associated variable on the probability of a price change and the share of individual observations showing this particular behaviour. The estimation technique does not allow to identify both parameters (impact, properly speaking, and share) separately. This does not mean however that the estimated coefficients are not relevant. Particularly, when β_1, β_2, β_3 and β_4 are not significantly different from 0, all firms follow Calvo pricing rules[10]. Alternatively, having significant estimates for β_1, β_2, β_3 or β_4 could be interpreted as a rejection of the Calvo model (or a mixture of different Calvo models[11]) and would provide evidence in favour of the existence of (at least) a statistically significant fraction of firms adopting a more state-dependent behaviour. For a given estimate (for instance β_1), it is however not possible to uncover whether this is the result of a very strong impact of the relevant variable (accumulated inflation in the case of β_1) on the probability to change prices for a relatively small fraction of firms or, alternatively, results from a milder, but more widespread impact. We leave the separate identification of impacts and shares for further research and concentrate in this paper on the fact whether their interaction affects the probability to change prices in a significant manner.

3.4 The baseline models

In this section, we propose our baseline model that extends Cecchetti's specification in order to allow for some of the time-dependent features described in subsections 3.1 and 3.2. We also propose separate specifications for price increases and price decreases.

[10] We used the term Calvo *rules* instead of Calvo rule, because the estimated specification allows that the constant term of the equation is different for the five main analytical components of the CPI and is moreover firm i- product j -specific (see below for the introduction of observed and non-observed heterogeneity).

[11] See also previous footnote.

3.4.1 Explaining the probability of observing a price change

As mentioned above, our specification of the probability of observing a price change after T periods extends Cecchetti's specification in several ways.

First, on the basis of the results of the non-parametric estimation, we want to allow for truncated pricing rules at specific durations in the time-dependent part of our specification. Therefore, we added one set of binary variables indicating that the actual duration of a price spell is equal to some specific durations. If these variables have a significant impact on the probability to observe a price change, they indicate that some firms automatically reset their price when their price spells reach some specific length. Such a behaviour is compatible with either the truncated Calvo model or the Taylor model. We tested for truncations after 1 month[12], 6, 12, 18, 24 and 36 months. Moreover, a set of (monthly) seasonal dummies has been added to the time-dependent part of our equation.

As to the state-dependent part, we extended Cecchetti's specification through substituting accumulated aggregate inflation by accumulated inflation, in absolute value, measured at the product category j level. In order to take changes in overall inflation into account, we included a set of year dummies. This set of year dummies can also reflect changes in global economic conditions or the impact of other common shocks not captured elsewhere. Thus, while these dummies could be seen, by definition, as being part of the time-dependent component of our model, their ex post interpretation might point out that they capture the impact of other unobserved state-dependent variables.

Because we do not observe demand conditions at the firm level, our specification does not incorporate changes in demand as an explanatory factor for price changes. However, we have in our model another variable that could reveal state-dependent pricing. Indeed, we test for the price reactions of firms to VAT rate shocks by adding to Cecchetti's specification two binary variables reflecting respectively increases and decreases in the VAT rate at the product category j level.

We also test for the potential impact of commercial practices on price-setting behaviour by measuring the impact of psychological pricing and the impact of the size and the sign of the last price change observed. The particular questions we address here are: (i) does a firm that uses psychological or attractive prices change its price less frequently than other firms and (ii) does a firm change its price more rapidly after a price reduction and/or after a major price increase/decrease.

[12] Truncation after 1 month is included in the equation in order to take into account that some price spells are very short-lived (especially for perishables and oil products).

Finally, we also took into account observed heterogeneity by including 5 sectoral dummies in our specification. According to the Eurostat classification, we ranged each product category under one of the 5 main components of the CPI: unprocessed food, processed food, energy, non-energy industrial goods and services.

Taking all these considerations into account, our specification of the probability that firm i changes the price of the product belonging to product category j at the end of period t, considering that the price was set T periods before, is provided by:

$$P[Y_{ijt} = 1] = \frac{\exp(X_{ijt}\beta)}{1+\exp(X_{ijt}\beta)} \tag{6}$$

with

$$\begin{aligned} X_{ijt}\beta = \; & \beta_0 + \beta_1 \log T_{ijt} + \beta_2 DUR1_{ijt} + \beta_3 DUR6_{ijt} + \beta_4 DUR12_{ijt} + \\ & \beta_5 DUR18_{ijt} + \beta_6 DUR24_{ijt} + \beta_7 DUR36_{ijt} + \beta_8 \pi_{acc,j,t-T,t} + \\ & \beta_9 VATUP_{jt} + \beta_{10} VATDW_{jt} + \beta_{11} ATTRAC_{ijt} + \\ & \beta_{12} LDP_{ijt}(1 - LDPDW_{ijt}) + \beta_{13} LDP_{ijt} LDPDW_{ijt} + \\ & \beta_{14} LDPDW_{ijt} + \sum_{k=1990}^{2000} \alpha_k YEARk_t + \\ & \sum_{k=1}^{11} \gamma_k MONTHk_t + \sum_{k=1}^{4} \delta_k SECk_j \end{aligned}$$

where :

- $DURx_{ijt}$ is a binary variable that takes value 1 if T_{ijt} equals x ;
- $\pi_{acc,j,t-T,t}$ represents the product category j-specific accumulated inflation in time t since the last price change made by firm i, in absolute value ;
- $VATUP_{jt}$ is a binary variable that takes value 1 if product category j experienced a VAT rate increase in month t ;
- $VATDW_{jt}$ is a binary variable that takes value 1 if product category j experienced a VAT rate decrease in month t ;
- $ATTRAC_{ijt}$ is a binary variable that takes value 1 if the price level[13] at time t of the product belonging to product category j sold by firm i ends in 0, 5 or 9 ;
- LDP_{ijt} represents the absolute value of the size of the previous price change, expressed in percentage points, made by firm i for the product belonging to product category j ;
- $LDPDW_{ijt}$ is a binary variable that takes value 1 if the previous price change made by firm i for the product belonging to product category j was a price reduction ;
- YEARk are year dummies (baseline year is 1989) ;
- MONTHk are seasonal dummies (baseline month is December) ;
- SECk are sectoral dummies (baseline sector is Services).

[13] The prices are expressed in Belgian francs.

Finally, even if the proposed specification already incorporates some elements of observed heterogeneity across firms (the sectoral dummies, the attractive price dummy), we also want to account for potentially important unobserved heterogeneity. As pointed out above, Aucremanne and Dhyne (2004) stressed that a high degree of the heterogeneity in observed price durations was observed within product categories. Dealing with unobserved heterogeneity across firms and product categories might be a key factor for obtaining non-declining hazard functions as mentioned in section 2. Therefore, we used a mixed Logit model (Train (2003)) or logistic normal regression (Allenby and Lenk (1994), Kuss (2002)) in order to take into account this characteristic.

We assume that the constant term of equation (6) can be specified as $\beta_0 + u_{ij}$, a constant term plus a random term associated with the specific firm i that sells a product belonging to the product category j. On that assumption, equation (6) becomes:

$$P\left[Y_{ijt} = 1\right] = \frac{\exp\left(X_{ijt}\beta + u_{ij}\right)}{1 + \exp\left(X_{ijt}\beta + u_{ij}\right)} \tag{7}$$

This random term u_{ij}, which is firm i - product category j specific, is assumed to be distributed as:

$$u_{ij} \approx N\left(\mu_1 1STDUR_{ij} + \mu_2 1STDP_{ij} \,;\quad \sum_{k=1}^{5} \sigma_k^2 SECk_j\right)$$

where $1STDUR_{ij}$ and $1STDP_{ij}$ represent respectively the duration of the first uncensored price spell of the price trajectory of the product belonging to product category j that is sold by firm i and the magnitude, in absolute value, of the first price change observed for that price trajectory.

This specification characterises the degree of price stickiness at the firm i - product category j level, $\beta_0 + u_{ij}$, by the characteristics of the first uncensored price spell of a particular price trajectory. If the first uncensored price spell observed is short, it is indeed more likely that we observe a firm that changes its price frequently, while the opposite holds in cases where the first uncensored price spell is long. This unobserved heterogeneity could be the result of differences in price adjustment costs and differences in the exposure to shocks, both across product categories and across firms. The proposed specification also introduces potential sectoral heteroscedasticity.

This choice is inspired by the methodological framework developed by Wooldridge (2003) for the treatment of initial conditions in dynamic non-linear unobserved effect models. Some details can be found in the Technical appendix. Empirical alternatives to that procedure could be those proposed by Heckman and Singer (1984) or by Honoré and Hu (2004).

3.4.2 Discriminating between price increases and price decreases

The models presented in the previous sub-section relate to price changes, without discriminating between price increases and price decreases. In this sub-section, we present two separated specifications for both events.

A decreasing hazard function associated with price changes might be the result of neglected heterogeneity, but not discriminating between price increases and price decreases might also play a role. It has been stated earlier that a declining hazard function associated with price changes is a counterintuitive result. This is due to the fact that, in a world characterised by positive inflation, it is hard to believe that the firm's incentives to change its prices decrease in time. However, this argument is clearly only relevant for price increases. It is not the case for price decreases, as accumulating positive inflation over time reduces the incentive to decrease prices. Therefore, one could obtain a decreasing hazard function for price changes as a result of the mix of a non-decreasing hazard function for price increases and a decreasing hazard function for price decreases. This effect is potentially important, as it was shown in Aucremanne and Dhyne (2004) that price decreases occur relatively frequently, except for services. This argument favours a separate analysis of price increases and decreases.

Let us define Y_{1ijt} and Y_{2ijt}, two binary variables that respectively indicate whether or not firm i increases / decreases the price of the product belonging to product category j at the end of period t.

$$Y_{1ijt} = \begin{cases} 1 & \text{if } P_{ijt} < P_{ij,t+1} \\ 0 & \text{otherwise} \end{cases} \tag{8}$$

$$Y_{2ijt} = \begin{cases} 1 & \text{if } P_{ijt} > P_{ij,t+1} \\ 0 & \text{otherwise} \end{cases} \tag{9}$$

We define the probability to observe a price increase after T periods as

$$P\left[Y_{1ijt} = 1\right] = \frac{\exp\left(X_{ijt}\beta_1 + u_{1ij}\right)}{1 + \exp\left(X_{ijt}\beta_1 + u_{1ij}\right)} \tag{10}$$

$$
\begin{aligned}
X_{ijt}\beta_1 = \; & \beta_{10} + \beta_{11}\log T_{ijt} + \beta_{12}DUR1_{ijt} + \beta_{13}DUR6_{ijt} + \beta_{14}DUR12_{ijt} + \\
& \beta_{15}DUR18_{ijt} + \beta_{16}DUR24_{ijt} + \beta_{17}DUR36_{ijt} + \beta_{18}^{up}\pi_{acc,j,t-T,t}^{up} + \\
& \beta_{18}^{dw}\pi_{acc,j,t-T,t}^{dw} + \beta_{19}VATUP_{jt} + \beta_{110}VATDW_{jt} + \beta_{111}ATTRAC_{ijt} + \\
\text{with} \quad & \beta_{112}LDP_{ijt}\left(1 - LDPDW_{ijt}\right) + \beta_{113}LDP_{ijt}LDPDW_{ijt} + \\
& \beta_{114}LDPDW_{ijt} + \sum_{k=1990}^{2000}\alpha_{1k}YEARk_t + \\
& \sum_{k=1}^{11}\gamma_{1k}MONTHk_t + \sum_{k=1}^{4}\delta_{1k}SECk_j
\end{aligned}
$$

and $$u_{1ij} \approx N\left(\mu_{11}1STDUR_{ij} + \mu_{12}1STDP_{ij}\,;\quad \sum_{k=1}^{5}\sigma_{1k}^2 SECk_j\right)$$

where :

- $\pi_{acc,j,t-T,t}^{up}$ represents the product category j-specific accumulated inflation in time t since the last price change made by firm i, when accumulated inflation is positive ;
- $\pi_{acc,j,t-T,t}^{dw}$ represents the product category j-specific accumulated inflation, in absolute terms, in time t since the last price change made by firm i, when accumulated inflation is negative.

Similarly, we define the probability to observe a price decrease after T periods as

$$P\left[Y_{2ijt} = 1\right] = \frac{\exp\left(X_{ijt}\beta_2 + u_{2ij}\right)}{1 + \exp\left(X_{ijt}\beta_2 + u_{2ij}\right)} \tag{11}$$

$$
\begin{aligned}
X_{ijt}\beta_2 = \; & \beta_{20} + \beta_{21}\log T_{ijt} + \beta_{22}DUR1_{ijt} + \beta_{23}DUR6_{ijt} + \beta_{24}DUR12_{ijt} + \\
& \beta_{25}DUR18_{ijt} + \beta_{26}DUR24_{ijt} + \beta_{27}DUR36_{ijt} + \beta_{28}^{up}\pi_{acc,j,t-T,t}^{up} + \\
& \beta_{28}^{dw}\pi_{acc,j,t-T,t}^{dw} + \beta_{29}VATUP_{jt} + \beta_{210}VATDW_{jt} + \beta_{211}ATTRAC_{ijt} + \\
\text{with} \quad & \beta_{212}LDP_{ijt}\left(1 - LDPDW_{ijt}\right) + \beta_{213}LDP_{ijt}LDPDW_{ijt} + \\
& \beta_{214}LDPDW_{ijt} + \sum_{k=1990}^{2000}\alpha_{2k}YEARk_t + \\
& \sum_{k=1}^{11}\gamma_{2k}MONTHk_t + \sum_{k=1}^{4}\delta_{2k}SECk_j
\end{aligned}
$$

and $$u_{2ij} \approx N\left(\mu_{21}1STDUR_{ij} + \mu_{22}1STDP_{ij}\,;\quad \sum_{k=1}^{5}\sigma_{2k}^2 SECk_j\right)$$

These two equations are estimated separately, neglecting potential correlation between u_{1ij} and u_{2ij}.

Positive and negative accumulated inflation are introduced separately in both equations in order to allow for potential asymmetric reactions of the firms. For instance, it might be the case that the incentive to change prices stemming from positive accumulated inflation is stronger than the disincentive to do so stemming from negative accumulated inflation. Similar arguments in favour of asymmetric reactions of firms motivate the discrimination between VAT rate increases and VAT rate decreases or the differentiated analysis of the impact of the relative size of the previous price change.

3.5 Analysing the truncation phenomena

Equations (7), (10) and (11) incorporate variables indicating that truncation occurs after some specific durations. However, such a specification does not allow us to disentangle a regular (or "Taylor contract") behaviour from an irregular (or "truncated Calvo") behaviour. Therefore, at a second stage, we estimate an alternative specification of our three baseline models. In this specification, we include a set of binary variables indicating that the actual duration is not only equal to a specific duration but is also equal to the duration of the previous price spell associated with the same individual product. The combined analysis of this set of binary variables with the set of binary variables that indicates truncation at specific durations allows us to make a distinction between the truncated Calvo model, on the one hand, and Taylor contracts, on the other hand. To that end, we test whether one of the two sets of binary variables is not jointly significant. We test for Taylor contracts of one, 6 and 12 months, of non-standard durations up to 1 year and of durations beyond 1 year.

Under this alternative specification, equation (7) is provided by:

$$P[Y_{ijt} = 1] = \frac{\exp(X_{ijt}\beta + u_{ij})}{1 + \exp(X_{ijt}\beta + u_{ij})} \tag{12}$$

with :

$$\begin{aligned} X_{ijt}\beta = \; & \beta_0 + \beta_1 \log T + \beta_2 DUR1_{ijt} + \beta_3 DUR6_{ijt} + \beta_4 DUR12_{ijt} + \\ & \beta_5 DUR18_{ijt} + \beta_6 DUR24_{ijt} + \beta_7 DUR36_{ijt} + \beta_8 \pi_{acc,j,t-T,t} + \\ & \beta_9 VATUP_{jt} + \beta_{10} VATDW_{jt} + \beta_{11} ATTRAC_{ijt} + \\ & \beta_{12} LDP_{ijt}(1 - LDPDW_{ijt}) + \beta_{13} LDP_{ijt} LDPDW_{ijt} + \\ & \beta_{14} LDPDW_{ijt} + \beta_{15} TAYLORNS_{ijt} + \beta_{16} TAYLOR1_{ijt} + \\ & \beta_{17} TAYLOR6_{ijt} + \beta_{18} TAYLOR12_{ijt} + \\ & \beta_{19} TAYLORUP12_{ijt} + \sum_{k=1990}^{2000} \alpha_k YEARk_t + \\ & \sum_{k=1}^{11} \gamma_k MONTHk_t + \sum_{k=1}^{4} \delta_k SECk_j \end{aligned}$$

where :

- $TAYLORNS_{ijt}$ is a binary variable that takes value 1 if the duration of the actual price spell in period t equals the duration of the previous price spell observed for the product belonging to product category j sold by firm i and if the duration of that previous price spell is less than 12 months and not equal to 1 or 6 month(s);
- $TAYLORx_{ijt}$ is a binary variable that takes value 1 if both the duration of the actual price spell in period t and of the previous price spell observed for the product belonging to product category j sold by firm i equal x months, with x having the values 1, 6 or 12;
- $TAYLORUP12_{ijt}$ is a binary variable that takes value 1 if the duration of the actual price spell in period t equals the duration of the previous price spell observed for the product belonging to product category j sold by firm i and if the duration of that previous price spell exceeds 12 months.

Similar expressions can be derived for the price increases and the price decreases equations.

4. ECONOMETRIC RESULTS

4.1 The baseline models

In this section, we present the econometric results associated with the estimation of equations (7), (10) and (11) based on the selected sample. The results based on this sample are shown in Table 1 and are referred to as "Full CPI" results because they are based on product categories that belong to the 5 main components of the CPI.

In order to document the impact of heterogeneity on the shape of the hazard function, we also estimate our 3 equations for different more homogenous sub-samples. We first confine ourselves to product categories belonging to the core part of the CPI (in other words, disregarding energy and unprocessed food). These results are referred to as "Core components" and are presented in Table 2. Finally, we also estimate equations (7), (10) and (11) for each of the main core components separately. These results are referred to as "Processed food", "Non-energy industrial goods" and "Services". They are summarised in Tables 3, 4 and 5 respectively. We did not estimate a specific equation for unprocessed food and energy.

As is well-known, the analysis of estimated results of non linear equations is not straightforward. In order to evaluate the impact of one particular variable on the probability to observe a price change /

increase / decrease, one has to compute either the impact of the increase of this variable evaluated at the sample mean or the risk ratio associated with that variable[14].

Traditionally, the impact at the sample mean has been used especially to evaluate the impact of the change of continuous variables on the probability of interest. However, when unobserved heterogeneity is introduced in the equation, one has to integrate the impact of one explanatory variable on the distribution of the unobserved heterogeneity parameter in order to evaluate correctly the impact of the change of that variable on the probability to observe a price change / increase / decrease. For mixed logit models, the computation of this impact becomes analytically cumbersome. Moreover, most of our variables are dichotomous. Therefore, we decided to use the risk ratio in the following sections. The risk ratio is also frequently used in duration analysis (Allison (1995)).

A first result that has to be put forward before analysing some particular aspects of our regressions is that our estimates clearly reject simple pure time-dependent representations of the price-setting behaviour such as pure Calvo, truncated Calvo or Taylor contracts or combinations of this type of models. Indeed, some of our state-dependent variables (such as accumulated inflation, VAT rate changes, year dummies) always affect the estimated probability to observe a price change / increase / decrease in a statistically significant way.

4.1.1 The level of the hazard function

Focusing on the estimation based on the "Full CPI" sample, it should be pointed out that the results, in terms of the relative degree of price stickiness across the 5 main CPI components, are in line with the basic facts that have been put forward in Aucremanne and Dhyne (2004).

Indeed, our results indicate that energy or more precisely oil products, constitute the product category showing the most flexible pricing behaviour, as their representative firm is characterised by the highest constant term (0.1309[15]) in Table 1. Energy is followed, in descending order, by unprocessed food (-1.1775[15]), processed food (-1.8748[15]), non-energy industrial goods (-2.6020[15]) and finally services (-3.0291[15]).

[14] The risk ratio is given by the exponential of the estimated coefficient.

[15] These constants are obtained by adding the coefficients of the binary variables for each component and the expected value of u_{ij} for the representative firm to the constant term of the "Full CPI" equation. The representative firm's expected value of u_{ij} is obtained considering an individual firm characterised by a duration of the first price spell and a size of the first price change equal to the average of these two variables computed for each component.

Considering the representative firm for each of the 5 main components, these constant terms would imply frequencies of price changes[16] of 53.3 p.c. for oil products, 23.6 p.c. for unprocessed food, 13.3 p.c. for processed food, 6.9 p.c. for non-energy industrial goods and 4.6 p.c. for services, if the other regressors were set equal to 0. The absolute level of these probabilities is not very meaningful per se, as they are computed irrespective of the impact of duration or accumulated inflation and they precisely refer to a representative firm characterised by a very specific description of its first price spell[17]. Nonetheless, they provide some reliable insights on the relative importance of price changes across the 5 main CPI components.

This sectoral comparison is made on the basis of the "Full CPI" results and not on the basis of estimates using more restricted samples, because the comparison of the estimated constant terms across samples is affected by changes in the other parameters. For instance, the constant will be affected by differences across samples in the role of the yearly dummies or the seasonal patterns. It should be noted, however, that an identical relative position of the three core components (in descending order: processed food, non-energy industrial goods and services) is found in the "Core components" sub-sample.

Our "Full CPI" results not only confirm the sectoral ranking presented in Aucremanne and Dhyne (2004), but also find that price decreases are not uncommon, except for services. Indeed, the constant term associated to the representative firm in the service sector using the "Full CPI" sample is estimated to be equal to -3.0052 (implied probability of 4.7 p.c. if the other regressors were set equal to 0) for the price increases equation and -6.2230 (implied probability of 0.2 p.c. if the other regressors were set equal to 0) for the price decreases equation. Based on these estimated probabilities, and, all other things being equal, the representative firm in the service sector was characterised in December 1989 by a probability to observe a price decrease which was much smaller than the probability to observe a price increase. This difference is much smaller for the other types of goods. The implied probabilities, if the other regressors were set equal to 0, of a price increase and a price decrease respectively amount to 36.2 p.c. and 36.7 p.c. for energy, 14.5 p.c. and 14.5 p.c. for unprocessed food, 8.8 p.c. and 6.0 p.c. for processed food, and 4.9 p.c. and 2.2 p.c. for non-energy industrial goods.

As mentioned above, these probabilities are not very meaningful per se, as they can be affected by differences in the seasonal or trend patterns associated with price increases or price decreases. For instance, price increases could be concentrated in December while price decreases could be concentrated in another month, or 1989 might be a period during which price decreases were particularly rare because of specific economic conditions. However, by focusing on the full CPI

16 These probabilities refer to the estimated probability of price changes, in December 1989, assuming all variables equal to 0, except the constant, the sectoral dummies and the expected value of u_{ij} for the representative firm using equation (7).

17 The characteristics of the first price spell for the representative firm are provided by the average values of the two determinants of the expected value of u_{ij}, computed for each component.

results, for which the seasonal and trend patterns are assumed to be the same for all the components, such an effect would affect the 5 main CPI components to the same extent. This implies that the relative position of the 5 main CPI components in terms of the difference in the probability to observe a price increase relative to a price decrease is meaningful.

The rarity of price decreases observed in services is not sufficient to conclude that specific downward rigidities are at work for these products. As is stressed below (see section 4.1.2.), our estimates seem to indicate that firms operating in the service sector can rapidly decrease their prices if they are hit by major negative price shocks.

4.1.2 The shape of the hazard functions

We now focus on what our results imply for the shape of the hazard function associated with price changes / increases / decreases.

Our estimation results first of all stress that taking (observed and unobserved) heterogeneity into account seriously reduces the downward sloping nature of the hazard functions, compared to the non-parametric estimate presented in Figure 1 and allows, under certain conditions, to produce increasing hazards. The latter result is obtained despite the fact that the direct impact of duration is always estimated to be negative, but clearly less so than in estimations which did not control for heterogeneity (not reported here).

It is important in this respect to stress that the apparent direct impact of duration on the probability to change prices in state-dependent pricing models, such as for instance Dotsey, King and Wolman (1999) (hereafter DKW), is obtained as a result of the role of accumulated *steady state* inflation. In case of steady state inflation there is indeed a simple, proportional relationship between duration, on the one hand, and accumulated inflation, on the other hand. Actually, inflation hovers however around its steady state level and this can disrupt the simple relationship between duration and actual accumulated inflation. That is why we analyse the impact of duration and the impact of accumulated inflation jointly in our discussion of the slope of the hazard function. Regression results measure the overall effect[18] of actual (as opposed to steady state) accumulated inflation, specific to each product category. The hazards discussed below are however obtained by applying the estimated results to constant product-specific inflation rates (referred to as sectoral trend inflation). Thus, we wanted to report results that are comparable with what is shown for theoretical models. It should however be noted that estimation results are obtained for product-specific or sectoral inflation, whereas DKW show graphs for different rates of aggregate steady state inflation, as they do not allow for sectoral shocks in their model.

[18] As explained before, the overall effect shown by the estimated parameters is the combined result of the impact of accumulated inflation and the share of firms this impact is relevant to.

In view of a greater comparability with DKW, hazards up to a 3 year horizon (36 months) are considered. For the same reason, the constant product-specific inflation rates that we consider - respectively 1, 2, and 5 p.c. positive and negative inflation - are in the range considered by DKW (namely 2.5, 5 and 10 p.c.). Moreover they are compatible with the (sectoral) inflation rates observed in Belgium over the sample period, as is shown in Figure 2. Indeed, over the entire sample (1989-2001) average sectoral inflation exceeded 1 p.c. in absolute value for 77 p.c. of the product categories and 2 p.c. for nearly half of them. Moreover, even in the more recent period characterised by lower inflation (1996-2001)[19], 62.3 p.c. of the CPI was characterised by an average sectoral inflation rate over 1 p.c. in absolute value, while 33 p.c. of the CPI sectoral inflation exceeded 2 p.c. in absolute value. In both periods a 5 p.c. sectoral average inflation was even observed for 6 p.c. of the product categories in the CPI.

For these sectoral inflation rates which are compatible with the current low inflation regime, our estimation results for the price changes equation are, in general, compatible with mildly upward sloping hazards, as is shown in Figures 3 to 7[20]. This is particularly true when we restrict our analysis to the "Core components" sample of the CPI (Figure 4) and the three sub-samples of the core CPI (Figures 5-7). We consider this as evidence in favour of state-dependent pricing, whereby the relevant "state" refers to economic conditions at the level of the product category, i.e. at a relatively disaggregated level. Sectoral and idiosyncratic shocks are therefore determinants of the relevant state. This does not mean however that aggregate inflation is not relevant at all, as changes in the steady state level of aggregate inflation would shift the entire distribution of sectoral inflation rates and affect the probabilities to change prices through this channel.

As has already been mentioned, Figures 4 to 7 suggest that using "more homogeneous" samples enhances the upward sloping character of the hazard for price changes. This is the result of a stronger effect of accumulated inflation. Indeed, the estimated risk ratios reveal that accumulating 1 percentage point additional sectoral inflation increases the hazard of a price change by 6.4 p.c., when considering the "Full CPI" sample. This effect should be interpreted with caution, as it means that the hazard of a price change is multiplied by a factor 1.064 for each additional percentage point of accumulated inflation. This implies for instance that when the conditional probability to observe a price change after T periods amounts to 10 p.c. under "normal" conditions, it will increase to 10.64 p.c. as a result of a 1 percentage point additional accumulated inflation. This effect on the hazard is more substantial or the "Core components" sample (15.4 p.c.) and reaches 20.7 p.c. for the processed food category.

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19 During the 1989-2001 period, average annual CPI inflation stood at 2.3 p.c., while it amounted to 1.7 p.c. during the 1996-2001 period.

20 What is called hazard in Figures 3 to 7 refers to the probability to observe a price change / increase / decrease after T periods at the end of month t conditional on the fact that no price changes were observed before the end of month t-1.

Figures 3 to 7 also show that the increasing hazard functions associated with price changes in case of positive (negative) sectoral trend inflation are the result of upward sloping hazard functions for price increases (decreases) and downward sloping hazard functions for price decreases (increases). As mentioned in section 3, observing decreasing hazards for price decreases (increases) in the case of positive (negative) sectoral inflation is not a counterintuitive result, as the firm's incentives to decrease (increase) its prices are indeed reduced in such a situation. Once this differentiated impact on respectively price increases and price decreases is acknowledged, the impact of accumulated inflation on the slope of the hazards is further magnified. In this respect, it is interesting to focus first on the impact of positive inflation on the probability to observe a price increase, as it is the one that is closest to the situation analysed in DKW.

The risk ratios associated with accumulated (positive) inflation are indeed more pronounced for the price increases equation than those discussed above for the price changes equation. Focusing on the price increases equation, accumulating 1 percentage point additional positive inflation increases the hazard of a price soar by 10.6 p.c. for the "Full CPI" sample. For the "Core components" sample it amounts to 23.3 p.c., while this effect ranges from 11.9 p.c. for services to 35.8 p.c. for processed food. The effect of accumulated inflation on the probability to observe a price increase is moreover asymmetric, as accumulating 1 percentage point negative inflation has a far milder impact. For the "Core components" sample it reduces the hazard of a price increase by 7.8 p.c. This impact amounts to 14.8 p.c. for processed food, while more negative inflation has no significant effect on the hazard of a price increase for the two other components of the core CPI (non-energy industrial goods and services).

The impact of negative inflation on the probability to observe a price reduction is comparable to the impact of positive inflation on the probability of a price increase. Accumulating 1 percentage point additional negative inflation increases the hazard of a price decrease by 8.9 p.c. for the "Full CPI". Here too, the effect is larger for the "Core components" sample, for which it amounts to 25.6 p.c. For the processed food category, the impact of a 1 percentage point additional negative inflation is 30.9 p.c. and for services it exceeds 200 p.c. The particularly large impact of negative accumulated inflation on the probability to observe a price decrease in the services sector has to be considered cautiously as price decreases and negative sectoral inflation are relatively uncommon phenomena in the service sector. Moreover, this large relative impact affects a very low probability to observe a price decrease in the service sector. Nevertheless, the estimated impact is statistically significant and implies that the rarity of price decreases for services is not necessarily symptomatic of specific downward rigidities. The effect of accumulated inflation on the probability to observe a price decrease is also asymmetric, as accumulating a 1 percentage point positive inflation has a far milder impact. For the "Core components" sample it reduces the hazard of a price increase by 2.2 p.c. This impact amounts to 9.5 p.c. for processed food, while more negative inflation has no significant effect on the hazard of a price increase for the two other components of the core CPI (non-energy industrial goods and services).

To summarise the impact of accumulated inflation on the shape of the hazard functions, our results first of all stress the fact that accumulated inflation plays a role in the characterisation of the hazard functions. This suggests that at least a significant (albeit unknown) share of price changes are motivated by state-dependent considerations. Even for relatively low levels of sectoral inflation, we obtain upward sloping hazard functions, when controlling for unobserved heterogeneity. Secondly, the impact of accumulated inflation is much larger when focusing on more homogeneous components of the CPI. Thirdly, when price increases are disentangled from price decreases, it is revealed that the (upward) slope of the hazard function for price changes is flattened by the fact that the hazard function of price decreases (increases) is being downward sloping in the case of positive (negative) sectoral inflation.

Figures 3 to 7 also reveal the existence of truncation phenomena at specific durations. Based on these graphs, truncation seems to occur particularly after 1 month and 12 or 24 months. Less pronounced spikes are observed after 6, 18 and 36 months and the estimated coefficients associated with truncation at 18 and 36 months are rarely significant.

As to the spike at 1 month, the coefficient associated with the DUR1 variable is smaller in the "Core components" sample than in the "Full CPI" sample. This indicates that the large spike at 1 month for the "Full CPI" is mostly associated with the very flexible components of the CPI, namely energy and unprocessed food. In the case of services, the spike at 1 month becomes even negative.

As far as the DUR6, DUR12 and DUR24 variables are concerned, their respective risk ratios are larger in the "Core components" sample than in the "Full CPI" sample. This suggests that these truncation patterns mostly occur in the stickier components of the CPI, particularly for non-energy industrial goods and for services. For the latter category, a particularly high risk ratio is found at the 12 months horizon.

A more detailed analysis of these truncations is presented in section 4.2.

4.1.3 The impact of the other variables

We now focus on the analysis of the remaining explanatory variables.

- **VAT rate changes**

As is shown in Tables 1 to 5, VAT rate changes seem to influence the price-setting behaviour of firms in a significant way. Both VAT rate increases or decreases positively affect the hazard associated with price increases and price decreases. However, VAT rate decreases seem to have the strongest impact, especially on the probability to observe a price decrease. This result may be partly related to the fact that the VAT rate cuts observed during the 1989-2001 period were in general larger in absolute value than the VAT rate increases.

Concerning VAT rate decreases, the impact of that variable on the probability to observe a price increase is not significant when using the more homogeneous samples. No impact of VAT rate cuts on price increases in the service sector has been estimated because such an event was not observed in our sample.

- **Attractive prices**

In this sub-section, we analyse the impact of the use of attractive prices. Their use slightly reduces the probability to change prices, both upward and downward, and the effect is somewhat more pronounced for price decreases. Except for services, the risk ratio associated with the change of an attractive price is at most 10 p.c. smaller than the one associated with the change of a non-attractive price. For services a larger impact is observed. In that case, the risk ratio is up to 35 p.c. lower when the price is set at an attractive level. This evidence suggests that, apart from services, the existence of pricing thresholds is not an important source of price stickiness, despite the widespread use of this type of pricing for consumer goods. This finding is confirmed by the results of an ad hoc survey on pricing behaviour in Belgium (see Aucremanne and Druant (2004)).

- **The relative size of the previous price change**

The impact of the size of the previous price change (in percentage points) on the probability to observe a price change/increase/decrease can give rise to two alternative interpretations. According to equation (5) and following the state-dependent framework of Cecchetti (1986), it provides an indication of the degree of price rigidity of the firm. Indeed, it may be assumed that a large price change at the end of the previous price spell indicates that the firm faces large price adjustment costs and, therefore, changes prices infrequently (by large amounts). On the contrary, a small price change at the end of the previous price spell could reveal that the firm faces small price adjustment costs and, therefore, re-sets prices frequently (by small amounts). This interpretation gives rise to a negative correlation between the size of the previous price change and the probability to observe a new one and particularly so for the probability to observe a new price change with the same sign.

Alternatively, the size of the previous price change can capture the fact that particularly large price changes are quickly followed by new price changes, either as a result of a sequence of very volatile shocks or as a result of the impact of (temporary) promotions. In the latter case, a large price decrease at the end of the previous price spell should reduce the duration of the current price spell and have a positive impact on the probability to observe a price increase and reduce the probability to observe a new price reduction.

In order to be able to disentangle both types of effects, the estimated models allow that the impact of the size of the previous price change depends on its sign. Moreover, the models include a dummy variable indicating that the previous price change was negative. Given the differentiated impact of the size on the probability of a price increase and on the probability of a price decrease, we only comment the results for equation (10) and equation (11) and ignore the price change equation (7). Notwithstanding this, it is not always possible to make a clear distinction between both types of effects.

There is, however, unambiguous evidence in favour of the second type of effect, particularly in the price increase models and this evidence is found consistently across all samples considered. In all samples, the mere fact that the previous price change was negative considerably increases the probability to observe a price increase. Such a sign reversal of two consecutive price changes is typical in the case of temporary promotions. Moreover, except for services, this impact is further enhanced by the size of the previous (negative) price change: the larger the price reduction, the more likely it is that it will be followed by a price increase. Large price reductions are also a typical phenomenon associated with promotions. A similar effect is systematically found in the price decrease models, except for services. Indeed, our estimates also indicate that a large price increase at the end of the previous price spell increases the probability of a price reduction. This result tends to suggest that observing a large price increase at the end of the previous price spell indicates that we potentially observe a firm that applies temporary promotions. Therefore, observing a large price increase at the end of the previous price spell might favour the observation of a subsequent price reduction.

As to the likelihood of a price decrease, it is reduced by the fact that the previous price change was negative. This is again compatible with the temporary promotions interpretation, as price decreases apparently do not tend to be repeated. It should however be noted that the estimated coefficient of this dummy is positive (but non-significant) in the price decrease equation for non-energy industrial goods. This seems to suggest that the (negative) impact of temporary promotions on this coefficient has been compensated by a positive impact for some product categories for which sequences of price decreases are observed. The latter might be compatible with downward trends in the prices of particular goods (for instance electronic devices).

The negative impact of the size of a negative price change on the probability to observe a new price decrease, could further strengthen the temporary promotions interpretation. A large price decrease could indeed reveal that a promotion occurred and therefore reduces the probability of a new decrease. However, in principle, this effect could also be compatible with the price rigidity interpretation. According to the latter interpretation, a large price decrease postpones the next price decrease (in a sequence of price decreases). The fact that the estimated coefficients for the price decrease dummies in equations (9) and (10) suggest that sequences of price decreases are, in general, not very common tends however to favour the temporary promotions interpretation.

As to the so-called rigidity interpretation, it is consistently found in the price increase models. Indeed, when the last price change is positive, its size considerably reduces the probability to observe a (new) price increase. Not surprisingly, this effect is more pronounced in the stickier "Core components" sample than in the "Full CPI" sample. It is particularly striking in the processed food category, while it turns out to be non-significant for services. For the price decrease models a similar, but generally somewhat less pronounced effect is found for the different core samples. When the last price change is negative, its size reduces the probability to observe a (new) price decrease. As has been pointed out before, the latter seems however to be more in line with the temporary promotions interpretation than with the rigidity interpretation.

- **The year dummies**

The estimated coefficients of the year dummies, which are shown in Figure 8, corroborate our interpretation that these dummies can reflect, to a considerable extent, the impact of the evolution of aggregate inflation on price setting.

This figure indeed suggests that the reduction in the aggregate inflation rate during the period under review coincides with a decrease of the probability to observe a price change, particularly for the non-energy industrial goods and services components. The increase in the inflation rate observed in 2000, which was due to negative supply shocks affecting the price level for unprocessed food and energy, did not lead to an increase in the frequency of price adjustment in 2000. This was also found in Aucremanne and Dhyne (2004). This might suggest that price-setters had already incorporated the low-inflation environment in EMU. Therefore, their frequency of adjustment did not react to what was considered as a temporary increase in aggregate inflation. In line with this interpretation, it is found that the values of these dummies for the core components are more consistent with core inflation. This is particularly true for the non-energy industrial goods. The link with inflation is less clear for processed food. The figure also illustrates that the temporal effect on the frequency of price changes is mainly driven by the trend for price increases, especially for the two stickier components of core inflation.

- **The seasonal patterns**

The seasonal patterns presented in Figure 9 are compatible with those in Aucremanne and Dhyne (2004). However, compared to the seasonal patterns in our previous analysis, these graphs lead by one month. Indeed, though we found relatively few price changes in July and August in our first paper, Figure 9 indicates that price changes rarely occur at the end of June and July. This difference is due to the fact that in Aucremanne and Dhyne (2004) a price change was recorded at the beginning of a price spell, while it is recorded at the end of the price spell in this paper.

These graphs also indicate that price changes in the service sector mostly occur at the end of December (or that new prices are recorded in January). Combined with the relative importance of the truncation observed after 12 months in that sector, this seasonal pattern tends to indicate that the assumption that firms follow price-setting rules à la Taylor in the service sector is supported by the data (see also below for more details in this respect).

Finally it should be pointed out that the seasonal patterns of price increases and price decreases are relatively well synchronised. This result is related to the fact that our price reports do not consider end-of-season sales, in which case one would have seen important seasonal effects for the months of December (new sales price in January) and in June (new sales price in July).

- **The distribution of the heterogeneity parameter**

Considering the parameters characterising the distribution of the heterogeneity parameter in each equation, we have to mention that they are in line with our priors. Indeed, based on our estimates, a firm-product combination, which is characterised by a relatively long first uncensored price spell will change its price less frequently than a firm-product combination which is characterised by a short first uncensored price spell. All other things being equal, firms that used to change their price (in)frequently always do so. The size of the first price change observed is of less help in characterising the degree of price stickiness of a firm. Finally, the assumption of heteroscedasticity conditional on the sectoral belonging of the product category j is supported by our findings.

4.2 Analysing the truncation phenomenon

As mentioned in section 3.2, the baseline specification does not allow to distinguish between price spells truncated after a specific duration because the firm changes its price on a regular basis (à la Taylor) and irregular truncation (à la truncated Calvo). To that end, the alternative specification presented in section 3.2 has been estimated. The results are summarised in Tables 6 to 10. In these equations the truncation phenomenon is split between price changes made on a regular basis (the "Taylor contract" variables) and others (the "Dur" variables).

The introduction of this new set of variables had only a marginal impact on the likelihood and on the estimated coefficients of most explanatory variables in the price changes / increases / decreases equations. However, and for obvious reasons, it clearly affected the estimated coefficients of the different truncation dummies and the seasonal dummies. This suggests that regularity in truncation - i.e. Taylor pricing - is potentially an important feature of the price-setting process. It also confirms our interpretation in the previous section that linked the pronounced seasonality for services with the occurrence of Taylor pricing for these products, seasonal effects for services being less important once Taylor pricing is explicitly modelled in the alternative specification.

Based on the regression results in Tables 6 to 10, it can be seen that Taylor pricing for durations beyond 12 months are often not significant, indicating that if there are regularities, they are predominantly of a periodicity of up to 1 year. Regularities of non-standard periodicity under 1 year are often estimated significantly, but the increase in the probability to observe a price change / increase / decrease associated with this type of regularity is not very important. Risk ratios are in general relatively mild compared to those associated with regular truncations at 6 and 12 months. The risk ratio associated to 1 month Taylor pricing is in general also relatively mild, indicating that the price changes occurring after 1 month are only to a relatively limited extent attributable to firms that systematically change their price at a monthly interval. Typical Taylor pricing therefore seems to be limited, to a large extent, to the 6 and 12 months duration. Particularly at the 12 months horizon high risk ratios are found and this phenomenon is very pronounced for services.

Overall, this tends to suggest that fixed durations are mainly a phenomenon in the services sector and that this type of price changing takes in particular place at the 12 months horizon. Combined with the pronounced seasonal effect and its particular pattern, which was estimated in the baseline specification, this suggests that these 12 months durations come to an end in a fairly synchronised way, i.e. at the end of the month December (new prices observed in January). This contrasts with the standard Taylor contracts in theoretical models, as they assume staggered pricing.

5. THE RELATIVE IMPORTANCE OF STATE- VERSUS TIME-DEPENDENCY

Before summarising our main conclusions, it seems useful to assess the relative importance of the state-dependent features of our specification. To that end, the likelihood associated with each of the baseline models (7), (10) and (11) is compared with the likelihood of 4 restricted specifications of each baseline model. These restricted specifications take on board, step by step, the different building blocks of the baseline, the aim being to assess the quality of adjustment associated with each block. We are particularly interested in the relative contribution of time- and state-dependent pricing rules to the adjustment. This type of analysis is carried out for the 5 different samples considered in this paper. The results are shown in Table 11.

The first constrained specification (referred to as "Constrained 1"), is a simple Calvo model without heterogeneity, similar to equation (2). This specification is used in order to evaluate the pseudo-R^2 of our baseline equation.

The second specification (referred to as "Constrained 2") is a Calvo model with the relevant sectoral dummies (5 in the "Full CPI" sample and 3 in the "Core components" sample) and unobserved heterogeneity. This specification allows us to identify the contribution of (observed) sectoral and unobserved heterogeneity.

The third specification (referred to as "Constrained 3") adds the logarithm of the duration, the seasonal dummies and the DURx variables to the second specification. This specification allows us to identify the further contribution of time-dependent pricing rules, particularly the contribution of the different types of truncation.

Finally, a fourth restricted specification (referred to as "Constrained 4") excludes the inflation variables from the baseline model. Comparing this last specification with the baseline model allows us to characterise the relative contribution of the accumulated inflation variables.

As to the results for the "Full CPI" sample, the baseline model is characterised by a pseudo-R^2 close to or slightly above 30 p.c. (ranging from 27.2 p.c. for the price increases equation to 32.3 p.c. for the price decreases equations). Because this sample mixes product categories such as oil products (the price of which changes very frequently) and services (characterised by infrequent price changes), it is not surprising to note that sectoral and firm level heterogeneity is by far the largest contributor to the quality of the adjustment. Indeed, from 83 p.c. (for the price increases equation) to 94 p.c. (for the price changes and price increases equations) of the pseudo-R^2 is related to this component of the model. This implies that an infinite mixture of Calvo pricing rules would produce relatively good results in explaining the price-setting behaviour of Belgian firms. In relation to this already large contribution of the mixed Calvo model, the importance of time-dependent pricing rules is further enhanced by the seasonal and truncation dummy variables. Their contribution to the pseudo-R^2 ranges from around 2 p.c. for the price decreases equations to 6 p.c. for the price increases equation (3 p.c. for the price changes equations).

Although the accumulated inflation variables are statistically significant in the estimated equations, their contribution to the pseudo-R^2 is relatively limited. It ranges from close to 2 p.c. for the price changes and price decreases equations to slightly less than 5 p.c. for the price increases equation. This indicates that, all in all, the state dependent component of the model is not very important. As to this "Full CPI" sample, it should however be mentioned that the results presented above are computed without checking for the weight of each product category in the Full CPI. As set out in section 2, this may lead to an over-representation of price reports associated with unprocessed food because the National Statistical Institute follows this type of goods more intensively. Focusing on the core parts of the CPI offsets this over-representation of unprocessed food products.

For the more homogenous core component samples of the CPI, the relative contribution to the pseudo-R^2 of sectoral and firm level heterogeneity is naturally reduced in favour of both the contributions of seasonal and truncation dummies and of the accumulated inflation variables. This is particularly so for the "Processed food" and "Services" samples. For non-energy industrial goods, unobserved heterogeneity still plays the largest role. This is probably due to the fact that this product category comprises goods with very different characteristics, such as for instance durable and non-durable goods.

For the "Processed food" sample, the contribution of the accumulated inflation variables exceeds the contribution of seasonality and truncation by 5 percentage points, at least when price increases and price decreases are considered separately. For this sample, the contribution of accumulated inflation ranges from close to 11 p.c. for the price decrease equation to 21 p.c. for the price increase equation. For non-energy industrial goods and services the contribution of accumulated inflation is less pronounced and ranges, in both cases, from slightly less than 4 p.c. for the price decrease equations to 7 p.c. for the price increase equations. On the basis of the risk ratios presented above, it was also found that the impact of accumulated inflation was more prominent for processed food. For the entire "Core components" sample the contribution of accumulated inflation ranges from nearly 6 p.c. for price decreases to nearly 14 p.c. for price increases.

Seasonality and truncation contribute by approximately 4 p.c. and 17 p.c. to the pseudo-R^2 of respectively the price decrease and the price increase equations for the "Core components" sample. Seasonality and truncation are particularly relevant for services. The contribution of these variables to the pseudo-R^2 of the price increase equation indeed amounts to nearly 50 p.c. for services. This corroborates earlier findings on the basis of the risk ratios, which pointed out that seasonality and truncation were important for services, particularly truncation at the 12 months horizon. Moreover, an important regular pattern - so-called Taylor contracting - was found for this type of truncation, which seems to occur in a fairly synchronised way at the end of each year (new prices observed in January).

6. CONCLUSIONS

This paper presents an analysis of the price-setting behaviour for a large sample of Belgian consumer prices. Our starting point is a model of price adjustment with state-dependent features similar to the one proposed by Cecchetti (1986). However, we also incorporate time-dependent features compatible with potential truncation phenomena in price setting at specific durations, as well as variables characterising some commercial practices at the retail level.

Using mixed Logit models (logistic normal regression), we analyse the probability to observe a price change on the basis of a sub-sample of 12,250 price trajectories that were randomly drawn from the price trajectories underlying the Belgian CPI over the January 1989 - January 2001 period. Mixed Logit models are considered as an alternative to duration models with unobserved heterogeneity. These models are used in order to take into account the heterogeneity of pricing behaviour that has been put forward by former empirical analyses (Álvarez and Hernando (2004), Aucremanne and

Dhyne (2004), Baudry et al. (2004), Bils and Klenow (2004), Dias, Dias and Neves (2004), Jonker et al. (2004)).

The econometric estimation allows us to tackle two not mutually exclusive questions. The first question addresses the issue whether state-dependent pricing is supported by the data, or, alternatively, whether it is sufficient to use only time-dependent factors to describe reality adequately. The second question tackles the shape of the hazard function and pays particular attention to its slope and to the truncation phenomena that are observed at specific horizons. While truncation is compatible with some theoretical time-dependent price-setting models (Taylor contracts and/or the truncated Calvo model), the generally downward sloping nature of the non-parametric estimates of the aggregate hazard is at odds with theoretical models. Hazards are typically non-decreasing for time-dependent models and smoothly increasing for state-dependent models, at least in steady state and with non-zero inflation, as in Dotsey, King and Wolman (1999).

Following the work done by Álvarez et al. (2004), our econometric results clearly show that the downward sloping nature of the hazard is considerably reduced when observed and non-observed heterogeneity is modelled and/or when results are based on samples that are more homogeneous in terms of price-setting behaviour. Also comparable with their results is the fact that we estimate significantly truncation at specific horizons, in particular after 6, 12 and 24 months. The estimated truncation after one month indicates that a non-negligible fraction of price spells are very short-lived.

The econometric results discussed so far would support a general pricing rule that is the result of aggregating some purely time-dependent theoretical pricing models. However, our further econometric results tend to reject this, as several other variables, including some that can clearly be considered as state variables, are found to be significant. This is particularly so for sectoral inflation. The probability to observe a price change is in a significant way an increasing function of the accumulated sectoral inflation since the last price change. This is particularly so when the role of accumulated inflation is estimated separately for price increases and for price decreases. As a result, we were able to construct upward sloping hazards under certain conditions, once the impact of duration incorporates this effect through accumulated inflation. It is indeed through this very channel that theoretical models of state-dependent price-setting produce increasing hazards. These hazards were obtained for sectoral inflation rates that are compatible with a low inflation environment.

This result tends to suggest that sectoral developments condition the relevant "state" for price setting. This does not mean, however, that aggregate inflation is not relevant at all, because movements in aggregate inflation would shift the entire distribution of sectoral inflation rates and, therefore, affect the probabilities to change prices through this channel. In addition, we found that a set of yearly dummies had a significant impact on the price changing probabilities and the coefficients associated with these dummies reflect, to some extent, a pattern which is comparable with the evolution of aggregate (core) inflation over the period considered. We also found that VAT rate changes increase the probability to observe a price change, whereas attractive pricing coincides with a somewhat smaller tendency to adjust prices.

Although our estimation results clearly reject a purely time-dependent representation of price-setting in a statistically significant way and therefore favour a more heterogeneous representation that mixes both time-dependent and state-dependent features, it is important to assess the contribution of each type of pricing in more quantitative terms. This has been done by evaluating the contribution of the time- and state-dependent features to the pseudo-R^2 of the estimated models, as the estimation technique used did not allow identifying the share of firms that apply a particular pricing rule in a more direct way.

In this respect, our results point out that the time-dependent characteristics are far more important than the state-dependent characteristics. This is particularly so for the "Full CPI" sample, where the contribution of time-dependent factors to the pseudo-R^2 of the estimated models ranges from 90 p.c. for price increases to 95 p.c. for price decreases. An infinite mixture of Calvo pricing rules is by itself representing somewhat more than 80 p.c. of the pseudo-R^2 of the price increase equation and more than 90 p.c. for the price decrease equation, while the remaining contribution of time-dependence stems from seasonality and truncation.

However, as very short price spells are over-represented in this sample, the analysis for the different core samples seems to provide better estimates of the importance of each type of price setting. For the core CPI, the contribution of an infinite mixture of Calvo rules attains somewhat more than 50 p.c. for price increases (but 82 p.c. for price decreases). For price increases, seasonality and truncation make a contribution of nearly 14 p.c., a level comparable to the contribution of accumulated inflation. For price decreases, the contribution of the two latter is less than half their contribution for price increases. The role of accumulated inflation is most important for processed food (somewhat more than 20 p.c. of the pseudo R^2 for price increases in that sector). Seasonality and truncation are very widespread in services. Their contribution to the pseudo R^2 for price increases is close to 50 p.c. in this sector.

As to the truncation phenomena, we were able to disentangle those that occur on a regular basis (Taylor pricing) from those that occur irregularly, for instance as a result of so-called truncated Calvo pricing. It turned out that Taylor pricing is estimated significantly, particularly at the 6 and 12 months horizons. However, it seems that this type of Taylor pricing is only relatively important for services, where price spells tend to come to an end at fixed durations of 12 months in a fairly synchronised way at the end of the year (new prices observed in January).

All in all, although our econometric results clearly reject a purely time-dependent representation of price-setting in a statistically significant way, it turns out that the importance of state-dependent features in price-setting is relatively limited. Indeed, at most 20 p.c. of the pseudo-R^2 of the estimated models for the core components of the CPI is attributable to the role of accumulated inflation. Although obtained from a different methodological angle, our results are remarkably in line with the findings of Klenow and Kryvstov (2004). They are also relatively well in line with the results of Álvarez et al (2004).

References

Allenby, G. and P. Lenk (1994): "Modeling Household Purchase Behavior With Logistic Normal Regression", *Journal of the American Statistical Association*, 89, pp. 1218-1231.

Allison, P. (1995): *Survival Analysis Using SAS® : A Practical Guide*, Cary, NC: SAS Institute Inc.

Álvarez, L. and I. Hernando (2004): " Price setting behaviour in Spain: stylised facts using consumer price micro data", *ECB Working Paper Series* No 416.

Álvarez, L., Burriel, P. and I. Hernando (2004): "Do Decreasing Empirical Hazard Functions Make Any Sense ?", Banco de España, Mimeo.

Aucremanne, L. and E. Dhyne (2004): "How Frequently Do Prices Change ? Evidence Based on the Micro-Data Underlying the Belgian CPI"; *ECB Working Paper Series* No 331.

Aucremanne, L. and M. Druant (2004): "Price setting behaviour in Belgium : what can be learned from an ad hoc survey?", *ECB Working Paper Series*, forthcoming.

Baudry, L., Le Bihan H., Sevestre, P. and S. Tarrieu (2004): "Price Rigidity : Evidence from the French CPI Micro-Data", *ECB Working Paper Series,* No 384.

Bils, M. and P. Klenow (2004): "Some Evidence on the Importance of Sticky Prices", Journal of Political Economy, 112, pp 947-985.

Burriel, P, Coenen G., Gali J. and A. Levin (2004): "Structural Models of Inflation and their Micro-Implications", IPN internal document.

Calvo, G. A. (1983): "Staggered Pricing in a Utility Maximizing Framework", *Journal of Monetary Economics*, 12, pp 383-398.

Campbell, J. and B. Eden (2004): "Rigid Prices: Evidence from U.S. Scanner Data", Mimeo.

Cecchetti, S. G. (1986): "The Frequency of Price Ajustment: A Study of the Newsstand Prices of Magazines", *Journal of Econometrics*, Vol 31, pp 255-274.

Cooper, R., Haltiwanger, J. and L. Power (1999): "Machine Replacement and the Business Cycle: Lumps and Bumps", *The American Economic Review*, pp 921-946.

Dias, D., Robalo Marques C. and J. Santos Silva (2004): " Testing price setting theories: evidence from Portuguese micro data ", Banco de Portugal, Mimeo.

Dias, M., Dias, D. and P. Neves (2004): "Stylised Features of Price Setting Behaviour in Portugal: 1992-2001", *ECB Working Paper Series,* No 332.

Dotsey, M., King R. G. and A. L. Wolman (1999): "State-dependent Pricing and the General Equilibrium Dynamics of Money and Output", *Quarterly Journal of Economics*, Vol 114, pp 655-690.

Fougère, D., Le Bihan, H. and P. Sevestre (2004): "Calvo, Taylor And The Estimated Hazard Function For Price Changes", Banque de France, Mimeo.

Galí, J., Gertler, M. and D. López-Salido (2001): "European Inflation Dynamics", *European Economic Review*, 45(6), 1237-1270.

Galí, J., Gertler, M. and D. López-Salido (2003): "Erratum to European Inflation Dynamics [European Economic Review 45 (2001), 1237-1270]", *European Economic Review*, 47(4), 759-760.

Heckman, J. and B. Singer (1984): "A Method for Minimizing the Impact of Distributional Assumptions in Econometric Models for Duration Data", Econometric, March 1984, 52 (2), pp 271-320.

Hsiao, C. (1986): *Analysis of Panel Data*, Cambridge University Press.

Honoré, B. and L. Hu (2004): "Estimation of Discrete Time Duration Models with Grouped Data", Princeton University, Mimeo.

Iway, K. (1981): *Disequilibrium Dynamics*, Yale University Press.

Jonker, N., Blijenberg H. and C. Folkertsma (2004): "An empirical analysis of price setting behaviour in the Netherlands in the period 1998-2003 using micro data ", *ECB Working Paper Series*, No 413.

Klenow, P. and O. Kryvtsov (2004): "State-Dependent or Time-Dependent Pricing: Does it Matter for Recent U.S. Inflation?", Mimeo.

Konieczny, J. D. and A. Skrzypacz (2003): "A New Test of the Menu Cost Model", Mimeo.

Kuss, O. (2002): "How to Use SAS® for Logistic Regression with Correlated Data, *Proceedings of the 27th Annual SAS Users Group International Conference*, Paper 261-27.

Lancaster, T. and S. Nickell (1980): ' "The analysis of re-employment probabilities for the unemployed", *Journal of the Royal Statistical Society*, 143, pp 141-165.

Lawless, J. (2003): *Statistical Models and Methods for Lifetime Data*, Wiley Series in Probability and Statistics, 2nd Edition.

Mash, R. (2004): "Optimising Microfoundations for Inflation Persistence", Mimeo.

Taylor, J. B. (1980): "Aggregate Dynamics and Staggered Contracts", *Journal of Political Economy*, No 88, pp 1-24.

Taylor, J. B. (1999): "Staggered Price and Wage Setting in Macroeconomics", in Taylor and Woodford (Editors), *Handbook of Macroeconomics*, Vol 1b, North Holland.

Train, K. E. (2003), *Discrete Choice Methods with Simulation*, Cambridge University Press.

Veronese G, Fabiani, S., Gattulli, A. and Sabbatini, R. (2004): "Consumer Price Behaviour In Italy: Evidence From Micro CPI Data", Banca d'Italia, Mimeo.

Wolman, A. L. (1999): "Sticky Prices, Marginal Cost, and the Behavior of Inflation", *Federal Reserve Bank of Richmond Economic Quarterly*, Vol 85/4, pp. 29-48.

Wooldridge, J. (2002): *Econometric Analysis of Cross Section and Panel Data*, MIT Press.

Wooldridge, J. (2003): "Simple Solutions to the Initial Conditions Problem in Dynamic, Nonlinear Panel Data Models With Unobserved Heterogeneity", Michigan State University, Mimeo.

Technical appendix

In this section, we briefly describe the econometric methodology used in this paper.

As mentioned in the introduction, the purpose of the paper is to explain the probability to observe a price change/increase/decrease in t accounting for the fact that the price level observed in t has been set by the firm T months before.

As the prices are observed on a discrete time basis, the methodology used is discrete time duration models.

These models are to some extent related to dynamic discrete response models of the type:

$$y_{it} = 1\{x'_{it}\beta + \rho y_{it-1} + \varepsilon_{it}\}$$

where 1{} is an index function.

This equation states that the response of individual i at time t depends not only on the values taken by the exogenous variables x_{it} but also on the response of this individual at time t-1.

If an unobserved effect is introduced, the function becomes:

$$y_{it} = 1\{x'_{it}\beta + \rho y_{it-1} + c_i + \varepsilon_{it}\}$$

One can always write the likelihood function associated with that model as:

$$L(\beta, \rho, c_i, y_{i0}) = \prod_{i=1}^{i_{max}} \prod_{t=1}^{t_{max}} G(x'_{it}\beta + \rho y_{it-1} + c_i)^{y_{it}} \left[1 - G(x'_{it}\beta + \rho y_{it-1} + c_i)\right]^{(1-y_{it})}$$

where G is the cumulated distribution that is associated with the index function 1{}.

As can be seen, this likelihood function depends on the initial value of y for each individual (y_{i0}).

If t_{max} is fixed, the direct estimation of this likelihood function leads to inconsistent estimates of the β and ρ parameters because of the presence of the random effect c_i To estimate those parameters consistently, the integration over c_i is needed, or :

$$L(\beta, \rho, c_i, y_{i0}) = \prod_{i=1}^{i_{max}} \prod_{t=1}^{t_{max}} \int_{-\infty}^{+\infty} G(x'_{it}\beta + \rho y_{it-1} + c_i)^{y_{it}} \left[1 - G(x'_{it}\beta + \rho y_{it-1} + c_i)\right]^{(1-y_{it})} dc_i$$

As pointed out in Wooldridge (2002), the need to integrate c raises the issue of the treatment of the initial conditions.

One can first treat these initial conditions as non-random but this is not desirable as it means that the value of y_{i0} is independent from c_i.

Alternatively, it may be considered that y_{io} are random and that the conditional distribution of y_{i0} on c_i and on all the future values of the exogenous variable is given by $h(y_{i0}|c_i, x_i, \theta)$. Jointly estimating this conditional distribution with the main model allows us to estimate the β and ρ parameters. However, finding the conditional distribution of y_{i0} on c_i and on all the future values of the exogenous variables is very difficult and may be impossible, as is discussed in Hsiao (1986).

Wooldridge (2003) proposes a third alternative based on the specification of the conditional distribution of the unobserved effect. He assumes that the conditional density is given by $h(c_i|y_{io}, x_i, \delta)$. On that assumption the likelihood becomes:

$$L(\beta, \rho, c_i, y_{i0}) = \prod_{i=1}^{i_{max}} \prod_{t=1}^{t_{max}} \int_{-\infty}^{+\infty} G(x'_{it}\beta + \rho y_{it-1} + c_i)^{y_{it}} \left[1 - G(x'_{it}\beta + \rho y_{it-1} + c_i)\right]^{(1-y_{it})} h(c_i|y_{io}, x_i, \delta) dc_i$$

where x_i represents all the past, present and future values observed of x_{it}.

As stated in Wooldridge (2003), specifying the conditional distribution of the unobserved effect on the initial conditions is as good as specifying the conditional distribution of the initial conditions on the unobserved effects. In some cases (such as Probit models with a normal distribution of the heterogeneity parameter), it leads to very simple expressions of the likelihood.

We decided to follow a similar way for the treatment of our duration models. Our specification of the probability to observe a price change is given by:

$$P[y_{it} = 1] = \frac{\exp(x'_{it}\beta + \tau T_{it-1} + c_i + \varepsilon_{it})}{1 + \exp(x'_{it}\beta + \tau T_{it-1} + c_i + \varepsilon_{it})}$$

where T is the duration of the current price spell.

As is put forward in the paper, the auto-regressive structure of the process is hidden by the use of the duration and of the size of the previous price change instead of the value taken by y_{it-1}. In our model, we define t=1 as the first observation of the second uncensored price spell of a price trajectory. Therefore, the value of y_{i0} is not informative, as it is always 1. However, the duration of the first uncensored price spell and the size of the first price change observed may be informative. It has to be mentioned that the auto-regressive structure of our specification is strengthened in our "Taylor" specifications which explicitly uses the duration of the previous price spell to derive some of the explanatory variables (the "Taylor" dummies). Therefore, we decided to use a similar approach as the one described above. Moreover, the estimation procedure (NLMIXED) in SAS explicitly uses a representation of the likelihood similar to the one proposed by Wooldridge (2003)..

The likelihood associated with our model is provided by :

$$\prod_{i=1}^{i_{max}}\prod_{t=1}^{t_{max}}\int_{-\infty}^{+\infty} G(x'_{it}\beta+\tau T_{it}+c_i)^{y_{it}}\left[1-G(x'_{it}\beta+\tau T_{it}+c_i)\right]^{(1-y_{it})} h(c_i|T_{io},Ldprix_{i0},x_i,\delta)dc_i$$

where $h(c_i|T_{io},Ldprix_{io},x_i,\delta)=N\left(\mu_1 T_{i0}+\mu_2 Ldprix_{i0};\sum_{i=1}^{5}\sigma_i^2 SECk_i\right)$

In section 3.4.1., T_{i0} and $Ldprix_{i0}$ are respectively designated as $1STDUR_i$ and $1STLDP_i$.

We used a simple functional form for the expression of the conditional distribution of the unobserved effect, which assumes that the exogenous x_i, which represents the past, present and future value of x at time t, do not affect the distribution of c_i because of the unbalanced structure of our panel. However, we allow for sectoral heteroscedasticity.

Statistical appendix

General remarks concerning Tables 1 to 10 :

1. the price changes equations have been estimated using all price reports available;

2. the price increases equations have been estimated disregarding price reports characterised by Y_{2ijt} equal to 1 or if $P_{ijt} > P_{ij;t+1}$;

3. the price decreases equations have been estimated disregarding price reports characterised by Y_{1ijt} equal to 1 or if $P_{ijt} < P_{ij,t+1}$;

4. the cells in light grey with results in italic indicate coefficients that are not significant at the 5 p.c. level.

Table 1 - Estimated results - Baseline model - Full CPI

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-2.2997	-	-2.3764	-	-5.1445	-
ln(Dur)	-0.2137	0.8076	-0.2988	0.7417	-0.2737	0.7606
Dur1	0.3536	1.4242	0.4629	1.5887	0.1359	1.1456
Dur6	0.1217	1.1294	0.1517	1.1638	0.0903	1.0945
Dur12	0.7852	2.1928	0.9475	2.5793	0.4527	1.5726
Dur18	0.1282	1.1368	*0.0973*	*1.1022*	*0.1787*	*1.1957*
Dur24	0.5610	1.7524	0.5937	1.8107	0.4566	1.5787
Dur36	0.3795	1.4616	0.4490	1.5667	*0.0861*	*1.0899*
Accumulated inflation in absolute terms	0.0616	1.0635	-	-	-	-
Positive accumulated inflation	-	-	0.1009	1.1061	0.0207	1.0209
Negative accumulated inflation	-	-	0.0299	1.0304	0.0851	1.0888
Increase of VAT rate	0.9927	2.6985	0.9298	2.5340	1.1202	3.0655
Decrease of VAT rate	2.8299	16.9438	1.0401	2.8295	3.7902	44.2653
Attractive pricing	-0.0512	0.9501	-0.0503	0.9509	-0.1038	0.9014
Size of previous price change if positive	0.1946	1.2148	-0.5853	0.5569	0.7219	2.0583
Size of previous price change if negative	0.3747	1.4546	0.4417	1.5553	*0.0009*	*1.0009*
Previous price change negative	0.3293	1.3900	0.7763	2.1734	-0.2964	0.7435
Year 1990	-0.1267	0.8810	-0.1768	0.8379	*-0.0910*	*0.9130*
Year 1991	*0.0238*	*1.0241*	*0.0098*	*1.0099*	*0.0069*	*1.0070*
Year 1992	-0.1196	0.8873	-0.1810	0.8344	*-0.0925*	*0.9117*
Year 1993	-0.2650	0.7672	-0.3149	0.7299	-0.2000	0.8187
Year 1994	-0.2985	0.7419	-0.3124	0.7317	-0.2588	0.7720
Year 1995	-0.3223	0.7245	-0.3035	0.7382	-0.2922	0.7466
Year 1996	-0.2702	0.7632	-0.2639	0.7681	-0.2121	0.8089
Year 1997	-0.3707	0.6903	-0.3767	0.6861	-0.2652	0.7671
Year 1998	-0.2797	0.7560	-0.3104	0.7332	-0.1705	0.8432
Year 1999	-0.2509	0.7781	-0.2911	0.7474	*-0.0893*	*0.9146*
Year 2000	-0.4138	0.6611	-0.4029	0.6684	-0.3399	0.7118
January	0.0508	1.0521	*-0.0256*	*0.9748*	0.0754	1.0783
February	0.2025	1.2245	0.1507	1.1626	0.2159	1.2410
March	0.2063	1.2291	0.1205	1.1281	0.3221	1.3800
April	0.1146	1.1214	*0.0305*	*1.0310*	0.2310	1.2599
May	0.0716	1.0742	*0.0490*	*1.0503*	0.1445	1.1555
June	*-0.0418*	*0.9591*	-0.2043	0.8152	0.1935	1.2135
July	*-0.0419*	*0.9590*	-0.1611	0.8512	0.1884	1.2073
August	0.1488	1.1604	*0.0512*	*1.0526*	0.2961	1.3446
September	0.2203	1.2465	0.1243	1.1324	0.3652	1.4408
October	0.1503	1.1622	*0.0235*	*1.0238*	0.3189	1.3756
November	0.0546	1.0561	-0.0756	0.9272	0.2656	1.3042
Unprocessed food	1.3818	3.9821	0.8240	2.2796	3.7500	42.5211
Processed food	0.8194	2.2691	0.3754	1.4556	2.9723	19.5368
Energy	2.5612	12.9513	1.9315	6.8999	4.7918	120.5181
Non-energy industrial goods	0.2639	1.3020	*-0.1066*	*0.8989*	2.1792	8.8392
Duration of first uncensored price spell	-0.0525	0.9488	-0.0452	0.9558	-0.0778	0.9251
Size of first price change observed	0.1215	1.1292	0.0956	1.1004	0.2124	1.2366
Unprocessed food - su2	1.5639	-	1.1028	-	1.9349	-
Processed food - su2	0.6225	-	0.4279	-	0.9605	-
Energy - su2	0.9976	-	0.5370	-	1.3103	-
Non-energy industrial goods - su2	0.8264	-	0.7157	-	1.7878	-
Services - su2	0.6104	-	0.5493	-	3.5442	-
Likelihood ratio	-154028	-	-108573	-	-86805	-

Table 2 - Estimated results - Baseline model - Core components

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-2.9718	-	-3.4296	-	-6.0137	-
ln(Dur)	-0.2599	0.7711	-0.3169	0.7284	-0.2888	0.7492
Dur1	0.2569	1.2929	0.3811	1.4639	*-0.0237*	*0.9766*
Dur6	0.1903	1.2096	0.2109	1.2348	0.1743	1.1904
Dur12	0.9138	2.4938	1.0951	2.9895	0.4762	1.6099
Dur18	*0.1327*	*1.1419*	*0.1489*	*1.1606*	*0.1344*	*1.1439*
Dur24	0.5920	1.8076	0.6657	1.9459	0.4267	1.5322
Dur36	0.2929	1.3403	*0.3124*	*1.3667*	*0.1296*	*1.1384*
Accumulated inflation in absolute terms	0.1434	1.1541	-	-	-	-
Positive accumulated inflation	-	-	0.2100	1.2337	-0.0220	0.9783
Negative accumulated inflation	-	-	-0.0815	0.9217	0.2278	1.2558
Increase of VAT rate	0.8410	2.3187	0.7402	2.0964	1.0059	2.7344
Decrease of VAT rate	2.8817	17.8446	1.5330	4.6321	3.6499	38.4708
Attractive pricing	-0.0656	0.9365	-0.0591	0.9426	-0.1141	0.8922
Size of previous price change if positive	-0.3441	0.7089	-1.7406	0.1754	0.6966	2.0069
Size of previous price change if negative	0.1804	1.1977	0.2186	1.2443	-0.6750	0.5092
Previous price change negative	0.4131	1.5115	0.9526	2.5924	-0.3241	0.7232
Year 1990	*-0.0725*	*0.9301*	*-0.0455*	*0.9555*	*-0.0796*	*0.9235*
Year 1991	0.2648	1.3032	0.3023	1.3530	0.2886	1.3346
Year 1992	0.1929	1.2128	0.1759	1.1923	0.3146	1.3697
Year 1993	0.0540	1.0555	*0.0719*	*1.0745*	0.2018	1.2236
Year 1994	-0.0008	0.9992	*0.0431*	*1.0441*	*0.1507*	*1.1626*
Year 1995	-0.0064	0.9937	*0.0416*	*1.0425*	*0.1313*	*1.1403*
Year 1996	*-0.0066*	*0.9935*	*-0.0021*	*0.9979*	0.2220	1.2486
Year 1997	-0.1397	0.8696	-0.1412	0.8683	*0.1255*	*1.1337*
Year 1998	*-0.0125*	*0.9875*	*0.0129*	*1.0130*	0.2360	1.2662
Year 1999	*-0.0033*	*0.9967*	*-0.0155*	*0.9846*	0.3269	1.3867
Year 2000	-0.1226	0.8846	*-0.1263*	*0.8814*	*0.1772*	*1.1939*
January	0.1517	1.1638	*0.0669*	*1.0692*	0.2883	1.3342
February	0.2908	1.3375	0.2019	1.2237	0.4056	1.5002
March	0.2428	1.2748	0.0803	1.0836	0.4885	1.6299
April	0.0651	1.0672	-0.1125	0.8936	0.3495	1.4184
May	*0.0262*	*1.0265*	-0.1117	0.8943	0.2596	1.2964
June	-0.1023	0.9028	-0.2234	0.7998	0.1194	1.1268
July	-0.1613	0.8510	-0.2253	0.7983	*-0.0138*	*0.9863*
August	0.1256	1.1338	*0.0264*	*1.0268*	0.2954	1.3437
September	0.1942	1.2143	0.1156	1.1225	0.3333	1.3956
October	0.1588	1.1721	*0.0424*	*1.0433*	0.3992	1.4906
November	*0.0343*	*1.0349*	-0.1097	0.8961	0.3057	1.3576
Unprocessed food	-	-	-	-	-	-
Processed food	1.1734	3.2330	1.0663	2.9046	3.3596	28.7777
Energy	-	-	-	-	-	-
Non-energy industrial goods	0.5333	1.7045	0.4480	1.5652	2.5931	13.3712
Duration of first uncensored price spell	-0.0429	0.9580	-0.0365	0.9642	-0.0591	0.9427
Size of first price change observed	*-0.0013*	*0.9987*	*-0.0023*	*0.9977*	*0.0568*	*1.0585*
Unprocessed food - su2	-	-	-	-	-	-
Processed food - su2	0.6288	-	0.4749	-	0.8851	-
Energy - su2	-	-	-	-	-	-
Non-energy industrial goods - su2	0.9216	-	1.0467	-	1.5903	-
Services - su2	1.1649	-	2.0362	-	4.3953	-
Likelihood ratio	-93116	-	-65306	-	-47133	-

Table 3 - Estimated results - Baseline model - Processed food

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-1.9218	-	-2.7229	-	-2.4522	-
ln(Dur)	-0.3049	0.7372	-0.3301	0.7189	-0.3208	0.7256
Dur1	0.2372	1.2677	0.4334	1.5425	-0.0999	0.9050
Dur6	0.0861	1.0900	*0.0538*	*1.0552*	0.1526	1.1649
Dur12	0.3207	1.3781	0.4139	1.5127	0.2050	1.2275
Dur18	*0.1352*	*1.1448*	*0.1666*	*1.1813*	*0.0936*	*1.0982*
Dur24	0.3599	1.4332	*0.2833*	*1.3275*	0.4633	1.5893
Dur36	*-0.1371*	*0.8719*	*-0.4177*	*0.6586*	*-0.1518*	*0.8592*
Accumulated inflation in absolute terms	0.1883	1.2072	-	-	-	-
Positive accumulated inflation	-	-	0.3063	1.3584	-0.0999	0.9049
Negative accumulated inflation	-	-	-0.1597	0.8524	0.2689	1.3085
Increase of VAT rate	1.1066	3.0241	0.7337	2.0828	1.5192	4.5686
Decrease of VAT rate	1.7203	5.5862	*-1.2784*	*0.2785*	1.9843	7.2740
Attractive pricing	*-0.0329*	*0.9676*	*-0.0343*	*0.9663*	-0.0618	0.9401
Size of previous price change if positive	-0.5811	0.5593	-2.5615	0.0772	0.4585	1.5817
Size of previous price change if negative	0.1499	1.1617	0.1669	1.1816	-0.5548	0.5742
Previous price change negative	0.4531	1.5732	1.1124	3.0416	-0.4933	0.6106
Year 1990	-0.1443	0.8656	*-0.1231*	*0.8842*	*-0.0931*	*0.9111*
Year 1991	0.3034	1.3545	0.4068	1.5020	0.2704	1.3105
Year 1992	0.2530	1.2879	0.3326	1.3946	0.2567	1.2927
Year 1993	*0.1161*	*1.1231*	0.2187	1.2445	*0.1691*	*1.1842*
Year 1994	*0.0818*	*1.0852*	0.2116	1.2357	*0.1190*	*1.1264*
Year 1995	0.1736	1.1896	0.3063	1.3584	*0.1778*	*1.1946*
Year 1996	0.1825	1.2002	0.3314	1.3929	*0.1691*	*1.1842*
Year 1997	*0.0443*	*1.0453*	*0.1380*	*1.1480*	*0.1019*	*1.1073*
Year 1998	*0.1027*	*1.1082*	0.2509	1.2852	*0.1498*	*1.1616*
Year 1999	0.1651	1.1795	0.2524	1.2871	0.3052	1.3569
Year 2000	*0.0266*	*1.0270*	*0.0917*	*1.0960*	*0.1856*	*1.2039*
January	0.1526	1.1649	0.1044	1.1100	0.2033	1.2254
February	0.2502	1.2843	0.1883	1.2072	0.3027	1.3535
March	0.2772	1.3194	0.1461	1.1573	0.4153	1.5148
April	0.1077	1.1137	*-0.0224*	*0.9779*	0.2732	1.3142
May	0.0808	1.0842	*-0.0004*	*0.9996*	0.1684	1.1834
June	*0.0353*	*1.0359*	*-0.0395*	*0.9613*	0.1125	1.1191
July	*-0.0433*	*0.9577*	*-0.0434*	*0.9575*	*-0.0620*	*0.9399*
August	0.0862	1.0900	*0.0026*	*1.0026*	0.1833	1.2012
September	0.1747	1.1909	0.1214	1.1291	0.2473	1.2806
October	0.2027	1.2247	0.1187	1.1260	0.3339	1.3964
November	*0.0697*	*1.0722*	*-0.0632*	*0.9387*	0.2544	1.2897
Unprocessed food	-	-	-	-	-	-
Processed food	-	-	-	-	-	-
Energy	-	-	-	-	-	-
Non-energy industrial goods	-	-	-	-	-	-
Duration of first uncensored price spell	-0.0429	0.9580	-0.0372	0.9635	-0.0564	0.9451
Size of first price change observed	*0.0091*	*1.0091*	*-0.0043*	*0.9957*	*0.0152*	*1.0153*
Unprocessed food - su2	-	-	-	-	-	-
Processed food - su2	0.5798	-	0.5126	-	0.8939	-
Energy - su2	-	-	-	-	-	-
Non-energy industrial goods - su2	-	-	-	-	-	-
Services - su2	-	-	-	-	-	-
Likelihood ratio	-57784	-	-38541	-	-31427	-

Table 4 - Estimated results - Baseline model - Non-energy industrial goods

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-2.6955	-	-2.9367	-	-3.9938	-
ln(Dur)	-0.1342	0.8744	-0.2423	0.7848	-0.1758	0.8388
Dur1	0.4090	1.5053	0.4086	1.5047	0.2548	1.2902
Dur6	0.3267	1.3864	0.3885	1.4748	0.2211	1.2474
Dur12	0.9263	2.5251	0.9967	2.7093	0.8121	2.2526
Dur18	0.0886	1.0927	0.0622	1.0642	0.1653	1.1797
Dur24	0.4807	1.6172	0.5335	1.7049	0.3930	1.4814
Dur36	0.3504	1.4196	0.3028	1.3536	0.3928	1.4811
Accumulated inflation in absolute terms	0.1003	1.1055	-	-	-	-
Positive accumulated inflation	-	-	0.1385	1.1485	0.0142	1.0143
Negative accumulated inflation	-	-	0.0035	1.0035	0.1664	1.1810
Increase of VAT rate	0.7386	2.0930	0.8201	2.2707	0.4818	1.6190
Decrease of VAT rate	3.1868	24.2108	1.4038	4.0706	3.9939	54.2661
Attractive pricing	-0.0881	0.9157	-0.0595	0.9422	-0.1926	0.8248
Size of previous price change if positive	0.2815	1.3251	-0.5837	0.5578	1.3461	3.8424
Size of previous price change if negative	1.1015	3.0087	2.5644	12.9929	-1.4662	0.2308
Previous price change negative	0.2937	1.3414	0.4690	1.5984	0.0841	1.0877
Year 1990	-0.0639	0.9381	-0.0449	0.9561	-0.1020	0.9030
Year 1991	0.1701	1.1854	0.1405	1.1508	0.2840	1.3284
Year 1992	-0.0507	0.9506	-0.1736	0.8406	0.2485	1.2821
Year 1993	-0.2263	0.7975	-0.2907	0.7477	0.0641	1.0662
Year 1994	-0.2981	0.7422	-0.3810	0.6832	0.0672	1.0695
Year 1995	-0.4766	0.6209	-0.5171	0.5962	-0.1490	0.8616
Year 1996	-0.3813	0.6830	-0.5870	0.5560	0.2081	1.2313
Year 1997	-0.5678	0.5668	-0.6869	0.5031	0.0022	1.0022
Year 1998	-0.2232	0.8000	-0.3479	0.7062	0.3103	1.3638
Year 1999	-0.3330	0.7168	-0.4277	0.6520	0.2396	1.2707
Year 2000	-0.4151	0.6603	-0.4430	0.6421	0.0502	1.0515
January	0.5354	1.7081	0.5348	1.7071	0.4390	1.5512
February	0.8449	2.3277	0.8045	2.2356	0.7217	2.0579
March	0.6316	1.8806	0.5194	1.6810	0.6782	1.9703
April	0.4371	1.5482	0.3035	1.3546	0.5449	1.7244
May	0.3943	1.4833	0.2229	1.2497	0.5706	1.7693
June	0.0732	1.0759	-0.0478	0.9534	0.2282	1.2563
July	0.0312	1.0317	-0.0736	0.9291	0.1705	1.1859
August	0.6142	1.8482	0.5573	1.7460	0.6109	1.8421
September	0.7012	2.0162	0.6747	1.9634	0.6061	1.8333
October	0.5070	1.6603	0.4075	1.5031	0.6356	1.8882
November	0.3759	1.4563	0.2991	1.3486	0.4577	1.5804
Unprocessed food	-	-	-	-	-	-
Processed food	-	-	-	-	-	-
Energy	-	-	-	-	-	-
Non-energy industrial goods	-	-	-	-	-	-
Duration of first uncensored price spell	-0.0434	0.9575	-0.0380	0.9627	-0.0571	0.9445
Size of first price change observed	0.2460	1.2789	0.2429	1.2749	0.4798	1.6158
Unprocessed food - su2	-	-	-	-	-	-
Processed food - su2	-	-	-	-	-	-
Energy - su2	-	-	-	-	-	-
Non-energy industrial goods - su2	0.9873	-	0.7199	-	1.6392	-
Services - su2	-	-	-	-	-	-
Likelihood ratio	-29177	-	-20857	-	-14327	-

Table 5 - Estimated results - Baseline model - Services

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-1.5750	-	-1.2834	-	-29.0972	-
ln(Dur)	-0.1689	0.8446	-0.1738	0.8405	*-0.2632*	*0.7686*
Dur1	-0.5030	0.6047	-0.3681	0.6920	-1.2467	0.2875
Dur6	0.4549	1.5760	0.4622	1.5876	*0.3856*	*1.4705*
Dur12	2.3473	10.4573	2.3782	10.7855	1.4012	4.0601
Dur18	*-0.1510*	*0.8598*	*-0.1713*	*0.8426*	*-0.0627*	*0.9392*
Dur24	1.4106	4.0984	1.5052	4.5051	-	-
Dur36	1.1473	3.1497	1.2320	3.4281	-	-
Accumulated inflation in absolute terms	0.1191	1.1264	-	-	-	-
Positive accumulated inflation	-	-	0.1126	1.1192	*0.0177*	*1.0178*
Negative accumulated inflation	-	-	*-0.1359*	*0.8729*	1.1146	3.0483
Increase of VAT rate	1.3404	3.8206	1.2824	3.6053	*0.7224*	*2.0594*
Decrease of VAT rate	*1.1713*	*3.2262*	-	-	4.1378	62.6648
Attractive pricing	-0.3773	0.6857	-0.4452	0.6407	-0.3947	0.6739
Size of previous price change if positive	*-0.3634*	*0.6953*	*-0.6801*	*0.5066*	*-0.0654*	*0.9367*
Size of previous price change if negative	*0.1662*	*1.1808*	*0.1283*	*1.1369*	-7.4621	0.0006
Previous price change negative	*0.0781*	*1.0812*	0.3169	1.3729	*-0.3173*	*0.7281*
Year 1990	0.6662	1.9468	*0.4308*	*1.5385*	24.8872	-
Year 1991	*0.1329*	*1.1421*	*-0.1795*	*0.8357*	25.0368	-
Year 1992	*0.1731*	*1.1890*	*-0.2072*	*0.8129*	24.9173	-
Year 1993	*0.1120*	*1.1185*	*-0.2952*	*0.7444*	25.3425	-
Year 1994	*-0.0735*	*0.9291*	*-0.4665*	*0.6272*	24.6490	-
Year 1995	*-0.2633*	*0.7685*	-0.6588	0.5175	24.9929	-
Year 1996	-0.6801	0.5066	-1.0782	0.3402	24.1526	-
Year 1997	*-0.4074*	*0.6654*	-1.1596	0.3136	25.7253	-
Year 1998	*-0.5939*	*0.5522*	-1.0253	0.3587	24.6880	-
Year 1999	*-0.5953*	*0.5514*	-0.9245	0.3967	23.0273	-
Year 2000	-0.7404	0.4769	-1.1082	0.3302	24.5912	-
January	-0.2746	0.7599	-0.4497	0.6378	*0.1381*	*1.1481*
February	-0.7391	0.4775	-0.8224	0.4394	-0.6265	0.5345
March	-0.9379	0.3914	-1.0121	0.3635	-0.8761	0.4164
April	-1.0250	0.3588	-1.2105	0.2980	-0.6123	0.5421
May	-1.2644	0.2824	-1.2385	0.2898	-1.6895	0.1846
June	-1.6074	0.2004	-1.5726	0.2075	-1.9781	0.1383
July	-1.4701	0.2299	-1.4580	0.2327	-1.4798	0.2277
August	-0.8250	0.4382	-0.7914	0.4532	-1.0334	0.3558
September	-1.0786	0.3401	-1.0526	0.3490	-1.3057	0.2710
October	-0.8688	0.4195	-0.8013	0.4487	-1.4114	0.2438
November	-0.9220	0.3977	-0.9691	0.3794	*-0.4787*	*0.6196*
Unprocessed food	-	-	-	-	-	-
Processed food	-	-	-	-	-	-
Energy	-	-	-	-	-	-
Non-energy industrial goods	-	-	-	-	-	-
Duration of first uncensored price spell	-0.0353	0.9654	-0.0265	0.9739	-0.1114	0.8946
Size of first price change observed	-4.7320	0.0088	-4.3291	0.0132	*0.6668*	*1.9480*
Unprocessed food - su2	-	-	-	-	-	-
Processed food - su2	-	-	-	-	-	-
Energy - su2	-	-	-	-	-	-
Non-energy industrial goods - su2	-	-	-	-	-	-
Services - su2	1.1848	-	0.7390	-	3.8028	-
Likelihood ratio	-5285	-	-4774	-	-1052	-

Table 6 - Estimated results - "Taylor" model - Full CPI

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-2.3766	-	-2.4653	-	-5.2062	-
ln(Dur)	-0.2154	0.8062	-0.3030	0.7386	-0.2965	0.7434
Dur1	0.2781	1.3206	0.3926	1.4808	-0.0801	0.9230
Dur6	0.0800	1.0833	0.1045	1.1102	*0.0607*	*1.0625*
Dur12	0.5532	1.7388	0.6563	1.9276	0.3945	1.4836
Dur18	0.1197	1.1272	*0.0894*	*1.0935*	*0.1776*	*1.1943*
Dur24	0.5490	1.7315	0.5846	1.7943	0.4547	1.5757
Dur36	0.3755	1.4557	0.4482	1.5655	*0.0926*	*1.0970*
Accumulated inflation in absolute terms	0.0619	1.0639	-	-	-	-
Positive accumulated inflation	-	-	0.1012	1.1065	0.0219	1.0222
Negative accumulated inflation	-	-	0.0307	1.0312	0.0866	1.0905
Increase of VAT rate	1.0149	2.7591	0.9540	2.5961	1.1237	3.0762
Decrease of VAT rate	2.8356	17.0406	1.0321	2.8070	3.8509	47.0354
Attractive pricing	-0.0528	0.9486	-0.0520	0.9494	-0.1061	0.8993
Size of previous price change if positive	0.1942	1.2143	-0.5969	0.5505	0.6915	1.9967
Size of previous price change if negative	0.3729	1.4519	0.4300	1.5373	*-0.0168*	*0.9833*
Previous price change negative	0.3358	1.3991	0.7804	2.1823	-0.2855	0.7516
Non-standard "Taylor contract" < 1 year	0.1582	1.1714	0.2052	1.2278	0.0902	1.0944
"Taylor contract" 1 month	0.2194	1.2453	0.2384	1.2692	0.4210	1.5235
"Taylor contract" 6 months	0.6864	1.9866	0.7373	2.0903	0.4917	1.6351
"Taylor contract" 12 months	2.1939	8.9701	2.2650	9.6311	1.1541	3.1712
"Taylor contract" > 1 year	0.4165	1.5166	0.4928	1.6369	*0.1916*	*1.2112*
Year 1990	-0.1168	0.8898	-0.1553	0.8562	*-0.0764*	*0.9265*
Year 1991	*0.0277*	*1.0281*	*0.0257*	*1.0260*	*0.0154*	*1.0155*
Year 1992	-0.1166	0.8899	-0.1663	0.8468	*-0.0730*	*0.9296*
Year 1993	-0.2546	0.7752	-0.2890	0.7490	-0.1719	0.8421
Year 1994	-0.2911	0.7474	-0.2920	0.7468	-0.2328	0.7923
Year 1995	-0.3143	0.7303	-0.2821	0.7542	-0.2638	0.7681
Year 1996	-0.2615	0.7699	-0.2412	0.7857	-0.1882	0.8284
Year 1997	-0.3593	0.6982	-0.3520	0.7033	-0.2384	0.7879
Year 1998	-0.2673	0.7654	-0.2852	0.7519	-0.1363	0.8726
Year 1999	-0.2403	0.7864	-0.2689	0.7642	*-0.0614*	*0.9404*
Year 2000	-0.4046	0.6672	-0.3826	0.6821	-0.3097	0.7337
January	0.0692	1.0716	*-0.0024*	*0.9976*	0.0776	1.0807
February	0.2205	1.2467	0.1734	1.1893	0.2172	1.2426
March	0.2242	1.2513	0.1424	1.1530	0.3223	1.3803
April	0.1347	1.1442	0.0554	1.0570	0.2340	1.2636
May	0.0891	1.0932	0.0701	1.0726	0.1422	1.1528
June	*-0.0231*	*0.9772*	-0.1818	0.8338	0.1983	1.2193
July	*-0.0257*	*0.9746*	-0.1389	0.8703	0.1845	1.2026
August	0.1608	1.1745	0.0660	1.0682	0.2982	1.3474
September	0.2399	1.2711	0.1483	1.1599	0.3698	1.4474
October	0.1694	1.1846	*0.0470*	*1.0481*	0.3203	1.3775
November	0.0747	1.0776	*-0.0495*	*0.9517*	0.2692	1.3089
Unprocessed food	1.4053	4.0767	0.8479	2.3347	3.7984	44.6297
Processed food	0.8675	2.3810	0.4267	1.5322	3.0605	21.3382
Energy	2.5376	12.6493	1.9558	7.0696	4.8026	121.8268
Non-energy industrial goods	0.3029	1.3538	*-0.0569*	*0.9447*	2.2949	9.9234
Duration of first uncensored price spell	-0.0520	0.9493	-0.0449	0.9561	-0.0778	0.9252
Size of first price change observed	0.1208	1.1284	0.0891	1.0932	0.2321	1.2612
Unprocessed food - su2	1.4552	-	0.9628	-	1.6211	-
Processed food - su2	0.5995	-	0.4008	-	0.9093	-
Energy - su2	0.9074	-	0.4854	-	1.0706	-
Non-energy industrial goods - su2	0.8129	-	0.7439	-	1.6433	-
Services - su2	0.6085	-	0.5015	-	3.3101	-
Likelihood ratio	-153627	-	-108232	-	-86634	-

Table 7 - Estimated results - "Taylor" model - Core components

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-3.0501	-	-3.3441	-	-6.0159	-
ln(Dur)	-0.2554	0.7746	-0.3115	0.7323	-0.2888	0.7492
Dur1	0.2189	1.2447	0.3367	1.4003	-0.1329	0.8756
Dur6	0.1481	1.1596	0.1594	1.1728	0.1510	1.1630
Dur12	0.6416	1.8995	0.7794	2.1802	0.4026	1.4957
Dur18	*0.1274*	*1.1359*	*0.1354*	*1.1450*	*0.1260*	*1.1343*
Dur24	0.5815	1.7887	0.6517	1.9188	0.4249	1.5294
Dur36	*0.2864*	*1.3316*	*0.3155*	*1.3709*	*0.1287*	*1.1373*
Accumulated inflation in absolute terms	0.1435	1.1543	-	-	-	-
Positive accumulated inflation	-	-	0.2047	1.2271	-0.0220	0.9782
Negative accumulated inflation	-	-	-0.0815	0.9217	0.2276	1.2556
Increase of VAT rate	0.8696	2.3860	0.7750	2.1706	1.0120	2.7511
Decrease of VAT rate	2.9148	18.4451	1.5435	4.6809	3.6595	38.8419
Attractive pricing	-0.0669	0.9353	-0.0630	0.9390	-0.1214	0.8857
Size of previous price change if positive	-0.3515	0.7036	-1.7517	0.1735	0.6194	1.8578
Size of previous price change if negative	0.1786	1.1955	0.2142	1.2389	-0.6957	0.4987
Previous price change negative	0.4173	1.5179	0.9488	2.5826	-0.3229	0.7240
Non-standard "Taylor contract" < 1 year	0.1789	1.1959	0.2111	1.2350	0.1466	1.1579
"Taylor contract" 1 month	0.1867	1.2053	0.2317	1.2607	0.3135	1.3682
"Taylor contract" 6 months	0.6718	1.9578	0.7537	2.1248	0.4220	1.5250
"Taylor contract" 12 months	2.2426	9.4178	2.2661	9.6417	1.1852	3.2713
"Taylor contract" > 1 year	0.3672	1.4437	0.4714	1.6022	*0.0554*	*1.0569*
Year 1990	*-0.0572*	*0.9444*	*-0.0282*	*0.9722*	*-0.0742*	*0.9285*
Year 1991	0.2769	1.3190	0.3268	1.3865	0.2847	1.3294
Year 1992	0.1949	1.2152	0.1813	1.1988	0.3107	1.3644
Year 1993	*0.0647*	*1.0669*	*0.0886*	*1.0926*	0.2000	1.2214
Year 1994	*0.0067*	*1.0068*	*0.0549*	*1.0565*	*0.1482*	*1.1597*
Year 1995	*-0.0017*	*0.9983*	*0.0490*	*1.0502*	*0.1300*	*1.1388*
Year 1996	*0.0014*	*1.0014*	*0.0092*	*1.0092*	0.2199	1.2460
Year 1997	-0.1295	0.8785	*-0.1275*	*0.8803*	*0.1245*	*1.1326*
Year 1998	*-0.0018*	*0.9982*	*0.0260*	*1.0264*	0.2351	1.2650
Year 1999	*0.0058*	*1.0059*	*-0.0059*	*0.9942*	0.3251	1.3842
Year 2000	*-0.1174*	*0.8892*	*-0.1203*	*0.8867*	*0.1744*	*1.1905*
January	0.1830	1.2008	0.1114	1.1178	0.2931	1.3406
February	0.3223	1.3803	0.2442	1.2766	0.4119	1.5097
March	0.2725	1.3132	0.1191	1.1265	0.4919	1.6354
April	0.0964	1.1012	-0.0741	0.9286	0.3538	1.4245
May	*0.0548*	*1.0564*	-0.0762	0.9266	0.2603	1.2973
June	-0.0744	0.9283	-0.1895	0.8274	0.1218	1.1295
July	-0.1325	0.8759	-0.1888	0.8280	*-0.0117*	*0.9884*
August	0.1458	1.1570	*0.0514*	*1.0528*	0.3013	1.3516
September	0.2266	1.2543	0.1576	1.1707	0.3411	1.4065
October	0.1923	1.2120	0.0842	1.0878	0.4062	1.5011
November	0.0659	1.0681	*-0.0689*	*0.9334*	0.3109	1.3647
Unprocessed food	-	-	-	-	-	-
Processed food	1.1992	3.3175	0.9291	2.5322	3.3597	28.7806
Energy	-	-	-	-	-	-
Non-energy industrial goods	0.5476	1.7291	0.3082	1.3610	2.5897	13.3258
Duration of first uncensored price spell	-0.0423	0.9586	-0.0363	0.9644	-0.0581	0.9436
Size of first price change observed	*-0.0011*	*0.9989*	*-0.0068*	*0.9932*	*0.0506*	*1.0519*
Unprocessed food - su2	-	-	-	-	-	-
Processed food - su2	0.6069	-	0.4470	-	0.8470	-
Energy - su2	-	-	-	-	-	-
Non-energy industrial goods - su2	0.9087	-	1.0330	-	1.6654	-
Services - su2	1.1372	-	1.7906	-	4.3927	-
Likelihood ratio	-92797	-	-65030	-	-47076	-

Table 8 - Estimated results - "Taylor" model - Processed food

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-1.9247	-	-2.7503	-	-2.4020	-
ln(Dur)	-0.3017	0.7396	-0.3267	0.7213	-0.3147	0.7300
Dur1	0.2167	1.2420	0.4041	1.4980	-0.1914	0.8258
Dur6	0.0939	1.0985	*0.0637*	*1.0658*	0.1568	1.1698
Dur12	0.2905	1.3371	0.3752	1.4553	0.1970	1.2177
Dur18	*0.1329*	*1.1421*	*0.1575*	*1.1706*	*0.0919*	*1.0962*
Dur24	0.3609	1.4346	*0.2667*	*1.3056*	0.4381	1.5498
Dur36	*-0.2750*	*0.7596*	*-0.4653*	*0.6279*	*-0.0197*	*0.9805*
Accumulated inflation in absolute terms	0.1883	1.2071	-	-	-	-
Positive accumulated inflation	-	-	0.3061	1.3581	-0.1150	0.8914
Negative accumulated inflation	-	-	-0.1597	0.8524	0.2485	1.2822
Increase of VAT rate	1.1146	3.0483	0.7421	2.1003	1.5264	4.6016
Decrease of VAT rate	2.0289	7.6057	*-1.2727*	*0.2801*	1.9070	6.7329
Attractive pricing	*-0.0335*	*0.9671*	*-0.0343*	*0.9663*	-0.0686	0.9337
Size of previous price change if positive	-0.6037	0.5468	-2.6733	0.0690	0.4042	1.4981
Size of previous price change if negative	0.1470	1.1584	0.1660	1.1806	-0.5259	0.5910
Previous price change negative	0.4553	1.5766	1.1084	3.0295	-0.4844	0.6161
Non-standard "Taylor contract" < 1 year	0.1191	1.1265	0.1663	1.1809	*0.0781*	*1.0812*
"Taylor contract" 1 month	0.1095	1.1157	0.1709	1.1864	0.2455	1.2783
"Taylor contract" 6 months	*-0.0025*	*0.9975*	*0.0328*	*1.0334*	*-0.0115*	*0.9885*
"Taylor contract" 12 months	0.8846	2.4220	0.9475	2.5793	*0.4772*	*1.6116*
"Taylor contract" > 1 year	0.4053	1.4998	0.4462	1.5624	*0.3854*	*1.4702*
Year 1990	-0.1540	0.8573	*-0.1102*	*0.8957*	*-0.1034*	*0.9018*
Year 1991	0.2891	1.3352	0.4154	1.5150	0.2520	1.2866
Year 1992	0.2370	1.2674	0.3394	1.4041	0.2360	1.2662
Year 1993	*0.1016*	*1.1069*	0.2296	1.2581	*0.1496*	*1.1614*
Year 1994	*0.0664*	*1.0686*	0.2186	1.2443	*0.0992*	*1.1043*
Year 1995	0.1581	1.1713	0.3118	1.3659	*0.1518*	*1.1639*
Year 1996	0.1680	1.1829	0.3398	1.4047	*0.1413*	*1.1518*
Year 1997	*0.0312*	*1.0316*	*0.1473*	*1.1587*	*0.0767*	*1.0797*
Year 1998	*0.0901*	*1.0942*	0.2602	1.2972	*0.1244*	*1.1325*
Year 1999	0.1515	1.1636	0.2608	1.2980	0.2783	1.3209
Year 2000	*0.0121*	*1.0122*	*0.0999*	*1.1051*	*0.1597*	*1.1732*
January	0.1555	1.1682	0.1072	1.1132	0.2015	1.2232
February	0.2522	1.2869	0.1889	1.2079	0.3003	1.3503
March	0.2780	1.3205	0.1467	1.1580	0.4113	1.5088
April	0.1099	1.1162	*-0.0210*	*0.9792*	0.2680	1.3073
May	0.0801	1.0834	*-0.0036*	*0.9964*	0.1626	1.1766
June	*0.0361*	*1.0368*	*-0.0392*	*0.9615*	0.1086	1.1147
July	*-0.0420*	*0.9588*	*-0.0423*	*0.9586*	*-0.0613*	*0.9405*
August	0.0885	1.0926	*0.0052*	*1.0052*	0.1828	1.2006
September	0.1773	1.1940	0.1241	1.1321	0.2457	1.2785
October	0.2053	1.2279	0.1213	1.1290	0.3322	1.3940
November	*0.0708*	*1.0733*	*-0.0617*	*0.9402*	0.2520	1.2866
Unprocessed food	-	-	-	-	-	-
Processed food	-	-	-	-	-	-
Energy	-	-	-	-	-	-
Non-energy industrial goods	-	-	-	-	-	-
Duration of first uncensored price spell	-0.0425	0.9584	-0.0364	0.9643	-0.0566	0.9450
Size of first price change observed	*0.0109*	*1.0109*	*-0.0032*	*0.9968*	*0.0249*	*1.0253*
Unprocessed food - su2	-	-	-	-	-	-
Processed food - su2	0.5686	-	0.4965	-	0.8322	-
Energy - su2	-	-	-	-	-	-
Non-energy industrial goods - su2	-	-	-	-	-	-
Services - su2	-	-	-	-	-	-
Likelihood ratio	-57765	-	-38518	-	-31424	-

Table 9 - Estimated results - "Taylor" model - Non-energy industrial goods

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-2.7509	-	-3.0041	-	-4.0024	-
ln(Dur)	-0.1376	0.8714	-0.2388	0.7876	-0.1714	0.8425
Dur1	0.3135	1.3682	0.3048	1.3564	*0.1242*	*1.1322*
Dur6	0.2294	1.2578	0.3003	1.3503	*0.1149*	*1.1218*
Dur12	0.8208	2.2723	0.8866	2.4269	0.7229	2.0604
Dur18	*0.0900*	*1.0941*	*0.0540*	*1.0554*	*0.1668*	*1.1815*
Dur24	0.4784	1.6135	0.5190	1.6803	*0.3996*	*1.4912*
Dur36	*0.3439*	*1.4104*	*0.3111*	*1.3649*	*0.3674*	*1.4440*
Accumulated inflation in absolute terms	0.1003	1.1055	-	-	-	-
Positive accumulated inflation	-	-	0.1377	1.1476	*0.0103*	*1.0104*
Negative accumulated inflation	-	-	*0.0030*	*1.0030*	0.1617	1.1755
Increase of VAT rate	0.7527	2.1227	0.8319	2.2977	0.4902	1.6326
Decrease of VAT rate	3.2118	24.8237	1.3788	3.9701	4.0333	56.4469
Attractive pricing	-0.0895	0.9144	*-0.0692*	*0.9332*	-0.2014	0.8176
Size of previous price change if positive	*0.2468*	*1.2799*	-0.7973	0.4505	1.3274	3.7712
Size of previous price change if negative	1.0903	2.9752	2.4139	11.1775	-1.4581	0.2327
Previous price change negative	0.2966	1.3453	0.4586	1.5819	*0.0882*	*1.0922*
Non-standard "Taylor contract" < 1 year	0.2818	1.3255	0.2746	1.3160	0.3079	1.3606
"Taylor contract" 1 month	0.3761	1.4566	0.3854	1.4702	0.4434	1.5580
"Taylor contract" 6 months	1.0019	2.7235	0.9549	2.5984	1.0459	2.8460
"Taylor contract" 12 months	1.0830	2.9535	1.0202	2.7737	0.9864	2.6816
"Taylor contract" > 1 year	*0.1370*	*1.1468*	*0.3562*	*1.4279*	*-0.5211*	*0.5939*
Year 1990	*-0.0342*	*0.9664*	*-0.0211*	*0.9792*	*-0.0644*	*0.9377*
Year 1991	*0.1863*	*1.2048*	*0.1520*	*1.1642*	*0.2975*	*1.3465*
Year 1992	*-0.0247*	*0.9756*	*-0.1610*	*0.8513*	*0.2707*	*1.3109*
Year 1993	*-0.1906*	*0.8265*	-0.2702	0.7632	*0.1032*	*1.1087*
Year 1994	-0.2646	0.7675	-0.3697	0.6909	*0.1067*	*1.1126*
Year 1995	-0.4429	0.6422	-0.5190	0.5951	*-0.0997*	*0.9051*
Year 1996	-0.3414	0.7108	-0.5770	0.5616	*0.2460*	*1.2789*
Year 1997	-0.5278	0.5899	-0.6788	0.5072	*0.0364*	*1.0371*
Year 1998	*-0.1843*	*0.8317*	-0.3122	0.7318	*0.3251*	*1.3842*
Year 1999	-0.2981	0.7422	-0.3996	0.6706	*0.2502*	*1.2843*
Year 2000	-0.3806	0.6835	-0.4235	0.6548	*0.0611*	*1.0630*
January	0.5382	1.7129	0.5420	1.7194	0.4339	1.5433
February	0.8569	2.3558	0.8314	2.2965	0.7257	2.0662
March	0.6359	1.8887	0.5356	1.7085	0.6724	1.9589
April	0.4474	1.5642	0.3206	1.3780	0.5465	1.7272
May	0.3998	1.4915	0.2405	1.2719	0.5664	1.7619
June	*0.0788*	*1.0820*	*-0.0344*	*0.9662*	0.2212	1.2476
July	*0.0352*	*1.0358*	*-0.0683*	*0.9340*	*0.1549*	*1.1675*
August	0.6164	1.8522	0.5636	1.7570	0.6030	1.8276
September	0.7181	2.0505	0.6995	2.0127	0.6091	1.8388
October	0.5248	1.6901	0.4338	1.5431	0.6409	1.8982
November	0.3877	1.4736	0.3193	1.3762	0.4553	1.5766
Unprocessed food	-	-	-	-	-	-
Processed food	-	-	-	-	-	-
Energy	-	-	-	-	-	-
Non-energy industrial goods	-	-	-	-	-	-
Duration of first uncensored price spell	-0.0418	0.9591	-0.0330	0.9676	-0.0568	0.9448
Size of first price change observed	*0.2362*	*1.2664*	*0.1538*	*1.1663*	0.4674	1.5958
Unprocessed food - su2	-	-	-	-	-	-
Processed food - su2	-	-	-	-	-	-
Energy - su2	-	-	-	-	-	-
Non-energy industrial goods - su2	0.9283	-	0.7271	-	1.6317	-
Services - su2	-	-	-	-	-	-
Likelihood ratio	-29093	-	-20795	-	-14285	-

Table 10 - Estimated results - "Taylor" model - Services

Variable	Price changes		Price increases		Price decreases	
	Coefficient	Risk ratio	Coefficient	Risk ratio	Coefficient	Risk ratio
Constant	-1.8034	-	-1.5480	-	-29.2338	-
ln(Dur)	-0.1757	0.8389	-0.1755	0.8390	*-0.2323*	*0.7927*
Dur1	-0.5051	0.6034	-0.3300	0.7189	-1.2497	0.2866
Dur6	*0.2207*	*1.2469*	*0.1786*	*1.1955*	*0.3642*	*1.4394*
Dur12	1.5867	4.8876	1.6493	5.2033	*0.6358*	*1.8885*
Dur18	*-0.1802*	*0.8351*	*-0.2006*	*0.8182*	*-0.1261*	*0.8815*
Dur24	1.3817	3.9817	1.4870	4.4238	-	-
Dur36	1.1480	3.1519	1.2398	3.4549	-	-
Accumulated inflation in absolute terms	0.1163	1.1234	-	-	-	-
Positive accumulated inflation	-	-	0.1141	1.1209	*0.0256*	*1.0259*
Negative accumulated inflation	-	-	*-0.1359*	*0.8729*	1.1172	3.0563
Increase of VAT rate	1.4415	4.2270	1.4351	4.2001	*0.7592*	*2.1366*
Decrease of VAT rate	*1.2282*	*3.4151*	-	-	4.0663	58.3407
Attractive pricing	-0.3865	0.6794	-0.4443	0.6413	*-0.3541*	*0.7018*
Size of previous price change if positive	*-0.1197*	*0.8872*	*-0.3726*	*0.6889*	*0.1377*	*1.1476*
Size of previous price change if negative	*0.1455*	*1.1566*	*0.1302*	*1.1391*	-7.5589	0.0005
Previous price change negative	*0.1192*	*1.1266*	0.3652	1.4408	*-0.3177*	*0.7278*
Non-standard "Taylor contract" < 1 year	0.4596	1.5834	0.5474	1.7288	*0.1752*	*1.1915*
"Taylor contract" 1 month	*0.2962*	*1.3447*	*0.2116*	*1.2357*	*0.4004*	*1.4924*
"Taylor contract" 6 months	1.8102	6.1117	1.8890	6.6128	-	-
"Taylor contract" 12 months	2.2164	9.1742	2.1251	8.3737	2.6946	14.7996
"Taylor contract" > 1 year	*0.4405*	*1.5535*	*0.5232*	*1.6874*	-	-
Year 1990	0.6125	1.8450	*0.3786*	*1.4602*	24.8833	-
Year 1991	*0.1883*	*1.2072*	*-0.1194*	*0.8875*	25.0336	-
Year 1992	*0.0838*	*1.0874*	*-0.2799*	*0.7559*	24.8887	-
Year 1993	*0.1013*	*1.1066*	*-0.2865*	*0.7509*	25.3889	-
Year 1994	*-0.1094*	*0.8964*	*-0.4770*	*0.6206*	24.6757	-
Year 1995	*-0.2871*	*0.7504*	-0.6720	0.5107	25.0508	-
Year 1996	-0.7851	0.4561	-1.1629	0.3126	24.0979	-
Year 1997	*-0.4332*	*0.6484*	-1.1654	0.3118	25.7602	-
Year 1998	*-0.6319*	*0.5316*	-1.0342	0.3555	24.6570	-
Year 1999	*-0.6091*	*0.5438*	-0.9201	0.3985	22.8930	-
Year 2000	-0.7610	0.4672	-1.1344	0.3216	24.5444	-
January	*-0.0034*	*0.9966*	*-0.1581*	*0.8538*	*0.2251*	*1.2524*
February	-0.4806	0.6184	-0.5420	0.5816	*-0.4988*	*0.6073*
March	-0.6896	0.5018	-0.7467	0.4739	-0.7390	0.4776
April	-0.7875	0.4550	-0.9533	0.3855	*-0.4954*	*0.6093*
May	-1.0095	0.3644	-0.9647	0.3811	-1.5967	0.2026
June	-1.4456	0.2356	-1.4063	0.2450	-1.8424	0.1584
July	-1.2389	0.2897	-1.2039	0.3000	-1.3659	0.2552
August	-0.6903	0.5014	-0.6510	0.5215	-0.9535	0.3854
September	-0.8261	0.4378	-0.7855	0.4559	-1.1861	0.3054
October	-0.6238	0.5359	-0.5441	0.5804	-1.2867	0.2762
November	-0.6702	0.5116	-0.7042	0.4945	*-0.3477*	*0.7063*
Unprocessed food	-	-	-	-	-	-
Processed food	-	-	-	-	-	-
Energy	-	-	-	-	-	-
Non-energy industrial goods	-	-	-	-	-	-
Duration of first uncensored price spell	-0.0330	0.9675	-0.0249	0.9754	-0.1416	0.8680
Size of first price change observed	-4.2554	0.0142	-4.2524	0.0142	*-1.6087*	*0.2001*
Unprocessed food - su2	-	-	-	-	-	-
Processed food - su2	-	-	-	-	-	-
Energy - su2	-	-	-	-	-	-
Non-energy industrial goods - su2	-	-	-	-	-	-
Services - su2	1.0728	-	0.7873	-	6.2011	-
Likelihood ratio	-5189	-	-4683	-	-1041	-

Table 11 - Contribution of time- and state-dependency to the likelihood

Price changes

Specification	Full CPI	Core components	Processed food	Non-energy industrial goods	Services
Constrained 1	-220681.8	-109715.3	-65752.8	-35263.0	-6657.9
Constrained 2	-158404.1	-96231.4.0	-59787.2	-30386.4	-6092.5
Constrained 3	-156451.0	-94997.5	-59070.5	-29789.5	-5503.5
Constrained 4	-155257.5	-93995.5	-58431.0	-29360.0	-5361.0
Baseline	-154028.0	-93116.0	-57784.0	-29177.0	-5285.0
Pseudo R² (baseline)	30.20	15.13	12.12	17.26	20.62
Contribution of					
Heterogeneity	93.43	81.23	74.86	80.13	41.18
Seasonality and truncation	2.93	7.43	8.99	9.81	42.90
Accumulated inflation	1.84	5.30	8.12	3.01	5.54

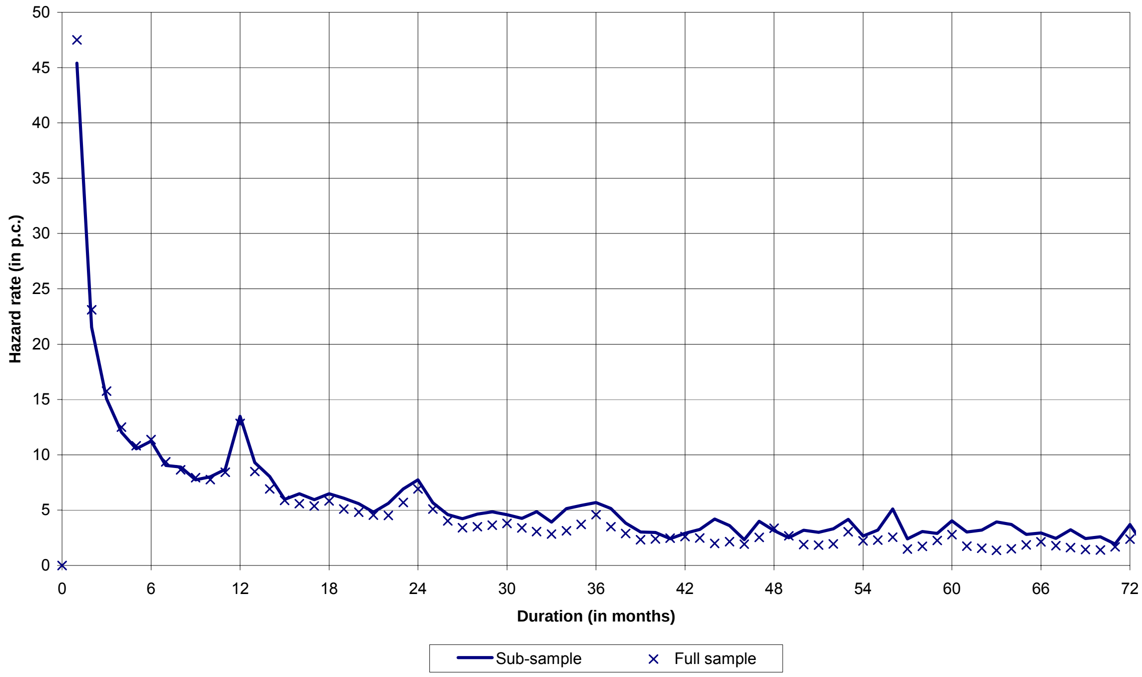
Price increases

Specification	Full CPI	Core components	Processed food	Non-energy industrial goods	Services
Constrained 1	-149209.0	-76313.4	-45165.6	-24355.7	-5899.3
Constrained 2	-115569.0	-70261.0	-42638.0	-22036.0	-5576.0
Constrained 3	-113075.0	-68736.0	-41567.5	-21557.5	-5013.5
Constrained 4	-110451.0	-66808.0	-39949.5	-21105.5	-4850.2
Baseline	-108573.0	-65306.0	-38541.0	-20857.0	-4774.0
Pseudo R² (baseline)	27.23	14.42	14.67	14.37	19.08
Contribution of					
Heterogeneity	82.78	54.98	38.15	66.30	28.73
Seasonality and truncation	6.14	13.85	16.16	13.68	49.99
Accumulated inflation	4.62	13.65	21.26	7.10	6.77

Price decreases

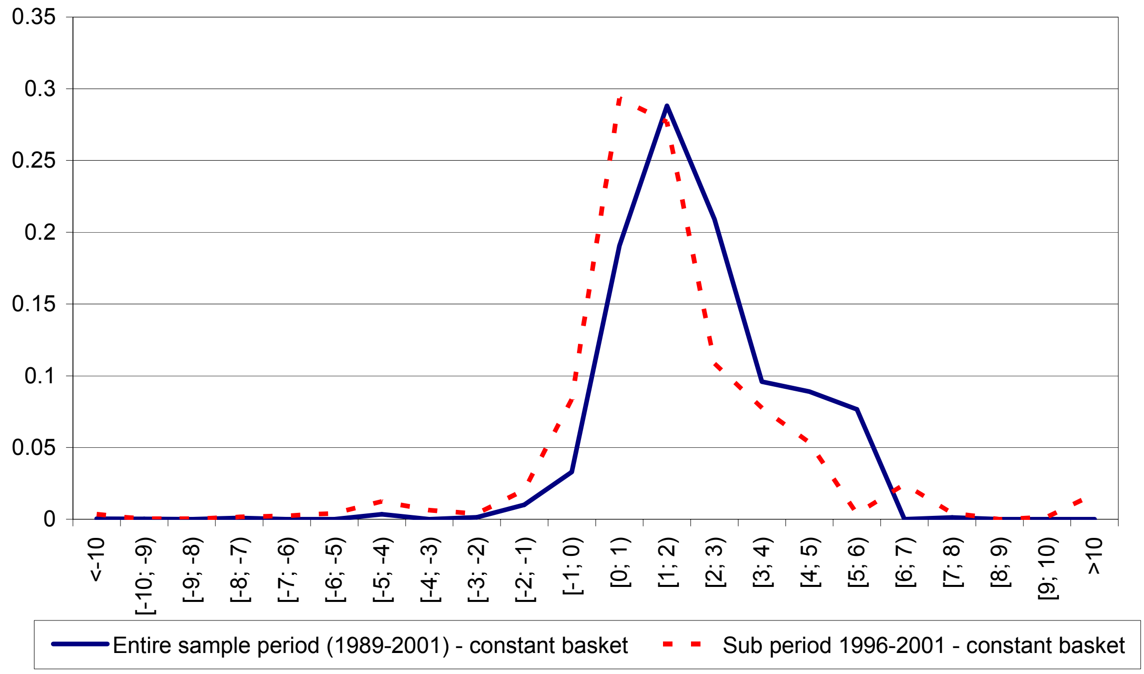
Specification	Full CPI	Core components	Processed food	Non-energy industrial goods	Services
Constrained 1	-128277.0	-56698.1	-35855.7	-17802.2	-1482.4
Constrained 2	-89229.0	-48669.0	-32511.0	-14925.0	-1173.0
Constrained 3	-88525.5	-48256.0	-32277.5	-14746.0	-1102.3
Constrained 4	-87688.5	-47688.5	-31899.5	-14458.5	-1070.2
Baseline	-86805.0	-47133.0	-31427.0	-14327.0	-1052.0
Pseudo R² (baseline)	32.33	16.87	12.35	19.52	29.03
Contribution of					
Heterogeneity	94.16	83.94	75.52	82.79	71.83
Seasonality and truncation	1.70	4.32	5.27	5.15	16.48
Accumulated inflation	2.13	5.81	10.67	3.78	4.24

Figure 1 - Kaplan Meier estimates of the probability to observe a price change after T periods



Source: NBB

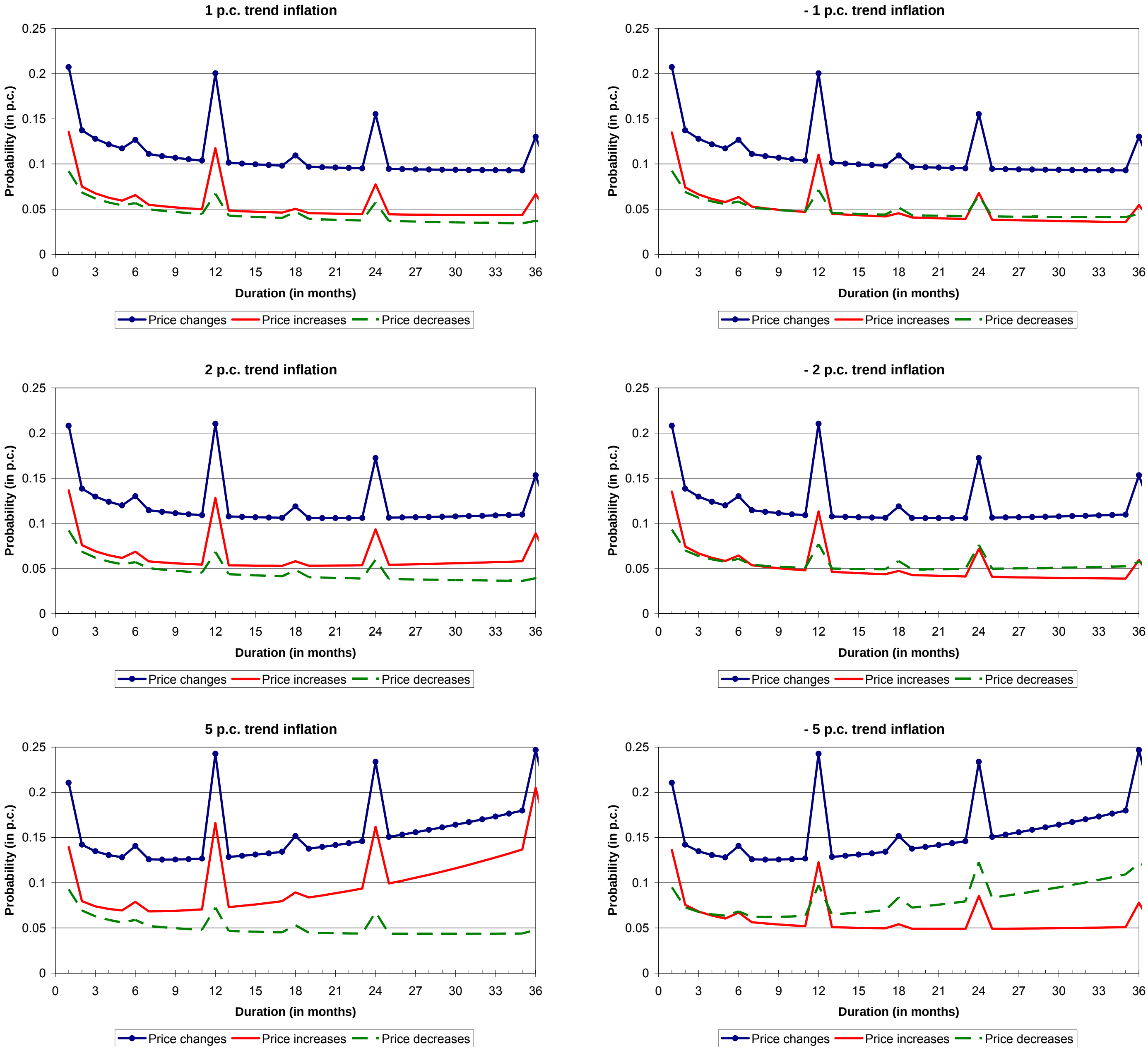
Figure 2 - Distribution of the average annual inflation rate per product category[21]



Source: NBB

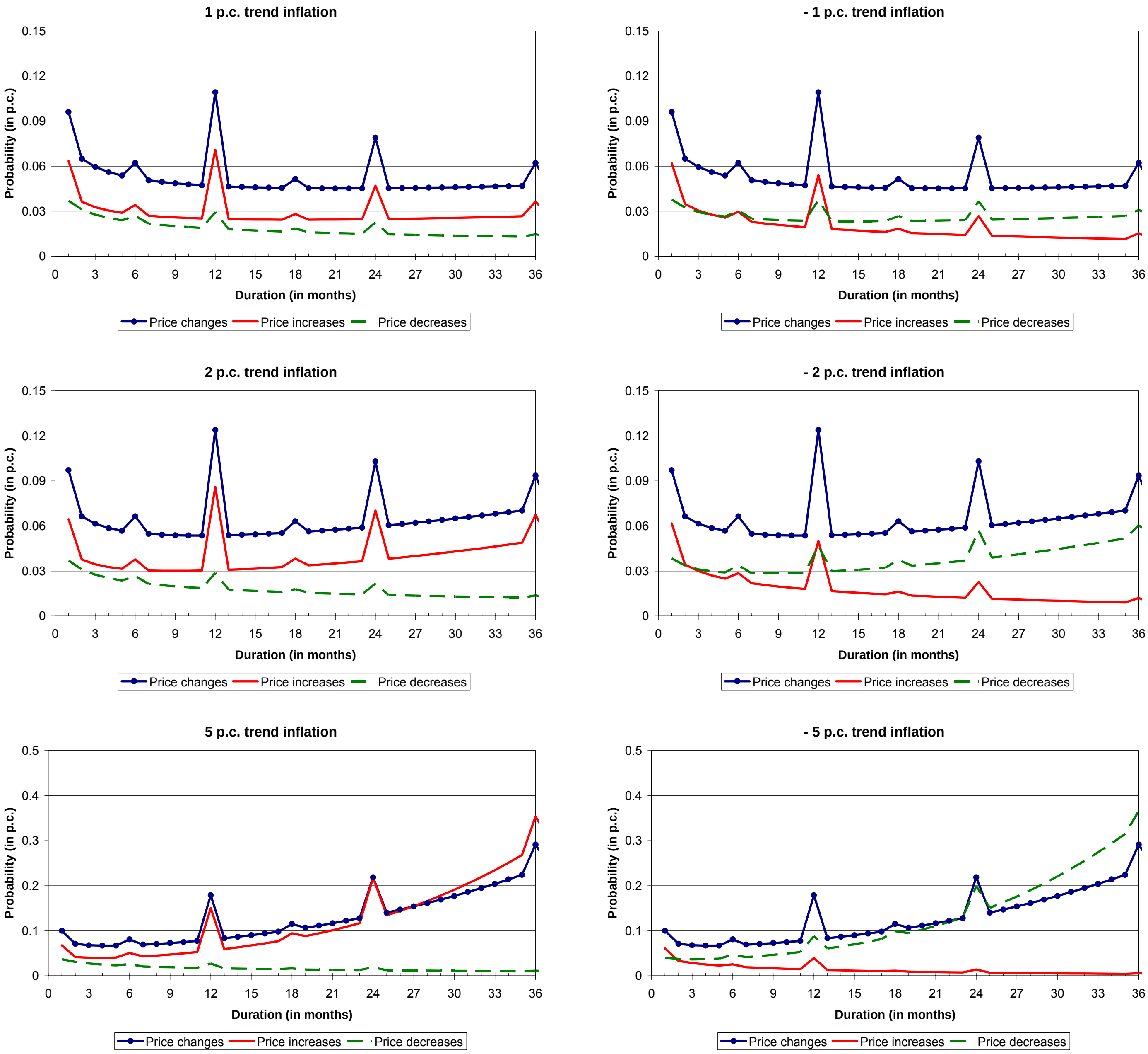
[21] The distribution of average annual inflation rate computed over the Jan 1989- Jan 2001 period refers to the distribution of the average annual inflation, considering product categories that were continuously observed during that period. The distribution of average annual inflation rate computed over the Jan 1996 - Jan 2001 period refers to the distribution of the average annual inflation considering product categories belonging to the CPI basket - Base year 1996 = 100.

Figure 3 - Hazard functions - Full CPI sample



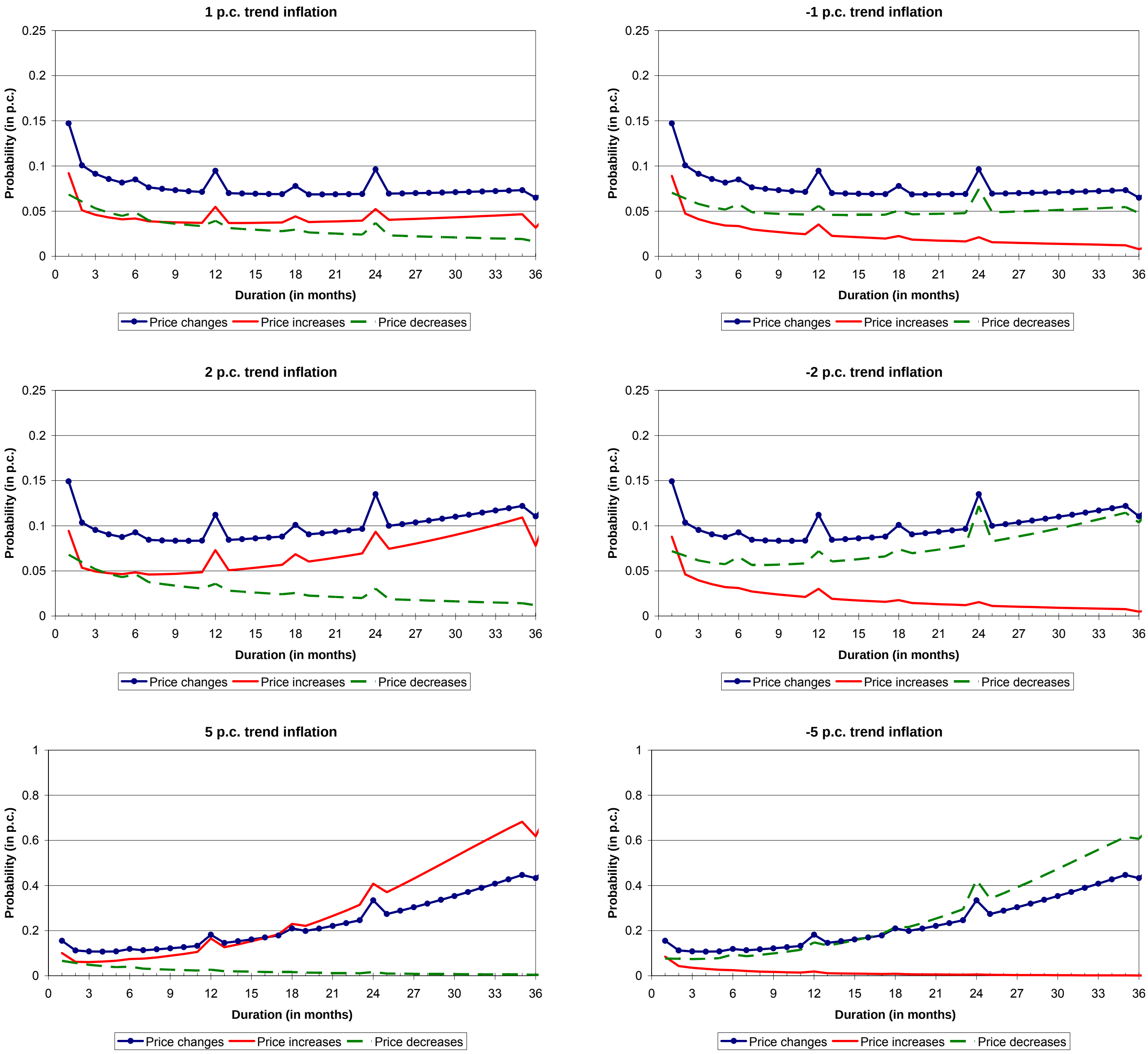
Source: NBB

Figure 4 - Hazard functions - Core components sample



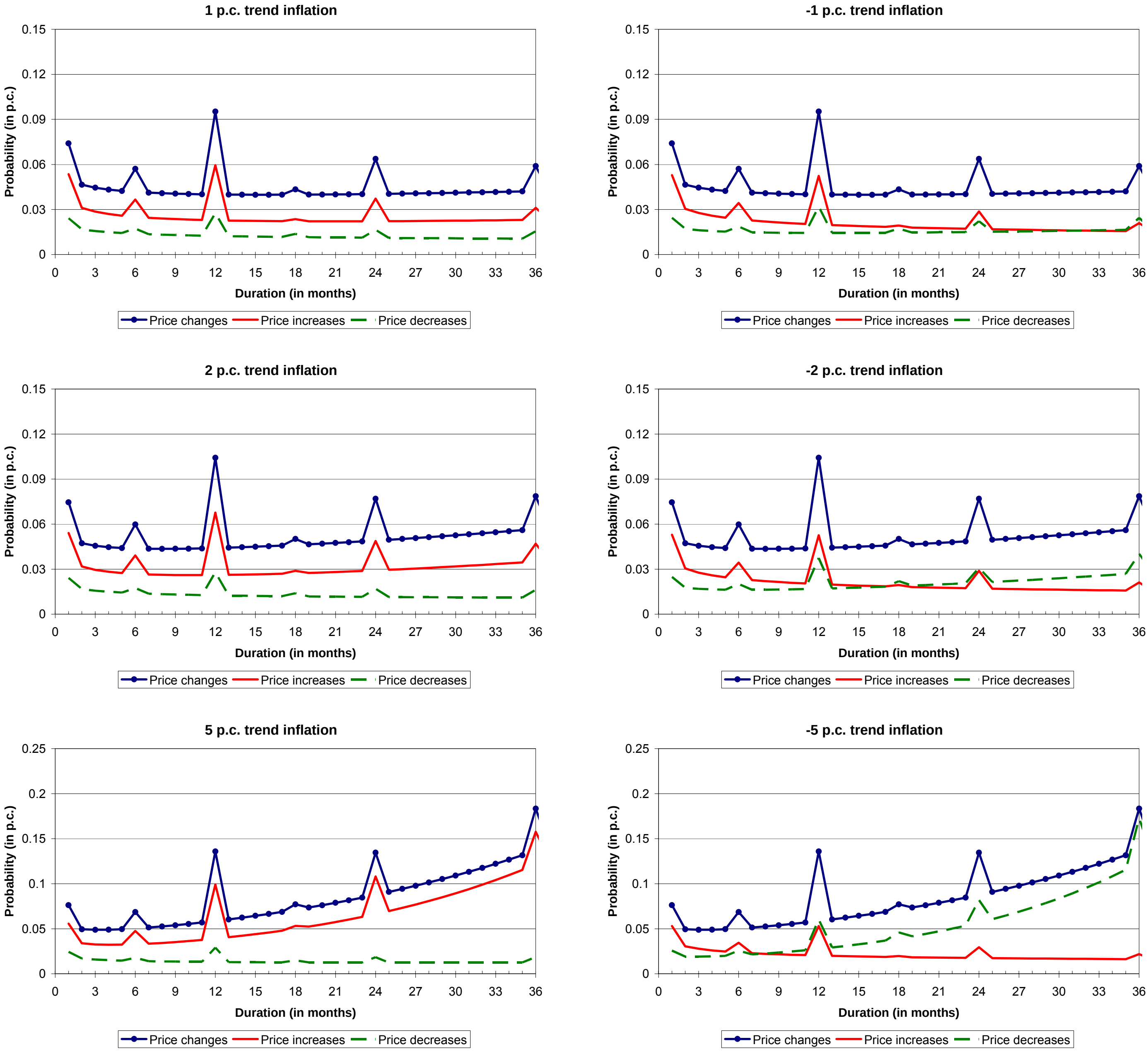
Source: NBB

Figure 5 - Hazard functions - Processed food sample



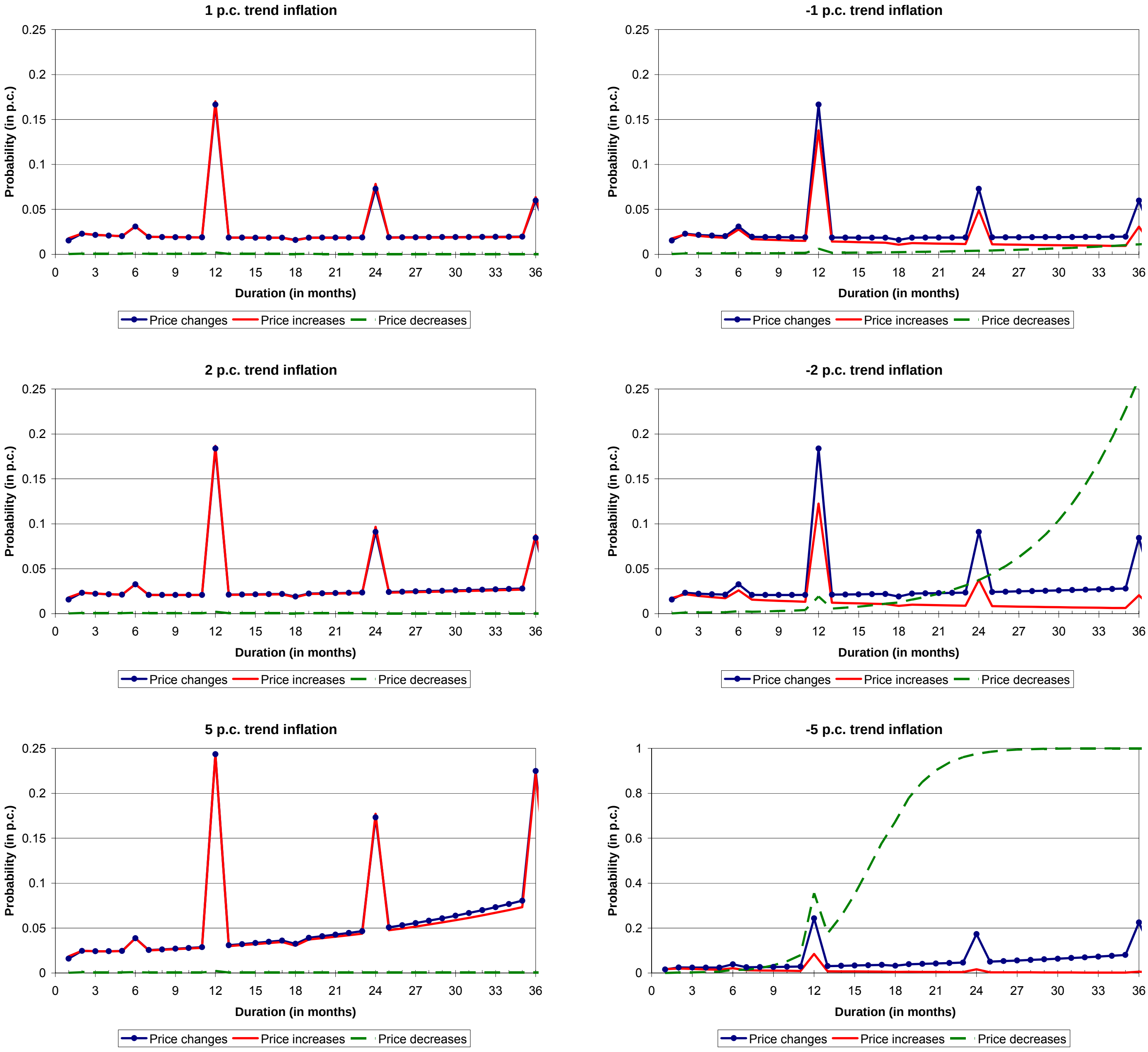
Source: NBB

Figure 6 - Hazard functions - Non-energy industrial goods sample



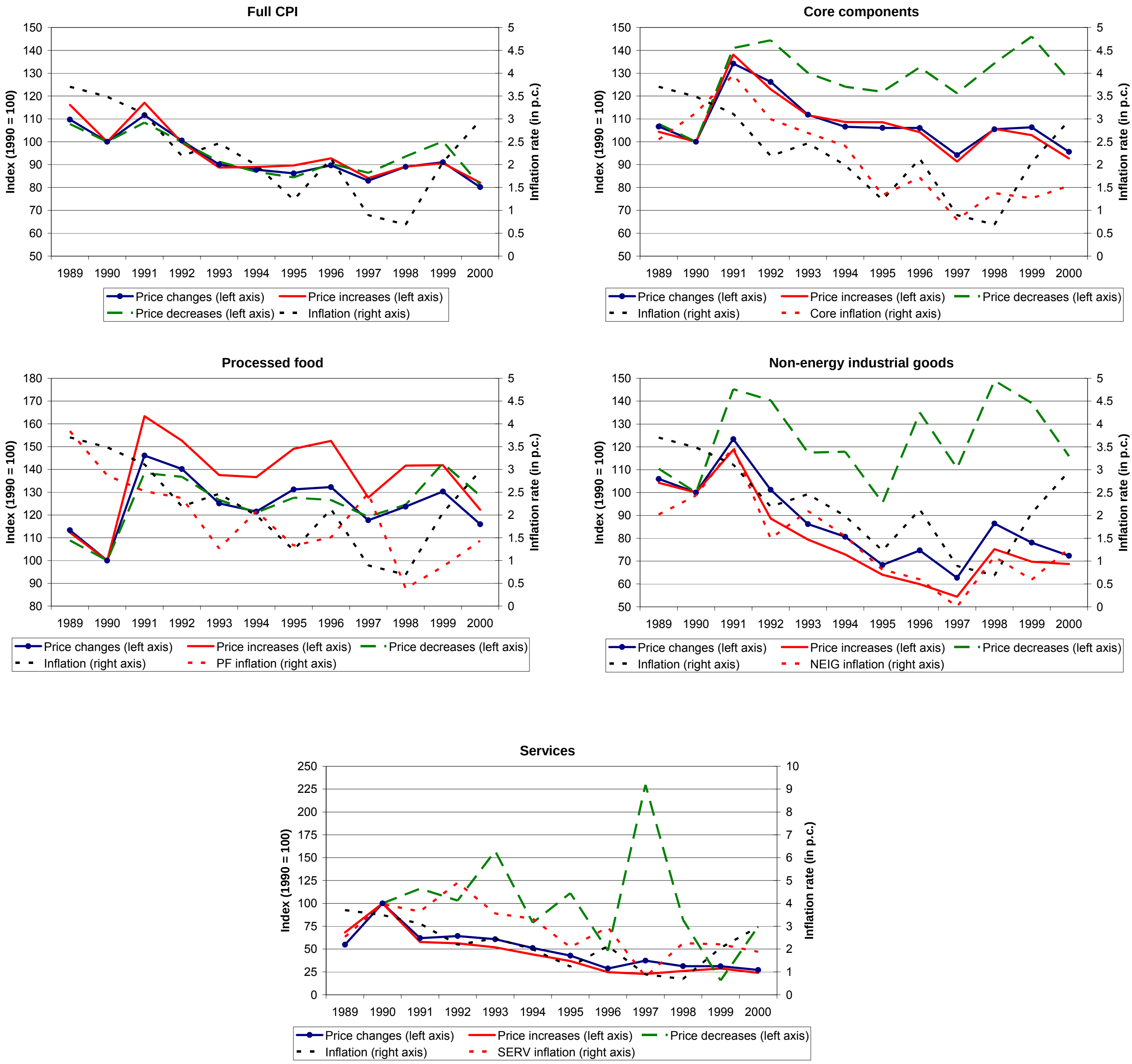
Source: NBB

Figure 7 - Hazard functions - Services sample



Source: NBB

Figure 8 - Time trends of the baseline hazards



Source: NBB

Figure 9 - Seasonal patterns of the baseline hazards

Source: NBB

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